MONTANA-DAKOTA
UTILITIES CO.
A Division of MDU Resources Group, Inc.

400 North Fourth Street
Bismarck, ND 58501
(701) 222-7900

July 23, 2013

Mr. Robert Nelson
Montana Consumer Counsel
111 North Last Chance Gulch, Suite 1B
PO Box 201703
Helena, MT 59620-1703

Re: General Gas Rate Application
Docket No. D2012.9.100

Dear Mr. Nelson:

Montana-Dakota Utilities Co. electronically submits its responses to the Montana Consumer Counsel's data requests dated July 12, 2013. Responses to request numbers MCC-194 through MCC-212 are attached.

Sincerely,

Rita A. Mulkern

Rita A. Mulkern
Director of Regulatory Affairs

Attachments
cc: Service List

Montana-Dakota Utilities Co.
Docket No. D2012.9.100
Service List

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MONTANA-DAKOTA UTILITIES CO.
MONTANA CONSUMER COUNSEL
MCC DATA REQUEST
DATED JULY 12, 2013
DOCKET NO. D2012.9.100

MCC-194 **Regarding: Returns on Natural Gas Utilities**
Witness: J. Stephen Gaske

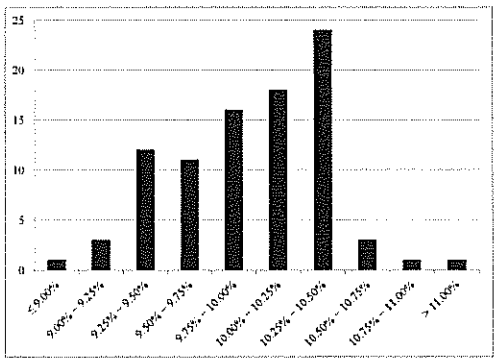
Please provide an electronic copy, with all source documents, including data and formulas, of Figure 1: Authorized Returns on Equity for Natural Gas Distribution Utilities (2010-2012), as shown at page 3 of your rebuttal testimony. Also, please provide the name of the utility, jurisdiction, requested and awarded ROE and date of filing and of Commission Order for each of the 90 rate proceedings shown on the chart.

Response:

Please see Attachment A.

Company Name	State	Requested Return on Equity (%)	Authorized Return on Equity (%)	Initial Filing Date (mm/dd/yyyy)	Rate Case Completion Date (mm/dd/yyyy)
Pacific Gas and Electric Company	CA	NA	NA	12/21/2009	5/13/2011
San Diego Gas & Electric Co.	CA	11.00	10.30	4/20/2012	12/20/2012
Southern California Gas Company	CA	10.90	10.10	4/20/2012	12/20/2012
Pacific Gas and Electric Company	CA	11.00	10.40	4/20/2012	12/20/2012
Mountaineer Gas Company	WV	10.89	NA	6/1/2009	3/19/2010
Central Hudson Gas & Electric Corporation	NY	10.00	10.00	7/31/2009	6/16/2010
New York State Electric & Gas Corporation	NY	11.43	10.00	9/17/2009	9/16/2010
Rochester Gas and Electric Corporation	NY	11.43	10.00	9/17/2009	9/16/2010
Consolidated Edison Company of New York, Inc.	NY	10.80	9.60	11/6/2009	9/16/2010
New Mexico Gas Company, Inc.	NM	11.25	10.00	3/25/2011	1/31/2012
Mountaineer Gas Company	WV	11.25	9.90	11/4/2011	10/31/2012
Atmos Energy Corporation	KY	11.00	NA	10/29/2009	5/28/2010
Louisville Gas and Electric Company	KY	11.50	NA	1/29/2010	7/30/2010
Delta Natural Gas Company, Inc.	KY	12.00	10.40	4/23/2010	10/21/2010
Louisville Gas and Electric Company	KY	11.00	10.25	6/29/2012	12/20/2012
Baltimore Gas and Electric Company	MD	11.65	9.56	5/7/2010	12/6/2010
Washington Gas Light Company	MD	10.45	9.60	4/15/2011	11/14/2011
Avista Corporation	ID	10.90	NA	3/23/2010	9/21/2010
Avista Corporation	ID	10.90	NA	7/5/2011	9/30/2011
Missouri Gas Energy	MO	11.25	10.00	4/2/2009	2/10/2010
Empire District Gas Company	MO	11.30	NA	6/5/2009	1/20/2010
Laclede Gas Company	MO	11.13	NA	12/4/2009	8/18/2010
Liberty Energy (Midstates) Corp	MO	10.90	NA	12/28/2009	8/18/2010
Union Electric Company	MO	10.50	NA	6/11/2010	1/19/2011
Columbia Gas of Virginia, Incorporated	VA	11.50	10.10	5/3/2010	12/17/2010
Washington Gas Light Company	VA	10.00	10.00	8/4/2010	4/21/2011
Washington Gas Light Company	VA	10.50	9.75	1/31/2011	7/2/2012
Virginia Natural Gas, Inc.	VA	10.95	10.00	2/8/2011	12/20/2011
Columbia Gas of Virginia, Incorporated	VA	NA	NA	6/1/2011	11/28/2011
Columbia Gas of Pennsylvania, Inc.	PA	11.70	NA	1/28/2010	8/18/2010
DTE Gas Company	MI	11.25	11.00	6/9/2009	6/3/2010
Consumers Energy Company	MI	11.00	10.55	5/22/2009	5/17/2010
SEMCO Energy, Inc.	MI	11.00	10.35	6/29/2010	1/6/2011
Consumers Energy Company	MI	11.00	10.50	8/13/2010	5/26/2011
Consumers Energy Company	MI	10.50	10.30	9/2/2011	6/7/2012
DTE Gas Company	MI	11.00	10.50	4/20/2012	12/20/2012
Northern Indiana Public Service Company	IN	11.75	NA	5/3/2010	11/4/2010
ONEOK, Inc.	OK	NA	NA	3/1/2012	7/19/2012
Wisconsin Electric Power Company	WI	10.40	10.40	3/23/2012	11/28/2012
Wisconsin Gas LLC	WI	10.50	10.50	3/23/2012	11/28/2012
Chattanooga Gas Company	TN	11.00	10.05	11/16/2009	5/24/2010
North Shore Gas Company	IL	11.87	10.33	2/13/2009	1/21/2010
Peoples Gas Light and Coke Company	IL	11.87	10.23	2/13/2009	1/21/2010
Ameren Illinois Company	IL	11.20	9.40	6/5/2009	4/29/2010
Ameren Illinois Company	IL	10.80	9.19	6/5/2009	4/29/2010
Ameren Illinois Company	IL	11.20	9.40	6/5/2009	4/29/2010
MidAmerican Energy Company	IL	11.25	10.13	6/2/2009	3/24/2010
Questar Gas Company	UT	10.60	10.35	12/3/2009	4/8/2010
Sierra Pacific Power Company	NV	10.80	10.10	6/1/2010	12/20/2010
Yankee Gas Services Company	CT	10.10	8.83	1/7/2011	6/29/2011
Delmarva Power & Light Company	DE	11.00	10.00	7/2/2010	6/21/2011
Atmos Energy Corporation	KS	11.40	NA	1/29/2010	7/30/2010
SourceGas Distribution LLC	CO	12.00	10.00	6/24/2010	12/1/2010
Public Service Company of Colorado	CO	10.80	10.10	12/17/2010	9/1/2011
Piedmont Natural Gas Company, Inc.	TN	11.25	10.20	9/2/2011	1/23/2012
North Shore Gas Company	IL	10.85	9.45	2/15/2011	1/10/2012
Peoples Gas Light and Coke Company	IL	10.85	9.45	2/15/2011	1/10/2012
Ameren Illinois Company	IL	10.75	9.06	2/18/2011	1/10/2012
Atmos Energy Corporation	TN	11.00	10.10	6/22/2012	11/8/2012
Southwest Gas Corporation	NV	10.65	9.30	4/4/2012	10/31/2012
Southwest Gas Corporation	NV	10.65	10.00	4/4/2012	10/31/2012
Atmos Energy Corporation	KS	10.90	NA	1/26/2012	8/22/2012
ONEOK, Inc.	KS	10.75	NA	5/18/2012	12/5/2012
South Carolina Electric & Gas Co.	SC	NA	NA	6/15/2011	10/13/2011
Northern Utilities, Inc.	ME	10.50	NA	5/6/2011	11/8/2011
South Carolina Electric & Gas Co.	SC	NA	NA	6/15/2012	10/11/2012
Cheyenne Light, Fuel and Power Company	WY	10.50	9.60	12/1/2011	6/18/2012
SourceGas Distribution LLC	WY	12.00	9.92	2/26/2010	12/23/2010
Atmos Energy Corporation	GA	11.00	10.70	10/1/2009	3/31/2010
Atlanta Gas Light Company	GA	11.00	10.75	5/3/2010	11/3/2010
Madison Gas and Electric Company	WI	10.40	10.30	4/22/2010	1/12/2011
Madison Gas and Electric Company	WI	10.30	10.30	3/23/2012	11/9/2012
Northern States Power Company - WI	WI	10.75	10.40	6/1/2011	12/22/2011
Northern States Power Company - WI	WI	10.40	10.40	6/1/2012	12/14/2012
Narragansett Electric Company	RI	10.75	9.50	4/27/2012	12/20/2012
Wisconsin Electric Power Company	WI	NA	NA	5/26/2011	10/6/2011
Wisconsin Gas LLC	WI	NA	NA	5/26/2011	10/6/2011
Wisconsin Power and Light Company	WI	10.40	10.40	5/3/2012	6/15/2012
Wisconsin Public Service Corporation	WI	11.25	10.30	4/1/2010	1/13/2011
Wisconsin Public Service Corporation	WI	10.30	10.30	3/30/2012	10/24/2012
NorthWestern Corporation	MT	10.90	10.25	10/16/2009	12/9/2010
Northern States Power Company - MN	MN	10.60	10.09	11/12/2009	12/6/2010
Minnesota Energy Resources Corporation	MN	10.75	9.70	11/30/2010	5/24/2012
CenterPoint Energy Resources Corp.	MN	11.00	10.24	11/3/2008	1/11/2010
Southwest Gas Corporation	AZ	11.00	9.50	11/12/2010	12/13/2011
UNS Gas, Inc.	AZ	11.00	9.50	11/7/2008	4/1/2010
UNS Gas, Inc.	AZ	10.50	9.75	4/8/2011	4/24/2012
South Jersey Gas Company	NJ	11.50	10.30	1/15/2010	9/16/2010
Public Service Electric and Gas Company	NJ	11.25	10.30	5/29/2009	6/18/2010
Atmos Energy Corporation	TX	11.05	NA	6/6/2012	10/2/2012
Atmos Energy Corporation	TX	11.50	10.40	4/24/2009	1/26/2010
CenterPoint Energy Resources Corp.	TX	11.25	10.50	7/31/2009	2/23/2010
Texas Gas Service Company	TX	11.00	10.33	6/3/2010	12/14/2010
CenterPoint Energy Resources Corp.	TX	11.00	10.05	12/3/2010	4/18/2011
Atmos Energy Corporation	TX	11.05	10.50	5/31/2012	12/4/2012

0.00	9.00	1	1%	≤ 9.00%
9.00	9.25	3	3%	9.00% – 9.25%
9.25	9.50	12	13%	9.25% – 9.50%
9.50	9.75	11	12%	9.50% – 9.75%
9.75	10.00	16	18%	9.75% – 10.00%
10.00	10.25	18	20%	10.00% – 10.25%
10.25	10.50	24	27%	10.25% – 10.50%
10.50	10.75	3	3%	10.50% – 10.75%
10.75	11.00	1	1%	10.75% – 11.00%
11.00	20.00	1	1%	> 11.00%
		90		



Company Name	State	Requested Return on Equity (%)	Authorized Return on Equity (%)	Initial Filing Date (mm/dd/yyyy)	Rate Case Completion Date (mm/dd/yyyy)
CenterPoint Energy Resources Corp.	TX	11.00	10.00	7/2/2012	12/4/2012
SourceGas Distribution LLC	NE	12.10	9.60	7/2/2009	3/9/2010
Black Hills Nebraska Gas Utility Company LLC	NE	11.50	10.10	12/1/2009	8/17/2010
SourceGas Distribution LLC	NE	10.62	9.60	9/30/2011	5/22/2012
PECO Energy Company	PA	11.75	NA	3/31/2010	12/16/2010
Peoples Natural Gas Company LLC	PA	11.50	NA	10/28/2010	6/9/2011
UGI Central Penn Gas, Inc.	PA	11.60	NA	1/14/2011	8/11/2011
Columbia Gas of Pennsylvania, Inc.	PA	11.60	NA	1/14/2011	10/14/2011
Peoples Natural Gas Company LLC	PA	11.25	NA	2/28/2012	9/27/2012
Black Hills Iowa Gas Utility Company, LLC	IA	11.25	NA	6/8/2010	2/10/2011
Interstate Power and Light Company	IA	10.90	10.00	5/25/2012	11/26/2012
ENSTAR Natural Gas Company	AK	14.75	12.55	6/1/2009	8/9/2010
Puget Sound Energy, Inc.	WA	10.10	NA	10/1/2010	3/15/2011
Avista Corporation	OR	10.90	10.10	9/30/2010	3/10/2011
Puget Sound Energy, Inc.	WA	10.80	10.10	5/8/2009	4/2/2010
Avista Corporation	WA	10.90	10.20	3/23/2010	11/19/2010
Avista Corporation	WA	10.90	NA	5/16/2011	12/16/2011
Puget Sound Energy, Inc.	WA	10.75	9.80	6/13/2011	5/7/2012
Avista Corporation	WA	10.90	9.80	4/2/2012	12/26/2012
Northwest Natural Gas Company	OR	10.30	9.50	12/30/2011	10/26/2012
Boston Gas Company	MA	11.30	9.75	4/16/2010	11/2/2010
Colonial Gas Company	MA	11.30	9.75	4/16/2010	11/2/2010
EnergyNorth Natural Gas, Inc.	NH	11.20	NA	2/26/2010	3/10/2011
Northern Utilities, Inc.	NH	10.50	9.50	5/4/2011	4/24/2012
New England Gas Company	MA	10.65	9.45	9/16/2010	3/31/2011
Fitchburg Gas and Electric Light Company	MA	10.70	9.20	1/14/2011	8/1/2011
Bay State Gas Company	MA	11.75	9.45	4/13/2012	11/1/2012

Source: Regulatory Research Associates

MONTANA-DAKOTA UTILITIES CO.
MONTANA CONSUMER COUNSEL
MCC DATA REQUEST
DATED JULY 12, 2013
DOCKET NO. D2012.9.100

MCC-195 **Regarding: Exhibit No._(JSG-04)**
Witness: J. Stephen Gaske

Please provide copies of the source documents used for preparing Exhibit No._(JSG-04).

Response:

Please see Attachment A.

South Jersey Industries, Inc.
Part II

	2011	2010	2009
Operating Income (Loss):			
Gas Utility Operations	$ 102,663	$ 90,700	$ 81,439
Wholesale Energy Operations	16,225	6,154	29,601
Retail Energy Operations:			
Retail Gas and Other Operations	478	1,065	78
Retail Electric Operations	1,755	13,051	(7,459)
On-Site Energy Production	2,254	4,063	5,309
Appliance Service Operations	(1,539)	432	1,361
Subtotal Retail Energy Operations	2,948	18,611	(711)
Corporate and Services	(229)	1,027	781
Total Operating Income	$ 121,607	$ 116,492	$ 111,110
Depreciation and Amortization:			
Gas Utility Operations	$ 41,959	$ 39,502	$ 34,507
Wholesale Energy Operations	253	252	304
Retail Energy Operations:			
Retail Gas and Other Operations	33	35	21
On-Site Energy Production	4,527	3,920	3,804
Appliance Service Operations	322	344	306
Subtotal Retail Energy Operations	4,882	4,299	4,131
Corporate and Services	665	641	505
Total Depreciation and Amortization	$ 47,759	$ 44,694	$ 39,447
Interest Charges:			
Gas Utility Operations	$ 18,922	$ 17,641	$ 16,442
Wholesale Energy Operations	54	94	339
Retail Energy Operations:			
Retail Gas and Other Operations	186	172	16
On-Site Energy Production	5,436	3,898	1,928
Subtotal Retail Energy Operations	5,622	4,070	1,944
Corporate and Services	2,648	1,076	780
Subtotal	27,246	22,881	19,505
Intersegment Borrowings	(3,168)	(985)	(513)
Total Interest Charges	$ 24,078	$ 21,896	$ 18,992
Income Taxes:			
Gas Utility Operations	$ 34,281	$ 29,676	$ 27,104
Wholesale Energy Operations	7,517	3,243	12,456
Retail Energy Operations:			
Retail Gas and Other Operations	511	678	20
Retail Electric Operations	717	5,353	(3,064)
On-Site Energy Production	(22,258)	(10,635)	(3,138)
Appliance Service Operations	1,480	174	601
Subtotal Retail Energy Operations	(19,550)	(4,430)	(5,581)
Corporate and Services	254	322	323
Total Income Taxes	$ 22,502	$ 28,811	$ 34,302

	2011	2010	2009
Property Additions:			
Gas Utility Operations	$ 141,918	$ 111,653	$ 110,694
Wholesale Energy Operations	83	682	27
Retail Energy Operations:			
Retail Gas and Other Operations	22	31	16
On-Site Energy Production	57,093	50,427	9,760
Appliance Service Operations	254	141	577
Subtotal Retail Energy Operations	57,369	50,599	10,353
Corporate and Services	605	1,029	290
Total Property Additions	$ 199,975	$ 163,963	$ 121,364

	2011	2010
Identifiable Assets:		
Gas Utility Operations	$ 1,615,723	$ 1,468,635
Wholesale Energy Operations	159,424	237,978
Retail Energy Operations:		
Retail Gas and Other Operations	14,659	21,532
Retail Electric Operations	31,225	37,383
On-Site Energy Production	275,053	147,064
Appliance Service Operations	11,335	18,528
Subtotal Retail Energy Operations	332,272	224,507
Discontinued Operations	224	890
Corporate and Services	156,252	168,788
Intersegment Assets	(16,385)	(24,183)
Total Identifiable Assets	$ 2,247,510	$ 2,076,615

9. LEASES:

The Company is considered to be the lessor of certain thermal energy generating property and equipment under an operating lease which expires in May 2027 . As of December 31, 2011 and 2010 the carrying costs of this property and equipment under operating lease was $74.0 million and $73.6 million , respectively, (net of accumulated depreciation of $17.5 million and $15.0 million , respectively) and is included in Nonutility Property and Equipment in the consolidated balance sheets.

Minimum future rentals to be received on non-cancelable leases as of December 31, 2011 for each of the next five years and in the aggregate are (in thousands):

Year ended December 31,	
2012	5,396
2013	5,396
2014	5,396
2015	5,396
2016	5,396
Thereafter	56,210
Total minimum future rentals	83,190

Minimum future rentals do not include additional amounts to be received based on actual use of the leased property.

MONTANA-DAKOTA UTILITIES CO.
MONTANA CONSUMER COUNSEL
MCC DATA REQUEST
DATED JULY 12, 2013
DOCKET NO. D2012.9.100

MCC-196 **Regarding: Earnings Growth Projections**
Witness: J. Stephen Gaske

Please provide a copy of the document referenced in footnote 7 of your rebuttal testimony: Vander Weide, J.H. and Carleton, W.T., "Investor Growth Expectations: Analysts vs. History," *The Journal of Portfolio Management*, Spring 1988.

Response:

Please see Attachment A.

Investor growth expectations: Analysts vs. history

Analysts' growth forecasts dominate past trends in predicting stock prices.

James H. Vander Weide and Willard T. Carleton

For the purposes of implementing the Discounted Cash Flow (DCF) cost of equity model, the analyst must know which growth estimate is embodied in the firm's stock price. A study by Cragg and Malkiel (1982) suggests that the stock valuation process embodies analysts' forecasts rather than historically based growth figures such as the ten-year historical growth in dividends per share or the five-year growth in book value per share. The Cragg and Malkiel study is based on data for the 1960s, however, a decade that was considerably more stable than the recent past.

As the issue of which growth rate to use in implementing the DCF model is so important to applications of the model, we decided to investigate whether the Cragg and Malkiel conclusions continue to hold in more recent periods. This paper describes the results of our study.

STATISTICAL MODEL

The DCF model suggests that the firm's stock price is equal to the present value of the stream of dividends that investors expect to receive from owning the firm's shares. Under the assumption that investors expect dividends to grow at a constant rate, g, in perpetuity, the stock price is given by the following simple expression:

$$P_s = \frac{D(1+g)}{k-g} \tag{1}$$

where:

P_s = current price per share of the firm's stock;

D = current annual dividend per share;

g = expected constant dividend growth rate; and

k = required return on the firm's stock.

Dividing both sides of Equation (1) by the firm's current earnings, E, we obtain:

$$\frac{P_s}{E} = \frac{D}{E} \cdot \frac{(1+g)}{k-g} \tag{2}$$

Thus, the firm's price/earnings (P/E) ratio is a nonlinear function of the firm's dividend payout ratio (D/E), the expected growth in dividends (g), and the required rate of return.

To investigate what growth expectation is embodied in the firm's current stock price, it is more convenient to work with a linear approximation to Equation (2). Thus, we will assume that:

$$P/E = a_0(D/E) + a_1 g + a_2 k. \tag{3}$$

(Cragg and Malkiel found this assumption to be reasonable throughout their investigation.)

Furthermore, we will assume that the required

JAMES H. VANDER WEIDE is Research Professor at the Fuqua School of Business at Duke University in Durham (NC 27706). WILLARD T. CARLETON is Karl Eller Professor of Finance at the University of Arizona in Tucson (AZ 85721).
Financial support for this project was provided by BellSouth and Pacific Telesis. The authors wish to thank Paul Blalock at BellSouth, Mohan Gyani at Pacific Telesis, Bill Keck at Southern Bell, and John Carlson, their programmer, for help with this project.

rate of return, k, in Equation (3) depends on the values of the risk variables B, Cov, Rsq, and Sa, where B is the firm's Value Line beta; Cov is the firm's pretax interest coverage ratio; Rsq is a measure of the stability of the firm's five-year historical EPS; and Sa is the standard deviation of the consensus analysts' five-year EPS growth forecast for the firm. Finally, as the linear form of the P/E equation is only an approximation to the true P/E equation, and B, Cov, Rsq, and Sa are only proxies for k, we will add an error term, e, that represents the degree of approximation to the true relationship.

With these assumptions, the final form of our P/E equation is as follows:

$$P/E = a_0(D/E) + a_1g + a_2B + a_3Cov + a_4Rsq + a_5Sa + e. \quad (4)$$

The purpose of our study is to use more recent data to determine which of the popular approaches for estimating future growth in the Discounted Cash Flow model is embodied in the market price of the firm's shares.

We estimated Equation (4) to determine which estimate of future growth, g, when combined with the payout ratio, D/E, and risk variables B, Cov, Rsq, and Sa, provides the best predictor of the firm's P/E ratio. To paraphrase Cragg and Malkiel, we would expect that growth estimates found in the best-fitting equation more closely approximate the expectation used by investors than those found in poorer-fitting equations.

DESCRIPTION OF DATA

Our data sets include both historically based measures of future growth and the consensus analysts' forecasts of five-year earnings growth supplied by the Institutional Brokers Estimate System of Lynch, Jones & Ryan (IBES). The data also include the firm's dividend payout ratio and various measures of the firm's risk. We include the latter items in the regression, along with earnings growth, to account for other variables that may affect the firm's stock price.

The data include:

Earnings Per Share. Because our goal is to determine which earnings variable is embodied in the firm's market price, we need to define this variable with care. Financial analysts who study a firm's financial results in detail generally prefer to "normalize" the firm's reported earnings for the effect of extraordinary items, such as write-offs of discontinued operations, or mergers and acquisitions. They also attempt, to the extent possible, to state earnings for different firms using a common set of accounting conventions.

We have defined "earnings" as the consensus analyst estimate (as reported by IBES) of the firm's earnings for the forthcoming year.[1] This definition approximates the normalized earnings that investors most likely have in mind when they make stock purchase and sell decisions. It implicitly incorporates the analysts' adjustments for differences in accounting treatment among firms and the effects of the business cycle on each firm's results of operations. Although we thought at first that this earnings estimate might be highly correlated with the analysts' five-year earnings growth forecasts, that was not the case. Thus, we avoided a potential spurious correlation problem.

Price/Earnings Ratio. Corresponding to our definition of "earnings," the price/earnings ratio (P/E) is calculated as the closing stock price for the year divided by the consensus analyst earnings forecast for the forthcoming fiscal year.

Dividends. Dividends per share represent the common dividends declared per share during the calendar year, after adjustment for all stock splits and stock dividends). The firm's dividend payout ratio is then defined as common dividends per share divided by the consensus analyst estimate of the earnings per share for the forthcoming calendar year (D/E). Although this definition has the deficiency that it is obviously biased downward — it divides this year's dividend by next year's earnings — it has the advantage that it implicitly uses a "normalized" figure for earnings. We believe that this advantage outweighs the deficiency, especially when one considers the flaws of the apparent alternatives. Furthermore, we have verified that the results are insensitive to reasonable alternative definitions (see footnote 1).

Growth. In comparing historically based and consensus analysts' forecasts, we calculated forty-one different historical growth measures. These included the following: 1) the past growth rate in EPS as determined by a log-linear least squares regression for the latest year,[2] two years, three years, . . ., and ten years; 2) the past growth rate in DPS for the latest year, two years, three years, . . ., and ten years; 3) the past growth rate in book value per share (computed as the ratio of common equity to the outstanding common equity shares) for the latest year, two years, three years, . . ., and ten years; 4) the past growth rate in cash flow per share (computed as the ratio of pretax income, depreciation, and deferred taxes to the outstanding common equity shares) for the latest year, two years, three years, . . ., and ten years; and 5) plowback growth (computed as the firm's retention ratio for the current year times the firm's latest annual return on common equity).

We also used the five-year forecast of earnings

per share growth compiled by IBES and reported in mid-January of each year. This number represents the consensus (i.e., mean) forecast produced by analysts from the research departments of leading Wall Street and regional brokerage firms over the preceding three months. IBES selects the contributing brokers "because of the superior quality of their research, professional reputation, and client demand" (IBES *Monthly Summary Book*).

Risk Variables. Although many risk factors could potentially affect the firm's stock price, most of these factors are highly correlated with one another. As shown above in Equation (4), we decided to restrict our attention to four risk measures that have intuitive appeal and are followed by many financial analysts: 1) B, the firm's beta as published by Value Line; 2) Cov, the firm's pretax interest coverage ratio (obtained from Standard & Poor's Compustat); 3) Rsq, the stability of the firm's five-year historical EPS (measured by the R^2 from a log-linear least squares regression); and 4) Sa, the standard deviation of the consensus analysts' five-year EPS growth forecast (mean forecast) as computed by IBES.

After careful analysis of the data used in our study, we felt that we could obtain more meaningful results by imposing six restrictions on the companies included in our study:

1. Because of the need to calculate ten-year historical growth rates, and because we studied three different time periods, 1981, 1982, and 1983, our study requires data for the thirteen-year period 1971-1983. We included only companies with at least a thirteen-year operating history in our study.
2. As our historical growth rate calculations were based on log-linear regressions, and the logarithm of a negative number is not defined, we excluded all companies that experienced negative EPS during any of the years 1971-1983.
3. For similar reasons, we also eliminated companies that did not pay a dividend during any one of the years 1971-1983.
4. To insure comparability of time periods covered by each consensus earnings figure in the P/E ratios, we eliminated all companies that did not have a December 31 fiscal year-end.
5. To eliminate distortions caused by highly unusual events that distort current earnings but not expected future earnings, and thus the firm's price/earnings ratio, we eliminated any firm with a price/earnings ratio greater than 50.
6. As the evaluation of analysts' forecasts is a major part of this study, we eliminated all firms that IBES did not follow.

Our final sample consisted of approximately sixty-five utility firms.[3]

RESULTS

To keep the number of calculations in our study to a reasonable level, we performed the study in two stages. In Stage 1, all forty-one historically oriented approaches for estimating future growth were correlated with each firm's P/E ratio. In Stage 2, the historical growth rate with the highest correlation to the P/E ratio was compared to the consensus analyst growth rate in the multiple regression model described by Equation (4) above. We performed our regressions for each of three recent time periods, because we felt the results of our study might vary over time.

First-Stage Correlation Study

Table 1 gives the results of our first-stage correlation study for each group of companies in each of the years 1981, 1982, and 1983. The values in this table measure the correlation between the historically oriented growth rates for the various time periods and the firm's end-of-year P/E ratio.

The four variables for which historical growth rates were calculated are shown in the left-hand column: EPS indicates historical earnings per share growth, DPS indicates historical dividend per share growth, BVPS indicates historical book value per share growth, and CFPS indicates historical cash flow per share growth. The term "plowback" refers to the product of the firm's retention ratio in the currennt year and its return on book equity for that year. In all, we calculated forty-one historically oriented growth rates for each group of firms in each study period.

The goal of the first-stage correlation analysis was to determine which historically oriented growth rate is most highly correlated with each group's year-end P/E ratio. Eight-year growth in CFPS has the highest correlation with P/E in 1981 and 1982, and ten-year growth in CFPS has the highest correlation with year-end P/E in 1983. In all cases, the plowback estimate of future growth performed poorly, indicating that — contrary to generally held views — plowback is not a factor in investor expectations of future growth.

Second-Stage Regression Study

In the second stage of our regression study, we ran the regression in Equation (4) using two different measures of future growth, g: 1) the best historically oriented growth rate (g_h) from the first-stage correlation study, and 2) the consensus analysts' forecast (g_a) of five-year EPS growth. The regression results, which are shown in Table 2, support at least

TABLE 1

Correlation Coefficients of All Historically Based Growth Estimates by Group and by Year with P/E

Historical Growth Rate Period in Years

Current Year	1	2	3	4	5	6	7	8	9	10
1981										
EPS	−0.02	0.07	0.03	0.01	0.03	0.12	0.08	0.09	0.09	0.09
DPS	0.05	0.18	0.14	0.15	0.14	0.15	0.19	0.23	0.23	0.23
BVPS	0.01	0.11	0.13	0.13	0.16	0.18	0.15	0.15	0.15	0.15
CFPS	−0.05	0.04	0.13	0.22	0.28	0.31	0.30	0.31	−0.57	−0.54
Plowback	0.19									
1982										
EPS	−0.10	−0.13	−0.06	−0.02	−0.02	−0.01	−0.03	−0.03	0.00	0.00
DPS	−0.19	−0.10	0.03	0.05	0.07	0.08	0.09	0.11	0.13	0.13
BVPS	0.07	0.08	0.11	0.11	0.09	0.10	0.11	0.11	0.09	0.09
CFPS	−0.02	−0.08	0.00	0.10	0.16	0.19	0.23	0.25	0.24	0.07
Plowback	0.04									
1983										
EPS	−0.06	−0.25	−0.25	−0.24	−0.16	−0.11	−0.05	0.00	0.02	0.02
DPS	0.03	−0.10	−0.03	0.08	0.15	0.21	0.21	0.21	0.22	0.24
BVPS	0.03	0.10	0.04	0.09	0.15	0.16	0.19	0.21	0.22	0.21
CFPS	−0.08	0.01	0.02	0.08	0.20	0.29	0.35	0.38	0.40	0.42
Plowback	−0.08									

two general conclusions regarding the pricing of equity securities.

First, we found overwhelming evidence that the consensus analysts' forecast of future growth is superior to historically oriented growth measures in predicting the firm's stock price. In every case, the R^2 in the regression containing the consensus analysts' forecast is higher than the R^2 in the regression containing the historical growth measure. The regression coefficients in the equation containing the consensus analysts' forecast also are considerably more significant than they are in the alternative regression. These results are consistent with those found by Cragg and Malkiel for data covering the period 1961-1968. Our results also are consistent with the hypothesis that investors use analysts' forecasts, rather than historically oriented growth calculations, in making stock buy–and–sell decisions.

TABLE 2

Regression Results
Model I

Part A: *Historical*

$P/E = a_0 + a_1 D/E + a_2 g_h + a_3 B + a_4 Cov + a_5 Rsq + a_6 Sa$

Year	$\hat{a}_0$	$\hat{a}_1$	$\hat{a}_2$	$\hat{a}_3$	$\hat{a}_4$	$\hat{a}_5$	$\hat{a}_6$	R^2	F Ratio
1981	−6.42*	10.31*	7.67*	3.24	0.54*	1.42*	57.43	0.83	46.49
	(5.50)	(14.79)	(2.20)	(2.86)	(2.50)	(2.85)	(4.07)		
1982	−2.90*	9.32*	8.49*	2.85	0.45*	−0.42	3.63	0.86	65.53
	(2.75)	(18.52)	(4.18)	(2.83)	(2.60)	(0.05)	(0.26)		
1983	−5.96*	10.20*	19.78*	4.85	0.44*	0.33	32.49	0.82	45.26
	(3.70)	(12.20)	(4.83)	(2.95)	(1.89)	(0.50)	(1.29)		

Part B: *Analysis*

$P/E = a_0 + a_1 D/E + a_2 g_a + a_3 B + a_4 Cov + a_5 Rsq + a_6 Sa$

Year	$\hat{a}_0$	$\hat{a}_1$	$\hat{a}_2$	$\hat{a}_3$	$\hat{a}_4$	$\hat{a}_5$	$\hat{a}_6$	R^2	F Ratio
1981	−4.97*	10.62*	54.85*	−0.61	0.33*	0.63*	4.34	0.91	103.10
	(6.23)	(21.57)	(8.56)	(0.68)	(2.28)	(1.74)	(0.37)		
1982	−2.16*	9.47*	50.71*	−1.07	0.36*	−0.31	119.05*	0.90	97.62
	(2.59)	(22.46)	(9.31)	(1.14)	(2.53)	(1.09)	(1.60)		
1983	−8.47*	11.96*	79.05*	2.16	0.56*	0.20	−34.43	0.87	69.81
	(7.07)	(16.48)	(7.84)	(1.55)	(3.08)	(0.38)	(1.44)		

Notes:
* Coefficient is significant at the 5% level (using a one-tailed test) and has the correct sign. T-statistic in parentheses.

Second, there is some evidence that investors tend to view risk in traditional terms. The interest coverage variable is statistically significant in all but one of our samples, and the stability of the operating income variable is statistically significant in six of the twelve samples we studied. On the other hand, the beta is never statistically significant, and the standard deviation of the analysts' five-year growth forecasts is statistically significant in only two of our twelve samples. This evidence is far from conclusive, however, because, as we demonstrate later, a significant degree of cross-correlation among our four risk variables makes any general inference about risk extremely hazardous.

Possible Misspecification of Risk

The stock valuation theory says nothing about which risk variables are most important to investors. Therefore, we need to consider the possibility that the risk variables of our study are only proxies for the "true" risk variables used by investors. The inclusion of proxy variables may increase the variance of the parameters of most concern, which in this case are the coefficients of the growth variables.[4]

To allow for the possibility that the use of risk proxies has caused us to draw incorrect conclusions concerning the relative importance of analysts' growth forecasts and historical growth extrapolations, we have also estimated Equation (4) with the risk variables excluded. The results of these regressions are shown in Table 3.

Again, there is overwhelming evidence that the consensus analysts' growth forecast is superior to the historically oriented growth measures in predicting the firm's stock price. The R^2 and t-statistics are higher in every case.

TABLE 3

Regression Results
Model II

Part A: *Historical*

$P/E = a_0 + a_1 D/E + a_2 g_h$

Year	$\hat{a}_0$	$\hat{a}_1$	$\hat{a}_2$	R^2	F Ratio
1981	−1.05	9.59	21.20	0.73	82.95
	(1.61)	(12.13)	(7.05)		
1982	0.54	8.92	12.18	0.83	167.97
	(1.38)	(17.73)	(6.95)		
1983	−0.75	8.92	12.18	0.77	107.82
	(1.13)	(12.38)	(7.94)		

Part B: *Analysis*

$P/E + a_0 + a_1 D/E + a_2 g_a$

Year	$\hat{a}_0$	$\hat{a}_1$	$\hat{a}_2$	R^2	F Ratio
1981	3.96	10.07	60.53	0.90	274.16
	(8.31)	(8.31)	(20.91)	(15.79)	
1982	−1.75	9.19	44.92	0.88	246.36
	(4.00)	(4.00)	(21.35)	(11.06)	
1983	−4.97	10.95	82.02	0.83	168.28
	(6.93)	(6.93)	(15.93)	(11.02)	

Notes:
* Coefficient is significant at the 5% level (using a one-tailed test) and has the correct sign. T-statistic in parentheses.

CONCLUSION

The relationship between growth expectations and share prices is important in several major areas of finance. The data base of analysts' growth forecasts collected by Lynch, Jones & Ryan provides a unique opportunity to test the hypothesis that investors rely more heavily on analysts' growth forecasts than on historical growth extrapolations in making security buy–and–sell decisions. With the help of this data base, our studies affirm the superiority of analysts' forecasts over simple historical growth extrapolations in the stock price formation process. Indirectly, this finding lends support to the use of valuation models whose input includes expected growth rates.

[1] We also tried several other definitions of "earnings," including the firm's most recent primary earnings per share prior to any extraordinary items or discontinued operations. As our results were insensitive to reasonable alternative definitions of "earnings" we report only the results for the IBES consensus.

[2] For the latest year, we actually employed a point-to-point growth calculation because there were only two available observations.

[3] We use the word "approximately," because the set of available firms varied each year. In any case, the number varied only from zero to three firms on either side of the figures cited here.

[4] See Maddala (1977).

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MCC-197 **Regarding: Growth Rates Estimates**
Witness: J. Stephen Gaske

Please provide a copy of the documents referenced in footnotes 8, 9, 10, 12 and 18 of your rebuttal testimony:

a. **F. Marston, R. Harris, and P. Crawford, "Risk and Return in Equity Markets: Evidence Using Financial Analysts' Forecasts," in Handbook of Security Analysts' Forecasting and Asset Allocation, J. Guerard and M. Gultekin (eds.), Greenwich, CT, JAI Press; as described in R. Harris and F. Marston, "Estimating Shareholder Risk Premia Using Analysts' Growth Forecasts," Financial Management, Summer 1992.**
b. **Advanced Research Center, Investor Growth Expectations, Summer, 2004.**
c. **Morin Roger T., New Regulatory Finance.**
d. **Sundaresh Ramnath, Steve Rock, and Philip Shane, A Review of Research Related to Financial Analysts' Forecasts and Stock Recommendations, Social Science Research Network, June 30, 2008.**
e. **Armen Hovakimian and Ekkachai Saenyasiri, Conflicts of Interest and Analyst Behavior: Evidence from Recent Changes in Regulation, Financial Analysts Journal, Volume 66, Number 4, July/August 2010.**

Response:

a. Please see Attachment A.
b. Please see Attachment B.
c. Please see Attachment C.
d. Please see Attachment D.
e. Please see Attachment E.

7

Estimating Shareholder Risk Premia Using Analysts' Growth Forecasts

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■ One of the most widely used concepts in finance is that shareholders require a risk premium over bond yields to bear the additional risks of equity investments. While models such as the two-parameter capital asset pricing model (CAPM) or arbitrage pricing theory offer explicit methods for varying risk premia across securities, the models are invariably linked to some underlying market (or factor-specific) risk premium. Unfortunately, the theoretical models provide limited practical advice on establishing empirical estimates of such a benchmark market risk premium. As a result, the typical advice to practitioners is to estimate the market risk premium based on historical realizations of share and bond returns (see Brealey and Myers [3]).

In this paper, we present estimates of shareholder required rates of return and risk premia which are derived using forward-looking analysts' growth forecasts. We update, through 1991, earlier work which, due to data availability, was restricted to the period 1982-1984 (Harris [12]). Using stronger tests, we also reexamine the efficacy of using such an expectational approach as an alternative to the use of historical averages. Using the S&P 500 as a proxy for the market portfolio, we find an average market risk premium (1982-1991) of 6.47% above yields on long-term U.S. government bonds and 5.13% above yields on corporate bonds. We also find that required returns for individual stocks vary directly with their risk (as proxied by beta) and that the market risk premium varies over time. In particular, the equity market premium over government bond yields is higher in low interest rate environments and when there is a larger spread between corporate and government bond yields. These findings show that, in addition to fitting the theoretical requirement of being forward-looking, the utilization of analysts' forecasts in estimating return requirements provides reasonable empirical results that can be useful in practical applications.

Section I provides background on the estimation of equity required returns and a brief discussion of related

Thanks go to Ed Buchmann, Bill Carleton, Pete Crawford, and Steve Osborn for their assistance on earlier research in this area. We thank Bell Atlantic for supplying data for this project. Financial support from the Darden Sponsors and from the Associates Program at the McIntire School of Commerce is gratefully acknowledged.

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literature on financial analysts' forecasts (FAF). In Section II, models and data are discussed. Following a comparison of the results to historical risk premia, the estimates are subjected to economic tests of both their time-series and cross-sectional characteristics in Section III. Finally, conclusions are offered in Section IV.

I. Background and Literature Review

In establishing economic criteria for resource allocation, it is often convenient to use the notion of a shareholder's required rate of return. Such a rate (k) is the minimum level of expected return necessary to compensate the investor for bearing risks and receiving dollars in the future rather than in the present. In general, k will depend on returns available on alternative investments (e.g., bonds or other equities) and the riskiness of the stock. To isolate the effects of risk, it is useful to work in terms of a risk premium (rp), defined as

$$rp = k - i, \tag{1}$$

where i = required return for a zero risk investment.[1]

Lacking a superior alternative, investigators often use averages of historical realizations to estimate a benchmark "market" risk premium which then may be adjusted for the relative risk of individual stocks (e.g., using the CAPM or a variant). The historical studies of Ibbotson Associates [13] have been used frequently to implement this approach.[2] This historical approach requires the assumptions that past realizations are a good surrogate for future expectations and, as typically applied, that risk premia are constant over time. Carleton and Lakonishok [5] demonstrate empirically some of the problems with such historical premia when they are disaggregated for different time periods or groups of firms.

As an alternative to historical estimates, the current paper derives estimates of k, and hence, implied values of rp, using publicly available expectational data. This expectational approach employs the dividend growth model (hereafter referred to as the discounted cash flow or DCF model) in which a consensus measure of financial analysts' forecasts (FAF) of earnings is used as a proxy for investor expectations. Earlier works by Malkiel [17], Brigham, Vinson, and Shome [4], and Harris [12] have used FAF in DCF models, and this approach has been employed in regulatory settings (see Harris [12]) and suggested by consultants as an alternative to use of historical data (e.g., Ibbotson Associates [13, pp. 127, 128]). Unfortunately, the published studies use data extending to 1984 at the latest. Our paper draws on this earlier work but extends it through 1991.[3] Our work is closest to that done by Harris [12], who reviews literature showing a strong link between equity prices and FAF and supporting the use of FAF as a proxy for investor expectations. Using data from 1982 to 1984, Harris' results suggest that this expectational approach to estimating equity risk premia is an encouraging alternative to the use of historical averages. He also demonstrates that such risk premia vary both cross-sectionally with the riskiness of individual stocks and over time with financial market conditions.

II. Models and Data

A. Model for Estimation

The simplest and most commonly used version of the DCF model to estimate shareholders' required rate of return, k, is shown in Equation (2):

$$k = \left(\frac{D_1}{P_0}\right) + g, \tag{2}$$

where D_1 = dividend per share expected to be received at time one, P_0 = current price per share (time 0), and g = expected growth rate in dividends per share. The limitations of this model are well known, and it is straightforward to derive expressions for k based on more general specifications of the DCF model.[4] The primary difficulty in using the DCF model is obtaining an estimate of g, since it should reflect market expectations of future perfor-

[1]Theoretically, i is a risk-free rate, though empirically its proxy (e.g., yield to maturity on a government bond) is only a "least risk" alternative that is itself subject to risk. In this development, the effects of tax codes on required returns are ignored.

[2]Many leading texts in financial management use such historical risk premia to estimate a market return. See, for example, Brealey and Myers [3]. Often a market risk premium is adjusted for the observed relative risk of a stock.

[3]See Harris [12] for a discussion of the earlier work and a detailed discussion of the approach employed here.

[4]As stated, Equation (2) requires expectations of either an infinite horizon of dividend growth at a rate g or a finite horizon of dividend growth at rate g and special assumptions about the price of the stock at the end of that horizon. Essentially, the assumption must ensure that the stock price grows at a compound rate of g over the finite horizon. One could alternatively estimate a nonconstant growth model, although the proxies for multistage growth rates are even more difficult to obtain than single stage growth estimates. Marston, Harris, and Crawford [19] examine publicly available data from 1982-1985 and find that plausible measures of risk are more closely related to expected returns derived from a constant growth model than to those derived from multistage growth models. These findings illustrate empirical difficulties in finding empirical proxies for multistage growth models for large samples.

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For [illegible]

mance. Without a ready source for measuring such expectations, application of the DCF model is fraught with difficulties. This paper uses published FAF of long-run growth in earnings as a proxy for g.

B. Data

FAF for this research come from IBES (Institutional Broker's Estimate System), which is a product of Lynch, Jones, and Ryan, a major brokerage firm.[5] Representative of industry practice, IBES contains estimates of (*i*) EPS for the upcoming fiscal years (up to five separate years), and (*ii*) a five-year growth rate in EPS. Each item is available at monthly intervals.

The mean value of individual analysts' forecasts of five-year growth rate in EPS will be used as a proxy for g in the DCF model.[6] The five-year horizon is the longest horizon over which such forecasts are available from IBES and often is the longest horizon used by analysts. IBES requests "normalized" five-year growth rates from analysts in order to remove short-term distortions that might stem from using an unusually high or low earnings year as a base.

Dividend and other firm-specific information come from COMPUSTAT. Interest rates (both government and corporate) are gathered from Federal Reserve Bulletins and *Moody's Bond Record*. Exhibit 1 describes key variables used in the study. Data collected cover all dividend paying stocks in the Standard & Poor's 500 stock (S&P 500) index, plus approximately 100 additional stocks of regulated companies. Since five-year growth rates are first available from IBES beginning in 1982, the analysis covers the 113-month period from January 1982 to May 1991.

III. Risk Premia and Required Rates of Return

A. Construction of Risk Premia

For each month, a "market" required rate of return is calculated using each dividend paying stock in the S&P 500 index for which data are available. The DCF model in

Exhibit 1. Variable Definitions

k	=	Equity required rate of return.
P_0	=	Average daily price per share.
D_1	=	Expected dividend per share measured as current indicated annual dividend from COMPUSTAT multiplied by $(1 + g)$.[a]
g	=	Average financial analysts' forecast of five-year growth rate in earnings per share (from IBES).
i_{lt}	=	Yield to maturity on long-term U.S. government obligations (source: Federal Reserve Bulletin, constant maturity series).
i_c	=	Yield to maturity on long-term corporate bonds: Moody's average.[b]
rp	=	Equity risk premium calculated as $rp = k - i$.
β	=	beta, calculated from CRSP monthly data over 60 months.

Notes:
[a]See footnote 7 for a discussion of the $(1 + g)$ adjustment.
[b]The average corporate bond yield across bond rating categories as reported by Moody's. See *Moody's Bond Survey* for a brief description and the latest published list of bonds included in the bond rating categories.

Equation (2) is applied to each stock and the results weighted by market value of equity to produce the market required return.[7] The return is converted to a risk premium

[5]Harris [12] provides a discussion of IBES data and its limitations. In more recent years, IBES has begun collecting forecasts for each of the next five years. Since this work was completed, the FAF used here have become available from IBES Inc., now a subsidiary of CitiBank.

[6]While the model calls for expected growth in dividends, no source of data on such projections is readily available. In addition, in the long run, dividend growth is sustainable only via growth in earnings. As long as payout ratios are not expected to change, the two growth rates will be the same.

[7]The construction of D_1 is controversial since dividends are paid quarterly and may be expected to change during the year; whereas, Equation (2), as is typical, is being applied to annual data. Both the quarterly payment of dividends (due to investors' reinvestment income before year's end, see Linke and Zumwalt [15]) and any growth during the year require an upward adjustment of the current annual rate of dividends to construct D_1. If quarterly dividends grow at a constant rate, both factors could be accommodated straightforwardly by applying Equation (2) to quarterly data with a quarterly growth rate and then annualizing the estimated quarterly required return. Unfortunately, with lumpy changes in dividends, the precise nature of the adjustment depends on both an individual company's pattern of growth during the calendar year and an individual company's required return (and hence reinvestment income in the risk class).
In this work, D_1 is calculated as D_0 $(1 + g)$. The full g adjustment is a crude approximation to adjust for both growth and reinvestment income. For example, if one expected dividends to have been raised, on average, six months ago, a "1/2 g" adjustment would allow for growth, and the remaining "1/2 g" would be justified on the basis of reinvestment income. Any precise accounting for both reinvestment income and growth would require tracking each company's dividend change history and making explicit judgments about the quarter of the next change. Since no organized "market" forecast of such a detailed nature exists, such a procedure is not possible. To get a feel for the magnitudes involved, during the sample period the dividend yield (D_1/P_0) and growth (market value weighted) for the S&P 500 were typically 4% to 6% and 11% to 13%, respectively. As a result, a "full g" adjustment on average increases the required return by 60 to 70 basis points (relative to no g adjustment).

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Exhibit 2. Bond Market Yields, Equity Required Return, and Equity Risk Premium,[a] 1982-1991

	Bond Market Yields[b]		Equity Market Required Return[c]	Equity Risk Premium	
Year	(1) U.S. Gov't	(2) Moody's Corporates	(3) S&P 500	U.S. Gov't (3) - (1)	Moody's Corporates (3) - (2)
1982	12.92	14.94	20.08	7.16	5.14
1983	11.34	12.78	17.89	6.55	5.11
1984	12.48	13.49	17.26	4.78	3.77
1985	10.97	12.05	16.32	5.37	4.28
1986	7.85	9.71	15.09	7.24	5.38
1987	8.58	9.84	14.71	6.13	4.86
1988	8.96	10.18	15.37	6.41	5.19
1989	8.46	9.66	15.06	6.60	5.40
1990	8.61	9.77	15.69	7.08	5.92
1991[d]	8.21	9.41	15.61	7.40	6.20
Average[e]	9.84	11.18	16.31	6.47	5.13

Notes:
[a]Values are averages of monthly figures in percent.
[b]Yields to maturity.
[c]Required return on value weighted S&P 500 index using Equation (1).
[d]Figures for 1991 are through May.
[e]Months weighted equally.

over government bonds by subtracting i_{ft}, the yield to maturity on long-term government bonds. A risk premium over corporate bond yields is also constructed by subtracting i_{ct}, the yield on long-term corporate bonds. Exhibit 2 reports the results by year (averages of monthly data).

The results are quite consistent with the patterns reported earlier (i.e., Harris [12]). The estimated risk premia in Exhibit 2 are positive, consistent with equity owners demanding additional rewards over and above returns on debt securities. The average expectational risk premium (1982 to 1991) over government bonds is 6.47%, only slightly higher than the 6.16% average for 1982 to 1984 reported earlier (Harris [12]). Furthermore, Exhibit 2 shows the estimated risk premia change over time, suggesting changes in the market's perception of the incremental risk of investing in equity rather than debt securities.

For comparison purposes, Exhibit 3 contains historical returns and risk premia. The average expectational risk premium reported in Exhibit 2 falls roughly midway between the arithmetic (7.5%) and geometric (5.7%) long-term differentials between returns on stocks and long-term government bonds. Note, however, that the expectational risk premia appear to change over time. In the following sections, we examine the estimated risk premia to see if they vary cross-sectionally with the risk of individual stocks and over time with financial market conditions.

B. Cross-Sectional Tests

Earlier, Harris [12] conducted crude tests of whether expectational equity risk premia varied with risk proxied by bond ratings and the dispersion of analysts' forecasts and found that required returns increased with higher risk. Here we examine the link between these premia and beta, perhaps the most commonly used measure of risk for equities.[8] In keeping with traditional work in this area, we adopt the methodology introduced by Fama and Macbeth [9] but replace realized returns with expected returns from Equation (2) as the variable to be explained. For this portion of our tests, we restrict our sample to 1982-1987

[8]For other efforts using expectational data in the context of the two-parameter CAPM, see Friend, Westerfield, and Granito [10], Cragg and Malkiel [7], Marston, Crawford, and Harris [19], Marston and Harris [20], and Linke, Kannan, Whitford, and Zumwalt [16]. For a more complete treatment of the subject, see Marston and Harris [20] from which we draw some of these results. Marston and Harris also investigate the role of unsystematic risk and the difference in estimates found when using expected versus realized returns.

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Exhibit 3. Average Historical Returns on Bonds, Stocks, Bills, and Inflation in the U.S., 1926-1989

Historical Return Realizations	Geometric	Arithmetic
Common stock	10.3%	12.4%
Long-term government bonds	4.6%	4.9%
Long-term corporate bonds	5.2%	5.5%
Treasury bills	3.6%	3.7%
Inflation rate	3.1%	3.2%

Source: Ibbotson Associates, Inc., *1990 Stocks, Bonds, Bills and Inflation*, 1990 Yearbook.

and in any month include firms that have at least three forecasts of earnings growth to reduce measurement error associated with individual forecasts.[9] This restricted sample still consists of, on average, 399 firms for each of the 72 months (or 28,744 company months).

For a given company in a given month, beta is estimated via the market model (using ordinary least squares) on the prior 60 months of return data taken from CRSP. Beta estimates are updated monthly and are calculated against an equally weighted index of all NYSE securities. For each month, we aggregate firms into 20 portfolios (consisting of approximately 20 securities each). The advantage of grouped data is the reduction in potential measurement error inherent in independent variables at the company level. Portfolios are formed based on a ranking of beta estimated from a prior time period (t = -61 to t = -120). Portfolio expected returns and beta are calculated as the simple averages for the individual securities.

Using these data, we estimate the following model for each of the 72 months:

$$R_p = \alpha_0 + \alpha_1 \beta_p + u_p, \quad p = 1...20, \qquad (3)$$

where:

R_p = Expected return for portfolio p in the given month,

β_p = Portfolio beta, estimated over 60 prior months, and

u_p = A random error term with mean zero.

As a result of estimating regression (3) for each month, 72 estimates of each coefficient (α_0 and α_1) are obtained. Using realized returns as the dependent variable, the traditional approach (e.g., Fama and Macbeth [9]) is to assume that realized returns are a fair game. Given this assumption, the mean of the 72 values of each coefficient is an unbiased estimate of the mean over that same time period if one could have actually used expected returns as the dependent variable. Note that if expected returns are used as the dependent variable the fair-game assumption is not required. Making the additional assumption that the true value of the coefficient is constant over the 72 months, a test of whether the mean coefficient is different from zero is performed using a t-statistic where the denominator is the standard error of the 72 values of the coefficient. This is the technique employed by Fama and Macbeth [9]. If one assumes the CAPM is correct, the coefficient α_1 is an empirical estimate of the market risk premium, which should be positive.

To test the sensitivity of the results, we also repeat our procedures using individual security returns rather than portfolios. To account, at least in part, for differences in precision of coefficient estimates in different months we also report results in which monthly parameter estimates are weighted inversely by the standard error of the coefficient estimate rather than being weighted equally (following Chan, Hamao, and Lakonishok [6]).

Exhibit 4 shows that there is a significant positive link between expectational required returns and beta. For instance, in Panel A, the mean coefficient of 2.78 on beta is significantly different from zero at better than the 0.001 level (t = 35.31), and each of the 72 monthly coefficients going into this average is positive (as shown by that 100% positive figure). Using individual stock returns, the significant positive link between beta and expected return remains, though it is smaller in magnitude than for portfolios.[10] Comparison of Panels A and B shows that the results are not sensitive to the weighting of monthly coefficients.

While the findings in Exhibit 4 suggest a strong positive link between beta and risk premia (a result often not supported when realized returns are used as a proxy for expectations; e.g., see Tinic and West [22]), the results do not support the predictions of a simple CAPM. In particular, the intercept is higher than a proxy for the risk-free rate over the sample period and the coefficient of beta is well below estimates of a market risk premium obtained from either expectational (Exhibit 2) or historical data (Exhibit

[9]Firms for which the standard deviation of individual FAF exceeded 20 in any month were excluded since we suspect some of these involve errors in data entry. This screen eliminated very few companies in any month. The 1982-1987 period was chosen due to the availability of data on betas.

[10]The smaller coefficients on beta using individual stock portfolio returns are likely due in part to the higher measurement error in measuring individual stock versus portfolio betas.

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Exhibit 4. Mean Values of Monthly Parameter Estimates for the Relationship Between Required Returns and Beta for Both Portfolios and Individual Securities (Figures in Parentheses are *t* Values and Percent Positive), 1982-1987

	Intercept	B	Adjusted R^2 [c]	F[c]
Panel A. Equal Weighting[a]				
Portfolio returns	14.06 (54.02, 100)	2.78 (35.31, 100)	0.503	25.4
Security returns	14.77 (58.10, 100)	1.91 (16.50, 99)	0.080	39.0
Panel B. Weighted by Standard Errors[b]				
Portfolio returns	13.86 (215.6, 100)	2.67 (35.80, 100)	0.503	25.4
Security returns	14.63 (398.9, 100)	1.92 (47.3, 99)	0.080	39.0

[a]Equally weighted average of monthly parameters estimated using cross-sectional data for each of the 72 months, January 1982 - December 1987.
[b]In obtaining the reported means, estimates of the monthly intercept and slope coefficients are weighted inversely by the standard error of the estimate from the cross-sectional regression for that month.
[c]Values are averages for the 72 monthly regressions.

3).[11] Nonetheless, the results show that the estimated risk premia conform to the general theoretical relationship between risk and required return that is expected when investors are risk-averse.

C. Time Series Tests — Changes in Market Risk Premia

A potential benefit of using ex ante risk premia is the estimation of changes in market risk premia over time. With changes in the economy and financial markets, equity investments may be perceived to change in risk. For instance, investor sentiment about future business conditions likely affects attitudes about the riskiness of equity investments compared to investments in the bond markets. Moreover, since bonds are risky investments themselves, equity risk premia (relative to bonds) could change due to changes in perceived riskiness of bonds, even if equities displayed no shifts in risk. For example, during the high interest rate period of the early 1980s, the high level of interest rate volatility made fixed income investments more risky holdings than they were in a world of relatively stable rates.

Studying changes in risk premia for utility stocks, Brigham, et al [4] conclude that, prior to 1980, utility risk premia increased with the level of interest rates, but that this pattern reversed thereafter, resulting in an inverse correlation between risk premia and interest rates. Studying risk premia for both utilities and the equity market generally, Harris [12] also reports that risk premia appear to change over time. Specifically, he finds that equity risk premia decreased with the level of government interest rates, increased with the increases in the spread between corporate and government bond yields, and increased with increases in the dispersion of analysts' forecasts. Harris' study is, however, restricted to the 36-month period, 1982 to 1984.

Exhibit 5 reports results of analyzing the relationship between equity risk premia, interest rates, and yield spreads between corporate and government bonds. Following Harris [12], these bond yield spreads are used as a time series proxy for equity risk. As the perceived riskiness of corporate activity increases, the difference between yields on corporate bonds and government bonds should increase. One would expect the sources of increased riskiness to corporate bonds to also increase risks to shareholders. All regressions in Exhibit 5 are corrected for serial correlation.[12]

[11]Estimation difficulties confound precise interpretation of the intercept as the risk-free rate and the coefficient on beta as the market risk premium (see Miller and Scholes [21], and Black, Jensen, and Scholes [2]). The higher than expected intercept and lower than expected slope coefficient on beta are consistent with the prior studies of Black, Jensen, and Scholes [2], and Fama and MacBeth [9] using historical returns. Such results are consistent with Black's [1] zero beta model, although alternative explanations for these findings exist as well (as noted by Black, Jensen, and Scholes [2]).

[12]Ordinary least squares regressions showed severe positive autocorrelation in many cases, with Durbin Watson statistics typically below one. Estimation used the Prais-Winsten method. See Johnston [14, pp. 321-325].

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Exhibit 5. Changes in Equity Risk Premia Over Time — Entries are Coefficient (t-value); Dependent Variable is Equity Risk Premium

Time period	Intercept	i_B	$i_C - i_B$	R^2
A. May 1991-1992	0.131	-0.651		0.53
	(19.82)	(-11.16)		
	0.092	-0.363	0.666	0.54
	(14.26)	(-6.74)	(5.48)	
B. 1982-1984	0.140	-0.637		0.43
	(8.15)	(-5.00)		
	0.064	-0.203	1.549	0.60
	(3.25)	(-1.63)	(4.84)	
C. 1985-1987	0.131	-0.739		0.74
	(7.73)	(-9.67)		
	0.110	-0.561	0.317	0.77
	(12.53)	(-7.30)	(1.87)	
D. 1988-1991	0.136	-0.793		0.68
	(16.23)	(-8.29)		
	0.130	-0.738	0.098	0.68
	(8.71)	(-4.96)	(0.40)	

Note: All variables are defined in Exhibit 1. Regressions were estimated using monthly data and were corrected for serial correlation using the Prais-Winsten method. For purposes of this regression, variables are expressed in decimal form, e.g., 14% = 0.14.

For the entire sample period, Panel A shows that risk premia are negatively related to the level of interest rates — as proxied by yields on government bonds, i_B. This negative relationship is also true for each of the subperiods displayed in Panels B through D. Such a negative relationship may result from increases in the perceived riskiness of investment in government debt at high levels of interest rates. A direct measure of uncertainty about investments in government bonds would be necessary to test this hypothesis directly.

For the entire 1982 to 1991 period, the addition of the yield spread risk proxy to the regressions dramatically lowers the magnitude of the coefficient on government bond yields, as can be seen by comparing Equations 1 and 2 of Panel A. Furthermore, the coefficient of the yield spread (0.666) is itself significantly positive. This pattern suggests that a reduction in the risk differential between investment in government bonds and in corporate activity is translated into a lower equity market risk premium. Further examination of Panels B through D, however, suggests that the yield spread variable is much more important in explaining changes in equity risk premia in the early portion of the 1980s than in the 1988 to 1991 period.

In summary, market equity risk premia change over time and appear inversely related to the level of government interest rates but positively related to the bond yield spread, which proxies for the incremental risk of investing in equities as opposed to government bonds.

IV. Conclusions

Shareholder required rates of return and risk premia are based on theories about investors' expectations for the future. In practice, however, risk premia are often estimated using averages of historical returns. This paper applies an alternate approach to estimating risk premia that employs publicly available expectational data. At least for the decade studied (1982 to 1991), the resultant average market equity risk premium over government bonds is comparable in magnitude to long-term differences (1926 to 1989) in historical returns between stocks and bonds. There is strong evidence, however, that market risk premia change over time and, as a result, use of a constant historical average risk premium is not likely to mirror changes in investor return requirements. The results also show that the expectational risk premia vary cross-sectionally with the relative risk (beta) of individual stocks.

The approach offers a straightforward and powerful aid in establishing required rates of return either for corporate investment decisions or in the regulatory arena. Since data are readily available on a wide range of equities, an investigator can analyze various proxy groups (e.g., portfolios of utility stocks) appropriate for a particular decision as well as analyze changes in equity return requirements over time.

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INVESTOR GROWTH EXPECTATIONS

Summer 2004

A study done by Vander Weide and Carleton in 1988[1] suggests that consensus analysts' forecast of future growth is superior to historically oriented growth measures in stock valuation process for domestic companies. We worked with one of the original authors of the study, Dr. James H. Vander Weide, and closely followed his suggestions and methodology to investigate whether the results still hold in more recent times (2001- 2003).

We used the following equation to determine which estimate of future growth (g) best predicts the firm's P/E ratio when combined with the dividend payout ratio, D/E, and risk variables, B, Cov, Stb, and Sa.

$$P/E = a_0(D/E) + a_1 g(\text{Growth}) + a_2 B(\text{Beta}) + a_3 Cov(\text{Interest Coverage Ratio}) + a_4 Stb(\text{Stability}) + a_5 Sa(\text{Std Dev}) + e$$

Data Description

Earnings Per Share: IBES consensus analyst estimate of the firm's earnings for the unreported year.

Price/Earnings Ratio: Closing stock price for the year divided by the consensus analyst earnings per share for the forthcoming year.

Dividends: Ratio of common dividends per share to the consensus analyst earnings forecast for the forthcoming fiscal year (D/E).

Historical Growth measures

- EPS Growth Rate: Determined by a log-linear least squares regression for the latest year, two years, three years, …, and ten years.
- Dividend per Share Growth Rate: Determined by a log-linear least squares regression for the latest year, two years, three years, …, and ten years.
- Book Value per Share Growth Rate: Common equity divided by the common shares outstanding. Determined by a log-linear least squares regression for the latest year, two years, three years, …, and ten years.
- Cash Flow per Share Growth Rate: Ratio of gross cash flow to common shares outstanding. Determined by a log-linear least squares regression for the latest year, two years, three years, …, and ten years.
- Plowback Growth: Firm's retention ratio for the current year times the firm's latest annual return on equity.
- 3yr Plowback Growth: Firm's three-year average retention ratio times the firm's three-year average return on equity.

Consensus Analysts' Forecasts

- Five-Year Earnings Per Share Growth: Mean analysts' forecast compiled by IBES.

[1] Vander Weide, J. H., and W. T. Carleton. "Investor Growth Expectations: Analysts vs. History." *The Journal of Portfolio Management*, Spring 1988, pp. 78-82.

Risk Variables

B: Beta, the firm's beta versus NYSE from Value Line.

Cov: The firm's pretax interest coverage ratio from Compustat.

Stb: Five-year historical earnings per share stability. Average absolute percentage difference between actual reported EPS and a 5yr historical EPS growth trend line from IBES.

Sa: The standard deviation of earnings per share estimate for the fiscal year from IBES.

We set five restrictions on the companies included in the study in order to be consistent with the original study and to obtain more meaningful results.

- Excluded all firms that IBES did not follow.
- Eliminated companies with:
  - Negative EPS during any of the years 1991-2003.
  - No dividend during any one of the years 1991-2003.
  - P/E ratio greater than 60 in years 2001-2003.
  - Less than five years of operating history.

The final universe consisted of 411 US firms, fifty-nine of which are utility companies.

Results

The study was performed in two stages.

Stage 1

In order to determine which historically oriented growth measure is most highly correlated with each firm's end-of-year P/E ratio, we computed spearman (rank) correlations between all forty-two historically oriented future growth measures and P/E.

The result of the stage 1 study is displayed in Table 1. Three-year plowback ratio has the highest correlation with P/E in 2001 and 2002, and five-year EPS growth rate has the highest correlation with P/E in 2003.

Table 1

Stage1 Results for Utility and Non-Utility Companies Combined

Correlations between Historically Based Growth Estimates by Year with P/E

Current Year		y1	y2	y3	y4	y5	y6	y7	y8	y9	y10
2001	EPS	0.232	0.210	0.145	0.122	0.059	0.034	-0.007	-0.076	-0.117	-0.154
	DPS	-0.243	-0.297	-0.296	-0.293	-0.313	-0.316	-0.336	-0.334	-0.329	-0.333
	BVPS	0.059	-0.017	-0.098	-0.138	-0.150	-0.182	-0.219	-0.259	-0.271	-0.273
	CFPS	0.092	0.092	0.087	0.042	-0.063	-0.102	-0.141	-0.193	-0.237	-0.262
	plowback	0.203									
	plowback3	0.308									
2002	EPS	-0.007	0.147	0.076	0.080	0.083	0.050	0.030	-0.018	-0.060	-0.089
	DPS	-0.126	-0.202	-0.251	-0.224	-0.215	-0.239	-0.232	-0.233	-0.211	-0.198
	BVPS	-0.036	-0.036	-0.078	-0.115	-0.114	-0.127	-0.152	-0.162	-0.175	-0.171
	CFPS	0.056	0.045	0.017	0.021	0.030	-0.024	-0.050	-0.080	-0.125	-0.162
	plowback	0.093									
	plowback3	0.180									
2003	EPS	0.073	0.084	0.214	0.231	0.244	0.228	0.182	0.158	0.104	0.049
	DPS	0.120	0.054	-0.001	-0.078	-0.090	-0.126	-0.152	-0.165	-0.163	-0.185
	BVPS	0.097	0.076	0.067	0.036	-0.045	-0.062	-0.063	-0.083	-0.105	-0.131
	CFPS	0.146	0.196	0.243	0.239	0.206	0.176	0.107	0.089	0.039	-0.022
	plowback	-0.017									
	plowback3	0.038									

We also independently examined utility and non-utility firms. Table 2 shows the result for the fifty-nine utility firms. Two-year growth in EPS has the highest correlation with P/E in 2001, four-year EPS has the highest correlation in 2002, and six-year EPS has the highest correlation in 2003.

Table 3 exhibits the result for the remaining non-utility firms. EPS one-year growth, two-year growth, and five-year growth has the highest correlation with P/E in 2001, 2002, and 2003, respectively.

Table 2

Stage1 Results for Utility Companies

Correlations between Historically Based Growth Estimates by Year with P/E

Current Year		y1	y2	y3	y4	y5	y6	y7	y8	y9	y10
2001	EPS	0.305	0.330	0.305	0.319	0.238	0.157	0.129	0.107	0.079	0.048
	DPS	-0.215	-0.321	-0.302	-0.294	-0.316	-0.281	-0.332	-0.414	-0.435	-0.429
	BVPS	0.164	0.137	0.147	-0.027	-0.072	-0.135	-0.117	-0.104	-0.106	-0.140
	CFPS	0.194	0.135	0.020	-0.018	-0.122	-0.157	-0.135	-0.134	-0.103	-0.219
	plowback	-0.143									
	plowback3	-0.027									
2002	EPS	-0.065	0.044	0.069	0.119	0.071	0.004	-0.038	-0.069	-0.061	-0.070
	DPS	-0.333	-0.327	-0.278	-0.313	-0.280	-0.321	-0.277	-0.226	-0.203	-0.210
	BVPS	-0.325	-0.239	-0.182	-0.177	-0.230	-0.237	-0.250	-0.247	-0.235	-0.235
	CFPS	-0.205	-0.132	-0.172	-0.166	-0.216	-0.289	-0.285	-0.265	-0.227	-0.218
	plowback	-0.151									
	plowback3	-0.133									
2003	EPS	0.010	0.136	0.186	0.263	0.365	0.367	0.344	0.343	0.309	0.302
	DPS	0.151	-0.029	-0.014	-0.022	-0.054	-0.117	-0.142	-0.137	-0.105	-0.092
	BVPS	0.212	0.060	0.047	0.019	0.003	0.040	0.022	0.005	0.003	-0.002
	CFPS	0.222	-0.046	0.173	0.115	0.165	0.100	0.017	0.077	0.057	0.077
	plowback	-0.365									
	plowback3	-0.403									

Table 3

Stage1 Results for Non-Utility Companies

Correlations between Historically Based Growth Estimates by Year with P/E

Current Year		y1	y2	y3	y4	y5	y6	y7	y8	y9	y10
2001	EPS	0.1843	0.1660	0.1293	0.1218	0.0873	0.0829	0.0618	0.0106	-0.0194	-0.0412
	DPS	-0.2036	-0.2211	-0.2042	-0.1935	-0.2098	-0.2066	-0.2186	-0.2155	-0.2046	-0.1975
	BVPS	0.0757	0.0084	-0.0791	-0.0997	-0.0916	-0.1146	-0.1388	-0.1783	-0.1866	-0.1823
	CFPS	0.0864	0.0710	0.0956	0.0704	-0.0033	-0.0162	-0.0366	-0.0747	-0.1186	-0.1325
	plowback	0.0781									
	plowback3	0.1781									
2002	EPS	0.0762	0.1767	0.0755	0.0817	0.0936	0.0757	0.0708	0.0316	-0.0011	-0.0254
	DPS	-0.0804	-0.1693	-0.2103	-0.1672	-0.1519	-0.1720	-0.1645	-0.1636	-0.1394	-0.1226
	BVPS	0.0527	0.0236	-0.0363	-0.0777	-0.0710	-0.0753	-0.0953	-0.1019	-0.1118	-0.1061
	CFPS	0.0905	0.0488	0.0143	0.0237	0.0563	0.0246	0.0097	-0.0079	-0.0458	-0.0821
	plowback	0.0634									
	plowback3	0.1306									
2003	EPS	0.1254	0.1783	0.2788	0.2689	0.2791	0.2622	0.2219	0.2039	0.1559	0.1090
	DPS	0.1810	0.1290	0.0655	-0.0128	-0.0101	-0.0400	-0.0630	-0.0772	-0.0930	-0.0952
	BVPS	0.1555	0.1740	0.1534	0.1056	0.0127	-0.0069	-0.0054	-0.0218	-0.0416	-0.0636
	CFPS	0.1479	0.2200	0.2512	0.2429	0.2004	0.1839	0.1349	0.1286	0.0892	0.0388
	plowback	-0.1109									
	plowback3	-0.0402									

Stage 2

We compared the multiple regression model of historical growth rate with the highest correlation to the P/E ratio from stage 1 to the five-year earnings per share growth forecast.

$$P/E = a_0(D/E) + a_1 g + a_2 B + a_3 Cov + a_4 Stb + a_5 Sa + e$$

The regression results are displayed in table 4. The results show that the consensus analysts' forecast of future growth better approximates the firm's P/E ratio, which is consistent with the results found by Vander Weide and Carleton. In both regressions, R^2 in the regression with the consensus analysts' forecast is higher than the R^2 in the regression with the historical growth.

Table 4
Stage2 Results for Utility and Non-Utility Companies Combined
Multiple Regression Results
P/E = a0 + a1 D/E + a2 g + a3 B + a4 Cov + a5 Stb + a6 Sa

Historical

	a0	a1	a2	a3	a4	a5	a6	Rsq	F Ratio
2001	10.43	8.46	10.79	6.79	0.02	-0.03	-18.83	0.20	13.90
	4.73	5.53	2.93	3.54	3.05	-3.06	-3.32		
2002	12.36	7.60	6.66	1.01	0.00	0.01	-32.48	0.15	9.46
	7.21	6.18	2.61	0.66	1.57	1.48	-4.04		
2003	13.34	5.96	9.87	5.27	0.01	-0.01	-20.46	0.24	17.61
	7.29	4.04	2.95	3.39	3.62	-1.31	-4.25		

Analysts' Forecasts

	a0	a1	a2	a3	a4	a5	a6	Rsq	F Ratio
2001	-1.26	16.14	144.75	-0.64	0.01	-0.03	-10.76	0.47	48.00
	-0.62	11.63	13.22	-0.38	3.07	-4.04	-2.29		
2002	3.37	13.37	106.07	-3.60	0.00	0.01	-21.85	0.35	29.73
	1.93	10.97	10.59	-2.57	1.25	1.50	-3.06		
2003	4.77	12.76	61.93	4.38	0.01	0.00	-19.41	0.33	26.38
	2.65	9.48	7.25	3.01	2.45	-0.81	-4.33		

*T-stats below the coefficients in smaller font

For utility companies shown in table 5, consensus analysts' forecast of future growth is superior to historically oriented growth in 2002 and 2003. R^2 is lower in the regression with the consensus analysts' forecast in 2001. For non-utility companies, we found that consensus analysts' forecast of future growth is superior to the alternative in all three years (table 6).

Table 5
Stage2 Results for Utility Companies

Multiple Regression Results

P/E = a0 + a1 D/E + a2 g + a3 B + a4 Cov + a5 Stb + a6 Sa

Historical

	a0	a1	a2	a3	a4	a5	a6	Rsq	F Ratio
2001	7.90	11.07	-11.19	-3.00	0.29	0.00	-9.37	0.44	6.38
	2.16	4.80	-5.71	-0.86	0.88	0.64	-1.51		
2002	13.87	7.00	-3.80	-6.89	0.56	0.00	-29.89	0.38	5.11
	4.02	3.54	-0.66	-2.01	1.48	0.42	-2.70		
2003	11.29	7.74	-1.65	-1.40	0.32	0.00	-5.69	0.25	2.68
	3.22	3.30	-0.23	-0.43	1.05	-0.73	-0.75		

Analysts' Forecasts

	a0	a1	a2	a3	a4	a5	a6	Rsq	F Ratio
2001	9.61	9.20	66.61	-7.92	0.50	-0.01	-12.83	0.27	2.95
	2.31	3.45	3.66	-1.86	1.31	-1.33	-1.76		
2002	12.43	7.86	50.74	-9.61	0.50	0.00	-24.94	0.48	7.56
	3.89	5.29	3.10	-2.94	1.50	0.17	-2.41		
2003	5.81	11.06	101.12	-1.69	-0.19	0.00	-4.75	0.50	7.81
	1.89	6.32	4.80	-0.58	-0.74	-0.22	-0.74		

*T-stats below the coefficients in smaller font

Table 6
Stage2 Results for Non-Utility Companies

Multiple Regression Results

P/E = a0 + a1 D/E + a2 g + a3 B + a4 Cov + a5 Stb + a6 Sa

Historical

	a0	a1	a2	a3	a4	a5	a6	Rsq	F Ratio
2001	15.90	8.39	2.82	3.53	0.02	-0.03	-21.05	0.21	12.45
	6.57	4.13	1.96	1.68	2.97	-2.14	-3.40		
2002	17.76	8.46	6.02	-3.06	0.00	0.02	-36.97	0.27	16.78
	9.39	5.19	3.28	-1.88	1.37	2.52	-4.31		
2003	14.24	9.86	8.85	3.46	0.01	0.00	-19.00	0.30	19.89
	7.49	5.89	2.49	2.11	3.23	-0.15	-3.73		

Analysts' Forecasts

	a0	a1	a2	a3	a4	a5	a6	Rsq	F Ratio
2001	-0.51	17.28	140.84	-1.06	0.01	-0.03	-8.63	0.44	36.00
	-0.22	11.21	10.73	-0.59	2.88	-2.62	-1.63		
2002	5.05	15.67	91.22	-4.06	0.00	0.02	-22.93	0.38	27.65
	2.48	11.23	7.66	-2.74	1.18	2.33	-2.87		
2003	7.25	14.47	45.60	3.47	0.01	0.00	-19.09	0.33	22.30
	3.56	9.42	4.68	2.20	2.36	-0.12	-3.89		

*T-stats below the coefficients in smaller font

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The average growth rate estimate from all the analysts that follow the company measures the consensus expectation of the investment community for that company. In most cases, it is necessary to use earnings forecasts rather than dividend forecasts due to the extreme scarcity of dividend forecasts compared to the widespread availability of earnings forecasts. Given the paucity and variability of dividend forecasts, using the latter would produce unreliable DCF results. In any event, the use of the DCF model prospectively assumes constant growth in both earnings and dividends. Moreover, as discussed below, there is an abundance of empirical research that shows the validity and superiority of earnings forecasts relative to historical estimates when estimating the cost of capital.

The uniformity of growth projections is a test of whether they are typical of the market as a whole. If, for example, 10 out of 15 analysts forecast growth in the 7%–9% range, the probability is high that their analysis reflects a degree of consensus in the market as a whole. As a side note, the lack of uniformity in growth projections is a reasonable indicator of higher risk. Chapter 3 alluded to divergence of opinion amongst analysts as a valid risk indicator.

Because of the dominance of institutional investors and their influence on individual investors, analysts' forecasts of long-run growth rates provide a sound basis for estimating required returns. Financial analysts exert a strong influence on the expectations of many investors who do not possess the resources to make their own forecasts, that is, they are a cause of g. The accuracy of these forecasts in the sense of whether they turn out to be correct is not at issue here, as long as they reflect widely held expectations. As long as the forecasts are typical and/or influential in that they are consistent with current stock price levels, they are relevant. The use of analysts' forecasts in the DCF model is sometimes denounced on the grounds that it is difficult to forecast earnings and dividends for only one year, let alone for longer time periods. This objection is unfounded, however, because it is present investor expectations that are being priced; it is the consensus forecast that is embedded in price and therefore in required return, and not the future as it will turn out to be.

Empirical Literature on Earnings Forecasts

Published studies in the academic literature demonstrate that growth forecasts made by security analysts represent an appropriate source of DCF growth rates, are reasonable indicators of investor expectations and are more accurate than forecasts based on historical growth. These studies show that investors rely on analysts' forecasts to a greater extent than on historic data only.

Academic research confirms the superiority of analysts' earnings forecasts over univariate time-series forecasts that rely on history. This latter category

includes many *ad hoc* forecasts from statistical models, ranging from the naive methods of simple averages, moving averages, etc. to the sophisticated time-series techniques such as the Box-Jenkins modeling techniques. The literature suggests that analysts' earnings forecasts incorporate all the public information available to the analysts and the public at the time the forecasts are released. This finding implies that analysts have already factored historical growth trends into their forecast growth rates, making reliance on historical growth rates somewhat redundant and, at worst, potentially double counting growth rates which are irrelevant to future expectations. Furthermore, these forecasts are statistically more accurate than forecasts based solely on historical earnings, dividends, book value equity, and the like.

Summary of Empirical Research

Important papers include Brown and Rozeff (1978), Cragg and Malkiel (1968, 1982), Harris (1986), Vander Weide and Carleton (1988), Lys and Sohn (1990), and Easterwood and Nutt (1999).

The study by Brown and Rozeff (1978) shows that analysts, as proxied by Value Line analysts, make better forecasts than could be obtained using only historical data, because analysts have available not only past data but also a knowledge of such crucial factors as rate case decisions, construction programs, new products, cost data, and so on. Brown and Rozeff test the accuracy of analysts' forecasts versus forecasts based on past data only, and conclude that their evidence of superior analyses means that analysts' forecasts should be used in studies of cost of capital. Their evidence supports the hypothesis that Value Line analysts consistently make better predictions than historical time-series models.

Using the IBES consensus earnings forecasts as proxies for investor expectation, Harris (1986) estimates the cost of equity using expected rather than historical earnings growth rates. In his review of the literature on financial analysts' forecasts, Harris concludes that a growing body of knowledge shows that analysts' earnings forecasts are indeed reflected in stock prices. Elton, Gruber, and Gultekin (1981) show that stock prices react more to changes in analysts' forecasts of earnings than they do to changes in earnings themselves, suggesting the usefulness of analysts' forecasts as surrogates for market expectations. In an extensive National Bureau of Economic Research study using analysts' earnings forecasts, Cragg and Malkiel (1982) present detailed empirical evidence that the average analyst's expectation is more similar to expectations being reflected in the marketplace than historical growth rates, and that it is the best possible source of DCF growth rates. The authors show that historical growth rates do not contain any information that is not already impounded in analysts' growth forecasts. They conclude that the expectations formed by Wall Street professionals get quickly and thoroughly impounded

Response No. MCC-197
Attachment D

A Review of Research Related to Financial Analysts' Forecasts and Stock Recommendations

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First Draft: 15 June 2005
This Draft: 30 June 2008

We greatly appreciate the research assistance of Kevin Hee and comments and suggestions from anonymous referees, Mark Bradshaw, Larry Brown, Donal Byard, Geoff Friesen, Steve Glover, Zhaoyang Gu, Rick Johnston, Dave Smith, Barry Spicer, Jake Thomas, and Jilnaught Wong.

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Electronic copy available at: http://ssrn.com/abstract=848248

A Review of Research Related to Financial Analysts' Forecasts and Stock Recommendations

Abstract: This paper reviews research regarding the role of financial analysts in capital markets. The paper builds on the perspectives provided by Schipper (1991) and Brown (1993). We categorize papers published mainly since 1992 and selectively discuss aspects of these papers that address or suggest key research topics of ongoing interest in seven broad areas: analysts' decision processes, the determinants of analyst expertise and distributions of individual analysts' forecasts, the informativeness of analysts' research outputs, analyst and market efficiency with respect to information, effects of analysts' economic incentives on their research outputs, effects of the institutional and regulatory environment (including cross-country comparisons), and the limitations of databases and various research paradigms.

Electronic copy available at: http://ssrn.com/abstract=848248

A Review of Research Related to Financial Analysts' Forecasts and Stock Recommendations

1. Introduction

This paper reviews research related to the role of financial analysts in the allocation of resources in capital markets.[1] Two important papers published in the early 1990s provide perspectives on the literature in this area, one appearing in *Accounting Horizons* (Schipper, 1991) and the other appearing in the *International Journal of Forecasting* (Brown, 1993). Our paper begins by summarizing the perspectives and directions for future research suggested in Schipper (1991) and Brown (1993).[2] We then take a look at the highlights of what we have learned and new questions that have emerged since 1992. Our goal is to provide an organized look at the literature, with particular attention to important questions that remain open for further research. Brown (1993) did not restrict his review of earnings forecasting research to the role of financial analysts. We focus more narrowly on research related to analysts' decision processes and the usefulness of their forecasts and stock recommendations. Kothari (2001) provides an excellent overall review of capital markets research, and we refer the reader to that paper for a broader perspective.

Since 1992 no less than 250 papers related to financial analysts have appeared in the nine major research journals that we used to launch our review of the literature. We also note that during the six months ending February 21, 2008 alone 154 new working papers with the word "analysts" in the abstract have posted to the Social Sciences Research Network, so the task for

[1] Much of this paper reprints, by permission, Ramnath et al. (2008). Ramnath et al. (2008) provide a detailed taxonomy listing and categorizing every paper published in 11 research journals since 1992. This paper differs from Ramnath et al. (2008) in that this paper selectively and critically reviews examples of the research in each of the seven broad areas of the Ramnath et al. (2008) taxonomy.

[2] Also see Givoly and Lakonishok (1983) for a review of analysts' forecasting research prior to 1983.

the next authors of a review article in this area will be even more daunting.

In our review of papers published since 1992, we find much progress in some of the areas identified by Schipper (1991) and Brown (1993) and less progress in other areas. In particular, the research has evolved from descriptions of the statistical properties of analysts' forecasts to investigations of the incentives and decision processes that give rise to those properties. However, in spite of this broader focus, much of analysts' decision processes and the market's mechanism of drawing a useful consensus from the combination of individual analysts' decisions remain hidden in a black box. Furthermore, we still have much to learn about the relevant valuation metrics and the mechanism by which analysts and investors translate forecasts into present equity values. For example, with renewed popularity of the earnings-based valuation model in the early 1990s, research turned toward an investigation of the model's role in the market's conversion of analysts' earnings forecasts into current equity values. Given the unexpected result that this model does a relatively poor job of explaining the variation in market prices and analysts' price forecasts and recommendations, researchers have turned their attention to examining heuristics that might better explain analyst and market decisions about firm value. We still have much to learn about the heuristics relied upon by analysts and the market and appropriateness of their use.

The rest of this paper draws attention to these and other issues that have arisen since 1992. The next section provides a summary of the questions identified in Schipper (1991) and Brown (1993) and the directions for future research suggested by those authors, as well as the authors of the four papers commenting on Brown (1993). Section 3 reviews papers published mainly since Schipper (1991) and Brown (1993) and also identifies new research questions that emerge from our reading of the literature. Section 4 provides concluding comments highlighting

the areas we consider most promising for future research.

2. Perspective from Schipper (1991) and Brown (1993)

Schipper's (1991) commentary makes two major points. First, she suggests that research regarding analysts' earnings forecasts tends to focus too narrowly on the statistical properties of the forecasts, without considering the full decision context and economic incentives that may affect those properties. She takes the perspective that the analyst's job is to provide buy-sell-hold recommendations and provide research reports to support those recommendations. Schipper (1991) describes analysts' earnings forecasts as one component of their research reports and a means to an end rather than an end in themselves. She suggests that a more complete description of analysts' economic incentives and the role of earnings forecasts in the full decision context of analysts should lead to richer hypotheses regarding the statistical properties of the earnings forecasts. The second major point (related to the first) is that research regarding the statistical properties of analysts' earnings forecasts focuses on the outputs from rather than the inputs to the analyst's decision process. The commentary calls for more research into how analysts actually use accounting information and their own earnings forecasts in making decisions.

From Brown's (1993) review paper, we glean four key points. First, he points out that models producing the most accurate forecasts of an earnings variable should also produce the best proxies for the market's expectations, assuming market efficiency and assuming the research design correctly models the valuation implications of the earnings variable. Under these assumptions, in Brown's words, "predictive ability and association are two sides of the same coin." Brown notes mixed results on this issue and calls for future research to sort out whether

the apparently conflicting results stem from research design problems or market inefficiency. Second, Brown admonishes researchers to carefully consider whether summary files of I/B/E/S consensus forecasts make sense for their studies. Although the date of the I/B/E/S report and coding of the forecast horizon indicates a timely consensus, the consensus may contain stale forecasts not updated since the information event on which the study intends to condition the forecasts (O'Brien, 1988). Brown suggests that using the I/B/E/S detail files can avoid this problem, and that forecast timeliness is a crucial attribute for effective proxies for analyst earnings expectations when evaluating the accuracy or price-relevance of those expectations.[3] Third, Brown (1993) calls for research to better understand the role of analysts' forecasts in post-earnings announcement drift. In particular, he calls for research into reasons for variation in the degree and speed of forecast convergence following earnings announcements and the effect, if any, of forecast convergence on post-earnings announcement drift. Finally, like Schipper (1991), Brown calls for research to better understand the decision processes of analysts and the roles of analysts' earnings forecasts, macroeconomic and industry factors, and other information in formulating stock price forecasts and recommendations.

In terms of research methodology, both Brown (1993) and Schipper (1991) indicate that behavioral research can play a more prominent role in understanding the uses of accounting and other information to make stock recommendations, within the full context of the analyst's decision environment and economic incentives. In Brown's words, "joint research by capital markets researchers and behavioralists to examine these issues more thoroughly would considerably enhance our understanding of the role of analysts in the price formation process."

[3] Most of the studies reviewed by Brown (1993) relied on either I/B/E/S consensus or *Value Line* data. With less frequency, studies also used Merrill Lynch's Opinion Alert, Standard and Poors Earnings Forecaster and Zack's Investment Research. Some used detail files from I/B/E/S and Zacks which, as the paper points out, only became readily available towards the end of the period reviewed.

It will be interesting to see the extent to which the role of behavioral research has expanded since 1992.

Four authors commented on Brown (1993), and each provides interesting insights and suggestions for future research. For example, O'Hanlon (1993) calls for investigations of the degree to which financial analysts' earnings forecasts effectively distinguish permanent from temporary earnings changes. Thomas (1993, p. 327) suggests that the importance of research into how analysts make earnings predictions depends on the answers to several questions, including: whether analysts' forecasts influence the marginal investor; whether they seek to predict reported earnings or a 'core' earnings number that will persist in the future; and whether their incentives are consistent with producing the most accurate forecasts possible. Brown (P.) (1993) calls for research into whether some analysts are better forecasters than others, whether the market's earnings expectations reflect these differences, and the degree to which consensus forecasts drawn from analyst tracking services such as I/B/E/S reflect investor expectations. Zmijewski (1993) focuses on the need for investigations of cross-country variation in properties of earnings forecasts and their roles in price formation in capital markets.

Based on our reading of Schipper (1991) and Brown (1993) and its related comment papers, along with an initial look at the research published over the last 12 years, we decided to organize our review around the following seven broad research areas: (1) What is the nature of analysts' decision processes and how do analysts rationalize the forecasts and recommendations contained in their research reports? (2) What is the nature of analyst expertise and what are the characteristics of distributions of individual analyst earnings forecasts? (3) How informative are the outputs from analyst research (including earnings forecasts, target price forecasts, stock recommendations and conceptual analysis)? (4) Do analysts' forecasts and recommendations

efficiently impound information about future earnings? Do stock prices efficiently impound the information in analysts' forecasts and recommendations? (5) How do analyst and management incentives affect the statistical properties of analysts' forecasts? (6) How does variation in the regulatory environment (over time and across countries) affect the behavior of analysts' forecasts and the role of analysts in capital markets? (7) What are some research design and database issues that threaten the validity of inferences from studies of the behavior of analysts and their forecasts and recommendations? The next section is divided into seven subsections that selectively review research papers addressing these questions, with selective focus on papers published since Brown (1993).

3. Selective review of research related to the role of financial analysts in capital markets

The questions at the end of section 2 above naturally arise from the picture of the analyst's reporting environment shown in figure 1. Figure 1 puts the analyst's overall reporting environment in perspective. The analyst obtains/develops information from various sources including: earnings and other information from SEC filings, such as proxy statements and quarterly and annual reports; industry reports and reports describing macro-economic conditions; and conference calls and other management communications. From this information the analyst produces earnings forecasts, target price forecasts and stock recommendations, along with a conceptual report describing the firm's prospects. Investors use these outputs from analyst research to make trading decisions that affect market prices. If the analyst forecasting process and capital markets are efficient, then market prices and analysts' forecasts immediately reflect all of the information described in the figure. Inefficiencies create predictable analyst forecast errors and stock price changes. The decision processes and analyst research output pictured in

figure 1 depend on overall governing forces including: regulatory and institutional factors that vary over time and across countries, and analysts' economic incentives. Finally, limitations associated with archival databases and econometric/analytical research technology create research design issues that constrain our views of the forces that ultimately drive market prices.

We launched our review by listing and categorizing all papers related to analysts and published since 1992 in the following nine major research journals: *The Accounting Review*, *Contemporary Accounting Research*, *International Journal of Forecasting*, *Journal of Accounting and Economics*, *Journal of Accounting Research*, *Journal of Finance*, *Journal of Financial Economics*, *Review of Accounting Studies*, and *Review of Financial Studies*. However, we expanded the set of papers as needed, given references to papers outside of the initial timeframe or published in other journals. We discuss papers as they apply to each of the seven research areas described at the end of section 2 above. These seven subsections are also shown in figure 1 and form our outline below. Our goal is not to provide exhaustive reviews of (or even references to) all of the papers published since 1992, but rather to selectively discuss papers addressing key issues and to suggest new questions that might become the subject of research during the foreseeable future.

3.1. Inputs to analysts' earnings forecasts and stock recommendations

This section reviews research regarding the role of earnings and other information in the broader context of the decision processes analysts use to produce their research reports and stock recommendations.

3.1.1 Inputs to research reports

Learning about the information analysts use and understanding analysts' decision processes is no easy matter. Researchers have used surveys to simply ask analysts how they process information (e.g., Block 1999), protocol analysis to record analysts' thought processes as they process information (e.g., Bouwman, et al. 1995), content analysis of analysts' research reports to infer the information analysts' rely upon to make forecasts and recommendations (e.g., Rogers and Grant 1997), and laboratory experiments to study how analysts use information (e.g., Maines, et al. 1997). As opposed to the research methods mentioned above, archival studies potentially offer more generalizable results, but are limited in their ability to penetrate the black box containing analysts' actual decision processes. The challenge is that analysts have a context-specific task that is very difficult to model. The rest of this section reviews research from each of the paradigms mentioned above.

Previts et al. (1994) examine a broad cross-section of 479 sell-side analyst reports (from Investext) to ascertain the information that analysts apparently use to make decisions. The content of the reports suggested heavy analyst use of earnings-related information and a strategy of disaggregating firm-level information into segments beyond the disaggregation level provided in GAAP-based segment reporting footnotes. The study also finds evidence of substantial analyst effort to extract non-recurring items and focus on "core" or "adjusted" earnings as a basis for forecasting future earnings.[4] The study also reports heavy analyst reliance on management for information, emphasizing the intermediation role of financial analysts. Interestingly, the authors also observe that analysts prefer following firms with effective strategies for presenting smooth earnings streams. The paper reports (p. 63) that "analysts most frequently refer to accounting earnings quality in terms of a company's ability to manage earnings through the

[4] The focus on "core" earnings relates to the substantial literature on 'pro-forma' earnings (e.g., Gu and Chen (2004) and Bradshaw and Sloan (2002)).

establishment and adjustment of conservative, discretionary reserves, allowances, and off-balance-sheet assets, which provide analysts a low-risk earnings platform for making stock price forecasts and buy/sell/hold recommendations..." It would be interesting to investigate whether analysts continue to endorse such views in the wake of Regulation Fair Disclosure (Reg. FD) and the Sarbanes-Oxley Act of 2002. Reg. FD changed the landscape for management correspondence with analysts so as to make information uniformly available to all analysts. We discuss the impact of Reg. FD and related research more fully in section 3.6.1. The Sarbanes-Oxley Act of 2002 called for expanded disclosure by firms and imposed senior executive responsibility for corporate disclosures. The Act also defines a code of conduct for analysts. Collectively, the various provisions of Reg. FD and the Sarbanes-Oxley Act likely decreased firms' ability to manage earnings and increased incentives and rewards for analysts to forecast unmanaged earnings without the help of private communications from management.

Rogers and Grant (1997) extend Previts et al. by using "context-specific" content analysis to examine 187 sell-side analyst reports issued between July 1993 and June 1994 (obtained from the One Source database). They report that only about one-half of the information in the analysts' research reports could be found in the corresponding corporate annual reports, consistent with analysts also using other, external, information . Further, even within the annual report about one-half of the information seems to have been obtained from the narrative sections (e.g., MD&A and president's letter) rather than the basic financial statements. Thus, examining analyst reports based solely on quantitative information may not capture the complex nature of the analyst's task.

Using protocol analysis, Bouwman et al. (1995) confirm the aforementioned findings from content analysis of research reports that analysts rely on a variety of sources in their

analysis.[5] They further find that, in making investment decisions, the nature of information used by buy-side analysts varies among the different stages of the decision process. In the familiarization stage, which involved getting to know the company and its general financial condition, subjects relied heavily on the GAAP-based materials, especially performance ratios and the historical summary provided in the 10-K. They relied most heavily on segment information (with more emphasis on non-GAAP materials) while performing an in-depth analysis of the company. Finally, while integrating material from the other phases and reaching a final screening decision, the analysts relied more heavily on non-GAAP general company information. Analysts also often verbalized a desire for information not included with the materials, such as more segment information, a 10-year rather than 5-year historical summary, more forward looking information about management's plans and one or two research reports by sell-side ("Street") analysts.

Maines, et al. (1997) use experimental methods to examine the importance of segment disclosures to analysts' decisions. Subjects were presented with a hypothetical firm's financial statements, which included results for two segments, one of which included two divisions.[6] The study manipulated two factors: congruence/incongruence between the segment definitions for internal and external reporting purposes; and similarity/dissimilarity between the product lines of the divisions combined in the segment with two divisions. As long as the analysts were told that the company's external reporting conformed to its internal reporting, there was no evidence to suggest that dissimilarity between divisions combined into a segment diminished analyst

[5] Bouwman et al. (1995) asked 12 buy-side analysts to "think out loud" as they examined a variety of information to determine whether a particular company merited further consideration as an investment opportunity. The analysts took an average of 52 minutes at the task and generated an average of 4,000 recorded words, which the researchers transcribed into nearly 400 pages of text. GAAP-based materials included the company's most recent 10-K, 10-Q and proxy statement; and non-GAAP materials included a COMPUSTAT report, an S&P stock report, an industry level report and current stock price information.

[6] Their subjects averaged 5.5 years of experience and included 42 buy-side analysts, 8 sell-side analysts and 6 analysts involved in other investment-related jobs.

confidence in their forecasts and recommendations. The results suggest that analysts suspect companies of obscuring the definitions of segments in order to hide performance issues, and that the FASB's new conformity rules for segment reporting should alleviate these suspicions. The same experiment was conducted with 60 MBA students. Unlike the experienced analysts, the MBA students were unable to discern from information in the MD&A section of the annual report whether or not the two divisions combined into one segment were similar or dissimilar. This calls into question the use of student subjects as surrogates for financial analysts and also suggests the need for research to better understand the development of analyst expertise.

Several experimental studies have evaluated whether classification issues affect expert analysts' judgments. For example, in an experiment involving 56 experienced buy-side analysts who specialize in banking industry stocks, Hirst et al. (2004) evaluate whether classification of gains and losses on financial assets and liabilities affected analysts' judgments. The researchers manipulated two variables: the reporting of gains and losses due to interest rate risk (full-fair-value versus piecemeal-fair-value); and the banks exposure to interest rate risk (fully hedged versus exposed). In the full-fair-value condition, fair values of all financial assets and liabilities were recorded on the balance sheet and all gains and losses were reported on a performance statement immediately following the income statement. Under piecemeal-fair-value accounting the same information was available, but deposits and liabilities were not marked to market on the balance sheet and gains and losses on those instruments were reported in the footnotes. The results indicate that the analysts were only able to effectively adjust their risk assessments and valuation judgments for the higher risk of exposed banks when the financial statements of those banks applied full-fair-value accounting. The study also examined the effect of the analyst's workload and found that analysts following less than 40 (the sample median) stocks in their

normal work environment performed significantly better in distinguishing the risk characteristics of exposed versus fully hedged banks. This suggests the need for further research into the effects of spreading analysts' workloads across larger numbers of stocks. It also suggests an important reason why buy-side analysts have an interest in the reports of sell-side analysts who typically follow far fewer stocks. Other experimental studies showing similar effects of accounting method or classification on the judgments of experienced analysts include Hopkins (1996), Hirst and Hopkins (1998) and Hopkins et al. (2000).

A number of archival studies suggest that complexity affects analyst forecast accuracy (e.g., Brown et al. 1987; Haw et al. 1994; Lang and Lundholm 1996; Duru and Reeb 2002; Clement 1999). Perhaps the most direct test of this proposition is in Plumlee (2003), which examines the effects of six aspects of the Tax Reform Act of 1986 (TRA86) on errors in Value Line analysts' effective tax rate forecasts for 355 firms. The six tax law changes vary in complexity from a simple decline from 48% to 36% in the corporate statutory tax rate to the highly complex implementation of an alternative minimum tax. The results indicate a significant positive relation between the absolute values of errors in forecasting effective tax rates and the research proxies for the expected effects of the three most complex tax law changes; and an insignificant relation with the three least complex tax law changes. Plumlee interprets her results as indicating that higher information complexity reduces analysts' use of the information, due either to analysts' processing limitations or time constraints. Since the research design did not predict the direction of the forecast errors, an alternative explanation is that analysts obtained and efficiently processed all possible information regarding the effects of the more complex tax law changes, but because those effects were highly uncertain, forecast errors were large in absolute

value for firms most affected. Further research is needed to distinguish between these explanations.

Research also suggests that analysts are more likely to cover firms that provide them with more (better quality) information required for analysis . Lang and Lundholm (1996) find that the quality of corporate disclosures (as rated by expert analyst committees) affects analysts' coverage decisions and the accuracy of their forecasts. In a more detailed study of the disclosure items that analysts value most, Healy and Hutton (1999) identify 97 firms whose disclosure quality ratings jumped significantly during the period between 1978 and 1991. They find that the following factors apparently play an important role in analysts' evaluations of firms' disclosures policies: segmental reporting quality, quality and candidness in the management discussion and analysis (MD&A) section of annual and quarterly reports; publication of supplemental disclosures outside of required periodic reports; and availability of management to analysts. Consistent with Healy and Hutton (1999), Botosan and Harris (2000) find that analyst following increased with firms' decisions to include information on segment activity as part of their quarterly (as opposed to only annual) reports.

In other examples of studies that relate disclosure quality to analysts' forecasting decisions, Bowen et al. (2002) and Barron et al. (1999) find that analysts' forecast accuracy depends on conference call information and the quality of managers' MD&A disclosures, respectively; and Williams (1996) finds that analysts' reliance on management earnings forecasts depends on the reliability of the forecast as measured by past management forecast accuracy. Chandra et al. (1999) find that analysts effectively use industry trade association disclosures to estimate the persistence of firm-specific sales changes. Ely and Mande (1996) find that analysts' earnings forecast revisions reflect the corroborative information in dividend and earnings

announcements, particularly when the earnings information is noisy. Thus, analysts appear to rely on detailed information in corporate reports, and their incentives to cover firms and report accurately appear to be related to the quality of corporate disclosures. We expect archival research to continue identifying associations between information variables and analysts' forecasting decisions. However, archival association studies cannot determine whether analysts *actually* use the specific information variables or whether analysts' private information production activities produce signals that are correlated with the public signals. Experimental research and protocol analysis can play important roles in evaluating the relative importance of each public information variable, as well as the importance of analysts' private information development activities, to analysts' actual forecasting decision processes.

3.1.2 The role of earnings forecasts and other information for analysts' recommendations

Bandyopadhyay et al. (1995) examine "the decision-usefulness of analysts' earnings forecasts for their price forecasts" (p. 430). They obtain analysts' price and earnings forecasts for 128 firms from Research Evaluation Service (RES), and Value Line price and earnings forecast data for the 46 firms in the group that are also covered by Value Line. RES supplies forecasts of prices over a 12 month forecast horizon and Value Line supplies price forecasts over a 3-5 year forecast horizon. Both databases provide current year and next year earnings forecasts, and Value Line also provides a 3-5 year earnings forecast. The primary results show that revisions (from June 30 to September 30) in RES forecasts of the next year's earnings explains about 30% of the variation in revisions in RES's 12-month ahead price forecast. In contrast, revisions in Value Line's 3-5 year earnings forecast explain about 60% of the variation in the revisions in Value Line's 3-5 year price forecasts. Thus, the authors conclude that

analysts' earnings forecasts provide important inputs to their price forecasting decisions, and the relation between the two variables increases with the forecast horizon. While these results provide insight into analysts' decision processes, they do not provide information as to the mechanism used by analysts to convert earnings forecasts into price forecasts.

Block (1999), in a survey of 880 members of the Association for Investment Management and Research (AIMR), finds that nearly half of the respondents never used present value techniques, while only 15% always used present value techniques in security valuation. Demirakos et al. (2004) use content analysis to examine sell-side analyst research reports (from Investext) and also find that analysts overwhelmingly refer to simple P/E multiples to support their stock recommendations. Bradshaw (2002) analyzes research reports of analysts and he too finds that analysts do not refer to present value techniques as a method they rely on to support their recommendations.[7] However, Bradshaw (2004, p. 27) points out that analysts may choose to communicate with investors in terms of simple heuristics that correlate with more sophisticated multiperiod present value models underlying analysts' valuation and recommendation judgments. He finds that a simple heuristic based on analysts' consensus long-term growth rate forecasts explains 23% of the variation in analysts' consensus stock recommendations (both variables from First Call over the 1994-98 time period). However, this simple heuristic (i.e., buy stocks with high long-term growth rate forecasts) is *negatively* correlated with value-to-price ratios derived from more sophisticated residual income valuation models. Furthermore, the long-term growth rate forecasts are also *negatively* correlated with abnormal returns over the year following the publication of the consensus forecasts and

[7] Consistent with the view that analysts rely more on multiples to value stocks, as opposed to relying on variants of DCF valuation models, only 23% of the reports contain earnings forecasts with horizons beyond the next fiscal year. Also, consistent with this view, the study finds that only 13% of the analyst reports refer to any variation of a DCF valuation model in computing price targets.

recommendations. This evidence is consistent with analysts pushing stocks whose high long-term growth rate forecasts have already been overpriced by the market. Bradshaw also finds evidence to suggest that the value-to-price ratios are *positively* associated with future abnormal returns but *negatively* associated with the analysts' recommendations. All in all, Bradshaw's evidence suggests that analysts do not use their own earnings forecasts efficiently in making recommendations. However, Bradshaw's sample period corresponded to a time period when the market was overheating, perhaps due to analysts' pushing their long-term growth forecasts. It will be interesting to examine if heuristics used by analysts to generate recommendations change over time, as well as the effects of these heuristics and recommendations on stock prices in different time periods.

If analysts' valuation judgments do not conform to finance theory, what models do analysts use to convert their forecasts into value judgments? Bradshaw (2002) examines approximately 100 analysts research reports from Investext, dated (primarily) in 1998 or 1999, and finds that analysts most frequently justify their recommendations with references to P/E ratios and long-term growth rate forecasts. Thus, it appears that analysts combine their long-term growth forecasts with the firm's P/E ratio to reach a valuation and recommendation decision. The PEG ratio, a popular Street heuristic, suggests that a firm's forward P/E ratio should equal 100 times its long-term growth rate forecast (Lynch 1989, p. 198). Bradshaw (2002) uses this heuristic to create pseudo price targets and finds that these pseudo price targets are highly correlated (r=0.39) with the level of analysts' buy/hold recommendations and with analysts' reported target prices (r=0.50). On the other hand, he finds smaller correlations of reported target prices with pseudo target prices based on industry P/E multiples (that do not consider differences across firms in long-term growth rates) and no correlation between these

P/E multiples-based pseudo target prices and recommendations. He concludes that the PEG ratio is an important heuristic used by analysts to convert their earnings forecasts into target prices forecasts and recommendations. Alternatively, analysts may use other valuation techniques, but simply use the PEG ratio to justify report recommendations and/or use models that are correlated with the PEG ratio.

While Bradshaw (2004) finds that consensus analyst recommendations based on analysts' consensus long-term earnings growth rate forecasts do not predict abnormal returns, he does not examine the association between the relative accuracy of an individual analyst's earnings forecasts and the profitability of the analyst's stock recommendations. Loh and Mian (2006) address this issue. Specifically, they compare the profitability of stock recommendations of relatively accurate earnings forecasters to those of poor earnings forecasters in any given firm-year. Relying on I/B/E/S earnings forecasts and recommendations related to over 32,000 firm-years between 1994 and 1999, they find that monthly abnormal returns on hedge portfolios based on recommendations of analysts in the top (bottom) quintile of earnings forecast accuracy are, on average, approximately 0.74% (-0.53%). The differences are highly significant, both statistically and economically. The authors infer that efforts by analysts to produce accurate earnings forecasts pay off in terms of the profitability of their stock recommendations. Thus, it appears that analysts use their earnings forecasts to produce stock recommendations, with more accurate forecasters providing more profitable recommendations. However, the model translating the earnings forecasts into valuation and recommendation judgments remains an elusive issue for further research.

3.2. The nature of analyst expertise and distributional characteristics of individual analyst earnings forecasts

In this section, we discuss research related to two commonly studied properties of analysts' earnings forecasts: forecast accuracy of individual analysts and the dispersion in forecasts provided by all analysts for a firm.

3.2.1 Forecast Accuracy

Studying earnings forecast accuracy of individual analysts is important for at least two reasons. First, investors benefit from identifying more accurate forecasts (and forecasters). Earnings forecasts are an input to analysts' stock recommendations; more accurate forecasters may provide more profitable stock recommendations, i.e., better input leads to better output (Loh and Mian 2006). Second, from a researcher's perspective identifying more accurate forecasts is important because in an efficient market, the market's expectation should reflect the best (most accurate) information available at any point in time. Studies that use analysts' forecasts as proxies for the market's earnings expectation should take into account investors' ability to identify and differentially weight earnings forecasts of individual analysts (Maines 1996).

Forecast accuracy is also important to analysts. More accurate forecasters are likely to be rewarded and less accurate forecasters may be forced to change brokerage houses or leave the profession. The reward to accurate forecasting may be in the form of recognition (for example, being selected to the Institutional Investor All American team) and/or career advancement (Hong and Kubik 2003). Mikhail et al. (1999) examine the relation between forecast accuracy and analyst turnover and find that analysts who contribute forecasts to the Zacks database are more likely to change brokerage firms or leave the database altogether when their forecast accuracy is lower relative to their peers. They find that the profitability of analysts' stock recommendations is unrelated to analyst turnover, suggesting that analysts may have more of an incentive to issue

accurate forecasts than to provide profitable stock recommendations. Forecast accuracy therefore seems important to both analysts and investors.

Research on analysts' earnings forecast accuracy has focused on two main attributes: 1) the nature of the forecast itself, for example, whether the forecast reflects new information or whether the analyst is merely herding with the consensus; and 2) characteristics of the analyst including his/her affiliation, for example, the analyst's prior experience and the brokerage firm that the analyst represents.

3.2.1.1. Forecast characteristics

Analyst forecasts for a firm differ on a number of dimensions like age (time between the forecast date and the related earnings announcement) and implied information in the forecast. It is fairly well established in the literature that recent forecasts are more accurate (O'Brien 1988). Gleason and Lee (2003) find that the price impact of forecasts depend on whether the analyst brings new information to the market. Analysts' earnings forecast revisions may bring their previous forecast closer to the current consensus (generally referred to as herding), or they may diverge from the existing consensus. Gleason and Lee (2003) show that forecast revisions are more informative, i.e., elicit bigger price responses, when they diverge from the consensus. Clement and Tse (2005) find that one reason why bold (diverging) forecasts have bigger price impact is because they are more accurate. Further, they find that bold forecast revisions tend to improve the previous forecast of the same analyst to a greater extent than herding forecasts. This is consistent with bold forecasts conveying more of the analyst's private information about the firm. Consistent with the predictions in Trueman (1994), they also find that smaller forecast revisions are more highly correlated with forecast errors (after the revision); herding analysts

revise their forecasts in the direction suggested by their information, but are less likely than bold analysts to fully incorporate this information in their forecasts.

Analysts may have to make a trade-off between timeliness of their forecasts and forecast accuracy, i.e., they could quickly issue forecasts in response to new information, or wait for additional information/analysis to provide more accurate forecasts. Cooper, Day and Lewis (2001) study the market response to forecast revisions by lead and follower analysts in one high-tech industry (semiconductors and printed circuit boards) and one low-tech industry (restaurants). They find that the price response to forecast revisions of lead analysts, defined as analysts who provide timely forecasts, in both industries is higher than the price response to follower analysts. Their results also suggest that timeliness is valued more by investors than ex post accuracy in the forecasts. Mozes (2003) also finds some evidence that timely forecasts are less accurate ex post, but do improve on the accuracy of the existing forecasts. From an investor usefulness standpoint, however, the timeliness of the forecast seems to be at least as important a criterion as forecast accuracy.

Together these results suggest that while accuracy may be an important criterion in evaluating forecasts, analysts could easily improve their accuracy by herding with (or improving on) existing forecasts of superior analysts. Thus studies that evaluate analyst forecast accuracy using *ex post* data may identify analysts as "superior" in terms of forecast accuracy, but these analysts may actually not be bringing any new information to the market place. Reliance on ex post accuracy measures may also explain why forecast accuracy increases as the age of the forecast decreases. Analysts providing forecasts later in the period have the advantage of observing the predictions of other analysts in addition to other information about the firm. Sinha et al. (1997) recognize the effect of forecast age on accuracy and find that forecast accuracy

differs across analysts, but only after controlling for the relative age of the forecasts. They further find that analysts identified as superior ex ante, on either firm-specific or industry levels, continue to provide more accurate forecasts in the subsequent year. However, Sinha et al. do not control for forecast age in their *predictive* tests, which would affect their conclusions if for example, superior forecasters systematically issue forecasts later in the holdout period. Interestingly, they find that inferior analysts do not provide poorer earnings estimates over the next year. However, their evidence is consistent with analysts differing in earnings forecast accuracy even after controlling for *ex post* bias in identifying accurate forecasters.

If analysts have superior information and bold forecasts are valued more by investors, why do some analysts choose to herd (and not fully convey their private information)?[8] Trueman (1994) suggests that forecast boldness is related to analysts' self-confidence. Analysts who have confidence in their forecasting abilities are more likely to issue bold forecasts while analysts who have lower confidence in their abilities are likely to herd. Hong et al (2000) find that analysts with less experience are more likely to herd, suggesting that career concerns may inhibit analyst boldness. Clarke and Subramanian (2006) argue that prior forecasting performance should be related to the degree of boldness in future forecasts. Specifically, they analytically demonstrate a U-shaped relation between forecast boldness and prior forecasting performance, i.e., analysts with either relatively good or poor performance in the past are more likely to issue bold forecasts than analysts whose past accuracy is closer to the average. This relation is driven by compensation and career concerns and is unaffected by the ability of the analyst or the quality of the analyst's private information. Consistent with their model's

[8] It is also possible that analysts issue similar forecasts (i.e., appear to herd) because they possess the same information. Welch (2000) in a study of analyst recommendations finds some evidence that herding towards the consensus does not appear to be information driven. Specifically, he finds that analyst recommendations do not herd to the consensus any stronger when the consensus recommendation turns out to be a good predictor (ex post) of future stock returns.

predictions, they find that I/B/E/S analysts (1988-2000) with relatively superior or poor prior performance are more likely to issue bold forecasts than analysts in the intermediate group.

Another reason why analysts may herd is because they may be concerned about their reputation, or their private information may be inconsistent with contemporaneously available public signals (Graham 1999). It is also possible that more uncertainty regarding a firm's future performance may lead to herding among analysts. An interesting extension of this stream of research will be to examine whether forecasting difficulty is associated with herding behavior. For example, is herding behavior more prevalent for firms that have higher variance in actual earnings? Higher dispersion in analysts' forecasts (which is inverse of herding behavior) is generally seen as an indication of analyst uncertainty with respect to firm earnings. It could well be the case that analysts may choose to herd when earnings are more uncertain, leading to lower dispersion in forecasts for firms with less predictable earnings.

3.2.1.2. Characteristics of the forecaster

Given that more accurate forecasts are more value relevant, identifying expert forecasters is a profitable strategy for investors. Characteristics that are associated with analyst superiority should also be of interest to brokerage houses (employers), in trying to enhance the quality of their output. If the quality of analyst forecasts and recommendations differ systematically based on analyst characteristics, then researchers could also use these characteristics to compute superior measures of earnings expectations.

One way to identify expert forecasters is by relying on external agencies, such as Institutional Investor (II), which annually recognizes, by industry, expert forecasters as All Star analysts. In fact, Stickel (1992) shows that II All Stars tend to outperform other analysts in terms

of forecast accuracy. Alternatively, one could identify expert analysts by studying analyst characteristics including the brokerage firm the analyst represents. Mikhail et al. (1997) use the Zacks Investment Research database and document that analysts' forecast errors decrease as their firm–specific experience increases, consistent with analysts learning over time. They also find that experienced analysts' forecast errors are more closely related to the market reaction around earnings announcements than the forecast errors of less experienced analysts. This is consistent with the market understanding the effect of experience on accuracy. However, they do not find consistent evidence that analysts with greater firm-specific forecasting experience also issue more profitable stock recommendations.

Using analyst forecast data from I/B/E/S, Clement (1999) examines the association between analysts' experience, affiliation, and specialization on their forecast accuracy. Like Mikhail et al. (1997), he also finds that analysts with more experience are more accurate in their earnings forecasts. Further, he finds that analysts affiliated with large brokerage houses, and who cover fewer firms and industries provide more accurate forecasts.[9] Thus resources available to the employer and specialization also seem to increase forecast accuracy.

Jacob et al. (1999) argue that certain analyst-company alignments (which they refer to as "analyst aptitude") may be more successful (in terms of accurate forecasting) because some analysts may have a natural aptitude in forecasting earnings for particular firms. Increases in firm-specific experience, as measured by the length of time over which analysts have made earnings forecasts for a firm, may merely be a manifestation of analysts' ability to better forecast earnings for that firm. They reexamine Clement's findings and find that the positive association

[9] Mikhail et al. (1997) do not find improvement in forecast accuracy for analysts following fewer industries. However, they use Zacks' forecasts whereas Clement (1999) uses I/B/E/S forecasts. Mikhail et al. define industry concentration as the number of firms followed by an analyst in the same 2-digit SIC as the subject firm, divided by the total number of firms followed by the same analyst on Zacks. On the other hand, Clement's measure of industry concentration is the number of 2-digit SICs for which an analyst issues forecasts on I/B/E/S.

between experience and forecast accuracy diminishes after controlling for analyst-company alignment. They conclude that it is more likely analysts' aptitude rather than overall experience that explains analyst forecasting superiority. They also find that analysts are more accurate in firm-quarters where they make more frequent forecasts.

It is certainly useful to understand *why* certain analysts are superior forecasters than others. However, if investors are merely interested in identifying superior analysts (in terms of forecasting accuracy), gathering information on these various analyst and brokerage firm characteristics may not be beneficial, especially if other (simpler) avenues to predict future performance are available. Brown (2001b) shows that a simple model that only takes into account past forecast accuracy of analysts does at least as well as the more complicated analyst characteristics model (Clement 1999) in predicting future forecast accuracy. Brown and Mohd (2003) also find that forecast age is as good a predictor of forecast accuracy, and is equally representative of investor expectations, as the five analyst characteristics model in Clement (1999). Further if as Cooper et al. (2001) show, timeliness is more important than accuracy when it comes to investment decisions, the importance of studying forecast accuracy for economic reasons may be diminished.

Analyst characteristics that are associated with forecast accuracy may also be important for brokerage houses that employ analysts. Notwithstanding the fact that brokerage houses have more information about the analyst (other than just forecast accuracy), Mikhail et al.'s (1999) finding that analyst turnover is higher when analysts are inferior relative to their peers suggests that brokerage houses do evaluate analysts based on forecast accuracy. This finding together with Jacob et al.'s (1999) finding that certain analyst-company alignments are more successful than others suggests that even if employers do not terminate inferior analysts (based on relatively

inaccurate forecasts for certain firms), the analyst could be reassigned to cover other firms. An interesting extension of the Jacob et al. finding will be to examine whether analysts who are accurate forecasters for certain companies but are not as accurate with others, continue with the same brokerage house but get reassigned out of the companies in which they are not very accurate.[10]

Another interesting issue for future research is to examine why certain firms assign their analysts to cover more firms and industries when analyst accuracy is improved by following fewer firms and industries. Clement (1999) reports that analysts in the 90th percentile of his sample cover 21 (seven) more firms (industries) than analysts in the 10th percentile. While a quick explanation is that these are most likely smaller brokerage firms that employ fewer analysts, the role of such "over worked" analysts in the market is still an interesting issue. Specifically, what is the role of these "inferior" analysts when other, presumably superior, analysts cover the same company for larger brokerage houses? Do investors respond to the forecasts of the inferior analysts, and if so, why?[11]

3.2.2. Dispersion in analyst forecasts as a measure of investor uncertainty

Forecast dispersion (measured as the standard deviation in analyst forecasts), which is a signal of the extent of analyst disagreement about a firm's upcoming earnings, is generally used as a proxy for investor uncertainty prior to information events. The reasoning (similar to the reasoning for using the mean/median earnings forecast as the market's *expectation* of earnings)

[10] Hong and Kubik (2003) provide some evidence on this issue. They find that relatively inferior analysts are more likely to be taken out of covering prestigious stocks (defined as stocks followed by at least 20 analysts and/or with market cap over $5 billion) even if they continue to work for the same brokerage firm.

[11] Mikhail et al. (2004) find that stock recommendations of superior analysts (measured based on profitability of prior year recommendations) continue to be more profitable than that of inferior analysts. However, even stock recommendations of historically inferior analysts are profitable (see their tables 3 and 4).

is that disagreement among analysts reflects general disagreement among investors. Based on the notion that investor disagreement is one factor that triggers trading in stock markets, forecast dispersion is generally used to study trading volume around information events such as earnings announcements.

Barron (1995) builds on prior research that had documented a positive relation between trading volume and the level of prior dispersion in analysts' forecasts as well as changes in analyst forecast dispersion (e.g., Ziebart 1990). He suggests that even if there is no change in the level of dispersion, trading may result because analysts change their relative positions from one forecast period to the next. He refers to this reordering of analyst beliefs as "belief jumbling." Using I/B/E/S annual forecast data from 1984-1990 he finds that belief jumbling is positively related to monthly trading volume even after controlling for other variables that have been posited to be related to trading volume such as the absolute price change (Abarbanell et al. 1995). Consistent with prior research, he also finds that both the level of prior dispersion in analysts' forecasts and changes in analyst forecast dispersion are positively related to monthly trading volume.

While Barron (1995) focuses on the general relation between monthly trading volume and monthly disagreement measures, Bamber et al. (1997) extend the analysis to earnings announcement periods. They restrict their sample to firm-quarters (I/B/E/S forecasts; 1984-1994) where at least five analysts issue annual earnings forecasts in the 45 days preceding interim announcements and also revise their annual earnings forecasts in the 30 days following the announcement. Bamber et al. find that all three measures of investor disagreement identified in Barron (1995) are also positively related to abnormal trading volume around earnings announcements. Prior research suggests that earnings announcements generally resolve pre-

announcement uncertainty; therefore, trading around earnings announcements should be related to prior dispersion in beliefs (i.e., forecast dispersion prior to the announcement). Bamber et al. show that earnings announcements may actually trigger further trading through two other mechanisms: belief jumbling and increases in forecast dispersion from before to after the announcement.

Barron et al. (1998) analytically show that the mean forecast error, together with forecast dispersion and number of forecasts, can be used to estimate analysts' total uncertainty as well as their consensus (common uncertainty relative to total uncertainty). Specifically, they show that higher dispersion implies higher total uncertainty but lower consensus. More importantly, they show that forecast dispersion is a measure of analysts' idiosyncratic uncertainty and therefore does not fully capture total earnings uncertainty. Total uncertainty is a combination of the common uncertainty shared by all analysts and forecast dispersion. They point out that decreases in forecast dispersion (after for example, earnings announcements), may not signal a decrease in overall uncertainty but rather a decrease in uncertainty related to the idiosyncratic component of analyst forecasts.

Barron et al. (2002) use the Barron et al. (1998) measures of consensus and uncertainty and examine the dispersion in analysts' annual earnings forecasts before and after interim earnings announcements to test whether earnings announcements spark gathering of private information by analysts (as observed through their forecast revision activity). Prior research suggested that release of public information such as earnings announcements may *decrease* the need for private information gathering. Barron et al. (2002) find that consensus among analysts actually decreases in the days following the earnings announcement consistent with analysts embedding more private information into their forecast revisions (for example through their

individualistic interpretation of the earnings news). They also show that the decrease in consensus is related to the number of analysts revising their earnings forecasts after an announcement which could explain the increased revision activity documented by Stickel (1989) following interim earnings announcements. Barron et al. (2005) further extend this line of research and find that trading volume around earnings announcements is related to the extent of private information gathering around earnings announcements.

Diether et al. (2002) use I/B/E/S data and find that stocks with high (low) earnings forecast dispersion earn negative (positive) returns in the subsequent month. The difference in returns between stocks in the highest and lowest quintile of forecast dispersion is 9.48% (annualized). They also find that the forecast dispersion effect is strongest for small stocks, even though return differences are observed in larger stocks as well. They interpret their results as consistent with Miller's (1977) prediction that in the presence of short sale constraints, investor disagreement will result in share prices reflecting the most optimistic valuations. This overvaluation of high dispersion stocks leads to negative returns in subsequent periods. Consistent with the return evidence, Diether et al. also find that analyst optimism is strongest when disagreement among analysts (forecast dispersion) is high. They conclude that their evidence is consistent with forecast dispersion being a proxy for investor disagreement and not as a proxy for risk. Forecast dispersion is negatively related to returns; if dispersion is a proxy for risk, dispersion should be *positively* related to future returns.

Johnson (2004) suggests that the finding in Diether et al. is consistent with a standard asset pricing model where forecast dispersion proxies for uncertainty about an upcoming signal of the value of the underlying asset (i.e., current earnings). He argues that this effect should be most evident in highly leveraged firms where the (option) value of equity should increase with

uncertainty, leading to lower returns in the future. Results based on IBES forecast data are consistent with the predictions; the negative relation between forecast dispersion and future returns documented by Diether et al. exists only for firms with risky debt. Alternatively, Chen and Jiambalvo (2005) argue that the findings of Diether et al. may be driven by post earnings announcement drift that has been well documented in the accounting literature. Higher forecast dispersion is typically associated with poor earnings performance, which is followed by negative price drifts.

Findings from the forecast dispersion studies described above suggest more interesting avenues for future research. For example, in their model of analyst uncertainty, Barron et al. (1998) assume that the precision of private information is the same across all analysts. This assumption appears too restrictive when we consider the empirical evidence regarding the distribution of analyst forecasts for a firm. It will be interesting to examine implications for analyst uncertainty and market trading if this assumption is relaxed and the precision of private information is allowed to vary across analysts.[12] Second, more research is necessary in tying the Barron et al. (1998) uncertainty measures to disclosure practices of firms. For example, do firms that have a reputation of providing higher quality disclosures have higher precision of common information and higher level of consensus as defined by Barron et al.? In an interesting foray in this direction, Byard and Shaw (2003) find that analyst forecast distributions for firms with a reputation for providing higher quality disclosures reflect a greater precision of both analysts' common and idiosyncratic (private) information.

3.3 Information content of analysts' research output

[12] In a recent working paper, Gu (2005) relaxes this assumption and provides generalized measures of the analysts' common and private information based on observable forecasts.

This section addresses questions related to informativeness of output from analysts' research. The output variables include earnings forecasts, target price forecasts, stock recommendations, and other information rationalizing the forecasts and recommendations. We first discuss research related to the information content of earnings surprise measured with reference to analysts' quarterly earnings forecasts. Then section 3.3.2 discusses research related to the informativeness of analysts' earnings forecast revisions. Section 3.3.3 discusses research related to the combination of all four output variables, and section 3.3.4 discusses research relying on analysts' long-term earnings growth forecasts.

3.3.1 Information content of analysts' earnings forecast errors

The association between returns and analysts' earnings forecast errors depends on the degree to which the forecast proxies for the market's expectations at the beginning of the return accumulation period, the value relevance of the earnings variable, and the model translating current earnings into value. An interesting puzzle emerged with evidence in O'Brien (1988) suggesting that, although more accurate than time-series model earnings predictions, analysts' quarterly earnings forecasts are not necessarily better proxies for the market's earnings expectations.[13] As noted by Brown (1993) this raises the question as to why accuracy and association approaches to comparing alternative earnings forecasts are not "two sides of the same coin." Wiedman (1996) builds on Brown et al. (BRS; 1987) to address this issue.

BRS argue that the analyst's forecasting advantage (relative to time-series models) should increase with the dimensionality of the information set (i.e., the number of available information signals beyond the time-series of historical earnings), and decrease with both the

[13] Although outside the scope of this paper, Foster (1977) similarly notes the curious result that simpler time series models of quarterly earnings (e.g., seasonal random walk) produce forecasts that outperform more descriptive models (e.g., seasonally differenced first order autoregressive) in producing forecasts that conform more closely to the market's earnings expectations. Bernard and Thomas (1989, 1990) and others build on this evidence by demonstrating profitable trading strategies that take advantage of this apparent inefficiency.

variance of the signals as unbiased estimators of the target earnings variable (i.e., signal imprecision) and the correlation (or common information) among the signals. Using a more recent sample (1988-91 versus 1977-82) of over 19,000 firm-quarters of I/B/E/S earnings forecasts, Wiedman confirms the BRS results that analysts' forecasting advantage increases with firm size (the proxy for dimensionality) and decreases with the dispersion in analysts' forecasts (the proxy for signal variance). Also similar to BRS, Wiedman fails to find the hypothesized positive relation between superior analyst accuracy and number of lines of business (proxying for lack of correlation among signals). In fact, she observes a significantly negative effect on both the accuracy and market expectations proxy dimensions of analysts' forecasting advantage. The interpretation of the apparently negative effect of lines of business on analyst's forecasting advantage awaits further research. Nevertheless, Wiedman's results support the notion that evaluations of earnings forecasts on the accuracy and association with returns dimensions are two sides of the same coin (p. 323).

In contrast, Walther (1997) finds no relation between forecast accuracy and the degree to which analysts' earnings forecasts proxy for the market's earnings expectations. Walther (1997) finds that the superiority of Zacks analysts' forecasts as proxies for the market's earnings expectations increases with firm size (and analyst following). However, Walther attributes this result to effects of investor sophistication rather than increased forecast accuracy. The results generally support her prediction in that for firm-quarters with high (low) investor sophistication, the mean 120-day analyst forecast (seasonal random walk forecast) serves as a better proxy for the market's earnings expectations. For example, she finds that the ratio of the ERC based on the analyst forecast error as the measure of earnings surprise to the ERC based on the forecast error from a seasonal random walk is 0.56 (1.83) in the lowest (highest) institutional ownership

quintile, where the lowest (highest) quintile corresponds to a mean of 11% (70%) of company shares held by institutional investors. Unlike Wiedman's investor sophistication proxies, *ex post* forecast accuracy does not explain the degree to which one forecast or the other reflects the market's earnings expectations. Thus, the puzzle remains: Why does the most accurate model not consistently produce the forecast most associated with the market's earnings expectations? Another question emerges from Walther (1997): If not for accuracy, why would a more sophisticated investor be more likely to rely on sell-side analysts' earnings forecasts? Whether and to what degree factors, in addition to (or instead of) forecast accuracy affect the marginal investor's reliance on one model or another to form earnings expectations remains an interesting avenue for further research.

Effective market association tests require an understanding of the valuation implications of unexpected earnings. The market impact of an earnings surprise depends on its persistence (e.g., Kormendi and Lipe, 1987; Easton and Zmijewski 1989), and persistence varies with the nature of the earnings components that create the surprise. In the mid-1990s analysts' revenue forecasts (and actuals) became readily available to academic researchers through I/B/E/S. An interesting question is whether the nature of the revenue surprise informs the market as to the expected persistence of the earnings surprise and, in turn, affects the market's response to the earnings surprise. Ertimur and Livnat (2003) provide evidence on this issue. In a sample of over 20,000 firm-quarters with I/B/E/S consensus revenue and earnings forecasts spanning the years 1994 to 2001, they find that the market responds positively (negatively) to increases in earnings of glamour firms when those earnings are driven by revenue increases (expense decreases that outweigh revenue decreases). On the other hand, the market does not disproportionately reward value firms' earnings increases driven by revenue increases. Ertimur and Livnat also find that

the market looks to the components of earnings to estimate effects of earnings management. They find that the market views suspiciously the earnings of a firm barely meeting analysts' forecasts while reporting disappointing revenues. Thus, details of the source of the earnings surprise explain differences in the intensity and direction of the market's response.

In addition to revenues, researchers might begin examining the information content of other components of analyst-based measures of earnings surprise. For example, cash flow forecasts are becoming more available on I/B/E/S, especially for international firms (DeFond and Hung, 2007). From the difference between analysts' earnings and cash flow forecasts, researchers might infer analysts' accruals forecasts, which can serve as a benchmark from which to measure unexpected accruals at earnings announcement dates. For example, Melendrez et al. (2005) derive unexpected accruals in this manner and report that the market overprices accruals, particularly for loss firms.

3.3.2 Analysts' earnings forecast revisions

In tests of the information content of analysts' forecasts of upcoming quarterly or annual earnings, the key variables are the market's RESPONSE and the NEWS the analyst brings to the market. RESPONSE is typically measured as market-adjusted returns around the forecast announcement date, and NEWS is measured as the analyst's new earnings forecast (F) minus a proxy for the market's expectations of that forecast immediately before the announcement date [i.e., NEWS=F-E(F)]. The literature includes a curious regularity indicating that the analyst's own most recent (i.e., current outstanding) forecast of the target earnings variable is the best proxy for E(F) (e.g., Stickel, 1991; Gleason and Lee, 2003). However, other papers use different proxies for E(F) but do not compare these proxies to the analyst's own most recent forecast to see which NEWS variable is most associated with returns (e.g., Chen et al. 2005).

Measurement error in the NEWS proxy potentially creates ambiguities in cross-sectional comparisons of the information content of forecast revisions. For example, Clement and Tse (2003) rely on the analyst's own most recent forecast as the proxy for market expectations and find lower market response to low innovation revisions (i.e., those moving towards the consensus). However, this result is consistent with the interpretation that the analyst's prior forecast is a particularly poor proxy for the market's expectations in "low innovation" situations and therefore the measurement error in the NEWS proxy is large and drives the market's response coefficient towards zero. More research is needed to better understand how the market forms its expectations regarding the timing and magnitude of an individual analyst's next earnings forecast.

A number of recent papers contain results suggesting that market participants care about analyst forecast accuracy (see section 3.2). This implies that the intensity of the market's response to earnings forecast revisions should increase with analyst forecasting superiority. With reference to a sample of over 15,000 analyst-firm-year observations from I/B/E/S detail files, Park and Stice (2000) find a positive correlation between past forecast usefulness and the market's response to individual analysts' forecast revisions. Williams (1996) introduces the usefulness concept and Park and Stice operationalize it as the percentage of times in the last two years that the analyst's forecast proved more accurate than the prevailing consensus. Interestingly, Park and Stice (2000) do not find a spillover effect of forecasting superiority from one firm to other firms followed by the same analyst. The authors interpret these results to suggest that analyst forecasting superiority stems more from access to inside information from management than from superior ability to analyze commonly available information. An interesting extension of Park and Stice (2000) would be to see if changes in the information

environment after Reg. FD and Sarbanes-Oxley affect the source of superior analysts' forecasting advantage.[14]

So, do investors recognize analyst characteristics that affect forecast accuracy and react accordingly to forecast revisions? In an application of the Brunswick lens model to an archival study (using Zacks data including over 200,000 analyst-firm-quarters spanning 1981-1999), Bonner et al. (2003) find that firm-quarters characterized by relatively high investor sophistication are also characterized by market response to analysts' quarterly earnings forecast revisions reflecting better understanding of *ex ante* expected analyst forecast accuracy. [15] Investor sophistication is associated with price responses that better reflect forecast age and prior forecast accuracy as the two dominant *ex ante* predictors of analyst accuracy.[16]

Chen et al. (2005) find that the market's response to analysts' forecast revisions is consistent with investors learning about analysts' forecasting ability in a Bayesian fashion as more observations of past forecast accuracy become available. However, the key variable in their study, the number of observations available to measure the analyst's average prior accuracy, is highly correlated with experience, a variable that prior research finds is associated with individual analyst forecast accuracy.[17] Further, the paper considers empirical proxies for the market's prior assessment of the analyst's forecasting ability (i.e., size of the brokerage's

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[14] Also see Ivkovic and Jegadeesh (2004) who "find that earnings forecast revisions are least informative in the week after an earnings announcement and that the information content generally increases over event time." They also find "a sharp increase in the information content of upward forecast revisions and recommendation upgrades in the week before earnings announcements, but ... do not find a similar increase for downward revisions or for recommendation downgrades."

[15] Investor sophistication is proxied using factor analysis of analyst following, institutional interest, and dollar value of shares traded.

[16] As discussed in section 3.2, Brown and Mohd (2003) and Brown (2001b) find that forecast age, and past forecast accuracy, respectively, predicts future forecast accuracy at least as well as more sophisticated models. Findings in Bonner at al. are consistent with sophisticated investors recognizing this.

[17] Chen et al. address this issue in two ways. First, they control for the effect of experience on *past* forecast accuracy. However, this analysis does not control for the effect of experience on *expected* forecast accuracy. Second, they find no evidence that analyst forecast accuracy improves with experience. However, these results are at odds with research showing that forecast accuracy is correlated with experience (Mikhail et al. 1997; Clement, 1999).

research department, brokerage house reputation and earnings predictability of the followed firm), but does not include variables to proxy for the precision of the market's prior earnings expectations. A likely candidate is the dispersion of the analysts' forecasts in the consensus that proxies for the market's prior earnings expectation (i.e., the benchmark from which the quantity of news is measured). Future research should consider the degree to which the market's response to analysts' forecast revisions depends on the precision of the market's prior earnings expectation, and should also include controls for effects of analyst experience on expected forecast accuracy.

While Chen et al. (2005, p. 5) interpret their results to indicate that investors "process information rationally and in a sophisticated manner", Clement and Tse (2003, p. 242) "conclude that return responses to forecast revisions are *inconsistent* with investors' use of rational expectations of forecast accuracy ..." (emphasis added). Using a sample of over 35,000 real time current year earnings forecast revisions from the First Call database (1994-98 sample period), Clement and Tse find that: individual analyst forecast accuracy increases with the analyst's prior year forecast accuracy, the size of the analyst's brokerage house, the number times the analyst issues a forecast for the firm in the current year (forecast frequency), and the number of prior years in which the analyst has issued a forecast for the firm (firm experience); and accuracy decreases with the forecast horizon and the number of companies and industries the analyst follows. These results are consistent with prior research. Clement and Tse also introduce a new variable, 'days elapsed since the most recent forecast' by any analyst following the firm, which they find is negatively related to forecast accuracy.

Consistent with their market irrationality inference, Clement and Tse's primary findings are: (1) the market attaches significantly more (positive) weight on brokerage house size than the

weight estimated in their forecast accuracy prediction model; and (2) the market weighs the forecast horizon and 'days since last forecast' variables positively, whereas the accuracy prediction model weighs them negatively.[18,19]

Variables omitted from the Clement and Tse study include "forecast immediacy" and "forecast usefulness." Mozes (2003) conceptualizes "forecast immediacy" as "the speed with which analysts respond to a significant change in the publicly available information set." Mozes predicts (and finds) that his proxies for forecast immediacy are positively related to forecast dispersion (i.e., uncertainty), negatively related to forecast accuracy, and *positively* related to forecast usefulness. Mozes (2003) measures usefulness as the relative improvement in forecast accuracy (i.e., resolution of uncertainty) from one forecast (or forecast cluster) to the next. Longer horizon forecasts and forecasts issued after more time has elapsed are likely to have greater forecast immediacy and usefulness. For example, longer horizon forecasts are those issued relatively early in the sequence and are therefore likely to be associated with greater forecast immediacy. Similarly, forecasts issued in a sequence following a long period without forecasts are likely to have greater forecast immediacy. In both cases, the analysts issuing the earlier forecasts in the sequence are likely to have been willing to trade off accuracy for timeliness in their responses that significantly inform the market's earnings expectations. This might explain why even though these forecasts tend to have low forecast accuracy (before conditioning on the accuracy of the previous forecast), Clement and Tse find (contrary to their predictions) that they have more impact on returns. Future research should consider resolution

[18] The tests do not convincingly reject the market rationality hypothesis with respect to any of the other characteristics studied, including prior forecast accuracy, forecast frequency, firm experience, and number of companies and industries followed by the analyst.

[19] Bonner et al. (2003) investigate very much the same issue as Clement and Tse (2003) and reach different conclusions. Clement and Tse use a Mishkin procedure to investigate investor rationality; whereas, Bonner et al. (2003) use a Brunswick lens model. Future research should explore the impact of the differences in method on the results of the two studies. Beyond method, the two studies differ in that Bonner et al. condition their analysis on investor sophistication, whereas Clement and Tse do not.

of uncertainty as a key ingredient in explaining the variation in the market's response to earnings forecast revisions.

3.3.3 Analysts' recommendations and research reports

Until very recently, research on the information content of information in analysts' research reports focused on either earnings forecasts or stock recommendations in isolation. Recent papers classify the information in analysts' research reports into four categories: earnings forecasts, target price forecasts, investment recommendations, and conceptual arguments supporting the forecasts and recommendations. While some view the recommendation as the bottom line of the research report (e.g., Schipper 1991), analysts' recommendations generally fall into three categories, strong buy, buy and hold, with sell and strong sell categories used much less frequently (McNichols and O'Brien, 1997). These coarse gradations leave room for the other variables in the report to provide information to investors over and above, or in some cases subsuming, the information in the summary recommendations. The four papers (one experimental and three archival) we discuss below provide examples of emerging research in this area.

Francis and Soffer (1997) examine 556 analyst research reports available in the Investext database between 1989 and 1991. They find that 3-day returns centered on the report announcement date are significantly associated with both the recommendation and earnings forecast revisions. The evidence suggests that the market responds more strongly to earnings forecast revisions accompanied by buy (versus hold or sell) recommendations. The authors argue that because analysts bias recommendations upward, investors turn to earnings forecast revisions for more information when analysts issue buy or strong buy recommendations. Hirst et

al. (1995) make the opposite argument. They suggest that doubting the veracity of the recommendation should carry over to other information in the research report. They predict that only when recommendations are unfavorable or unexpectedly revised downward will investors expend the effort to analyze any of the information in the report and impound that information in their decisions. Hirst et al. find support for their prediction in an experiment using student subjects. In an archival study, Asquith et al. (2005) also find evidence consistent with the Hirst et al. prediction. They find a stronger association between the strength of analysts' remarks and returns around the release of analyst reports containing recommendation downgrades as compared to reiterations or recommendation upgrades.

To reconcile the findings in the three studies described above, we offer a slightly different perspective on investor perceptions of information credibility. All three studies consider investor response to information incremental to the recommendation. However, the incremental information variable in Francis and Soffer (1997) is the earnings forecast revision; whereas, Asquith et al. (2005) and Hirst et al. (1995) consider strength of arguments variables. Analysts' reputations depend directly on their earnings forecast accuracy (e.g., Mikhail et al., 1999), and records of forecast accuracy are carefully maintained by interested observers; whereas, accuracy regarding the strength of the analyst's arguments is harder to verify. For these reasons, investors may view earnings forecast revisions as more credible than the strength of analysts' remarks in support of buy recommendations. On the other hand, given analysts' incentives to bias recommendations upward, investors may attach more credibility to analysts' *arguments* in support of hold and sell recommendations. Further empirical research (both experimental and archival) can enhance our understanding of the interaction between the type of

recommendation and investor use of the various types of other information in analysts' research reports.

Brav and Lehavy (2003) extend this literature by adding analysts' target price forecasts to earnings forecasts revisions and stock recommendations. They find that all three variables contribute significantly to explaining variation in returns around research report dates.[20] While earnings and target price forecast revisions are most significant for reiterations, the coefficients on these variables are also significantly positive in cases of recommendation upgrades and downgrades. Brav and Lehavy find that only two-thirds of all analysts' reports include target prices, and reports containing buy or strong buy recommendations are more likely to contain target price forecasts. The authors speculate that analysts may provide target prices to stimulate the purchase of stocks in conjunction with their buy recommendations. Short-selling restrictions constrain trading commissions associated with target price forecasts to stimulate sell recommendations, and the authors suggest that lowering price targets to stimulate sell orders could jeopardize already strained relations with managers of the followed firms. These conjectures warrant examination in further research.[21] Brav and Lehavy also find that when the analyst revises the recommendation in a direction opposite to (same as) the direction of the target price revision, the association between returns and the recommendation revision declines (increases) dramatically. In addition, the evidence indicates a significantly larger market response to target price forecast revisions accompanied by downward (versus upward) earnings forecast revisions. Why this is the case and, more generally, understanding the interactive

[20] Brav and Lehavy (2003) obtain all earnings forecasts, target price forecasts and recommendations from the First Call real time database. The sample includes over 70,000 research reports during the years 1997-99.

[21] Reports with reiterations are also more likely than other report types to contain a target price forecast, suggesting that analysts use their target price forecasts to provide finer gradations of information than they are able to provide with their recommendations. Similarly, Asquith et al. find a significant market reaction to target price forecast revision and strength of arguments variables included in reports with reiterations of stock recommendations. Thus, reports contain valuable detailed information even when the analyst does not see fit to change the recommendation.

effects between all combinations of the three variables warrants further research. Given the peculiar time period of the study (1997-99) and difficult research design issues associated with the earnings forecast revision variable (see section 3.3.2), further research is needed to validate the Brav and Lehavy findings that target price forecast revisions contain information incremental to concurrently released earnings forecast revisions. If investors have more information from alternative sources about earnings than about target prices, the forecast revision variable may contain less surprise than the target price revision variable. Furthermore, the Brav and Lehavy regression assumes a constant revision response coefficient (RRC) across all firms. Simply allowing for industry differences could generate a stronger RRC and leave less unexplained variation for the target price forecast revision variable to pick up in terms of incremental analyst expertise.

The two most important/prominent summary statistics associated with equity securities are earnings per share and stock price. Studies like Brav and Lehavy (2003) of the informativeness of target price forecast revisions, conditional on the informativeness of earnings forecast revisions, potentially provide insight into analyst expertise in modeling the relation between earnings and returns. Opening the black box containing the process by which analysts convert earnings forecasts into price forecasts could provide interesting insights into valuation models most relevant to investors and to the allocation of scarce resources in capital markets. However, the persistent explanatory power of the earnings variable with the target price variable in the regression suggests that the market's translation of earnings forecasts into current equity value differs from the analyst's. An interesting question for future research is why earnings forecast revisions have a significant relation with returns conditional on both recommendations and target prices.

Asquith et al. (2005) evaluate the information content of the strength of analysts' arguments in conjunction with all three forecast and recommendation variables studied by Brav and Lehavy (2003). Asquith et al. develop their strength of arguments variable based on the research team's coding of 14 different types of analyst remarks as either positive or negative in 390 actual analyst research reports issued between 1997 and 1999.[22] A multivariate regression model explains 25% of the variation in returns during 3-day windows centered on the research report announcement date, and most of the explanatory power appears to come from the target price change variable. However, the strength of arguments variable is highly significant, suggesting that market prices are sensitive to detailed information in the analyst's research report, controlling for the information summarized by the other three variables. The market response appears most sensitive to negative analyst remarks regarding the company's prospects for earnings or revenue growth and whether the company met expectations, and to positive analyst remarks regarding the company's new projects, stock repurchases, and risk.

An alternative explanation for the relative importance of the Asquith et al. strength of arguments and target price forecast revision variables is that these represent unique information that the individual analyst brings to the market. The earnings forecast revision and recommendation variables (which appear to be subsumed by the other two variables) are more likely to mimic the forecasts and recommendations of other analysts already in the public domain, particularly for larger firms, and the firm-size effects documented in the study seems to support this explanation.

[22] The small sample (gathered from Investext) is largely due needing the analyst's most recent prior research report for purposes of computing revision variables. Asquith et al. report that during their time period analysts' reports rarely included prior forecasts and recommendations. On the other hand, Francis and Soffer report that about half of the reports in their sample from the years 1989-1991 included the analyst's prior earning forecast and recommendation. This raises the question as to the factors that explain analysts' decisions to include comparison forecasts and recommendations from prior reports. Apparently this decision varies over time and across firm-analyst pairs.

Asquith et al. (2005) also analyze the interactive effect of recommendation type (reiteration, upgrade or downgrade) on the market's response to the four information variables and a variable indicating the existence of an underwriting relationship. From this analysis, it appears that target price forecast revisions and the strength of argument variable matter most in the case of downgrades, and this is the only analysis where the underwriting variable is significant in the expected direction. It appears that the downgrade is only informative when the analyst works for a brokerage with an underwriting relationship with the followed firm. This is consistent with the view that the market considers the underwriting relationship to create incentives for the analyst to bias recommendations upward, and so views a downward recommendation in this situation as very bad news. However, it is unclear why the market would not also consider downgrades as bad news when analysts do not have such incentives.[23] Further research is needed to see if the results hold in larger samples and in a variety of economic conditions (1997-99 were boom years).

Our reading of Asquith et al. (2005) suggests other questions for further research. The authors point out that the earnings forecast revision and strength of argument variables are highly correlated, and "this relation suggests that positive (negative) earnings forecast revisions are generally supported by more optimistic (pessimistic) analyst statements." This begs the question as to the interactive effect of the strength of arguments variable on the market's reaction to earnings forecast revisions. Interactions between other variables would also be interesting to examine. For example, does the market respond more negatively to information in a stock downgrade accompanied by a negative earnings forecast or target price revision than to a stock

[23] The sample size is small, including only 193 observations, and the study is silent on the handling of outliers which could significantly affect interactions depending on very few observations. Brav and Lehavy (2003), for example, winsorize their earnings and target price forecast revision variables at the 1st and 99th percentiles and further examine the regression results for effects of influential observations.

downgrade accompanied by an upward earnings forecast or target price revision? Finally, it is not clear what analysts are trying to communicate through their stock recommendations. In particular, what does a reiteration of a strong buy or a downgrade from a strong buy to a buy really mean? In the 17.6% of the Asquith et al. sample where analysts reiterated a strong buy, the target price forecast increased by only 1%. Why would analysts reiterate a strong buy if they can only increase their target price forecast by 1%? One explanation might be that the price has not yet increased from the last strong buy recommendation and therefore the analyst still views the firm as undervalued. However, abnormal returns around these report release dates are relatively small. Francis and Soffer (1997) find that the change in the recommendation has a significant association with returns after controlling for the level of the recommendation. Future research will perhaps shed more light on the nature of the information in recommendation *changes* that are not subsumed by the information in recommendation *levels*.

3.3.4 Long-term earnings forecasts and earnings-based valuation models

In connection with Ohlson (1995) and the revival of academic and practitioner interest in earnings-based valuation, recent studies examine the usefulness of analysts' long-term forecasts in: explaining current market-to-book ratios (Cheng, 2005a); inferring firms' cost of equity capital (e.g., Botosan and Plumlee, 2005); evaluating the value-relevance of accounting earnings (Liu and Thomas, 2000); evaluating the theoretical assumptions and operational characteristics of the earnings-based valuation model (Begley and Feltham, 2002); and identifying mispriced securities (e.g., Frankel and Lee, 1998). In its most basic forms, the model appears as follows:

$$V_0 = BV_0 + E_0\left[\sum_{t=1}^{T}\frac{EARN_t - r\cdot BV_{t-1}}{(1+r)^t} + \frac{EARN_{T+1} - r\cdot BV_T}{(r-g)\cdot(1+r)^T}\right] \quad (1)$$

$$V_0/BV_0 = 1 + E_0\left[\sum_{t=1}^{T}\frac{(ROE_t - r)\cdot\left(BV_{t-1}/BV_0\right)}{(1+r)^t} + \frac{(ROE_{T+1} - r)\cdot\left(BV_T/BV_0\right)}{(r-g)\cdot(1+r)^T}\right] \quad (2)$$

The first format expresses current equity value as a function of current book value of equity and the present value of all future periods' expected abnormal returns (i.e., earnings in excess of prior period's book value times the equity cost of capital). The second format divides through by current book value of equity to provide an equivalent expression for the estimated current market to book ratio. T refers to the forecast horizon terminal year, after which time abnormal earnings are assumed zero, constant (g=0), growing at a constant rate, or (with slight modification) fading at a constant rate. Researchers combine analysts' earnings and long-term growth rate forecasts, current book value, an assumed dividend payout rate and either terminal value assumptions about g or price forecasts in order to estimate expected abnormal earnings over the forecast horizon and the value of the equity at time 0.

With reference to model (2) above, Cheng (2005a) evaluates the degree to which analysts' earnings and long-term growth rate forecasts captures and adds to the explanatory power of other information variables in explaining firms' market-to-book ratios.[24] With reference to a sample of 6,737 firm-years spanning 1991-2000, the results indicate that current ROE_t alone explains about 19% of the cross-sectional variation in firms' market-to-book ratios. Adding only the analysts' earnings forecasts for years t+1 and t+2 and long-term growth rate forecasts increases the r-square from 0.19 to 0.682, and analysts' earnings and long-term growth rate forecasts parsimoniously reflect 83% of the collective information in the other information

[24] The other information variables include ROE_t (i.e., return on equity for year t), risk and growth proxies (including industry cost of capital, financial leverage and year t sales growth), and various proxies for: the firm's potential to earn economic rents, the degree of conservatism reflected in the firm's current book value, and quality of earnings signals that prior research has found useful in predicting future earnings (e.g., Fairfield et al., 1996; Abarbanell and Bushee, 1997, 1998; Cheng, 2005a).

variables. Furthermore, the analysts' forecasts explain 22% of the variation in market-to-book ratios not captured by the other information variables (whereas, the other information variables only uniquely explain 5.5% of the variation). Cheng's results indicating the usefulness of analysts' forecasts improve even further in a subset of firms with characteristics found by prior research to make current earnings less predictive of future earnings (e.g., firms in high tech industries, firms with large amounts of R&D capital and firms with relatively high abnormal accruals).

Cheng's (2005) results represent very good news for analysts, as two research design issues probably bias tests against finding that analysts' forecasts subsume the other information variables. Cheng gathers consensus analysts' forecasts from the first I/B/E/S report following the year t earnings announcement. As described in O'Brien (1988) and Brown (1993), this consensus is likely to include stale individual analysts' forecasts that are not conditioned on the information in the year t earnings announcement or financial statements. Therefore, it is unlikely to find that analysts' consensus forecasts completely subsume that information. Second, the relation between the other information variables and current market-to-book ratios are estimated in-sample, effectively requiring the analyst to anticipate relationships between future values of those variables and future years' market-to-book ratios. Further research might investigate the effects of these research design issues by: (a) relying on the I/B/E/S detail files to generate analysts' earnings and long-term growth rate forecasts, and (b) estimating the parameters to attach to the other information variables based only on year t and prior years' information. Nonetheless, analysts' long-term forecasts appear to effectively reflect market expectations in the context of explaining market-to-book ratios.

Liu and Thomas (2000) rely on I/B/E/S consensus analysts' earnings forecasts over a 5-

year forecast horizon and consensus long-term growth forecasts to better understand the value-relevance of accounting earnings. The study begins with model (1) above and derives a theoretical relation between unexpected returns, unexpected current year earnings and unexpected revisions in forecasts of future years' earnings. Liu and Thomas find that the r-squares in regressions of annual returns on unexpected earnings dramatically increases from about 5% to about 30% when the forecast revision terms are added.[25] Furthermore, the coefficient relating unexpected returns to the unexpected earnings and forecast revision variables take on estimated values very close to their theoretical values. These results provide evidence confirming a joint hypothesis that: (a) analysts' forecasts effectively proxy for the market's earnings expectations; (b) the earnings target is value-relevant; and (c) the earnings-based valuation model has descriptive validity with respect to capital market price formation. Thus, while analysts' target price forecasts and recommendations may not emerge from a multiperiod valuation model (Block, 1999; Bradshaw, 2002), the Liu and Thomas results suggest that the market combines analysts' earnings and long-term growth forecasts into a multiperiod model to rationalize current price. On the other hand, Frankel and Lee (1998) find that value-to-price ratios derived using a variant of model (1) predict abnormal returns. Thus, the market mechanism that converts forecasts into current equity values remains an elusive concept for future research.

A number of recent studies rely on various forms of model (1) to infer firms' costs of equity capital (e.g., Botosan and Plumlee, 2002; Baginski and Wahlen, 2003; Claus and Thomas, 2001; Easton, 2004; Easton and Monahan, 2005; Gebhardt et al., 2001; and Gode and

[25] Note that the impact of adding analysts' forecast revisions may be even greater using forecast data from the I/B/E/S detail files, since many of the forecasts entering the consensus used in this study may not have been issued after the announcement of the year's earnings and other financial statement information on which they are intended to be conditioned.

Mohanram, 2003). Botosan and Plumlee (2005) conclude that two very simple approaches to estimating the cost of equity capital perform best in terms of the sign and magnitude of multivariate regression coefficients relating the cost of capital proxies to a battery of generally accepted risk factors. The first of the two best behaved risk proxies solves for the cost of equity capital that equates the discounted present value of Value Line's forecasts of the firm's dividends for the next five years and Value Line's target price forecast at the end of those five years with the firm's current stock price. The second well-behaved risk proxy extends from the approach developed in Easton (2004) and simply equals the square root of a fraction equal to Value Line's projected growth in the firm's *EPS* between year's 4 and 5 in the forecast horizon divided by the firm's current stock price.

As pointed out by Botosan and Plumlee, the risk estimation procedures evaluated in their study depend critically on the assumption that analysts' earnings and/or price forecasts mirror the market's expectations. An important corollary to this assumption is that the current stock price mirrors the analyst's assessment of the firm's intrinsic equity value. Since analysts are in the business of identifying mispriced stocks, this corollary is unlikely to hold.[26] Future research might reevaluate the Botosan and Plumlee results on a subsample of stocks with timeliness ranks corresponding to hold recommendations. Moreover, research regarding divergence between analyst and market expectations can help researchers evaluate various approaches to estimating the cost of equity capital, make appropriate adjustments to the forecasts, or choose subsamples where the critical assumption of similar analyst and market expectations most likely holds.

3.4 Market and analyst inefficiency

[26] We are grateful to Jake Thomas for discussions leading us to this insight.

A number of studies have examined analysts' forecasts as a means to understanding the broader issue of whether investors efficiently respond to new information.[27] Analysts have long been seen as sophisticated processors of financial information who are less likely (compared to naïve investors) to misunderstand the implications of financial information. Thus evidence of inefficient information processing by analysts is seen as strong evidence of overall inefficiency by market participants. A second reason to examine analyst forecasts for possible biases is that evidence of market inefficiency based on "abnormal" stock returns is always open to the criticism that the expected return benchmark used in measuring abnormal returns may be misspecified (Fama 1998). Analyst forecasts do not suffer from benchmark issues and thus provide a neat avenue for studying general market inefficiencies.

3.4.1 Short-term forecasts

Kang et al. (1994) examine the rationality of analyst forecasts for a given quarter over different time horizons. Their purpose is to understand whether analysts follow similar forecast rules across different forecasting horizons. Their results, based on Value Line forecasts over multiple quarters from 1980-1985, indicate that the coefficient relating forecasts to actual earnings is significantly less than one in all horizons. More importantly, the coefficient increases as the forecast horizon decreases. They interpret this evidence as inconsistent with analysts following a uniform prediction model over different horizons. They also carefully rule out the possibility that the differing relation between forecast and actual earnings over time is driven by an adaptive process that adjusts to new information. They conclude that the observed forecasting

[27] If analysts efficiently revise their forecasts in response to a piece of new information, then the error in their revised forecasts should be unrelated to that information. A positive (negative) relation between the information item and the revised forecast error will imply under (over) reaction by analysts with respect to that piece of information.

behavior of analysts is more likely because of incentive issues (Dugar and Nathan 1995) or due to cognitive/processing biases.

Elliott et al. (1995) examine sequential forecast revisions of IBES analysts and find that analysts underweight new information, which leads to a systematic positive relationship between their earnings forecast revisions and subsequently realized forecast errors. They find that this positive association between intra-year revisions and year-end forecast errors persist even after controlling for other known factors that affect the forecasting environment such as earnings persistence, analyst forecast dispersion and prior stock price changes.

Easterwood and Nutt (1999) attempt to reconcile conflicting conclusions in prior research of analyst overreaction (De Bondt and Thaler 1990), as well as analyst underreaction to past earnings information (Abarbanell and Bernard 1992). They hypothesize and find that inefficiency in analyst forecasts is not characterized by uniform overreaction or underreaction to information, but is more appropriately described as general optimism. Specifically, analysts seem to overreact (underreact) to good (bad) news in prior year earnings, consistent with incentive-based explanations of analyst optimism (Lin and McNichols 1998; Dugar and Nathan 1995). While the finding of analyst optimism in Easterwood and Nutt is consistent with incentive-driven analyst behavior, the sensitivity of their results to truncation rules suggests the need for future research on this issue.[28]

Dechow et al. (1999) find that analysts fail to fully take into account the mean-reversion of abnormal earnings in their year-ahead earnings forecasts. They find weak evidence that the

[28] Some papers point out that the findings in Easterwood and Nutt do not appear to be robust and are sensitive to how outliers are handled (see, Ahmed et al. 2000, Mikhail et al. 2003). Abarbanell and Lehavy (2003) also caution that tests of over/underreaction by analysts are affected by the distributional properties of analyst forecast errors. In a recent working paper, Gu and Xue (2005) argue that in the presence of high uncertainty, analyst overreaction to extreme good news is rational (rather than a result of cognitive bias). They report that extreme earnings news, whether good or bad, are associated with higher earnings uncertainty. The overreaction to good news documented by Easterwood and Nutt (1999) disappears when they control for the resulting earnings uncertainty.

systematic error in analysts' forecasts is also reflected in stock prices, suggesting that investors do not adjust for the predictable errors in analysts' forecasts. Ettredge et al. (1995) find that analysts' forecast revisions partially adjust for transitory effects of earnings misstatements even before those misstatements are announced. Analysts' forecast revisions around announcements of earnings containing (unknown at the time) earnings overstatements appear to effectively adjusted for an estimated 27% of the overstatement amounts. The authors attribute this result to analysts' gathering other information that causes them to doubt the persistence of the overstated earnings numbers. In contrast to Ettredge et al., Bradshaw et al. (2001) find evidence suggesting that analysts do not fully adjust their forecasts for transitory working capital accrual components of earnings. Specifically, they find a significant negative relation between annual working capital accruals and errors in subsequent consensus analyst forecasts of next year's earnings.

Elgers et al. (2003) extend Bradshaw et al. by evaluating whether analysts' underestimation of the transitory component of current accruals is as severe as the underestimate reflected in market prices. They find that analysts' forecasts can explain at most about 40% of the market's apparent underestimation of the transitory component of current accruals. Thus, analysts at least partially (and more effectively than investors) recognize the difference in persistence of accruals and cash flow components of earnings. Furthermore, the study's estimate of analyst inefficiency is probably overstated, because the forecast variable is the consensus forecast from the I/B/E/S summary file released in April of the year following the fiscal year containing the accruals variable. Given typical March 10-K filing dates for 12/31 yearend firms, many forecasts in the consensus may not have been updated in light of the financial statements containing the accruals variable that is the subject of the study, and this effect is exacerbated by the study's focus on firms with low analyst coverage.

Burgstahler and Eames (2003) study the degree to which analysts successfully anticipate earnings management. The paper develops a conceptual framework where analysts must estimate the point at which it becomes cost effective for firms with premanaged earnings falling below zero to manage earnings upward to avoid reporting a loss. If the analyst believes the firm is at or above this point, then the analyst should forecast earnings at or barely above zero. Different firms have different threshold points where it becomes cost effective to manage earnings to avoid reporting a loss, and analysts are faced with the daunting task of identifying this point on a firm-by-firm basis. The results indicate that the distributions of both forecasted and actual earnings contain a disproportionate number of observations at or barely above zero. This suggests that firms manage earnings to avoid losses, and analysts anticipate this behavior. However, the distribution of forecast errors indicates that forecasts are systematically pessimistic at zero *reported* earnings and systematically optimistic at zero *forecasted* earnings. Thus, analysts do not accurately identify which firms will and which firms will not successfully manage earnings to avoid losses.

Burgstahler and Eames' results suggest a number of avenues for further research. First, it would be interesting to learn more about whether and how analysts attempt to identify firms that will, or will not, manage earnings to avoid losses. Second, it would be interesting to explore the reasons why firms apparently cannot simultaneously manage earnings *and* guide forecasts. Third, as noted by the authors, "future research might examine whether alternative forms of earnings management … are also reflected in analysts' forecasts." For example, Shane and Stock (2006) investigate whether analysts successfully anticipate income effects of incentives for firms to manage earnings in order to take advantage of declining statutory tax rates.

Gu and Chen (2004) hypothesize that if analysts effectively identify temporary components of earnings, then the non-recurring components that they choose to exclude from their forecasts and actual earnings reports will have less persistence and lower price-multiples than non-recurring earnings components that they choose to include.[29] The study's data (spanning 1990-2003) come from the First Call actuals database footnotes, which explicitly describe the items analysts identified as non-recurring and which of these items analysts chose to include and exclude from actual quarterly earnings. It appears that First Call generally obtains its list of potentially non-recurring items from management press releases and then decides which management-identified non-recurring items to include and which to exclude. These decisions conform to the majority views of analysts following the company in that firm-quarter. First Call asks analysts forecasting an earnings number inconsistent with the majority view to adjust their forecasts for the inconsistency. If these analysts do not adjust their forecasts, they are dropped from the database (these drop decisions have implications for consensus forecasts derived from the database and await future research). The results of the study indicate that the non-recurring items First Call chooses to include in actual earnings have greater persistence and higher valuation multiples than excluded items. Furthermore, the authors infer that these higher valuation multiples seem appropriate as neither the included nor the excluded items are associated with future abnormal returns. However, as Bradshaw and Sloan (2002) point out, if special items do indeed explain a large portion of the differences between Street and GAAP earnings, then the evidence in Burgstahler et al. (2002) that the market does not efficiently

[29] Gu and Chen build on Bradshaw and Sloan (2002) which documents a dramatic increase between 1985 and 1997 in: the excess of Street over GAAP *EPS*; earnings press release emphasis on Street (or pro-forma) earnings; frequency and amount of negative special items included in company's income statements; the degree to which these special items explain the difference between GAAP and Street earnings; annual *EPS* growth rates computed with reference to Street versus GAAP *EPS*; and the sensitivity of stock price changes to Street relative to GAAP earnings numbers. The authors interpret their evidence to suggest that managers have taken a proactive role in reorienting the investment community in general and analysts' forecast tracking services in particular towards a focus on Street-based versus GAAP-based *EPS* performance measures.

incorporate negative valuation implications of special items is inconsistent with Gu and Chen's analyst market efficiency conclusion.

The results in Gu and Chen raise an interesting question. As the authors point out, the items analysts choose to exclude, while less persistent and value relevant than included items, still have statistically significant price multiples and are significantly related to future earnings. Furthermore, the persistence and price-multiples associated with included "non-recurring" items are less persistent and have lower price-multiples than "core earnings" which conforms closely to management's view of pro-forma earnings. How and why analysts draw the line in including some items and excluding others remains an issue for further research.

While a number of studies examine analyst efficiency with respect to a firm's own past time-series of earnings, other research has focused on non-earnings information for the firm and/or information external to the firm. Bartov and Bodnar (1994) examine whether analysts adequately account for past exchange rate changes in their quarterly earnings forecasts. They report that errors in IBES analyst forecasts are negatively related to the lagged change of the trade-weighted value of the dollar, consistent with analysts not fully adjusting their forecasts for past exchange rate changes.

Abarbanell and Bushee (1997) examine how analysts use other financial information found to be relevant for predicting future earnings. They use nine accounting related fundamental signals available from annual financial statements and identified in Lev and Thiagarajan (1993). These signals include commonly used ratios in financial analysis such as changes in receivables compared to changes in sales. AB first examine how the non-earnings financial information relates to one and five-year ahead earnings changes and in a second step study how IBES analyst forecast revisions (from 1983-1990) and revised forecast errors are

related to the information. They find that some of the signals are related to year-ahead earnings changes and five-year growth in earnings, even after controlling for the effect of current year's change in earnings. Financial analysts seem to revise their year-ahead forecasts consistent with the information in the fundamental signals. However, even after the revision, analysts' earnings forecast errors are related to the fundamental signals in similar fashion as the year-ahead earnings changes, suggesting that the revisions do not fully incorporate the information about future earnings contained in these signals. Interestingly, they find that once underreaction to the non-earnings financial information variables is controlled for, previously documented analyst underreaction to aggregate earnings changes in the previous year (Ali et al. 1992), is no longer evident in the data. Long window returns (-11, +1; 0 is the earnings announcement month) are associated with both the fundamental signals and the analyst forecast revision consistent with investors adjusting for the information contained in the fundamental signals that are not fully impounded in the analyst forecast revision. Abarbanell and Bushee (1998) follow up on AB and find size-adjusted abnormal returns of 13.2% over 12 months following formation of portfolios of stocks based on the predictions of the nine fundamental signals used in AB. This finding is consistent with investors, much like the analysts in AB, not fully recognizing the information contained in the fundamental signals.

Shane and Brous (2001) re-examine Abarbanell and Bernard's (1992) finding that serial correlation in analyst forecast errors (i.e., analyst inefficiency) explains part of the post earnings announcement drift. They find that analysts seem to use non-earnings-surprise information to correct their underreaction to past earnings and that not controlling for this information overestimates the market's underreaction to the prior quarter's earnings news. They conclude

that a larger portion of the post-earnings-announcement and post-forecast-revision drift is attributable to inefficiency in analysts' forecasts than documented in prior research.

Chaney et al. (1999) examine the effect of restructuring charges on future analyst forecast revisions and revised forecast errors. They find that even though IBES analysts revise their one- and two-year ahead forecasts downward following a restructuring charge, their revisions do not fully take into account the overall negative signal from the restructuring charge. Forecast errors after the revision are still negative, suggesting that analysts continue to be optimistic regarding the earnings of firms that have recently taken restructuring charges.

Ramnath (2002) examines analysts' forecast revision activities following the first earnings announcement in an industry. He finds that analyst forecast revisions for firms that are yet to announce quarterly earnings do not fully reflect the information from the first announcement in the industry. This underreaction to information from other firms leads to predictable stock price reactions around the earnings announcements of the later announcers. Ikenberry and Ramnath (2002) reexamine the well documented long run positive drift in stock prices following stock splits. They confirm previous findings of positive abnormal returns to firms in the year following the stock split announcement (Ikenberry et al. 1996; Desai and Jain, 1997). More importantly, they find that the abnormal returns following split announcements are not spurious, but are at least partially attributable to a gradual positive revision in analysts' earnings forecasts in the months following the announcement.

Teoh and Wong (2002) examine whether analysts systematically misinterpret information in abnormal accruals. Using IBES forecast data they find that analysts tend to over estimate the earnings for stock-issuing firms that have abnormally positive accruals in the year preceding the stock issue. Investors seem to buy into these lofty analyst expectations and subsequent

correction of this overoptimism, leads to underperformance in the returns of issuers over the next four years. While not the focus of their study, TW also find similar analyst behavior in non-issuing firms. Louis (2004) similarly finds that analysts do not completely adjust their post-merger forecasts for acquiring firms' aggressive accrual behavior in the quarter prior to a stock-for-stock acquisition. Results from both studies suggest that even sophisticated financial statement users, like financial analysts, generally fail to recognize the lower persistence of abnormal accruals.

Systematic errors in analysts' earnings forecasts documented thus far could be the product of analyst incentives (i.e., may be rational) or could be attributed to inefficient processing of information by analysts. We defer discussion of research in support of these arguments to section 3.5.

3.4.2. Long-Run Forecasts

Unlike short-term forecasts, where the evidence is mixed as to whether analyst forecasts can be characterized as an underreaction to information, or as overall optimism, there is general consensus that long-term growth forecasts are, on average, overly optimistic. La Porta (1996) finds that firms with high (low) analysts' long-term growth forecasts are associated with negative (positive) future returns in subsequent periods. He interprets this as evidence that analysts' systematic misestimation of future growth explains returns to contrarian strategies. Dechow and Sloan (1997) report that analysts' forecasts of long-term earnings growth are, on average, twice as high as the (ex post) realized growth in earnings. Stock prices initially seem to reflect this optimism in forecasted growth which explains a significant portion of the returns to contrarian strategies. Doukas et al. (2002) report that the year-ahead forecast errors and revisions for high

and low book-to-market firms do not support the argument of naïve extrapolation of growth by investors put forth in La Porta (1996). However, their principal focus is on near-term (year-ahead) forecasts and not on systematic differences in long-term growth forecasts across the different book-to-market groups. Chan et al. (2003) also provide evidence consistent with overly optimistic analysts' forecasts of long-term growth.

Subsequent research has also documented analyst optimism in growth forecasts in different contexts and various sub-groups of analysts. Rajan and Servaes (1997) study analyst following and analyst forecasts in the context of initial public offerings (IPOs). Using data from IBES, they find that analysts' long-term growth forecasts are incrementally optimistic for IPO firms than for peer firms (in the same industry). Interestingly, they also find that when analysts are over-optimistic about growth prospects, more firms come to the market. Thus, firms seem to take advantage of prevailing optimism about their prospects and strategically time their stock offering. Future period returns are lower as investors begin to realize that the long-run growth expectations are unlikely to be achieved.

Purnanandam and Swaminathan (2004) find that IPO firms with high price-to-value (P/V) ratios have higher first-day returns, but significantly lower returns than low P/V firms in the long-run.[30] Analysts' growth forecasts of high P/V firms are much more optimistic than low P/V firms whereas realized future sales growth is no different between the two groups by the fifth year after the IPO. They conclude that investors seem to place too much weight on the (optimistic) growth forecasts of the high P/V firms and too little weight on their current profitability. Dechow et al. (2000) examine both IPOs and seasoned equity offerings and find analyst optimism in long-term growth forecasts for both types of equity offerings. More

[30] P is the offer price of the IPO. Three different measures of V are computed based on sales, EBITDA and earnings multiples of non-IPO industry peers.

importantly, they find that analysts employed by lead underwriters are the most optimistic and the long-run underperformance of new equity issues is related to this over-optimism.

Analysts' forecasts of long-run growth have also come under increased scrutiny with the increasing popularity of the residual income model (Ohlson 1995). Frankel and Lee (1998) use analyst forecasts of annual earnings and long-term growth to compute measures of firm value applying the residual income methodology proposed by Ohlson (1995). They report that the ratio of firm value (V_f) to prevailing market prices (P) is a good predictor of future returns. Specifically, firms with high (low) V_f/P ratios outperform (underperform) the market over the next three years suggesting that investors do not fully utilize the information available in analyst forecasts of future earnings and growth. Frankel and Lee also find that analysts make predictable errors in their longer term (three-year ahead) forecasts; specifically, analysts are overly optimistic in their three-year ahead earnings forecasts for high market-to-book (proxy for growth) firms, and firms with high past sales growth (consistent with Lakonishok et al. 1994). Correcting for these predictable errors leads to additional returns (over and above the V_f/P strategy) over the following three-year period, suggesting that investors fail to correct for predictable errors in analysts forecasts.

Ali et al. (2003) carefully examine whether the V_f/P anomaly documented by Frankel and Lee (1998) can be explained by V_f/P capturing unknown risk factors. They find that even after controlling for known measures of risk (such as B/M, size, stock return volatility, financial distress), the V_f/P strategy yields positive abnormal returns over a three-year period after portfolio formation. Further, a portion of the V_f/P returns are concentrated around subsequent earnings announcements which also suggests that the V_f/P strategy is more likely related to information rather than a premium for systematic risk. Billings and Morton (2001) also report

systematic biases in analysts' long-term growth forecasts. Specifically, they find a positive relation between lagged price changes and analysts' long-term growth forecast errors.

Bandyophadhyay et al. (1995) find that about 60% of the variation in Value Line's price forecasts can be traced to their 3-5 year earnings forecasts. Considering the importance of long-term forecasts to firm value (and stock prices), there has been surprisingly little research on individual analyst forecast accuracy with respect to long-term forecasts. Further, as Dechow et al. (2000, 6) note "analysts are frequently evaluated on the accuracy of their buy-sell recommendations and annual earnings forecasts, but not on their long-term growth forecasts." Thus, both the market place and researchers have largely ignored factors that affect long-term forecast accuracy of individual analysts. Identifying analysts who consistently provide more accurate long-term growth forecasts is an interesting avenue for future research. This should also be appealing to investors given the research findings (mentioned above) that suggest a significant amount of mispricing attributable to overly optimistic long-term growth forecasts. Future research could examine if some of the same characteristics determined by prior research to be associated with superior short-term forecasters (e.g., Clement 1999) also apply to analysts who are more accurate with their long-term forecasts.

3.4.3. Market inefficiency with respect to analysts' forecasts and recommendations

Chan et al. (1996) find that a moving average of the forecast revision for the last six months is a good predictor of firms' returns over the next 6-12 months. This relation between past earnings forecast revisions and future returns persists even after controlling for recent price momentum and prior quarter's unexpected earnings. Their findings are consistent with a market in which analysts are slow to revise their expectations and investors are even slower in responding to both actual and forecasted earnings. Elgers et al. (2001) examine whether the

sluggish response of investors to analysts' earnings forecasts is related to characteristics of firms' information environment. Using early-in-the-year forecasts, they show that the market is slow to react to analyst forecasts for thinly followed firms, which leads to predictable returns in the following months. Elgers et al. provide evidence that the market generally places lower weight on the analyst forecast than warranted; however, the delayed market reaction is most pronounced for firms with low analyst following. This is consistent with information flow to, and assimilation by, the market being slower for thinly followed firms.

Gleason and Lee (2003) find that investor underreaction to analysts' earnings forecast revision is related to the boldness of the forecast. Specifically, they find that the drift in stock prices following forecast revisions are of higher magnitude (in the same direction as the revision) for high-innovation revisions as opposed to low-innovation (or herding) forecasts. Further, the drift is lower for high-innovation revisions of All-Star analysts, suggesting that investors seem to underreact more to bolder forecasts from equally accurate, but lesser known, analysts. Consistent with slower dissemination of information for thinly followed firms, Gleason and Lee also find that the drift is stronger for firms followed by fewer analysts. A significant portion of the drift is concentrated around the next six forecast revisions and subsequent earnings announcements consistent with investors delaying their responses until further confirmation of the information already provided by analysts' forecast revisions.[31]

Barth and Hutton (2004) use the information in past accruals and analyst forecast revisions to study their combined effect on future returns. Previous research had documented abnormal returns to trading on past accruals (e.g., Sloan 1996) and earnings forecast revisions

[31] Clement and Tse (2005) find that bold forecast revisions exhibit less underreaction in the sense that they are less correlated with forecast errors. Based on this finding, they suggest that the result in Gleason and Lee of bold forecasts being associated with more delayed reaction around earnings announcement dates may not be because the analyst under-revised, but more likely because the market does not quickly respond to the analyst revision.

(e.g., Chan et al. 1996). Barth and Hutton find that buying and holding (selling short) stocks that have the lowest (highest) accruals in the prior period and positive (negative) forecast revisions of current year earnings, yields returns of 28.5% over the next year. The combined return is significantly greater than the return to either the accruals or the forecast revision strategy in isolation. Approximately one-half of the annual returns is concentrated around earnings announcements, again suggesting that investors miss the information conveyed by accruals and reinforced through analyst forecast revisions.

Investors have also been shown to be slow in responding to analysts' stock recommendations. Barber et al. (2001) revisit Womack's (1996) finding that stock recommendations of analysts are associated with abnormal returns both around the recommendation period and in the subsequent six (one) months for downgrades (upgrades). They use a calendar time strategy and find significantly positive abnormal returns in the post recommendation period to portfolios formed on the basis of analysts' consensus recommendations drawn from the Zacks Investment Research database. However, in additional tests Barber et al. find that implementing trading strategies based on analysts' stock recommendations involve frequent reshuffling in the portfolio (turnover of over 300 percent annually) which negates the abnormal returns that can be earned through the strategy. However, they use recommendations of all analysts in their analysis and caution that there may be other strategies that could produce abnormal returns net of transaction costs (for example, picking recommendations of analysts/brokerage houses that have superior past performance).

Mikhail et al. (2004) explore this issue. Specifically, they examine whether analysts' whose recommendations yielded positive returns in prior periods, command more attention from investors (i.e., elicit bigger price reactions) when they recommend stocks in the future. They

find that both the market reaction around the recommendation change as well as the post-recommendation drift are higher, in the direction of the recommendation change, for superior analysts (based on past performance). While the abnormal returns to taking a long (short) position in upgrades (downgrades) of analysts with the best prior performance yields positive abnormal returns, it is insufficient to cover round-trip transaction costs. Thus, Mikhail et al. find no evidence in support of one of the conjectures in Barber et al. (2001) that differentiating between analysts may alter their conclusions regarding trading on analyst recommendations.

As discussed in this section, it is fairly well accepted in the literature that the market is slow in responding to analysts' earnings forecast revisions and stock recommendations. Some studies contend that exploiting this market inefficiency is unprofitable because of transaction costs. However, it is still intriguing that investors continue to systematically underreact to a direct signal, like analysts' recommendations and revisions, despite numerous research studies having consistently documented this phenomenon over a number of years.[32] Explaining such (continued) anomalous behavior on the part of investors is a challenging task for future research.

3.5 Analysts' incentives

This section discusses research investigating effects of analysts' incentives on analysts' decision processes and the properties of their forecasts and recommendations. Early work in this area documents what appears to be underreaction in analysts' earnings forecasts and optimism bias in their forecasts and recommendations (e.g., Brown et al. 1985, O'Brien 1988, Womack 1996). Recent work focuses on economic incentives (and/or psychological judgment errors) that

[32] Givoly and Lakonishok (1979) is one of the earliest studies to document predictable stock returns following analysts' earnings forecast revisions.

might create analyst optimism and underreaction. In addition, recent research investigates the impact of management incentives and communications on analysts' forecasts.

3.5.1 Incentives affecting analyst effort and accuracy

Stickel (1992) finds evidence suggesting that forecast accuracy is an important variable influencing *Institutional Investor*'s selection of All-American analysts. Recent studies confirm the link between reputation and forecast accuracy and more broadly identify forecast accuracy as an important variable affecting analysts' careers. For example, Mikhail et al. (1999) find an inverse relation between turnover and analyst forecast accuracy, relative to the accuracy of other analysts covering firms in the same industry. Hong, et al. (2000) confirm that result and also find that forecast accuracy is directly related to the likelihood of promotion, especially for less experienced analysts. However, controlling for forecast accuracy, less experienced analysts are more likely to be fired for being bold (i.e., deviating from the consensus). Thus, less experienced analysts have incentives to trade off some accuracy and timeliness for the safety of being close to the consensus. One interpretation of these results is that analysts gain experience by watching the consensus, while at the same time testing their own models privately. Once they become confident in their own models, they become bolder and attempt to lead rather than follow the consensus. Future research might investigate the descriptive validity of this interpretation.

Recent theoretical and empirical research establishes a link between analyst effort and optimism. McNichols and O'Brien (1997) provide evidence suggesting that analysts tend to cover firms for which they genuinely have optimistic views.[33] Analysts also expend more effort

[33] McNichols and O'Brien rely on data from Research Holdings, Ltd., a proprietary database providing individual analyst stock recommendations, annual earnings forecasts, and actual *EPS* data, along with (particularly relevant to

covering these firms and, therefore, forecast accuracy improves with optimism. McNichols and O'Brien find that when analysts initiate coverage of a stock their forecasts and recommendations are relatively optimistic, but that optimism is justified by actual results, and Branson et al. (1998) find evidence of significantly positive abnormal returns at the time of coverage initiation announcements. McNichols and O'Brien also find that the most recent forecasts and recommendations of analysts dropping coverage tend to be relatively low but overly optimistic when viewed from the perspective of the subsequent information that causes the analysts to drop coverage. Including these outdated forecasts and recommendations in a consensus creates the appearance of overly optimistic analysts. In other words, the negative views of analysts dropping coverage (or failing to update) are censored, leaving a distribution with optimistically biased consensus forecasts. Overall, the study's evidence suggests that analysts' economic incentives create an asymmetry that leaves firms for which analysts have unfavorable views with less analyst following, less analyst scrutiny and less accurate earnings forecasts.

The results in McNichols and O'Brien (1997) confirm predictions in the theoretical model developed in Hayes (1998). Given the model's assumptions that a risk averse investor relies on an analyst for information and pays commissions to the analyst's brokerage house, the analyst has more incentive to gather and report information about securities that the analyst expects to perform well. The driving force behind the model is that, ceteris paribus, precision increases present values and, therefore, providing more precise information to support buy (sell) recommendations increases (decreases) investors' incentives to buy (sell) shares. Many of the model's predictions are borne out in the results reported by McNichols and O'Brien. However, some of Hayes's predictions await empirical testing. For example, she predicts that the effect of

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their study) detailed information indicating the timing of analysts' decisions to drop coverage of a security. The sample includes approximately 500 analysts chosen randomly from a group of about 1,800 analysts meeting the data requirements of the study, and the sample time-frame spans 1990-94.

asymmetric benefits associated with providing more precise information to support buy (versus sell) recommendations should increase with the extent to which investors already own shares of the stock, which in turn should increase with the size of the followed firm and the extent/influence of analysts' recent buy recommendations. Hayes also predicts that the asymmetry should increase with short selling restrictions on the stock and the dispersion of ownership among investors.

The evidence in McNichols and O'Brien and theory in Hayes (1998) could help explain the role of financial analysts in the Hong, et al. (2000) tests of Hong and Stein's (1999) "gradual information diffusion" theory of market inefficiency. Consistent with this theory, Hong et al. hypothesize and find that returns momentum increases with decreases in analyst following (which proxies for the speed of information diffusion across investors). The study also documents "an interesting regularity": the effect of low analyst coverage on returns momentum is most pronounced in stocks that are past losers. The authors speculate that this regularity is due to managers having less incentive to quickly publicize bad news and therefore low analyst coverage more severely restricts the flow of bad news. Hong et al.'s results are also consistent with: Hayes's (1998) theory and McNichols and O'Brien's empirical results suggesting that analysts expend less effort in their coverage of bad news stocks; and Hayes and Levine's (2000) evidence that the market does not appear to adjust its expectations for the selection bias documented by McNichols and O'Brien. Thus, the incentives described by Hayes (1998), when combined with the results in Hong et al. (2000), McNichols and O'Brien (1997) and Hayes and Levine (2000), contribute to the theory of returns momentum developed in Hong and Stein (1999). Further research is needed to understand the role of financial analysts in theories of

market inefficiency described in Hong and Stein (1999) and other behavioral finance studies, such as Barberis et al. (1998) and Daniel et al. (1998).[34]

3.5.2 Optimism bias?

In contrast to the theories described above, where forecast optimism arises endogenously from analyst coverage decisions, some prior research suggests that analysts may strategically add optimistic bias to their forecast reports. We first summarize research regarding whether, in fact, forecasts are generally optimistic and then discuss studies attempting to explain relative forecast and recommendation optimism including: rational responses to incentives including the need to curry favor with managers in order to maintain the supply of information and stimulating investment banking business and trading commissions.

Though researchers routinely assert that analyst forecasts are optimistic (e.g., research reviewed by Brown 1993), the evidence supporting overall optimism is contextually confined and sample-period specific.[35] The following overlapping factors contribute to conclusions regarding overall bias: (1) the sample period considered, (2) whether the mean or median forecast is used as evidence (and, related, treatment of outliers), (3) the forecast horizon, (4) the realization employed to determine optimism (actual data), and (5) the conduct of statistical tests establishing bias.

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[34] Barberis, et al. (1998) and Daniel, et al. (1998) provide theories of market inefficiency assuming psychological biases affect market prices. Friesen and Weller (2006) develop a model of behaviorally biased analyst forecasts due to overconfidence and cognitive dissonance of individual analysts. Further empirical, theoretical and experimental research is needed to determine whether analysts' forecasts reflect psychological biases, and whether these biases, in turn, affect market prices.

[35] Abarbanell and Lehavy (2003, 142) note that "[a]fter four decades of research on the rationality of analysts' forecasts it is somewhat disconcerting that the most definitive statements observers and critics of earnings forecasters appear willing to agree on are ones for which there is only tenuous empirical support."

Recent sample periods do not provide compelling support for forecast optimism. Brown (2001) notes that *median* forecast errors have decreased over time from slightly negative (optimistic) to zero to slightly positive (pessimistic) for all types of earnings (profits, losses, and zero net income) over the 1984-1999 time period. Similarly, Richardson, et al (2004) study an apparent walk-down in analyst forecasts to beatable targets and show that *median* price-scaled forecast errors are not optimistic within four months of announcement dates for sample periods from 1992-2001 using annual data (likewise, there is no apparent optimism using quarterly data from any bi-weekly horizon after 1991). Ramnath, et al (2005), using data from 1993-97 compare Value Line and IBES quarterly forecasts and note that while mean forecast errors from Value Line are negative (optimistic), errors resulting from consensus forecasts constructed from the IBES detail file are not different from zero. They report that median forecast errors are positive (pessimistic) using either Value Line or IBES.

The aforementioned studies evidencing a change over time tend to reference *median* forecast errors or distributions as opposed to the *mean* forecast errors. Abarbanell and Lehavy (2003) consider the distribution of forecast errors using Zacks data from 1985-1998 and report that two relatively small asymmetries drive conclusions regarding overall bias, which can alternatively be interpreted as supporting optimism, pessimism, or no bias (see section 3.7 for a detailed discussion). Evidence also supports a strong positive relation between optimism and forecast horizon. Longer-horizon within-quarter and, particularly, within-year forecasts (which exclude consideration of interim realizations) are optimistic (Kang, et al 1994). However, forecasts made closer to earnings announcement dates do not appear to be optimistic and, as

above, in recent periods appear to be slightly pessimistic.[36] Richardson, et al (2004) provide evidence of a walk-down pattern: forecasts tend to be optimistic at the beginning of the forecast horizon and approach realizations from above, ultimately ending slightly below reported EPS. This pattern is apparent across their sample sub-periods (1984-2001).

The forecast target proxy can also influence conclusions regarding overall optimism. For example, Keane and Runkle (1998) and Lim (2001) use Compustat actual EPS and truncate observations based on the magnitude of special (non-recurring) items.[37] Keane and Runkle are unable to reject unbiased forecasting with respect to bias *only* when 1) observations are truncated based on the magnitude of price-scaled non-recurring charges, and 2) standard errors are estimated using GMM which accounts for the presence of correlation across analysts' forecast errors.[38] Another challenge with use of benchmarks, as noted by Abarbanell and Lehavy (2003), is that analysts may not be expected or motivated to forecast the managed portion of the reported outcome.

In sum, conclusions regarding overall bias in analyst forecasts are subject to a variety of caveats. Forecasts appear to be trending toward pessimism and factors such as the staleness of the forecast, the actual data used, treatment of outliers, and how bias is assessed and tested must be considered before defensible assertions can be made. Another stream of literature seeks to ascribe explanations for relative optimism or pessimism. These explanations fall into two broad categories: 1) strategic optimism to improve management relations, and 2) potential biases caused by incentives.

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[36] E.g., Brown (2001) uses the most recent individual forecast and Richardson et al. (2004), note that *median* price-scaled forecast errors provide no evident optimism within four months of announcement dates using annual data, though longer-horizon forecasts appear optimistic.

[37] IBES, First Call, and Zacks all purportedly remove these items from their reported actuals and, while researchers can never be certain that individual analysts attempt to forecast the reported outcome from these databases, using database-defined outcomes appears to be a superior approach to estimating forecast errors.

[38] Based on arguments above, they were otherwise likely to reject rationality because their forecasts are relatively stale and the sample period largely predates the aforementioned distributional shift.

Francis and Philbrick (1993) develop a story and test the prediction that to cultivate management relations, analysts report optimistic forecasts. Using Value Line data, they provide evidence that less favorable stock recommendation are accompanied by relatively optimistic forecasts for the same securities. Francis and Philbrick conclude that focus on forecasting earnings in a multi-task environment is misguided to the extent that interaction among activities influences forecasting decisions. Motivated by Francis and Phibrick, Das, et al (1998) argue that analysts have increased incentive to issue optimistic forecasts to curry favor with management in an effort to obtain private information when the benefits of doing so are greatest (when earnings are unpredictable). Using Value Line data, they find evidence consistent with this argument in that forecast errors and measures of predictability are negatively correlated, controlling for other factors.

Eames et al. (2002) replicate the Francis and Philbrick (1993) result with Zacks data and find that the relationship between recommendation optimism and forecast optimism is reversed when including actual earnings in the equation or when excluding extreme observations. They conclude that their results are consistent with trade boosting.[39] Similarly, Eames and Glover (2003) find that the Das et al. (1998) result disappears with actual earnings included. Gu (2005) suggests that including actual earnings as a control variable can bias the coefficient on the test variable. Since forecast errors (the dependent variable) are defined as actual earnings minus forecasts, Gu suggests that with actual earnings included again as an independent variable, the coefficient on any other independent variable that is correlated with forecasts will be mechanically affected regardless of forecast efficiency. However, Eames et al. (2002) and Eames and Glover (2003) support their conclusions with alternative analyses that do not depend

[39] Eames et al. (2006) provide experimental evidence suggesting that cognitive biases lead to unintentially optimistic (pessimistic) earnings forecasts in the wake of buy (sell) stock recommendations.

on regression, and Eames et al. (2002) show that (without actual earnings on the RHS) the Francis and Philbrick (2003) regression analysis is not robust to exclusion of extreme observations. Overall, biasing earnings forecasts upward to curry favor with management seems tenuous since upward biased forecasts lead to negative forecast errors, and surely management prefers to avoid negative earnings surprise. This issue warrants further theoretical, archival and experimental research.

Lim (2001) draws on the management relations hypothesis to argue that historically documented optimism is not irrational because unbiasedness does not necessarily correspond to the "best" or most accurate forecast from the perspective of minimizing mean square error. Analysts may trade off positive bias for improved accuracy resulting from access to better information. Lim hypothesizes and finds that both firm and analyst characteristics correspond to optimism. Firm size and / or analyst following reduce optimism and target-specific uncertainty is positively related to optimism (proxies for demand for private information) whereas analyst employer size reduces relative optimism. Hong and Kubik (2003) claim that career concerns are a function of forecasting ability but controlling for accuracy, relative optimism in forecasts is valued for promotions. For investment bank analysts, promotions / demotions depend relatively more on optimism than accuracy.

Incentives related to attracting or maintaining underwriting relations and related to trading commissions are possible causes of relative optimism. Lin and McNichols (1998) explore the incentive impacts of underwriting business by studying earnings forecasts and recommendations on firms making seasoned equity offerings from 1985-1994. They report no evidence that short horizon earnings forecasts by affiliated analysts (those working for underwriters or co-underwriters) are optimistic relative to unaffiliated analysts; however, they

note that affiliated analysts issue relatively optimistic long-term earnings growth and stock recommendations. Dugar and Nathan, (1995) using a small sample from an earlier era, find the same result for recommendations, but report that analysts employed by target firm investment bankers issue relatively optimistic forecasts of various horizons. Dechow et al (2000) find that long-term earnings growth forecasts from underwriter analysts are significantly more optimistic than those of unaffiliated analysts. Jackson (2005) argues that optimistic forecasts generate more trading commissions for brokerage firms, but so does high reputation. As such, there is tension between making unbiased forecasts to build reputation and issuing optimistic forecasts to generate trading commissions. Jackson demonstrates that in equilibrium, forecast optimism can exist.

Relatedly, Michaely and Womack (1999) find that in the month following the quiet period after an IPO, lead underwriter analysts issue 50% more buy recommendations on IPO firms than do unaffiliated analysts. However, as described in the previous section, distinguishing optimistic bias from the issuer's underwriter selection based on favorable predisposition toward the issuer is difficult. Kadan, et al (2004) study whether the recent Global Analyst Research Settlement appears to have affected recommendation and earnings forecast behavior of affiliated vis-à-vis unaffiliated analysts. They conclude that affiliated-analyst recommendations remain relatively optimistic, however the level of relative optimism decreased following the settlement, and they suggest that residual relative optimism can be attributed to selection bias. Ljungqvist, et al (2006) investigate whether analyst recommendation behavior is influenced by potential underwriting business and report that recommendation behavior is influenced by economic incentives, but that behavior is ineffective in gaining business. They suggest that analysts are concerned about reputation capital and that aggressive recommendations are counterproductive.

A number of recent working papers revisit and extend research related to incentives affected by investment banking and brokerage relations. Because the previous results related to relative earnings forecast optimism are mixed, the results of these studies are important. While the studies differ in focus and execution, Cowen et al (2006), Jacob et al (2008), Agrawal and Chen (2007), and Clarke et al (2006) all report that investment bank analysts issue relatively *pessimistic* earnings forecasts. These results may differ from Dugar and Nathan (1998) because of larger, more recent samples, and the inclusion of more control variables and because Dugar and Nathan compare forecasts from analysts employed by the target's underwriter to all other analysts (as opposed to comparing analysts employed by investment bankers generally to all others).[40] Both Cowan, et al (2004) and Agrawal and Chen (2007) conclude that relative forecast optimism is associated with brokerage activities.[41] However, Irvine (2004) reports that while forecasts that differ from consensus generate significant brokerage trading, biasing earnings forecasts does not. Irvine concludes that greater trading commissions are generated through optimistic stock recommendations than via earnings forecasts.

Assessing absolute optimism in the context of recommendations is tenuous because there is no corresponding, mutually-agreed upon "actual," though empirically we observe a paucity of "sell" recommendations (Lin and McNichols, 1999; Michaely and Womack, 1999). Approaches to assess absolute recommendation optimism include comparing returns over intervals contemporaneous with and/or subsequent to the recommendation date[42] and evaluating market reactions for correspondence with the recommendation. For example, Lin and McNichols find

[40] However, Jacob, et al reach a conclusion similar to Lin and McNichols (1998) in that within the set of analysts employed by investment banks, affiliated and unaffiliated analysts issue comparably optimistic forecasts.
[41] In an experimental paper using a small sample of financial analysts from one brokerage firm, Hunton and McEwen (1997) report the opposite: an underwriter treatment group issues the most optimistic forecasts, then the brokerage treatment group, and finally, the control group issues the least optimistic forecasts.
[42] In an efficient market, all the information in a recommendation should be immediately impounded in securities prices, so the correspondence between longer-run returns and recommendations (if any) may reflect market inefficiency.

that though the market reactions to affiliated and unaffiliated analyst "strong buy" and "buy" recommendations are comparable, the reaction to "hold" recommendations is more negative for affiliated analysts leading to the conclusion that the market interprets a "hold" recommendation to mean "sell." Malloy (2005) studies the influence of geographic proximity on incentives of affiliated analysts. He replicates Lin and McNichols' results, but notes that the relative optimism of affiliated analysts is concentrated in distant, not local, affiliated analysts and also concludes that the negative market reaction to "hold" recommendations by affiliated analysts is prominent for distant affiliated analysts. Michaely and Womack (1999) report that securities assigned "buy" recommendations by underwriters under-perform "buy" recommendations from unaffiliated analysts before, at, and after recommendation dates and conclude that underwriter analysts are biased. Barber, et al (2007) report that while the abnormal return reaction to "buy" recommendations is significantly stronger for independent analysts than for investment bank analysts, the reverse is true for "hold" and "sell" recommendations (abnormal returns are more negative for investment bank analyst recommendations). They find that the underperformance related to "buy" recommendations is concentrated in a period after the NASDAQ market peak, suggesting that investment bank analysts were reluctant to downgrade stocks, consistent with the Global Analyst Research Settlement. However, the underperformance applies generally to investment bank analyst recommendations and is not confined to the ten investment banks sanctioned as part of the settlement.[43] More research is needed to determine the causes of relative optimism and pessimism in analyst forecasts and recommendations and how they relate to analyst and firm incentives.

[43]Another cause of absolute optimism given little attention in the literature relates to cognitive biases of analysts. For example, Sedor (2002) explores a non-rational explanation for forecast optimism in an experimental setting and reports evidence consistent with unintentional optimism about prior loss firms due to how information is provided to analysts by managers. She suggests that firms release information in a narrative manner causing scenario thinking among analysts, thereby inflating beliefs about the plausibility of future plans mapping to favorable outcomes.

3.5.3 Underreaction

While a plethora of studies have suggested and investigated various incentives that might create optimism in analysts' earnings forecasts, relatively few have addressed potential economic reasons for underreaction in analysts' forecasts. Empirical research has documented what appears to be underreaction in analysts' earnings forecasts at least since Mendenhall (1991) and Abarbanell and Bernard (1992) observed positive serial correlation in analysts' earnings forecast errors, and since Givoly and Lakonishok (1979) and Brown et al. (1985) observed a relation between forecast revisions and lagged changes in stock prices. Since underreaction was first documented empirically, some papers explain the phenomenon as a result of psychological biases in analysts' judgments, and a few others attribute the phenomenon to analysts' economic incentives. We review some of this research below.

Studies of forecasting efficiency with respect to information about future earnings generally rely on student subjects, and these studies generally view this as research design strength rather than a weakness. For example, Maines and Hand (1996) study whether M.B.A. students efficiently recognize the time-series properties of past quarterly earnings in forecasting the next quarter's earnings. They presented subjects with quarterly *EPS* data with seasonal first-order autoregressive and fourth-order moving average terms of various strengths. In cases where these terms were missing altogether, the time series conformed to a seasonal random walk. The authors summarize the results as follows: "... subjects overweighted the autoregressive component when it was small or non-existent ... appropriately weighted it when it took on a moderate value ... and underweighted it when it was large ... (subjects) correctly weighted (the moving average term) when it was not present ... and underweighted it when it was present."

Thus, subjects apparently paid too little attention to the temporary component of the change in earnings of the same quarter last year and understood the importance of the previous quarter's seasonal earnings change but were generally poor at distinguishing differences in it's importance between firms.

In a similar experiment, Calegari and Fargher (1997) found that business students generally underweighted the earnings change at the first lag (by about 40%), and they found no evidence that the students misweighted the earnings change at the fourth lag. Further research is needed to resolve this mixed evidence on the nature of the failings of business students' earnings forecasts to fully impound the time-series properties of quarterly earnings. One possibility is that student subjects are not appropriate surrogates for trained financial analysts. Some of it is evident from the results. For example, surprisingly, Calegari and Fargher (1997) find that increased experience did not improve the forecasting ability of student subjects (unlike archival findings as in Clement 1999 and Mikhail et al. 2003).[44] Future experiments could investigate whether/how financial analysts with professional experience overcome whatever cognitive limitations affect the students' judgments, and whether economic incentives or psychological biases are responsible for underreaction in professional analysts' earnings forecasts.

Turning to economic explanations for underreaction, Trueman (1990) derives a theoretical model from analysts' incentives. In this model, investors and analysts begin with unbiased expectations and analysts have three opportunities to update their forecasts before the earnings announcement. The first and third opportunities reflect analysts' private information, and in the middle analysts can respond to a public information release. Analysts have either strong or weak forecasting ability, and they are rewarded based on their forecasting reputations. Investors assume analysts only issue forecasts when their private information justifies Bayesian

[44] Maines and Hand (1996) also do not find improvements based on experience (see their footnote 16).

updates, so earlier updating signals greater ability to quickly develop/obtain useful private information. Investors increase their prior probabilities that analysts have strong ability based on both the timeliness and accuracy of analysts' forecasts. In this scenario, weak analysts sometimes rationally issue early forecasts even when private information does not justify a change in the analyst's true earnings expectations. In that case, since investors observe the weak analyst's first forecast and know analysts update their priors in a Bayesian manner, the weak analyst must underreact to the public information in order to avoid revealing that s/he really did not have substantive information on which to base the prior forecast. Thus, underreaction arises out of analysts' incentives to convince investors that they are skilled in timely development of private information useful for predicting future earnings. Whether the equilibrium conditions necessary for weak analysts to survive hold true in financial markets is an issue awaiting empirical examination.[45]

In contrast to Trueman (1990), Abarbanell (1991) suggests an economic explanation for underreaction that does not depend on analysts misreporting their true expectations. Abarbanell's explanation hinges on the idea that analysts' "private information is more easily inferred by investors if it is not combined with other signals whose information content is open to individual interpretation (p. 164)." However, analysts' incentives to provide useful (Mozes 2003) and accurate (Mikhail et al. 1999) forecasts suggest that they would want to include the effects of all available information in their forecasts. Raedy et al. (2006) provide an economic explanation for underreaction that does not require analysts to ignore public information. The explanation hinges on analysts facing an asymmetric loss function, whereby analysts incur

[45] Trueman (1994) develops a different theoretical model that predicts underreaction (and herding). However, that model assumes an earnings distribution with only four discrete outcomes, and the paper does not address the limitations of this assumption. It seems that the underreaction predicted by the model should approach zero as the number of possible earnings outcomes approaches a continuous distribution.

greater (less) reputation cost of forecast error when the error has the opposite (same) sign as the analyst's prior earnings forecast revision. Given this asymmetry in the penalty for forecast inaccuracy, analysts rationally underreact to information about future earnings (public and private), and the underreaction increases with the uncertainty of the earnings distribution and the asymmetric cost of forecast inaccuracy. Raedy et al. argue further that if market frictions (e.g., De Long et al. 1990, Slezak 2003) prevent the market from quickly unraveling analysts' underreaction tactics, then investors taking trading positions consistent with the direction of analysts' earnings forecast revisions should earn abnormal returns on their trades. These investors rationally prefer analyst underreaction and this preference could lead to the asymmetric cost of forecast inaccuracy driving the underreaction. Future research might investigate whether the nature of the analyst's clientele affects underreaction. For example, does the percentage of shares held by institutional investors with momentum trading strategies affect sell-side analyst underreaction?

Ideas in Stickel (1989) and Mozes (2003) suggest a potential extension of the Raedy et al. (2006) incentives-based explanation for underreaction. Stickel (1989) finds significantly increased forecast revision activity shortly after earnings announcements. However, as pointed out by Mozes (2003), competition to revise forecasts quickly after public information releases comes at the cost of lower forecast accuracy. Analysts might strategically manage these costs and benefits by releasing a forecast quickly in response to new information, but intentionally revise less than their information suggests. Then after taking the necessary time to thoroughly analyze the information, analysts release revised forecasts which, because of the previous underreaction, most likely create a change in earnings expectations in the same direction as implied by the previous revision. Given market frictions limiting arbitrage, investors taking

trading positions consistent with timely analyst forecast revisions earn abnormal returns and investors' incentives line up with analyst underreaction. Thus, both investors and analysts may be better off if analysts respond quickly to new information but with some restraint. Future research might investigate whether underreaction occurs in response to incentives encouraging analysts to trade off forecast accuracy for timeliness. More generally, continuing research is needed to better understand how analyst's economic incentives affect the timing, accuracy and bias of their earnings forecast revisions.

3.5.4 Management Communications and Incentives to Guide Analysts' Forecasts

In this section we describe research on management communications with analysts and discuss managers' incentives to guide forecasts and/or manage earnings to exceed forecasts. Managers may communicate to analysts through formal corporate presentations, through preannouncements of earnings, or informally by providing guidance. Clearly, there are strategic aspects of management communication with analysts. For example, managers may use earnings pre-announcements strategically to set expectations.[46] Soffer et al. (2000) suggest that there is an asymmetry in the tendency to preannounce earnings. Specifically, Soffer, et al examine management preannouncement disclosures shortly before earnings announcements and document that managers with bad news release almost all of their news at the preannouncement date, while those with good news release about half of the news. The benefit to such a strategy, consistent with research detailed below, is that firms with negative announcement date unexpected earnings experience significantly lower excess returns for the entire interval from just prior to the

[46] As another example, Fischer and Stocken (2000) consider the case where there is a misalignment of incentives between information senders and receivers and suggest that firms may choose to reduce the quality of disclosure to analysts in order to maximize the amount of information that analysts communicate to investors (effectively making the disclosure to analysts more coarse).

preannouncement to immediately after the earnings announcement date, controlling for the combined news released at both dates. This suggests that the manner of information presentation affects the market's reaction in that the market appears to focus more on earnings announcement news than to news released at other times. Tan et al. (2002) conduct a controlled experiment with analyst subjects and find that analysts forecast higher (lower) *future* earnings when earnings preannouncements are too conservative (optimistic). Preannouncements are too conservative (optimistic) when either positive news is understated (overstated) or negative news is overstated (understated).[47] Their finding suggests that not only current, but future earnings expectations can be influenced by preannouncements.

Apart from formal preannouncements, firms may choose to provide qualitative warnings or engage in corporate presentations as information venues. Libby and Tan (1999) conduct an experiment about how the timing of information release conditions the impact of qualitative earnings warnings. They provided subjects with information simultaneously and sequentially (as well as no warning) and found that analysts in the simultaneous warning condition made higher future earnings estimates than those in the no warning condition who, in turn, made higher estimates than those in the sequential warning condition. The implication of the study is that any positive effect related to analysts' preference for firms to provide warning is offset by the effect of sequentially processing a warning followed by an actual announcement. Francis et al. (1997) explore the benefits of communications made at corporate presentations to securities analysts for both senders and receivers. Firms may benefit via increases in analyst following or by mispricing correction, while analysts may benefit from information acquisition improving the frequency or quality of their forecasts. The results indicate increases in analyst following (but

[47] From debriefing the small sample of subjects, Tan, et al. report that analysts rate managers who accurately preannounce earnings highly and that analysts are aware that managers tends to introduce bias into preannouncements, they but do not appear to factor bias into their forecasts.

not more so for small versus large firms as suspected) and positive abnormal returns at the presentation date, with larger reactions observed for "underpriced" securities, generally consistent with firms benefiting from engaging in corporate presentations. However, analysts seem not to gain as Francis et al. provide no evidence that analysts' post-presentation forecasts are less disperse, more accurate, or less biased than forecasts made prior to presentations.

As noted in section 3.5.2, there is an increasing tendency for analysts to issue optimistic forecasts at the beginning of a forecast horizon and revise them downward such that slightly pessimistic forecasts manifest by the earnings announcement date. Two overlapping streams of literature related to managing analyst expectations downward and/or managing earnings to meet or barely exceed consensus forecasts have evolved. One set explores market consequences of missing versus meeting or beating expectations (MBE). The second considers related incentives, including management compensation consequences. Among studies documenting negative consequences to missing earnings forecasts are Bartov et al. (2002), Kasznik and McNichols (2002) and Brown (2003).[48] Bartov et al. examine the MBE phenomena by controlling for absolute earnings performance across a quarterly horizon and determine whether a residual market premium associated with MBE at period ends remains. They find evidence consistent with such a market premium, in spite of dampening expectations to exceed targets. Though the market appears to discount the effect when expectations or earnings management is deployed to meet or exceed expectations, a MBE reward still exists. Bartov, et al consider explanations for the phenomena and dismiss investor over-reaction, but note that future accounting performance is stronger for firms that MBE and that MBE may be a signal of future performance. Kasznik

[48] Brown and Caylor (2005) report on temporal patterns in propensities to MBE vis-à-vis other benchmarks and the market reward (penalty) to meeting (missing) various thresholds. They report that since the mid-1990s, analyst forecasts appear to be the appropriate threshold and that the market reward for exceeding analyst forecasts is much greater than that for exceeding a seasonal random walk or zero profit target.

and McNichols also report on rewards to MBE, finding that returns are higher for firms that MBE, increasing in the number of prior years that they do (from one to three years). MBE firms also experience higher future earnings forecasts and earnings realizations than other firms, controlling for other factors. Kasznik and McNichols interpret their results to suggest that the premium may reflect the market's perception that firms consistently meeting or beating earnings targets are less risky.

Skinner and Sloan (2002) explore explanations for why growth or "glamour" stocks appear to underperform the market and report that the inferior returns relative to value stocks result from expectational errors about future earnings. Specifically, growth stocks are punished more severely for the same magnitude of negative earnings surprise than are value stocks, providing incentives for growth companies to engage in expectation/earnings management. They report that if the asymmetric price response is controlled for, no return differential between growth and value stocks remains. Similarly, Brown (2003) documents the *decreased* propensity for firms to just miss expectations notes that it is concentrated in growth firms. Like Skinner and Sloan (2002), Brown finds that the negative valuation consequences of just missing forecasts are stronger for growth firms than others. Speculating upon incentives, he notes that over his sample period, stock and option compensation has increased substantially, particularly for growth firms, as has the influence of momentum investors and litigation against firms whose stocks drop appreciatively. Brown suggests that these temporal trends increasingly provide managers with incentives to avoid slightly negative earnings surprises.

Matsumoto (2002) more directly considers incentives for firms to avoid negative earnings surprises and the means to achieve such avoidance. She reports that firms with higher transient institutional ownership, greater reliance on implicit claims with their stakeholders, and higher

value-relevance of earnings are more likely to MBE. She relates the mechanisms to MBE (managing earnings upward or guide analysts' forecasts downward) to the same firm characteristics. Matsumoto finds that firms with more transient institutional ownership (consistent patterns of prior losses) are more (less) likely to use both means, while firms with more value-relevant earnings and those relying more on implicit contracts with stakeholders are more likely to manage expectations. Growth firms are more likely to use earnings management, not forecast guidance to MBE.

Another stream of literature relates to managers' personal incentives to guide forecasts (and/or manage earnings) to levels that can be met or beaten. Richardson et al. (2004) provide an incentive-based explanation that appears to comport to analyst forecast patterns within a quarter or year. They suggest that forecasts are guided downward to beatable targets so that there is a price increase in response to positive earnings news. This pattern is linked to institutional and regulatory changes (Insider Trading and Securities Fraud Enforcement act of 1988 and lifting of the short-selling rule for insiders in 1991) that created incentives for management guidance to boost stock prices in the post-earnings-announcement period. They report that the walk-down to beatable expectations is most pronounced for firms with stock issuances and firms with insiders selling their own shares in post-earnings-announcement period.[49] In another study about manager's incentives to MBE, Matsunaga and Park (2001) argue that the portion of executive annual bonus left to board discretion (rather than formulaic) could be adversely impacted by missing quarterly targets. They consider consensus forecasts, seasonal random walk expectations, and zero profit thresholds and find that CEO annual bonuses

[49] Firm benefits resulting from increased price levels prior to seasoned equity offerings is not necessarily self-interested but capitalizing via individual stock sales appears so.

are reduced if quarterly consensus forecasts and seasonal random walk expectations are missed for at least two quarters, after controlling for the general pay-for-performance relation.

In sum, a sizable body of research studies management communications with analysts and most relates to incentives to provide earnings guidance. Disproportionate negative market reactions to missing analyst forecasts at period ends are well established, but more work is needed to explore the rationality of such market reactions and to further explore the apparent walk-down pattern in earnings expectations.

3.6 Regulatory and cross-country comparisons

In this section we review literature related to the impact of regulatory changes (specifically the impact of Reg. FD, the Global Settlement Agreement with ten of the largest investment banking firms, and other exchange rulings) on analyst activities and the properties of their outputs and analyst activities from an international perspective. Both Reg. FD, the Global Settlement Agreement are relatively recent developments that could potentially dictate reinterpretation / reinvestigation of many prior findings. Also, increased access to international data offers new opportunities for cross-sectional comparisons of analyst behavior in different regulatory environments.

3.6.1 The Impact of Regulation Fair Disclosure (Reg FD) on Analyst Activities

In approving Reg FD in August 2000 (effective in October 2000), the SEC intended to proscribe selective disclosure of important information to particular (preferred) analysts. In effect, the regulation is intended to level the informational playing field. Prior to its passage there was broad speculation upon its likely impacts with respect to levels of information asymmetry across analysts, forecast accuracy, forecast dispersion, forecast and recommendation

informativeness, managers' propensity to communicate with analysts and the form of communication, and volatility in prices. A sizable body of literature seeks to fill this void.

Regarding forecast dispersion, directional hypotheses hinge on whether analysts forecasts rely more heavily on public versus private information in the post-Reg FD period. If public information is more heavily weighted, there should be less dispersion because the information provided by management is common to all. Alternatively, if analysts seek to gain advantage via their own analysis because public information is common, then dispersion could increase after Reg FD. The results related to forecast dispersion are mixed. Heflin, et al (2003) and Shane, et al (2001) report no evidence consistent with a change in forecast dispersion or accuracy following Reg FD. However, both Bailey, et al (2003) and Irani and Karamanou (2003) report that dispersion has increased following Reg FD. Irani and Karamanou suggest that the Heflin, et al (2003) result may be attributable to the brief post-Reg FD period (two months) in their sample. Irani and Karamanou also report a reduction in analyst following post-Reg FD, though Gintschel and Markov (2004) suggest that reduced analyst coverage could result from a downturn in the brokerage business.

With respect to pricing effects, Gintschel and Markov (2004) study the informativeness of outputs and find that the absolute price impact of forecasts and recommendations are lower by 28% after Reg FD. The decrease in pricing impact relates to the level of selective disclosure pre-Reg FD as proxied by brokerage and firm characteristics. Gintschel and Markov conclude that Reg FD is effective at eliminating selective disclosure. The results regarding return volatility are less conclusive as Shane, et al (2001) and Heflin, et al (2003) report reductions in return volatility after Reg FD, but Bailey, et al (2003) find no reduction in return volatility after controlling for decimalization of securities trading.

Bailey et al (2003) also report increases in trading volume attributable to difference in opinions following Reg FD. They also find that companies increased the level of disclosure, but the increase is confined to information about the immediate quarter. They conclude that the quantity of public information increased as a result of Reg FD. Eleswarapu et al (2004) conclude that aggregate information flow is unchanged by Reg FD, but information asymmetry as reflected in the adverse selection component of trading costs has decreased. Heflin, et al (2003) concur that there appears to be an increase in company disclosure, but the increase relates to forward-looking information. They also report that the information gap, proxied by deviations between pre- and post-earnings announcement stock prices, decreased after Reg FD. Further research is needed to reconcile the conflicting findings regarding the nature of increased public disclosure following Reg FD.

Bushee, et al (2004) use a unique database to study how Reg FD influences conference calls and the impact conference calls have on securities prices and trading activity. They compare timing, use and information content changes of the calls between firms that held closed conference calls prior to Reg FD and firms that previously held open calls. In terms of influence on the conduct of calls, Bushee, et al report that managers are more likely to discontinue conference calls in the post-Reg FD period and that proportionately more (previously) closed-call firms shifted away from always conducting calls after hours, though the impact of these changes is not substantial. They find no evidence that the amount of information disclosed during conference calls decreased after Reg FD, but report that price volatility increased for closed-call firms relative to open-call firms. Finally, they report that the amount of individual investor trading increased following the regulation change. This may result because individual

investors perceive that they are at less of an information disadvantage following the regulation change, and are therefore more willing to trade.

Jorion, et al (2005) take a different approach to assessing the impact of Reg FD by studying information providers exempted from the change. Specifically, they reason that because credit ratings agencies retained the right of restricted access to private information following Reg FD, the information content of ratings changes may have increased because new private information is revealed which is inaccessible to equity analysts. Jorion, et al report results consistent with their conjecture in that the positive (negative) market reactions to rating upgrades (downgrades) are greater after Reg FD.

In sum, a large body of research investigates the impact of Reg FD on analyst activities. Many of the results related to impacts are mixed and await further investigation. A challenge to many findings is that the regulation impacted all firms at the same time and, as such, there is no control group. Researchers need to exercise care in dismissing macro-economic (e.g., market downturn) and firm-specific effects that happened concurrently with the implementation of Reg FD.

3.6.2. The Global Settlement and other regulatory changes

In April of 2003, U.S. and state regulators, led by New York Attorney General Elliott Spitzer, announced a settlement (Global Analyst Research Settlement) with ten of the largest investment banking firms settling enforcement actions related to alleged conflicts of interest between research and investment banking operations. The settlement was intended to curb such conflicts by severing investment banking from research analysis, proscribing analysts from being evaluated or rewarded based on investment banking business and from being involved in 'sales

pitches' or 'road shows'. The settlement also requires firms' research management to make all company coverage decisions, provide 'independent' research to investor clients in addition to their own, and disclose historical ratings and price targets.[50] The Global Settlement along with regulations NASD 2711 and NYSE 472, which also relate to separation of investment banking and analysis, have likely impacted analyst activities substantially. Research related to these regulatory impacts, though largely unpublished at this point, is summarized in this section.

Barber et al (2006) report that after NASD Rule 2711 stock recommendations issued by analysts became more pessimistic on average and that recommendations departing from historical distributions presented substantial return opportunities (for example, going long on buy recommendations after Rule 2711 for brokers that previously issued few buy recommendations). Kadan et al (2008) also study the impact of the settlement and the two rule changes on recommendations for the ten firms agreeing to the settlement as well as other brokers. Overall, Kadan et al find that the regulations appear to impact analyst recommendations. Specifically, like Barber et al, they note that recommendations became more pessimistic post-regulation, more so for affiliated analysts (those whose employers had a past underwriting relationship with the firm), and particularly more pessimistic for the ten firms agreeing to the Global Settlement.[51] In a similar vein, Clarke et al (2006) compare recommendation distributions across investment firms that are investment banks, brokerages, or independent research firms (which became more prominent as a result of the Global settlement) and note that while recommendations became more pessimistic across all types of firms, the effect was the greatest for investment banking firms. Kadan et al note that this overall shift could be due to a modified rating system that was implemented at the same time as the regulations.

[50] See www.sec.gov/news/press/2003-54.htm for full details of the settlement.
[51] Boni (2005) reaches the opposite conclusion regarding the ten firms covered by the settlement.

Regarding the impact of these regulations on market reactions to recommendations, Kadan, et al find that in the post-regulation period, the market seems to regard recommendations as more credible. The decreased preponderance of 'buy' recommendations are met with higher returns than before, and the division between 'hold' and 'sell' recommendations is better reflected in returns post-settlement (i.e., only "sell' recommendations result in lower returns after regulations, but not 'hold' recommendations which were also interpreted as 'sell' pre-regulation). Regarding the type of analyst, Clarke et al conclude that the market interprets investment bank analyst recommendation changes to be the most informative both before and after the Global Settlement, implying that investors considered potential conflicts in assessing the valuation implications of recommendation changes.

Recent studies have also considered the impact of the regulations on analyst following. Kolasinski (2006) concludes that the new regulatory restrictions did not adversely impact analyst coverage, whereas Boni (2005) indicates that the ten firms included in the settlement reduced coverage post-regulation. The distinction between the results could be due to Kolasinski's focus on SEO and IPO coverage, whereas Boni bases her conclusion on the average total coverage of the ten firms agreeing to the settlement. Clarke et al also assess the impact of the settlement on analyst forecast bias and accuracy and conclude that both before and after the settlement investment bank analysts' earnings forecasts are less optimistic (consistent with Cowen et al, 2006 and Jacob et al 2008 before the settlement) but that analysts in each of the three categories are comparably accurate both before and after the settlement (inconsistent with Jacob et al 2008), implying that the Global Settlement did not substantially impact earnings forecasting. Future research might explore the impacts of the regulations on longer-horizon analyst forecasts and market responses to forecast revisions pre-and post-regulation.

3.6.3. International Forecasting

With the expanded access to international forecasts provided by IBES and other data providers recently, researchers have increased ability to study new research questions about the impacts of accounting standards and practices in various countries on analyst activities. Hope (2003a) and Hope (2003b) address issues related to the impact of disclosure practices and enforcement standards and accounting policy disclosures, respectively, on analyst forecasting behavior. Hope (2003a) focuses on forecast accuracy and conjectures that enforcement of accounting standards provides incentives for managers to follow rules, thereby reducing analyst uncertainty about managers' disclosures. He argues that reduced uncertainty results in better forecasting, reflected in lower forecast errors. Because there is presumably little, if any, variation in enforcement of standards across firms in the U.S., the international setting is appropriate. Hope (2003a) reports that strong enforcement is associated with increased accuracy and that disclosures gain importance with scant analyst following. He also notes that enforcement is more critical when managers have a menu of accounting methods from which to select. Hope (2003b) argues that information about accounting policies likewise reduces analyst uncertainty regarding the interpretation of accounting disclosures. After controlling for firm- and country-level variables, he reports that the level of disclosure about accounting policies is inversely related to forecast errors and forecast dispersion, consistent with reduced uncertainty.

Lang et al (2003) pursue a different tactic to study international issues by considering whether cross listing in the U.S. appears to provide informational and market benefits. They provide cross-sectional and time-series evidence that analyst following and forecast accuracy both improve as a result of cross listing. Firms with higher following and more predictable

earnings also benefit from higher market valuations and Lang, et al claim this is consistent with cross listing improving firm value via the effect on the information environment.

This review of forecasting of international companies listing locally, subject to country-specific reporting requirements, as well as those that cross list in the U.S. is far from exhaustive. The increased flow of international capital coupled with the convergence of international accounting standards makes this line of research important and we expect this area of research to expand substantially in the future.

3.7. Research design and database issues

Recent research questions some of the conclusions drawn by studies that examine the properties of analyst forecasts, especially with respect to analyst inefficiency and bias. Much of this criticism stems from the effect of temporal variations in analyst forecasting behavior, and firms' reporting and earnings guidance strategies, on conclusions related to analyst behavior. Other studies caution that the distributional properties of earnings forecast errors and assumptions about analysts' loss functions affect conclusions about the rationality of analysts' forecasts. While we discuss these issues in other sections of the paper, where appropriate, in this section we provide more detail on a few recent studies that focus on these issues.

Abarbanell and Lehavy (AL; 2003) argue that two asymmetries in the distribution of analyst forecast errors drive most of the prior findings of analyst optimism and analysts' under- and over-reaction to information. First, there are a larger number of greater magnitude observations in the extreme negative tail of the forecast error distribution than in the extreme positive tail, which AL refer to as the tail asymmetry. Second, there is a greater preponderance of slightly positive forecast errors than slightly negative forecast errors (consistent with

Degeorge, et al 1999). AL refer to this as the middle asymmetry. The tail asymmetry could be caused by accounting conservatism (Keane and Runkle 1998) resulting in more latitude to managers in recognizing non-recurring losses than gains (though this explanation is less plausible for studies using database-defined actuals which are presumably purged of such charges). Or, as Abarbanell and Lehavy suggest, the tail asymmetry could also be the result of managers engaging in substantial income-decreasing earnings management (big bath behavior) coupled with analysts not anticipating this behavior. The asymmetry near the middle may result from management guidance of forecasts to report slightly positive news.

AL argue that forecast errors computed in the usual way (i.e., actual earnings less forecasted earnings) presumes that the "actual" earnings number is the true value of the variable forecasted. This assumption may not hold if management strategically reports earnings and will affect inferences about analysts' forecasting bias and inefficiency. They find that unexpected accruals contribute more heavily to both the middle and tail asymmetries suggesting that the asymmetries are in fact influenced by earnings management. AL caution that econometric attempts at attenuating the effect of the tail asymmetry will lead to misleading inferences regarding analyst behavior. Researchers routinely winsorize or truncate observations at 1% and 99% of the forecast error distribution, ostensibly to eliminate potential data errors. Symmetric winsorization or truncation results in larger values of the negative tail being left out of the tests, but still leaves a number of high value negative forecast errors, as compared to positive errors, in the tests. This is because of the existence of more negative than positive extreme observations in the forecast error distribution. However, if the tail asymmetry reflects legitimate forecasting

errors, the effect of symmetric truncation or winsorization is to underweight (overweight) tail (middle) asymmetry observations resulting in an invalid pessimism inference.[52]

Kim et al. (2001) argue that mean analysts' forecasts do not efficiently aggregate forecasts across analysts. This inefficiency is increasing in the number of forecasts used to compute the mean because the common information across analysts gets more weight as the number of forecasts increase while the private information of analysts (impounded in the forecasts) get correspondingly lower weight. Kim et al. demonstrate that one way to overcome the misweighting of forecasts is to use the current mean forecast, but add a positive multiple of the change in the mean forecast from the prior period.

Gu and Wu (2003) analytically show that if analysts' objective is to minimize the mean absolute error in their predictions, their optimal forecast is the median, rather than the mean, of the earnings distribution. If analysts choose to forecast the median, then the resulting (observed) forecast bias will be related to the differences between the mean and median of the distributions of actual earnings for firms. They find that earnings distributions of firms are generally skewed and analyst forecast bias is positively related to the skewness in the earnings distributions. They also report that investors seem to (partially) adjust for this bias in earnings forecasts resulting from skewness in the earnings distribution.

The Gu and Wu (2004) work leads Basu and Markov (2004) to question the typical least squares-based research design in studying analyst forecast efficiency or rationality. The least squares regressions assume that analysts have a quadratic loss function. If analysts' objective is to minimize mean absolute forecast errors, then the proper research design would be regressions that minimize absolute errors rather than squared errors. Indeed, when using linear absolute deviation regressions, Basu and Markov find the surprising result that most prior documented

[52] Cohen and Lys (2003) provide a detailed critique of Abarbanell and Lehavy (2003).

forecast inefficiencies no longer hold. Their results highlight the importance of ensuring that assumptions implicit in the research design are recognized and validated. More research, however, is needed on understanding the type of loss functions analysts have and if there is systematic variation in this regard across analysts (for example, based on experience, brokerage house characteristics, etc.).

Research has also addressed specific issues related to analyst forecast databases. Payne and Thomas (2003) study the effect of split-adjusted forecast data provided by I/B/E/S. They point out that when I/B/E/S adjusts data for stock splits, they round off forecast and actual amounts, after the split, to two decimal places. While this may not affect the conclusions of most research, Payne and Thomas caution that some studies are especially prone to the split adjustment problem (for example, studies that examine zero forecast errors).

Ramnath et al. (2005) compare the properties of forecasts across the two premier sources of analyst forecast data, Value Line and I/B/E/S. With more recent data, they reexamine the conclusions in Philbrick and Ricks (1991) that Value Line and I/B/E/S provide forecasts of comparable quality, but the actual earnings numbers reported by Value Line provides a superior measure of earnings surprise. Ramnath et al. find, contrary to Philbrick and Ricks, that *actual* EPS data from Value Line and I/B/E/S are comparable, but I/B/E/S consensus forecasts are more accurate and are a better proxy for market's earnings expectations than Value Line forecasts. The superiority of I/B/E/S consensus forecasts can be traced to two attributes: 1) the ability to provide forecasts that are closer to the earnings announcement date (Value Line only publishes one forecast per quarter; I/B/E/S consensus is updated monthly), and 2) the advantage of being able to aggregate across forecasts from multiple analysts (Value Line forecasts are issued by a single analyst, whereas the I/B/E/S consensus is based on forecasts contributed by multiple

analysts). Finally, Brown and Sivakumar (2003) emphasize that the source of actual earnings used in measuring forecast errors is also important and could affect inferences. Specifically, they show that relative to operating earnings from Compustat, actual EPS from I/B/E/S better predicts future earnings, is more closely associated with stock prices, and also produces higher correlation between forecast errors and returns. .

Issues raised in the papers discussed above, about the reliability of inferences from prior research, seem substantial enough to motivate replication and validation of findings from past research. Further research is necessary especially in areas where the research design and database choices could have affected the conclusions of the study (for example, analysts' forecasting bias). Moreover, the findings in Abarbanell and Lehavy (2003) also point to the need for researchers to provide detailed discussion of the sensitivity of results reported in their studies to the treatment of outliers.

4. Summary and conclusion

Discovering the information and valuation models that determine prices of equity securities in capital markets is a formidable task, worthy of substantial research effort for many years to come. Analysts may collectively hold the key, but no single analyst can tell you what it is. Instead, the key lies in the way the market derives a consensus from the distribution of extant individual analysts' forecasts of a company's future earnings, the characteristics of the information impounded in that consensus, and the additional information the market incorporates into its model for valuing a company's equity securities. Important insights can be gained from research regarding analysts' decision processes, the determinants of analyst expertise and distributions of individual analysts' forecasts, the informativeness of analysts' research outputs,

market and analyst efficiency with respect to value-relevant information, effects of analysts' economic incentives on their research outputs, effects of the institutional and regulatory environment, and the limitations of databases and various research paradigms. In this paper, we hope we have provided some perspective on the research in each of these important areas.

The areas of future research that seem most promising to us include the following. First Schipper's (1991) and Brown's (1993) calls for research providing more insight into analysts' decision processes are still relevant today. We look forward to research clarifying the distinction between analysts' roles as interpreters of public information and developers of private information useful in determining prices of equity securities. The decision processes of analysts in distinguishing permanent from more temporary components of earnings reports (including temporary components due to earnings management) also remains a critical area for future research. We also expect to see research that clarifies the role of heuristics in the price-setting process and the degree to which these heuristics function as effective substitutes for rigorous multiperiod valuation models. More research is needed to understand the interaction between analysts' economic incentives and frictions that limit investors' abilities to arbitrage away any inefficiencies or biases in forecasts and prices resulting from those incentives, and we expect this research to have implications for emerging behavioral finance theories of market inefficiency.

We also predict that researchers will continue exploring factors that make some analysts better forecasters than others and the mechanism that develops consensus market expectations from individual analyst forecasts. Further research is also required on forecasts that have higher price impact, such as long-term growth forecasts and target prices. Given evidence of the informativeness of earnings in the presence of analysts' price target forecasts, recommendations, and other information in analysts' research reports, it is not clear that earnings forecasts are

simply a means to an end (Schipper, 1991); further research is needed to explore the importance of analysts' earnings forecasts and actual earnings reports in the allocation of resources in capital markets. Finally, we expect to see more international research describing institutional and regulatory factors that create cross-country differences in the role of analysts and the properties of their forecasts.

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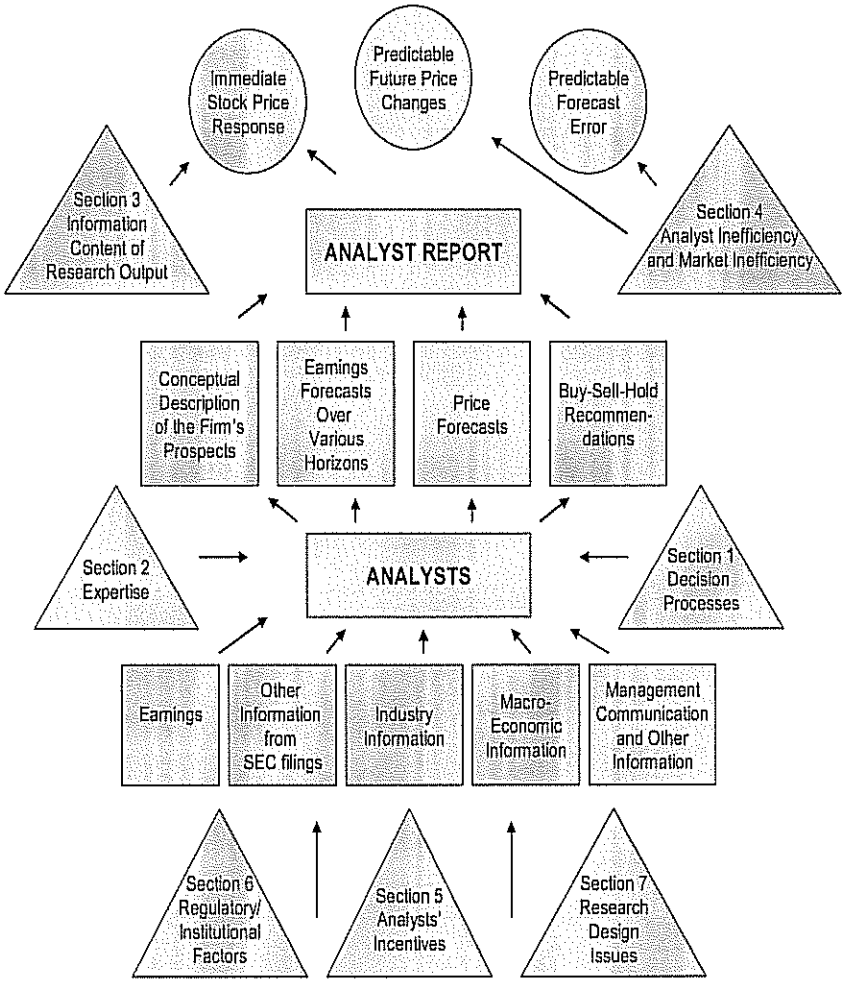
Figure 1
Analysts' Reporting Environment



FAJ Financial Analysts Journal
Volume 66 • Number 4
©2010 CFA Institute

Conflicts of Interest and Analyst Behavior: Evidence from Recent Changes in Regulation

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Regulation FD made analysts less dependent on insider information and diminished analysts' motives to inflate their forecasts. The Global Research Analyst Settlement had an even bigger impact on analyst behavior: The mean forecast bias declined significantly, whereas the median forecast bias essentially disappeared. These results are similar for all analysts.

Our investigation of the impact of recent changes in regulation on analysts' forecasting behavior follows a number of studies that argued that analysts were motivated to produce research reports that did not reflect their true opinions. Analysts tended to make excessive "buy" recommendations and inflated earnings forecasts for several reasons, two of which gained considerable attention from regulators in the United States. First, analysts may have felt compelled to favor managers in covered companies in order to gain privileged access to information flow (Lim 2001). Second, although analysts are supposed to provide investors with accurate and truthful research reports, conflicts of interest could occur because analysts' compensation was tied to profits generated from investment banking business and brokerage commissions (Lin and McNichols 1998; Carleton, Chen, and Steiner 1998).

In the early part of the first decade of this century, in an effort to restore public confidence in U.S. capital markets, U.S. regulators enacted several rules and regulations, prosecuted analysts whose research reports were tainted by conflicts of interest, and fined banks that failed to prevent research analysts' conflicts of interest. Two of the main regulatory developments during this period were (1) Regulation Fair Disclosure (Reg FD), which became effective on 23 October 2000, and (2) the Global Research Analyst Settlement (Global Settlement), which was announced on 20 December 2002.[1]

Although the primary goals of these two regulatory actions are different, they both have the potential to improve the quality of analyst forecasts. One of the stated goals of Reg FD is to prohibit private communication between companies and analysts, thereby helping to level the playing field so that market participants can have equal access to information and making analysts less dependent on such communication. In prohibiting companies from selectively disclosing private information to analysts, Reg FD may reduce analyst forecast bias by eliminating the incentive for analysts to inflate their earnings forecasts in order to gain access to insider information.

The Global Settlement is an important enforcement agreement between U.S. regulators and 12 large investment banks (the Big-12 banks) designed to eliminate research analysts' conflicts of interest. If successful, the Global Settlement should reduce optimistic bias in analyst forecasts.

Our study considered whether these two actions by U.S. regulators reduced the bias in analysts' earnings forecasts documented in previous studies. We focused on annual earnings forecast bias for several reasons. First, investors may use analyst forecasts to form expectations of earnings and cash flows, both of which are important inputs for stock valuation models. Inflated earnings forecasts can drive stock prices above their fair values if investors fail to adjust for the bias.[2]

Second, given the flurry of new regulations, regulators clearly consider analyst behavior an important factor in maintaining investor confidence in financial markets. Regulation is costly because of the significant expenses associated with analyzing problematic situations and developing remedies. Moreover, restrictions and reporting requirements imposed on various market participants result in ongoing compliance costs. These costs can be justified only if the new regulations help reduce analysts' conflicts of interest and thereby generate an important benefit for financial markets.

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Third, most studies that have examined the impact of Reg FD and the Global Settlement on analyst behavior focused on forecast accuracy and forecast dispersion (Bailey, Li, Mao, and Zhong 2003; Agrawal, Chadha, and Chen 2006).[3] These aspects of analyst behavior, however, are little affected by conflicts of interest, the focus of our study.

Other studies have examined forecast bias. Clarke, Khorana, Patel, and Rau (2006) found that the Global Settlement had no impact on relative bias in analyst forecasts. Focusing on the impact of Reg FD on bias in quarterly earnings forecasts between October 1999 and December 2001, Mohanram and Sunder (2006) found that these forecasts became more optimistic after Reg FD but attributed the increase to unexpectedly low realized earnings during the 2001 recession. Our longer study period (1996–2006) allowed us to control for macroeconomic conditions in our regression analysis. Furthermore, we examined longer-term (up to 24 months) earnings forecasts in which the forecast bias is more apparent (Richardson, Teoh, and Wysocki 2004). Although Herrmann, Hope, and Thomas (2008) found some evidence of decline in forecast bias following Reg FD, they focused on internationally diversified companies only; we examined all U.S. companies, and our primary focus was on changes in forecast bias after the Global Settlement.

Lastly, the ability of analysts to forecast earnings accurately can be easily and straightforwardly verified because actual earnings are observed at the end of the forecast period. Barber, Lehavy, McNichols, and Trueman (2006) studied the change in distribution of stock recommendations made from 1996 to 2003. They found that the percentage of buys decreased starting in mid-2000.[4] How unbiased the new distribution of stock recommendations is, however, remains uncertain. But we know that the bias should be zero at the aggregate level when analysts make their forecasts on the basis of their true opinions.

Institutional Background

Historically—and especially before recent regulations—analysts have tended to make unduly optimistic earnings forecasts. In this section, we discuss the possible reasons for this optimistic bias and the potential impacts of the recent regulations on such bias.

Why Do Analysts Make Overoptimistic Earnings Forecasts? A number of studies have documented that analysts regularly make overoptimistic earnings forecasts (Brown 1997; Chopra 1998; Beckers, Steliaros, and Thomson 2004). Optimistic bias tends to be larger for longer-term forecasts and smaller for forecasts made closer to the earnings announcement date. This phenomenon is usually referred to as the walk-down trend (Richardson, Teoh, and Wysocki 2004). Several explanations have been offered for analyst optimism.

First, analysts may be influenced by conflicts of interest if their compensation is tied to investment banking fees and brokerage commissions. Lin and McNichols (1998) found that analysts affiliated with underwriters make more favorable stock recommendations and long-term earnings growth forecasts than analysts not so affiliated. Agrawal and Chen (2005) discovered that optimism in long-term earnings growth forecasts is high when analysts work for financial institutions whose revenues come mainly from brokerage business. Carleton, Chen, and Steiner (1998) found that stock recommendations made by brokerage firms are more optimistic than those of nonbrokerage firms. Using Australian data, Jackson (2005) noted that optimistic analysts generate more trades for their brokerage firms than do less optimistic analysts. Chan, Karceski, and Lakonishok (2007) showed that analysts' earnings forecasts are influenced by their desire to win investment banking clients. Doukas, Kim, and Pantzalis (2005) reported that stocks with excess analyst coverage yield lower future returns, consistent with the conflict-of-interest hypothesis. Hong and Kubik (2003) found that brokerage houses reward optimistic analysts; optimistic analysts at low-status brokerage houses are more likely to move up to higher-status brokerage houses than are less optimistic analysts.

Second, analysts may feel compelled to maintain good relations with company management in order to gain access to insider information that can help improve the accuracy of their forecasts (Lim 2001). Third, analysts may tend to cover stocks for which they have positive views and drop or avoid stocks for which they have negative views, which can induce a self-selection bias (McNichols and O'Brien 1997). Fourth, analysts may have a cognitive bias that leads them to overreact to good earnings information and underreact to bad earnings information (Easterwood and Nutt 1999; Nutt, Easterwood, and Easterwood 1999). Finally, the walk-down trend may be driven by the "earnings guidance game," in which analysts issue optimistic forecasts at the start of the fiscal year and then revise their estimates until the company can beat the forecast at the earnings announcement date (Richardson, Teoh, and Wysocki 2004).

Recent Regulations. Before Reg FD, analysts and institutional investors often had an informational advantage over small investors through private communications with management and conference calls in which company managers discussed past performance and provided guidance on future prospects. Such timely information gave these investment professionals an unfair advantage that allowed them to trade stocks profitably at the expense of uninformed investors.

To gain access to this information flow, analysts may have had to maintain good relations with insiders by making optimistic forecasts and buy recommendations in their research reports. Analysts' excessively optimistic views of the stocks were misleading and contributed to the deterioration of investor confidence in capital market integrity. Through Reg FD, which was introduced in October 2000, the U.S. SEC intended to improve fairness and restore public confidence in the markets by requiring U.S. public companies to disclose material information simultaneously to all market participants.

Other sources of conflicts of interest, however, remained unaddressed by Reg FD. For instance, analysts could be pressured to make optimistic forecasts and buy recommendations in order to favor investment banking clients and generate trading volume. The SEC and such self-regulatory organizations (SROs) as the National Association of Securities Dealers (NASD; now the Financial Industry Regulatory Authority [FINRA]) and the NYSE paid significant attention to this issue and introduced a number of new rules and regulations to curb the negative consequences of these conflicts of interest.

The Sarbanes–Oxley Act of 2002 (SOA), also known as the Public Company Accounting Reform and Investor Protection Act of 2002, became law on 30 July 2002. The SOA is a broad piece of legislation that covers various business practices, including auditor independence, corporate responsibility, enhanced financial disclosure, analysts' conflicts of interest, and corporate and criminal fraud accountability. The SOA amended the Securities Exchange Act of 1934 by creating Section 15D, which requires FINRA and the NYSE to adopt rules reasonably designed to address research analysts' conflicts of interest.

To comply with the SOA, the NASD released Rule 2711 (Research Analysts and Research Reports) and the NYSE amended its Rule 351 (Reporting Requirements) and Rule 472 (Communications with the Public). Most provisions of these rules went into effect on 9 July 2002. These rules mitigate analysts' conflicts of interest by separating research analysts from the influence of the investment banking and brokerage businesses. Research analysts' compensation can no longer be tied to the performance of these businesses. In addition, analysts are restricted from personal trading in the stocks they cover.

On 6 February 2003, the SEC adopted Regulation Analyst Certification (Reg AC).[5] Reg AC provides guidelines for proper disclosure of potential conflicts of interest of sell-side analysts, including their association with investment banking clients and the structure of their compensation.

Regulatory objectives have also received support from rigorous enforcement actions. Following a joint investigation by the SEC, NASD, NYSE, and New York State Attorney General, 10 large U.S. and multinational investment banks agreed to pay a fine of $1.435 billion in the Global Research Analyst Settlement for their failure to adequately address research analysts' conflicts of interest. Announced on 20 December 2002, the terms of the Global Settlement initially covered 10 banks.[6] The final agreement was announced on 28 April 2003. Two more banks reached settlements on 26 August 2004.[7] The Global Settlement and the SRO rules share the same spirit in that their mutual objective is to eliminate analysts' conflicts of interest.

The introduction of these rules and regulations allows us to differentiate among the alternative explanations for analyst forecast bias proposed in the literature. First, a reduction in forecast bias after Reg FD would support the argument that analysts were overoptimistic owing to their need for insider information, especially if such a reduction were stronger for informationally more opaque companies. Second, a reduction in bias after the Global Settlement and Rule 2711 would be consistent with the hypothesis that analyst behavior was unduly influenced by conflicts of interest.[8] In contrast, self-selection and cognitive biases may exist even in a world without conflicts of interest. Therefore, if these biases are the main reasons for analysts' overoptimistic forecasts, then these regulatory changes should have no effect on forecast bias.[9]

Sample and Variables

We downloaded sell-side analysts' earnings forecasts for fiscal year-end dates between 1996 and 2006 from the Detail file of the I/B/E/S database. We used forecasts for current- and subsequent-year earnings per share (EPS), which are made for the upcoming and following years' earnings announcement dates.[10] **Figure 1** illustrates the timeline of analyst forecasts. The earliest analyst forecasts for a specific fiscal year-end EPS are made 24 months before the forecast fiscal year-end (in forecast month –23). For each EPS, analysts can

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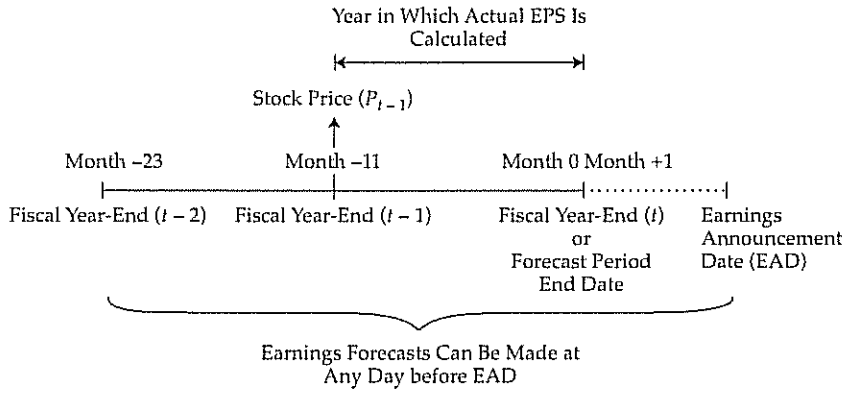
Figure 1. Timeline of Analyst Forecasts



make multiple forecasts over the course of the next 24 months. Some analysts may continue to make forecasts after the forecast fiscal year ends because companies announce their annual earnings after a delay of several months. Because the length of the EPS announcement delay could be affected by how high or low the realized EPS is relative to the consensus, we retained only those forecasts made no more than one month after the forecast fiscal year-end (in forecast month +1), which left us with a total of 2,297,792 forecasts.

For each forecast, I/B/E/S provides actual earnings, forecast date, forecast period (fiscal year) end, earnings announcement date, analyst code identity, broker code identity, and number of analysts used for consensus calculation.[11] We used the I/B/E/S Broker Translation file to convert broker codes into brokers' names, which we used to identify analysts who worked for the Big-12 banks. Stock prices are from the I/B/E/S Summary file.[12] We downloaded real GDP growth rates from the website of the U.S. Bureau of Economic Analysis. We downloaded SIC codes from the CRSP monthly file.

We defined analyst forecast bias, the focus of our analysis, as the average analyst forecast error and calculated it as follows:

$$Bias_{j,t,m} = 100\left(F_{j,t,m} - A_{j,t}\right)/P_{j,t-1}, \quad (1)$$

$$F_{j,t,m} = \frac{1}{I_{j,t,m}} \sum_{i=1}^{I} F_{j,t,m,i}, \quad (2)$$

and

$$F_{j,t,m,i} = \frac{1}{K_{j,t,m,i}} \sum_{k=1}^{K} F_{j,t,m,i,k}, \quad (3)$$

where

$A_{j,t}$ = the actual earnings per share for company j in fiscal year t

$F_{j,t,m,i}$ = the average of annual earnings forecasts for fiscal year-end t of company j, made in month m by analyst i

$K_{j,t,m,i}$ = the number of forecasts made in month m by the same analyst i for the same company j and fiscal year t

$I_{j,t,m}$ = the number of analysts making forecasts in month m for company j and fiscal year t

$P_{j,t-1}$ = the stock price of company j one year before the fiscal year-end t[13]

Note that all EPS forecasts made for the same company and the same fiscal year are normalized by the same stock price. Using the same stock price as the denominator guarantees that any changes in forecast bias across forecast months (m) are the result of changes in analyst forecasts, not of changes in the stock price. In our calculations according to Equations 1–3, we used only new forecasts made in month m. Stale forecasts from earlier months ($m - 1$, etc.) were not carried over into month m. In other words, each forecast participated in the calculation of the forecast bias only once, in the month in which the forecast was made. In our sample, an average analyst made 4.5 forecasts for each annual EPS. Because for each annual EPS we tracked 25-month forecasts (from month –23 to month +1), the implication is that an average analyst in our sample made a forecast for each covered company about once every six months.

To minimize the influence of outliers and misreported data in our analysis, we replaced with missing values any extreme observations of forecast bias, company size, market-to-book ratio, the number of stocks, and the number of industry analysts following.[14] We dropped from the sample all forecasts made in October 2000 and December 2002 (1.5 percent of our sample) and observations with missing values of any relevant variable. We were

left with 1,586,000 individual analyst forecasts, which we used to calculate 434,268 average forecast errors. For each fiscal year and for each of our 7,315 sample companies, our sample contained up to 25 monthly observations of forecast bias ($Bias_{j,t,m}$).

Table 1 reports the summary statistics for the overall sample of 434,268 observations and for each of the three subperiods. The period before Reg FD represents 53 percent of our sample observations, with the period between Reg FD and the Global Settlement and the period after the Global Settlement representing 18 percent and 29 percent of the sample observations, respectively. The mean forecast bias across all sample observations is 1.39 percent of stock price. This result is consistent with prior evidence that analysts' forecasts are optimistically biased (Brown 1997; Chopra 1998). No significant difference exists between the mean forecast bias before Reg FD (1.72) and the mean forecast bias between Reg FD and the Global Settlement (1.97). The mean forecast bias is more than four times smaller after the Global Settlement (0.41), with the difference statistically significant at the 1 percent level.

The average market capitalization of companies in our sample was $4.5 billion, and the average market-to-book ratio was 3.57. On average, 8.41 analysts covered a company in any particular month. The analysts in our sample worked for brokers that, on average, each employed 65.7 analysts. A typical analyst followed 16.30 stocks from 4.78 industries and, at the time of the forecast, had been in the I/B/E/S database for 6.24 years and making forecasts for the covered stock for 2.5 years. Around 17 percent of forecasts were made for companies with negative earnings, and 36 percent of forecasts were made for companies whose earnings were declining relative to earnings in the prior fiscal year.

Test Results

In this section, we present the results of the univariate tests and of the regression analysis of the effects of Reg FD and the Global Settlement on bias in analyst forecasts.

Univariate Results by Forecast Month. **Table 2** presents the median forecasts by the month in which the forecasts were made and by the fiscal year for which they were made. The numbers in the leftmost column represent the month (relative to the fiscal year-end) of the forecast. The numbers in the top row represent the fiscal years for which the

Table 1. Summary Statistics

				Number of Observations			Mean		
Description	Variable	Number of Observations	Mean	Before Reg FD	Between Reg FD and GS	After GS	Before Reg FD	Between Reg FD and GS	After GS
Forecast bias	Bias	434,268	1.39	231,096	77,305	125,867	1.72	1.97	0.41
Reg FD indicator	RegFD	434,268	0.18	231,096	77,305	125,867	0.00	1.00	0.00
Global Settlement indicator	Glob	434,268	0.29	231,096	77,305	125,867	0.00	0.00	1.00
Company characteristics									
Analyst coverage	NumA	434,268	8.41	231,096	77,305	125,867	8.21	8.23	8.88
Market cap ($ millions)	CompanySize	434,268	4,470.00	231,096	77,305	125,867	3,480.00	5,250.00	5,800.00
Market-to-book ratio	MB	434,268	3.57	231,096	77,305	125,867	3.78	3.47	3.23
Negative EPS	EPSLoss	434,268	0.17	231,096	77,305	125,867	0.16	0.26	0.14
Declining EPS	EPSDecline	434,268	0.36	231,096	77,305	125,867	0.37	0.45	0.27
Litigation	Litigation	434,268	0.27	231,096	77,305	125,867	0.25	0.30	0.27
Labor intensive	Labor	434,268	0.61	231,096	77,305	125,867	0.60	0.63	0.63
Analyst characteristics									
Company-specific experience	YearStk	434,268	2.50	231,096	77,305	125,867	2.55	2.43	2.46
General experience	YearIBES	434,268	6.24	231,096	77,305	125,867	6.45	6.19	5.87
No. of stocks covered	NumStk	434,268	16.30	231,096	77,305	125,867	18.18	14.31	14.06
No. of industries covered	NumInd	434,268	4.78	231,096	77,305	125,867	5.46	4.15	3.93
Broker size	BrokerSize	434,268	65.70	231,096	77,305	125,867	54.98	89.03	71.06

Note: This table presents the summary statistics for the overall sample and for the three subperiods.

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Table 2. Forecast Bias by Fiscal Year and Forecast Month

	Forecast Period End Year										
Month	96	97	98	99	00	01	02	03	04	05	06
–23	0.1	0.4	1.4	1.6	–0.3	1.9	2.3	1.2	–0.2	–0.3	–0.3
–22	0.3	0.5	0.9	1.3	0.5	2.2	2.7	1.3	0.0	–0.1	0.0
–21	0.3	0.5	1.1	1.6	0.5	2.1	2.6	1.3	0.0	0.0	0.2
–20	0.4	0.5	1.1	1.3	0.6	2.2	2.2	1.4	–0.1	0.0	0.0
–19	0.5	0.7	1.1	1.6	0.5	2.1	2.1	1.3	–0.1	0.0	0.1
–18	0.5	0.4	1.2	1.4	0.6	2.1	1.8	1.1	–0.2	0.0	0.1
–17	0.4	0.4	1.2	1.1	0.5	2.1	1.4	1.0	–0.2	0.0	0.1
–16	0.4	0.5	1.3	1.3	0.6	2.0	1.5	1.1	–0.1	0.0	0.2
–15	0.4	0.4	1.1	0.8	0.4	1.7	0.9	0.8	–0.3	0.0	0.2
–14	0.4	0.3	0.9	0.6	0.4	FD	0.6	0.4	–0.2	0.0	0.1
–13	0.4	0.3	1.0	0.6	0.4	1.5	0.5	0.3	–0.2	0.1	0.2
–12	0.3	0.2	0.8	0.4	0.3	1.6	0.4	GS	–0.2	–0.1	0.1
–11	0.3	0.3	0.8	0.3	0.1	1.3	0.3	0.1	–0.1	0.0	0.1
–10	0.2	0.2	0.5	0.1	0.2	1.1	0.2	0.0	–0.1	–0.1	–0.1
–9	0.2	0.2	0.6	0.1	0.1	1.1	0.2	0.0	–0.1	0.0	–0.1
–8	0.1	0.1	0.5	0.1	0.1	0.7	0.2	–0.1	–0.1	–0.1	–0.1
–7	0.1	0.0	0.5	0.1	0.1	0.6	0.2	–0.1	0.0	0.0	0.0
–6	0.1	0.1	0.4	0.0	0.1	0.5	0.2	–0.1	–0.1	–0.1	0.0
–5	0.0	0.0	0.2	0.0	0.0	0.2	0.1	–0.1	0.0	–0.1	0.0
–4	0.0	0.0	0.1	0.0	0.0	0.2	0.1	–0.1	0.0	0.0	0.0
–3	0.0	0.0	0.1	0.0	0.0	0.1	0.0	–0.1	–0.1	0.0	0.0
–2	0.0	0.0	0.0	0.0	FD	0.0	0.0	–0.1	–0.1	0.0	0.0
–1	0.0	0.0	0.0	0.0	0.0	0.0	0.0	–0.1	0.0	0.0	–0.1
0	0.0	0.0	0.0	–0.1	0.0	0.0	GS	–0.1	0.0	0.0	–0.1
1	0.0	–0.1	0.0	–0.1	–0.1	0.0	0.0	–0.1	–0.1	–0.1	–0.3
Median bias	0.2	0.2	0.8	0.3	0.1	1.2	0.4	0.0	–0.1	0.0	0.0
Mean bias	1.2	1.1	1.8	2.2	1.4	3.0	2.1	1.6	0.1	0.5	0.3
Mean forecast	6.2	5.3	4.6	5.1	5.3	3.7	3.0	4.0	4.4	4.2	5.0
Mean actual earnings	5.0	4.1	2.8	2.9	3.9	0.7	0.9	2.4	4.2	3.7	4.7
Mean stock return (%)	0.2	0.2	0.0	0.2	0.0	0.0	–0.2	0.6	0.2	0.1	0.2
GDP (%)	3.7	4.5	4.2	4.5	3.7	0.8	1.6	2.5	3.9	3.2	3.3

Notes: Forecast bias is the difference between the mean of all forecasts made in a particular month for a particular company and a particular fiscal year and the realized EPS, scaled by the stock price and multiplied by 100. *Forecast period end year* is the fiscal year for which the forecast was made. *Month* is the month of the forecast relative to the fiscal year-end. *FD* is the month in which Reg FD became effective (October 2000). *GS* is the month in which the Global Settlement was announced (December 2002). Stock returns were calculated from our samples.

forecasts were made. For example, forecasts made in September 2000 for the fiscal year ended December 2000 (i.e., three months before the fiscal year-end) are in row –3 and column 00. The two solid lines separate the forecasts made before and after Reg FD and the forecasts made before and after the Global Settlement. The six bottom rows present forecast bias for each fiscal year averaged across all forecast months, along with the realized earnings per share, average forecasts, annual stock returns, and real GDP growth rates.[15] To align fiscal year-end dates with annual variables, such as real GDP growth rates, we used only forecasts for companies with December fiscal year-ends.

For each year before the Global Settlement, the median forecast errors are significantly positive. Furthermore, for each year before the Global Settlement, we observe the walk-down trend with forecast bias steadily declining as forecasts are made closer to the fiscal year-end. After the Global Settlement, we observe a significant drop in the forecast bias. The results show a total absence of bias in the median forecast errors for 2004–2006 (–0.1 percent, 0.0 percent, and 0.0 percent, respectively). The walk-down trend in median forecast errors is also practically nonexistent for fiscal years 2004–2006.

These results suggest that analysts' conflicts of interest indeed led to excess optimism in earnings forecasts before the Global Settlement and that the Global Settlement has been effective in neutralizing analysts' conflicts of interest. Alternative interpretations of the forecast bias, such as self-selection, cognitive bias, and need for insider information, cannot explain these findings because the Global Settlement should have no effect on these factors.

Unusually high stock valuations and/or realized earnings, rather than less optimistic forecasts, could be responsible for the decline in the average forecast errors after the Global Settlement. A quick look at the actual and forecasted EPS, stock returns, and real GDP growth rates before and after the Global Settlement, however, does not seem to support this idea. Neither aggregate economic performance nor stock valuations seem to be out of the ordinary in the post-settlement years. The actual earnings, stock returns, and GDP growth rates seem to be unusually low in the period between Reg FD and the Global Settlement. We controlled for the effects of these and other potentially relevant factors by examining the effects of Reg FD and the Global Settlement in a regression framework.

Regression Analysis. To examine how Reg FD and the Global Settlement affect bias in analyst forecasts while controlling for the confounding effects of company and analyst characteristics, as well as economic conditions, we estimated the following regression model:

$$\begin{aligned} Bias_{j,t,m} = {} & \alpha_0 + \alpha_1 RegFD_{t,m} + \alpha_2 Glob_{t,m} + \alpha_3 NumA_{j,t,m} \\ & + \alpha_4 CompanySize_{j,t,m-1} + \alpha_5 MB_{j,t,m-1} \\ & + \alpha_6 YearStk_{j,t,m} + \alpha_7 YearIBES_{j,t,m} \\ & + \alpha_8 NumStk_{j,t,m} + \alpha_9 NumInd_{j,t,m} \\ & + \alpha_{10} BrokerSize_{j,t,m} + \alpha_{11} EPSLoss_{j,t} \\ & + \alpha_{12} EPSDecline_{j,t} + \alpha_{13} Litigation_j \\ & + \alpha_{14} Labor_{j,t,m-1} + \alpha_{15} ActualGDP_t \\ & + \alpha_{16} UnexpectedGDP_{t,m} + \beta Month_m + \gamma Year_t \\ & + \delta_j \sum DCompany_j + \varepsilon_{j,t,m}. \end{aligned} \tag{4}$$

In Equation 4, $Bias_{j,t,m}$ is the mean forecast error for all forecasts for company j made in month m relative to the end of fiscal year t, calculated according to Equations 1–3. $RegFD_{t,m}$ equals 1 for forecasts made between 23 October 2000 and 20 December 2002. $Glob_{t,m}$ equals 1 for forecasts made after 20 December 2002. A negative sign for the coefficient of $RegFD_{t,m}$ or $Glob_{t,m}$ would indicate a decline in the bias following, respectively, Reg FD and the Global Settlement.

Lim (2001) argued that the forecast bias is higher when a company's information environment is less transparent—for example, when the company is small and has less analyst coverage. Beckers, Steliaros, and Thomson (2004) showed that the number of analysts following a stock affects the accuracy of the consensus earnings forecast. Hence, we used analyst coverage and company size as proxies for the degree of information transparency. Analyst coverage, $NumA_{j,t,m}$, is defined as the number of outstanding forecasts used in I/B/E/S's monthly consensus calculation. Analyst coverage represents the number of analysts following company j in month m for fiscal year t. $CompanySize_{j,t,m-1}$ is defined as the natural log of the company's market capitalization at the end of the previous month.

Analysts tend to forecast more accurately when they have more experience and resources (Clement 1999; Lim 2001). We measured company-specific experience as the number of years analyst i has been following company j ($YearStk_{j,t,m}$). We measured general experience as the number of years since analyst i first appeared in the I/B/E/S database ($YearIBES_{j,t,m}$). $BrokerSize_{j,t,m}$ is the number of analysts who work for the same employer during the same forecast year as the analyst who makes the forecast. Analysts who work for larger firms tend to have more resources at their disposal.

Clement (1999) found that analysts' forecasts are less accurate the more stocks and the more industries they follow. $NumStk_{j,t,m}$ is the number of stocks for which analyst i supplies at least one forecast within the calendar year. $NumInd_{j,t,m}$ is the number of two-digit SIC industries for which analyst i supplies at least one forecast within the calendar year.

Previous studies have found that forecasting is more difficult when companies report a loss or a decline in earnings (Brown 2001). The $EPSLoss_{j,t}$ indicator equals 1 when the corresponding actual earnings of company j are negative. The $EPSDecline_{j,t}$ indicator equals 1 when actual earnings in fiscal year t are lower than actual earnings in the previous year.

Matsumoto (2002) argued that companies in industries with a higher risk of shareholder lawsuits and/or greater reliance on implicit claims with stakeholders are more likely to avoid missing analyst forecasts. The $Litigation_j$ indicator equals 1 for companies in high-litigation-risk industries: SIC codes 2833–2836 (biotechnology), 3570–3577 and 7370–7374 (computers), 3600–3674 (electronics), and 5200–5961 (retailing).

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Matsumoto (2002) also argued that labor-intensive companies try to avoid missing analyst forecasts because their stakeholders are concerned about company credit risk. Labor intensity, $Labor_{j,t,m-1}$, is defined as 1 minus the ratio of gross plant, property, and equipment (PPE) to total gross assets, where gross PPE is the quarterly Compustat item 118 and total gross assets is item 44 plus item 41. $Labor_{j,t,m-1}$ is measured at the end of the last quarter preceding forecast month m.

Richardson, Teoh, and Wysocki (2004) found lower forecast bias for companies with high growth opportunities. We used the market-to-book ratio ($MB_{j,t,m-1}$) at the end of the last quarter preceding the forecast month as a proxy for growth opportunities. The ratio is calculated as the market value of equity divided by the book value of common equity (Compustat quarterly data item 14 multiplied by item 61 and divided by item 59).

We used both the real GDP growth rate and the unexpected change in the real GDP growth rate to capture analysts' inability to forecast earnings accurately if the state of the economy changes substantially. $ActualGDP_t$ is the actual real GDP growth rate in fiscal year t. $UnexpectedGDP_{t,m}$ is defined as the difference between the expected real GDP growth rate and the actual real GDP growth rate in fiscal year t. For earnings forecasts made more than nine months before the fiscal year-end date, the expected real GDP growth rate in fiscal year t is defined as the real GDP growth rate in the quarter for which analysts made earnings forecasts. For forecasts made in Q2 (seven to nine months before the fiscal year-end date), we calculated the expected real GDP growth rate as (Growth in Q1 + 3 × Growth in Q2)/4. For forecasts made in Q3 (four to six months before the fiscal year-end date), we calculated the expected real GDP growth rate as (Growth in Q1 + Growth in Q2 + 2 × Growth in Q3)/4. For forecasts made within the three months before the fiscal year-end date, $UnexpectedGDP_{t,m}$ is set to zero.

Prior research and our results in Table 2 show that forecasts made earlier in the fiscal year are less accurate (Richardson, Teoh, and Wysocki 2004). To control for forecast horizon, we used $Month_m$, defined as the number of months until the fiscal year-end date. For example, for an analyst forecast made in October 1999 for the fiscal year ended December 1999, $Month_m$ equals 2. Richardson, Teoh, and Wysocki (2004) found that forecast bias has been declining gradually since the early 1990s. To address the concern that our results may be driven by this trend, we included a calendar year variable, $Year_t$, in the regression model (Equation 4). To control for unobserved company effects, we estimated the regressions with fixed company effects ($DCompany_j$).

The first set of estimation results in **Table 3** is for the regression model (Equation 4). The results imply that forecast bias declined by 0.24 percent of the stock price after the introduction of Reg FD. This finding confirms our earlier conjecture that the increase in forecast bias following Reg FD (observed in our univariate results) was driven by unexpectedly poor macroeconomic conditions. The decline in forecast bias following Reg FD is consistent with Lim's prediction (2001) that analysts become less optimistic when they rely less on insider information.

After the Global Settlement, the forecast bias is lower by 0.96 percent of the stock price compared with the forecast bias before Reg FD. This result is consistent with our univariate findings and implies that the Global Settlement and related regulations successfully neutralized analysts' conflicts of interest. The positive coefficient on *Month* suggests the presence of the walk-down trend. Forecast bias is high for earlier forecasts and becomes lower over time. On average, forecast bias increases by 0.14 percent of the stock price per month with the length of the forecast horizon.

Because the Global Settlement is an enforcement agreement between U.S. regulators and the Big-12 banks, we next examined whether the impact of the Global Settlement is limited to the Big-12 banks or whether there are spillover effects on other analysts.[16] In a recent study, Barber, Lehavy, McNichols, and Trueman (2006) reported that the proportion of buy recommendations declined significantly among all analysts after the implementation of NASD Rule 2711. They also documented that the decline was stronger for the sanctioned banks. Whether the Global Settlement has had a differential impact on analyst forecast bias, however, remains an open question.

To identify the differential impacts of Reg FD and the Global Settlement on Big-12 analysts, we compared the bias in the forecasts of Big-12 analysts with the bias in the forecasts of other analysts. In a univariate comparison, we found that, on average, the forecasts of analysts working for the Big-12 banks are statistically significantly less biased than the forecasts of their counterparts in each of the three periods. The differences, however, are economically trivial. For example, the difference between the mean forecast bias of Big-12 analysts and that of other analysts is –0.04 percent of the share price in the pre–Reg FD period, –0.09 percent after Reg FD, and –0.05 percent after the Global Settlement.

Table 3. The Impact of Reg FD and the Global Settlement on Forecast Bias

	(1)		(2)	
	Coefficient	*t*-Statistic	Coefficient	*t*-Statistic
RegFD	−0.24**	−3.29	−0.16*	−2.05
Glob	−0.96**	−10.68	−0.86**	−9.51
CompanySize	0.65**	16.89	0.67**	17.52
NumA	0.02**	3.39	0.01**	2.68
MB	−0.03**	−5.97	−0.03**	−5.59
YearStk	0.01	1.58	0.01**	2.59
YearIBES	0.00	1.54	0.00	0.78
NumStk	0.00*	−2.38	0.00*	−2.05
NumInd	−0.01	−1.18	−0.01	−1.40
BrokerSize	0.00	−1.64	0.00	−0.41
EPSLoss	5.40**	43.20	5.23**	40.53
EPSDecline	2.40**	62.82	2.38**	60.63
Litigation	−0.03	−0.24	−0.08	−0.66
Labor	0.52	2.12	0.47	1.89
ActualGDP	−0.04*	−2.05	−0.03	−1.23
UnexpectedGDP	−0.03**	−6.26	−0.04**	−6.61
Big12			−0.06**	−3.05
Big12 × RegFD			−0.07*	−2.04
Big12 × Glob			0.03	1.34
Month	0.14**	51.70	0.13**	47.76
Year	0.03*	2.16	0.02	1.09
Adjusted R^2	0.46		0.45	
No. of observations	434,268		434,268	
No. of companies	7,315		7,315	

Notes: This table presents the coefficients obtained from Equation 4. The dependent variable is earnings forecast bias, defined as the difference between the mean of all forecasts made in a particular month for a particular company and a particular fiscal year and the realized EPS, scaled by the stock price and multiplied by 100. The RegFD indicator equals 1 for forecasts made between 23 October 2000 and 20 December 2002. The Glob indicator equals 1 for forecasts made after 20 December 2002. Analyst coverage, NumA, is the number of outstanding forecasts used by I/B/E/S to calculate monthly consensus. CompanySize is the natural log of a company's market capitalization. Market-to-book ratio, MB, is the market value of equity divided by the book value of common equity. Company-specific experience, YearStk, is the number of years since the analyst made her first forecast for a particular stock. General experience, YearIBES, is the number of years since the first day the analyst appeared in I/B/E/S. NumStk and NumInd are the number of stocks and the number of industries covered by the analyst, respectively. The EPSLoss indicator equals 1 when the corresponding actual earnings of company *j* are negative. The EPSDecline indicator equals 1 when the realized earnings in fiscal year *t* are lower than the realized earnings in the previous year. BrokerSize is the number of analysts working for the employer of the analyst who makes the forecast. The litigation risk indicator, Litigation, equals 1 for companies in high-litigation-risk industries. Labor intensity, Labor, is (1 − Gross PPE/Total gross assets). The regressions are estimated with fixed company effects. The reported *t*-statistics reflect robust standard errors adjusted for heteroscedasticity and clustering by company.

*Significant at the 5 percent level.
**Significant at the 1 percent level.

To see whether the differential impacts of Reg FD and the Global Settlement on Big-12 and other analysts change when we control for company and analyst characteristics, as well as economic conditions, we re-estimated the regression model (Equation 4) with the Big-12 indicator and its interactions with the Reg FD and Global Settlement indicators included as additional independent variables.[17] The second set of results in Table 3 is for this regression. Consistent with our univariate results, the Big-12 indicator and its interaction with Reg FD are significant in statistical but not in economic terms. More importantly, the interaction of the Big-12 indicator with the Glob indicator is insignificant, both statistically and economically.

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These results imply that both Big-12 and other analyst forecasts were biased before Reg FD, which is consistent with Lin and McNichols (1998), who found no difference between the earnings forecasts of analysts affiliated with banks involved in underwriting deals with the covered companies and the forecasts of unaffiliated analysts. These results also imply that the impact of the Global Settlement and related regulations is the same among Big-12 and other analysts. This finding may reflect the fear of non-Big-12 firms that they may become targets of similar investigations. In addition, because Big-12 banks no longer reward optimism, the incentive for lower-tier analysts to make optimistic forecasts as a means of moving up to the bigger banks has also been reduced. Finally, the rules and regulations introduced by the SEC, NYSE, and NASD around the time of the Global Settlement covered all analysts.

We checked the robustness of our main conclusion—that forecast bias declined after both Reg FD and the Global Settlement—in a number of ways. First, we used an alternative definition of the forecast bias by normalizing it by the book value of equity per share.[18] Second, we changed the cutoff dates for each period by using the effective date of Rule 2711 instead of the announcement date of the Global Settlement. Third, to ensure that our conclusions were unaffected by changes in the sample composition across the three subperiods, we required at least one forecast by the same analyst for the same company in all three periods. Fourth, we dropped observations with stock prices under $5 to avoid any potential biases induced when the scaling factor is a small number. Fifth, we extended our sample period to include an earlier period (January 1984–December 1995). In all these cases, the results (not reported here) remain qualitatively the same as those reported in Table 3, confirming that forecast bias declined after Reg FD and especially after the Global Settlement.

We also examined the breadth of these effects by estimating forecast bias regressions (Equation 4) separately for 12 business sectors and for subsamples formed on the basis of annual quintile sorts by company size and analyst coverage.[19] The results (not reported here) show that the effects of the Global Settlement are negative for 11 of 12 sectors and are statistically significant for 9 sectors. The effects of Reg FD are negative for 8 of 12 sectors, but significantly so for only 6 sectors. Our results also show that the effect of Reg FD is concentrated among smaller companies and companies with low analyst coverage, whereas the effect of the Global Settlement is more widespread, with no clear cross-sectional pattern.

Conclusion

Analysts' conflicts of interest were evident before the Global Research Analyst Settlement and were not limited to the 12 banks covered by it. Reg FD made analysts less dependent on insider information and thus diminished analysts' motives to favor company managers by inflating their earnings forecasts. The impact of Reg FD is more significant for companies with a less transparent information environment in which insider information has the most value.

Introduced in 2002, the Global Settlement and related regulations had an even bigger impact than Reg FD on analyst behavior. After the Global Settlement, the mean forecast bias declined significantly, whereas the median forecast bias essentially disappeared. Although disentangling the impact of the Global Settlement from that of related rules and regulations aimed at mitigating analysts' conflicts of interest is impossible, forecast bias clearly declined around the time the Global Settlement was announced. These results suggest that the recent efforts of regulators have helped neutralize analysts' conflicts of interest.

We thank Donal Byard, Terrence Martell, and seminar participants at Baruch College for helpful comments. Armen Hovakimian gratefully acknowledges the financial support of the PSC-CUNY Research Foundation of the City University of New York.

This article qualifies for 1 CE credit, inclusive of 1 SER credit.

Notes

1. Several rules and regulations were enacted around the Global Research Analyst Settlement—for example, NASD Rule 2711, NYSE Rule 472, and Regulation Analyst Certification. Because they were introduced over a relatively short period, determining the separate impact of each one of these regulatory actions is impossible. Nevertheless, all these rules and regulations share the same goal of reducing analysts' conflicts of interest. Therefore, we use the term Global Settlement to represent all the rules and regulations enacted around the Global Research Analyst Settlement to address analysts' conflicts of interest.
2. Scherbina (2004) found a negative relationship between the estimated bias that arises from self-selection in coverage and subsequent stock returns. Her results suggest that retail

investors fail to adjust for the bias. Malmendier and Shanthikumar (2007) found that retail investors react to stock recommendations literally. Institutional investors buy stocks that have "strong buy" ratings and sell stocks that have "buy" ratings, whereas retail investors buy in both cases. Kwag and Shrieves (2006) found that persistence in forecast errors can lead to potentially profitable trading strategies.

3. Overall, these studies found either no change (Bailey, Li, Mao, and Zhong 2003) or a decrease in forecast accuracy (Agrawal, Chadha, and Chen 2006; Mohanram and Sunder 2006) and forecast dispersion (Agrawal, Chadha, and Chen 2006) following Reg FD.
4. Kadan, Madureira, Wang, and Zach (2009) documented that stock recommendations have become less optimistic since the Global Settlement. Furthermore, they found that the likelihood of an optimistic recommendation is no longer associated with analyst affiliation. Ferreira and Smith (2006) found that investors have not changed the way they respond to analysts' changes in recommendations since Reg FD. Examining bid–ask spreads and trading activity following Reg FD, Lee, Rosenthal, and Gleason (2004) found no significant increase in volatility or in the adverse-selection component of bid–ask spreads.
5. Reg AC took effect on 14 April 2003. See the joint report by the NASD and NYSE (2005) for the effectiveness of the new rules.
6. The 10 investment banks are Bear Stearns, Citigroup, Credit Suisse First Boston, Goldman Sachs, J.P. Morgan, Lehman Brothers, Morgan Stanley, Merrill Lynch, UBS, and U.S. Bancorp Piper Jaffray. In 2008, Bear Stearns and Merrill Lynch were taken over because of their deteriorating financial positions, whereas Lehman Brothers ended up in bankruptcy. Because our sample period ends in 2006, these events did not affect our results.
7. These two investment banks are Deutsche Bank and Thomas Weisel Partners.
8. Because prior studies (e.g., Lin and McNichols 1998) found no cross-sectional differences in forecast bias between affiliated and unaffiliated analysts, one would not reasonably expect cross-sectional differences in the impact of the Global Settlement on these two analyst types.
9. Therefore, one would not reasonably expect cross-sectional differences in the impact of the Global Settlement on self-selection bias.
10. Forecasts for current-year EPS are the forecasts in I/B/E/S with code FPI 1. Forecasts for subsequent-year EPS are the forecasts in I/B/E/S with code FPI 2.
11. We excluded forecasts in the I/B/E/S Excluded Estimates file and forecasts for which actual earnings figures were missing.
12. The I/B/E/S Summary file contains monthly snapshots of consensus-level data and corresponding stock prices. The snapshots are as of the Thursday before the third Friday of every month. The reported stock prices in this file are the last available prices before the Thursday. I/B/E/S's earnings-related data and stock prices are split adjusted.
13. Using stock price to normalize forecast bias is common (see, e.g., Richardson, Teoh, and Wysocki 2004). Later in the article, we discuss the robustness of our findings to alternative scaling of analyst forecast errors.
14. We defined extreme values as those in 1 percent of both tails of the distribution. Variables that took only positive (negative) values were trimmed only on the right (left) tail of the distribution.
15. Realized earnings and forecasts are scaled by the stock price, consistent with the scaling of the bias measure.
16. Other regulations, such as NASD Rule 2711, affect all analysts.
17. In this analysis, for each forecast month of each sample company-year, the mean forecast bias is calculated separately for Big-12 and other analysts.
18. This step also ruled out the possibility that such events as the decimalization of stock prices in August 2000–April 2001 affected our findings.
19. The sector classification for each company is from the I/B/E/S Identifier file.

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MONTANA-DAKOTA UTILITIES CO.
MONTANA CONSUMER COUNSEL
MCC DATA REQUEST
DATED JULY 12, 2013
DOCKET NO. D2012.9.100

MCC-198 Regarding: Dividend Yield Adjustment
Witness: J. Stephen Gaske

a. Please specify the journals or periodicals to which the paper provided as Exhibit No._(JSG-08) was submitted for possible publication. If it was published, please provide a copy of the published version.

b. Please state whether this paper or arguments substantially similar to those in this paper were presented as evidence in one or more FERC proceedings. If so, please identify those proceedings by name and docket number, and provide copies of any FERC orders or decisions addressing these arguments.

Response:

a. Exhibit No.__(JSG-08) was part of a longer paper, "The Nature of Estimation Biases in DCF Analyses of Equity Cost in Rate Regulation," that was presented at the 1988 Annual Meeting of the Financial Management Association. Exhibit No.__(JSG-08) was subsequently submitted to *Public Utilities Fortnightly*, however they declined to publish it because they indicated that it is too technical for their general readership.

b. The analysis which ends on page 6 of the paper in Exhibit No.__(JSG-08) was presented as Appendix C of Comments by New England Power Company (NEPCo) in FERC's generic rate of return Docket No. RM86-12-000.

In generic rate of return orders prior to that proceeding the FERC had described the assumptions that the Commission considered to be most reasonable concerning the timing of dividend payments and growth in dividends. In Order No. 442-A the Commission showed the algebra that it believed would reflect the Commission's assumptions, and used this algebra to supposedly show that the current dividend yield in the DCF model should be multiplied by [1 + **0.50**g] in order to reflect the Commission's assumptions.

The NEPCo Appendix C comments in Docket No. RM86-12-000 demonstrated that there had been a mathematical error in the Commission's prior analysis and that when the Commission's original algebra was corrected, the Commission's assumptions indicated that the dividend yield should be multiplied by [1 + 0.**75**g]. That result is shown in Equation {6} on page 6 of Exhibit No. __(JSG-8).

The Commission acknowledged this comment, but did not explicitly respond to the comment when it issued its decision in Docket No. RM86-12-000. Instead, it ignored its prior erroneous algebraic approach to deriving the dividend yield adjustment factor and attempted to use a numerical example instead of algebra to justify using a factor of [1 + 0.**50**]. However, the numerical example unveiled in the

text of Order No. 461 incorrectly truncated several numbers to reach an erroneous conclusion. When calculated correctly, the example in Order No. 461indicates that the dividend yield adjustment factor should be [1 + 0.**625**].

Pages 6-8 of the paper in Exhibit No.__(JSG-08) address the error in the new reasoning presented by the Commission in Order No. 461. To Dr. Gaske's knowledge, the arguments on pages 6-8 of the paper have never been introduced as evidence in an FERC proceeding.

A copy of Order No. 461 is provided as Attachment A.

Response No. MCC-198
Attachment A

UNITED STATES OF AMERICA
FEDERAL ENERGY REGULATORY COMMISSION

[18 C.F.R. Part 37]

Generic Determination of Rate of Return on Common Equity For Public Utilities

[Docket No. RM86-12-000]

ORDER NO. 461

(Issued December 24, 1986)

AGENCY: Federal Energy Regulatory Commission

ACTION: Final Rule

SUMMARY: The Federal Energy Regulatory Commission (Commission) determines that the average cost of common equity for the jurisdictional operation of electric utilities during the year ending June 30, 1986, was 13.05 percent.

The Commission also modifies the quarterly indexing procedure which establishes and updates the benchmark rates of return. The quarterly updates will no longer be subject to a cap on the charges from quarter to quarter. New benchmarks will be established for filings made on or after February 1, 1987.

These benchmark rates of return will remain advisory only. The benchmark rates of return established as a result of this proceeding are intended to guide companies and intervenors in individual rate cases and to serve as a reference point for the

Commission in its deliberations. The Commission may take official notice of them in individual rate proceedings.

EFFECTIVE: The final rule is effective [30 days after publication in the Federal Register].

FOR FURTHER TECHNICAL INFORMATION CONTACT:

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SUPPLEMENTAL INFORMATION:

UNITED STATES OF AMERICA
FEDERAL ENERGY REGULATORY COMMISSION

[18 C.F.R. Part 37]

Before Commissioners: Martha O. Hesse, Chairman;
Anthony G. Sousa, Charles G. Stalon,
Charles A. Trabandt and C. M. Naeve.

Generic Determination of)
Rate of Return on Common) Docket No. RM86-12-000
Equity for Public Utilities)

ORDER NO. 461

FINAL RULE

(Issued December 24, 1986)

TABLE OF CONTENTS

	Page
I. INTRODUCTION	4
II. SUMMARY AND ANALYSIS OF COMMENTS	9
A. Status of the Rule	9
1. Introduction	9
2. Advisory Only or Presumptive Benchmark Rate	10
a. Comment Summary	10
b. Analysis and Findings	11
3. Rebuttable Presumption Standard	11
a. Introduction	11
b. Comment Summary	12
c. Analysis and Findings	13
4. Termination of the Rulemaking	13
a. Comment Summary	13
b. Analysis and Findings	15

	Page
B. Base Year Cost	17
1. DCF Model Formulation	17
a. Introduction	17
b. Comment Summary and Analysis	17
2. Sample	24
a. Comment Summary	26
b. Analysis and Conclusion	27
3. Dividend Yield	29
a. Introduction	29
b. Comment Summary	30
c. Analysis and Findings	31
4. Growth Rate	33
a. Introduction	33
b. Comment Summary	33
c. Analysis and Findings	43
5. Corroborative Evidence	55
a. Introduction	55
b. Comment Summary and Analysis	55
(i) Risk Premium Analyses	55
(ii) Earnings-Price (E/P) and Earnings-Book (E/B) Ratios	58
6. Flotation Costs	62
a. Introduction	62
b. Comment Summary and Analysis	63
(i) Types of Costs to Be Recovered	63
(ii) Method of Recovery	65
(iii) Form of Recovery	66
c. Flotation Cost Adjustment	67
7. Jurisdictional Risk	69
a. Comment Summary	69
b. Analysis and Conclusion	70

C. Quarterly Indexing Procedure..... 71

1. Introduction..... 71

2. Comment Summary and Analysis..... 71

a. The Six-month Dividend Yield and the 50 Basis Point Cap..... 72
b. Changes in Growth Expectations..... 76

D. Ratemaking Rate of Return..... 78

1. Introduction..... 78
2. Comment Summary..... 81
3. Conclusion..... 83

E. The DCF Method/Cost of Capital Standard..... 84

1. Introduction..... 84
2 Comment Summary..... 84
3. Analysis and Findings..... 87

III. SUMMARY OF CHANGES..... 94

IV. REGULATORY FLEXIBILITY ACT CERTIFICATION..... 94

V. TIMING OF QUARTERLY UPDATES AND EFFECTIVE DATE OF RULE..... 95

VI. REGULATORY TEXT..... 96

APPENDICES
A. List of Commenters
B. Sample of Companies Used for Base Year Dividend Yield Calculation
C. Sample of Companies for Dividend Yield Update

I. INTRODUCTION

In accordance with § 37.4 of its Regulations, the Federal Energy Regulatory Commission (Commission) is determining in this order: (1) the average cost of common equity for the jurisdictional operations of public utilities 1/ for the year ending June 30, 1986 (hereafter the "base year"); and (2) a quarterly indexing procedure to update the cost estimate and establish benchmark rates of return on common equity for use in individual rate cases. This is the third annual proceeding. 2/ The benchmark rate of return established in this proceeding was to have been accorded rebuttable presumption status, but, as discussed below, will remain advisory only.

1/ The terms "public utilities" and "electric utilities" are used interchangeably.

2/ The annual proceedings were first established by Order No. 389, Generic Determination of Rate of Return on Common Equity for Electric Utilities, 49 Fed. Reg. 29,946 (July 25, 1984) (Docket No. RM80-36-000) (Final Rule) (Issued July 19, 1984). The first annual proceeding resulted in Order No. 420, 50 Fed. Reg. 21,802 (May 29, 1985) (Docket No. RM84-15-000) (Final Rule) (Issued May 20, 1985) and Order No. 420-A, 50 Fed. Reg. 34,086 (August 23, 1985) (Docket Nos. RM84-15-001, et al.) (Order Denying Rehearing) (Issued Aug. 20, 1985). The second proceeding resulted in Order No. 442, 51 Fed. Reg. 343 (Jan. 6, 1986) (Docket No. RM85-19-000) (Final Rule) (Issued December 26, 1985) and Order No. 442-A, 51 Fed. Reg. 22,505 (June 20, 1986) (Docket No. RM85-19-001, et al.) (Order Granting in Part and Denying in Part Requests for Rehearing) (Issued June 11, 1986).

The Commission's intent is to produce more accurate and consistent rate of return decisions, to involve the Commission on an ongoing basis in consideration of the financial and operating circumstances of the industry, and, ultimately, to reduce the resources directed to this issue by applicants, intervenors, and the Commission. 3/ The Commission has previously discussed the statutory requirements applicable to electric utility rate filings subject to the Commission's jurisdiction and its reasons for attempting to develop a generic or benchmark approach to the measurement of the cost of common equity for individual electric utilities in rate cases. 4/

On July 21, 1986, the Commission issued a Notice of Proposed Rulemaking (Notice or NOPR) in Docket No. RM86-12-000 5/ proposing to determine: (1) the average cost of common equity for the jurisdictional operations of public utilities for the base year and (2) a quarterly indexing procedure to establish benchmark rates of return on common equity for use in individual rate cases.

3/ Order No. 420, 50 Fed. Reg. at 21,803.

4/ Id.

5/ Generic Determination of Rate of Return on Common Equity for Public Utilities, 51 Fed. Reg. 27,050 (July 29, 1986) (Docket No. RM86-12-000) (Notice of Proposed Rulemaking) (Issued July 21, 1986).

In the NOPR the Commission proposed to use the same Discounted Cash Flow (DCF) approach to determining and updating the generic rate of return on equity that it had previously adopted in Order No. 420 and in Order No. 442-A. 6/ The Commission further proposed to adopt the finding of Order No. 442 that there is no appreciable difference in risk between the wholesale and retail operations of electric utilities. 7/ The Commission sought comment on a number of issues with regard to the ratemaking rate of return concept. The Commission also proposed a codification of the procedure for determining the generic cost of equity and reduced filing requirements for rate filings that make use of the generic rate of return on equity. Finally, the Commission proposed to make the generic rate of return on equity applicable on a binding basis to rate filings by electric utilities that possessed certain bond rating characteristics. The Commission asked for comment on this proposal, particularly, on how the proposal would work in conjunction with the burden of proof that the Federal Power Act places upon a utility filing a rate increase. 8/

6/ 51 Fed. Reg. at 27,051-52.

7/ 51 Fed. Reg. at 27,052.

8/ 16 U.S.C. § 824d(e) (1982).

In response to the NOPR, 29 persons submitted comments: 3 regulatory commission staffs, 19 individual utilities or groups of utilities, an electric utility trade association, 5 individual utility customers or groups of utility customers or representatives of utility customers, and 1 individual. 9/ Most comments favored primary reliance on the DCF approach to estimate the cost of common equity and several included comprehensive studies estimating the cost during the base year. Most commenters also favored the Commission's proposal to incorporate an estimate of the industry average flotation cost in the benchmark rate of return. There was general support for the use of a dividend yield-based indexing mechanism. Finally, no commenter favored the proposal to make the benchmark applicable on a binding basis. In fact, most commenters actively opposed it.

In response to the comments and after consideration of the issues involved, the Commission has decided to adopt the procedure used in the second annual proceeding, as described in Order No. 442-A, for determining and updating the benchmark

9/ These commenters are referred to by acronyms in the text that follows. See Appendix B for a list of the acronyms and the parties that they signify. Initial comments are hereinafter referred to as "IC" and reply comments as "RC."

rate of return, except that the updates will no longer be subject to a 50 basis point cap.

As detailed below, the Commission estimates that the average cost of common equity for the jurisdictional operation of electric utilities during the base year was 13.05 percent. This is based on a required rate of return of 13.02 percent and a flotation cost adjustment of .03 percent.

The benchmark rate of return will be updated on a quarterly basis for use in individual rate cases. In January 1987, the Commission will announce the first benchmark rate of return, based on the dividend yields for the last two quarters of 1986.

As discussed more specifically below, the benchmark rates of return will remain advisory only. The benchmark rates established as a result of this proceeding are intended to guide companies and intervenors in individual rate cases and to serve as a reference point for the Commission in its deliberations. The Commission may take official notice of them in individual rate proceedings and the Commission will determine the weight to accord the applicable benchmark rates based on the record in each case. In this regard, the Commission urges participants in rate cases to evaluate the reasonableness of the applicable benchmarks in light of any special circumstances of the filing utility. 10/

10/ The primary exception to the application of the benchmark rate of return to a utility during a rate case is when the utility is significantly more or less risky than the average utility.

Table 1. Estimates of the Average Cost of Common Equity to Electric Utilities for the Year Ending June 30, 1986

Commenter 1/	Sample Size	Model Used 2/	Current Dividend Yield	Quarterly Dividend Adjustment	Adjusted Dividend Yield	Constant Growth Rate	Required Rate of Return	Flotation Cost Adjustment	Cost of Common Equity
PEPCO	107	3/	NR	NR	8.20%	6.25%	14.45%	.30%	14.75%
	107	3/	NR	NR	8.20	6.00	14.20	.30	14.50
EEI	99	420	8.35	.23	8.58	5.50	14.08	.02	14.10
	99	420	8.35	.20	8.55	4.80	13.35	.02	13.37
NEP	99	420	8.20	.23	8.43	5.50	13.93	NR	NR
	99	420	8.20	.20	8.40	5.00	13.40	NR	NR
AUS	99	420	8.28	.19	8.47	4.61	13.08	.11	13.19
Southern	NR	4/	NR	NR	NR	4.01	NR	NR	13.13
	NR	420	8.03	NR	NR	4.01	NR	NR	12.63
Commission	99	420	8.25	.17	8.42	4.60	13.02	.03	13.05
FA Staff	99	420	8.18	.19	8.37	4.60	12.97	.03	13.00
Cooperatives	84	420	7.37	.18	7.55	4.76	12.31	NR	NR

NR = Not reported.

1/ See Appendix A for identification of commenters.
2/ "420" refers to the Commission's "420 Model".
3/ "420 Model" with quarterly dividend adjustments factor (1+ .625n).
4/ Effective rate DCF Model.

II. SUMMARY AND ANALYSIS OF COMMENTS

A. Status of the Rule

1. Introduction

In Order No. 389, 11/ the Commission adopted a transitional provision, 18 C.F.R. § 37.8, which made the first two annual proceedings advisory only. The Commission stated that during this period the benchmark rates of return "are intended to provide guidance to parties and serve as a point of departure for the Commission in setting allowed rates of return [and to provide a valid test of the potential consequences of moving to a rebuttable presumption standard...." 12/ Order No. 389 established a presumption that beginning with the third annual proceeding the allowed rate of return in an individual case would be the benchmark rate of return in effect at the time a rate schedule is filed.

Commenters address three issues regarding the status of the rule. First, commenters address whether the Commission should make the benchmark rate presumptively the rate of return in individual rate cases. Second, commenters address the standard the Commission should adopt for rebutting the presumption if the benchmark rate is changed from an advisory only to a presumptive rate. Third, commenters address whether the rulemaking should be terminated. These issues are addressed in turn.

11/ 49 Fed. Reg. at 29,946.
12/ Id. at 29,954.

2. Advisory Only or Presumptive Benchmark Rate

a. Comment Summary

All commenters that address this question recommend that the benchmark rate remain advisory. 13/ Some commenters argue that the Commission has not had sufficient experience with which to judge the rule. 14/ Commenters note that the Commission made substantial revisions in the second proceeding. In this regard, one utility states that there are still unresolved issues, e.g., the ratemaking rate of return. 15/ Another commenter states that the cost calculations of the commenters still vary significantly due to differences in the methods for measuring the components of the cost. 16/

One commenter questions the need for a rebuttable presumption since rate of return is so seldom the subject of litigation. 17/ Another alleges that during the advisory period, the benchmark rates of return have not been used in any significant way. As a result this commenter questions the value in adding another rate of return analysis which would have to be addressed in individual cases. 18/

13/ AEP IC at 4; NEP IC at 22-24; CGE IC at 8-9; BEC RC at 15; RP at 14.

14/ PEPCO IC at 3; BEC IC at 2.

15/ CGE IC at 8.

16/ AEP IC at 4.

17/ AEP IC at 4.

18/ SCE RC at 3.

b. Analysis and Findings

The Commission agrees that it has not had enough experience with the rule to justify moving to the rebuttable presumption standard at this time. While the Commission made changes in its estimation procedures over the last year, they were not the cause for the lack of experience with the rule. Rather, the Commission has yet to decide a case in which the benchmark was available and rate of return is an issue. With respect to the differences among commenters relative to the proper method for estimating the cost of common equity, the Commission's findings in the last two proceedings have reduced the number of issues and narrowed these differences significantly.

However, questions remain regarding implementation of a rebuttable presumption rule. Therefore, the Commission has decided not to move beyond the advisory only status for the benchmark in this proceeding. 19/

3. Rebuttable Presumption Standard

a. Introduction

The Commission proposed a significant risk difference standard. A utility would receive the benchmark rate if it

19/ In the Notice, the Commission proposed to amend its regulations to reduce the filing requirements in individual rate cases. Since the benchmark rates of return will not be implemented as rebuttable presumptions in individual rate cases, however, changing the filing requirements would not be appropriate. Thus the Commission is not changing its filing requirements in this proceeding.

fell within the midrange for the aggregate risk measures established in this proceeding. In particular, the Commission proposed to use bond ratings as the aggregate risk measure and sought comment on alternative measures.

b. Comment Summary

The consensus among the commenters is that the rule should allow significant flexibility for challenging the presumption. Many commenters viewed the Commission's proposal as setting up an "irrebuttable" presumption. 20/ Commenters generally argue that no single measure can adequately distinguish the relative risk of all utilities. 21/ One commenter argues that the Commission's proposed significant risk difference standard is similar to the risk categorization scheme that was proposed in Docket No. RM80-36 and rejected in Order No. 389. 22/

Most commenters argue that there is not enough homogeneity of risk in the industry to support the use of a rebuttable presumption standard. 23/ They argue that the record supports a finding of risk heterogeneity not homogeneity. Two utilities

20/ See e.g., PEPCO IC at 3; BEC IC at 2.

21/ See e.g., AEP IC at 3 AUS IC at 61; EEI IC at A-9.

22/ See BGE IC at 2-3.

23/ See e.g., AUS IC at 63-66; EEI IC at Attachment A; Cooperatives IC at 35.

recommend that the Commission consider dividing the industry into four risk categories. 24/

Commenters also raise concerns with the Commission's proposed use of bond ratings. 25/ They argue that there is inadequate support for the assumption of a correlation between bond ratings and common equity risk. 26/ Some present empirical evidence to show a weak relationship. 27/ Finally, one commenter argues that the bond rating bands within which utilities would be presumed to be of average risk are arbitrary and unsupportable in the context of determining the allowed rate of return. 28/

c. Analysis and Findings

As discussed above, the Commission has decided to keep the rule advisory only. By keeping the rule advisory only, the Commission will have an opportunity to further evaluate the implementation issues associated with distinguishing individual company risks from industry average risks.

4. Termination of the Rulemaking

a. Comment Summary

A number of commenters recommend that the Commission terminate or give serious consideration to terminating the

24/ See DE IC at 10; APS IC at 2.

25/ See AUS IC at 61, WCG IC at 11.

26/ See AUS IC at 61; WCG IC at 11.

27/ See e.g., WCG IC at 13.

28/ See SCE IC at 8.

rulemaking. 29/ One commenter supports the Commission's generic approach and others appear to endorse the concept. 30/

Critics state that the objectives of the rule are not or will not be met. Commenters allege that using the benchmark rate of return during the advisory period would not have produced more accurate and consistent rate of return decisions. According to FA Staff, in only three of the eight cases in which it filed rate of return testimony did the benchmark rate come reasonably close to the Commission Trial Staff's recommendation. 31/ Some commenters also claim that, because of the degree of heterogeneity in the industry, the benchmark rate would only reasonably be applicable to a small number of companies. 32/

29/ Cooperatives IC at 24-34; FA Staff IC at 29 31; AEP RC at 1-2; AUS RC at 36-37; BEC RC at 14-15; EEI RC at 2-4; PEPCO RC at 1-7; PSC RC at 1-3; SCE RC at 2-4; SW RC at 1; VEPCO RC at 1-2; Second Cooperatives RC at 12-14.

30/ NSP IC at 8 MINN IC at 10; WVCD IC at 2-4; AWW RC at 1.

31/ FA Staff IC at 29-31.

32/ Cooperatives IC at 24-34; FA Staff IC at 29-31; Second Cooperatives RC at 12-14; Cooperatives RC at 7-12. Cooperatives repeat the argument they made in the prior proceeding that the industry is too heterogenous for a benchmark to be useful. They contend that the "industry's lack of homogeneity is obvious from the numerous risk groups within the industry that are recognized by investment analysts and investors." See Cooperatives IC at 35. The essence of this argument is that the cost of equity is so widely diverse that a single benchmark is not appropriate. Given the Commission's decision to maintain the benchmark rates of return in an

(FOOTNOTE CONTINUED ON NEXT PAGE)

Commenters argue that the generic rule has not and probably will not save resources. 33/ One commenter states that the cost of service for most cases is settled and the advisory benchmark rates have not had any impact on the rate of settlements. 34/ In fact, one commenter also alleges that, during the past advisory period, the benchmark has not been used to any significant degree by rate case participants or the Commission. 35/

Finally, commenters allege that the annual generic proceedings have not resulted in more direct and current Commission involvement in the industry's financial and operating circumstances. 36/

b. Analysis and Findings

While few commenters support the rule, many commenters actively oppose it. The central areas of disagreement in the evaluation of the generic approach to the rate of return on common equity issue are three. The first concerns the extent

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(FOOTNOTE CONTINUED FROM PREVIOUS PAGE)

advisory only status, there is no presumption that the benchmark is applicable to any segment of the industry. However, an industry-average benchmark is useful as a point of departure, regardless of the distribution of the costs of equity in the industry.

33/ Second Cooperatives RC at 14; Cooperatives IC at 30.

34/ Cooperatives IC at 27-28.

35/ SCE RC at 2.

36/ Cooperatives IC at 30-31.

to which the quality of rate of return decisions (including the accuracy of estimates of the cost of common equity for the utility in question) is enhanced by the case-by-case approach. Second, there is disagreement on the impacts on interested parties. The third area of dispute is in the valuation of the resources devoted to this issue, including the time and cost to all parties and the Commission.

All three of these issues are factual. However, little or no new empirical evidence is presented in this proceeding bearing on the merits of these issues.

With regard to the resource issue, the Commission observes the cost of service portion of a high percentage of rate filings is settled. To date, no rate filing made since the benchmark rate has been in effect has received a final decision on rate of return. However, the potential benefit from generic resolution of recurrent rate issues remains a desirable objective. Therefore, the Commission believes that the benefits of further analysis of the generic approach outweigh the costs.

In addition, estimates of the industry average cost of common equity can be used by the Commission in its analysis in individual rate cases. The existence of industry average cost estimates also gives parties an indication of the Commission's current view of capital costs. If the Commission places weight on this analysis in individual rate proceedings, it could narrow the differences among parties.

B. Base Year Cost

1. DCF Model Formulation

a. Introduction

In the Notice, the Commission expressed its intention to rely on the discounted cash flow (DCF) method for estimating the rate of return on common equity. 37/ The particular formulation of the DCF model that the Commission proposed to rely on is the same one used in Order Nos. 420 and 442-A:

$$(1) \qquad k = \frac{D_0}{P_0}(1 + .5g) + g$$

where:

k = market required rate of return

$\frac{D_0}{P_0}$ = current dividend yield (current annual dividend rate divided by current market price)

g = annual dividend growth rate

$(1 + .5g)$ = adjustment factor for quarterly dividend payments

This model, first adopted in Order No. 420, will hereinafter be referred to as the "420 Model."

The Commission requested comments on this model.

b. Comment Summary and Analysis

CGE argues that, in the dividend yield computation of the DCF model, "it would be more correct to use the next period dividend (D_1) instead of the 'indicated dividend rate.'" The DCF model is prospective, and, as such, the future

37/ 51 Fed. Reg. at 27,050.

income stream can only be considered by utilizing the next period dividend." 38/

In Order No. 420, the Commission rejected CGE's model because it "assumes that investors receive dividends once a year. Clearly, [it does not] correctly characterize the real world where dividends are generally paid out to investors on a quarterly basis." 39/ For the same reasons, we reject CGE's model in this proceeding.

NEP argues for a modification to model (1), where, instead of multiplying the current dividend yield by one plus .5 times the growth rate, it is multiplied by one plus .75 times the growth rate. NEP concludes that model (1) is "inconsistent with the assumption that, on average, annual dividend increases are a half year away." 40/ Similarly, PEPCO argues for a .625 factor rather than the .5 factor. 41/

Model (1) was originally adopted in Order No. 420, wherein the Commission stated that it attempted:

> to approximate the average expected annual dividends received during the first year. Assuming that some companies will increase their dividend rate within the first quarter, some during the second quarter, etc., [this model] attempts to approximate

38/ CGE IC at 3.

39/ Order No. 420, 50 Fed. Reg. at 21,805.

40/ NEP IC at Appendix C.

41/ PEPCO IC at Attachment A, page 2.

> the average amount of dividends that the average investor (or, equivalently, investors in the average company) would expect to receive during the first year. 42/

The Commission made a similar assumption in deriving its effective rate of return model in Order No. 442. 43/

Assume that a public utility stock is purchased on January 1 and dividends are paid the following March 31, June 30, September 30 and December 31. If the quarterly dividend is $ .25 and the dividend growth rate is 5 percent, the four equally likely scenarios are modeled below:

3/31	6/30	9/30	12/31	Total	Dividend Payment Date
------Dividend Received-------					
$.25	.25	.25	.26	1.01	Rate Increased During 4th Quarter
$.25	.25	.26	.26	1.02	Rate Increased During 3rd Quarter
$.25	.26	.26	.26	1.03	Rate Increased During 2nd Quarter
$.26	.26	.26	.26	1.04	Rate Increased During 1st Quarter
				$ 1.025	Average

As can be seen, the average yearly dividend received under the four equally likely scenarios is equal to the current annual dividend rate ($ .25 x 4 = $1.00) multiplied by one plus one half of the dividend growth rate (1 + .025). This demonstrates the reasonableness of the Commission's choice of model (1), which implements its stated goal of approximating "the average expected annual dividends received during the first year."

42/ Order No. 420, 50 Fed. Reg. at 21,806.

43/ Order No. 442, 51 Fed. Reg. at 348.

AUS argues for a DCF model which measures the investors' effective required rate of return: 44/

$$(2) \qquad k = \frac{P_0}{4P_0}\left[(1+k)^{.75} + (1+k)^{.5} + (1+g)(1+k)^{.25} + (1+g)\right] + g$$

where the variables are generally defined the same as in model (1).

In Order No. 442, Model (2) was used by the Commission to measure the investors' effective market required rate of return on common equity. 45/ The Commission described the formula as reflecting "the benefits to investors of getting the dividends in four quarterly installments rather than in a lump sum at the end of the first year. These benefits are, of course, the additional return investors may obtain by reinvesting the dividends received quarterly in the same or another comparable investment until the end of the year." 46/ Because the investor retains the quarterly payments and can reinvest them, the utility's cost should reflect only the nominal amount of the dividends. Model (2) would thus compensate the investor twice -- once by the utility and once through the investor's reinvestment of the dividends in some

44/ AUS IC at 42. See also VEPCO IC at 1.

45/ 51 Fed. Reg. at 348.

46/ Id.

other alternative investment. Therefore, model (2) does not accurately reflect the cost to the utility and is rejected. 47/

Southern argues for the following DCF model: 48/

$$(3) \qquad k = \left[\frac{D_{10}(1+k)^{.75} + D_{20}(1+k)^{.5} + D_{30}(1+k)^{.25} + D_{40}}{P_0\,(1-f)}\right](1+g) + g$$

where:

k = cost of common equity

D_{10}, D_{20}, D_{30}, D_{40} = average dividend for 1st through 4th quarters, respectively

P_0 = stock price

g = growth rate

f = flotation cost rate

Southern's proposed model is another variation of the investors' effective required rate of return model. It will be rejected for the same reasons the Commission rejects model (2) proposed by AUS.

Southern also makes the following comment in support of its version of the DCF model: "When a utility pays dividends

47/ "The utility is only required to provide the quarterly dividends which give the investor the opportunity to earn these additional earnings through reinvestment." Order No. 442, 51 Fed. Reg. at 349. "The return that investors expect from the firm does not include the income that they expect to receive from the reinvestment of dividends. Investors have the opportunity to produce this income by their own actions in reinvesting the dividend portion of their return." Order No. 442-A, 51 Fed. Reg. at 22,508.

48/ Southern IC at 14.

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on a quarterly basis rather than an annual basis, the utility is unable to use those funds in the course of the year. This is an opportunity cost resulting from the time value of money. Therefore, it is more expensive for a utility to pay dividends on a quarterly basis than an annual basis." 49/

Southern's argument is flawed. Contrary to Southern's assertion, it is more expensive, in nominal terms, for a utility to pay dividends on an annual basis rather than on a quarterly basis, if a constant dividend payout ratio is to be maintained. To illustrate this, consider two identical utilities, both of which intend to pay all of their earnings in dividends (i.e., have 100 percent dividend payout ratios), but one pays a single annual dividend ("Utility A") and the other pays dividends quarterly ("Utility Q").

When Utility Q pays its dividends out of earnings quarterly, the investor has the opportunity to reinvest these dividends at the required rate of return and earn an effective rate of return which would be equal to that calculated by Order No. 442's "effective rate model." This is similar to compound interest in a savings account. Conversely, since Utility A does not pay dividends during the course of the year, it has the opportunity to reinvest its earnings. At the end of the year, Utility A will pay a single dividend

49/ Id. at 14.

which, since it is paid out of these higher earnings, is higher than the sum of Utility Q's four quarterly dividends. However, Utility A's rate of return, using this higher end-of-year dividend, is identical to that of Utility Q. The value of Utility A's single end-of-year dividend is equal to the sum of Utility Q's four quarterly dividends plus the investors' reinvestment income. 50/

The investor is indifferent to the utility's dividend payment practice. Utilities which pay dividends more often during the year, but at lesser nominal amounts, will tend to have lower nominal earnings per share than utilities which pay dividends less often, but at slightly higher nominal amounts. However, because both the utility and investor have reinvestment opportunities, the yield to the investor will be the same regardless of the dividend payment policy, assuming equal payout ratios. Therefore, Southern's argument fails.

In conclusion, the Commission believes that the model (1), proposed in the Notice, provides the best approximation of the cost of common equity for purposes of this proceeding.

50/ The "ratemaking rate of return" issue, discussed elsewhere in this order, raises the question of whether these "earnings on earnings" should be accounted for in the revenue requirement analysis.

2. Sample

The Commission proposed to use a sample of 99 electric utilities 51/ based on the standards adopted in its first two annual proceedings for three reasons. First, the sample is representative of the electric utility industry as a whole. Second, the relevant price and dividend data are generally available for all of these companies. Finally, the data is readily accessible from more than one source. The sample would consist essentially of those publicly traded electric utilities or combination companies that meet explicit standards:

(1) the utility is predominantly electric; 52/

51/ See Order No. 420, 50 Fed. Reg. at 21,831. As a result of a recent merger between Cleveland Electric Illuminating Company and Toledo Edison Company to form Centerior Energy Corporation, the number of companies in the sample has been reduced to 99 from the 100 company sample previously used.

52/ Operationally, the Commission has selected all companies classified in the industry groupings "Electric Service" or "Electric and Other Services Combined" by Standard and Poor's Compustat Services, Inc. These industry groupings are supposed to conform as nearly as possible to the Office of Management and Budget Standard Industry Classification Codes. The Compustat "Electric Services" industry grouping (Industry Classification Number 4911) is defined as establishments engaged in the generation, transmission and distribution of electric energy for sale where these services constitute 90% or more of revenues. The industry grouping "Electric and Other Services Combined" (Industry Classification Number 4931) is defined as establishments primarily engaged in providing electric services in combination with other services, with electric services as the major part, though less than 90% of revenues. (Standard and Poor's Compustat Services, Inc., Utility Compustat II User Manual (1985)).

(2) the utility has its stock traded on either the New York or American Stock Exchange;

(3) the utility is included in the Utility Compustat II data base; and

(4) the utility is not excluded by the Commission on a case-by-case basis, based on unique circumstances. 53/

The fourth standard would give the Commission the discretion to eliminate companies for which data may be unavailable or inappropriate.

The Commission also proposed to continue using the following screening criteria in each quarterly calculation to ensure that the data for each company is available and that it can reasonably be employed in a mechanical fashion without

53/ In Order No. 442, three companies which meet the first three standards were eliminated from the sample. Southwestern Public Service Company was eliminated because it uses a non-standard fiscal year. This causes its dividend yield to be out of time with the rest of the companies. CP National was deleted because, in spite of its being listed as a predominantly electric company, only 17.7 percent of its revenues in 1985 were derived from electric sales. Finally, UNITIL was eliminated because it is a new utility and insufficent data are available. 51 Fed. Reg. at 351. The same considerations eliminate these three companies from our base year calculation. Moreover, because of the merger of Toledo Edison Company and Cleveland Electric Illuminating Company to form Centerior Energy Corporation, the sample for the base year dividend yield calculation had decreased from 100 companies to 99 companies. However, adequate data on UNITIL is now available which permits us to include this company in the sample used for purposes of the quarterly indexing procedure. With the addition of UNITIL, the sample for the quarterly indexing procedure is increased to 100 companies in this proceeding.

producing distorted statistics. Companies would be dropped from the sample for the following reasons:

(i) the company's common stock, through merger or other action, no longer is publicly traded;

(ii) the company has decreased or omitted a common dividend payment in the current or prior three quarters; or

(iii) the Commission determines on a case-by-case basis that some other occurrence causes the dividend yield for that company to be substantially misleading and bias the resulting quarterly average.

The first screen would ensure data availablity. If a company is no longer publicly traded, it will not have a current market price (and yield). The second screen would eliminate companies for which data would probably be inappropriate in a constant growth DCF model. The third screen would give the Commission the discretion to further eliminate atypical companies when necessary.

a. Comment Summary

Four commenters support the proposed sampling procedure. 54/ PEPCO uses five screens, which it states are in accord with the proposed criteria, to reduce the sample to 65 companies. 55/ Cooperatives state that the sample used to determine the average dividend yield should be the same as that used to determine

54/ Southern IC at 12-13; MINN IC at 6; NEP IC at 88; PA Staff IC at 12.

55/ PEPCO IC at A-4.

the growth rate and that only companies which are reviewed quarterly by both Salomon Brothers and Value Line should be included. 56/ EEI states that under the proposed screening criteria only the riskiest companies were excluded from the sample; thus, the average is understated. 57/

b. Analysis and Conclusion

The Cooperatives' arguments as to the same sample being used for the dividend yield and the growth estimate have been raised in previous proceedings, and we reject them for the same reasons that we gave in these previous proceedings. 58/

Concerning EEI's arguments as to the exclusion of only the riskiest companies, given our decision to continue primary reliance on a DCF model, it would be inappropriate to use data for firms that are not currently paying dividends. 59/ EEI is correct, however, that the second screening criteria tends to eliminate only the riskiest companies from the sample. Nevertheless, the two other screens adopted by the Commission provide an appropriate balance. The Commission recently used the first screen when two companies on the sample merged. In addition, the third screen provides sufficient flexibility to eliminate

56/ Cooperatives IC at 93-99.

57/ EEI IC at B-1 to B-2; EEI RC at B-9.

58/ See Order No. 442, 51 Fed. Reg. at 351-52.

59/ See Order No. 442, 51 Fed. Reg. at 352.

individual companies that might bias the quarterly average. Moreover, the reliance on a median dividend yield further mitigates the effect of extremes.

PEPCO uses two sampling criteria in addition to those proposed by the Commission in the Notice. The two additional criteria proposed by PEPCO are: (1) whether a company paid out more than 100 percent of earnings in any year since 1980; (2) whether a company is viewed as a speculative investment by Value Line. 60/ The Commission finds that it would not be appropriate for purposes of a forward-looking DCF analysis to exclude utilities on the basis of events as far back as 1980. PEPCO's additional criteria will therefore not be adopted.

The Commission determines that for its base year estimation it will use the 99-company sample, subject to the screening criteria listed above. As noted supra, as of the first quarter of 1987, UNITIL meets the criteria for inclusion in the sample used for quarterly indexing. This sample is, therefore, 100 companies. A list of the companies included in the base year sample appears in Appendix B; those included in the quarterly indexing sample appear in Appendix C.

60/ PEPCO IC at 4-5.

3. Dividend Yield

a. Introduction

In the Notice, the Commission proposed to continue the dividend yield policy adopted in Order Nos. 420 and 442. This policy is to use the median dividend yields of the 100 company sample. 61/ The Commission stated that the distribution of dividend yields (and, by inference, the distribution of the cost of common equity) for electric utilities is skewed rather than symmetrical. 62/ Under these circumstances, the dividend yields for a greater number of utilities are closer to the median than the mean. The Commission also stated its belief that, compared to the mean, the median is less likely to be affected by extreme values in the data.

In computing the dividend yield the Commission proposed to continue its current policy: (1) the dividend rate used would be the "indicated dividend rate," which is the last declared quarterly dividend times four; and (2) the price used would be the simple average of the three monthly high and low prices for the quarter. The computation of the base year dividend yield would use the average of four quarterly median yields. The computation of the dividend yield used in the quarterly indexing procedure would use the average of two quarterly median yields.

61/ See Order No. 420, 50 Fed. Reg. at 21,812; Order No. 442, 51 Fed. Reg. at 352-53.

62/ Id.

b. Comment Summary

Few commenters directly address the method proposed by the Commission to compute the dividend yield. Most commenters use the method proposed without discussion.

EEI proposes that the Commission change its method of calculating the dividend yield. 63/ EEI asserts that using the midpoint of the mean and median dividend yields would further reduce the influence of extremes.

Southern computes dividend yields by weighting each utility's market price and quarterly dividend by its number of common shares outstanding. 64/

Cooperatives suggest that a six-month yield be used in the calculation of the base year dividend yield. 65/ Cooperatives argue that the DCF model is a long-run expectations model. They state that it is often necessary to smooth out short-term dividend yields to reduce the effects of variations which do not reflect investors' long-run expectations, but the yields should not be over-smoothed. Cooperatives recommend using a six-month dividend yield for the base year because the six-month yield is consistent with the yield used in the quarterly indexing procedure, and because there have been dramatic changes in interest rates since the end of 1985.

63/ EEI IC at B-2.

64/ Southern IC at 14-15.

65/ Cooperatives IC at 100-107; Cooperatives RC at 20-21.

c. Analysis and Findings

The midpoint of the mean and the median, proposed by EEI, in computing quarterly dividend yields is an "ad hoc" statistic unsupported by statistical theory and rarely, if ever, used in statistical analysis. The mean is not resistant to the influence of extreme data while the median is resistant. 66/ Thus, the midpoint of the mean and median would be less resistant to the influence of extreme data than the median. Therefore, the Commission will continue to rely on the median.

Southern's weighting scheme for dividend yields gives proportionately more weight to the larger companies. The Commission addressed the issue of weighted average dividend yields in Order No. 420. 67/ It found that such weighting conflicts with its objective of establishing a rate of return that is representative of most utilities.

One of the reasons cited by Cooperatives for proposing a six month dividend yield for the base period computation is consistency with the yield used in the quarterly indexing procedure.

66/ See, e.g., Frederick Mosteller and John W. Tukey, Data Analysis and Regression, Addison-Wesley, Reading, MA; 1977.

67/ Order No. 420, 50 Fed. Reg. at 21,814.

The other reason cited by Cooperatives is that there have been dramatic changes in interest rates since the end of 1985. In these generic proceedings, the Commission adopts a "base year" as the time period over which commenters are requested to make estimates of the cost of common equity. The intention is to have all parties focus on and estimate this cost for the same time period. The Commission could have adopted a "base 6 months". It did not. Use of a dividend yield over only half of the base year will not provide a reasonable estimate of the cost of common equity for the entire year since the cost during the first half of the year may be different than that for the second half. Cooperatives' proposal is rejected.

The Commission finds no reason to change its policy concerning the use of the median dividend yield or the time period for computing the dividend yield for the base year. The base year dividend yield using the average of the four quarterly median yields for the year ended June 1986 is 8.25 percent. 68/

68/ The median dividend yields for the third and fourth quarter of 1985 and for the first and second quarter of 1986 are 9.13, 8.92, 7.79, and 7.16 percent, respectively. See Appendix B for a listing of the companies included in the sample and the companies excluded in each of the four base year quarters.

4. Growth Rate

a. Introduction

The Commission proposed to rely on both a fundamental analysis and a two-stage growth analysis to estimate the constant growth rate in this proceeding. With a fundamental analysis, the underlying components of dividend growth -- retained earnings and new stock sales -- are evaluated separately. 69/ Similarly, in the two-stage growth analysis, near-term and long-term growth expectations are evaluated separately. Both historical and forecast data are relied on for these analyses. The Commission also proposed to consider other data and methods to estimate expected growth as a check on the reasonableness of the above analyses.

b. Comment Summary

Ten commenters make growth rate recommendations. The range of these growth rate recommendations is from no more than 3.66 percent (WCG) to as much as 6.25 percent (PEPCO). See Table 2, below. The lowest recommendation by a utility or utility group is 4.01 percent (Southern). The highest recommendation by a utility customer or customer group is 4.76 percent (Cooperatives). Three of the 10 commenters (AUS, Second Cooperatives, and FA Staff) recommend a growth rate about 4.60 percent.

69/ The first component of this growth -- growth from retained earnings -- is a function of the expected rate of return of common equity (r) and the expected retention rate (b). The second component -- growth from new common stock sales -- is a function of the amount of new stock sales (s) and the price of the new stock sales relative to book value (v). The latter factor (v) is often referred to as the equity accretion factor.

Table 2. Summary of Growth Rate Recommendations

Commenter	Growth Rate	Basis for Recommendation
Utilities		
1. PEPCO	6-6.25%	1. Historical DPS growth rates 2. Corroborated with analyst forecasts
2. NEP	5-5.5%	1. Historical DPS growth rates 2. Base year fundamental analysis
3. EEI	4.8-5.5%	1. Hist. EPS and DPS growth rates 2. Base year fundamental analysis 3. Projected fundamental analysis 4. Analyst forecasts 5. Two-stage growth analysis
4. AUS	4.61%	1. Hist. EPS and DPS growth rates 2. Analyst forecasts 3. Base year fundamental analysis
5. Southern	4.01%	1. Historical EPS growth rates 2. Base year fundamental analysis 3. Analyst forecasts
Customers		
6. Cooperatives	4.76%	1. Projected fundamental analysis 2. Corroborated with: a. Historical EPS, DPS, and BVPS growth rates b. Analyst forecasts c. Two-stage growth analysis
7. Second Cooperatives	4.60%	1. Projected fundamental analysis
8. WCG	2.6-3.66%	1. Multistage DCF growth analysis
9. GSA	3.50%	1. Historical growth in dividends and price appreciation 2. Analyst forecasts
Other Commenters		
10. FA Staff	4.60%	1. Projected fundamental analysis

DPS = Dividends Per Share
EPS = Earnings Per Share
BVPS = Book Value Per Share

Most commenters support their recommendations with more than one approach. Growth rate recommendations above 5 percent are based primarily on extrapolation of past growth in dividends and earnings. The bottom end of the growth rate recommendations is based on a variety of data and methods.

The highest recommended growth rate is recommended by PEPCO. For a 65 company sample, PEPCO evaluates dividend per share (DPS) growth rates for all 3, 5, 7, and 10 year periods ending in years 1980 through 1985. 70/ PEPCO places most weight on the industry's most recent 3 and 5 year experienced growth rates and chooses a range of expected growth rates of 6 - 6.5 percent. This commenter compares these rates to projections by Value Line, Merrill Lynch, and Salomon Brothers and finds them to be higher but consistent. Using data for the end of the base year, PEPCO's study shows that the medians for these projections range from 5.4 to 6.0 percent for the 65 company sample. Based on these statistics and its judgement, PEPCO recommends a growth rate in the range of 6.00-6.25 percent. 71/

70/ While PEPCO presents the data for all of these periods, it focuses on the periods ending in years 1983 through 1985. The range of rates so determined is from 5.11 to 6.45 percent. Generally, PEPCO's data shows that the rates are higher the shorter and the more current the time period.

71/ PEPCO IC at 10-12, Schedules 5-7. In its comments, PEPCO criticizes the Commission for relying on a fundamental (br + sv) analysis. The criticisms, however, all relate to a type of fundamental analysis that relies exclusively on actual data or extrapolations of historical data. The Commission has clearly emphasized the use of forecast data as well as the evaluation of recent trends in its fundamental analyses. The Commission has explained that the intention is to estimate the expected long term growth rate. Further, the Commission has never specifically proposed using historical data. Other commenters criticize the fundamental approach for other reasons. See e.g., VEPCO IC at 2; MINN IC at 7. The Commission has addressed such criticisms in previous proceedings. 50 Fed. Reg. at 21,818-20; 51 Fed. Reg. at 355-7.

NEP recommends a growth rate in the range of 5.00 to 5.50 percent. This range is based on two types of data. First, NEP looks at median 5 and 10 year historical growth rates in dividends per share for its 89 company sample -- 5.56 and 5.46 percent, respectively. Second, NEP estimates the median retention growth at the end of the base year -- 4.59 percent. NEP states it was not necessary to adjust the retention growth rate for accretion from new stock sales because market prices exceeded book values by a small percentage. 72/

NEP also criticizes the use of a two-stage model to evaluate growth rates for this proceeding. According to this commenter, the two-stage model is only useful when two different growth rates are expected and the timing is predictable. NEP believes that accurate predictions in such detail are unlikely for a large group of companies such as the industry as a whole and investors are unlikely to make such refinements in their expectations.

EEI looks at a greater variety of information in its evaluation of the growth rate and recommends a rate between 4.80 and 5.50 percent. For its analyses, EEI relies on a sample of 91 companies. First, EEI looks at 5 and 10 year historical growth rates in earnings and dividends per share. The range of means and medians for this measures 5.31 to 7.84 percent. Second, EEI uses a fundamental analysis of retention and new common stock growth and estimates the actual growth during the base year as 5.1 percent and the forecasted growth as 5.2

72/ NEP IC at 6-8, Schedules E-6 and E-7.

percent. 73/ Third, the forecasts of analysts from Value Line, Merrill Lynch, Salomon Brothers, and I/B/E/S (Institutional Brokers Estimation System) -- whose means and medians ranged from 4.0 to 5.3 percent -- are taken into consideration. Fourth, EEI uses a two-stage growth DCF model to determine a composite average rate. For the first and second stages of growth, EEI uses the 5.1 percent estimated actual base year growth rate and the average of the analyst forecasts of 4.8 percent, respectively. The result of this study is a composite growth rate of 5.00 percent. 74/

73/ EEI determines fundamental growth during the base year by calculating 12 month moving retention ratios (b) and rates of return on common equity (r) for its sample companies for each quarter during the base year. Over the four quarters, the means of the retention growth rates (b x r) for its sample of companies range from 4.24 to 4.71 percent and the medians range from 4.50 to 4.90. EEI places greater weight on the median values and chooses 4.80 as the representative growth rate.

To this value, EEI adds a base year growth rate from new stock sales of .31 percent. This is based on an median market-to-book ratio of 1.22 times, which implies an equity accretion factor (v) of .18 (or 1 - 1/1.22). The base year new stock sales growth rate (s) is estimated as 1.7 percent based on subtracting the 4.8 retention growth rate (b x r) from the estimated aggregate 1985 common equity growth rate (G) of 6.5 percent. Thus, EEI estimates base year new stock sales growth of .31 (s x v = 1.7 x .18).

For its determination of projected fundamental growth, EEI relies on the 3-5 year Value Line industry composite average projection of a 5.0 percent retention growth (adjusted from end of year measures to average year measures). Projected new stock sales growth (s) is estimated as .75 based on its own projections from another study (1.6 to 1.7 percent) and projections from Value Line (according to EEI, .5 percent). The equity accretion factor (v) is based on the implied industry average market-to-book ratio projection by Value Line of 1.19 times. Thus, v equals .16 (or 1 - 1/1.19) and sv = .1 (or .75 x .16).

74/ EEI IC at Attachment B, B-6 to B-27, Appendices 4-18.

AUS also uses a variety of methods and data. AUS states that historical and analyst forecast growth measures indicate a range of plausible growth rates between 4.00 and 6.50 percent. Median 5 and 10 year historical growth rates in dividends and earnings per share for its sample of companies during the base year ranges from 4.38 to 6.50. Median values for analyst forecasts from Value Line, Merrill Lynch, and I/B/E/S range from 4.00 to 4.80 percent. The mean and median of the whole set of 10 historical and forecast estimates are 4.77 and 4.57 percent, respectively. AUS uses these mean and median values in a two stage growth DCF analysis to produce a composite growth rate of 4.63 percent. AUS also determines that actual fundamental growth from retention during the base year is 4.59 percent for the sample. 75/ AUS's recommended growth rate of 4.61 percent is the simple average of these latter two rates. 76/

Southern is the final utility commenter to make a growth rate recommendation. Southern recommends 4.01 percent based on the simple average of three growth rate measures: (1) a 5-year historical earnings per share (EPS) growth rate of 4.06 percent, (2) a sustainable (retention) growth rate of 3.65 percent, and

75/ AUS determines the base year fundamental growth rate by multiplying the median retention ratio (.319) by the median earnings-to-book ratio (14.39 percent) for its sample of companies. AUS makes no attempt to estimate growth attributable to new stock sales. In reply comments, AUS recommends that the Commission reject the sv component in principle. AUS states that when the market-to-book ratio is less than unity, the sv adjustment is negative, which lowers the growth rate. AUS argues that this lowering of the growth rate lowers the market cost of capital estimate at a time when the marketplace indicates a need for greater growth rates. Thus, AUS believes the Commission should reject the sv component. AUS RC at 10.

76/ AUS IC at 25-37, Schedules 3-5.

(3) an I/B/E/S average of analysts forecasts of 4.33 percent. Each of these growth rates are calculated on a company-by-company basis and then weighted by each company's total assets. Southern argues that this weighted average is a more representative industry growth rate than the simple average growth rate. 77/

Cooperatives recommend a growth rate of 4.76 percent based primarily on a kind of fundamental analysis (referred to as intrinsic") using data for the second half of the base year. 78/

77/ Southern IC at 15-17.

78/ Cooperatives estimate the mean and median return on common equity (r) for their sample as 14.74 and 14.87 percent, respectively. These values are based on Value Line 3-5 year forecasts during the latter half of the base year. The comparable mean and median retention rates (b) based on Value Line forecasts of earnings and dividends per share are .298 and .296, respectively. Using these mean and median values produces retention growth rates of 4.39 percent (.298 x 14.74) and 4.40 (.296 x 14.87). The mean and median values for the retention growth rates derived from the distribution of individual company calculations are 4.45 and 4.34 percent, respectively. Cooperatives recommend a retention growth rate of 4.40 percent since it is the mid-point of these four different estimates.

To this 4.40 percent retention growth rate, Cooperatives recommend adding a new stock financing growth rate of .36 percent. This value is based on a new stock growth rate (s) of 1.30 percent and an accretion factor (v) of .280. The mean and median market-to-book ratios for its sample of companies are 1.38 and 1.39 times for the latter half of the base year. From these values, an equity accretion rate (v) of .28 is estimated (1 - 1/1.38 or 1 - 1/1.39). The growth rate in new common stock (s) is determined by subtracting the retention growth rate of 4.40 percent, above, from the Value Line forecasts of annual growth in total common equity (G). The mean and median values of these G values are 5.75 and 5.68 percent, respectively. Using these values produces new stock growth rates of 1.35 and 1.28 percent. The mean and median of the distribution of the same calculations made for the individual companies in the sample are 1.34 and 0.84 percent, respectively. Based on these four estimates, Cooperatives recommend using a new stock growth rate of 1.30 percent and, thus, a new stock financing growth rate of .36 percent. This rate, added to the retention growth rate of 4.40, results in a 4.76 overall rate.

For their 84 company sample, Cooperatives look at historical data, analyst forecasts and a two-stage growth analysis for corroboration. Cooperatives' estimates of average 5 and 10 year historical growth rates in earnings, dividends, and book value per share range from 2.71 to 7.43 percent. The means and medians of analyst forecasts from Salomon Brothers, Value Line, Merrill Lynch, and Zack's for the end of the base year range from 4.0 to 5.31 percent. Finally, Cooperatives state that the growth rate to use in a constant growth DCF model should be somewhere between their estimated range of near-term growth rates (4.5 to 5.25 percent) and their range of steady-state growth rates (3.8 to 4.2 percent). These latter ranges are based on the averages of the analyst short-term and long-term forecasts for selected company groupings. 79/

GSA recommends a growth rate of 3.5 percent. This rate is based on (1) average concomitant growth in dividends and stock prices over 20 and 30 year holding periods of about 3 percent for Moody's Utility Index and between 3 and 4 percent for unregulated companies and (2) an average Merrill Lynch Steady State earnings per share growth forecast of 4 percent for utility companies during the base year. 80/

79/ Cooperatives IC at 109-135, Schedules 12-15.

80/ GSA IC at 8-11, Exhibits II-V.

In Reply Comments, commenters criticize the GSA analysis. 81/ NEP argues that GSA misunderstands the theory underlying discounted cash flow analyses when GSA argues that that price growth must be considered along with dividend growth in growth rate analyses. NEP states that the DCF model, as commonly used, incorporates price growth because future prices are assumed to be determined by dividend growth beyond the date of those prices. NEP also argues that the GSA data is biased by its choice of beginning and ending dates and that if 1985 were included in the analysis different conclusions would result. EEI also states that GSA provides little data in the record to support its conclusions. While stating that its recommendation is based on 20 year holding periods, GSA shows statistics relating only to 5 and 10 year holding periods. Further, SCE criticizes GSA for basing its recommendation only on long-term forecasts of growth while advocating the use of multi-stage DCF models.

In Reply Comments, a second group of cooperatives (Second Cooperatives) recommends a growth rate of 4.6 percent. This recommendation comes from a fundamental analysis based on an evaluation of the data and analyses of the Initial Comments of other commenters. 82/

81/ AUS RC at 28-31; EEI RC at 28-31; NEP RC at 3-5; SCE RC at 6-10.

82/ Second Cooperatives RC at 19-24. Second Cooperatives project a retention ratio (b) of .30, a return on common equity (r) of 14.5 percent, a new stock sales growth rate

(FOOTNOTE CONTINUED ON NEXT PAGE)

Also in Reply Comments, the WCG customer group states that a growth rate in excess of between 2.6 to 3.66 percent cannot be justified. WCG bases this statement on single and multiple stage growth DCF analyses where the growth rates incorporate declines in expected rates of return from the current 14.7 percent rate. 83/

Finally, FA Staff relies solely on a forecasted fundamental analysis to support its recommended growth rate of 4.60 percent. 84/

(FOOTNOTE CONTINUED FROM PREVIOUS PAGE)

(s) of 1.3 percent, and an equity accretion rate (v) of .194. The (s) term is based on the Value Line projected common equity growth rate (G) of 5.6 percent less the retention growth rate of 4.35 percent (or .30 x 14.5). The equity accretion rate of .194 is based on a market-to-book ratio of 1.24 times, which Second Cooperatives state is the industry median for the year ending June 30, 1986.

83/ WCG RC at 9-16, Appendices A-D.

84/ FA Staff IC at 2-10, 14-22, and Attachments B, D and E. The components of FA Staff's fundamental analysis -- b, r, s, and v -- are .30, 14.25, 1.3, and .194, respectively. The retention ratio (b) of .30 is based on (1) mean and median payout ratios for July 1986 of .69 and .68, respectively, (2) average payout ratios for the 1981-1985 period of .747 to .721, and (3) Value Line projections during the first half of 1986 generally between .696 and .708. FA Staff's projected rate of return on common equity (r) is based on its review and judgement of (1) Value Line 3-5 year projections generally between 14.7 and 14.9 percent, (2) a Duff and Phelps prediction of a decline in earned rates of return, (3) an attrition analysis finding that earned rates may be .6 percentage points below the current average allowed rate of 15.2 percent, and (4) a "sustainable rate of return" analysis which produces a rate of 14.83 percent. The new common stock growth rate (s) of 1.3 percent is based on a Value Line forecasted common equity growth rate (G) of 5.6 percent less the above-determined retention growth rate of 4.3 percent (or .30 x 14.25). Finally, FA Staff estimates the average market-to-book ratio during the base year is 1.24 times and projects an equity accretion factor (v) of .194 (or 1 - 1/1.24).

c. Analysis and Findings

In the previous generic rate of return proceeding (Docket No. RM85-19), the Commission estimated the expected growth rate during the year ending June 30, 1985 to be 4.5 percent. 85/ On review of the record in this proceeding, the Commission finds a 4.6 percent rate estimate appropriate for the year ending June 30, 1986.

In evaluating the growth rate question, the Commission follows the same general approach it used in the previous generic proceeding. In fact, given that few of the underlying facts have changed over the course of the last year, much of that analysis still applies. The Commission reviews and evaluates the recommendations of the commenters. Table 2, above, summarizes these recommendations and describes the bases for them. The Commission also reviews and evaluates the data underlying commenters' recommendations for use in a fundamental analysis and a two-stage growth analysis. Table 3 categorizes this raw data for comparison purposes. The Commission considers both historical and forecast data relevant and useful for these analyses. As the Commission stated in the last proceeding, "all relevant data should be used and any apparent inconsistencies explained to the extent possible." 86/

85/ 51 Fed. Reg. at 355.

86/ Id.

Table 3. Raw Growth Rate Data

Rate(s)	Type of Rate	Commenter
	Historical DPS Growth Rates	
5.25	5-year median	AUS
5.37/5.52	5-year mean/median	EEI
5.56	5-year median	NEP
5.64	5-year mean	Cooperatives
4.63	10-year median	AUS
5.58/5.33	10-year mean/median	EEI
5.46	10-year median	NEP
5.30	10-year mean	Cooperatives
5-11-6.45	range of medians for selected time periods - see text	PEPCO
	Historical EPS Growth Rates	
6.50	5-year median	AUS
7.84/7.18	5-year mean/median	EEI
4.06	5-year average	Southern
7.43	5-year mean	Cooperatives
4.38	10-year median	AUS
5.38/5.31	10-year mean/median	EEI
5.15	10-year mean	Cooperatives
	Base Year Fundamental Growth Rates	
	(b)(r) + (s)(v)	
4.59	(.319)(14.39); no sv term	AUS
5.1	4.80 + (1.71)(.18)	EEI
4.59	br only; no other data given	NEP
3.65	br only; no other data given	Southern
	Projected Fundamental Growth Rates	
	(b)(r) + (s)(v)	
5.2	5.10 + (.75)(.16)	EEI
4.76	(.297)(14.8) + (1.3)(.28)	Cooperatives
4.60	(.30)(14.5) + (1.3)(.194)	Second Cooperatives
4.60	(.30)(14.25) + (1.3)(.194)	FA Staff

- Continued -

Table 3. Basic Growth Rate Data (continued)

Rate(s)	Type of Rate	Commenter
	Analyst Near Term Forecasts	
4.45	I/B/E/S median	AUS
4.3/4.0	I/B/E/S mean/median	EEI
4.33	I/B/E/S average	Southern
4.73/4.67	Zack's mean/median	Cooperatives
4.50	Value Line DPS median	AUS
4.8/4.9	Value Line DPS mean/median	EEI
6.0	Value Line DPS median	PEPCO
4.94/5.00	Value Line DPS mean/median	Cooperatives
4.38	Value Line EPS median	AUS
4.2/4.4	Value Line EPS mean/median	EEI
4.80	Merrill Lynch DPS median	AUS
4.9/4.8	Merrill Lynch DPS mean/median	EEI
5.4	Merrill Lynch DPS median	PEPCO
4.71/4.90	Merrill Lynch DPS mean/median	Cooperatives
4.80	Merrill Lynch EPS median	AUS
4.7/4.8	Merrill Lynch EPS mean/median	EEI
5.3/5.3	Salomon Brothers' Normalized Growth mean/median	EEI
5.5	Salomon Brothers' Normalized Growth median	PEPCO
5.31/5.00	Salomon Brothers' Normalized Growth mean/median	Cooperatives
	Analyst Long Term Forecasts	
4.00	Merrill Lynch Steady State EPS median	AUS
4.00/4.00	Merrill Lynch Steady State EPS mean/median	Cooperatives
4.00	Merrill Lynch Steady State EPS average	GSA

DPS = Dividends per share
EPS = Earnings per share

The Commission also reiterates the following:

> The determination of the growth rate involves substantial judgment on the Commission's part. While the Commission's perspective is different from that of a security analyst or a prospective stock buyer, it has the same data available to it. It must infer from that data the expectation of investors on the future prospects of companies implied by current market prices. Thus, the Commission's analysis is no more precise than any other judgmental exercise. The Commission's analysis therefore determines a range for the growth rate based on the best available data and within the context of each analytical approach used. The Commission must then decide on a specific rate within that range. 87/

At the high end of the range of growth rate recommendations, the Commission finds those of PEPCO and NEP excessive. Those recommendations are based primarily on past trends in dividends per share. However, as in the last proceeding, analyst forecast data, in which the Commission places greater credence, suggests significantly lower growth in both the near term and long term. Obviously, simple extrapolation of past trends is not adequate in current times.

PEPCO corroborates its recommendation with analyst forecast data suggesting a range of 5.4 to 6.0 percent. However, these data are substantially different than the comparable data of other commenters. This suggests that the smaller samples of companies used by PEPCO for its analyses is unrepresentative of the industry as a whole.

87/ Id.

NEP presents a base year retention growth rate "br" calculation of 4.59 percent in addition to its historical dividend per share growth rates. This does not support a 5 to 5.5 percent growth rate recommendation.

At the other end of the recommendations are those of GSA and WCG. Notwithstanding the criticisms that could be made as to the specifics (or lack thereof) of the approaches and data these commenters used to support their recommendations, 88/ the Commission believes that the preponderance of evidence in this record supports a growth rate in excess of 3.7 percent.

In general, the Commission finds no appreciable changes in the various measures of the growth rate between the last proceeding and the current one. The 4.6 percent rate adopted here is within the range of 4.3 to 4.7 percent found reasonable in the last proceeding, the same range which the Commission believes is reasonable for this proceeding. The Commission's

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88/ WCG presents a statistical study which is purported to be a two stage growth analysis. Based on the scant information provided, it looks more like it incorporates annual changes in the growth rate during the first 5 years. Further, the model gives counterintuitive results. For example, WCG's study begins with a retention growth rate (b x r) of 4.41 percent (or .30 x 14.7) which declines over 5 year to 3.66 percent (or .30 x 12.2) but, according to WCG, averages 2.6 percent. WCG's study also shows negative growth in dividends and earnings over the first 5 years but a 5-year book value per share growth rate of 4.1 percent and a 5-year "b x r" growth rate of 4.0 percent. Because of the inadequate support for the study and the questions that arise from reviewing it the Commission places little probative value on its results. With respect to the GSA analysis, the Commission generally concurs with the criticisms made by commenters.

judgment as to which rate to adopt within this range is influenced by the fact that three of the 10 commenters that made growth rate proposals recommended 4.6 percent. The fact that three commenters -- AUS, Second Cooperatives and FA Staff -- represent different interests in this proceeding lends credence to the reasonableness of the Commission's determination.

In the Commission's fundamental analysis for this proceeding, the expected growth from earnings retention may have fallen slightly, but this reduction appears offset by an increase in the expected growth from new stock sales. With regard to the Commission's two-stage growth analysis, none of the underlying data used to support the analysis of the last proceeding has changed in any measurable degree. In the last proceeding, historical 5 and 10 year DPS growth rate estimates of commenters ranged from 4.5 to 5.6 percent, with the majority of estimates in the range of 5.2 to 5.6 percent. In this proceeding, the estimates range generally from 4.6 to 5.6 percent with the majority in the 5.2 to 5.6 percent range. Just as the data in the last proceeding indicated a wide range of historical 5 and 10 year EPS growth rate estimates (5.3 to 8.9 percent), the data in this proceeding suggests a wide range (4.1 to 7.8 percent). In the last proceeding near-term analyst forecasts ranged from 4.2 to 4.9 percent. In this proceeding, the range is generally from 4.2 to 5.0 percent. Finally, the

average long-term analyst forecast (Merrill Lynch's Steady State EPS growth rate) is 4.0 percent in both proceedings. 89/

The Commission's fundamental analysis suggests a long-term expected growth rate of 4.7 percent. The components of this analysis are addressed in turn below.

For the retention ratio (b), the Commission sees no reason to depart from the range of .28 to .32 that it adopted in the last proceeding and the .30 value as its best estimate. No party has explicitly recommended a value outside this range. FA Staff shows average retention ratios for its 86 company sample of .279 to .317 during the period 1983 to 1985. 90/ The mean and median retention ratio forecasts by Value Line based on individual company data are .292 and .30. 91/

AUS, which determined an actual base year retention ratio of .319, argues that a Value Line projection of higher common equity ratios supports the notion of the retention ratio rising from its current level. 92/ However, FA Staff presents contrary Value Line data suggesting that, at the end of the base year, Value Line reduced its projected retention ratio from .304 to

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89/ Cf. 51 Fed. Reg. at 358; Table 3, above.

90/ FA Staff IC at 14 and 15. These values are calculated from FA Staff's reported payout ratios since the retention ratio is equal to one minus the payout ratio.

91/ FA Staff IC at 15. (Calculated from the reported payout ratios per previous footnote)

92/ AUS IC at 36; AUS RC at 8.

.240. 93/ In the Commission's judgement, a .30 retention ratio is a reasonable estimate of the investors' long-term expectations.

The Commission adopts a range of 14.5 to 14.8 percent as its best estimate of the average expected long-term rate of return on common equity. Since the use of any value within this range combined with a .30 retention ratio produces a 4.4 percent retention growth rate (rounded to the nearest tenth of a percent), it is unnecessary to choose any best estimate within that range. 94/ These values are supported by recent average earned rates of return and by Value Line near-term forecasts of 14.7 to 14.9 percent. 95/ It is also supported by the recommendations of Cooperatives and Second Cooperatives, which are based to a large extent on the Value Line forecasts.

93/ FA Staff IC at 15.

94/ The Commission believes that the conversion of rates of return on average common equity -- the predominant type of return data referred to -- to rates of return on beginning-of-year common equity may still be warranted for the DCF model adopted by the Commission. See Order No. 442, 51 Fed. Reg. at 357. However, since the magnitude of the adjustment -- about 40 basis points for the range of returns considered reasonable -- has only a minor impact on the retention growth rate once a retention ratio is adopted, this issue is subsumed in the Commission's consideration of the appropriate expected long-term rate of return on common equity.

95/ FA Staff IC at 16 and 17. FA Staff's data shows average earned rates of return of 14.5 to 14.9 during the last few years. Also, AUS presents a base year earned rate of return of 14.39 percent. AUC IC at 36.

FA Staff projects an expected return of 14.25 percent but the Commission believes that this rate is not adequately explained or supported by FA Staff's data. 96/ Notwithstanding this criticism of FA Staff's low estimate of the expected rate of return -- which, if adopted, would only lower the estimated retention growth rate by about 10 basis points -- the Commission is sensitive to the effect of lower interest rates and, with a lag, lower allowed rates of return. The record in this proceeding supports the notion that the high allowed rates of return of recent years are not expected to continue indefinitely. The Commission believes that the dramatic fall in interest rates over recent years, reflected in the fall in the Commission's quarterly estimates of the cost of and benchmark rates of return on common equity, are likely to be reflected in lower allowed rates of return and, eventually, in lower earned rates of return. 97/ These trends lend some credence to average

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96/ FA Staff seems to have been unduly influenced by the change in Value Line's projected industry average rate of return from 14.9 percent (per issue dated 6/6/86) to 13.9 (per issue dated 6/27/86). The appropriate basis for the expected return are the average expectations over the whole of the base year not the expectations during the last few days of the year. The Value Line rate of return projections for the base year prior to the last week of June suggest average expected rates above 14.5 for the year. The Commission believes that FA Staff's conclusions from its attrition and sustainable rate of return analyses are also unduly influenced by projected declines in earned rates of return since all of the data indicates rates of return above 14.25 percent.

97/ Duff and Phelps reports that the average allowed return for 29 rate orders issued in the first half of 1986 was 14.5 percent. FA Staff IC at 18.

expected long-term rates of return below 14.5 percent. But the Commission sees little evidence that significant declines in rates of return are actually projected to any great extent in investor expectations during the base year.

With regard to the average expected long-term rate of new stock sales (s), the Commission adopts a rate of 1.3 percent. This rate was proposed by three commenters and was within the range of values used by a fourth. 98/ It is also consistent with a projected growth rate in aggregate common equity of 5.5 to 5.8 percent less the above-determined retention growth rate of 4.4 percent. 99/

The last component of the fundamental analysis is the average expected long-term equity accretion rate (v), which is based solely on the average expected market-to-book ratio. 100/ The range of recommended values for "v" is .160 to .280 based on a projected range in market-to-book ratios of 1.19 to 1.39 times. 101/ The Commission adopts a value of "v" of .160 (rounded

98/ Cooperatives IC at 119-120; FA Staff IC at 19-21; EEI IC at B-18 to B-22; Second Cooperatives RC at 20-21.

99/ EEI IC at Appendix 12; Cooperatives IC at 120; FA Staff IC at 21; Second Cooperatives RC at 20.

100/ The term "v" is defined as one minus the reciprocal of the market-to-book ratio.

101/ EEI IC at B-21; Cooperatives IC at 119; FA Staff IC at 21; Second Cooperatives RC at 21.

to .2) near the bottom end of the range since this estimate is based on projections from Value Line. 102/ The other recommendations are based on actual market-to-book ratios during the base year.

Putting the above components together produces an estimate of the average expected long term growth rate of 4.7 percent -- 4.4 percent from retention growth (.3 times 14.5 to 14.8 percent) and .3 percent from sale of new common shares (1.3 times .2).

In evaluating the two-stage growth analysis, the Commission reiterates that there has been little or no change in the measures for the year ending June 1985 to the year ending June 1986. 103/ See Table 3, above. 104/ As a result, the Commission's analysis in Order No. 442 which used a first stage (5 years) growth rate of 4.8 percent and a second stage growth

102/ EEI IC at 13-21. See also 51 Fed. Reg. at 357.

103/ The Commission finds little merit to NEP's criticism of the use of two-stage models. The issue is not one of accuracy, as NEP puts it, but rather one of laying open to the greatest extent possible the implications of the growth rate recommendations. The constant growth rate is an assumption that simplifies the analysis. As such it is a composite of different growth rates into the future. Just as it is a worthwhile endeavor to break down analysts' assumptions as to the factors underlying their expectations of fundamental growth, it is also helpful to break down the assumptions as to near term and long term growth. There clearly must be greater confidence placed in near term growth forecasts but that does not mean that they must be assumed to apply to the long term as well. The object of the exercise is to make explicit that which is implicit in growth rate analyses so that a more reasoned evaluation is possible.

104/ Cf. 51 Fed. Reg. at 358.

rate of 4.0 percent remains appropriate. The Commission determines a best estimate of 4.3 percent from its two stage growth analysis. 105/

In summary, the Commission finds the same range of growth rates as in the last proceeding -- 4.3 to 4.7 percent. The latter is based primarily on a fundamental analysis, the former on a two-stage growth analysis. Within this range, the Commission adopts a growth rate of 4.6 percent.

105/ In light of its understanding that the Merrill Lynch Steady State EPS growth rates are generally projected to apply to periods beginning 10-15 years in the future, the Commission analyzed the effect of lengthening the first stage in its analysis. Assuming that the 4.8 percent first stage rate applied, on average, for 10 years instead of only 5 years, the composite average growth rate applicable to a constant growth DCF model is estimated as 4.45 percent. Alternatively, assuming a 4.8 percent average rate for the first 15 years with the 4.0 percent growth rate beyond produces a composite average growth rate of 4.6 percent. The Commission believes this further supports its finding of a 4.6 percent growth rate in this proceeding.

Cooperatives present a similar two-stage growth analysis for comparison with their constant growth analysis. Cooperatives IC at 140-146. These results are generally consistent with the Commission's.

5. Corroborative Evidence

a. Introduction

In the Notice, the Commission requested that commenters support their market required rate of return estimates with corroborative evidence. The Commission did not specify any particular types of corroborative evidence. Commenters were requested to provide comprehensive explanations of alternative models they propose along with their assumptions.

Few commenters offered corroborative evidence. Two commenters submitted risk premium analyses and two others submitted earnings-price and earnings-book ratio analyses. These studies are summarized below along with other evidence that the Commission believes corroborates its findings.

b. Comment Summary and Analysis

i. Risk Premium Analyses

Two commenters present risk premium analyses for corroboration.

AUS offers a collection of seven different risk premium studies. For each of these studies, the commenter states that it adjusts the resulting premium to place each on the same basis, a "basis compatible with A-rated public utility bonds." The range of adjusted risk premiums is from 2.6 to 6.1 percent. 106/

106/ AUS IC at 43-49, Schedule 7.

NEP submits two risk premium studies, which it refers to as "interest premium" studies. In these studies, average risk premiums over 1976-1985 were estimated for each of 89 electric utilities based on constant growth DCF cost estimates and the yields to maturity for specific bonds of the individual companies. This commenter used two different constant growth DCF models based on different methods for estimating the growth rates. The studies produced average risk premiums of 2.15 and 2.53 percent. NEP adds these premiums to its estimate of the average yield on the individual utility bonds for the base year ended June 30, 1986 of 10.50 percent. The resulting range of investor return requirements is from 12.74 to 13.08 percent. 107/

WCG raises two criticisms about risk premium analyses generally. First, WCG states that the accuracy of the risk premiums are dependent on the accuracy of the cost of common equity estimates used to derive the risk premiums. Second, WCG claims that the premiums are not constant over time and that recent studies have shown that long term debt may at times be more risky than common equity. 108/

In the last generic proceeding, the Commission reviewed risk premium analyses similar to some of those submitted by

107/ NEP IC at 9-11, Schedules E-8 through E-10.

108/ WCG RC at 17-20.

AUS and NEP. Generally, the Commission questioned the stability of risk premiums for recent years and the historical relationship between debt and equity securities. The Commission concluded that it "is reluctant to place any great weight on risk premium analyses in general other than those based on a simple ranking of securities." 109/

The Commission is concerned with the validity of the specific risk premiums found in the AUS and NEP studies. Some of the same criticisms identified in the last proceeding apply to the studies prepared by AUS and NEP. In addition to those criticisms, the Commission believes that the risk premiums must be consistent with some DCF analysis. 110/ As a result, the AUS studies that imply growth rates significantly above the level supported by the Commission's analysis are suspect. The Commission remains concerned with the applicability of historical risk premiums. Therefore, the Commission sees no merit in pursuing the technicalities of the various

109/ 51 Fed. Reg. at 359-60.

110/ For example, applying AUS's range of risk premiums on A-rated bonds produces required rate of return estimates from 13.1 to 16.6 percent. Given an average dividend yield of 8.28 percent, the growth rates implied by returns in this range are from 4.7 to about 8.2 percent. As the Commission finds in Section III.B.4, above, there is no reasonable evidence supporting long run industry average growth rates much above 5 percent. The average yield on Moody's new A-rated public utility bonds is 10.52 percent for the year ending June 30, 1986. See Table 4, below. Adding the range of risk premiums noted earlier -- 2.6 to 6.1 percent -- to this yield produces a range of required returns from 13.1 to 16.6 percent.

studies submitted in this proceeding.

While the Commission has concerns with the quantification of specific risk premiums, it continues to believe in the ranking of securities based on relative risk. The higher the risk associated with a security, the higher is the investors' return requirement. Table 4, below, presents selected interest rates and risk premiums for a wide range of securities. 111/ A review of these rates in comparison to the industry average required rates of return shows rates that the Commission believes are consistent over time and consistent across securities based on risk differences. The Commission believes these statistics corroborate its finding in this proceeding.

ii. Earnings-Price (E/P) and Earnings-Book (E/B) Ratios

FA Staff and Cooperatives submitted E/P ratio analyses identical to those they submitted in the last proceeding. 112/ Both commenters provide estimates of the industry average E/P ratio for comparison with their DCF-derived estimates. These analyses are based on the notion that when the price-book (P/B, or market-to-book) ratio is greater than one, the E/P ratio understates the market cost.

111/ See Order No. 442 for a discussion of nominal and effective interest rates and for methods of converting the rates to make them consistent. Since the Commission is here mostly concerned with the relative ranking of the securities, it does not make the conversions in this order. 51 Fed. Reg. at 360-61.

112/ FA Staff IC at 10-11, 22-24; Cooperatives IC at 147-152. See also Order No. 442, 51 Fed. Reg. at 361-62.

Table 4. Selected Interest Rates and Risk Premiums

Security	Year Ending 6/30/84	6/30/85	6/30/86
Selected Interest Rates*			
Treasury Bills (New 3 month).....	9.24%	8.76%	6.82%
Commercial Paper (New 3 month).....	9.65	9.17	7.41
Treasury Bonds			
10 Year Constant Maturity.....	12.11	11.75	9.06
20 Year Constant Maturity.....	12.25	11.89	9.39
Moody's Public Utility A-Rated			
Preferred Stock.....	12.82	12.45	10.05
Moody's Public Utility Bonds			
Aaa.....	12.84	12.47	10.08
Aa.....	13.44	13.10	10.45
A.....	13.80	13.53	10.81
Baa.....	14.41	13.96	11.31
Composite Average.....	13.79	13.27	10.66
Yields on Recently Issued Bonds			
Moody's New A-Rated.....	12.98	12.37	10.52
Composite Average.....	13.52	13.11	10.52
Average Market Required Rate of Return on Common Equity for Electric Utilities:			
Nominal Rate (using 420 Model)...	15.25	14.73	13.02
Effective Rate (using 442 Model).	15.90	15.32	13.45
Selected Risk Premiums**	-- (rounded values) --		
Treasury Bills (New 3 Month).....	6.0	6.0	6.2
Treasury Bonds (10 Year Constant).....	3.1	3.0	4.0
Moody's Public Utility Preferred.....	2.4	2.3	3.0
Moody's New A-Rated Bonds.....	2.3	2.4	2.5

* Rates are average of monthly rates for specified periods.
** Risk premiums are determined by subtracting the average yield for the specified security from Nominal Rate determined from the 420 Model.
Sources: Federal Reserve Statistical Release G.13 (various dates)
Moody's 1986 Public Utility Manual
FERC Order Nos. 420 and 442

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FA Staff also compares its estimate of the expected E/B ratio (or rate of return on common equity) with its DCF-derived estimate of the cost. This analysis is based on the theory that when the P/B ratio is greater than one, the E/B ratio overstates the market cost.

AUS criticizes FA Staff's E/P analysis primarily for the same reasons reported in the last proceeding. 113/

In Order No. 442, the Commission extensively reviewed the comparable E/P studies. Generally, the Commission agreed with the criticisms made by AUS. 114/ Those shortcomings in FA Staff's and Cooperatives' implementation of this corroborative test remain.

The record in this proceeding does not contain an estimate of the actual average E/P ratio for the base year ending June 30, 1986. As a result, the Commission will not evaluate the consistency of E/P ratios with the final determined base year's cost of common equity estimate.

The E/B ratio test was also reviewed by the Commission in the last proceeding. 115/ With an average base year P/B ratio about 1.24 times, the E/B ratio should overstate the

113/ AUS RC at 11-13. See also Order No. 442, 51 Fed. Reg. at 362.

114/ 51 Fed. Reg. at 363.

115/ Id.

market required rate of return estimate. Investors expect to earn a greater return on the book value of their investment than on their market value. The Commission estimates the long-term expected rate of return on book to be 14.3-14.8 percent. These values exceed the final required rate of return estimate of 13.03 percent. Thus, the E/B ratio test corroborates the Commission's finding in this proceeding.

6. Flotation Costs

a. Introduction

In the Notice, the Commission proposed to use the flotation cost policy adopted in Orders Nos. 420 and 442:

(1) utilities would be compensated only for issuance costs, such as underwriters' compensation and legal and printing fees;

(2) this cost would be reflected in an industry average adjustment to the market required rate of return; and

(3) adjustments for flotation costs would be made through the following formula which reflects recovery of the average annual cost incurred:

$$k^* = \frac{fs}{(1 + s)}$$

where:

k^* = flotation cost adjustment to required rate of return

f = industry average flotation cost as a percentage of offering price

s = proportion of new common stock expected to be issued annually to total common equity

The Commission asked commenters to submit estimates of the parameters for the above formula.

b. Comment Summary and Analysis

In the previous proceeding commenters addressed three primary issues. These same issues are raised in this proceeding. One, whether the Commission should make any allowance for costs other than issuance costs, such as costs due to "market pressure" or "market break." Two, whether the recovery of flotation costs should be reflected in the allowed rate of return on common equity or through some other method. Three, whether flotation costs should be recovered through a form of current cost recovery or a form of perpetual amortization.

i. Type of Costs to be Recovered

All commenters who address recovery of issuance costs argue for some form of recovery. 116/

A number of commenters state that market pressure 117/ occurs and that public utilities should be compensated for these costs. 118/ Cooperatives state that market pressure does not exist. 119/ One commenter, NSP, requests the Commission to perform its own study of market pressure costs.

116/ AUS IC at 37-40; CGE IC at 3-4; DE IC at 5-7; EEI IC at 8-30; NEP IC at 11; NSP IC at 4; PEPCO IC at A-13; Southern IC at 17-19; UPL IC at 3-4; VEPCO IC at 2-3; FA Staff IC at 11-12; Cooperatives IC at 92; GSA IC at 13; AWW IC at 33; Second Cooperatives RC at 27-28.

117/ "Market pressure" cost is the alleged decline in the price of a stock at the time of the news of a new issue of stock.

118/ AUS IC at 38; EEI IC at 8-30; NEP IC at 11; NSP IC at 4-5; PEPCO IC at A-13; Southern IC at 18; AWW IC at 33-34.

119/ Cooperatives RC at 33-34.

In past proceedings the Commission reviewed a number of market pressure studies and found that they did not demonstrate the existence of market pressure costs. No new market pressure studies are submitted in this proceeding. The Commission finds insufficient evidence of market pressure to initiate its own study or to change its policy regarding market pressure costs.

Two commenters raise the issue of market break. 120/ NEP states that market break costs exist and that they should be recovered. AUS is the only commenter to submit evidence on market break. 121/ AUS claims that the "short-term" market variability of the Dow-Jones Utility Average for the five years ending in 1985 was 3 percent and for the year ending June 30, 1986 was 3.1 percent. 122/

In Order No. 420 the Commission found that the theoretical argument made by WCG of an equal likelihood of a market break "cost" and a market break "profit" was reasonable. 123/ The Commission finds that the evidence provided by AUS shows only that utility stock prices vary, not that there is a market

120/ AUS IC at 40; NEP IC at 11.

121/ "Market break" cost is the alleged effect of the reduced price received by a utility when it sells stock during a period of short-term market decline.

122/ AUS defines "short-term market variability" as the ratio of the low price for a given month to the high price of the prior two months.

123/ Order No. 420, 50 Fed. Reg. at 21,824.

break cost. The Commission finds no evidence in the record to support a change in its policy on market break costs.

ii. Method of Recovery

Five commenters propose case-by-case methods of flotation cost recovery. 124/ DE suggests that current issuance costs and amortized amounts of past unrecovered issuance costs be recovered as cost-of-service items. Cooperatives and Second Cooperatives propose that only current issuance costs be recovered as cost-of-service items.

VEPCO proposes that flotation costs be recovered through a rate base adjustment but did not provide an example of how the adjustment would be applied. 125/

GSA opposes a generic approach to flotation cost adjustment, preferring instead that costs be recovered using the Commission's adjustment formula on a case-by-case approach. 126/

The Commission addressed the issue of company specific flotation cost adjustments in Order No. 420. 127/ The Commission continues to believe that an industry average adjustment to

124/ DE IC at 6-7; VEPCO IC at 2-3; Cooperatives IC at 92; GSA IC at 13; Second Cooperatives RC at 27-28.

125/ VEPCO IC at 3.

126/ GSA IC at 11-14.

127/ See 50 Fed. Reg. at 21,826.

34

the market required rate is the best way of dealing with flotation costs: (1) they have a relatively small quantitative impact, (2) any adjustments are subject to forecasting errors, and (3) overrecovery and underrecovery of these costs by individual utilities should be offset over time.

iii. Form of Recovery

In this proceeding, as well as in the previous proceedings, most commenters who argue for the perpetual amortization method also argue that the resulting flotation cost adjustment be applied to all equity. 128/

In Order No. 442, the same arguments were dealt with in detail. 129/ Basically, there are two methods of recovering flotation costs, amortization and current cost recovery. The Commission explained that for new companies the perpetual amortization method and current cost recovery methods lead to the same recovery of costs. Once the perpetual amortization method is adopted, it must be continued and entails recovery each year on all outstanding stock. Similarly, once the current cost recovery is adopted, it also must be continued. This method recovers costs as they occur and overrecovery would result if costs of past issues were recovered each year.

128/ AUS IC at 40; CGE IC at 3-4; EEI IC at 32-34; NEP IC at 11; NSP IC at 4; PEPCO IC at A-13; Southern IC at 18-19; UPL IC at 3-4.

129/ 51 Fed. Reg. at 364-365.

Once either method is adopted it should be followed. In Order No. 442 the Commission chose to continue using a form of current cost recovery stating:

> When justified, the Commission has allowed flotation costs in the past. However, it is not clear whether past recovery has been the amount that would be permitted by either the current method or the amortization method. With the generic proceedings, the Commission wishes to start with a clean slate. Thus, the Commission adopted a policy of current cost recovery in Order No. 420 and will continue this policy in the current proceeding. 130/

The record in this proceeding does not support a change in this policy.

c. Flotation Cost Adjustment

As explained in Order No. 420, the following formula determines an increment to the cost of common equity which reflects, on average, the annualized amount of flotation cost incurred by the industry: 131/

$$k^* = \frac{fs}{(1 + s)}$$

where:

k^* = flotation cost adjustment to required rate of return

f = industry average flotation cost as a percentage of offering price

s = proportion of new common stock expected to be issued annually to total common equity

130/ Order No. 442, 51 Fed. Reg. at 365.

131/ Order No. 420, 50 Fed. Reg. at 21,826.

The range of estimates for "f", issuance costs as a percent of gross sales price, are in a narrow range from 2.05 to 2.59 percent. 132/ The differences are due to the company samples used in commenters' analyses.

The Commission finds the analyses of EEI and FA staff, which include the same twelve new issues, to be the most complete and adopts their estimate of 2.4 percent.

The expected proportion of new common equity issued annually, "s," was found in the growth rate section (III.B.4, above) to be 1.3 percent. Applying the 2.4 percent estimate of issuance costs, f, and the 1.3 percent estimate of new equity financing, s, to the above formula, the Commission finds a flotation cost adjustment of 3 basis points. 133/

132/ AUS IC at 38; EEI IC at B-30; PEPCO IC at 13; UPL IC at 2 and FA Staff IC at 24. PEPCO performed a study of these costs between 1970 and 1986 and found the average yearly cost to be 4.1 percent. It notes that the costs declined in recent years due to changes in industry financing practices, more intensive competition among underwriters and higher per share stock prices. For the year ending June 30, 1986, PEPCO found a median cost of 2.5 percent.

133/ Flotation Cost Adjustment $= \frac{0.024(0.013)}{1.013}$

$= 0.0003$

7. Jurisdictional Risk

Concerning the question of whether there is a difference in risk between the wholesale and retail operations of electric utilities, the Commission proposed to adopt the finding of Order No. 442 that there is no appreciable difference in risk due to this factor. 134/

a. Comment Summary

AUS and NEP support the proposed finding that there is no difference in juridsictional risk. 135/ APPA argues that there are differences in jurisdictional risk due to differences in rates of return allowed by different regulatory commissions and differences in risks of providing different kinds of service. 136/ BEC claims that wholesale service is riskier than retail service because there is more risk due to uncertainty in the level of the customer's load since the customer can serve its load from other sources or from its own generation. 137/ WCG argues that the Commission has recognized that its policies with regard to Construction Work in Progress (CWIP) in rate base result in a transfer of risk from investors to ratepayers.

134/ See 51 Fed. Reg. at 366.

135/ AUS IC at 60; NEP IC at 19.

136/ APPA IC at 15-17.

137/ BEC RC at 12-13.

30

WCG argues that to the extent that different jurisdictions have different policies regarding CWIP in rate base, there are differences in jurisdictional risk. 138/

b. Analysis and Conclusion

All of the arguments made by the commenters as to the existence of a difference in overall risk between utility operations subject to this Commission's jurisdiction and the nonjurisdictional operations of utilities have been raised in prior proceedings. No commenter has pointed out any change in circumstances which would change the basis for the Commission's prior finding that there is no significant difference in overall risk between operations of utilities under the Commission's jurisdiction and nonjurisdictional operations. The Commission therefore continues to find that there is no significant difference in jurisdictional risk. 139/

138/ WCG IC at 6.

139/ See Order No. 442, 51 Fed. Reg. at 366.

C. Quarterly Indexing Procedure

1. Introduction

In the Notice, the Commission proposed the quarterly indexing procedure established in Order No. 442 as modified on rehearing by Order No. 442-A. 140/ In that indexing procedure, quarterly changes in the cost of common equity are tied to changes in utility dividend yields. The average cost of common equity is indexed to the average of the median dividend yields for the two most recent calendar quarters for the company sample. The benchmark rate of return on common equity is set equal to the cost of common equity except where the quarter-to-quarter changes exceed 50 basis points. Thus, the quarter-to-quarter changes in the benchmark rates of return are capped at 50 basis points. The intent of the cap was to smooth out fluctuations in the benchmark rates of return and, by implication, allowed rates of return, over time. The initial benchmark rate established in each annual proceeding is not subject to the 50 basis point cap.

The Commission requested comments on any changes that would improve the proposed indexing procedure.

2. Comment Summary And Analysis

While some commenters supported the current indexing procedure, 141/ other commenters suggested four kinds of changes: (1) the use

140/ See 51 Fed. Reg. at 366; 51 Fed. Reg. at 22,509.

141/ NEP IC at 19; EEI RC at 5; NEP RC at 19; EEI RC at 5; NEP RC at 8; SCE RC at 10.

of a period different than two quarters to calculate the dividend yield applied in the indexing, (2) elimination of the 50 basis point cap, (3) the use of the cap as a "trigger" mechanism which, when exceeded, would cause the indexing procedure to be suspended, and (4) an adjustment for changes in growth expectations in addition to adjustment for changes in the dividend yield.

a. The Six-Month Dividend Yield and the 50 Basis Point Cap

Two commenters support the use of a six-month dividend yield for indexing but reject the 50 basis point cap. 142/ Three commenters support the proposed use of both a six-month yield and a 50 basis point cap. 143/ MINN recommends the use of the average dividend yield over the latest 20 trading days and WVCAD proposes the most recent month's average of daily closing prices. 144/ VEPCO proposes a three-month period. 145/ The shorter time periods are proposed primarily to make the benchmark rates more current. WVCAD, in proposing a time period shorter than two quarters, argues that the two-

142/ FA Staff IC 25; Cooperatives IC at 153.

143/ NEP IC at 19; EEI RC at 5; NEP RC at 8; SCE RC at 10. EEI states that it "does not object to updating the benchmark with the industry average dividend yield using the two most recent quarters of market data provided that the use of the 50-basis point cap be continued."

144/ MINN IC at 7; WVCAD IC at 10. In times of "market ambivalence", WVCAD suggests it might propose a longer period, up to the most recent quarter.

145/ VEPCO IC at 3.

quarter option (1) violates fundamental financial principles underlying the DCF method, (2) offers no more rate stability than the "cap" already provides, and (3) results in less accuracy in the rates of return.

Three commenters state that a period longer than two quarters should be used to calculate the dividend yield employed in the indexing procedure. 146/ These commenters propose using a twelve-month period for the yield for the following reasons: (1) it reduces the mismatching of the time frame used for the dividend yield calculation and the growth rate calculation, (2) it is more consistent with the use of a twelve-month dividend yield in the annual proceedings, (3) it provides a more stable benchmark, and (4) it minimizes the frequency of the application of the cap.

Four commenters express the view that the 50 basis point cap should be eliminated. 147/ CGE argues that the cap prohibits adjustments to the benchmark rate of return to reflect current market conditions. Moreover, CGE argues that since the cap is not applied to the annual proceeding, it should be eliminated from the quarterly adjustment. PEPCO wants to abandon the cap but extend the period for calculating the dividend yield from six months to twelve months. FA

146/ AUS IC at 57-59; NSP IC at 5; PEPCO IC at A-15.

147/ CGE IC at 4; PEPCO IC at A-15; and FA Staff IC at 25. Second Cooperatives RC at 31.

Staff argues that the use of a cap is contrary to one of the goals of the generic rate of return, namely, 'more accurate' rate of return decisions.

Other commenters state that the cap is beneficial. 148/ Cooperatives, however, support the cap only as a "trigger" mechanism which, when exceeded, would cause the indexing procedure to be suspended and a new benchmark rate of return to be established. They argue that "[w]hen large movements in stock prices and dividend yields occur, there is a good reason to suspect that the growth rate in the constant growth model may have also significantly changed." 149/

As the comments demonstrate, the 50 basis point cap has distorted the benchmark rates of return by limiting adjustments that would reflect current market conditions, and current capital costs. During 1986, the cap was applied to the second, third, and fourth quarterly benchmark rates of return under Docket No. RM85-19-000 to the point where it now exceeds the estimated market cost of common equity by .87 percentage points. Although the estimated cost of common equity fell 2.32 percentage points during 1986, the cap limited the

148/ AUS IC at 57-59; NEP IC at 19; NSP IC at 5; WVCAD IC at 5; REI PC at 5; SCE RC at 10; Southern RC at 4. MINN believes that the 50 basis point cap should be used as a limit on upward adjustment but removed as a limit on downward adjustments. MINN IC at 8.

149/ Cooperatives IC at 155; see also Southern RC at 4-5; Second Cooperatives RC at 31.

reduction in the benchmark to 1.50 percentage points. Similar situations may occur in the future during changing cost conditions.

The Commission believes that reconsideration of the cap is warranted. A presumption behind the use of the cap was that it would not come into play very often. The benchmark was not intended to diverge from the cost of common equity significantly or for very long periods.

The use of a six-month dividend yield should provide the measure of stability that led to the 50 basis point cap, which was originally coupled with a three-month dividend yield. As evidence from three comments, the decision as to the length of the time period over which the dividend yield should be computed is a matter of judgement. The arguments presented in this proceeding are not substantially different from those presented in the previous proceedings and which the Commission has considered and evaluated. 150/ The Commission thus reaffirms the use of a six-month dividend yield in the quarterly indexing procedure.

150/ See 51 Fed. Reg. at 357.

b. Changes in Growth Expectations

Some commenters express the view that the quarterly indexing procedure should reflect changes in investors' growth rate expectations. 151/ These commenters point out that, under the proposed procedure, the base year growth rate estimate is used in estimating the cost of common equity for periods up to one and half years after the base period ends. They state that there is an inverse relationship between dividend yield and expected growth. According to AUS, the changes over time in the expected growth rate are so significant that the use of the base year's estimated growth rate, together with a more current dividend yield, results in a "mismatch" that could lead to substantial errors in the estimated cost of common equity. This mismatch is said to prevent the updated return from reflecting current capital market conditions. 152/

It should be noted that, in Order No. 420, the Commission found that the long-run constant growth rate for the base year ending June 1984 was 4.3 percent. In Order 442, for the base year ending June 1985, the long-run constant growth rate was found to be 4.5 percent. In this proceeding, the base

151/ AUS IC at 57-59; NSP IC at 5-6; VEPCO IC at 3; DE IC at 4.

152/ AUS urges the Commission to adopt a method of time-matched (i.e., synchronized) dividend yields and growth rates each with a quarterly cap of 50 basis points variation. Moreover, AUS suggests that the indexing be done quarterly using a twelve-month moving average. AUS IC at 57-59.

year constant growth rate is found to be 4.6 percent. These small differences between the growth rates for the three base years are consistent with the view that the industry's expected growth rate changes slowly.

The Commission continues to believe that investors' growth rate expectations are relatively stable over the length of time at issue. In addition, the specific proposals of commenters to incorporate changes in growth rate expectations into an indexing procedure are inadequate.

No new arguments are presented which would cause the Commission to modify its position not to subject the growth rate to the quarterly indexing procedure. 153/

153/ 51 Fed. Reg. at 367-68. AUS claims that the Commission has misunderstood its concern about the mismatching of the dividend yield and the growth rate. On the contrary, the Commission has expressed its problems with determining a procedure for updating the growth rate component in the DCF model on a quarterly basis. The Commission has also indicated that, in its judgement, the expected growth rate changes very slowly in comparison with the dividend yield and that it is not unreasonable to assume the base year growth rate is a good estimate for the growth rate applicable to the following year. The Commission believes its judgement has been confirmed by the relatively modest changes in the growth rates it has adopted in the first three annual proceedings.

D. Ratemaking Rate of Return

1. Introduction

The Notice sought comments with regard to a concept that the Commission has previously identified as the "ratemaking rate of return." In Order No. 442 the Commission described the ratemaking rate as the rate of return which, when applied to the particular rate base determined by the regulatory agency, allows the electric utility to provide the investors with their effective required return. 154/ The distinction between the investors' expected return from a utility common stock and the ratemaking rate of return is based upon the recognition that the investors' rate of return from an investment in a utility's common stock may be treated as having three different components. The first component is the payment received as return on funds provided by investors for the utility's rate base.

The second component of the investors' expected return is the investors' return from the investors' reinvestment of dividend payments made by the utility during a given year. In the Notice, the Commission proposed to use the Order No. 420 version of the DCF model, which would not include this second component in the benchmark rate. In Order No. 442-A, the Commission recognized

154/ See 51 Fed. Reg. at 349.

that it was unnecessary to include this component of return because "[b]y paying dividends quarterly, the firm makes it possible for the investor to reinvest the dividends during the year"; thus "the firm does not have to pay out the income received from this reinvestment of dividends since investors produce this income by their own actions." 155/

The third component of the ratemaking rate of return concept is that a utility has an opportunity to reinvest intra-year retained earnings, which reduces the rate of return ratepayers must pay to allow the utility an opportunity to pay out the amount required. The Notice cites the Commission's Staff Report which states that "in a fashion analogous to the investors' opportunity for intra-year reinvestment of dividends, the firm can increase income through the intrayear reinvestment of its earnings." 156/ The Notice concludes that "if the ratepayers paid in at the estimated 'payout rate' [of return] the firm would have the opportunity to earn more than it is required to pay out." 157/

The Notice sought comments on three questions in connection with the third component of the expected rate of return: (1) does a utility have an opportunity to earn a higher rate of return

155/ 51 Fed. Reg. at 27,052.

156/ Id.

157/ Id.

than the Commission allows through the utility's ability to reinvest its intrayear retained earnings through an inconsistency in the way rate base is defined or estimated for cost of service purposes or through some other mechanism; (2) if the utility does in fact have such an opportunity, what should the Commission do about it; and (3) if it is determined that the allowed rate of return should be adjusted, how should this adjustment be accomplished? 158/ The Commission commented that if "the concept involves the firm's intrayear reinvestment of earnings, this determination entails the empirical questions of how often a company compounds its earnings and at what rate." 159/

In Order No. 442, the Commission initially adjusted the investors' effective rate of return to take account of the imputed return component from intrayear reinvestment of retained earnings. The adjustment was referred to as the ratemaking rate of return adjustment. 160/ On rehearing, however, the Commission ultimately

158/ A Staff Report by the Commission's Office of Regulatory Analysis suggested that not only is the firm's "pay out" rate of return "less than the investors' effective required [rate of] return [because of the opportunity for the investor to receive income from the reinvestment of dividends], but also that the rate [of return] which ratepayers have to pay in is less than the firm's required 'pay out' rate [of return]." 51 Fed. Reg. at 27,053.

159/ Id.

160/ See 51 Fed. Reg. at 350.

decided that "there are a number of unresolved questions with regard to some of the stated purposes of the ratemaking rate of return." 161/ The Commission determined then to adopt instead the model that it had previously adopted in Order No. 420 since that model excluded from the allowed rate of return "the income that . . . [investors] . . . expect to receive from the reinvestment of dividends." 162/ At the same time, the Order No. 420 model did not involve attempting to exclude from the allowed rate of return the return associated with reinvestiment of retained earnings.

2. Comment Summary

In response to the Notice, most of the commenters opposed implementation of the ratemaking rate of return concept beyond the use of the Order No. 420 model. The following are among the major criticisms: (1) the application of the concept requires the assumption that retained earnings are invested periodically in rate base; 163/ (2) the concept is incompatible with the DCF method because the DCF method assumes that the market price of a company's stock already reflects the investors' awareness of the

161/ 51 Fed. Reg. at 22,500.

162/ 51 Fed. Reg. at 22,508.

163/ AUS IC at 53; EEI IC at 25.

fact that reinvestment of dividends is occurring or may occur; 164/ and (3) the concept involves the unrealistic assumptions that the cash that accrues temporarily prior to being paid out as dividends is reinvested at the allowed rate of return, and the income from such investments is tax-free. 165/ Because the Commission has determined not to implement the ratemaking rate of return concept beyond the Order No. 420 model at this time, the criticisms of the concept will not be dealt with in detail.

Some commenters do urge the Commission to proceed with the full implementation of the concept as set forth in the Staff Report. FLA supports the concept generally but does not address the issues of implementation discussed in the Notice. 166/ Although GSA agrees with the concept, it would support adjusting the cash working capital allowance rather than the allowed return. 167/ MINN takes a similar position. 168/ WVCAD discusses why it supports the concept behind adjusting the generic rate of return to recognize the firm's ability to

164/ REC IC at 9; EEI IC at 24.

165/ NEP IC at 16.

166/ FLA IC at 1-9.

167/ GSA IC at 14.

168/ MINN IC at 9.

reinvest intrayear retained earnings, but does not address the issues regarding implementation set forth in the NOPR. 169/ WCG asserts that the ratemaking rate of return should be adopted, but does not address the question of how the concept should be implemented other than simply to assert that the Commission should use a daily compounding model because a utility receives a portion of its earnings every day. 170/

3. Conclusion

The comments on the ratemaking rate of return support our conclusion that a number of issues regarding the implementation of ratemaking rate of return remain. These include such empirical issues as the assumptions to be made concerning the tax treatment of returns from reinvested retained earnings, how frequently such earnings are to be compounded and at what rate, and whether other aspects of the utility's cost-of-service are affected. Commenters that favorably address this issue have failed to provide adequate evidence for resolution of these issues. These are formidable issues which we believe are not susceptible to resolution at this time. The Commission has therefore determined not to apply the ratemaking rate of return concept to adjust the generic rate of return.

169/ WVCAD IC at 32.

170/ WCG IC at 10.

43

E. The DCF Method/Cost of Capital Standard

1. Introduction

In the Notice, the Commission proposed to adopt the same DCF model as proposed and ultimately adopted in the first two annual generic rate of return proceedings. 171/ It also requested comments on whether there are reasons for the Commission to depart from placing primary reliance on the DCF method. 172/

2. Comment Summary

Four commenters question the Commission's use of the DCF method for purposes of determining allowed rates of return. Both APPA and AWW argue that a rate of return based solely on the cost of capital, which is what the DCF method attempts to estimate, is inadequate and unsupportable. 173/ APPA argues that the cost of capital is but one factor the Commission may consider in establishing a fair rate of return. That factor may be outweighed by other factors, and in some cases need not be considered. 174/ AWW contends that the allowed return on common equity "must enable the utility to: (1) attract capital on reasonable terms, and (2) realize a return on book equity comparable to other enterprises." 175/

171/ 51 Fed. Reg. at 27,050.

172/ Id. at 27,051.

173/ APPA IC at 3; AWW IC at 25.

174/ APPA IC at 3.

175/ AWW IC at 26.

APPA also argues, as it has in the past, that it is incorrect to apply a DCF based allowed rate of return to a book value rate base.

> APPA's concern is that there is a fundamental difference between an economic rate of return and an accounting rate of return that precludes their use in the manner proposed by the Commission. 176/

In APPA's view, there has been substantial research since the 1970's which supports its conclusion. 177/

AWW and AUS take a somewhat different tack in raising questions about the use of the DCF method. According to AWW, the "DCF formula is premised on the assumption that the market price of the utility's stock reflects the stock's underlying value." 178/ AWW argues, however, that recent studies show that this assumption is a "myth" and that this "new evidence requires the Commission to reexamine its proposed reliance on the DCF methodology." 179/ The studies cited by AWW call into question the validity of the efficient market hypothesis upon which AWW contends DCF theory is founded. AUS also questions the usefulness of the DCF method by arguing that "[w]hen it can be shown that significant uncertainties face the industry in the future, coupled with an equity market characterized by euphoric investor expectations which cannot be sustained, a DCF calculation of the

176/ APPA IC at 7.

177/ Id. at 9.

178/ AWW IC at 10.

179/ Id.

cost rate of common equity capital should be given less weight than under normal circumstances." 180/ AUS reviews historical levels of price-earnings ratios and concludes that "the market for equities is substantially influenced by undue investor optimism which has resulted in an over-valued stock market." 181/

AWW also contends that the Commission cannot rely solely on DCF evidence because it "provides no information about what comparable firms are earning on their book equity." 182/ As a result, AWW argues that "the Commission needs, at a minimum, comparable earnings data to verify the results of its DCF formula" and proposes a comparable earnings approach that would provide such data. 183/ APPA raises a somewhat related issue by contending that "it appears that the Commission does not feel a need to check the results of its methodology against other economic and financial evidence." 184/ Notwithstanding its criticism of the DCF method, AWW suggests that it could produce a reasonable result if it were modified "to produce a market-to-book ratio equal to that of unregulated, comparable risk companies." 185/

180/ AUS IC at 16.

181/ Id. at 18. SCE also argues that the stock market is not "properly priced." SCE IC at 2.

182/ AWW IC at viii.

183/ Id. at viii, 36-44.

184/ APPA IC at 12-13.

185/ AWW IC at 19.

Finally, APPA interprets "the Commission's proposal to guarantee existing equity holders the current market cost of capital." 186/ Its position is based on the fact that "current equity holders in firms in competitive markets are not guaranteed the current market cost of capital on their investments." 187/ APPA also claims that the Commmission's approach "places primary focus on the wrong group of investors." According to the APPA, the Commission should be concerned with compensating existing stockholders rather than potential stockholders. 188/

3. Analysis and Findings

The Commission believes that there is compelling economic justification for relying on the market cost of capital as the standard for rate of return decisions. 189/ Nonetheless, the Commission is prepared to take into account non-cost factors in setting an allowed rate of return in an individual case if circumstances warrant.

186/ APPA IC at 10.

187/ Id.

188/ Id. at 11.

189/ "Since by definition the cost of capital of a regulated firm represents precisely the expected return that investors could anticipate from other investments while bearing no more and no less risk, and since investors will not provide capital unless the investment is expected to yield its opportunity cost of capital, the correspondence of the definition of the cost of capital with the court's definition of legally required earnings appears clear. Hope refers to both: commensurate earnings and the attraction of capital. These two approaches are harmonized when the allowed rate of return is set equal to the cost of capital". A. L. Kolbe, and J. Read, Jr. with G. Hall, The Cost of Capital: Estimating the Rate of Return for Public Utilities (1984), at 21.

Although comparable earnings data has been offered before as corroborative evidence of the cost of capital, the Commission has found fault with its use in this regard for essentially two reasons. 190/ First, unlike the relationship between risk and market required rates of return, the relationship between risk and accounting rates of return is not clear. In other words, companies with high risk don't necessarily earn high book returns, and vice versa for companies with low risk. In contrast, investors will expect/require a high market rate of return from companies with high risk and a lower market rate of return from lower risk companies. Second, and more fundamentally, the Commission stated:

> Accounting rates of return are not reliable measures of the current cost of capital, since they do not reflect the current market prices that are determined in competitive capital markets. 191/

With respect to APPA's argument that "economic returns and accounting returns are conceptually and numerically different," 192/ the Commission notes that it has never disputed this particular point. What the Commission said in Order No. 420 is that it has not been adequately demonstrated why this fact makes it inappropriate to apply a DCF-based allowed rate of return to a book value rate

190/ 50 Fed. Reg. at 21,823.

191/ Id.

192/ APPA IC at 7.

base. 193/ APPA has not explained why the more recent literature it cites should change the Commission's view of this matter. It appears as if this literature addresses the differences between accounting and economic rates of return and not whether the application of a DCF-based allowed rate of return to a book value rate base in a regulatory environment is inappropriate. In any event, it is clear that mere citations to the literature are not enough to make one's case. As a result, the Commission finds that APPA has not offered any new evidence that would cause the Commission to change its approach to setting allowed rates of return.

AWW offers some new evidence regarding the efficient market theory and its relationship to the DCF method. The efficient market theory is founded upon the proposition that "all relevant information is widely and cheaply available to investors and that all relevant and ascertainable information is already reflected in security prices." 194/ Although there have been some studies that have pointed to specific inefficiencies that exist in the stock market, the general proposition still seems to reflect mainstream thinking:

> The concept of an efficient market is astonishingly simple and remarkably well supported by the facts. Less than 20 years ago any suggestion that security investment is a fair game was generally regarded as bizarre. Today it is not only widely accepted

193/ 49 Fed. Reg. at 21,829.

194/ See Brealey, R. and Myers S., Principles of Corporate Finance, McGraw Hill (1984) at 266.

> in business schools, but it also permeates investment practice and government policy toward the security markets. 195/

AWW cites some recent literature questioning the appropriateness of relying on a DCF formula, specifically the market price that is a primary input in such a formula. The Commission, however, is unable to conclude from so little evidence that the efficient market theory has been so discredited that one cannot rely on the market prices of electric utility common stocks to reasonably reflect the cash flows expected by investors. It may be that further research in this area will convincingly demonstrate that the evidence cited by AWW substantially undercuts the validity of the efficient market hypothesis or the appropriateness of using a DCF analysis to estimate the cost of capital. In the Commission's judgment, it is premature to make that finding now.

AUS's contention that the stock market is overvalued is fundamentally inconsistent with the efficient markets theory. As evidence that the current market has overvalued electric utilities' common stock, AUS calculates a cost rate for utilities' common equity by dividing the recent 14.39% earned rate of return on common equity by the 151.1% market-to-book ratio. AUS compares the 9.5% result with current bond yields, and concludes that the result is too low. The Commission notes, however, that the 9.5% is

195/ Id. at 281.

merely the industry average earnings-price ratio, which is not necessarily equal to the market cost of capital, especially when price-to-book ratios differ from unity. All that can be concluded from AUS's calculation is that the market cost of capital exceeds 9.5%.

With respect to APPA's concern that the Commission check its results in some way, it appears that APPA is concerned more with having the Commission consider evidence on why the cost of capital is not the appropriate standard to use for rate of return decisions and less with having the Commission look to corroborative evidence on whether the DCF-generated estimate of the cost of capital is reasonable. APPA's point seems to be that the use of a cost of capital standard may produce results that are inconsistent with the prevailing economic environment. For example, during a recessionary period when unregulated companies are experiencing earning declines, a cost of a capital standard may support and perhaps increase utility earnings.

While this may be true, it works both ways. During boom times, when unregulated companies are experiencing significant increases in earnings, a cost of capital standard will limit utility earnings. The fact of the matter is that unregulated companies may earn less than their cost of capital during bad times and more than their

cost of capital during good times. This is why the Commission must also reject the modified DCF formula offered by AWW. By trying to force an equivalence between the market-book ratios of unregulated companies and those of public utilities, AWW would have the Commission depart from a cost of capital standard. Moreover, its ad hoc adjustment is based on neither financial theory nor empirical research.

As to APPA's concern that the Commission's proposal guarantees existing equity holders the current market cost of capital, the Commission thinks otherwise. Setting an allowed rate of return equal to the current cost of capital does not guarantee that the rate of return allowed will in fact be earned. The earned rate of return may be above or below that which is allowed and will depend on numerous factors. Among these are whether the Commission finds reason to adjust the estimated cost-of-service for purposes of determining a just and reasonable rate and whether actual sales and costs turn out to be above or below those used in establishing this rate. In short, utility ratemaking does not guarantee that the allowed rate of return will be earned, regardless of what it is based on.

The Commission also disagrees with APPA that the use of a cost of capital standard focuses inappropriately on potential stockholders rather than existing stockholders. In fact, it is somewhat difficult to reconcile this concern with APPA's other concern described above. In any event, the Commission believes that the consistent use of a

cost of capital standard over time is fair and equitable to both existing and potential stockholders. Not only has APPA not demonstrated otherwise, but it has not offered a superior substitute.

III. SUMMARY OF CHANGES IN REGULATORY TEXT

This rule makes certain changes in the text of the Commission's regulations that deal with the generic rate of return. These changes reflect two decisions by the Commission in this proceeding. The first decision is to continue the advisory status of the generic rate of return for another year. The language of § 37.8 of the regulations is therefore being changed to refer to the first three proceedings rather than the first two proceedings. The second decision is to remove the 50 basis point cap that has previously been part of the quarterly update procedure. The language of § 37.9(a)(3) is therefore being changed to eliminate the references to the cap. Section 37.9(a)(4) is being eliminated because it is the 50 basis point cap provision.

IV. REGULATORY FLEXIBILITY ACT CERTIFICATION

The Regulatory Flexibility Act 196/ requires the Commission to describe the impact that a proposed rule would have on small entities or to certify that the rule will not have a significant economic impact on a substantial number of small entities. In the NOPR, the Commission found that the proposed rule would not impose any regulatory or administrative burdens on a significant number of small entities and that it would

196/ 5 U.S.C. §§ 601-612 (1982).

not require an expenditure of resources by such entities. 197/ No comments were received on this finding and the modifications adopted in the final rule do not materially affect the earlier conclusion.

Accordingly, the Commission certifies that the rule does not have a significant economic impact on a substantial number of small entities.

V. TIMING OF QUARTERLY UPDATES AND EFFECTIVE DATE OF RULE

The Commission establishes a procedure which will be used to establish quarterly updates. The benchmark rates of return will be published on or before the fifteenth of the month following the close of a calendar quarter.

The first quarter following the close of an annual proceeding will run from February 1 to April 30. The second quarter will run from May 1 to August 31, etc.

List of Subjects contained in 18 C.F.R. Part 37: Electric Power Rates, Electric Utilities, Rate of Return.

In consideration of the foregoing, the Commission amends Chapter I, Title 18 of the Code of Federal Regulations, as set forth below, effective 30 days after publication in the Federal Register.

By the Commission.

(S E A L)

Kenneth F. Plumb

Kenneth F. Plumb,
Secretary.

197/ 51 Fed. Reg. at 27,055.

Part 37--- GENERIC DETERMINATION OF RATE OF RETURN ON COMMON EQUITY FOR PUBLIC UTILITIES

1. The authority citation for Part 37 continues to read as follows:

AUTHORITY: Federal Power Act, 16 U.S.C. §§ 791a-825r (1982): Department of Energy Organization Act, 42 U.S.C. §§ 7101-7352 (1982).

2. Section 37.8 is revised to read as follows:

§ 37.8 Transitional provision.

The benchmark rates of return resulting from the first three annual proceedings under this Part will be advisory only. During the advisory period, the Commission may take official notice of the benchmark rates of return in individual rate proceedings if they are not otherwise made a part of the record.

3. Section 37.9 is amended by revising paragraph (a)(3) and removing (a)(4) to read as follows:

§ 37.9 Quarterly Indexing Procedure.

(a) * * *

(3) The benchmark rate of return on common equity for subsequent quarters prior to the conclusion of the next

annual proceeding will be set equal to the average cost of common equity for the jurisdictional operations of public utilities as determined by the formula of paragraph (a)(1) of this section.

* * * * *

LIST OF COMMENTERS

Commenter	Abbreviation
Companies	
1. American Electric Power Service Corp.	AEP
2. Arizona Public Service Co.	APS
3. Associated Utility Services, Inc.	AUS
4. Baltimore Gas and Electric Co.	BGE
5. Boston Edison Co., et al.	BEC
6. Cincinnati Gas and Electric	CGE
7. Detroit Edison	DE
8. Edison Electric Institute	EEI
9. Florida Power & Light Co.	FPL
10. New England Power Co.	NEP
11. Northern States Power Co.	NSP
12. Ocean State Power	OSP
13. Potomac Electric Power Co.	PEPCO
14. Public Service Company of Colorado	PSC
15. Southern California Edison	SCE
16. Southern Company	Southern
17. Southwestern Public Service Co.	SW
18. Utah Power & Light	UPL
19. Virginia Electric & Power	VEPCO
Customers	
20. Alabama Electric Coop., et al.	Cooperatives
21. Allegheny Electric Coop., et al.	Second Cooperatives
22. American Public Power Association	APPA
23. General Services Administration	GSA
24. Wholesale Customer Group	WCG
Other	
25. Financial Analysis Branch Office of Electric Power Regulation Federal Energy Regulatory Commission	FA Staff
26. Florida Public Service Commission	FLA
27. Minnesota Department of Public Service	MINN
28. West Virginia Public Service Commission (Consumer Advocate Division)	WVCAD
29. Alfred W. Whittaker	AWW

SAMPLE OF COMPANIES USED FOR BASE YEAR DIVIDEND YIELD CALCULATION

1. ALLEGHENY POWER SYSTEM
2. AMERICAN ELECTRIC POWER
3. ATLANTIC CITY ELECTRIC
4. AZP GROUP INC
5. BALTIMORE GAS & ELECTRIC
6. BLACK HILLS CORP
7. BOSTON EDISON CO
8. CAROLINA POWER & LIGHT
9. CENTERIOR ENERGY CORP
10. CENTRAL & SOUTH WEST CORP
11. CENTRAL HUDSON GAS & ELEC
12. CENTRAL ILL PUBLIC SERVICE
13. CENTRAL LOUISIANA ELECTRIC
14. CENTRAL MAINE POWER CO
15. CENTRAL VERMONT PUB SERV
16. CINCINNATI GAS & ELECTRIC
17. COMMONWEALTH EDISON
18. COMMONWEALTH ENERGY SYSTEM
19. CONSOLIDATED EDISON OF NY
20. CONSUMERS POWER CO
21. DELMARVA POWER & LIGHT
22. DETROIT EDISON CO
23. DOMINION RESOURCES INC-VA
24. DPL INC
25. DUKE POWER CO
26. DUQUESNE LIGHT CO
27. EASTERN UTILITIES ASSOC
28. EMPIRE DISTRICT ELECTRIC CO
29. FITCHBURG GAS & ELEC LIGHT
30. FLORIDA PROGRESS CORP
31. FPL GROUP INC
32. GENERAL PUBLIC UTILITIES
33. GREEN MOUNTAIN POWER CORP
34. GULF STATES UTILITIES CO
35. HAWAIIAN ELECTRIC INDS
36. HOUSTON INDUSTRIES INC
37. I E INDUSTRIES INC
38. IDAHO POWER CO
39. ILLINOIS POWER CO
40. INTERSTATE POWER CO
41. IOWA RESOURCES INC
42. IOWA-ILLINOIS GAS & ELEC
43. IPALCO ENTERPRISES INC
44. KANSAS CITY POWER & LIGHT
45. KANSAS GAS & ELECTRIC
46. KANSAS POWER & LIGHT
47. KENTUCKY UTILITIES CO
48. LONG ISLAND LIGHTING
49. LOUISVILLE GAS & ELECTRIC
50. MAINE PUBLIC SERVICE
51. MIDDLE SOUTH UTILITIES
52. MIDWEST ENERGY CO
53. MINNESOTA POWER & LIGHT
54. MONTANA POWER CO
55. NEVADA POWER CO
56. NEW ENGLAND ELECTRIC SYSTEM
57. NEW YORK STATE ELEC & GAS
58. NEWPORT ELECTRIC CORP
59. NIAGARA MOHAWK POWER
60. NORTHEAST UTILITIES
61. NORTHERN INDIANA PUBLIC SERV
62. NORTHERN STATES POWER-MN
63. OHIO EDISON CO
64. ORANGE & ROCKLAND UTILITIES
65. PACIFIC GAS & ELECTRIC
66. PACIFICORP
67. PENNSYLVANIA POWER & LIGHT
68. PHILADELPHIA ELECTRIC CO
69. PORTLAND GENERAL CO
70. POTOMAC ELECTRIC POWER
71. PUBLIC SERVICE CO OF COLO
72. PUBLIC SERVICE CO OF IND
73. PUBLIC SERVICE CO OF N H
74. PUBLIC SERVICE ENTERPRISES
75. PUGET SOUND POWER & LIGHT
76. ROCHESTER GAS & ELECTRIC
77. SAN DIEGO GAS & ELECTRIC
78. SAVANNAH ELEC & POWER
79. SCANA CORP
80. SIERRA PACIFIC RESOURCES
81. SOUTHERN CALIF EDISON CO
82. SOUTHERN CO
83. SOUTHERN INDIANA GAS & ELEC
84. ST JOSEPH LIGHT & POWER
85. TECO ENERGY INC
86. TEXAS UTILITIES CO
87. TNP ENTERPRISES INC
88. TUCSON ELECTRIC POWER CO
89. UNION ELECTRIC CO
90. UNITED ILLUMINATING CO
91. UTAH POWER & LIGHT
92. UTILICORP UNITED INC
93. WASHINGTON WATER POWER
94. WISCONSIN ELECTRIC POWER
95. WISCONSIN POWER & LIGHT
96. WISCONSIN PUBLIC SERVICE

UTILITIES EXCLUDED FROM THE SAMPLE FOR THE INDICATED QUARTER
DUE TO EITHER ZERO DIVIDENDS OR A CUT IN DIVIDENDS FOR
THIS QUARTER OR THE PRIOR THREE QUARTERS

YEAR=85 QUARTER=3

TICKER SYMBOL	UTILITY	REASON FOR EXCLUSION
CMS	CONSUMERS POWER CO	DIVIDEND RATE WAS ZERO FOR THE QUARTER ENDING 09/30/85
CTP	CENTRAL MAINE POWER CO	DIVIDEND RATE REDUCED IN THE QUARTER ENDING 12/31/84
FGE	FITCHBURG GAS & ELEC LIGHT	DIVIDEND RATE WAS ZERO FOR THE QUARTER ENDING 09/30/85
GPU	GENERAL PUBLIC UTILITIES	DIVIDEND RATE WAS ZERO FOR THE QUARTER ENDING 09/30/85
LIL	LONG ISLAND LIGHTING	DIVIDEND RATE WAS ZERO FOR THE QUARTER ENDING 09/30/85
MAP	MAINE PUBLIC SERVICE	DIVIDEND RATE WAS ZERO FOR THE QUARTER ENDING 09/30/85
MSU	MIDDLE SOUTH UTILITIES	DIVIDEND RATE WAS ZERO FOR THE QUARTER ENDING 09/30/85
MTP	MONTANA POWER CO	DIVIDEND RATE REDUCED IN THE QUARTER ENDING 03/31/85
PNM	PUBLIC SERVICE CO OF N M	DIVIDEND RATE WAS ZERO FOR THE QUARTER ENDING 09/30/85

N= 9

YEAR=85 QUARTER=4

TICKER SYMBOL	UTILITY	REASON FOR EXCLUSION
CMS	CONSUMERS POWER CO	DIVIDEND RATE WAS ZERO FOR THE QUARTER ENDING 12/31/85
FGE	FITCHBURG GAS & ELEC LIGHT	DIVIDEND RATE WAS ZERO FOR THE QUARTER ENDING 12/31/85
GPU	GENERAL PUBLIC UTILITIES	DIVIDEND RATE WAS ZERO FOR THE QUARTER ENDING 12/31/85
KGE	KANSAS GAS & ELECTRIC	DIVIDEND RATE REDUCED IN THE QUARTER ENDING 12/31/85
LIL	LONG ISLAND LIGHTING	DIVIDEND RATE WAS ZERO FOR THE QUARTER ENDING 12/31/85
MAP	MAINE PUBLIC SERVICE	DIVIDEND RATE WAS ZERO FOR THE QUARTER ENDING 12/31/85
MSU	MIDDLE SOUTH UTILITIES	DIVIDEND RATE WAS ZERO FOR THE QUARTER ENDING 12/31/85
MTP	MONTANA POWER CO	DIVIDEND RATE REDUCED IN THE QUARTER ENDING 03/31/85
NI	NORTHERN INDIANA PUBLIC SERV	DIVIDEND RATE WAS ZERO FOR THE QUARTER ENDING 12/31/85
PNM	PUBLIC SERVICE CO OF N M	DIVIDEND RATE WAS ZERO FOR THE QUARTER ENDING 12/31/85

N= 10

UTILITIES EXCLUDED FROM THE SAMPLE FOR THE INDICATED QUARTER
DUE TO EITHER ZERO DIVIDENDS OR A CUT IN DIVIDENDS FOR
THIS QUARTER OR THE PRIOR THREE QUARTERS

YEAR=86 QUARTER=1

TICKER SYMBOL	UTILITY	REASON FOR EXCLUSION
CMS	CONSUMERS POWER CO	DIVIDEND RATE WAS ZERO FOR THE QUARTER ENDING 03/31/86
FGE	FITCHBURG GAS & ELEC LIGHT	DIVIDEND RATE WAS ZERO FOR THE QUARTER ENDING 03/31/86
GPU	GENERAL PUBLIC UTILITIES	DIVIDEND RATE WAS ZERO FOR THE QUARTER ENDING 03/31/86
KGE	KANSAS GAS & ELECTRIC	DIVIDEND RATE REDUCED IN THE QUARTER ENDING 12/31/85
LIL	LONG ISLAND LIGHTING	DIVIDEND RATE WAS ZERO FOR THE QUARTER ENDING 03/31/86
MAP	MAINE PUBLIC SERVICE	DIVIDEND RATE WAS ZERO FOR THE QUARTER ENDING 03/31/86
MSU	MIDDLE SOUTH UTILITIES	DIVIDEND RATE WAS ZERO FOR THE QUARTER ENDING 03/31/86
NI	NORTHERN INDIANA PUBLIC SERV	DIVIDEND RATE WAS ZERO FOR THE QUARTER ENDING 03/31/86
PIN	PUBLIC SERVICE CO OF IND	DIVIDEND RATE WAS ZERO FOR THE QUARTER ENDING 03/31/86
PNM	PUBLIC SERVICE CO OF N M	DIVIDEND RATE WAS ZERO FOR THE QUARTER ENDING 03/31/86

N= 10

YEAR=86 QUARTER=2

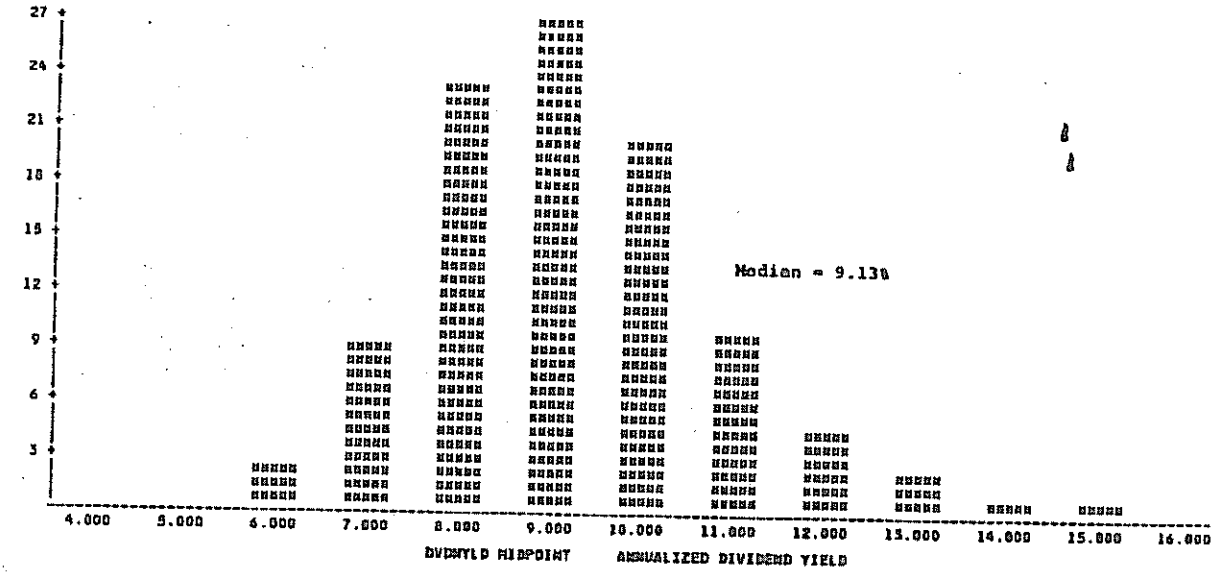
TICKER SYMBOL	UTILITY	REASON FOR EXCLUSION
CMS	CONSUMERS POWER CO	DIVIDEND RATE WAS ZERO FOR THE QUARTER ENDING 06/30/86
DQU	DUQUESNE LIGHT CO	DIVIDEND RATE REDUCED IN THE QUARTER ENDING 06/30/86
FGE	FITCHBURG GAS & ELEC LIGHT	DIVIDEND RATE WAS ZERO FOR THE QUARTER ENDING 06/30/86
GPU	GENERAL PUBLIC UTILITIES	DIVIDEND RATE WAS ZERO FOR THE QUARTER ENDING 06/30/86
GSU	GULF STATES UTILITIES CO	DIVIDEND RATE REDUCED IN THE QUARTER ENDING 06/30/86
KGE	KANSAS GAS & ELECTRIC	DIVIDEND RATE REDUCED IN THE QUARTER ENDING 12/31/85
KLT	KANSAS CITY POWER & LIGHT	DIVIDEND RATE REDUCED IN THE QUARTER ENDING 06/30/86
LIL	LONG ISLAND LIGHTING	DIVIDEND RATE WAS ZERO FOR THE QUARTER ENDING 06/30/86
MAP	MAINE PUBLIC SERVICE	DIVIDEND RATE WAS ZERO FOR THE QUARTER ENDING 03/31/86
MSU	MIDDLE SOUTH UTILITIES	DIVIDEND RATE WAS ZERO FOR THE QUARTER ENDING 06/30/86
NI	NORTHERN INDIANA PUBLIC SERV	DIVIDEND RATE WAS ZERO FOR THE QUARTER ENDING 06/30/86
PIN	PUBLIC SERVICE CO OF IND	DIVIDEND RATE WAS ZERO FOR THE QUARTER ENDING 06/30/86
PNM	PUBLIC SERVICE CO OF N M	DIVIDEND RATE WAS ZERO FOR THE QUARTER ENDING 06/30/86

N= 13

ANNUALIZED DIVIDEND YIELDS FOR THE INDICATED QUARTER FOR
UTILITIES RETAINED IN THE SAMPLE

YEAR=85 QUARTER=3

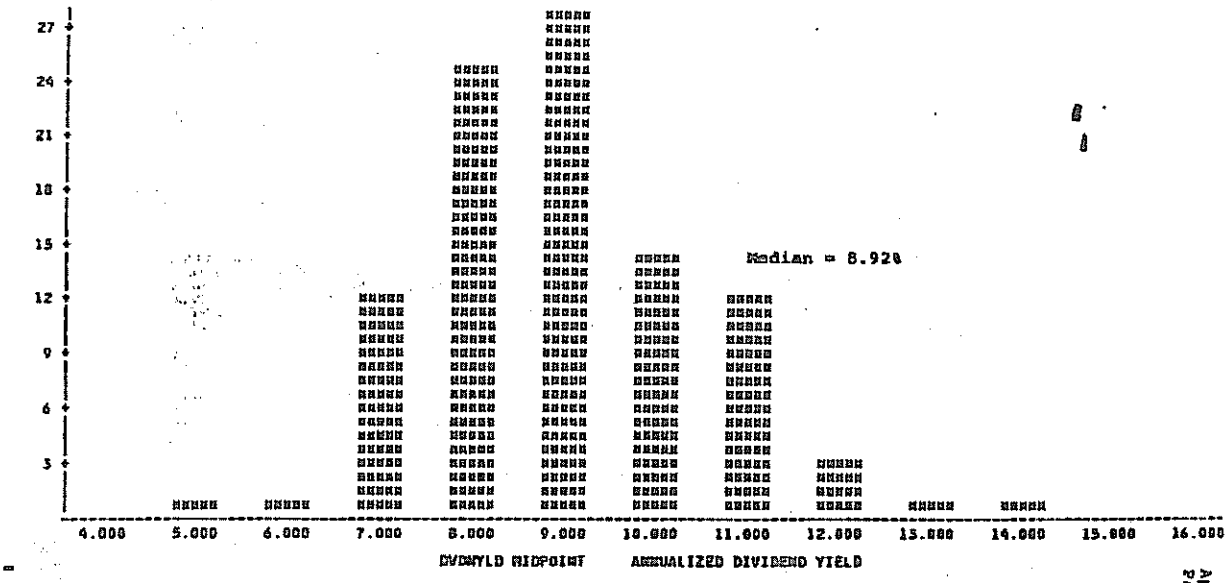
PERCENTAGE BAR CHART



ANNUALIZED DIVIDEND YIELDS FOR THE INDICATED QUARTER FOR
UTILITIES RETAINED IN THE SAMPLE

YEAR=85 QUARTER=4

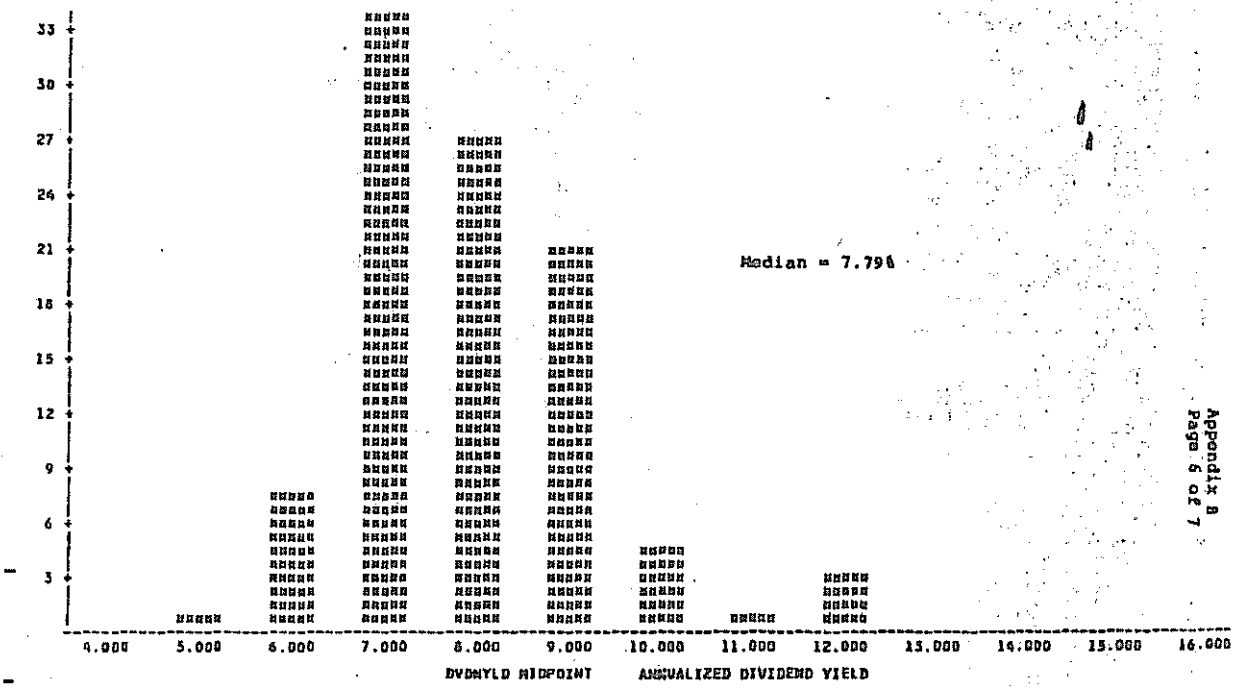
PERCENTAGE BAR CHART



ANNUALIZED DIVIDEND YIELDS FOR THE INDICATED QUARTER FOR
UTILITIES RETAINED IN THE SAMPLE

YEAR=86 QUARTER=1

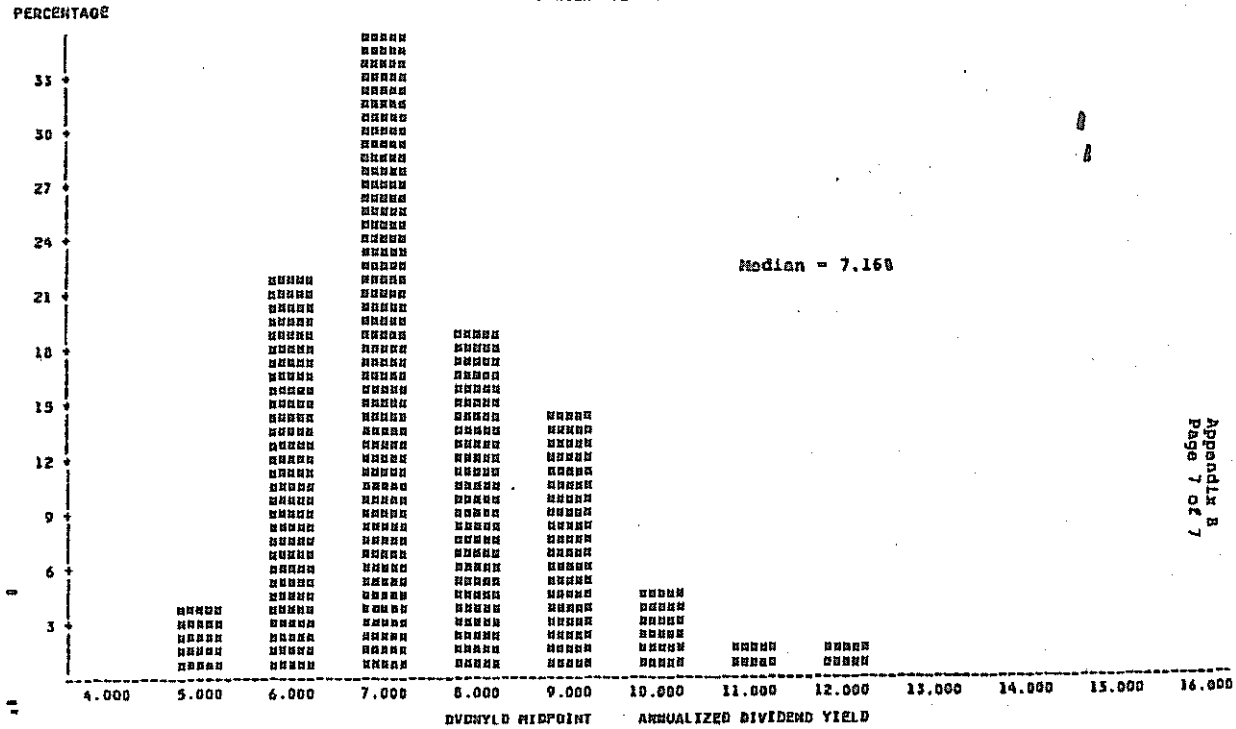
PERCENTAGE BAR CHART



ANNUALIZED DIVIDEND YIELDS FOR THE INDICATED QUARTER FOR
UTILITIES RETAINED IN THE SAMPLE

YEAR=86 QUARTER=2

PERCENTAGE BAR CHART



SAMPLE OF COMPANIES FOR DIVIDEND YIELD UPDATES

1. ALLEGHENY POWER SYSTEM
2. AMERICAN ELECTRIC POWER
3. ATLANTIC CITY ELECTRIC
4. AZP GROUP INC
5. BALTIMORE GAS & ELECTRIC
6. BLACK HILLS CORP
7. BOSTON EDISON CO
8. CAROLINA POWER & LIGHT
9. CENTERIOR ENERGY CORP
10. CENTRAL & SOUTH WEST CORP
11. CENTRAL HUDSON GAS & ELEC
12. CENTRAL ILL PUBLIC SERVICE
13. CENTRAL LOUISIANA ELECTRIC
14. CENTRAL MAINE POWER CO
15. CENTRAL VERMONT PUB SERV
16. CILCORP INC
17. CINCINNATI GAS & ELECTRIC
18. COMMONWEALTH EDISON
19. COMMONWEALTH ENERGY SYSTEM
20. CONSOLIDATED EDISON OF NY
21. CONSUMERS POWER CO
22. DELMARVA POWER & LIGHT
23. DETROIT EDISON CO
24. DOMINION RESOURCES INC-VA
25. DPL INC
26. DUKE POWER CO
27. DUQUESNE LIGHT CO
28. EASTERN UTILITIES ASSOC
29. EMPIRE DISTRICT ELECTRIC CO
30. FITCHBURG GAS & ELEC LIGHT
31. FLORIDA PROGRESS CORP
32. FPL GROUP INC
33. GENERAL PUBLIC UTILITIES
34. GREEN MOUNTAIN POWER CORP
35. GULF STATES UTILITIES CO
36. HAWAIIAN ELECTRIC INDS
37. HOUSTON INDUSTRIES INC
38. I E INDUSTRIES INC
39. IDAHO POWER CO
40. ILLINOIS POWER CO
41. INTERSTATE POWER CO
42. IOWA RESOURCES INC
43. IOWA-ILLINOIS GAS & ELEC
44. IPALCO ENTERPRISES INC
45. KANSAS CITY POWER & LIGHT
46. KANSAS GAS & ELECTRIC
47. KANSAS POWER & LIGHT
48. KENTUCKY UTILITIES CO
49. LONG ISLAND LIGHTING
50. LOUISVILLE GAS & ELECTRIC
51. MAINE PUBLIC SERVICE
52. MIDDLE SOUTH UTILITIES
53. MIDWEST ENERGY CO
54. MINNESOTA POWER & LIGHT
55. MONTANA POWER CO
56. NEVADA POWER CO
57. NEW ENGLAND ELECTRIC SYSTEM
58. NEW YORK STATE ELEC & GAS
59. NEWPORT ELECTRIC CORP
60. NIAGARA MOHAWK POWER
61. NORTHEAST UTILITIES
62. NORTHERN INDIANA PUBLIC SERV
63. NORTHERN STATES POWER-MN
64. OHIO EDISON CO
65. OKLAHOMA GAS & ELECTRIC
66. ORANGE & ROCKLAND UTILITIES
67. PACIFIC GAS & ELECTRIC
68. PACIFICORP
69. PENNSYLVANIA POWER & LIGHT
70. PHILADELPHIA ELECTRIC CO
71. PORTLAND GENERAL CO
72. POTOMAC ELECTRIC POWER
73. PUBLIC SERVICE CO OF COLO
74. PUBLIC SERVICE CO OF IND
75. PUBLIC SERVICE CO OF N H
76. PUBLIC SERVICE CO OF N MEX
77. PUBLIC SERVICE ENTERPRISES
78. PUGET SOUND POWER & LIGHT
79. ROCHESTER GAS & ELECTRIC
80. SAN DIEGO GAS & ELECTRIC
81. SAVANNAH ELEC & POWER
82. SCANA CORP
83. SIERRA PACIFIC RESOURCES
84. SOUTHERN CALIF EDISON CO
85. SOUTHERN CO
86. SOUTHERN INDIANA GAS & ELEC
87. ST JOSEPH LIGHT & POWER
88. TECO ENERGY INC
89. TEXAS UTILITIES CO
90. TNP ENTERPRISES INC
91. TUCSON ELECTRIC POWER CO
92. UNION ELECTRIC CO
93. UNITED ILLUMINATING CO
94. UNITIL CORP
95. UTAH POWER & LIGHT
96. UTILICORP UNITED INC
97. WASHINGTON WATER POWER
98. WISCONSIN ELECTRIC POWER
99. WISCONSIN POWER & LIGHT
100. WISCONSIN PUBLIC SERVICE

Appendix C

MONTANA-DAKOTA UTILITIES CO.
MONTANA CONSUMER COUNSEL
MCC DATA REQUEST
DATED JULY 12, 2013
DOCKET NO. D2012.9.100

MCC-199 **Regarding: Exhibit No._(JSG-09)**
Witness: J. Stephen Gaske

Please provide an electronic copy (Excel) of your Exhibit No._(JSG-09), including all data, supporting documents and worksheets, with formulas and links intact.

Response:

Please see the Excel file on the enclosed CD labeled "MCC-199 JSG-09.xlxs".

MONTANA-DAKOTA UTILITIES CO.
MONTANA CONSUMER COUNSEL
MCC DATA REQUEST
DATED JULY 12, 2013
DOCKET NO. D2012.9.100

MCC-200 **Regarding: Capital Asset Pricing Model**
Witness: J. Stephen Gaske

Please provide a copy of the documents referenced in footnotes 25, 26, 27, 29 and 30 of your rebuttal testimony:

a. **Eugene F. Fama and Kenneth R. French, "The Capital Asset Pricing Model: Theory and Evidence," Journal of Economic Perspectives, Volume 18, Number 3, Summer 2004.**
b. **Fama and French, "The Cross-Section of Expected Stock Returns," Journal of Finance, Vol. XLVII, No. 2, June 1992.**
c. **Value Line Industry Review, March 13, 1992.**
d. **Levhari, D. and Levy, H., "The Capital Asset Pricing Model and the Investment Horizon," Review of Economics and Statistics (February 1977).**
e. **Hawawini, G., "Why Beta Shifts as the Return Interval Changes," Financial Analysts Journal (May-June 1983).**

Response:

a. Please see Attachment A.
b. Please see Attachment B.
c. Please see Attachment C.
d. Please see Attachment D.
e. Please see Attachment E.

Journal of Economic Perspectives—Volume 18, Number 3—Summer 2004—Pages 25–46

The Capital Asset Pricing Model: Theory and Evidence

Eugene F. Fama and Kenneth R. French

The capital asset pricing model (CAPM) of William Sharpe (1964) and John Lintner (1965) marks the birth of asset pricing theory (resulting in a Nobel Prize for Sharpe in 1990). Four decades later, the CAPM is still widely used in applications, such as estimating the cost of capital for firms and evaluating the performance of managed portfolios. It is the centerpiece of MBA investment courses. Indeed, it is often the only asset pricing model taught in these courses.[1]

The attraction of the CAPM is that it offers powerful and intuitively pleasing predictions about how to measure risk and the relation between expected return and risk. Unfortunately, the empirical record of the model is poor—poor enough to invalidate the way it is used in applications. The CAPM's empirical problems may reflect theoretical failings, the result of many simplifying assumptions. But they may also be caused by difficulties in implementing valid tests of the model. For example, the CAPM says that the risk of a stock should be measured relative to a comprehensive "market portfolio" that in principle can include not just traded financial assets, but also consumer durables, real estate and human capital. Even if we take a narrow view of the model and limit its purview to traded financial assets, is it

[1] Although every asset pricing model is a capital asset pricing model, the finance profession reserves the acronym CAPM for the specific model of Sharpe (1964), Lintner (1965) and Black (1972) discussed here. Thus, throughout the paper we refer to the Sharpe-Lintner-Black model as the CAPM.

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legitimate to limit further the market portfolio to U.S. common stocks (a typical choice), or should the market be expanded to include bonds, and other financial assets, perhaps around the world? In the end, we argue that whether the model's problems reflect weaknesses in the theory or in its empirical implementation, the failure of the CAPM in empirical tests implies that most applications of the model are invalid.

We begin by outlining the logic of the CAPM, focusing on its predictions about risk and expected return. We then review the history of empirical work and what it says about shortcomings of the CAPM that pose challenges to be explained by alternative models.

The Logic of the CAPM

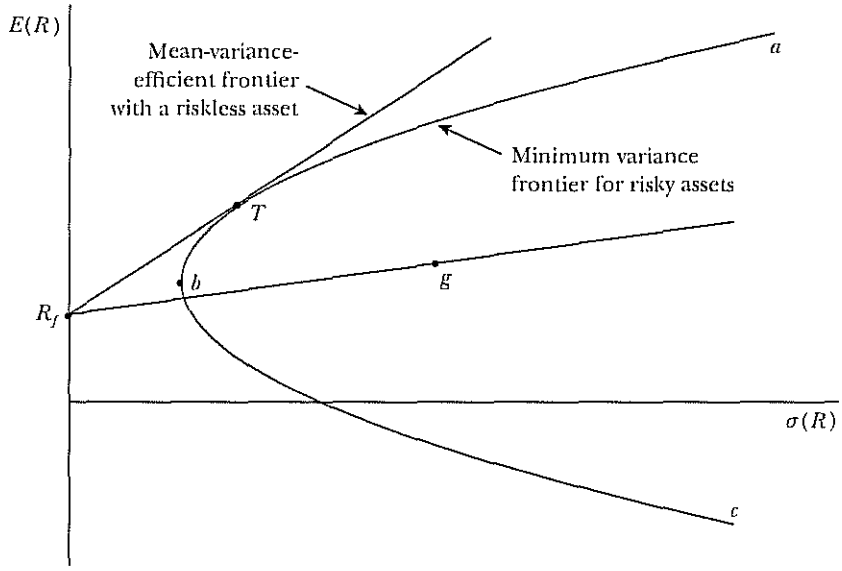
The CAPM builds on the model of portfolio choice developed by Harry Markowitz (1959). In Markowitz's model, an investor selects a portfolio at time $t-1$ that produces a stochastic return at t. The model assumes investors are risk averse and, when choosing among portfolios, they care only about the mean and variance of their one-period investment return. As a result, investors choose "mean-variance-efficient" portfolios, in the sense that the portfolios 1) minimize the variance of portfolio return, given expected return, and 2) maximize expected return, given variance. Thus, the Markowitz approach is often called a "mean-variance model."

The portfolio model provides an algebraic condition on asset weights in mean-variance-efficient portfolios. The CAPM turns this algebraic statement into a testable prediction about the relation between risk and expected return by identifying a portfolio that must be efficient if asset prices are to clear the market of all assets.

Sharpe (1964) and Lintner (1965) add two key assumptions to the Markowitz model to identify a portfolio that must be mean-variance-efficient. The first assumption is *complete agreement*: given market clearing asset prices at $t-1$, investors agree on the joint distribution of asset returns from $t-1$ to t. And this distribution is the true one—that is, it is the distribution from which the returns we use to test the model are drawn. The second assumption is that there is *borrowing and lending at a risk-free rate*, which is the same for all investors and does not depend on the amount borrowed or lent.

Figure 1 describes portfolio opportunities and tells the CAPM story. The horizontal axis shows portfolio risk, measured by the standard deviation of portfolio return; the vertical axis shows expected return. The curve *abc*, which is called the minimum variance frontier, traces combinations of expected return and risk for portfolios of risky assets that minimize return variance at different levels of expected return. (These portfolios do not include risk-free borrowing and lending.) The tradeoff between risk and expected return for minimum variance portfolios is apparent. For example, an investor who wants a high expected return, perhaps at point *a*, must accept high volatility. At point *T*, the investor can have an interme-

Figure 1
Investment Opportunities



diate expected return with lower volatility. If there is no risk-free borrowing or lending, only portfolios above b along abc are mean-variance-efficient, since these portfolios also maximize expected return, given their return variances.

Adding risk-free borrowing and lending turns the efficient set into a straight line. Consider a portfolio that invests the proportion x of portfolio funds in a risk-free security and $1 - x$ in some portfolio g. If all funds are invested in the risk-free security—that is, they are loaned at the risk-free rate of interest—the result is the point R_f in Figure 1, a portfolio with zero variance and a risk-free rate of return. Combinations of risk-free lending and positive investment in g plot on the straight line between R_f and g. Points to the right of g on the line represent borrowing at the risk-free rate, with the proceeds from the borrowing used to increase investment in portfolio g. In short, portfolios that combine risk-free lending or borrowing with some risky portfolio g plot along a straight line from R_f through g in Figure 1.[2]

[2] Formally, the return, expected return and standard deviation of return on portfolios of the risk-free asset f and a risky portfolio g vary with x, the proportion of portfolio funds invested in f, as

$$R_p = xR_f + (1 - x)R_g,$$

$$E(R_p) = xR_f + (1 - x)E(R_g),$$

$$\sigma(R_p) = (1 - x)\sigma(R_g),\ x \leq 1.0,$$

which together imply that the portfolios plot along the line from R_f through g in Figure 1.

To obtain the mean-variance-efficient portfolios available with risk-free borrowing and lending, one swings a line from R_f in Figure 1 up and to the left as far as possible, to the tangency portfolio T. We can then see that all efficient portfolios are combinations of the risk-free asset (either risk-free borrowing or lending) and a single risky tangency portfolio, T. This key result is Tobin's (1958) "separation theorem."

The punch line of the CAPM is now straightforward. With complete agreement about distributions of returns, all investors see the same opportunity set (Figure 1), and they combine the same risky tangency portfolio T with risk-free lending or borrowing. Since all investors hold the same portfolio T of risky assets, it must be the value-weight market portfolio of risky assets. Specifically, each risky asset's weight in the tangency portfolio, which we now call M (for the "market"), must be the total market value of all outstanding units of the asset divided by the total market value of all risky assets. In addition, the risk-free rate must be set (along with the prices of risky assets) to clear the market for risk-free borrowing and lending.

In short, the CAPM assumptions imply that the market portfolio M must be on the minimum variance frontier if the asset market is to clear. This means that the algebraic relation that holds for any minimum variance portfolio must hold for the market portfolio. Specifically, if there are N risky assets,

$$\text{(Minimum Variance Condition for } M) \quad E(R_i) = E(R_{ZM}) + [E(R_M) - E(R_{ZM})]\beta_{iM}, \; i = 1, \ldots, N.$$

In this equation, $E(R_i)$ is the expected return on asset i, and β_{iM}, the market beta of asset i, is the covariance of its return with the market return divided by the variance of the market return,

$$\text{(Market Beta)} \quad \beta_{iM} = \frac{\text{cov}(R_i, R_M)}{\sigma^2(R_M)}.$$

The first term on the right-hand side of the minimum variance condition, $E(R_{ZM})$, is the expected return on assets that have market betas equal to zero, which means their returns are uncorrelated with the market return. The second term is a risk premium—the market beta of asset i, β_{iM}, times the premium per unit of beta, which is the expected market return, $E(R_M)$, minus $E(R_{ZM})$.

Since the market beta of asset i is also the slope in the regression of its return on the market return, a common (and correct) interpretation of beta is that it measures the sensitivity of the asset's return to variation in the market return. But there is another interpretation of beta more in line with the spirit of the portfolio model that underlies the CAPM. The risk of the market portfolio, as measured by the variance of its return (the denominator of β_{iM}), is a weighted average of the covariance risks of the assets in M (the numerators of β_{iM} for different assets).

Thus, β_{iM} is the covariance risk of asset i in M measured relative to the average covariance risk of assets, which is just the variance of the market return.[3] In economic terms, β_{iM} is proportional to the risk each dollar invested in asset i contributes to the market portfolio.

The last step in the development of the Sharpe-Lintner model is to use the assumption of risk-free borrowing and lending to nail down $E(R_{ZM})$, the expected return on zero-beta assets. A risky asset's return is uncorrelated with the market return—its beta is zero—when the average of the asset's covariances with the returns on other assets just offsets the variance of the asset's return. Such a risky asset is riskless in the market portfolio in the sense that it contributes nothing to the variance of the market return.

When there is risk-free borrowing and lending, the expected return on assets that are uncorrelated with the market return, $E(R_{ZM})$, must equal the risk-free rate, R_f. The relation between expected return and beta then becomes the familiar Sharpe-Lintner CAPM equation,

$$\text{(Sharpe-Lintner CAPM)} \quad E(R_i) = R_f + [E(R_M) - R_f]\beta_{iM},\ i = 1, \ldots, N.$$

In words, the expected return on any asset i is the risk-free interest rate, R_f, plus a risk premium, which is the asset's market beta, β_{iM}, times the premium per unit of beta risk, $E(R_M) - R_f$.

Unrestricted risk-free borrowing and lending is an unrealistic assumption. Fischer Black (1972) develops a version of the CAPM without risk-free borrowing or lending. He shows that the CAPM's key result—that the market portfolio is mean-variance-efficient—can be obtained by instead allowing unrestricted short sales of risky assets. In brief, back in Figure 1, if there is no risk-free asset, investors select portfolios from along the mean-variance-efficient frontier from a to b. Market clearing prices imply that when one weights the efficient portfolios chosen by investors by their (positive) shares of aggregate invested wealth, the resulting portfolio is the market portfolio. The market portfolio is thus a portfolio of the efficient portfolios chosen by investors. With unrestricted short selling of risky assets, portfolios made up of efficient portfolios are themselves efficient. Thus, the market portfolio is efficient, which means that the minimum variance condition for M given above holds, and it is the expected return-risk relation of the Black CAPM.

The relations between expected return and market beta of the Black and Sharpe-Lintner versions of the CAPM differ only in terms of what each says about $E(R_{ZM})$, the expected return on assets uncorrelated with the market. The Black version says only that $E(R_{ZM})$ must be less than the expected market return, so the

[3] Formally, if x_{iM} is the weight of asset i in the market portfolio, then the variance of the portfolio's return is

$$\sigma^2(R_M) = Cov(R_M, R_M) = Cov\left(\sum_{i=1}^{N} x_{iM}R_i,\ R_M\right) = \sum_{i=1}^{N} x_{iM}Cov(R_i, R_M).$$

premium for beta is positive. In contrast, in the Sharpe-Lintner version of the model, $E(R_{ZM})$ must be the risk-free interest rate, R_f, and the premium per unit of beta risk is $E(R_M) - R_f$.

The assumption that short selling is unrestricted is as unrealistic as unrestricted risk-free borrowing and lending. If there is no risk-free asset and short sales of risky assets are not allowed, mean-variance investors still choose efficient portfolios—points above *b* on the *abc* curve in Figure 1. But when there is no short selling of risky assets and no risk-free asset, the algebra of portfolio efficiency says that portfolios made up of efficient portfolios are not typically efficient. This means that the market portfolio, which is a portfolio of the efficient portfolios chosen by investors, is not typically efficient. And the CAPM relation between expected return and market beta is lost. This does not rule out predictions about expected return and betas with respect to other efficient portfolios—if theory can specify portfolios that must be efficient if the market is to clear. But so far this has proven impossible.

In short, the familiar CAPM equation relating expected asset returns to their market betas is just an application to the market portfolio of the relation between expected return and portfolio beta that holds in any mean-variance-efficient portfolio. The efficiency of the market portfolio is based on many unrealistic assumptions, including complete agreement and either unrestricted risk-free borrowing and lending or unrestricted short selling of risky assets. But all interesting models involve unrealistic simplifications, which is why they must be tested against data.

Early Empirical Tests

Tests of the CAPM are based on three implications of the relation between expected return and market beta implied by the model. First, expected returns on all assets are linearly related to their betas, and no other variable has marginal explanatory power. Second, the beta premium is positive, meaning that the expected return on the market portfolio exceeds the expected return on assets whose returns are uncorrelated with the market return. Third, in the Sharpe-Lintner version of the model, assets uncorrelated with the market have expected returns equal to the risk-free interest rate, and the beta premium is the expected market return minus the risk-free rate. Most tests of these predictions use either cross-section or time-series regressions. Both approaches date to early tests of the model.

Tests on Risk Premiums

The early cross-section regression tests focus on the Sharpe-Lintner model's predictions about the intercept and slope in the relation between expected return and market beta. The approach is to regress a cross-section of average asset returns on estimates of asset betas. The model predicts that the intercept in these regressions is the risk-free interest rate, R_f, and the coefficient on beta is the expected return on the market in excess of the risk-free rate, $E(R_M) - R_f$.

Two problems in these tests quickly became apparent. First, estimates of beta

for individual assets are imprecise, creating a measurement error problem when they are used to explain average returns. Second, the regression residuals have common sources of variation, such as industry effects in average returns. Positive correlation in the residuals produces downward bias in the usual ordinary least squares estimates of the standard errors of the cross-section regression slopes.

To improve the precision of estimated betas, researchers such as Blume (1970), Friend and Blume (1970) and Black, Jensen and Scholes (1972) work with portfolios, rather than individual securities. Since expected returns and market betas combine in the same way in portfolios, if the CAPM explains security returns it also explains portfolio returns.[4] Estimates of beta for diversified portfolios are more precise than estimates for individual securities. Thus, using portfolios in cross-section regressions of average returns on betas reduces the critical errors in variables problem. Grouping, however, shrinks the range of betas and reduces statistical power. To mitigate this problem, researchers sort securities on beta when forming portfolios; the first portfolio contains securities with the lowest betas, and so on, up to the last portfolio with the highest beta assets. This sorting procedure is now standard in empirical tests.

Fama and MacBeth (1973) propose a method for addressing the inference problem caused by correlation of the residuals in cross-section regressions. Instead of estimating a single cross-section regression of average monthly returns on betas, they estimate month-by-month cross-section regressions of monthly returns on betas. The times-series means of the monthly slopes and intercepts, along with the standard errors of the means, are then used to test whether the average premium for beta is positive and whether the average return on assets uncorrelated with the market is equal to the average risk-free interest rate. In this approach, the standard errors of the average intercept and slope are determined by the month-to-month variation in the regression coefficients, which fully captures the effects of residual correlation on variation in the regression coefficients, but sidesteps the problem of actually estimating the correlations. The residual correlations are, in effect, captured via repeated sampling of the regression coefficients. This approach also becomes standard in the literature.

Jensen (1968) was the first to note that the Sharpe-Lintner version of the

[4] Formally, if x_{ip}, $i = 1, \ldots, N$, are the weights for assets in some portfolio p, the expected return and market beta for the portfolio are related to the expected returns and betas of assets as

$$E(R_p) = \sum_{i=1}^{N} x_{ip}E(R_i), \text{ and } \beta_{pM} = \sum_{i=1}^{N} x_{ip}\beta_{pM}.$$

Thus, the CAPM relation between expected return and beta,

$$E(R_i) = E(R_f) + [E(R_M) - E(R_f)]\beta_{iM},$$

holds when asset i is a portfolio, as well as when i is an individual security.

relation between expected return and market beta also implies a time-series regression test. The Sharpe-Lintner CAPM says that the expected value of an asset's excess return (the asset's return minus the risk-free interest rate, $R_{it} - R_{ft}$) is completely explained by its expected CAPM risk premium (its beta times the expected value of $R_{Mt} - R_{ft}$). This implies that "Jensen's alpha," the intercept term in the time-series regression,

$$\text{(Time-Series Regression)} \quad R_{it} - R_{ft} = \alpha_i + \beta_{iM}(R_{Mt} - R_{ft}) + \varepsilon_{it},$$

is zero for each asset.

The early tests firmly reject the Sharpe-Lintner version of the CAPM. There is a positive relation between beta and average return, but it is too "flat." Recall that, in cross-section regressions, the Sharpe-Lintner model predicts that the intercept is the risk-free rate and the coefficient on beta is the expected market return in excess of the risk-free rate, $E(R_M) - R_f$. The regressions consistently find that the intercept is greater than the average risk-free rate (typically proxied as the return on a one-month Treasury bill), and the coefficient on beta is less than the average excess market return (proxied as the average return on a portfolio of U.S. common stocks minus the Treasury bill rate). This is true in the early tests, such as Douglas (1968), Black, Jensen and Scholes (1972), Miller and Scholes (1972), Blume and Friend (1973) and Fama and MacBeth (1973), as well as in more recent cross-section regression tests, like Fama and French (1992).

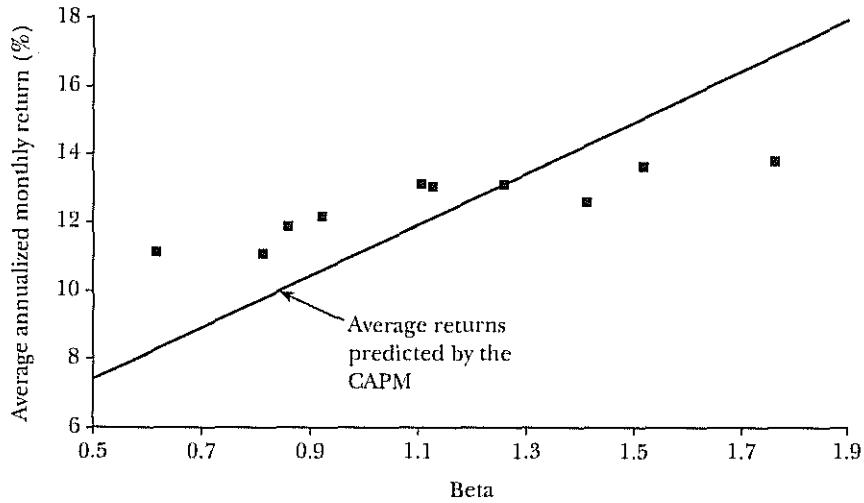
The evidence that the relation between beta and average return is too flat is confirmed in time-series tests, such as Friend and Blume (1970), Black, Jensen and Scholes (1972) and Stambaugh (1982). The intercepts in time-series regressions of excess asset returns on the excess market return are positive for assets with low betas and negative for assets with high betas.

Figure 2 provides an updated example of the evidence. In December of each year, we estimate a preranking beta for every NYSE (1928–2003), AMEX (1963–2003) and NASDAQ (1972–2003) stock in the CRSP (Center for Research in Security Prices of the University of Chicago) database, using two to five years (as available) of prior monthly returns.[5] We then form ten value-weight portfolios based on these preranking betas and compute their returns for the next twelve months. We repeat this process for each year from 1928 to 2003. The result is 912 monthly returns on ten beta-sorted portfolios. Figure 2 plots each portfolio's average return against its postranking beta, estimated by regressing its monthly returns for 1928–2003 on the return on the CRSP value-weight portfolio of U.S. common stocks.

The Sharpe-Lintner CAPM predicts that the portfolios plot along a straight

[5] To be included in the sample for year t, a security must have market equity data (price times shares outstanding) for December of $t - 1$, and CRSP must classify it as ordinary common equity. Thus, we exclude securities such as American Depository Receipts (ADRs) and Real Estate Investment Trusts (REITs).

Figure 2
Average Annualized Monthly Return versus Beta for Value Weight Portfolios Formed on Prior Beta, 1928–2003



line, with an intercept equal to the risk-free rate, R_f, and a slope equal to the expected excess return on the market, $E(R_M) - R_f$. We use the average one-month Treasury bill rate and the average excess CRSP market return for 1928–2003 to estimate the predicted line in Figure 2. Confirming earlier evidence, the relation between beta and average return for the ten portfolios is much flatter than the Sharpe-Lintner CAPM predicts. The returns on the low beta portfolios are too high, and the returns on the high beta portfolios are too low. For example, the predicted return on the portfolio with the lowest beta is 8.3 percent per year; the actual return is 11.1 percent. The predicted return on the portfolio with the highest beta is 16.8 percent per year; the actual is 13.7 percent.

Although the observed premium per unit of beta is lower than the Sharpe-Lintner model predicts, the relation between average return and beta in Figure 2 is roughly linear. This is consistent with the Black version of the CAPM, which predicts only that the beta premium is positive. Even this less restrictive model, however, eventually succumbs to the data.

Testing Whether Market Betas Explain Expected Returns

The Sharpe-Lintner and Black versions of the CAPM share the prediction that the market portfolio is mean-variance-efficient. This implies that differences in expected return across securities and portfolios are entirely explained by differences in market beta; other variables should add nothing to the explanation of expected return. This prediction plays a prominent role in tests of the CAPM. In the early work, the weapon of choice is cross-section regressions.

In the framework of Fama and MacBeth (1973), one simply adds predetermined explanatory variables to the month-by-month cross-section regressions of

returns on beta. If all differences in expected return are explained by beta, the average slopes on the additional variables should not be reliably different from zero. Clearly, the trick in the cross-section regression approach is to choose specific additional variables likely to expose any problems of the CAPM prediction that, because the market portfolio is efficient, market betas suffice to explain expected asset returns.

For example, in Fama and MacBeth (1973) the additional variables are squared market betas (to test the prediction that the relation between expected return and beta is linear) and residual variances from regressions of returns on the market return (to test the prediction that market beta is the only measure of risk needed to explain expected returns). These variables do not add to the explanation of average returns provided by beta. Thus, the results of Fama and MacBeth (1973) are consistent with the hypothesis that their market proxy—an equal-weight portfolio of NYSE stocks—is on the minimum variance frontier.

The hypothesis that market betas completely explain expected returns can also be tested using time-series regressions. In the time-series regression described above (the excess return on asset i regressed on the excess market return), the intercept is the difference between the asset's average excess return and the excess return predicted by the Sharpe-Lintner model, that is, beta times the average excess market return. If the model holds, there is no way to group assets into portfolios whose intercepts are reliably different from zero. For example, the intercepts for a portfolio of stocks with high ratios of earnings to price and a portfolio of stocks with low earning-price ratios should both be zero. Thus, to test the hypothesis that market betas suffice to explain expected returns, one estimates the time-series regression for a set of assets (or portfolios) and then jointly tests the vector of regression intercepts against zero. The trick in this approach is to choose the left-hand-side assets (or portfolios) in a way likely to expose any shortcoming of the CAPM prediction that market betas suffice to explain expected asset returns.

In early applications, researchers use a variety of tests to determine whether the intercepts in a set of time-series regressions are all zero. The tests have the same asymptotic properties, but there is controversy about which has the best small sample properties. Gibbons, Ross and Shanken (1989) settle the debate by providing an F-test on the intercepts that has exact small-sample properties. They also show that the test has a simple economic interpretation. In effect, the test constructs a candidate for the tangency portfolio T in Figure 1 by optimally combining the market proxy and the left-hand-side assets of the time-series regressions. The estimator then tests whether the efficient set provided by the combination of this tangency portfolio and the risk-free asset is reliably superior to the one obtained by combining the risk-free asset with the market proxy alone. In other words, the Gibbons, Ross and Shanken statistic tests whether the market proxy is the tangency portfolio in the set of portfolios that can be constructed by combining the market portfolio with the specific assets used as dependent variables in the time-series regressions.

Enlightened by this insight of Gibbons, Ross and Shanken (1989), one can see

a similar interpretation of the cross-section regression test of whether market betas suffice to explain expected returns. In this case, the test is whether the additional explanatory variables in a cross-section regression identify patterns in the returns on the left-hand-side assets that are not explained by the assets' market betas. This amounts to testing whether the market proxy is on the minimum variance frontier that can be constructed using the market proxy and the left-hand-side assets included in the tests.

An important lesson from this discussion is that time-series and cross-section regressions do not, strictly speaking, test the CAPM. What is literally tested is whether a specific proxy for the market portfolio (typically a portfolio of U.S. common stocks) is efficient in the set of portfolios that can be constructed from it and the left-hand-side assets used in the test. One might conclude from this that the CAPM has never been tested, and prospects for testing it are not good because 1) the set of left-hand-side assets does not include all marketable assets, and 2) data for the true market portfolio of all assets are likely beyond reach (Roll, 1977; more on this later). But this criticism can be leveled at tests of any economic model when the tests are less than exhaustive or when they use proxies for the variables called for by the model.

The bottom line from the early cross-section regression tests of the CAPM, such as Fama and MacBeth (1973), and the early time-series regression tests, like Gibbons (1982) and Stambaugh (1982), is that standard market proxies seem to be on the minimum variance frontier. That is, the central predictions of the Black version of the CAPM, that market betas suffice to explain expected returns and that the risk premium for beta is positive, seem to hold. But the more specific prediction of the Sharpe-Lintner CAPM that the premium per unit of beta is the expected market return minus the risk-free interest rate is consistently rejected.

The success of the Black version of the CAPM in early tests produced a consensus that the model is a good description of expected returns. These early results, coupled with the model's simplicity and intuitive appeal, pushed the CAPM to the forefront of finance.

Recent Tests

Starting in the late 1970s, empirical work appears that challenges even the Black version of the CAPM. Specifically, evidence mounts that much of the variation in expected return is unrelated to market beta.

The first blow is Basu's (1977) evidence that when common stocks are sorted on earnings-price ratios, future returns on high E/P stocks are higher than predicted by the CAPM. Banz (1981) documents a size effect: when stocks are sorted on market capitalization (price times shares outstanding), average returns on small stocks are higher than predicted by the CAPM. Bhandari (1988) finds that high debt-equity ratios (book value of debt over the market value of equity, a measure of leverage) are associated with returns that are too high relative to their market betas.

Finally, Statman (1980) and Rosenberg, Reid and Lanstein (1985) document that stocks with high book-to-market equity ratios (B/M, the ratio of the book value of a common stock to its market value) have high average returns that are not captured by their betas.

There is a theme in the contradictions of the CAPM summarized above. Ratios involving stock prices have information about expected returns missed by market betas. On reflection, this is not surprising. A stock's price depends not only on the expected cash flows it will provide, but also on the expected returns that discount expected cash flows back to the present. Thus, in principle, the cross-section of prices has information about the cross-section of expected returns. (A high expected return implies a high discount rate and a low price.) The cross-section of stock prices is, however, arbitrarily affected by differences in scale (or units). But with a judicious choice of scaling variable X, the ratio X/P can reveal differences in the cross-section of expected stock returns. Such ratios are thus prime candidates to expose shortcomings of asset pricing models—in the case of the CAPM, shortcomings of the prediction that market betas suffice to explain expected returns (Ball, 1978). The contradictions of the CAPM summarized above suggest that earnings-price, debt-equity and book-to-market ratios indeed play this role.

Fama and French (1992) update and synthesize the evidence on the empirical failures of the CAPM. Using the cross-section regression approach, they confirm that size, earnings-price, debt-equity and book-to-market ratios add to the explanation of expected stock returns provided by market beta. Fama and French (1996) reach the same conclusion using the time-series regression approach applied to portfolios of stocks sorted on price ratios. They also find that different price ratios have much the same information about expected returns. This is not surprising given that price is the common driving force in the price ratios, and the numerators are just scaling variables used to extract the information in price about expected returns.

Fama and French (1992) also confirm the evidence (Reinganum, 1981; Stambaugh, 1982; Lakonishok and Shapiro, 1986) that the relation between average return and beta for common stocks is even flatter after the sample periods used in the early empirical work on the CAPM. The estimate of the beta premium is, however, clouded by statistical uncertainty (a large standard error). Kothari, Shanken and Sloan (1995) try to resuscitate the Sharpe-Lintner CAPM by arguing that the weak relation between average return and beta is just a chance result. But the strong evidence that other variables capture variation in expected return missed by beta makes this argument irrelevant. If betas do not suffice to explain expected returns, the market portfolio is not efficient, and the CAPM is dead in its tracks. Evidence on the size of the market premium can neither save the model nor further doom it.

The synthesis of the evidence on the empirical problems of the CAPM provided by Fama and French (1992) serves as a catalyst, marking the point when it is generally acknowledged that the CAPM has potentially fatal problems. Research then turns to explanations.

One possibility is that the CAPM's problems are spurious, the result of data dredging—publication-hungry researchers scouring the data and unearthing contradictions that occur in specific samples as a result of chance. A standard response to this concern is to test for similar findings in other samples. Chan, Hamao and Lakonishok (1991) find a strong relation between book-to-market equity (B/M) and average return for Japanese stocks. Capaul, Rowley and Sharpe (1993) observe a similar B/M effect in four European stock markets and in Japan. Fama and French (1998) find that the price ratios that produce problems for the CAPM in U.S. data show up in the same way in the stock returns of twelve non-U.S. major markets, and they are present in emerging market returns. This evidence suggests that the contradictions of the CAPM associated with price ratios are not sample specific.

Explanations: Irrational Pricing or Risk

Among those who conclude that the empirical failures of the CAPM are fatal, two stories emerge. On one side are the behavioralists. Their view is based on evidence that stocks with high ratios of book value to market price are typically firms that have fallen on bad times, while low B/M is associated with growth firms (Lakonishok, Shleifer and Vishny, 1994; Fama and French, 1995). The behavioralists argue that sorting firms on book-to-market ratios exposes investor overreaction to good and bad times. Investors overextrapolate past performance, resulting in stock prices that are too high for growth (low B/M) firms and too low for distressed (high B/M, so-called value) firms. When the overreaction is eventually corrected, the result is high returns for value stocks and low returns for growth stocks. Proponents of this view include DeBondt and Thaler (1987), Lakonishok, Shleifer and Vishny (1994) and Haugen (1995).

The second story for explaining the empirical contradictions of the CAPM is that they point to the need for a more complicated asset pricing model. The CAPM is based on many unrealistic assumptions. For example, the assumption that investors care only about the mean and variance of one-period portfolio returns is extreme. It is reasonable that investors also care about how their portfolio return covaries with labor income and future investment opportunities, so a portfolio's return variance misses important dimensions of risk. If so, market beta is not a complete description of an asset's risk, and we should not be surprised to find that differences in expected return are not completely explained by differences in beta. In this view, the search should turn to asset pricing models that do a better job explaining average returns.

Merton's (1973) intertemporal capital asset pricing model (ICAPM) is a natural extension of the CAPM. The ICAPM begins with a different assumption about investor objectives. In the CAPM, investors care only about the wealth their portfolio produces at the end of the current period. In the ICAPM, investors are concerned not only with their end-of-period payoff, but also with the opportunities

they will have to consume or invest the payoff. Thus, when choosing a portfolio at time $t - 1$, ICAPM investors consider how their wealth at t might vary with future *state variables*, including labor income, the prices of consumption goods and the nature of portfolio opportunities at t, and expectations about the labor income, consumption and investment opportunities to be available after t.

Like CAPM investors, ICAPM investors prefer high expected return and low return variance. But ICAPM investors are also concerned with the covariances of portfolio returns with state variables. As a result, optimal portfolios are "multifactor efficient," which means they have the largest possible expected returns, given their return variances and the covariances of their returns with the relevant state variables.

Fama (1996) shows that the ICAPM generalizes the logic of the CAPM. That is, if there is risk-free borrowing and lending or if short sales of risky assets are allowed, market clearing prices imply that the market portfolio is multifactor efficient. Moreover, multifactor efficiency implies a relation between expected return and beta risks, but it requires additional betas, along with a market beta, to explain expected returns.

An ideal implementation of the ICAPM would specify the state variables that affect expected returns. Fama and French (1993) take a more indirect approach, perhaps more in the spirit of Ross's (1976) arbitrage pricing theory. They argue that though size and book-to-market equity are not themselves state variables, the higher average returns on small stocks and high book-to-market stocks reflect unidentified state variables that produce undiversifiable risks (covariances) in returns that are not captured by the market return and are priced separately from market betas. In support of this claim, they show that the returns on the stocks of small firms covary more with one another than with returns on the stocks of large firms, and returns on high book-to-market (value) stocks covary more with one another than with returns on low book-to-market (growth) stocks. Fama and French (1995) show that there are similar size and book-to-market patterns in the covariation of fundamentals like earnings and sales.

Based on this evidence, Fama and French (1993, 1996) propose a three-factor model for expected returns,

$$\text{(Three-Factor Model)} \quad E(R_{it}) - R_{ft} = \beta_{iM}[E(R_{Mt}) - R_{ft}] + \beta_{is}E(SMB_t) + \beta_{ih}E(HML_t).$$

In this equation, SMB_t (small minus big) is the difference between the returns on diversified portfolios of small and big stocks, HML_t (high minus low) is the difference between the returns on diversified portfolios of high and low B/M stocks, and the betas are slopes in the multiple regression of $R_{it} - R_{ft}$ on $R_{Mt} - R_{ft}$, SMB_t and HML_t.

For perspective, the average value of the market premium $R_{Mt} - R_{ft}$ for 1927–2003 is 8.3 percent per year, which is 3.5 standard errors from zero. The

average values of SMB_t, and HML_t are 3.6 percent and 5.0 percent per year, and they are 2.1 and 3.1 standard errors from zero. All three premiums are volatile, with annual standard deviations of 21.0 percent $(R_{Mt} - R_{ft})$, 14.6 percent (SMB_t) and 14.2 percent (HML_t) per year. Although the average values of the premiums are large, high volatility implies substantial uncertainty about the true expected premiums.

One implication of the expected return equation of the three-factor model is that the intercept α_i in the time-series regression,

$$R_{it} - R_{ft} = \alpha_i + \beta_{iM}(R_{Mt} - R_{ft}) + \beta_{is}SMB_t + \beta_{ih}HML_t + \varepsilon_{it},$$

is zero for all assets i. Using this criterion, Fama and French (1993, 1996) find that the model captures much of the variation in average return for portfolios formed on size, book-to-market equity and other price ratios that cause problems for the CAPM. Fama and French (1998) show that an international version of the model performs better than an international CAPM in describing average returns on portfolios formed on scaled price variables for stocks in 13 major markets.

The three-factor model is now widely used in empirical research that requires a model of expected returns. Estimates of α_i from the time-series regression above are used to calibrate how rapidly stock prices respond to new information (for example, Loughran and Ritter, 1995; Mitchell and Stafford, 2000). They are also used to measure the special information of portfolio managers, for example, in Carhart's (1997) study of mutual fund performance. Among practitioners like Ibbotson Associates, the model is offered as an alternative to the CAPM for estimating the cost of equity capital.

From a theoretical perspective, the main shortcoming of the three-factor model is its empirical motivation. The small-minus-big (SMB) and high-minus-low (HML) explanatory returns are not motivated by predictions about state variables of concern to investors. Instead they are brute force constructs meant to capture the patterns uncovered by previous work on how average stock returns vary with size and the book-to-market equity ratio.

But this concern is not fatal. The ICAPM does not require that the additional portfolios used along with the market portfolio to explain expected returns "mimic" the relevant state variables. In both the ICAPM and the arbitrage pricing theory, it suffices that the additional portfolios are well diversified (in the terminology of Fama, 1996, they are multifactor minimum variance) and that they are sufficiently different from the market portfolio to capture covariation in returns and variation in expected returns missed by the market portfolio. Thus, adding diversified portfolios that capture covariation in returns and variation in average returns left unexplained by the market is in the spirit of both the ICAPM and the Ross's arbitrage pricing theory.

The behavioralists are not impressed by the evidence for a risk-based explanation of the failures of the CAPM. They typically concede that the three-factor model captures covariation in returns missed by the market return and that it picks

up much of the size and value effects in average returns left unexplained by the CAPM. But their view is that the average return premium associated with the model's book-to-market factor—which does the heavy lifting in the improvements to the CAPM—is itself the result of investor overreaction that happens to be correlated across firms in a way that just looks like a risk story. In short, in the behavioral view, the market tries to set CAPM prices, and violations of the CAPM are due to mispricing.

The conflict between the behavioral irrational pricing story and the rational risk story for the empirical failures of the CAPM leaves us at a timeworn impasse. Fama (1970) emphasizes that the hypothesis that prices properly reflect available information must be tested in the context of a model of expected returns, like the CAPM. Intuitively, to test whether prices are rational, one must take a stand on what the market is trying to do in setting prices—that is, what is risk and what is the relation between expected return and risk? When tests reject the CAPM, one cannot say whether the problem is its assumption that prices are rational (the behavioral view) or violations of other assumptions that are also necessary to produce the CAPM (our position).

Fortunately, for some applications, the way one uses the three-factor model does not depend on one's view about whether its average return premiums are the rational result of underlying state variable risks, the result of irrational investor behavior or sample specific results of chance. For example, when measuring the response of stock prices to new information or when evaluating the performance of managed portfolios, one wants to account for known patterns in returns and average returns for the period examined, whatever their source. Similarly, when estimating the cost of equity capital, one might be unconcerned with whether expected return premiums are rational or irrational since they are in either case part of the opportunity cost of equity capital (Stein, 1996). But the cost of capital is forward looking, so if the premiums are sample specific they are irrelevant.

The three-factor model is hardly a panacea. Its most serious problem is the momentum effect of Jegadeesh and Titman (1993). Stocks that do well relative to the market over the last three to twelve months tend to continue to do well for the next few months, and stocks that do poorly continue to do poorly. This momentum effect is distinct from the value effect captured by book-to-market equity and other price ratios. Moreover, the momentum effect is left unexplained by the three-factor model, as well as by the CAPM. Following Carhart (1997), one response is to add a momentum factor (the difference between the returns on diversified portfolios of short-term winners and losers) to the three-factor model. This step is again legitimate in applications where the goal is to abstract from known patterns in average returns to uncover information-specific or manager-specific effects. But since the momentum effect is short-lived, it is largely irrelevant for estimates of the cost of equity capital.

Another strand of research points to problems in both the three-factor model and the CAPM. Frankel and Lee (1998), Dechow, Hutton and Sloan (1999), Piotroski (2000) and others show that in portfolios formed on price ratios like

book-to-market equity, stocks with higher expected cash flows have higher average returns that are not captured by the three-factor model or the CAPM. The authors interpret their results as evidence that stock prices are irrational, in the sense that they do not reflect available information about expected profitability.

In truth, however, one can't tell whether the problem is bad pricing or a bad asset pricing model. A stock's price can always be expressed as the present value of expected future cash flows discounted at the expected return on the stock (Campbell and Shiller, 1989; Vuolteenaho, 2002). It follows that if two stocks have the same price, the one with higher expected cash flows must have a higher expected return. This holds true whether pricing is rational or irrational. Thus, when one observes a positive relation between expected cash flows and expected returns that is left unexplained by the CAPM or the three-factor model, one can't tell whether it is the result of irrational pricing or a misspecified asset pricing model.

The Market Proxy Problem

Roll (1977) argues that the CAPM has never been tested and probably never will be. The problem is that the market portfolio at the heart of the model is theoretically and empirically elusive. It is not theoretically clear which assets (for example, human capital) can legitimately be excluded from the market portfolio, and data availability substantially limits the assets that are included. As a result, tests of the CAPM are forced to use proxies for the market portfolio, in effect testing whether the proxies are on the minimum variance frontier. Roll argues that because the tests use proxies, not the true market portfolio, we learn nothing about the CAPM.

We are more pragmatic. The relation between expected return and market beta of the CAPM is just the minimum variance condition that holds in any efficient portfolio, applied to the market portfolio. Thus, if we can find a market proxy that is on the minimum variance frontier, it can be used to describe differences in expected returns, and we would be happy to use it for this purpose. The strong rejections of the CAPM described above, however, say that researchers have not uncovered a reasonable market proxy that is close to the minimum variance frontier. If researchers are constrained to reasonable proxies, we doubt they ever will.

Our pessimism is fueled by several empirical results. Stambaugh (1982) tests the CAPM using a range of market portfolios that include, in addition to U.S. common stocks, corporate and government bonds, preferred stocks, real estate and other consumer durables. He finds that tests of the CAPM are not sensitive to expanding the market proxy beyond common stocks, basically because the volatility of expanded market returns is dominated by the volatility of stock returns.

One need not be convinced by Stambaugh's (1982) results since his market proxies are limited to U.S. assets. If international capital markets are open and asset prices conform to an international version of the CAPM, the market portfolio

should include international assets. Fama and French (1998) find, however, that betas for a global stock market portfolio cannot explain the high average returns observed around the world on stocks with high book-to-market or high earnings-price ratios.

A major problem for the CAPM is that portfolios formed by sorting stocks on price ratios produce a wide range of average returns, but the average returns are not positively related to market betas (Lakonishok, Shleifer and Vishny, 1994; Fama and French, 1996, 1998). The problem is illustrated in Figure 3, which shows average returns and betas (calculated with respect to the CRSP value-weight portfolio of NYSE, AMEX and NASDAQ stocks) for July 1963 to December 2003 for ten portfolios of U.S. stocks formed annually on sorted values of the book-to-market equity ratio (B/M).[6]

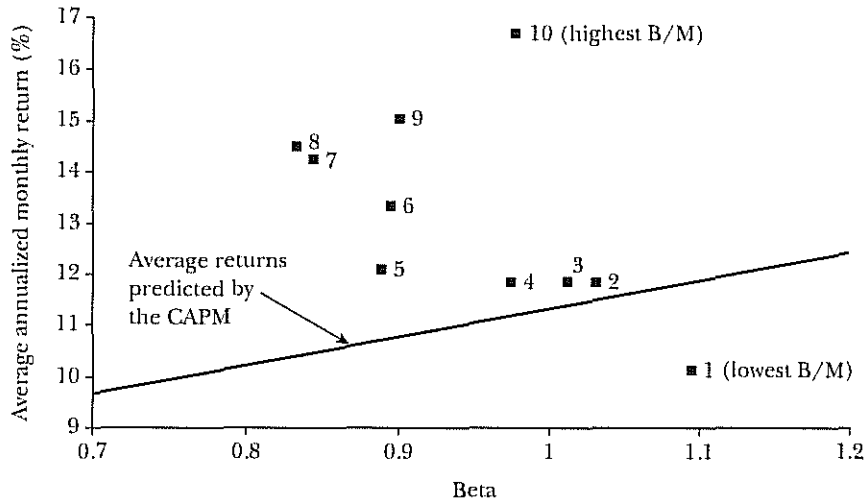
Average returns on the B/M portfolios increase almost monotonically, from 10.1 percent per year for the lowest B/M group (portfolio 1) to an impressive 16.7 percent for the highest (portfolio 10). But the positive relation between beta and average return predicted by the CAPM is notably absent. For example, the portfolio with the lowest book-to-market ratio has the highest beta but the lowest average return. The estimated beta for the portfolio with the highest book-to-market ratio and the highest average return is only 0.98. With an average annualized value of the riskfree interest rate, R_f, of 5.8 percent and an average annualized market premium, $R_M - R_f$, of 11.3 percent, the Sharpe-Lintner CAPM predicts an average return of 11.8 percent for the lowest B/M portfolio and 11.2 percent for the highest, far from the observed values, 10.1 and 16.7 percent. For the Sharpe-Lintner model to "work" on these portfolios, their market betas must change dramatically, from 1.09 to 0.78 for the lowest B/M portfolio and from 0.98 to 1.98 for the highest. We judge it unlikely that alternative proxies for the market portfolio will produce betas and a market premium that can explain the average returns on these portfolios.

It is always possible that researchers will redeem the CAPM by finding a reasonable proxy for the market portfolio that is on the minimum variance frontier. We emphasize, however, that this possibility cannot be used to justify the way the CAPM is currently applied. The problem is that applications typically use the same

[6] Stock return data are from CRSP, and book equity data are from Compustat and the Moody's Industrials, Transportation, Utilities and Financials manuals. Stocks are allocated to ten portfolios at the end of June of each year t (1963 to 2003) using the ratio of book equity for the fiscal year ending in calendar year $t - 1$, divided by market equity at the end of December of $t - 1$. Book equity is the book value of stockholders' equity, plus balance sheet deferred taxes and investment tax credit (if available), minus the book value of preferred stock. Depending on availability, we use the redemption, liquidation or par value (in that order) to estimate the book value of preferred stock. Stockholders' equity is the value reported by Moody's or Compustat, if it is available. If not, we measure stockholders' equity as the book value of common equity plus the par value of preferred stock or the book value of assets minus total liabilities (in that order). The portfolios for year t include NYSE (1963–2003), AMEX (1963–2003) and NASDAQ (1972–2003) stocks with positive book equity in $t - 1$ and market equity (from CRSP) for December of $t - 1$ and June of t. The portfolios exclude securities CRSP does not classify as ordinary common equity. The breakpoints for year t use only securities that are on the NYSE in June of year t.

Figure 3

Average Annualized Monthly Return versus Beta for Value Weight Portfolios Formed on B/M, 1963–2003



market proxies, like the value-weight portfolio of U.S. stocks, that lead to rejections of the model in empirical tests. The contradictions of the CAPM observed when such proxies are used in tests of the model show up as bad estimates of expected returns in applications; for example, estimates of the cost of equity capital that are too low (relative to historical average returns) for small stocks and for stocks with high book-to-market equity ratios. In short, if a market proxy does not work in tests of the CAPM, it does not work in applications.

Conclusions

The version of the CAPM developed by Sharpe (1964) and Lintner (1965) has never been an empirical success. In the early empirical work, the Black (1972) version of the model, which can accommodate a flatter tradeoff of average return for market beta, has some success. But in the late 1970s, research begins to uncover variables like size, various price ratios and momentum that add to the explanation of average returns provided by beta. The problems are serious enough to invalidate most applications of the CAPM.

For example, finance textbooks often recommend using the Sharpe-Lintner CAPM risk-return relation to estimate the cost of equity capital. The prescription is to estimate a stock's market beta and combine it with the risk-free interest rate and the average market risk premium to produce an estimate of the cost of equity. The typical market portfolio in these exercises includes just U.S. common stocks. But empirical work, old and new, tells us that the relation between beta and average return is flatter than predicted by the Sharpe-Lintner version of the CAPM. As a

result, CAPM estimates of the cost of equity for high beta stocks are too high (relative to historical average returns) and estimates for low beta stocks are too low (Friend and Blume, 1970). Similarly, if the high average returns on value stocks (with high book-to-market ratios) imply high expected returns, CAPM cost of equity estimates for such stocks are too low.[7]

The CAPM is also often used to measure the performance of mutual funds and other managed portfolios. The approach, dating to Jensen (1968), is to estimate the CAPM time-series regression for a portfolio and use the intercept (Jensen's alpha) to measure abnormal performance. The problem is that, because of the empirical failings of the CAPM, even passively managed stock portfolios produce abnormal returns if their investment strategies involve tilts toward CAPM problems (Elton, Gruber, Das and Hlavka, 1993). For example, funds that concentrate on low beta stocks, small stocks or value stocks will tend to produce positive abnormal returns relative to the predictions of the Sharpe-Lintner CAPM, even when the fund managers have no special talent for picking winners.

The CAPM, like Markowitz's (1952, 1959) portfolio model on which it is built, is nevertheless a theoretical tour de force. We continue to teach the CAPM as an introduction to the fundamental concepts of portfolio theory and asset pricing, to be built on by more complicated models like Merton's (1973) ICAPM. But we also warn students that despite its seductive simplicity, the CAPM's empirical problems probably invalidate its use in applications.

■ *We gratefully acknowledge the comments of John Cochrane, George Constantinides, Richard Leftwich, Andrei Shleifer, René Stulz and Timothy Taylor.*

[7] The problems are compounded by the large standard errors of estimates of the market premium and of betas for individual stocks, which probably suffice to make CAPM estimates of the cost of equity rather meaningless, even if the CAPM holds (Fama and French, 1997; Pastor and Stambaugh, 1999). For example, using the U.S. Treasury bill rate as the risk-free interest rate and the CRSP value-weight portfolio of publicly traded U.S. common stocks, the average value of the equity premium $R_{Mt} - R_{ft}$ for 1927–2003 is 8.3 percent per year, with a standard error of 2.4 percent. The two standard error range thus runs from 3.5 percent to 13.1 percent, which is sufficient to make most projects appear either profitable or unprofitable. This problem is, however, hardly special to the CAPM. For example, expected returns in all versions of Merton's (1973) ICAPM include a market beta and the expected market premium. Also, as noted earlier the expected values of the size and book-to-market premiums in the Fama-French three-factor model are also estimated with substantial error.

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Response No. MCC-200
Attachment B

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THE JOURNAL OF FINANCE • VOL. XLVII, NO. 2 • JUNE 1992

The Cross-Section of Expected Stock Returns

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ABSTRACT

Two easily measured variables, size and book-to-market equity, combine to capture the cross-sectional variation in average stock returns associated with market β, size, leverage, book-to-market equity, and earnings-price ratios. Moreover, when the tests allow for variation in β that is unrelated to size, the relation between market β and average return is flat, even when β is the only explanatory variable.

THE ASSET-PRICING MODEL OF Sharpe (1964), Lintner (1965), and Black (1972) has long shaped the way academics and practitioners think about average returns and risk. The central prediction of the model is that the market portfolio of invested wealth is mean-variance efficient in the sense of Markowitz (1959). The efficiency of the market portfolio implies that (a) expected returns on securities are a positive linear function of their market βs (the slope in the regression of a security's return on the market's return), and (b) market βs suffice to describe the cross-section of expected returns.

There are several empirical contradictions of the Sharpe-Lintner-Black (SLB) model. The most prominent is the size effect of Banz (1981). He finds that market equity, ME (a stock's price times shares outstanding), adds to the explanation of the cross-section of average returns provided by market βs. Average returns on small (low ME) stocks are too high given their β estimates, and average returns on large stocks are too low.

Another contradiction of the SLB model is the positive relation between leverage and average return documented by Bhandari (1988). It is plausible that leverage is associated with risk and expected return, but in the SLB model, leverage risk should be captured by market β. Bhandari finds, however, that leverage helps explain the cross-section of average stock returns in tests that include size (ME) as well as β.

Stattman (1980) and Rosenberg, Reid, and Lanstein (1985) find that average returns on U.S. stocks are positively related to the ratio of a firm's book value of common equity, BE, to its market value, ME. Chan, Hamao, and Lakonishok (1991) find that book-to-market equity, BE/ME, also has a strong role in explaining the cross-section of average returns on Japanese stocks.

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Finally, Basu (1983) shows that earnings-price ratios (E/P) help explain the cross-section of average returns on U.S. stocks in tests that also include size and market β. Ball (1978) argues that E/P is a catch-all proxy for unnamed factors in expected returns; E/P is likely to be higher (prices are lower relative to earnings) for stocks with higher risks and expected returns, whatever the unnamed sources of risk.

Ball's proxy argument for E/P might also apply to size (ME), leverage, and book-to-market equity. All these variables can be regarded as different ways to scale stock prices, to extract the information in prices about risk and expected returns (Keim (1988)). Moreover, since E/P, ME, leverage, and BE/ME are all scaled versions of price, it is reasonable to expect that some of them are redundant for describing average returns. Our goal is to evaluate the joint roles of market β, size, E/P, leverage, and book-to-market equity in the cross-section of average returns on NYSE, AMEX, and NASDAQ stocks.

Black, Jensen, and Scholes (1972) and Fama and MacBeth (1973) find that, as predicted by the SLB model, there is a positive simple relation between average stock returns and β during the pre-1969 period. Like Reinganum (1981) and Lakonishok and Shapiro (1986), we find that the relation between β and average return disappears during the more recent 1963–1990 period, even when β is used alone to explain average returns. The appendix shows that the simple relation between β and average return is also weak in the 50-year 1941–1990 period. In short, our tests do not support the most basic prediction of the SLB model, that average stock returns are positively related to market βs.

Unlike the simple relation between β and average return, the univariate relations between average return and size, leverage, E/P, and book-to-market equity are strong. In multivariate tests, the negative relation between size and average return i[illegible]ist to the inclusion of other variables. The positive relation between book-to-market equity and average return also persists in competition with other variables. Moreover, although the size effect has attracted more attention, book-to-market equity has a consistently stronger role in average returns. Our bottom-line results are: (a) β does not seem to help explain the cross-section of average stock returns, and (b) the combination of size and book-to-market equity seems to absorb the roles of leverage and E/P in average stock returns, at least during our 1963–1990 sample period.

If assets are priced rationally, our results suggest that stock risks are multidimensional. One dimension of risk is proxied by size, ME. Another dimension of risk is proxied by BE/ME, the ratio of the book value of common equity to its market value.

It is possible that the risk captured by BE/ME is the relative distress factor of Chan and Chen (1991). They postulate that the earning prospects of firms are associated with a risk factor in returns. Firms that the market judges to have poor prospects, signaled here by low stock prices and high ratios of book-to-market equity, have higher expected stock returns (they are penalized with higher costs of capital) than firms with strong prospects. It is

also possible, however, that BE/ME just captures the unraveling (regression toward the mean) of irrational market whims about the prospects of firms.

Whatever the underlying economic causes, our main result is straightforward. Two easily measured variables, size (ME) and book-to-market equity (BE/ME), provide a simple and powerful characterization of the cross-section of average stock returns for the 1963–1990 period.

In the next section we discuss the data and our approach to estimating β. Section II examines the relations between average return and β and between average return and size. Section III examines the roles of E/P, leverage, and book-to-market equity in average returns. In sections IV and V, we summarize, interpret, and discuss applications of the results.

I. Preliminaries

A. Data

We use all nonfinancial firms in the intersection of (a) the NYSE, AMEX, and NASDAQ return files from the Center for Research in Security Prices (CRSP) and (b) the merged COMPUSTAT annual industrial files of income-statement and balance-sheet data, also maintained by CRSP. We exclude financial firms because the high leverage that is normal for these firms probably does not have the same meaning as for nonfinancial firms, where high leverage more likely indicates distress. The CRSP returns cover NYSE and AMEX stocks until 1973 when NASDAQ returns also come on line. The COMPUSTAT data are for 1962–1989. The 1962 start date reflects the fact that book value of common equity (COMPUSTAT item 60), is not generally available prior to 1962. More important, COMPUSTAT data for earlier years have a serious selection [illegible] the pre-1962 data are tilted toward big historically successful firms.

To ensure that the accounting variables are known before the returns they are used to explain, we match the accounting data for all fiscal yearends in calendar year $t-1$ (1962–1989) with the returns for July of year t to June of $t+1$. The 6-month (minimum) gap between fiscal yearend and the return tests is conservative. Earlier work (e.g., Basu (1983)) often assumes that accounting data are available within three months of fiscal yearends. Firms are indeed required to file their 10-K reports with the SEC within 90 days of their fiscal yearends, but on average 19.8% do not comply. In addition, more than 40% of the December fiscal yearend firms that do comply with the 90-day rule file on March 31, and their reports are not made public until April. (See Alford, Jones, and Zmijewski (1992).)

We use a firm's market equity at the end of December of year $t-1$ to compute its book-to-market, leverage, and earnings-price ratios for $t-1$, and we use its market equity for June of year t to measure its size. Thus, to be included in the return tests for July of year t, a firm must have a CRSP stock price for December of year $t-1$ and June of year t. It must also have monthly returns for at least 24 of the 60 months preceding July of year t (for

"pre-ranking" β estimates, discussed below). And the firm must have COMPUSTAT data on total book assets (A), book equity (BE), and earnings (E), for its fiscal year ending in (any month of) calendar year $t-1$.

Our use of December market equity in the E/P, BE/ME, and leverage ratios is objectionable for firms that do not have December fiscal yearends because the accounting variable in the numerator of a ratio is not aligned with the market value in the denominator. Using ME at fiscal yearends is also problematic; then part of the cross-sectional variation of a ratio for a given year is due to market-wide variation in the ratio during the year. For example, if there is a general fall in stock prices during the year, ratios measured early in the year will tend to be lower than ratios measured later. We can report, however, that the use of fiscal-yearend MEs, rather than December MEs, in the accounting ratios has little impact on our return tests.

Finally, the tests mix firms with different fiscal yearends. Since we match accounting data for all fiscal yearends in calendar year $t-1$ with returns for July of t to June of $t+1$, the gap between the accounting data and the matching returns varies across firms. We have done the tests using the smaller sample of firms with December fiscal yearends with similar results.

B. Estimating Market βs

Our asset-pricing tests use the cross-sectional regression approach of Fama and MacBeth (1973). Each month the cross-section of returns on stocks is regressed on variables hypothesized to explain expected returns. The time-series means of the monthly regression slopes then provide standard tests of whether different explanatory variables are on average priced.

Since size, E/P, leverage, and BE/ME are measured precisely for individual stocks, there is no reason to smear the information in these variables by using portfolios in the Fama-MacBeth (FM) regressions. Most previous tests use portfolios because estimates of market βs are more precise for portfolios. Our approach is to estimate βs for portfolios and then assign a portfolio's β to each stock in the portfolio. This allows us to use individual stocks in the FM asset-pricing tests.

B.1. β Estimation: Details

In June of each year, all NYSE stocks on CRSP are sorted by size (ME) to determine the NYSE decile breakpoints for ME. NYSE, AMEX, and NASDAQ stocks that have the required CRSP-COMPUSTAT data are then allocated to 10 size portfolios based on the NYSE breakpoints. (If we used stocks from all three exchanges to determine the ME breakpoints, most portfolios would include only small stocks after 1973, when NASDAQ stocks are added to the sample.)

We form portfolios on size because of the evidence of Chan and Chen (1988) and others that size produces a wide spread of average returns and βs. Chan and Chen use only size portfolios. The problem this creates is that size and the βs of size portfolios are highly correlated (-0.988 in their data), so

asset-pricing tests lack power to separate size from β effects in average returns.

To allow for variation in β that is unrelated to size, we subdivide each size decile into 10 portfolios on the basis of pre-ranking βs for individual stocks. The pre-ranking βs are estimated on 24 to 60 monthly returns (as available) in the 5 years before July of year t. We set the β breakpoints for each size decile using only NYSE stocks that satisfy our COMPUSTAT-CRSP data requirements for year $t-1$. Using NYSE stocks ensures that the β breakpoints are not dominated after 1973 by the many small stocks on NASDAQ. Setting β breakpoints with stocks that satisfy our COMPUSTAT-CRSP data requirements guarantees that there are firms in each of the 100 size-β portfolios.

After assigning firms to the size-β portfolios in June, we calculate the equal-weighted monthly returns on the portfolios for the next 12 months, from July to June. In the end, we have post-ranking monthly returns for July 1963 to December 1990 on 100 portfolios formed on size and pre-ranking βs. We then estimate βs using the full sample (330 months) of post-ranking returns on each of the 100 portfolios, with the CRSP value-weighted portfolio of NYSE, AMEX, and (after 1972) NASDAQ stocks used as the proxy for the market. We have also estimated βs using the value-weighted or the equal-weighted portfolio of NYSE stocks as the proxy for the market. These βs produce inferences on the role of β in average returns like those reported below.

We estimate β as the sum of the slopes in the regression of the return on a portfolio on the current and prior month's market return. (An additional lead and lag of the market have little effect on these sum βs.) The sum βs are meant to adjust for nonsynchronous trading (Dimson (1979)). Fowler and Rorke (1983) show that sum βs are biased when the market return is autocorrelated. The 1st- and 2nd-order autocorrelations of the monthly market returns for July 1963 to December 1990 are 0.06 and -0.05, both about 1 standard error from 0. If the Fowler-Rorke corrections are used, they lead to trivial changes in the βs. We stick with the simpler sum βs. Appendix Table AI shows that using sum βs produces large increases in the βs of the smallest ME portfolios and small declines in the βs of the largest ME portfolios.

Chan and Chen (1988) show that full-period β estimates for portfolios can work well in tests of the SLB model, even if the true βs of the portfolios vary through time, if the variation in the βs is proportional,

$$\beta_{jt} - \beta_j = k_t(\beta_j - \beta), \tag{1}$$

where β_{jt} is the true β for portfolio j at time t, β_j is the mean of β_{jt} across t, and β is the mean of the β_j. The Appendix argues that (1) is a good approximation for the variation through time in the true βs of portfolios (j) formed on size and β. For diehard β fans, sure to be skeptical of our results on the weak role of β in average stock returns, we can also report that the results stand up to robustness checks that use 5-year pre-ranking βs, or 5-year post-ranking βs, instead of the full-period post-ranking βs.

We allocate the full-period post-ranking β of a size-β portfolio to each stock in the portfolio. These are the βs that will be used in the Fama-MacBeth cross-sectional regressions for individual stocks. We judge that the precision of the full-period post-ranking portfolio βs, relative to the imprecise β estimates that would be obtained for individual stocks, more than makes up for the fact that true βs are not the same for all stocks in a portfolio. And note that assigning full-period portfolio βs to stocks does not mean that a stock's β is constant. A stock can move across portfolios with year-to-year changes in the stock's size (ME) and in the estimates of its β for the preceding 5 years.

B.2. β Estimates

Table I shows that forming portfolios on size and pre-ranking βs, rather than on size alone, magnifies the range of full-period post-ranking βs. Sorted on size alone, the post-ranking βs range from 1.44 for the smallest ME portfolio to 0.92 for the largest. This spread of βs across the 10 size deciles is smaller than the spread of post-ranking βs produced by the β sort of *any* size decile. For example, the post-ranking βs for the 10 portfolios in the smallest size decile range from 1.05 to 1.79. Across all 100 size-β portfolios, the post-ranking βs range from 0.53 to 1.79, a spread 2.4 times the spread, 0.52, obtained with size portfolios alone.

Two other facts about the βs are important. First, in each size decile the post-ranking βs closely reproduce the ordering of the pre-ranking βs. We take this to be evidence that the pre-ranking β sort captures the ordering of true post-ranking βs. (The appendix gives more evidence on this important issue.) Second, the β sort is not a refined size sort. In any size decile, the average values of ln(ME) are similar across the β-sorted portfolios. Thus the pre-ranking β sort achieves its goal. It produces strong variation in post-ranking βs that is unrelated to size. This is important in allowing our tests to distinguish between β and size effects in average returns.

II. β and Size

The Sharpe-Lintner-Black (SLB) model plays an important role in the way academics and practitioners think about risk and the relation between risk and expected return. We show next that when common stock portfolios are formed on size alone, there seems to be evidence for the model's central prediction: average return is positively related to β. The βs of size portfolios are, however, almost perfectly correlated with size, so tests on size portfolios are unable to disentangle β and size effects in average returns. Allowing for variation in β that is unrelated to size breaks the logjam, but at the expense of β. Thus, when we subdivide size portfolios on the basis of pre-ranking βs, we find a strong relation between average return and size, but no relation between average return and β.

A. Informal Tests

Table II shows post-ranking average returns for July 1963 to December 1990 for portfolios formed from one-dimensional sorts of stocks on size or β. The portfolios are formed at the end of June each year and their equal-weighted returns are calculated for the next 12 months. We use returns for July to June to match the returns in later tests that use the accounting data. When we sort on just size or 5-year pre-ranking βs, we form 12 portfolios. The middle 8 cover deciles of size or β. The 4 extreme portfolios (1A, 1B, 10A, and 10B) split the bottom and top deciles in half.

Table II shows that when portfolios are formed on size alone, we observe the familiar strong negative relation between size and average return (Banz (1981)), and a strong positive relation between average return and β. Average returns fall from 1.64% per month for the smallest ME portfolio to 0.90% for the largest. Post-ranking βs also decline across the 12 size portfolios, from 1.44 for portfolio 1A to 0.90 for portfolio 10B. Thus, a simple size sort seems to support the SLB prediction of a positive relation between β and average return. But the evidence is muddied by the tight relation between size and the βs of size portfolios.

The portfolios formed on the basis of the ranked market βs of stocks in Table II produce a wider range of βs (from 0.81 for portfolio 1A to 1.73 for 10B) than the portfolios formed on size. Unlike the size portfolios, the β-sorted portfolios do not support the SLB model. There is little spread in average returns across the β portfolios, and there is no obvious relation between β and average returns. For example, although the two extreme portfolios, 1A and 10B, have much different βs, they have nearly identical average returns (1.20% and 1.18% per month). These results for 1963–1990 confirm Reinganum's (1981) evidence that for β-sorted portfolios, there is no relation between average return and β during the 1964–1979 period.

The 100 portfolios formed on size and then pre-ranking β in Table I clarify the contradictory evidence on the relation between β and average return produced by portfolios formed on size or β alone. Specifically, the two-pass sort gives a clearer picture of the separate roles of size and β in average returns. Contrary to the central prediction of the SLB model, the second-pass β sort produces little variation in average returns. Although the post-ranking βs in Table I increase strongly in each size decile, average returns are flat or show a slight tendency to decline. In contrast, within the columns of the average return and β matrices of Table I, average returns and βs decrease with increasing size.

The two-pass sort on size and β in Table I says that variation in β that is tied to size is positively related to average return, but variation in β unrelated to size is not compensated in the average returns of 1963–1990. The proper inference seems to be that there is a relation between size and average return, but controlling for size, there is no relation between β and average return. The regressions that follow confirm this conclusion, and they produce another that is stronger. The regressions show that when one allows

Table I

Average Returns, Post-Ranking βs and Average Size For Portfolios Formed on Size and then β: Stocks Sorted on ME (Down) then Pre-Ranking β (Across): July 1963 to December 1990

Portfolios are formed yearly. The breakpoints for the size (ME, price times shares outstanding) deciles are determined in June of year t (t = 1963–1990) using all NYSE stocks on CRSP. All NYSE, AMEX, and NASDAQ stocks that meet the CRSP-COMPUSTAT data requirements are allocated to the 10 size portfolios using the NYSE breakpoints. Each size decile is subdivided into 10 β portfolios using pre-ranking βs of individual stocks, estimated with 2 to 5 years of monthly returns (as available) ending in June of year t. We use only NYSE stocks that meet the CRSP-COMPUSTAT data requirements to establish the β breakpoints. The equal-weighted monthly returns on the resulting 100 portfolios are then calculated for July of year t to June of year $t + 1$.

The post-ranking βs use the full (July 1963 December 1990) sample of post-ranking returns for each portfolio. The pre- and post-ranking βs (here and in all other tables) are the sum of the slopes from a regression of monthly returns on the current and prior month's returns on the value-weighted portfolio of NYSE, AMEX, and (after 1972) NASDAQ stocks. The average return is the time-series average of the monthly equal-weighted portfolio returns, in percent. The average size of a portfolio is the time-series average of monthly averages of ln(ME) for stocks in the portfolio at the end of June of each year, with ME denominated in millions of dollars.

The average number of stocks per month for the size-β portfolios in the smallest size decile varies from 70 to 177. The average number of stocks for the size-β portfolios in size deciles 2 and 3 is between 15 and 41, and the average number for the largest 7 size deciles is between 11 and 22.

The All column shows statistics for equal-weighted size-decile (ME) portfolios. The All row shows statistics for equal-weighted portfolios of the stocks in each β group.

	All	Low-β	β-2	β-3	β-4	β-5	β-6	β-7	β-8	β-9	High-β
					Panel A: Average Monthly Returns (in Percent)						
All	1.25	1.34	1.29	1.36	1.31	1.33	1.28	1.24	1.21	1.25	1.14
Small-ME	1.52	1.71	1.57	1.79	1.61	1.50	1.50	1.37	1.63	1.50	1.42
ME-2	1.29	1.25	1.42	1.36	1.39	1.65	1.61	1.37	1.31	1.34	1.11
ME-3	1.24	1.12	1.31	1.17	1.70	1.29	1.10	1.31	1.36	1.26	0.76
ME-4	1.25	1.27	1.13	1.54	1.06	1.34	1.06	1.41	1.17	1.35	0.98
ME-5	1.29	1.34	1.42	1.39	1.48	1.42	1.18	1.13	1.27	1.18	1.08
ME-6	1.17	1.08	1.53	1.27	1.15	1.20	1.21	1.18	1.04	1.07	1.02
ME-7	1.07	0.95	1.21	1.26	1.09	1.18	1.11	1.24	0.62	1.32	0.76
ME-8	1.10	1.09	1.05	1.37	1.20	1.27	0.98	1.18	1.02	1.01	0.94
ME-9	0.95	0.98	0.88	1.02	1.14	1.07	1.23	0.94	0.82	0.88	0.59
Large-ME	0.89	1.01	0.93	1.10	0.94	0.93	0.89	1.03	0.71	0.74	0.56

Table I—*Continued*

	All	Low-β	β-2	β-3	β-4	β-5	β-6	β-7	β-8	β-9	High-β
Panel B: Post-Ranking βs											
All		0.87	0.99	1.09	1.16	1.26	1.29	1.35	1.45	1.52	1.72
Small-ME	1.44	1.05	1.18	1.28	1.32	1.40	1.40	1.49	1.61	1.64	1.79
ME-2	1.39	0.91	1.15	1.17	1.24	1.36	1.41	1.43	1.50	1.66	1.76
ME-3	1.35	0.97	1.13	1.13	1.21	1.26	1.28	1.39	1.50	1.51	1.75
ME-4	1.34	0.78	1.03	1.17	1.16	1.29	1.37	1.46	1.51	1.64	1.71
ME-5	1.25	0.66	0.85	1.12	1.15	1.16	1.26	1.30	1.43	1.59	1.68
ME-6	1.23	0.61	0.78	1.05	1.16	1.22	1.28	1.36	1.46	1.49	1.70
ME-7	1.17	0.57	0.92	1.01	1.11	1.14	1.26	1.24	1.39	1.34	1.60
ME-8	1.09	0.53	0.74	0.94	1.02	1.13	1.12	1.18	1.26	1.35	1.52
ME-9	1.03	0.58	0.74	0.80	0.95	1.06	1.15	1.14	1.21	1.22	1.42
Large-ME	0.92	0.57	0.71	0.78	0.89	0.95	0.92	1.02	1.01	1.11	1.32
Panel C: Average Size (ln(ME))											
All	4.11	3.86	4.26	4.33	4.41	4.27	4.32	4.26	4.19	4.03	3.77
Small-ME	2.24	2.12	2.27	2.30	2.30	2.28	2.29	2.30	2.32	2.25	2.15
ME-2	3.63	3.65	3.68	3.70	3.72	3.69	3.70	3.69	3.69	3.70	3.68
ME-3	4.10	4.14	4.18	4.12	4.15	4.16	4.16	4.18	4.14	4.15	4.15
ME-4	4.50	4.53	4.53	4.57	4.54	4.56	4.55	4.52	4.58	4.52	4.56
ME-5	4.89	4.91	4.91	4.93	4.95	4.93	4.92	4.93	4.92	4.92	4.95
ME-6	5.30	5.30	5.33	5.34	5.34	5.33	5.33	5.33	5.33	5.34	5.36
ME-7	5.73	5.73	5.75	5.77	5.76	5.73	5.77	5.77	5.76	5.72	5.76
ME-8	6.24	6.26	6.27	6.26	6.24	6.24	6.27	6.24	6.24	6.24	6.26
ME-9	6.82	6.82	6.84	6.82	6.82	6.81	6.81	6.81	6.81	6.80	6.83
Large-ME	7.93	7.94	8.04	8.10	8.04	8.02	8.02	7.94	7.80	7.75	7.62

Table II

Properties of Portfolios Formed on Size or Pre-Ranking β: July 1963 to December 1990

At the end of June of each year t, 12 portfolios are formed on the basis of ranked values of size (ME) or pre-ranking β. The pre-ranking βs use 2 to 5 years (as available) of monthly returns ending in June of t. Portfolios 2–9 cover deciles of the ranking variables. The bottom and top 2 portfolios (1A, 1B, 10A, and 10B) split the bottom and top deciles in half. The breakpoints for the ME portfolios are based on ranked values of ME for all NYSE stocks on CRSP. NYSE breakpoints for pre-ranking βs are also used to form the β portfolios. NYSE, AMEX, and NASDAQ stocks are then allocated to the size or β portfolios using the NYSE breakpoints. We calculate each portfolio's monthly equal-weighted return for July of year t to June of year $t+1$, and then reform the portfolios in June of $t+1$.

BE is the book value of common equity plus balance-sheet deferred taxes, A is total book assets, and E is earnings (income before extraordinary items, plus income statement deferred taxes, minus preferred dividends). BE, A, and E are for each firm's latest fiscal year ending in calendar year $t-1$. The accounting ratios are measured using market equity ME in December of year $t-1$. Firm size ln(ME) is measured in June of year t, with ME denominated in millions of dollars.

The average return is the time-series average of the monthly equal-weighted portfolio returns, in percent. ln(ME), ln(BE/ME), ln(A/ME), ln(A/BE), E/P, and E/P dummy are the time-series averages of the monthly average values of these variables in each portfolio. Since the E/P dummy is 0 when earnings are positive, and 1 when earnings are negative, E/P dummy gives the average proportion of stocks with negative earnings in each portfolio.

β is the time-series average of the monthly portfolio βs. Stocks are assigned the post-ranking β of the size-β portfolio they are in at the end of June of year t (Table I). These individual-firm βs are averaged to compute the monthly βs for each portfolio for July of year t to June of year $t+1$.

Firms is the average number of stocks in the portfolio each month.

	1A	1B	2	3	4	5	6	7	8	9	10A	10B
Panel A: Portfolios Formed on Size												
Return	1.64	1.16	1.29	1.24	1.25	1.29	1.17	1.07	1.10	0.95	0.88	0.90
β	1.44	1.44	1.39	1.34	1.33	1.24	1.22	1.16	1.08	1.02	0.95	0.90
ln(ME)	1.98	3.18	3.63	4.10	4.50	4.89	5.30	5.73	6.24	6.82	7.39	8.44
ln(BE/ME)	−0.01	−0.21	−0.23	−0.26	−0.32	−0.36	−0.36	−0.44	−0.40	−0.42	−0.51	−0.65
ln(A/ME)	0.73	0.50	0.46	0.43	0.37	0.32	0.32	0.24	0.29	0.27	0.17	−0.03
ln(A/BE)	0.75	0.71	0.69	0.69	0.68	0.67	0.68	0.67	0.69	0.70	0.68	0.62
E/P dummy	0.26	0.14	0.11	0.09	0.06	0.04	0.04	0.03	0.03	0.02	0.02	0.01
E(+)/P	0.09	0.10	0.10	0.10	0.10	0.10	0.10	0.10	0.10	0.10	0.09	0.09
Firms	772	189	236	176	144	140	128	125	119	114	60	64

Table II—*Continued*

	1A	1B	2	3	4	5	6	7	8	9	10A	10B
						Panel B: Portfolios Formed on Pre-Ranking β						
Return	1.20	1.20	1.32	1.26	1.31	1.30	1.30	1.23	1.23	1.33	1.34	1.18
β	0.81	0.79	0.92	1.04	1.13	1.19	1.26	1.32	1.41	1.52	1.63	1.73
ln(ME)	4.21	4.86	4.75	4.68	4.59	4.48	4.36	4.25	3.97	3.78	3.52	3.15
ln(BE/ME)	−0.18	−0.13	−0.22	−0.21	−0.23	−0.22	−0.22	−0.25	−0.23	−0.27	−0.31	−0.50
ln(A/ME)	0.60	0.66	0.49	0.45	0.42	0.42	0.45	0.42	0.47	0.46	0.46	0.31
ln(A/BE)	0.78	0.79	0.71	0.66	0.64	0.65	0.67	0.67	0.70	0.73	0.77	0.81
E/P dummy	0.12	0.06	0.09	0.09	0.08	0.09	0.10	0.12	0.12	0.14	0.17	0.23
E(+)/P	0.11	0.12	0.10	0.10	0.10	0.10	0.10	0.09	0.10	0.09	0.09	0.08
Firms	116	80	185	181	179	182	185	205	227	267	165	291

for variation in β that is unrelated to size, the relation between β and average return is flat, even when β is the only explanatory variable.

B. Fama-MacBeth Regressions

Table III shows time-series averages of the slopes from the month-by-month Fama-MacBeth (FM) regressions of the cross-section of stock returns on size, β, and the other variables (leverage, E/P, and book-to-market equity) used to explain average returns. The average slopes provide standard FM tests for determining which explanatory variables on average have non-zero expected premiums during the July 1963 to December 1990 period.

Like the average returns in Tables I and II, the regressions in Table III say that size, ln(ME), helps explain the cross-section of average stock returns. The average slope from the monthly regressions of returns on size alone is -0.15%, with a t-statistic of -2.58. This reliable negative relation persists no matter which other explanatory variables are in the regressions; the average slopes on ln(ME) are always close to or more than 2 standard errors from 0. The size effect (smaller stocks have higher average returns) is thus robust in the 1963–1990 returns on NYSE, AMEX, and NASDAQ stocks.

In contrast to the consistent explanatory power of size, the FM regressions show that market β does not help explain average stock returns for 1963–1990. In a shot straight at the heart of the SLB model, the average slope from the regressions of returns on β alone in Table III is 0.15% per month and only 0.46 standard errors from 0. In the regressions of returns on size and β, size has explanatory power (an average slope -3.41 standard errors from 0), but the average slope for β is negative and only 1.21 standard errors from 0. Lakonishok and Shapiro (1986) get similar results for NYSE stocks for 1962–1981. We can also report that β shows no power to explain average returns (the average slopes are typically less than 1 standard error from 0) in FM regressions that use various combinations of β with size, book-to-market equity, leverage, and E/P.

C. Can β Be Saved?

What explains the poor results for β? One possibility is that other explanatory variables are correlated with true βs, and this obscures the relation between average returns and measured βs. But this line of attack cannot explain why β has no power when used alone to explain average returns. Moreover, leverage, book-to-market equity, and E/P do not seem to be good proxies for β. The averages of the monthly cross-sectional correlations between β and the values of these variables for individual stocks are all within 0.15 of 0.

Another hypothesis is that, as predicted by the SLB model, there is a positive relation between β and average return, but the relation is obscured by noise in the β estimates. However, our full-period post-ranking βs do not seem to be imprecise. Most of the standard errors of the βs (not shown) are

Table III

Average Slopes (t-Statistics) from Month-by-Month Regressions of Stock Returns on β, Size, Book-to-Market Equity, Leverage, and E/P: July 1963 to December 1990

Stocks are assigned the post-ranking β of the size-β portfolio they are in at the end of June of year t (Table I). BE is the book value of common equity plus balance-sheet deferred taxes, A is total book assets, and E is earnings (income before extraordinary items, plus income-statement deferred taxes, minus preferred dividends). BE, A, and E are for each firm's latest fiscal year ending in calendar year $t-1$. The accounting ratios are measured using market equity ME in December of year $t-1$. Firm size ln(ME) is measured in June of year t. In the regressions, these values of the explanatory variables for individual stocks are matched with CRSP returns for the months from July of year t to June of year $t+1$. The gap between the accounting data and the returns ensures that the accounting data are available prior to the returns. If earnings are positive, E(+)/P is the ratio of total earnings to market equity and E/P dummy is 0. If earnings are negative, E(+)/P is 0 and E/P dummy is 1.

The average slope is the time-series average of the monthly regression slopes for July 1963 to December 1990, and the t-statistic is the average slope divided by its time-series standard error.

On average, there are 2267 stocks in the monthly regressions. To avoid giving extreme observations heavy weight in the regressions, the smallest and largest 0.5% of the observations on E(+)/P, BE/ME, A/ME, and A/BE are set equal to the next largest or smallest values of the ratios (the 0.005 and 0.995 fractiles). This has no effect on inferences.

β	ln(ME)	ln(BE/ME)	ln(A/ME)	ln(A/BE)	E/P Dummy	E(+)/P
0.15						
(0.46)						
	−0.15					
	(−2.58)					
−0.37	−0.17					
(−1.21)	(−3.41)					
		0.50				
		(5.71)				
			0.50	−0.57		
			(5.69)	(−5.34)		
					0.57	4.72
					(2.28)	(4.57)
	−0.11	0.35				
	(−1.99)	(4.44)				
	−0.11		0.35	−0.50		
	(−2.06)		(4.32)	(−4.56)		
	−0.16				0.06	2.99
	(−3.06)				(0.38)	(3.04)
	−0.13	0.33			−0.14	0.87
	(−2.47)	(4.46)			(−0.90)	(1.23)
	−0.13		0.32	−0.46	−0.08	1.15
	(−2.47)		(4.28)	(−4.45)	(−0.56)	(1.57)

0.05 or less, only 1 is greater than 0.1, and the standard errors are small relative to the range of the βs (0.53 to 1.79).

The β-sorted portfolios in Tables I and II also provide strong evidence against the β-measurement-error story. When portfolios are formed on pre-ranking βs alone (Table II), the post-ranking βs for the portfolios almost perfectly reproduce the ordering of the pre-ranking βs. Only the β for portfolio 1B is out of line, and only by 0.02. Similarly, when portfolios are formed on size and then pre-ranking βs (Table I), the post-ranking βs in each size decile closely reproduce the ordering of the pre-ranking βs.

The correspondence between the ordering of the pre-ranking and post-ranking βs for the β-sorted portfolios in Tables I and II is evidence that the post-ranking βs are informative about the ordering of the true βs. The problem for the SLB model is that there is no similar ordering in the average returns on the β-sorted portfolios. Whether one looks at portfolios sorted on β alone (Table II) or on size and then β (Table I), average returns are flat (Table II) or decline slightly (Table I) as the post-ranking βs increase.

Our evidence on the robustness of the size effect and the absence of a relation between β and average return is so contrary to the SLB model that it behooves us to examine whether the results are special to 1963–1990. The appendix shows that NYSE returns for 1941–1990 behave like the NYSE, AMEX, and NASDAQ returns for 1963–1990; there is a reliable size effect over the full 50-year period, but little relation between β and average return. Interestingly, there is a reliable simple relation between β and average return during the 1941–1965 period. These 25 years are a major part of the samples in the early studies of the SLB model of Black, Jensen, and Scholes (1972) and Fama and MacBeth (1973). Even for the 1941–1965 period, however, the relation between β and average return disappears when we control for size.

III. Book-to-Market Equity, E/P, and Leverage

Tables I to III say that there is a strong relation between the average returns on stocks and size, but there is no reliable relation between average returns and β. In this section we show that there is also a strong cross-sectional relation between average returns and book-to-market equity. If anything, this book-to-market effect is more powerful than the size effect. We also find that the combination of size and book-to-market equity absorbs the apparent roles of leverage and E/P in average stock returns.

A. Average Returns

Table IV shows average returns for July 1963 to December 1990 for portfolios formed on ranked values of book-to-market equity (BE/ME) or earnings-price ratio (E/P). The BE/ME and E/P portfolios in Table IV are formed in the same general way (one-dimensional yearly sorts) as the size and β portfolios in Table II. (See the tables for details.)

The relation between average return and E/P has a familiar U-shape (e.g., Jaffe, Keim, and Westerfield (1989) for U.S. data, and Chan, Hamao, and Lakonishok (1991) for Japan). Average returns decline from 1.46% per month for the negative E/P portfolio to 0.93% for the firms in portfolio 1B that have low but positive E/P. Average returns then increase monotonically, reaching 1.72% per month for the highest E/P portfolio.

The more striking evidence in Table IV is the strong positive relation between average return and book-to-market equity. Average returns rise from 0.30% for the lowest BE/ME portfolio to 1.83% for the highest, a difference of 1.53% per month. This spread is twice as large as the difference of 0.74% between the average monthly returns on the smallest and largest size portfolios in Table II. Note also that the strong relation between book-to-market equity and average return is unlikely to be a β effect in disguise; Table IV shows that post-ranking market βs vary little across portfolios formed on ranked values of BE/ME.

On average, only about 50 (out of 2317) firms per year have negative book equity, BE. The negative BE firms are mostly concentrated in the last 14 years of the sample, 1976–1989, and we do not include them in the tests. We can report, however, that average returns for negative BE firms are high, like the average returns of high BE/ME firms. Negative BE (which results from persistently negative earnings) and high BE/ME (which typically means that stock prices have fallen) are both signals of poor earning prospects. The similar average returns of negative and high BE/ME firms are thus consistent with the hypothesis that book-to-market equity captures cross-sectional variation in average returns that is related to relative distress.

B. Fama-MacBeth Regressions

B.1. BE/ME

The FM regressions in Table III confirm the importance of book-to-market equity in explaining the cross-section of average stock returns. The average slope from the monthly regressions of returns on ln(BE/ME) alone is 0.50%, with a t-statistic of 5.71. This book-to-market relation is stronger than the size effect, which produces a t-statistic of -2.58 in the regressions of returns on ln(ME) alone. But book-to-market equity does not replace size in explaining average returns. When both ln(ME) and ln(BE/ME) are included in the regressions, the average size slope is still -1.99 standard errors from 0; the book-to-market slope is an impressive 4.44 standard errors from 0.

B.2. Leverage

The FM regressions that explain returns with leverage variables provide interesting insight into the relation between book-to-market equity and average return. We use two leverage variables, the ratio of book assets to market equity, A/ME, and the ratio of book assets to book equity, A/BE. We interpret A/ME as a measure of market leverage, while A/BE is a measure

Table IV

Properties of Portfolios Formed on Book-to-Market Equity (BE/ME) and Earnings-Price Ratio (E/P): July 1963 to December 1990

At the end of each year $t-1$, 12 portfolios are formed on the basis of ranked values of BE/ME or E/P. Portfolios 2–9 cover deciles of the ranking variables. The bottom and top 2 portfolios (1A, 1B, 10A, and 10B) split the bottom and top deciles in half. For E/P, there are 13 portfolios; portfolio 0 is stocks with negative E/P. Since BE/ME and E/P are not strongly related to exchange listing, their portfolio breakpoints are determined on the basis of the ranked values of the variables for all stocks that satisfy the CRSP-COMPUSTAT data requirements. BE is the book value of common equity plus balance-sheet deferred taxes, A is total book assets, and E is earnings (income before extraordinary items, plus income-statement deferred taxes, minus preferred dividends). BE, A, and E are for each firm's latest fiscal year ending in calendar year $t-1$. The accounting ratios are measured using market equity ME in December of year $t-1$. Firm size ln(ME) is measured in June of year t, with ME denominated in millions of dollars. We calculate each portfolio's monthly equal-weighted return for July of year t to June of year $t+1$, and then reform the portfolios at the end of year t.

Return is the time-series average of the monthly equal-weighted portfolio returns (in percent). ln(ME), ln(BE/ME), ln(A/ME), ln(A/BE), E(+)/P, and E/P dummy are the time-series averages of the monthly average values of these variables in each portfolio. Since the E/P dummy is 0 when earnings are positive, and 1 when earnings are negative, E/P dummy gives the average proportion of stocks with negative earnings in each portfolio.

β is the time-series average of the monthly portfolio βs. Stocks are assigned the post-ranking β of the size-β portfolio they are in at the end of June of year t (Table I). These individual-firm βs are averaged to compute the monthly βs for each portfolio for July of year t to June of year $t+1$.

Firms is the average number of stocks in the portfolio each month.

Portfolio	0	1A	1B	2	3	4	5	6	7	8	9	10A	10B
Panel A: Stocks Sorted on Book-to-Market Equity (BE/ME)													
Return		0.30	0.67	0.87	0.97	1.04	1.17	1.30	1.44	1.50	1.59	1.92	1.83
β		1.36	1.34	1.32	1.30	1.28	1.27	1.27	1.27	1.27	1.29	1.33	1.35
ln(ME)		4.53	4.67	4.69	4.56	4.47	4.38	4.23	4.06	3.85	3.51	3.06	2.65
ln(BE/ME)		−2.22	−1.51	−1.09	−0.75	−0.51	−0.32	−0.14	0.03	0.21	0.42	0.66	1.02
ln(A/ME)		−1.24	−0.79	−0.40	−0.05	0.20	0.40	0.56	0.71	0.91	1.12	1.35	1.75
ln(A/BE)		0.94	0.71	0.68	0.70	0.71	0.71	0.70	0.68	0.70	0.70	0.70	0.73
E/P dummy		0.29	0.15	0.10	0.08	0.08	0.08	0.09	0.09	0.11	0.15	0.22	0.36
E(+)/P		0.03	0.04	0.06	0.08	0.09	0.10	0.11	0.11	0.12	0.12	0.11	0.10
Firms		89	98	209	222	226	230	235	237	239	239	120	117

Table IV—*Continued*

Portfolio	0	1A	1B	2	3	4	5	6	7	8	9	10A	10B
Panel B: Stocks Sorted on Earnings-Price Ratio (E/P)													
Return	1.46	1.04	0.93	0.94	1.03	1.18	1.22	1.33	1.42	1.46	1.57	1.74	1.72
β	1.47	1.40	1.35	1.31	1.28	1.26	1.25	1.26	1.24	1.23	1.24	1.28	1.31
ln(ME)	2.48	3.64	4.33	4.61	4.64	4.63	4.58	4.49	4.37	4.28	4.07	3.82	3.52
ln(BE/ME)	−0.10	−0.76	−0.91	−0.79	−0.61	−0.47	−0.33	−0.21	−0.08	0.02	0.15	0.26	0.40
ln(A/ME)	0.90	−0.05	−0.27	−0.16	0.03	0.18	0.31	0.44	0.58	0.70	0.85	1.01	1.25
ln(A/BE)	0.99	0.70	0.63	0.63	0.64	0.65	0.64	0.65	0.66	0.68	0.71	0.75	0.86
E/P dummy	1.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00
E(+)/P	0.00	0.01	0.03	0.05	0.06	0.08	0.09	0.11	0.12	0.14	0.16	0.20	0.28
Firms	355	88	90	182	190	193	196	194	197	195	195	95	91

of book leverage. The regressions use the natural logs of the leverage ratios, ln(A/ME) and ln(A/BE), because preliminary tests indicated that logs are a good functional form for capturing leverage effects in average returns. Using logs also leads to a simple interpretation of the relation between the roles of leverage and book-to-market equity in average returns.

The FM regressions of returns on the leverage variables (Table III) pose a bit of a puzzle. The two leverage variables are related to average returns, but with opposite signs. As in Bhandari (1988), higher market leverage is associated with higher average returns; the average slopes for ln(A/ME) are always positive and more than 4 standard errors from 0. But higher book leverage is associated with lower average returns; the average slopes for ln(A/BE) are always negative and more than 4 standard errors from 0.

The puzzle of the opposite slopes on ln(A/ME) and ln(A/BE) has a simple solution. The average slopes for the two leverage variables are opposite in sign but close in absolute value, e.g., 0.50 and −0.57. Thus it is the difference between market and book leverage that helps explain average returns. But the difference between market and book leverage is book-to-market equity, ln(BE/ME) = ln(A/ME) − ln(A/BE). Table III shows that the average book-to-market slopes in the FM regressions are indeed close in absolute value to the slopes for the two leverage variables.

The close links between the leverage and book-to-market results suggest that there are two equivalent ways to interpret the book-to-market effect in average returns. A high ratio of book equity to market equity (a low stock price relative to book value) says that the market judges the prospects of a firm to be poor relative to firms with low BE/ME. Thus BE/ME may capture the relative-distress effect postulated by Chan and Chen (1991). A high book-to-market ratio also says that a firm's market leverage is high relative to its book leverage; the firm has a large amount of market-imposed leverage because the market judges that its prospects are poor and discounts its stock price relative to book value. In short, our tests suggest that the relative-distress effect, captured by BE/ME, can also be interpreted as an involuntary leverage effect, which is captured by the difference between A/ME and A/BE.

B.3. E/P

Ball (1978) posits that the earnings-price ratio is a catch-all for omitted risk factors in expected returns. If current earnings proxy for expected future earnings, high-risk stocks with high expected returns will have low prices relative to their earnings. Thus, E/P should be related to expected returns, whatever the omitted sources of risk. This argument only makes sense, however, for firms with positive earnings. When current earnings are negative, they are not a proxy for the earnings forecasts embedded in the stock price, and E/P is not a proxy for expected returns. Thus, the slope for E/P in the FM regressions is based on positive values; we use a dummy variable for E/P when earnings are negative.

The U-shaped relation between average return and E/P observed in Table IV is also apparent when the E/P variables are used alone in the FM regressions in Table III. The average slope on the E/P dummy variable (0.57% per month, 2.28 standard errors from 0) confirms that firms with negative earnings have higher average returns. The average slope for stocks with positive E/P (4.72% per month, 4.57 standard errors from 0) shows that average returns increase with E/P when it is positive.

Adding size to the regressions kills the explanatory power of the E/P dummy. Thus the high average returns of negative E/P stocks are better captured by their size, which Table IV says is on average small. Adding both size and book-to-market equity to the E/P regressions kills the E/P dummy and lowers the average slope on E/P from 4.72 to 0.87 (t = 1.23). In contrast, the average slopes for ln(ME) and ln(BE/ME) in the regressions that include E/P are similar to those in the regressions that explain average returns with only size and book-to-market equity. The results suggest that most of the relation between (positive) E/P and average return is due to the positive correlation between E/P and ln(BE/ME), illustrated in Table IV; firms with high E/P tend to have high book-to-market equity ratios.

IV. A Parsimonious Model for Average Returns

The results to here are easily summarized:

(1) When we allow for variation in β that is unrelated to size, there is no reliable relation between β and average return.
(2) The opposite roles of market leverage and book leverage in average returns are captured well by book-to-market equity.
(3) The relation between E/P and average return seems to be absorbed by the combination of size and book-to-market equity.

In a nutshell, market β seems to have no role in explaining the average returns on NYSE, AMEX, and NASDAQ stocks for 1963–1990, while size and book-to-market equity capture the cross-sectional variation in average stock returns that is related to leverage and E/P.

A. Average Returns, Size and Book-to-Market Equity

The average return matrix in Table V gives a simple picture of the two-dimensional variation in average returns that results when the 10 size deciles are each subdivided into 10 portfolios based on ranked values of BE/ME for individual stocks. Within a size decile (across a row of the average return matrix), returns typically increase strongly with BE/ME: on average, the returns on the lowest and highest BE/ME portfolios in a size decile differ by 0.99% (1.63% − 0.64%) per month. Similarly, looking down the columns of the average return matrix shows that there is a negative relation between average return and size: on average, the spread of returns across the size portfolios in a BE/ME group is 0.58% per month. The average return matrix gives life to the conclusion from the regressions that,

Table V

Average Monthly Returns on Portfolios Formed on Size and Book-to-Market Equity; Stocks Sorted by ME (Down) and then BE/ME (Across): July 1963 to December 1990

In June of each year t, the NYSE, AMEX, and NASDAQ stocks that meet the CRSP-COMPUSTAT data requirements are allocated to 10 size portfolios using the NYSE size (ME) breakpoints. The NYSE, AMEX, and NASDAQ stocks in each size decile are then sorted into 10 BE/ME portfolios using the book-to-market ratios for year $t-1$. BE/ME is the book value of common equity plus balance-sheet deferred taxes for fiscal year $t-1$, over market equity for December of year $t-1$. The equal-weighted monthly portfolio returns are then calculated for July of year t to June of year $t+1$.

Average monthly return is the time-series average of the monthly equal-weighted portfolio returns (in percent).

The All column shows average returns for equal-weighted size decile portfolios. The All row shows average returns for equal-weighted portfolios of the stocks in each BE/ME group.

	Book-to-Market Portfolios										
	All	Low	2	3	4	5	6	7	8	9	High
All	1.23	0.64	0.98	1.06	1.17	1.24	1.26	1.39	1.40	1.50	1.63
Small-ME	1.47	0.70	1.14	1.20	1.43	1.56	1.51	1.70	1.71	1.82	1.92
ME-2	1.22	0.43	1.05	0.96	1.19	1.33	1.19	1.58	1.28	1.43	1.79
ME-3	1.22	0.56	0.88	1.23	0.95	1.36	1.30	1.30	1.40	1.54	1.60
ME-4	1.19	0.39	0.72	1.06	1.36	1.13	1.21	1.34	1.59	1.51	1.47
ME-5	1.24	0.88	0.65	1.08	1.47	1.13	1.43	1.44	1.26	1.52	1.49
ME-6	1.15	0.70	0.98	1.14	1.23	0.94	1.27	1.19	1.19	1.24	1.50
ME-7	1.07	0.95	1.00	0.99	0.83	0.99	1.13	0.99	1.16	1.10	1.47
ME-8	1.08	0.66	1.13	0.91	0.95	0.99	1.01	1.15	1.05	1.29	1.55
ME-9	0.95	0.44	0.89	0.92	1.00	1.05	0.93	0.82	1.11	1.04	1.22
Large-ME	0.89	0.93	0.88	0.84	0.71	0.79	0.83	0.81	0.96	0.97	1.18

controlling for size, book-to-market equity captures strong variation in average returns, and controlling for book-to-market equity leaves a size effect in average returns.

B. The Interaction between Size and Book-to-Market Equity

The average of the monthly correlations between the cross-sections of ln(ME) and ln(BE/ME) for individual stocks is -0.26. The negative correlation is also apparent in the average values of ln(ME) and ln(BE/ME) for the portfolios sorted on ME or BE/ME in Tables II and IV. Thus, firms with low market equity are more likely to have poor prospects, resulting in low stock prices and high book-to-market equity. Conversely, large stocks are more likely to be firms with stronger prospects, higher stock prices, lower book-to-market equity, and lower average stock returns.

The correlation between size and book-to-market equity affects the regressions in Table III. Including ln(BE/ME) moves the average slope on ln(ME) from -0.15 ($t = -2.58$) in the univariate regressions to -0.11 ($t = -1.99$) in the bivariate regressions. Similarly, including ln(ME) in the regressions

lowers the average slope on ln(BE/ME) from 0.50 to 0.35 (still a healthy 4.44 standard errors from 0). Thus, part of the size effect in the simple regressions is due to the fact that small ME stocks are more likely to have high book-to-market ratios, and part of the simple book-to-market effect is due to the fact that high BE/ME stocks tend to be small (they have low ME).

We should not, however, exaggerate the links between size and book-to-market equity. The correlation (−0.26) between ln(ME) and ln(BE/ME) is not extreme, and the average slopes in the bivariate regressions in Table III show that ln(ME) and ln(BE/ME) are both needed to explain the cross-section of average returns. Finally, the 10 × 10 average return matrix in Table V provides concrete evidence that, (a) controlling for size, book-to-market equity captures substantial variation in the cross-section of average returns, and (b) within BE/ME groups average returns are related to size.

C. Subperiod Averages of the FM Slopes

The message from the average FM slopes for 1963–1990 (Table III) is that size on average has a negative premium in the cross-section of stock returns, book-to-market equity has a positive premium, and the average premium for market β is essentially 0. Table VI shows the average FM slopes for two roughly equal subperiods (July 1963–December 1976 and January 1977–December 1990) from two regressions: (a) the cross-section of stock returns on size, ln(ME), and book-to-market equity, ln(BE/ME), and (b) returns on β, ln(ME), and ln(BE/ME). For perspective, average returns on the value-weighted and equal-weighted (VW and EW) portfolios of NYSE stocks are also shown.

In FM regressions, the intercept is the return on a standard portfolio (the weights on stocks sum to [illegible] which the weighted averages of the explanatory variables are 0 (Fama (1976), chapter 9). In our tests, the intercept is weighted toward small stocks (ME is in millions of dollars so ln(ME) = 0 implies ME = $1 million) and toward stocks with relatively high book-to-market ratios (Table IV says that ln(BE/ME) is negative for the typical firm, so ln(BE/ME) = 0 is toward the high end of the sample ratios). Thus it is not surprising that the average intercepts are always large relative to their standard errors and relative to the returns on the NYSE VW and EW portfolios.

Like the overall period, the subperiods do not offer much hope that the average premium for β is economically important. The average FM slope for β is only slightly positive for 1963–1976 (0.10% per month, $t = 0.25$), and it is negative for 1977–1990 (−0.44% per month, $t = -1.17$). There is a hint that the size effect is weaker in the 1977–1990 period, but inferences about the average size slopes for the subperiods lack power.

Unlike the size effect, the relation between book-to-market equity and average return is so strong that it shows up reliably in both the 1963–1976 and the 1977–1990 subperiods. The average slopes for ln(BE/ME) are all more than 2.95 standard errors from 0, and the average slopes for the

Table VI

Subperiod Average Monthly Returns on the NYSE Equal-Weighted and Value-Weighted Portfolios and Subperiod Means of the Intercepts and Slopes from the Monthly FM Cross-Sectional Regressions of Returns on (a) Size (ln(ME)) and Book-to-Market Equity (ln(BE/ME)), and (b) β, ln(ME), and ln(BE/ME)

Mean is the time-series mean of a monthly return, Std is its time-series standard deviation, and t(Mn) is Mean divided by its time-series standard error.

	7/63–12/90 (330 Mos.)			7/63–12/76 (162 Mos.)			1/77–12/90 (168 Mos.)		
Variable	Mean	Std	t(Mn)	Mean	Std	t(Mn)	Mean	Std	t(Mn)
NYSE Value-Weighted (VW) and Equal-Weighted (EW) Portfolio Returns									
VW	0.81	4.47	3.27	0.56	4.26	1.67	1.04	4.66	2.89
EW	0.97	5.49	3.19	0.77	5.70	1.72	1.15	5.28	2.82
$R_{it} = a + b_{2t}\ln(ME_{it}) + b_{3t}\ln(BE/ME_{it}) + e_{it}$									
a	1.77	8.51	3.77	1.86	10.10	2.33	1.69	6.67	3.27
b_2	−0.11	1.02	−1.99	−0.16	1.25	−1.62	−0.07	0.73	−1.16
b_3	0.35	1.45	4.43	0.36	1.53	2.96	0.35	1.37	3.30
$R_{it} = a + b_{1t}\beta_{it} + b_{2t}\ln(ME_{it}) + b_{3t}\ln(BE/ME_{it}) + e_{it}$									
a	2.07	5.75	6.55	1.73	6.22	3.54	2.40	5.25	5.92
b_1	−0.17	5.12	−0.62	0.10	5.33	0.25	−0.44	4.91	−1.17
b_2	−0.12	0.89	−2.52	−0.15	1.03	−1.91	−0.09	0.74	−1.64
b_3	0.33	1.24	4.80	0.34	1.36	3.17	0.31	1.10	3.67

subperiods (0.36 and 0.3[illegible]) are close to the average slope (0.35) for the overall period. The subperiod results thus support the conclusion that, among the variables considered here, book-to-market equity is consistently the most powerful for explaining the cross-section of average stock returns.

Finally, Roll (1983) and Keim (1983) show that the size effect is stronger in January. We have examined the monthly slopes from the FM regressions in Table VI for evidence of a January seasonal in the relation between book-to-market equity and average return. The average January slopes for ln(BE/ME) are about twice those for February to December. Unlike the size effect, however, the strong relation between book-to-market equity and average return is not special to January. The average monthly February-to-December slopes for ln(BE/ME) are about 4 standard errors from 0, and they are close to (within 0.05 of) the average slopes for the whole year. Thus, there is a January seasonal in the book-to-market equity effect, but the positive relation between BE/ME and average return is strong throughout the year.

D. β and the Market Factor: Caveats

Some caveats about the negative evidence on the role of β in average returns are in order. The average premiums for β, size, and book-to-market

equity depend on the definitions of the variables used in the regressions. For example, suppose we replace book-to-market equity (ln(BE/ME)) with book equity (ln(BE)). As long as size (ln(ME)) is also in the regression, this change will not affect the intercept, the fitted values or the R^2. But the change, in variables increases the average slope (and the t-statistic) on ln(ME). In other words, it increases the risk premium associated with size. Other redefinitions of the β, size, and book-to-market variables will produce different regression slopes and perhaps different inferences about average premiums, including possible resuscitation of a role for β. And, of course, at the moment, we have no theoretical basis for choosing among different versions of the variables.

Moreover, the tests here are restricted to stocks. It is possible that including other assets will change the inferences about the average premiums for β, size, and book-to-market equity. For example, the large average intercepts for the FM regressions in Table VI suggest that the regressions will not do a good job on Treasury bills, which have low average returns and are likely to have small loadings on the underlying market, size, and book-to-market factors in returns. Extending the tests to bills and other bonds may well change our inferences about average risk premiums, including the revival of a role for market β.

We emphasize, however, that different approaches to the tests are not likely to revive the Sharpe-Lintner-Black model. Resuscitation of the SLB model requires that a better proxy for the market portfolio (a) overturns our evidence that the simple relation between β and average stock returns is flat and (b) leaves β as the only variable relevant for explaining average returns. Such results seem unlikely, given Stambaugh's (1982) evidence that tests of the SLB model do not seem to be sensitive to the choice of a market proxy. Thus, if there is a role for β in average returns, it is likely to be found in a multi-factor model that transforms the flat simple relation between average return and β into a positively sloped conditional relation.

V. Conclusions and Implications

The Sharpe-Lintner-Black model has long shaped the way academics and practitioners think about average return and risk. Black, Jensen, and Scholes (1972) and Fama and MacBeth (1973) find that, as predicted by the model, there is a positive simple relation between average return and market β during the early years (1926–1968) of the CRSP NYSE returns file. Like Reinganum (1981) and Lakonishok and Shapiro (1986), we find that this simple relation between β and average return disappears during the more recent 1963–1990 period. The appendix that follows shows that the relation between β and average return is also weak in the last half century (1941–1990) of returns on NYSE stocks. In short, our tests do not support the central prediction of the SLB model, that average stock returns are positively related to market β.

Banz (1981) documents a strong negative relation between average return and firm size. Bhandari (1988) finds that average return is positively related to leverage, and Basu (1983) finds a positive relation between average return

and E/P. Stattman (1980) and Rosenberg, Reid, and Lanstein (1985) document a positive relation between average return and book-to-market equity for U.S. stocks, and Chan, Hamao, and Lakonishok (1992) find that BE/ME is also a powerful variable for explaining average returns on Japanese stocks.

Variables like size, E/P, leverage, and book-to-market equity are all scaled versions of a firm's stock price. They can be regarded as different ways of extracting information from stock prices about the cross-section of expected stock returns (Ball (1978), Keim (1988)). Since all these variables are scaled versions of price, it is reasonable to expect that some of them are redundant for explaining average returns. Our main result is that for the 1963–1990 period, size and book-to-market equity capture the cross-sectional variation in average stock returns associated with size, E/P, book-to-market equity, and leverage.

A. Rational Asset-Pricing Stories

Are our results consistent with asset-pricing theory? Since the FM intercept is constrained to be the same for all stocks, FM regressions always impose a linear factor structure on returns and expected returns that is consistent with the multifactor asset-pricing models of Merton (1973) and Ross (1976). Thus our tests impose a rational asset-pricing framework on the relation between average return and size and book-to-market equity.

Even if our results are consistent with asset-pricing theory, they are not economically satisfying. What is the economic explanation for the roles of size and book-to-market equity in average returns? We suggest several paths of inquiry.

(a) The intercepts and slopes in the monthly FM regressions of returns on ln(ME) and ln(BE/ME) are returns on portfolios that mimic the underlying common risk factors in returns proxied by size and book-to-market equity (Fama (1976), chapter 9). Examining the relations between the returns on these portfolios and economic variables that measure variation in business conditions might help expose the nature of the economic risks captured by size and book-to-market equity.

(b) Chan, Chen, and Hsieh (1985) argue that the relation between size and average return proxies for a more fundamental relation between expected returns and economic risk factors. Their most powerful factor in explaining the size effect is the difference between the monthly returns on low- and high-grade corporate bonds, which in principle captures a kind of default risk in returns that is priced. It would be interesting to test whether loadings on this or other economic factors, such as those of Chen, Roll, and Ross (1986), can explain the roles of size and book-to-market equity in our tests.

(c) In a similar vein, Chan and Chen (1991) argue that the relation between size and average return is a relative-prospects effect. The earning prospects of distressed firms are more sensitive to economic

conditions. This results in a distress factor in returns that is priced in expected returns. Chan and Chen construct two mimicking portfolios for the distress factor, based on dividend changes and leverage. It would be interesting to check whether loadings on their distress factors absorb the size and book-to-market equity effects in average returns that are documented here.

(d) In fact, if stock prices are rational, BE/ME, the ratio of the book value of a stock to the market's assessment of its value, should be a direct indicator of the relative prospects of firms. For example, we expect that high BE/ME firms have low earnings on assets relative to low BE/ME firms. Our work (in progress) suggests that there is indeed a clean separation between high and low BE/ME firms on various measures of economic fundamentals. Low BE/ME firms are persistently strong performers, while the economic performance of high BE/ME firms is persistently weak.

B. Irrational Asset-Pricing Stories

The discussion above assumes that the asset-pricing effects captured by size and book-to-market equity are rational. For BE/ME, our most powerful expected-return variable, there is an obvious alternative. The cross-section of book-to-market ratios might result from market overreaction to the relative prospects of firms. If overreaction tends to be corrected, BE/ME will predict the cross-section of stock returns.

Simple tests do not confirm that the size and book-to-market effects in average returns are due to market overreaction, at least of the type posited by DeBondt and Thaler (1985). One overreaction measure used by DeBondt and Thaler is a stock's most recent 3-year return. Their overreaction story predicts that 3-year losers have strong post-ranking returns relative to 3-year winners. In FM regressions (not shown) for individual stocks, the 3-year lagged return shows no power even when used alone to explain average returns. The univariate average slope for the lagged return is negative, −6 basis points per month, but less than 0.5 standard errors from 0.

C. Applications

Our main result is that two easily measured variables, size and book-to-market equity, seem to describe the cross-section of average stock returns. Prescriptions for using this evidence depend on (a) whether it will persist, and (b) whether it results from rational or irrational asset-pricing.

It is possible that, by chance, size and book-to-market equity happen to describe the cross-section of average returns in our sample, but they were and are unrelated to expected returns. We put little weight on this possibility, especially for book-to-market equity. First, although BE/ME has long been touted as a measure of the return prospects of stocks, there is no evidence that its explanatory power deteriorates through time. The 1963–1990 relation between BE/ME and average return is strong, and remarkably similar

for the 1963-1976 and 1977-1990 subperiods. Second, our preliminary work on economic fundamentals suggests that high-BE/ME firms tend to be persistently poor earners relative to low-BE/ME firms. Similarly, small firms have a long period of poor earnings during the 1980s not shared with big firms. The systematic patterns in fundamentals give us some hope that size and book-to-market equity proxy for risk factors in returns, related to relative earning prospects, that are rationally priced in expected returns.

If our results are more than chance, they have practical implications for portfolio formation and performance evaluation by investors whose primary concern is long-term average returns. If asset-pricing is rational, size and BE/ME must proxy for risk. Our results then imply that the performance of managed portfolios (e.g., pension funds and mutual funds) can be evaluated by comparing their average returns with the average returns of benchmark portfolios with similar size and BE/ME characteristics. Likewise, the expected returns for different portfolio strategies can be estimated from the historical average returns of portfolios with matching size and BE/ME properties.

If asset-pricing is irrational and size and BE/ME do not proxy for risk, our results might still be used to evaluate portfolio performance and measure the expected returns from alternative investment strategies. If stock prices are irrational, however, the likely persistence of the results is more suspect.

Appendix
Size Versus β: 1941-1990

Our results on the absence of a relation between β and average stock returns for 1963-1990 are so contrary to the tests of the Sharpe-Lintner-Black model by Black, Jensen, and Scholes (1972), Fama and MacBeth (1973), and (more recently) Chan and Chen (1988), that further tests are appropriate. We examine the roles of size and β in the average returns on NYSE stocks for the half-century 1941-1990, the longest available period that avoids the high volatility of returns in the Great Depression. We do not include the accounting variables in the tests because of the strong selection bias (toward successful firms) in the COMPUSTAT data prior to 1962.

We first replicate the results of Chan and Chen (1988). Like them, we find that when portfolios are formed on size alone, there are strong relations between average return and either size or β; average return increases with β and decreases with size. For size portfolios, however, size (ln(ME)) and β are almost perfectly correlated (−0.98), so it is difficult to distinguish between the roles of size and β in average returns.

One way to generate strong variation in β that is unrelated to size is to form portfolios on size and then on β. As in Tables I to III, we find that the resulting independent variation in β just about washes out the positive simple relation between average return and β observed when portfolios are formed on size alone. The results for NYSE stocks for 1941-1990 are thus much like those for NYSE, AMEX, and NASDAQ stocks for 1963-1990.

This appendix also has methodological goals. For example, the FM regressions in Table III use returns on individual stocks as the dependent variable. Since we allocate portfolio βs to individual stocks but use firm-specific values of other variables like size, β may be at a disadvantage in the regressions for individual stocks. This appendix shows, however, that regressions for portfolios, which put β and size on equal footing, produce results comparable to those for individual stocks.

A. Size Portfolios

Table AI shows average monthly returns and market βs for 12 portfolios of NYSE stocks formed on the basis of size (ME) at the end of each year from 1940 to 1989. For these size portfolios, there is a strong positive relation between average return and β. Average returns fall from 1.96% per month for the smallest ME portfolio (1A) to 0.93% for the largest (10E) and β falls from 1.60 to 0.95. (Note also that, as claimed earlier, estimating β as the sum of the slopes in the regression of a portfolio's return on the current and prior month's NYSE value-weighted return produces much larger βs for the smallest ME portfolios and slightly smaller βs for the largest ME portfolios.)

The FM regressions in Table AI confirm the positive simple relation between average return and β for size portfolios. In the regressions of the size-portfolio returns on β alone, the average premium for a unit of β is 1.45% per month. In the regressions of individual stock returns on β (where stocks are assigned the β of their size portfolio), the premium for a unit of β is 1.39%. Both estimates are about 3 standard errors from 0. Moreover, the βs of size portfolios do not leave a residual size effect; the average residuals from the simple regressions of returns on β in Table AI show no relation to size. These positive SLB results for 1941–1990 are like those obtained by Chan and Chen (1988) in tests on size portfolios for 1954–1983.

There is, however, evidence in Table AI that all is not well with the βs of the size portfolios. They do a fine job on the relation between size and average return, but they do a lousy job on their main task, the relation between β and average return. When the residuals from the regressions of returns on β are grouped using the pre-ranking βs of individual stocks, the average residuals are strongly positive for low-β stocks (0.51% per month for group 1A) and negative for high-β stocks (−1.05% for 10B). Thus the market lines estimated with size-portfolio βs exaggerate the tradeoff of average return for β; they underestimate average returns on low-β stocks and overestimate average returns on high-β stocks. This pattern in the β-sorted average residuals for individual stocks suggests that (a) there is variation in β across stocks that is lost in the size portfolios, and (b) this variation in β is not rewarded as well as the variation in β that is related to size.

B. Two-Pass Size-β Portfolios

Like Table I, Table AII shows that subdividing size deciles using the (pre-ranking) βs of individual stocks results in strong variation in β that is

Table AI

Average Returns, Post-Ranking βs and Fama-MacBeth Regression Slopes for Size Portfolios of NYSE Stocks: 1941–1990

At the end of each year $t-1$, stocks are assigned to 12 portfolios using ranked values of ME. Included are all NYSE stocks that have a CRSP price and shares for December of year $t-1$ and returns for at least 24 of the 60 months ending in December of year $t-1$ (for pre-ranking β estimates). The middle 8 portfolios cover size deciles 2 to 9. The 4 extreme portfolios (1A, 1B, 10A, and 10B) split the smallest and largest deciles in half. We compute equal-weighted returns on the portfolios for the 12 months of year t using all surviving stocks. Average Return is the time-series average of the monthly portfolio returns for 1941–1990, in percent. Average firms is the average number of stocks in the portfolios each month. The simple βs are estimated by regressing the 1941–1990 sample of post-ranking monthly returns for a size portfolio on the current month's value-weighted NYSE portfolio return. The sum βs are the sum of the slopes from a regression of the post-ranking monthly returns on the current and prior month's VW NYSE returns.

The independent variables in the Fama-MacBeth regressions are defined for each firm at the end of December of each year $t-1$. Stocks are assigned the post-ranking (sum) β of the size portfolio they are in at the end of year $t-1$. ME is price times shares outstanding at the end of year $t-1$. In the individual-stock regressions, these values of the explanatory variables are matched with CRSP returns for each of the 12 months of year t. The portfolio regressions match the equal-weighted portfolio returns with the equal-weighted averages of β and ln(ME) for the surviving stocks in each month of year t. Slope is the average of the (600) monthly FM regression slopes and SE is the standard error of the average slope. The residuals from the monthly regressions for year t are grouped into 12 portfolios on the basis of size (ME) or pre-ranking β (estimated with 24 to 60 months of data, as available) at the end of year $t-1$. The average residuals are the time-series averages of the monthly equal-weighted portfolio residuals, in percent. The average residuals for regressions (1) and (2) (not shown) are quite similar to those for regressions (4) and (5) (shown).

	Portfolios Formed on Size											
	1A	1B	2	3	4	5	6	7	8	9	10A	10B
Ave. return	1.96	1.59	1.44	1.36	1.28	1.24	1.23	1.17	1.15	1.13	0.97	0.93
Ave. firms	57	56	110	107	107	108	111	113	115	118	59	59
Simple β	1.29	1.24	1.21	1.19	1.16	1.13	1.13	1.12	1.09	1.05	1.00	0.98
Standard error	0.07	0.05	0.04	0.03	0.02	0.02	0.02	0.02	0.01	0.01	0.01	0.01
Sum β	1.60	1.44	1.37	1.32	1.26	1.23	1.19	1.17	1.12	1.06	0.99	0.95
Standard error	0.10	0.06	0.05	0.04	0.03	0.03	0.03	0.02	0.02	0.01	0.01	0.01

Table AI—*Continued*

	Portfolio Regressions				Individual Stock Regressions			
	(1) β	(2) ln(ME)	(3) β and ln(ME)		(4) β	(5) ln(ME)	(6) β and ln(ME)	
Slope	1.45	−0.137	[illegible].05	0.149	1.39	−0.133	0.71	−0.060
SE	0.47	0.044	51	0.115	0.46	0.043	0.81	0.062

	Average Residuals for Stocks Grouped on Size											
	1A	1B	2	3	4	5	6	7	8	9	10A	10B
Regression (4)	0.17	0.00	−0.04	−0.06	−0.05	−0.04	0.00	−0.03	0.03	0.08	0.01	0.04
Standard error	0.11	0.06	0.04	0.04	0.04	0.04	0.03	0.03	0.03	0.03	0.05	0.06
Regression (5)	0.30	0.02	−0.05	−0.06	−0.08	−0.07	−0.03	−0.04	0.02	0.08	0.01	0.13
Standard error	0.14	0.07	0.04	0.04	0.04	0.04	0.04	0.03	0.03	0.03	0.04	0.07
Regression (6)	0.20	0.02	−0.05	−0.07	−0.08	−0.06	−0.01	−0.02	0.04	0.09	0.00	0.06
Standard error	0.10	0.06	0.04	0.04	0.04	0.04	0.03	0.03	0.03	0.03	0.05	0.05

	Average Residuals for Stocks Grouped on Pre-Ranking β											
	1A	1B	2	3	4	5	6	7	8	9	10A	10B
Regression (4)	0.51	0.61	0.38	0.32	0.16	0.12	0.03	−0.10	−0.27	−0.31	−0.66	−1.05
Standard error	0.21	0.19	0.13	0.08	0.04	0.03	0.04	0.05	0.09	0.11	0.18	0.23
Regression (5)	−0.10	0.00	0.02	0.09	0.05	0.07	0.05	0.00	−0.03	−0.01	−0.11	−0.33
Standard error	0.11	0.10	0.07	0.05	0.04	0.03	0.03	0.04	0.05	0.07	0.10	0.13
Regression (6)	0.09	0.25	0.13	0.19	0.11	0.14	0.09	0.01	−0.11	−0.12	−0.38	−0.70
Standard error	0.41	0.37	0.24	0.14	0.07	0.04	0.04	0.09	0.16	0.21	0.34	0.43

Table AII

Properties of Portfolios Formed on Size and Pre-Ranking β: NYSE Stocks Sorted by ME (Down) then Pre-Ranking β (Across): 1941–1990

At the end of year $t-1$, the NYSE stocks on CRSP are assigned to 10 size (ME) portfolios. Each size decile is subdivided into 10 β portfolios using pre-ranking βs of individual stocks, estimated with 24 to 60 monthly returns (as available) ending in December of year $t-1$. The equal-weighted monthly returns on the resulting 100 portfolios are then calculated for year t. The average returns are the time series averages of the monthly returns, in percent. The post-ranking βs use the full 1941–1990 sample of post-ranking returns for each portfolio. The pre- and post-ranking βs are the sum of the slopes from a regression of monthly returns on the current and prior month's NYSE value-weighted market return. The average size for a portfolio is the time-series average of each month's average value of ln(ME) for stocks in the portfolio. ME is denominated in millions of dollars. There are, on average, about 10 stocks in each size-β portfolio each month. The All column shows parameter values for equal-weighted size-decile (ME) portfolios. The All rows show parameter values for equal-weighted portfolios of the stocks in each β group.

	All	Low-β	β-2	β-3	β-4	β-5	β-6	β-7	β-8	β-9	High-β
					Panel A: Average Monthly Return (in Percent)						
All		1.22	1.30	1.32	1.35	1.36	1.34	1.29	1.34	1.14	1.10
Small-ME	1.78	1.74	1.76	2.08	1.91	1.92	1.72	1.77	1.91	1.56	1.46
ME-2	1.44	1.41	1.35	1.33	1.61	1.72	1.59	1.40	1.62	1.24	1.11
ME-3	1.36	1.21	1.40	1.22	1.47	1.34	1.51	1.33	1.57	1.33	1.21
ME-4	1.28	1.26	1.29	1.19	1.27	1.51	1.30	1.19	1.56	1.18	1.00
ME-5	1.24	1.22	1.30	1.28	1.33	1.21	1.37	1.41	1.31	0.92	1.06
ME-6	1.23	1.21	1.32	1.37	1.09	1.34	1.10	1.40	1.21	1.22	1.08
ME-7	1.17	1.08	1.23	1.37	1.27	1.19	1.34	1.10	1.11	0.87	1.17
ME-8	1.15	1.06	1.18	1.26	1.25	1.26	1.17	1.16	1.05	1.08	1.04
ME-9	1.13	0.99	1.13	1.00	1.24	1.28	1.31	1.15	1.11	1.09	1.05
Large-ME	0.95	0.99	1.01	1.12	1.01	0.89	0.95	0.95	1.00	0.90	0.68

Table AII—*Continued*

	All	Low-β	β-2	β-3	β-4	β-5	β-6	β-7	β-8	β-9	High-β
Panel B: Post-Ranking β											
All		0.76	0.95	1.05	1.14	1.22	1.26	1.34	1.38	1.49	1.69
Small-ME	1.52	1.17	1.40	1.31	1.50	1.46	1.50	1.69	1.60	1.75	1.92
ME-2	1.37	0.86	1.09	1.12	1.24	1.39	1.42	1.48	1.60	1.69	1.91
ME-3	1.32	0.88	0.96	1.18	1.19	1.33	1.40	1.43	1.56	1.64	1.74
ME-4	1.26	0.69	0.95	1.06	1.15	1.24	1.29	1.46	1.43	1.64	1.83
ME-5	1.23	0.70	0.95	1.04	1.10	1.22	1.32	1.34	1.41	1.56	1.72
ME-6	1.19	0.68	0.86	1.04	1.13	1.20	1.20	1.35	1.36	1.48	1.70
ME-7	1.17	0.67	0.88	0.95	1.14	1.18	1.26	1.27	1.32	1.44	1.68
ME-8	1.12	0.64	0.83	0.99	1.06	1.14	1.14	1.21	1.26	1.39	1.58
ME-9	1.06	0.68	0.81	0.94	0.96	1.06	1.11	1.18	1.22	1.25	1.46
Large-ME	0.97	0.65	0.73	0.90	0.91	0.97	1.01	1.01	1.07	1.12	1.38
Panel C: Average Size (ln(ME))											
All		4.39	4.39	4.40	4.40	4.39	4.40	4.38	4.37	4.37	4.34
Small-ME	1.93	2.04	1.99	2.00	1.96	1.92	1.92	1.91	1.90	1.87	1.80
ME-2	2.80	2.81	2.79	2.81	2.83	2.80	2.79	2.80	2.80	2.79	2.79
ME-3	3.27	3.28	3.27	3.28	3.27	3.27	3.28	3.29	3.27	3.27	3.26
ME-4	3.67	3.67	3.67	3.67	3.68	3.68	3.67	3.68	3.66	3.67	3.67
ME-5	4.06	4.07	4.06	4.05	4.06	4.07	4.06	4.05	4.05	4.06	4.06
ME-6	4.45	4.45	4.44	4.46	4.45	4.45	4.45	4.45	4.44	4.45	4.45
ME-7	4.87	4.86	4.87	4.86	4.87	4.87	4.88	4.87	4.87	4.85	4.87
ME-8	5.36	5.38	5.38	5.38	5.35	5.36	5.37	5.37	5.36	5.35	5.34
ME-9	5.98	5.96	5.98	5.99	6.00	5.98	5.98	5.97	5.95	5.96	5.96
Large-ME	7.12	7.10	7.12	7.16	7.17	7.20	7.29	7.14	7.09	7.04	6.83

independent of size. The β sort of a size decile always produces portfolios with similar average ln(ME) but much different (post-ranking) βs. Table AII also shows, however, that investors are not compensated for the variation in β that is independent of size. Despite the wide range of βs in each size decile, average returns show no tendency to increase with β. AII

The FM regressions in Table AIII formalize the roles of size and β in NYSE average returns for 1941-1990. The regressions of returns on β alone show that using the βs of the portfolios formed on size and β, rather than size alone, causes the average slope on β to fall from about 1.4% per month (Table AI) to about 0.23% (about 1 standard error from 0). Thus, allowing for variation in β that is unrelated to size flattens the relation between average return and β, to the point where it is indistinguishable from no relation at all.

The flatter market lines in Table AIII succeed, however, in erasing the negative relation between β and average residuals observed in the regressions of returns on β alone in Table AI. Thus, forming portfolios on size and β (Table AIII) produces a better description of the simple relation between average return and β than forming portfolios on size alone (Table AI). This improved description of the relation between average return and β is evidence that the β estimates for the two-pass size-β portfolios capture variation in true βs that is missed when portfolios are formed on size alone.

Unfortunately, the flatter market lines in Table AIII have a cost, the emergence of a residual size effect. Grouped on the basis of ME for individual stocks, the average residuals from the univariate regressions of returns on the βs of the 100 size-β portfolios are strongly positive for small stocks and negative for large stocks (0.60% per month for the smallest ME group, 1A, and -0.27% for the largest, 10B). Thus, when we allow for variation in β that is independent of size, the resulting βs leave a large size effect in average returns. This residual size effect is much like that observed by Banz (1981) with the βs of portfolios formed on size and β.

The correlation between size and β is -0.98 for portfolios formed on size alone. The independent variation in β obtained with the second-pass sort on β lowers the correlation to -0.50. The lower correlation means that bivariate regressions of returns on β and ln(ME) are more likely to distinguish true size effects from true β effects in average returns.

The bivariate regressions (Table AIII) that use the βs of the size-β portfolios are more bad news for β. The average slopes for ln(ME) are close to the values in the univariate size regressions, and almost 4 standard errors from 0, but the average slopes for β are negative and less than 1 standard error from 0. The message from the bivariate regressions is that there is a strong relation between size and average return. But like the regressions in Table AIII that explain average returns with β alone, the bivariate regressions say that there is no reliable relation between β and average returns when the tests use βs that are not close substitutes for size. These uncomfortable SLB results for NYSE stocks for 1941-1990 are much like those for NYSE, AMEX, and NASDAQ stocks for 1963-1990 in Table III.

C. Subperiod Diagnostics

Our results for 1941–1990 seem to contradict the evidence in Black, Jensen, and Scholes (BJS) (1972) and Fama and MacBeth (FM) (1973) that there is a reliable positive relation between average return and β. The βs in BJS and FM are from portfolios formed on β alone, and the market proxy is the NYSE equal-weighted portfolio. We use the βs of portfolios formed on size and β, and our market is the value-weighted NYSE portfolio. We can report, however, that our inference that there isn't much relation between β and average return is unchanged when (a) the market proxy is the NYSE EW portfolio, (b) portfolios are formed on just (pre-ranking) βs, or (c) the order of forming the size-β portfolios is changed from size then β to β then size.

A more important difference between our results and the earlier studies is the sample periods. The tests in BJS and FM end in the 1960s. Table AIV shows that when we split the 50-year 1941-1990 period in half, the univariate FM regressions of returns on β produce an average slope for 1941–1965 (0.50% per month, $t = 1.82$) more like that of the earlier studies. In contrast, the average slope on β for 1966–1990 is close to 0 (-0.02, $t = 0.06$).

But Table AIV also shows that drawing a distinction between the results for 1941–1965 and 1966–1990 is misleading. The stronger tradeoff of average return for β in the simple regressions for 1941–1965 is due to the first 10 years, 1941–1950. This is the only period in Table AIV that produces an average premium for β (1.26% per month) that is both positive and more than 2 standard errors from 0. Conversely, the weak relation between β and average return for 1966–1990 is largely due to 1981–1990. The strong negative average slope in the univariate regressions of returns on β for 1981–1990 (-1.01, $t = -2.10$) offsets a positive slope for 1971–1980 (0.82, $t = 1.27$).

The subperiod variation in the average slopes from the FM regressions of returns on β alone seems moot, however, given the evidence in Table AIV that adding size always kills any positive tradeoff of average return for β in the subperiods. Adding size to the regressions for 1941–1965 causes the average slope for β to drop from 0.50 ($t = 1.82$) to 0.07 ($t = 0.28$). In contrast, the average slope on size in the bivariate regressions (-0.16, $t = -2.97$) is close to its value (-0.17, $t = -2.88$) in the regressions of returns on ln(ME) alone. Similar comments hold for 1941–1950. In short, any evidence of a positive average premium for β in the subperiods seems to be a size effect in disguise.

D. Can the SLB Model Be Saved?

Before concluding that β has no explanatory power, it is appropriate to consider other explanations for our results. One possibility is that the variation in β produced by the β sorts of size deciles in just sampling error. If so, it is not surprising that the variation in β within a size decile is unrelated to average return, or that size dominates β in bivariate tests. The standard errors of the βs suggest, however, that this explanation cannot save the SLB

Table AIII

Average Slopes, Their Standard Errors (SE), and Average Residuals from Monthly FM Regressions for Individual NYSE Stocks and for Portfolios Formed on Size and Pre-Ranking β: 1941–1990

Stocks are assigned the post-ranking β of the size-β portfolio they are in at the end of year $t-1$ (Table AII). ln(ME) is the natural log of price times shares outstanding at the end of year $t-1$. In the individual-stock regressions, these values of the explanatory variables are matched with CRSP returns for each of the 12 months in year t. The portfolio regressions match the equal-weighted portfolio returns for the size-β portfolios (Table AII) with the equal-weighted averages of β and ln(ME) for the surviving stocks in each month of year t. Slope is the time-series average of the monthly regression slopes from 1941–1990 (600 months); SE is the time-series standard error of the average slope.

The residuals from the monthly regressions on year t are grouped into 12 portfolios on the basis of size or pre-ranking β (estimated with 24 to 60 months of returns as available) as of the end of year $t-1$. The average residuals are the time-series averages of the monthly equal-weighted averages of the residuals in percent. The average residuals (not shown) from the FM regressions (1) to (3) that use the returns on the 100 size-β portfolios as the dependent variable are always within 0.01 of those from the regressions for individual stock returns. This is not surprising given that the correlation between the time-series of 1941–1990 monthly FM slopes on β or ln(ME) for the comparable portfolio and individual stock regressions is always greater than 0.99.

	Portfolio Regressions				Individual Stock Regressions			
	(1) β	(2) ln(ME)	(3) β and ln(ME)		(4) β	(5) ln(ME)	(6) β and ln(ME)	
Slope	0.22	−0.128	−0.13	−0.143	0.24	−0.133	−0.14	−0.147
SE	0.24	0.043	0.21	0.039	0.23	0.043	0.21	0.039

	Average Residuals for Stocks Grouped on Size											
	1A	1B	2	3	4	5	6	7	8	9	10A	10B
Regression (4)	0.60	0.26	0.13	0.06	−0.01	−0.03	−0.03	−0.09	−0.10	−0.11	−0.25	−0.27
Standard error	0.21	0.10	0.06	0.04	0.04	0.04	0.04	0.04	0.04	0.05	0.06	0.08
Regression (5)	0.30	0.02	−0.05	−0.06	−0.08	−0.07	−0.03	−0.04	0.02	0.08	0.01	0.13
Standard error	0.14	0.07	0.04	0.04	0.04	0.04	0.04	0.03	0.03	0.03	0.04	0.07
Regression (6)	0.31	0.02	−0.05	−0.06	−0.09	−0.07	−0.03	−0.04	0.02	0.08	0.01	0.13
Standard error	0.14	0.07	0.04	0.04	0.04	0.04	0.04	0.03	0.03	0.03	0.04	0.07

Table AIII—*Continued*

	Portfolio Regressions						Individual Stock Regressions					
	(1) β		(2) ln(ME)		(3) β and ln(ME)		(4) β		(5) ln(ME)		(6) β and ln(ME)	
	Average Residuals for Stocks Grouped on Pre-Ranking β											
	1A	1B	2	3	4	5	6	7	8	9	10A	10B
Regression (4)	−0.08	0.03	−0.01	0.08	0.04	0.08	0.04	0.02	−0.03	0.02	−0.11	−0.32
Standard error	0.07	0.05	0.03	0.03	0.03	0.03	0.04	0.04	0.04	0.04	0.06	0.07
Regression (5)	−0.10	0.00	0.02	0.09	0.05	0.07	0.05	0.00	−0.03	−0.01	−0.11	−0.33
Standard error	0.11	0.10	0.07	0.05	0.04	0.03	0.03	0.04	0.05	0.07	0.10	0.13
Regression (6)	−0.17	−0.07	−0.02	0.07	0.04	0.06	0.05	0.03	0.00	0.04	−0.04	−0.23
Standard error	0.05	0.04	0.03	0.03	0.03	0.03	0.03	0.03	0.04	0.04	0.06	0.07

Table AIV

Subperiod Average Returns on the NYSE Value-Weighted and Equal-Weighted Portfolios and Average Values of the Intercepts and Slopes for the FM Cross-Sectional Regressions of Individual Stock Returns on β and Size (ln(ME))

Mean is the average VW or EW return or an average slope from the monthly cross-sectional regressions of individual stock returns on β and/or ln(ME). Std is the standard deviation of the time-series of returns or slopes, and t(Mn) is Mean over its time-series standard error. The average slopes (not shown) from the FM regressions that use the returns on the 100 size-β portfolios of Table AII as the dependent variable are quite close to those for individual stock returns. (The correlation between the 1941–1990 month-by-month slopes on β or ln(ME) for the comparable portfolio and individual stock regressions is always greater than 0.99

Panel A

	1941–1990 (600 Mos.)			1941–1965 (300 Mos.)			1966–1990 (300 Mos.)		
Variable	Mean	Std	t(Mn)	Mean	Std	t(Mn)	Mean	Std	t(Mn)
NYSE Value-Weighted (VW) and Equal-Weighted (EW) Portfolio Returns									
VW	0.93	4.15	5.49	1.10	3.58	5.30	0.76	4.64	2.85
EW	1.12	5.10	5.37	1.33	4.42	5.18	0.91	5.70	2.77
$R_{it} = a + b_{1t}\beta_{it} + e_{it}$									
a	0.98	3.93	6.11	0.84	3.18	4.56	1.13	4.57	4.26
b_1	0.24	5.52	1.07	0.50	4.75	1.82	−0.02	6.19	−0.06
$R_{it} = a + b_{2t}\ln(ME_{it}) + e_{it}$									
a	1.70	8.24	5.04	1.88	6.43	5.06	1.51	9.72	2.69
b_2	−0.13	1.06	−3.07	−0.17	1.01	−2.88	−0.10	1.11	−1.54
$R_{it} = a + b_{1t}\beta_{it} + b_{2t}\ln(ME_{it}) + e_{it}$									
a	1.97	6.16	7.84	1.80	4.77	6.52	2.14	7.29	5.09
b_1	−0.14	5.05	−0.66	0.07	4.15	0.28	−0.34	5.80	−1.01
b_2	−0.15	0.96	−3.75	−0.16	0.94	−2.97	−0.13	0.99	−2.34

Table AIV—*Continued*

Panel B:

	1941–1950		1951–1960		1961–1970		1971–1980		1981–1990	
Return	Mean	t(Mn)	Mean	t(Mn)	Mean	t(Mn)	Mean	t(Mn)	Mean	t(Mn)
NYSE Value-Weighted (VW) and Equal-Weighted (EW) Portfolio Returns										
VW	1.05	2.88	1.18	3.95	0.66	1.84	0.72	1.67	1.04	2.40
EW	1.59	3.16	1.13	3.76	0.88	1.96	1.04	1.82	0.95	2.01
$R_{it} = a + b_{1t}\beta_{it} + e_{it}$										
a	0.24	0.66	1.41	6.36	0.64	1.94	0.27	0.62	2.35	5.99
b_1	1.26	2.20	−0.19	−0.63	0.32	0.72	0.82	1.27	−1.01	−2.10
$R_{it} = a + b_{2t}\ln(ME_{it}) + e_{it}$										
a	2.63	3.47	1.08	2.73	1.78	2.50	2.18	2.03	0.82	1.20
b_2	−0.37	−2.90	0.03	0.53	−0.17	−2.19	−0.20	−1.57	0.04	0.57
$R_{it} = a + b_{1t}\beta_{it} + b_{2t}\ln(ME_{it}) + e_{it}$										
a	2.14	3.93	1.38	4.03	2.01	4.16	1.50	2.12	2.84	4.25
b_1	0.34	0.75	−0.17	−0.53	−0.11	−0.27	0.41	0.75	−1.14	−2.16
b_2	−0.34	−2.92	0.01	0.20	−0.13	−2.89	−0.16	−1.50	−0.07	−0.84

model. The standard errors for portfolios formed on size and β are only slightly larger (0.02 to 0.11) than those for portfolios formed on size alone (0.01 to 0.10, Table AI). And the range of the post-ranking βs within a size decile is always large relative to the standard errors of the βs.

Another possibility is that the proportionality condition (1) for the variation through time in true βs, that justifies the use of full-period post-ranking βs in the FM tests, does not work well for portfolios formed on size and β. If this is a problem, post-ranking βs for the size-β portfolios should not be highly correlated across subperiods. The correlation between the half-period (1941–1965 and 1966–1990) βs of the size-β portfolios is 0.91, which we take to be good evidence that the full-period β estimates for these portfolios are informative about true βs. We can also report that using 5-year βs (pre- or post-ranking) in the FM regressions does not change our negative conclusions about the role of β in average returns, as long as portfolios are formed on β as well as size, or on β alone.

Any attempt to salvage the simple positive relation between β and average return predicted by the SLB model runs into three damaging facts, clear in Table AII. (a) Forming portfolios on size and pre-ranking βs produces a wide range of post-ranking βs in every size decile. (b) The post-ranking βs closely reproduce (in deciles 2 to 10 they exactly reproduce) the ordering of the pre-ranking βs used to form the β-sorted portfolios. It seems safe to conclude that the increasing pattern of the post-ranking βs in every size decile captures the ordering of the true βs. (c) Contrary to the SLB model, the β sorts do not produce a similar ordering of average returns. Within the rows (size deciles) of the average return matrix in Table AII, the high-β portfolios have average returns that are close to or less than the low-β portfolios.

But the most damaging evidence against the SLB model comes from the univariate regressions of returns on β in Table AIII. They say that when the tests allow for variation in β that is unrelated to size, the relation between β and average return for 1941–1990 is weak, perhaps nonexistent, even when β is the only explanatory variable. We are forced to conclude that the SLB model does not describe the last 50 years of average stock returns.

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Recent Sector and

Industry Betas: A Closer Look

Most of today's capital market theory is driven by the fundamental relationship between risk and return. The basic premise is that most investors are rational, and that a rational investor will only accept an investment with greater risk if it offers an expectation of greater return. With the advent of William Sharpe's Capital Asset Pricing Model (CAPM) in the early 1960s, the risk of an individual stock is generally accepted to be measured by its Beta, the historical sensitivity of a stock's price to market price movements. This article reviews a recent study challenging the empirical relevance of Beta, then investigates whether these results can be extended to the industry level.

Is This The End Of Beta?

Two of the most prestigious researchers in the financial community, professors Eugene F. Fama and Kenneth R. French from the University of Chicago have challenged the traditional relationship between Beta and return in a recent paper published by the Center for Research In Security Prices. In this study, the duo traced the performance of thousands of stocks over 50 years, but found no statistical support for the hypothesis that the relationship between volatility and return is significantly different from random. Indeed, professor Fama concluded, 'The fact is that Beta, as the sole variable explaining returns on stocks, is dead.' These findings support previous studies that have called into question the real-world applicability of the CAPM Beta, including papers by Keim (Financial Analysts Journal, 1986), and Roll (Journal of Financial Economics, 1977). Never before, however, has the lack of a statistically significant relationship between Beta and return been so rigorously and dramatically established. The line graph on the top half of the opposite page illustrates the startling degree to which the theoretical straight-line relationship is refuted by actual data. This new documentation has investors all over the world paraphrasing Edward G. Robinson's line from the 1930 classic gangster movie, *Little Caesar*, by asking 'Mother of Mercy, is this the end of Beta?'

Empirical Evidence Is Much More Convincing On The Industry Level

Although the argument of Beta's relevancy as a determinant of stock price's is a long way from conclusion, there is no debate that its veneer of incontestability has been severely damaged. A logical extension of these findings, then, is to question the merit of the industry-level Beta as an explanatory variable in the relationship between industry price indexes and historical price performance. While not nearly as much academic literature has been devoted to industry Betas as individual stock Betas, the articles that we could find all demonstrate the existence of much stronger relationships between Beta and return by industry class. Carleton and Lakonishok (Financial Analysts Journal, 1985) found a coefficient of determination of 0.86, indicated a very strong near-linear relationship, for 15 industry groups measured between 1971 and 1980. Furthermore, Levhari and

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Recent Sector and Industry Trends

Levy (Review of Economics and Statistics, 1977) demonstrated that the error terms in beta estimation intervals are greatly reduced when industry aggregates are employed.

A 12-Month Examination Of Value Line Industry Betas

Since this issue marks the one-year publication anniversary of the *Value Line Industry Review*, we selected the twelve-month period ending this past Friday, March 6th, for a test of Value Line industry Betas. Accordingly, we aggregated our 104 industry groups by their Betas that appeared in our March 8, 1991 Edition, and divided the universe into five quintiles. First quintile industries had Betas between 1.22 and 1.54. The ranges for the other four groupings were 1.11 – 1.20, 1.06 – 1.10, 0.96 – 1.05, and 0.68 – 0.95 respectively. Given that this was generally a Bull Market period, marked by a Dow Jones Industrial Average rise of 9% (from 2955.2 to 3168.8) and a NASDAQ jump of 29.7% (from 475.1 to 616.0), we expected that high-Beta industries, most sensitive to market price movements, should experience price gains that were greater than low-Beta industries. The bar graph featured on the bottom half of page 1 – 9 verifies vividly that this was, indeed, the case. Quintile 1, representing the industries with the 21 highest Betas, enjoyed a 37.3% price rise while the low-Beta groups comprising Quintile 5 produced a meager 5.2% gain. The average industry group gained 22.2%. The most prosperous Quintile 1 groups include: Retail/Building Supply, Homebuilding, Securities Brokers, and Computer Software. On the other extreme, Foreign Electronics, Petroleum Producers, and Canadian Energy were all among the worst-performing, low-Beta industries.

Implications For Top-Down investors

A major problem with the high-Beta strategy is that, while a greater-than-average rise can be expected in an up market, an above-average decline is probable in a down market. Since market-timing is very difficult at best, an investor cannot know with certainty whether this is a good time to emphasize high-Beta industry groups. Our largely neutral asset allocation position of 60% – 70% in equities does little to help decide the situation. Alternatively, we have shown in past studies that emphasizing industry groups with high Timeliness ranks has been a profitable strategy in both up and down markets. For those who expect the bull market to continue, timely high-Beta choices include Telecom/Cellular, Semiconductors, Retail/Building Supply, and Shoes. Investors with a bearish outlook may prefer Telecom/Foreign, Insurance/Brokers, and Beverages/Soft Drink. Whatever the choice, you will most likely find the industry-Beta-to-return relationship to be more reliable than its counterpart on the individual stock level.

THE CAPITAL ASSET PRICING MODEL AND THE INVESTMENT HORIZON

David Levhari and Haim Levy*

I. Introduction

IN following the mean-variance analysis developed by Markowitz (1952) and Tobin (1958), Sharpe (1964), Lintner (1965a, b) and Treynor (1961) have developed the theory for determination of asset prices under conditions of uncertainty. The equilibrium asset pricing model, and its implication for measuring ex post performance of individual securities, have been empirically tested by Lintner,[1] Jensen (1968, 1972), Miller and Scholes (1972), Douglas (1969), Roll (1969) and others. The empirical results obtained by both Douglas and Lintner deviated from the theory of the model. Moreover, Miller and Scholes have run empirical tests similar to those of Douglas and Lintner and found a significant disparity between the theoretical model and the empirical evidence. They maintain that part of this discrepancy can be explained by possible statistical biases, measurement errors, and consideration of the skewness of the distribution of returns. Black, Jensen, and Scholes (1972) (hereafter B-J-S) confirm the systematic bias. Using monthly data covering a 35 year period, they discovered that, on average, high risk securities earned less than the amount predicted by the model. Similarly, earnings on low risk securities exceeded the amount predicted.[2]

Although these disparities clearly suggest some systematic empirical bias, we are not examining a possible statistical bias but a mathematical bias stemming from one of the assumptions underlying the capital asset pricing model (CAPM). To be more specific, the model assumes that all investors are single period, expected utility of terminal wealth maximizers. There is no particular restriction on the length of this period as long as it is identical for all investors. Clearly, the length of the "true" investment horizon affects asset prices under conditions of uncertainty. We claim that the disparities noted above may result from using data calculated for an investment horizon that differs from the "true" investment horizon.

In the various empirical tests, the investment horizon has been selected arbitrarily. For example, Lintner and Miller and Scholes use annual data (i.e., they implicitly assume a one-year horizon), Douglas uses quarterly and annual data; Black, Jensen and Scholes, as well as Friend and Blume (1970), use monthly rates of return in their empirical tests, while Roll uses weekly data. It has been shown elsewhere by Levy (1972) that the Reward to Variability index (developed by Sharpe, 1966) is a function of the investment horizon assumed. Hence, there exists a systematic mathematical bias that is a function of the horizon assumed. The above theoretical findings are related to the theory of pricing capital assets, but deal only with *efficient portfolios* and not individual stocks.

In this paper we illustrate that the assumed horizon plays a crucial role in empirical testing. Any deviation from the "true" horizon causes a systematic bias in the regression coefficient (i.e., in the security systematic risk). This in turn causes a systematic bias in the performance measures of each security, and hence the deviation between the theoretical model and the empirical evidence. The results of this paper are not limited to the theory of pricing capital assets; they are applicable to any econometric study in which the variables have multiplicative rather than additive properties. In such a case the regression coefficients will have a mathe-

Received for publication February 27, 1975. Revision accepted for publication March 8, 1976.

* The authors acknowledge the technical assistance of Moshe Smith and two anonymous referees. The first author has been partially financed by the Maurice Falk Foundation, and the second author has been financed by the Ford Foundation.

[1] Lintner's paper "Security Prices and Risk: The Theory and a Comparative Analysis of A.T.&T. and Leading Industrials" was presented at the Conference on Economics of Regulated Public Utilities, June 24, 1965, Chicago.

[2] Miller and Scholes show that the presence of certain biases could have accounted for the Douglas and Lintner findings. B-J-S maintain that the assumption of borrowing at riskless interest rates could have accounted for these deviations.

matical bias which is a function of the unit of time for which empirical data are collected.

In section II we analyze the bias of using data for a longer period than the true horizon for estimating the systematic risk. In section III similar analysis is carried out for horizons smaller than the true investment horizon. In section IV the bias in measuring performance indices by deviation from the "true" horizon is discussed. Section V deals with the implication of the bias discussed on the empirical evidence. Section VI presents some empirical findings which relate to the theoretical results of the previous sections.

II. The Bias in Measuring Systematic Risk Using Data for a Longer Period than the "True" Horizon

Let us first assume that the investor horizon is one time unit and hence the true systematic risk will be denoted by β_1. The data are collected for a longer period, n-time units. The calculated systematic risk for the n period data will be denoted by β_n. In order to examine the bias stemming from using a longer horizon let us now calculate the n periods' covariance between the ith security and the market rate of return on the basis of one period parameters of the distribution. It is assumed that between periods, the rates of return have an identical independent distribution. Denote the t period rate of return of the ith security by x_{it}, and that of the market portfolio by x_{mt}.

If we denote the n periods covariance by cov_n we get by definition $\text{cov}_n = \text{cov}\{x_{i,1} \cdots x_{i,n}, x_{m,1} \cdots x_{m,n}\} = E\{x_{i,1} \cdots x_{i,n} \cdot x_{m,1} \cdots x_{m,n}\} - E\{x_{i,1} \cdots x_{i,n}\} \cdot E\{x_{m,1} \cdots x_{m,n}\}$. Using the assumption of independent distributed returns we get

$$\text{cov}_n = Ex_{i,1}x_{m,1} \cdot Ex_{i,2}x_{m,2} \cdots Ex_{i,n}x_{m,n} - Ex_{i,1} \cdots Ex_{i,n} \cdot Ex_{m,1} \cdots Ex_{m,n}$$

and from the fact of identical distribution of returns we can ignore the second index of each term to obtain

$$\text{cov}_n = (Ex_{i,1}x_{m,1})^n - (Ex_{i,1})^n(Ex_{m,1})^n.$$

But since $Ex_{i,1}x_{m,1} = \text{cov}(x_{i,1}x_{m,1}) + Ex_{i,1}Ex_{m,1}$, we obtain

$$\text{cov}_n = [\text{cov}(x_{i,1}, x_{m,1}) + Ex_{i,1}Ex_{m,1}]^n - (Ex_{i,1})^n(Ex_{m,1})^n.$$

If we denote $\mu_i = Ex_{i,1}$, $\mu_m = Ex_{m,1}$ and $\text{cov}(x_{i,1}x_{m,1}) = \text{cov}(x_i, x_m)$, we find

$$\text{cov}_n = [\text{cov}(x_i, x_m) + \mu_i\mu_m]^n - \mu_i^n\mu_m^n. \quad (1)$$

The nth period variance of the ith security will be a special case of (1) given by

$$\sigma_n^2 = [\sigma_i^2 + \mu_i^2]^n - \mu_i^{2n}. \quad (2)$$

Now, let us examine the relation between β_1 and β_n where the true investment horizon is assumed to be one period while the data collected and the regression analysis are based on n periods.

Since we assume that the investment horizon is one period we obtain the following for the market equilibrium relationship:

$$\mu_1 = r_F + \beta_1(\mu_m - r_F) \quad (3)$$

where μ_1 is the expected return of the security under consideration, r_F is the riskless interest rate, μ_m is the expected rate of return on the market portfolio and $\beta_1 = \text{cov}(x_1, x_m)/\sigma_m^2$, i.e., the regression coefficient of x_1—the rate of return on the given security—with respect to x_m—the rate of return on the market portfolio.

Now, suppose that in the empirical test, n-period data are used rather than the one period data. Namely, the rate of return of the security under consideration is given by $x_1x_2 \cdots x_n$, and the rate of return on the market portfolio is given by $x_{m,1}x_{m,2} \cdots x_{m,n}$.

The n-period regression coefficient is given by

$$\tilde{\beta}_n = \frac{\text{cov}(x_1 \cdots x_n; x_{m,1} \cdots x_{m,n})}{\text{var}\{x_{m,1} \cdots x_{m,n}\}} \quad (4)$$

where the tilde denotes the measured coefficient (not necessarily the right one). Using relations (1) and (2) we find

$$\tilde{\beta}_n = \frac{[\text{cov}(x_1, x_m) + \mu_1\mu_m]^n - (\mu_1\mu_m)^n}{(\sigma_m^2 + \mu_m^2)^n - \mu_m^{2n}} = \frac{\sum_{i=0}^{n-1}\binom{n}{i}[\text{cov}(x_1, x_m)]^{n-i}(\mu_1\mu_m)^i}{\sum_{i=0}^{n-1}\binom{n}{i}(\sigma_m^2)^{n-i}(\mu_m^2)^i}. \quad (5)$$

Let us now divide and multiply the ith term ($i = 0, \ldots, n-1$) of the numerator by $(\sigma_m^2)^{n-i}$ and using $\beta_1 = \text{cov}(x_1, x_m)/\sigma_m^2$ we obtain

$$\tilde{\beta}_n = \frac{\sum_{i=0}^{n-1} \binom{n}{i} \beta_1^{n-i} (\sigma_m^2)^{n-i} (\mu_1 \mu_m)^i}{\sum_{i=0}^{n-1} \binom{n}{i} (\sigma_m^2)^{n-i} (\mu_m^2)^i}. \quad (6)$$

Substituting for μ_1 from equation (3), we find $\mu_1 = \mu_m + (\beta_1 - 1)(\mu_m - r)$ and we get

$$\tilde{\beta}_n = \frac{\sum_{i=0}^{n-1} \binom{n}{i} \beta_1^{n-i} (\sigma_m^2)^{n-i} [(\mu_m + (\beta_1 - 1)(\mu_m - r))\mu_m]^i}{\sum_{i=0}^{n-1} \binom{n}{i} (\sigma_m^2)^{n-i} (\mu_m^2)^i} \quad (7)$$

or

$$\tilde{\beta}_n = \frac{\sum_{i=0}^{n-1} \binom{n}{i} \beta_1^{n-i} (\sigma_m^2)^{n-i} \left[1 + (\beta_1 - 1)\frac{\mu_m - r}{\mu_m}\right]^i (\mu_m^2)^i}{\sum_{i=0}^{n-1} \binom{n}{i} (\sigma_m^2)^{n-i} (\mu_m^2)^i}. \quad (8)$$

Denote

$$a_i = \binom{n}{i} (\sigma_m^2)^{n-i} (\mu_m^2)^i$$

and

$$\alpha = (\beta_1 - 1)\frac{\mu_m - r}{\mu_m};$$

thus,

$$\tilde{\beta}_n = \frac{\sum_{i=0}^{n-1} a_i \beta_1^{n-i} (1 + \alpha)^i}{\sum_{i=0}^{n-1} a_i}. \quad (9)$$

To analyze the possible bias from the use of n period data we should discuss the characteristics of the prospect under consideration.

A. *Neutral Stock*

A stock with $\beta_1 = 1$ is defined as neutral stock. In this case $\alpha = 0$ and $\tilde{\beta}_n = \beta_1^n = 1$. Thus for a neutral stock we obtain the same regression coefficient regardless of the investment horizon.

B. *Aggressive Stock*

Aggressive stock is expected to go upward more than the market portfolio in a bull market and downward more than the market in a bear market, i.e., $\beta_1 > 1$. In this case $\alpha > 0$ and therefore $\sum a_i(1 + \alpha)^i / \sum a_i > 1$ and since $\beta_1^{n-i} \geqslant \beta_1$, $i = 0, \ldots, n-1$, $\tilde{\beta}_n > \beta_1$.[3]

C. *Defensive Stock*

In this case $\beta_1 < 1$, $\alpha < 0$, $\beta_1^{n-i} \leqslant \beta_1$, $i = 0, \ldots, n-1$. Therefore, $\sum a_i(1 + \alpha)^i / \sum a_i < 1$ and

$$\tilde{\beta}_n = \frac{\sum a_i (1 + \alpha)^i \beta_1^{n-i}}{\sum a_i} \leqslant \beta_1 \frac{\sum a_i (1 + \alpha_i)}{\sum a_i} \leqslant \beta_1$$

where equality holds only for $\beta_1 = 1$.

III. The Bias in Measuring the Systematic Risk Using Data for a Shorter Period than the "True" Horizon

In this case the market equilibrium equation written in terms of the one-period parameters is given by

$$\mu_1^n = r^n + \beta_n[\mu_m^n - r^n] \quad (10)$$

where β_n is the true coefficient, which indicates the systematic risk of the stock under consideration. However, in the empirical test one uses data of a shorter period, say, one-period. Hence we calculate $\tilde{\beta}_1 = \text{cov}(x_1, x_m)/\sigma_m^2$ rather than β_n where the tilde denotes the estimated rather than the "true" coefficient. In order to measure the possible deviation between $\tilde{\beta}_1$ and β_n, let us examine first the relationship between the n-period and the one-period covariance. Since $\text{cov}(x_1 \cdots x_n, x_{m,1} \cdots x_{m,n}) = [\text{cov}(x_1, x_m) + \mu_1 \mu_m]^n - (\mu_1 \mu_m)^n$ we obtain

[3] Note that for $n = 1$, we obtain $\tilde{\beta}_n = \beta_1$.

$$\operatorname{cov}(x_1, x_m) = [\operatorname{cov}(x_1 \cdots x_n, x_{m,1} \cdots x_{m,n}) + (\mu_1 \mu_m)^n]^{1/n} - \mu_1 \mu_m.$$

Recalling the definition of β_n and the nth period variance we get

$$\operatorname{cov}(x_1 \cdots x_n, x_{m,1} \cdots x_{m,n}) = \beta_n[(\sigma_m^2 + \mu_m^2)^n - \mu_m^{2n}]. \quad (11)$$

Hence,

$$\operatorname{cov}(x_1, x_m) = \{\beta_n[(\sigma_m^2 + \mu_m^2)^n - \mu_m^{2n}] + (\mu_1 \mu_m)^n\}^{1/n} - \mu_1 \mu_m \quad (12)$$

$$\tilde{\beta}_1 = \{\{\beta_n[(\sigma_m^2 + \mu_m^2)^n - \mu_m^{2n}] + (\mu_1 \mu_m)^n\}^{1/n} - \mu_1 \mu_m\}/\sigma_m^2. \quad (13)$$

The equilibrium market condition is

$$\mu_1^n = r^n + \beta_n(\mu_m^n - r^n) = \mu_m^n + (\beta_n - 1)(\mu_m^n - r^n). \quad (14)$$

If $\beta_n = 1$, then $\mu_1 = \mu_m$, and

$$\tilde{\beta}_1 = \frac{\{(\sigma_m^2 + \mu_m^2)^n - \mu_m^{2n} + \mu_m^{2n}\}^{1/n} - \mu_m^2}{\sigma_m^2} = 1 = \beta_n.$$

If we deal with an aggressive stock, i.e., $\beta_n > 1$, we get, by (10), that $\mu_1 > \mu_m$.

Equation (13) can be rewritten as

$$(\tilde{\beta}_1 \sigma_m^2 + \mu_1 \mu_m)^n = \beta_n[(\sigma_m^2 + \mu_m^2)^n - \mu_m^{2n}] + (\mu_1 \mu_m)^n \quad (15)$$

or after cancellation of $(\mu_1 \mu_m)^n$ we find

$$\sum_{i=0}^{n-1} \binom{n}{i} \tilde{\beta}_1^{n-i} (\sigma_m^2)^{n-i} (\mu_1 \mu_m)^i = \beta_n \sum_{i=0}^{n-1} \binom{n}{i} (\sigma_m^2)^{n-i} (\mu_m^2)^i. \quad (16)$$

For $\beta_n > 1$, $\mu_1 > \mu_m$ and, therefore,

$$\sum_{i=0}^{n-1} \binom{n}{i} \tilde{\beta}_1^{n-i} (\sigma_m^2)^{n-i} (\mu_m^2)^i < \beta_n \sum_{i=0}^{n-1} \binom{n}{i} (\sigma_m^2)^{n-i} (\mu_m^2)^i. \quad (17)$$

Let us denote

$$\binom{n}{i} (\sigma_m^2)^{n-i} (\mu_m^2)^i = a_i.$$

Hence,

$$\sum_{i=0}^{n-1} a_i \tilde{\beta}_1^{n-i} < \beta_n \sum_{i=0}^{n-1} a_i$$

or

$$\beta_n > \frac{\sum_{i=0}^{n-1} a_i \tilde{\beta}_1^{n-i}}{\sum_{i=0}^{n-1} a_i}.$$

From this it is clear that $\beta_n > \tilde{\beta}_1$ since if $\tilde{\beta}_1 < 1$ there is nothing to prove; but if $\tilde{\beta}_1 > 1$ then the right hand side, being a weighted average of $\tilde{\beta}_1^i$, $i = 1, \ldots, n$, must be greater than $\tilde{\beta}_1$, i.e.,

$$\beta_n > \frac{\sum_{i=0}^{n-1} a_i \tilde{\beta}_1^{n-i}}{\sum_{i=0}^{n-1} a_i} > \tilde{\beta}_1. \quad (18)$$

For defensive stock $\beta_n < 1$ and $\mu_1 < \mu_m$; therefore,

$$\sum_{i=0}^{n-1} \binom{n}{i} \tilde{\beta}_1^{n-i} (\sigma_m^2)^{n-i} (\mu_m^2)^i > \beta_n \sum_{i=0}^{n-1} \binom{n}{i} (\sigma_m^2)^{n-i} (\sigma_m^2)^i \quad (19)$$

or

$$\sum_{i=0}^{n-1} a_i \tilde{\beta}_1^{n-i} > \beta_n \sum_{i=0}^{n-1} a_i$$

and thus

$$\beta_n < \frac{\sum_{i=0}^{n-1} a_i \tilde{\beta}_1^{n-i}}{\sum_{i=0}^{n-1} a_i}.$$

Again if $\tilde{\beta}_1 > 1$ then trivially $\beta_n < \tilde{\beta}_1$ as $\beta_n < 1$, and if $\tilde{\beta}_1 < 1$ then

$$\beta_n < \frac{\sum_{i=0}^{n-1} a_i \tilde{\beta}_1^{n-i}}{\sum a_i} < \tilde{\beta}_1.$$

Thus, in general, we obtain a reverse symmetrical case to the one discussed before.

IV. The Biases in Measuring Performance Indices by Deviation from the "True" Investment Horizon

As has been suggested by Treynor (1965), the relevant performance index of individual secu-

rities as well as a portfolio is given by the reward to volatility index,

$$I = \frac{\mu_j - r_F}{\beta_j}. \tag{20}$$

In this section we analyze the possible error stemming from choosing an investment horizon that differs from the "true" investment horizon. We will start with the case where the horizon used in the empirical data is longer than the "true" horizon and later on we will analyze the opposite case.

Assuming that the investment horizon is one-period, we expect in equilibrium that

$$I_1 = \frac{\mu_j - r_F}{\beta_j} = c \text{ for all } j. \tag{21}$$

(We shall delete the subscript F in cases where no confusion arises.) Thus $\mu_j = \beta_j c + r$. Suppose now that we measure the performance index using a longer period than the horizon taken into account in the equilibrium model; i.e., we calculate the following n-period performance index:

$$I_n = \frac{\mu_j^n - r^n}{\beta_{jn}}. \tag{22}$$

where

$$A = \sum_{i=0}^{n-1} \binom{n}{i} (\sigma_m^2)^{n-i} (\mu_m^2)^i.$$

Hence, the n-period performance measure is

$$I_n = \frac{\mu_j^n - r^n}{(\beta_j \sigma_m^2 + \mu_j \mu_m)^n - (\mu_j \mu_m)^n} A. \tag{24}$$

We would now like to find whether the "new" index will preserve the property of being identical for all securities using the n-period data. We shall prove that $\partial I_n / \partial \beta_{1j} < 0$; thus the higher the β_{1j}, the smaller the measured performance index will be. Notice that A is independent of β_{1j}, and hence one can calculate $(\partial I_n / A)/\partial \beta_j$. Substituting $\beta_j c + r$ for μ_j and ignoring the subscript 1 we get

$$I_n / A = \frac{(\beta_j c + r)^n - r^n}{\sum_{i=0}^{n-1} \binom{n}{i} \beta_j^{n-i} (\sigma_m^2)^{n-i} (\beta_j c + r)^i \mu_m^i}$$

and, expanding the numerator and the denominator,

$$I_n / A = \frac{\sum_{k=0}^{n-1} \binom{n}{k} \beta_j^{n-k} c^{n-k} r^k}{\sum_{i=0}^{n-1} \binom{n}{i} \beta_j^{n-i} (\sigma_m^2)^{n-i} \mu_m^i \sum_{s=0}^{i} \binom{i}{s} \beta_j^{i-s} c^{i-s} r^s}$$

or

$$I_n / A = \frac{\sum_{k=0}^{n-1} \binom{n}{k} \beta_j^{n-k} c^{n-k} r^k}{\sum_{i=0}^{n-1} \sum_{s=0}^{i} \binom{n}{i} \binom{i}{s} \beta_j^{n-s} (\sigma_m^2)^{n-i} \mu_m^i c^{i-s} r^s}. \tag{25}$$

Using equation (6) the denominator can be written as

$$\beta_{jn} = \frac{\sum_{i=0}^{n-1} \binom{n}{i} \beta_{1j}^{n-i} (\sigma_m^2)^{n-i} (\mu_j \mu_m)^i}{\sum_{i=0}^{n-1} \binom{n}{i} (\sigma_m^2)^{n-i} (\mu_m^2)^i} \tag{23}$$

$$= \frac{(\beta_{1j} \sigma_m^2 + \mu_j \mu_m)^n - (\mu_j \mu_m)^n}{A}.$$

We would like to prove that $\partial I_n / \partial \beta_j < 0$; that is, when we use a period longer than the proper horizon data we get a systematic bias in the performance measures. Thus, for a more aggressive stock (i.e., the higher β_j) the lower is the performance measure. Hence, the measured performance index, which is biased, will discriminate against aggressive stocks. Differentiating (25) and denoting the denominator by D we get

$$\begin{aligned}\frac{D^2}{A}\frac{\partial I_n}{\partial \beta_j} = &\left[\sum_{i=0}^{n-1}\binom{n}{k}(n-k)\beta_j^{n-k-1}c^{n-k}r^k\right]\\ &\left[\sum_{i=0}^{n-1}\sum_{s=0}^{i}\binom{n}{i}\binom{i}{s}\beta_j^{n-s}(\sigma_m^2)^{n-i}\mu_m^i c^{i-s}r^s\right]\\ &-\left[\sum_{k=0}^{n-1}\binom{n}{k}\beta_j^{n-k}c^{n-k}r^k\right]\\ &\left[\sum_{i=0}^{n-1}\sum_{s=0}^{i}\binom{n}{i}\binom{i}{s}(n-s)\beta_j^{n-s-1}(\sigma_m^2)^{n-i}\mu_m^i c^{i-s}r^s\right]\end{aligned} \tag{26}$$

or

$$\begin{aligned}&\frac{D^2}{A}\frac{\partial I_n}{\partial \beta_j}\\ &= \sum_{k=0}^{n-1}\sum_{i=0}^{n-1}\sum_{s=0}^{i}\binom{n}{k}\binom{n}{i}\binom{i}{s}(n-k)\beta_j^{2n-k-s-1}c^{n+i-k-s}r^{k+s}(\sigma_m^2)^{n-i}\mu_m^i \\ &- \sum_{k=0}^{n-1}\sum_{i=0}^{n-1}\sum_{s=0}^{i}\binom{n}{k}\binom{n}{i}\binom{i}{s}(n-s)\beta_j^{2n-k-s-1}c^{n+i-k-s}r^{k+s}(\sigma_m^2)^{n-i}\mu_m^i\end{aligned} \tag{27}$$

or

$$\frac{D^2}{A}\frac{\partial I_n}{\partial \beta_j} = \sum_{k=0}^{n-1}\sum_{i=0}^{n-1}\sum_{s=0}^{i}\binom{n}{k}\binom{n}{i}\binom{i}{s}(s-k)\beta_j^{2n-k-s-1}c^{n+i-k-s}r^{k+s}(\sigma_m^2)^{n-i}\mu_m^i. \tag{28}$$

Note that this is a polynomial in β_j, c, r, σ_m^2, μ_m, and for $k + s = \theta$ and any i we get the same polynomial term. Hence to prove that $(D^2/A)(\partial I/\partial \beta_j) < 0$ we shall prove that the coefficient of this polynomial term is negative. For, given θ and i, the coefficient of $\beta_j^{2n-\theta-1}c^{n+i-\theta}r^\theta(\sigma_x^2)^{n-i}\mu_m^i$ will be shown to be negative for all θ and i. Let us now distinguish between two cases: (i) $\theta \leqslant i$ and (ii) $\theta > i$. The term $\binom{n}{i}$ is, of course, positive and can be ignored. In case (i) the coefficient of the polynomial term (denoted by $\Gamma_{i,\theta}$) is given by (substituting $s = \theta - k$)

$$\begin{aligned}\Gamma_{i,\theta} &= \sum_{k=0}^{\theta}\binom{n}{k}\binom{i}{\theta-k}(s-k)\\ &= \sum_{k=0}^{\theta}\binom{n}{k}\binom{i}{\theta-k}(\theta-2k)\\ &= \theta\sum_{k=0}^{\theta}\binom{n}{k}\binom{i}{\theta-k}\\ &\quad - 2\sum_{k=0}^{\theta}\binom{n}{k}\binom{i}{\theta-k}k.\end{aligned}$$

Recalling that for $k = 0$ the first term of the second summation vanishes and we obtain

$$\begin{aligned}\Gamma_{i,\theta} &= \theta\sum_{k=0}^{\theta}\binom{n}{k}\binom{i}{\theta-k}\\ &\quad - 2\sum_{k=1}^{\theta}\binom{n}{k}\binom{i}{\theta-k}\cdot k.\end{aligned}$$

However, using

$$\binom{n}{k}\binom{i}{\theta-k}k = n\binom{n-1}{k-1}\binom{i}{\theta-k},$$

we find

$$\begin{aligned}\Gamma_{i,\theta} &= \theta\sum_{k=0}^{\theta}\binom{n}{k}\binom{i}{\theta-k}\\ &\quad - 2n\sum_{k=1}^{\theta}\binom{n-1}{k-1}\binom{i}{\theta-k}.\end{aligned} \tag{29}$$

Now using the relation

$$\binom{x}{0}\binom{y}{m} + \binom{x}{1}\binom{y}{m-1} + \cdots + \binom{x}{m}\binom{y}{0} = \binom{x+y}{m}$$

(Feller, 1957, p. 62), we get

$$\Gamma_{i,\theta} = \theta\binom{n+i}{\theta} - 2n\binom{n-1+i}{\theta-1}$$
$$= \theta\binom{n+i}{\theta} - \frac{2n\theta}{n+i}\binom{n+i}{\theta}$$
$$= \theta\binom{n+i}{\theta}$$
$$= \theta\binom{n+i}{\theta}\left[1 - \frac{2n}{n+i}\right].$$

But since $i \leqslant n-1$, $1 - (2n/n+i) < 0$ for all i. Thus, for the case $\theta \leqslant i$ we have proved that the coefficient of the polynomial term is negative. Let us now turn to (ii), i.e., the case where $\theta > i$. In this case since $s \leqslant i$, and $k + s = \theta$, k and s can assume the values

$$k = \theta - i, \theta - i + 1, \ldots, \theta$$
$$s = i, i - 1 \ldots, 0.$$

Hence, in this case the polynomial coefficient is given by

$$\sum_{k=\theta-i}^{\theta}\binom{n}{k}\binom{i}{\theta-k}(\theta - 2k).$$

Let us substitute $\omega = k - \theta + i$. Thus the above term can be rewritten as

$$\sum_{\omega=0}^{i}\binom{n}{\omega+\theta-i}\binom{i}{i-\omega}(2i - 2\omega - \theta)$$
$$= 2\sum_{\omega=0}^{i}\binom{n}{\omega+\theta-i}\binom{i}{i-\omega}(i-\omega)$$
$$-\theta\sum_{\omega=0}^{i}\binom{n}{\omega+\theta-i}\binom{i}{i-\omega}.$$

Note that

$$\binom{i}{i-\omega}(i-\omega) = i\binom{i-1}{i-\omega-1}$$

and again using the same identity (Feller, 1957) we get

$$2i\binom{n+i-1}{\theta-1} - \theta\binom{n+i}{\theta}$$
$$= 2i\binom{n+i-1}{\theta-1} - \theta\binom{n+i-1}{\theta-1}\frac{n+i}{\theta}$$
$$= \binom{n+i-1}{\theta-1}[2i - (n+i)]$$
$$= \binom{n+i-1}{\theta-1}(i-n),$$

which is of course non-positive as $i \leqslant n$.

This finally implies that $(D^2/A)(\partial I_n/\partial\beta_j) < 0$, or $\partial I_n/\partial\beta_j < 0$, as has been maintained. The other possible bias which we believe to be more relevant is that the performance index is calculated for a shorter horizon than the "true" one. In virtually all the empirical research, a horizon period of one year or even less has been assumed, which seems to be shorter than the "true" horizon. In this case we expect to have in equilibrium

$$I_n = \frac{\mu_j^n - r^n}{\beta_{jn}} = c$$

where c is a constant for all securities. When using a shorter horizon one calculates the following performance index, which might differ between securities, $I_1 = (\mu_j - r)/\beta_{j,1}$. We investigate the possible bias by examining the sign $\partial I_1/\partial\beta_j$. We have proved that this term is positive but for brevity we delete the proof.

Another interesting phenomenon is that for a given β_j, or a given β_j^n, the higher the investment horizon for which the performance index is calculated, the higher the measured index. Since this theoretical result is valid independent of the "true" horizon, it can be tested empirically, (see section VI) even though the true investment horizon is unknown.

V. Implication of the Mathematical Bias in the Empirical Evidence

Douglas, Lintner, Miller and Scholes, Jensen, and Black-Jensen-Scholes indicate the following three major inconsistencies between the capital asset pricing model and the empirical evidence.

A. *The Residual Variance*

The most disturbing empirical result is the fact that the residual variance is positively correlated with the average return. To be more specific, using cross-section data the following regression has been estimated:

$$\overline{R}_j = \hat{\gamma}_0 + \hat{\gamma}_1\hat{\beta}_j + \hat{\gamma}_2\hat{S}_{e,j}^2$$

where $\overline{R}_j$ is the mean rate of return of the jth

security in excess of the risk free interest rate; $\hat{\beta}_j$ is the systematic risk of the jth security estimated from time series, using the following regression:

$$R_{jt} = \alpha_j + \hat{\beta}_j R_{mt} + e_{jt}$$

where R_{mt} is the rate of return on the market portfolio in excess of the risk free interest rate; $S_{e,j}^2$ is the residual variance taken from the time series regression.

Using the capital asset pricing model one expects (in equilibrium) that $\gamma_0 = 0$, $\gamma_1 = \bar{R}_m$ and $\gamma_2 = 0$. However, virtually all empirical evidence indicates that γ_0 and γ_2 are significantly positive.

According to the capital asset pricing model, γ_2 should be equal to zero since the jth security variance (and hence its residual variance) is a very small component of its total risk, and even this small component is already taken into account in the systematic risk, i.e., in the $\hat{\beta}_j$ coefficient. We claim that the fact that γ_2 is significantly positive does not necessarily indicate that the capital asset pricing model is wrong but merely that the investment horizon assumed in the empirical research mentioned above is shorter than the "true" investment horizon.[4] In this case the relationship between the true systematic risk (β_j) and the estimated systematic risk $\hat{\beta}_j$ is illustrated in figure 1. Thus one can write

$$\beta_j = \hat{\beta}_j + (\beta_j - \hat{\beta}_j) = \hat{\beta}_j + \delta$$

where $\delta < 0$ for all values $\beta_j < 1$ and $\delta > 0$ for all values $\beta_j > 1$. The jth security variance $\sigma_{R,j}^2$ can be decomposed to the following two components

$$\sigma_{R,j}^2 = \beta_j^2 \sigma_{R,m}^2 + S_{e,j}^2$$

and substituting $\beta_j = \hat{\beta}_j + \delta$ we obtain

$$\sigma_{R,j}^2 = (\hat{\beta}_j + \delta)^2 \sigma_{R,m}^2 + S_{e,j}^2$$
$$= \hat{\beta}_j^2 \sigma_{R,m}^2 + \sigma_{R,m}^2(\beta_j^2 - \hat{\beta}_j^2) + S_{e,j}^2.$$

Thus,

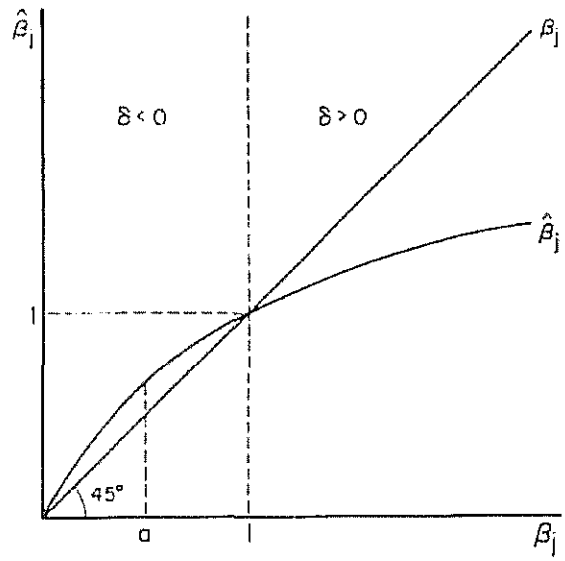
$$\sigma_{R,j}^2 = \hat{\beta}_j^2 \sigma_{R,m}^2 + \hat{S}_{e,j}^2$$

where

$$\hat{S}_{e,j}^2 \equiv S_{e,j}^2 + (\beta_j^2 - \hat{\beta}_j^2)\sigma_{R,m}^2.$$

The estimated residual variance of the time series regression, $\hat{S}_{e,j}$, is biased and functionally related to the jth security systematic risk. Moreover, we claim that $\hat{S}_{e,j}^2$ and β_j are positively correlated. Looking again at figure 1 we can see

FIGURE 1. — ESTIMATED SYSTEMATIC RISK ($\hat{\beta}_j$) AS A FUNCTION OF THE TRUE SYSTEMATIC RISK (β_j)



that through most of the relevant domain $\hat{S}_{e,j}$ and β_j increase simultaneously. To the right of point a, δ as well as β_j increase, hence $\hat{S}_{e,j}^2$ and β_j increase together. The opposite holds to the left of point a. But this part is relatively small and cannot change the positive correlation. Moreover, this part will include only a few securities since most securities are characterized by a systematic risk greater than 0.5.[5]

B. *Bias in the Regression Coefficients*

The second bias mentioned in the literature is that using the cross-section regression

$$\bar{R}_j = \hat{\gamma}_0 + \hat{\gamma}_1 \hat{\beta}_j + e_j$$

(where again $\hat{\beta}_j$ is taken from an estimate using

[4] In all the empirical research mentioned in this paper, the assumed investment horizon is less or equal to 1 year (1 week, 1 month, 3 months, and 1 year).

[5] It is worth mentioning that Miller and Scholes (1972, p. 62) find a positive correlation between the residual variances and the estimated systematic risk.

time series data.) It has been found that γ_0 is significantly positive and γ_1 is less than what one would expect. That is, the $\hat{\gamma}_1$ estimate turns out to be significantly less than the market portfolio average rate of return in excess of the risk-free rate.

We claim that the biases in γ_0 and γ_1 may stem from errors in the choice of a proper horizon, causing a systematic error in the estimated values of β_j. This may be explained as follows:

$$\hat{\beta}_j = \beta_j - \delta_j$$

and using our results on the consistent mathematic bias (see figure 1), $\text{cov}[\beta, \delta] > 0$. Thus, the regression estimate is

$$\hat{\gamma}_1 = \frac{\sum (\bar{R}_j - \bar{\bar{R}})(\hat{\beta}_j - \bar{\hat{\beta}})}{\sum (\hat{\beta}_j - \bar{\hat{\beta}})^2}$$

where $\bar{\hat{\beta}} = \bar{\beta} - \bar{\delta}$. Hence,

$$\hat{\gamma}_1 = \frac{\sum (\bar{R}_j - \bar{\bar{R}})[\beta_j - \bar{\beta} - \delta_j + \bar{\delta}]}{\sum (\beta_j - \bar{\beta} - \delta_j + \bar{\delta})^2}$$

or

$$\hat{\gamma}_1 = \frac{\sum (\bar{R}_j - \bar{\bar{R}})(\beta_j - \bar{\beta}) - \sum (\bar{R}_j - \bar{\bar{R}})(\delta_j - \bar{\delta})}{\sum (\beta_j - \bar{\beta})^2 + \sum (\delta_j - \bar{\delta})^2 - 2 \sum (\beta_j - \bar{\beta})(\delta_j - \bar{\delta})}$$

and, dividing the numerator and the denominator by $\sum (\beta_j - \bar{\beta})^2$ and looking at the probability limit, we get

$$\text{plim } \hat{\gamma}_1 = \frac{\gamma_1 - \dfrac{\text{cov}(\bar{R}_j, \delta_j)}{\sigma_\beta^2}}{1 + \dfrac{\sigma_\delta^2}{\sigma_\beta^2} - 2\dfrac{\text{cov}(\beta, \delta)}{\sigma_\beta^2}}.$$

Since $\text{cov}[\bar{R}, \delta] = \gamma_1 \cdot \text{cov}[\beta, \delta]$ we get

$$\text{plim } \hat{\gamma}_1 = \gamma_1 \frac{1 - \dfrac{\text{cov}(\beta, \delta)}{\sigma_\beta^2}}{1 - \dfrac{2\,\text{cov}(\beta, \delta)}{\sigma_\beta^2} + \dfrac{\sigma_\delta^2}{\sigma_\beta^2}}.$$

Clearly, if $\text{cov}(\beta, \delta) = 0$, i.e., if there is a *random* error in the measurement of β_j, then $\text{plim } \hat{\gamma}_1 = \gamma_1(1/1 + (\sigma_\delta^2/\sigma_\beta^2))$. This means that $\hat{\gamma}_1$ is biased downward (see also Miller and Scholes, 1972). Using our analysis and assuming that the data are collected for shorter than the true horizon, then $\text{cov}(\delta, \beta) > 0$, and we cannot be sure of the direction of the bias. The empirical evidence lends support to the hypothesis that $(\sigma_\delta^2/\sigma_\beta^2) - (\text{cov}(\beta, \delta)/\sigma_\beta^2) > 0$, and $(\text{cov}(\beta, \delta)/\sigma_\beta^2) < 1$, and hence $\hat{\gamma}_1$ is biased downward.

The possibility that $\hat{\gamma}_0$ is significantly positive follows immediately. The exact relationship is

$$\bar{R}_j = \gamma_0 + \gamma_1 \beta_j + e_j$$

while the measured regression is

$$\bar{R}_j = \hat{\gamma}_0 + \hat{\gamma}_1 \hat{\beta}_j + \hat{e}_j.$$

Taking the probability limits of the averages of both sides of the above two regressions we get

$$\gamma_0 + \gamma_1 \bar{\beta}_j = \hat{\gamma}_0 + \hat{\gamma}_1 \bar{\hat{\beta}}_j.$$

But since $\hat{\beta}_j = \beta_j - \delta_j$, this yields

$$\gamma_0 + \gamma_1 \bar{\beta}_j = \hat{\gamma}_0 + \hat{\gamma}_1 \bar{\beta}_j - \hat{\gamma}_1 \bar{\delta}.$$

But $\hat{\gamma}_1 < \gamma_1$ and looking again at figure 1 one may realize that $\delta > 0$ hence $\hat{\gamma}_0 > \gamma_0$ as has been found in the empirical papers mentioned above.

C. *Measuring Ex Post Performance*

Black, Jensen and Scholes (1972) examined the following time series regression:

$$R_{jt} = \alpha_j + \beta_j R_{mt} + e_{jt}$$

where α_j indicates the performance of the jth portfolio. It has been found that in three periods α_j is positive for low risk securities ($\beta_j < 1$) and $\alpha_j < 0$ for high risk securities ($\beta_j > 1$). In the fourth subperiod analyzed the opposite relationship was found. B-J-S explain the results of three out of the four periods by the "two-factor model." Namely, they maintain that α_j estimates, in this case the term $E(\tilde{\gamma}_z)(1 - \beta_j)$ where $\tilde{\gamma}_z$ is an asset whose return is uncertain but whose covariance with the market portfolio

is zero. The fourth period results, which cannot be explained by the "two factor model," are quite easily explained in our framework. If we assume that the horizon used in the empirical research is too short, then the following regression is analyzed:

$$R_{jt}^{(1)} = \alpha_j^{(1)} + \beta_j^{(1)} R_{mt} + e_{jt}^{(1)}$$

where the superscript denotes the length of the horizon, assumed to be one time-unit. Dividing through by $\beta_j^{(1)}$ we get

$$\frac{\overline{R}_{jt}^{(1)}}{\beta_j^{(1)}} = \frac{\alpha_j^{(1)}}{\beta_j^{(1)}} + \overline{R}_m .$$

As $\overline{R}_j^{(1)}/\beta_j^{(1)} = I_1$ and we have proved that $\partial I_1/\partial \beta_j^{(n)} > 0$, we learn that on the right hand side $\partial(\alpha_j^{(1)}/\beta_j^{(1)})/\partial \beta_j^{n} > 0$ must hold as $\overline{R}_m$ is independent of $\beta_j^{(n)}$. Increasing $\beta_j^{(n)}$ (and holding all other parameters constant), $\beta_j^{(1)}$ also increases and hence $\alpha_j^{(1)}$ must increase with $\beta_j^{(n)}$. For the value $\beta_j^{(n)} = 1$, $\beta_j^{(1)}$ is also unity; hence there is no mathematical bias and in equilibrium we expect α_j to be equal to zero. Thus, we expect to find negative α_j for low risk securities ($\beta_j < 1$) and positive α_j for high risk securities ($\beta_j > 1$) as has been found by B-J-S in one of the subperiods that they studied.

We would like to note that when the data collected are for longer than the true horizon, the results are ambiguous. In this case we proved that $\partial I^{(n)}/\partial \beta_j^{(1)} < 0$, i.e., $R_{jt}^{(n)}/\beta_j^{(n)} = \alpha_j^{(n)}/\beta_j^{(n)} + \overline{R}_m^{(n)}$ and $\partial(\alpha_j^{(n)}/\beta_j^{(n)})/\partial \beta_j^{(1)} < 0$ since $\overline{R}_m^{(n)}$ is independent of $\beta_j^{(1)}$. Unlike the previous case, it is not clear whether $\alpha_j^{(n)}$ decreases with $\beta_j^{(1)}$ since $\beta_j^{(n)}$ increases with $\beta_j^{(1)}$ and it might be that $\alpha_j^{(n)}$ also increases, but at a slower rate. Thus, in the regression of the type

$$R_{jt}^{(n)} = \alpha_j^{n} + \beta_j^{n} R_{mt}^{(n)} + e_{jt}^{(n)}$$

one cannot predict the sign of $\hat{\alpha}_j^{(n)}$ merely on the basis of the mathematical consideration given above.

In fact, there are several forces causing a deviation between the theoretical results and the empirical findings. The possibility that riskless assets do not exist causes a bias in one direction. Choosing a relatively short horizon causes a bias in an opposite direction. These two forces may explain the fact that sometimes α_j is related negatively to β_j and sometimes related positively to β_j.

VI. The Empirical Findings

The monthly rates of return for a sample of 101 stocks traded on the New York Stock Exchange were calculated for the period 1948-68. That is, for each security there are 240 observations: thus if $R^*_{i,1}, R^*_{i,2}, \ldots, R^*_{i,240}$ are monthly rates of return on the ith security, one can calculate the bimonthly rates of return,

$$R^*_{i,1}, R^*_{i,2}, \ldots, R^*_{i,120}$$

where $(1 + R_{i,1})(1 + R_{i,2}) = 1 + R^*_{i,1}$, $(1 + R_{i,3})(1 + R_{i,4}) = 1 + R^*_{i,2}$, etc. Note that by using a horizon of two months, although we subdivide the period 1948-68 to 120 time units rather than 240 time units, there is no change in the length of the period covered by the empirical research, which remains twenty years. Similarly, if we use annual rates of return we have only twenty observations.[6]

In order to examine the theoretical findings, one should identify defensive stocks and aggressive stocks and examine how the estimate of the systematic risk of each stock varies with the investment horizon. Since many stocks are characterized by a systematic risk which is close to 1, we illustrate the relationship between the systematic risk and the investment horizon with ten stocks that were undoubtedly defensive ($\beta_j < 1$) and ten stocks that were clearly aggressive ($\beta_j > 1$).[7] Table 1 indicates that the estimates $\hat{\beta}_{j(n)}$ of the ten defensive stocks generally decrease with n, as expected by theoretical findings. Moreover, the change in the estimate of the systematic risk is so sharp that some of the defensive stocks are characterized by a

[6] As a proxy to the market portfolio we use a Fisher arithmetic index. The rates of return on riskless assets for shorter periods are taken to be the rates of return on Treasury bills, while for longer periods the rates of return on riskless assets are estimated by the yield to maturity on Government bonds in accordance with the assumed investment horizon. The rates of return on Treasury bills as well as on Government bonds were taken from various issues of the Federal Reserve Bulletin. The sample of shares was taken from the return file of the CRSP tapes.

[7] Out of the 101 stocks included in the sample only 10 stocks have systematic risk that is below 1 for almost all the horizons and about the same number have been found with systematic risk greater than 1 for almost all horizons.

TABLE 1. — THE ESTIMATE OF THE SYSTEMATIC RISK OF TEN DEFENSIVE STOCKS

Horizon (in months)	Idaho Power Corp.	American Can Corp.	National Dairy Products	P. Lorillard Corp.	American Tobacco	Borden, Inc.	Abbott Laboratory	Standard Brands	Grey-hound Corp.	Continental Can
1	.4282	.5167	.5281	.6166	.6296	.6172	.6576	.6650	.6752	.7807
2	.4012	.4886	.4655	.5711	.4652	.5912	.5717	.6147	.6651	.7081
3	.3796	.3755	.4475	.3496	.4993	.5684	.5892	.5578	.5773	.6520
4	.3329	.3311	.3400	.4881	.3697	.6142	.5284	.6397	.5340	.6109
5	.1881	.2631	.4428	.2604	.3283	.3449	.6319	.4331	.6709	.5882
6	.3862	.3402	.4119	.4253	.3706	.4330	.3811	.6112	.5294	.6616
8	.4322	.0621	.5309	.4815	.3020	.4627	.2398	.7087	.4907	.3967
10	.2312	.1236	.4777	—.0656	.2438	.4272	.4729	.5325	.4800	.4209
12	.2367	—.0118	.3511	—.4615	.0364	.3390	.4227	.4289	.6188	.2834
15	.1556	.0702	.4544	—1.0612	—.0365	—.0561	.1243	.2008	.1541	.3526
16	.3016	.2049	.5016	—1.0387	.1400	.2723	.1463	.7473	.1719	.4753
20	.1142	—.2563	.3283	—1.1855	—.1060	.2336	.0247	.4002	.2378	.3307
24	.1068	—.2690	.3996	—2.0036	.1657	.0849	.2474	.3771	.7826	.5319
30	.2210	.0101	.2781	—2.8251	.1187	.1360	—.3863	—.0150	—.5545	.3536

negative regression coefficient for large investment horizons.

Table 2 demonstrates the response of the systematic risk estimate for ten aggressive stocks to changes in the assumed investment horizon. For example, the estimate of the systematic risk of Evans-Products increases from 1.82 for a horizon of one month, to 3.49 for a twelve months horizon and to more than 8 for a thirty month horizon. Hence, the estimated required rate of return and the determination of which stock is underpriced and which is overpriced, is a function of the assumed investment horizon.

Another unambiguous theoretical result is that the performance measure for the ith security increases with the investment horizon n. Table 3 shows that the empirical findings conform perfectly with the theoretical results. The average performance measure increases from 0.012 when it is calculated on the basis of monthly rates of return to 1.683 should one calculate rates of return for time units of thirty months. Note that in accordance with the theoretical results, $\bar{I}_{j(n)}$ increases with n for aggressive as well as for defensive stocks. The average performance index for defensive stocks is significantly greater than the average performance measure for the aggressive stocks, over all n. Nevertheless, as expected, I_j increases with n for both categories.

The performance index I_j is an important index employed in evaluating the ex post performance of individual securities or mutual funds. We have seen that the index I_j increases with n. However, if the performance index increases at the same rate for all evaluated securities, the damage induced by selecting the investment horizon arbitrarily would be very small, since the ranking of the securities accord-

TABLE 2. — THE ESTIMATE OF THE SYSTEMATIC RISK OF TEN AGGRESSIVE STOCKS

Horizon (in months)	Evans Products Corp.	Cerro Corp.	Colt Industries	Conelco Inc.	Anaconda Corp.	Bethlehem Steel	United States Steel	Carpenter Steel	Hooker Chemical	Medusa Portland
1	1.8252	1.5876	1.5353	1.4091	1.2904	1.1664	1.1657	1.1093	1.0650	1.0108
2	1.8002	1.6050	1.7738	1.5725	1.3379	1.1525	1.1778	1.0790	1.1205	1.1431
3	2.1009	1.5389	2.0234	1.6586	1.1890	1.2046	1.1777	1.1384	1.1241	1.1164
4	1.9293	1.6299	1.8991	1.7332	1.3045	.9997	1.0952	1.0051	1.1947	1.3197
5	1.8935	1.6583	2.2407	2.2077	1.1977	1.0340	1.1037	.9627	1.1016	1.2757
6	1.9414	1.6056	2.1994	2.2806	1.2368	1.1495	1.2811	1.1078	1.0907	1.0505
8	2.0450	1.7603	1.7112	2.6780	1.2632	1.1889	1.2164	1.1081	1.2390	1.5629
10	2.0446	1.4159	2.3771	3.2640	1.1683	1.2068	1.1751	1.1476	1.1098	1.4558
12	3.4909	1.6676	1.3451	3.4419	1.4399	1.3672	1.6645	1.6671	1.4250	1.8580
15	3.2813	2.0732	1.5511	2.2901	1.5936	1.9100	1.7893	1.8270	1.3184	2.7472
16	2.2692	1.8526	1.8232	3.4655	1.3207	1.5274	1.4989	1.0993	1.4488	2.0514
20	1.8549	2.0076	4.0385	3.0123	.9667	1.2587	1.0467	.6608	1.1361	1.7524
24	4.1552	3.1044	2.2391	2.6956	1.5806	1.2426	1.3561	1.4828	1.6444	1.2072
30	8.1045	4.0495	2.2277	5.2007	2.1563	3.6042	2.9569	1.7334	1.4476	3.2152

TABLE 3.—THE AVERAGE PERFORMANCE INDEX (I_j)

Investment Horizon	All Stocks (101 stocks)	All Defensive Stocks (61 stocks)	All Aggressive Stocks (40 stocks)
1	0.012	.014	.009
2	0.025	.031	.018
3	0.039	.045	.035
4	0.054	.064	.039
5	0.073	.088	.046
6	0.085	0.100	.051
8	0.120	0.150	.076
10	0.146	0.164	.099
12	0.133	0.152	.109
15	0.159	0.175	.131
16	0.265	0.336	.169
20	0.575	0.778	.214
24	0.526	0.709	.221
30	1.683	3.281	.234

ing to their ex post success should not change when one varies the investment horizon. Unfortunately, I_j does not change at the same rate, and the performance measure is a direct function of the selected investment horizon.

Using ex post data we ran the following regression for various investment horizons,

$$\overline{R}_j = \gamma_0 + \gamma_1 \hat{\beta}_j + \gamma_2 \hat{S}_{e,j}^2$$

where $\hat{\beta}_j$ and the residual variance $\hat{S}_{e,j}^2$ are taken from the first-pass regression. If the capital asset pricing model provides satisfactory results we expect $\gamma_0 = 0$, $\gamma_1 = \overline{R}_m$, $\gamma_2 = 0$, and a relatively high R^2. We found that γ_0 is positive, and that one cannot eliminate this factor by changing the investment horizon. For horizons less than two months γ_2 is not significantly different from zero and the R^2 is quite low. For horizons of three and four months, we conclude that γ_2 is not significant, a fact that is consistent with the capital asset pricing model. However, in this case the R^2 values are only 13% and 15%, respectively, and hence only a small part of the variability is explained, an unsatisfactory feature. For larger horizons ($n \geqslant 5$), both γ_2 and γ_1 are positive and significantly different from zero.

Thus, by changing the investment horizon we could not reconcile this disparity between the capital asset pricing model and the empirical findings. However, in spite of the fact that we cannot find a horizon that yields satisfactory results, the coefficients γ_0, γ_1, γ_2 are not invariant to the investment horizon. In particular, note that R^2 increases with the horizon; it is only 6% for a one month horizon and 34%-43% for horizons greater than one year.

Concluding Remarks

Theoretical models and empirical results may differ for many possible reasons. However, in some respects, the empirical evidence taken from the market deviates systematically from the theoretical model, a phenomenon that lends support to the hypothesis that there is some common error in the way in which empirical research has been carried out. Indeed, we have found that the investment horizon for which data are collected plays a crucial role and has a great impact on both the regression coefficients and the performance indices. The investment horizon, implicitly assumed in the empirical research, ranges from one week to one year. When the "true" investment horizon is no longer than the investment horizon used in the empirical research, one can explain many of the deviations between the empirical results and the capital asset pricing model by the mathematical bias stemming from the fact that relatively short horizons have been selected.

Using an investment horizon shorter than the true horizon causes a bias in the systematic risk. However, the closer the true systematic risk is to 1, the smaller the deviation. Since this deviation is positively correlated with the residual variance of each security, it also provides a theoretical explanation for the fact that the coefficient of the residual variance may be significantly positive.

Deviation from the true horizon also explains the empirical biases in the intercept as well as in the coefficient of the systematic risk, as measured from a cross-section regression.

Ranking securities according to the Reward to Volatility is also biased as this performance index is functionally related to the investment horizon assumed. Thus, one security might show a better performance than another on the basis of monthly data, while the opposite relationship would be obtained if yearly data had been selected.

From the empirical part of this paper we can draw two main conclusions. First, the assumed investment horizon has a crucial impact on estimates of the systematic risk of securities as

well as on their performance indices over the 1948-68 period. The systematic risk of defensive stocks tends to decline while that for aggressive stocks tends to increase with increases in the investment horizon. The impact of an increase in that horizon on the performance index is in the same direction for aggressive and defensive stocks; the higher the assumed investment horizon, the higher the performance index. The second conclusion has rather negative implications. By trying various horizons we could not reconcile the gap between the capital asset pricing model and empirical evidence. In second pass regressions the intercept is positive and significant for most horizons, the coefficient for systematic risk is smaller than hypothesized by the CAPM and the coefficient for the residual variance is positive and significant. Yet, the estimate of the parameters as well as the coefficient of correlation are functions of the assumed investment horizon and one cannot choose arbitrarily the time units for which rates of return are calculated.

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by Gabriel Hawawini

Why Beta Shifts as the Return Interval Changes

A security's beta may vary substantially depending upon whether it is estimated on the basis of daily, weekly or monthly returns. For instance, for the four-year period January 1970 to December 1973, Eastman Kodak had a beta of 1.25 based on daily returns, but a beta of 0.93 based on monthly returns.

In general, the betas of securities with a smaller market value than the average of all securities outstanding (the market) will decrease as the return interval is shortened, whereas the betas of securities with a large market value relative to the market will increase. This suggests that betas measured over return intervals of arbitrary length will tend to be biased. In particular, securities with relatively small market values may appear to be less risky than they truly are, whereas securities with relatively large market values may appear to be more risky than they truly are.

MODERN Portfolio Theory (MPT) tells us that a security's historical rates of return can be used to estimate its systematic risk—i.e., its beta coefficient. MPT does not tell us, however, if these rates of return should be measured over a day, a week, a month or any other length of time. Several researchers have shown that this is not a trivial issue.[1]

Beta coefficients have in the past generally been estimated using monthly returns, mainly because these data were the most readily available.[2] Today, betas may be estimated using weekly or even daily returns. As Table I indicates, however, a beta based on daily returns may differ substantially from a beta based on monthly returns, even if both betas are estimated over the same fixed calendar period.

Here, betas for the four-year period January 1970 to December 1973 have been estimated using 50 monthly returns, 1,009 daily returns and various combinations of weekly returns. For the first firm on the list, Wayne-Gossard, beta based on monthly returns is 0.976, whereas the daily beta is 0.459—a decrease of about 53 per cent. In the case of Eastman Kodak, on the other hand, beta increases by 34 per cent—from 0.932 to 1.251—when the return interval is shortened from a month to a day.

This article presents a simple model that explains why estimates of beta depend upon the length of the return interval. This model also predicts the direction and the strength of the variations in estimated betas. Specifically, I will show that when the return interval is shortened, the following occurs: Securities with a smaller market value than the average of all securities outstanding (the market) will generally have a decreasing beta, whereas securities with a larger market value than the average of all securities outstanding will generally have an increasing beta. The implications of this phenomenon for portfolio performance evaluation and estimation of the Security Market Line and the market price of risk reduction are examined.

What Causes Beta to Shift?

A major factor responsible for shifts in estimated betas is the existence of intertemporal (noncontemporaneous) relationships between the daily returns of individual securities and those of the

1. Footnotes appear at end of article.

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Table I Betas Estimated on the Basis of Various Return Intervals*
(January 1970–December 1973)

	Monthly Beta	*Triweekly Beta*	*Biweekly Beta*	*Weekly Beta*	*Daily Beta*
Wayne-Gossard	0.976	0.692	0.986	0.654	0.459
Mich. Seamless Tube	0.973	0.883	0.917	0.784	0.433
Publicker Inds.	1.521	1.491	1.513	1.277	1.006
Great West. United	2.496	2.311	2.122	1.911	1.442
Family Finance	1.268	1.324	1.212	0.821	0.795
Bobbie Brooks	1.874	1.889	1.818	1.592	1.405
Monogram Inds.	2.950	2.887	2.844	2.403	2.144
Faberge	1.882	1.511	1.511	1.416	1.449
Dillingham Corp.	1.004	1.164	0.990	0.750	0.725
Vornado	2.329	1.628	2.170	1.823	1.765
Big Three Inds.	1.339	0.970	1.283	0.969	0.712
Cabot Corp.	0.752	0.898	0.844	0.805	0.756
General Development	1.423	1.628	1.657	1.382	1.358
Addresso-Multigraph	2.094	2.341	1.566	1.414	1.733
Great West. Financial	2.246	1.820	2.043	2.158	1.917
Colgate-Palmolive	1.131	1.002	1.011	0.958	0.850
Aluminium Co. of America	1.115	1.221	1.118	1.150	1.118
Shell Oil	0.930	1.093	0.827	0.860	0.742
Kresge, S.S.	1.190	1.326	1.299	1.308	1.237
Eastman Kodak	0.932	0.859	0.958	1.166	1.251

*Returns are measured as the logarithm of investment relatives. Market returns are those of the S&P 500. All betas are statistically significant at the 5 per cent level.

general market. Securities' daily prices do not move in unison; some stocks may lag behind the general market movement, others may lead it. Previous research has suggested that the fundamental cause of these intertemporal cross-correlations is friction in the trading process, which delays the response of securities' prices to new information.[3]

Table II presents some evidence of this phenomenon. Here the first column gives the contemporaneous correlation coefficient (ρ_{im}) between daily security returns and the daily returns of the Standard & Poor's (S&P) 500, used as a proxy for the market, m. This coefficient measures the simultaneous association between a security's daily price movements and those of the S&P 500. The second column gives the intertemporal correlation coefficient with a lag of one day (ρ_{im}^{-1}). This measures the strength of the association between today's securities' returns and those of the S&P 500 one day earlier; in this case, changes in individual prices lag movements in the S&P 500. The third column gives the intertemporal correlation coefficient with a lead of one day (ρ_{im}^{+1}).[4] The last column gives the ratio of the sum of the two intertemporal correlation coefficients to the contemporaneous correlation coefficient; it is a measure of the strength of intertemporal cross-correlation relative to contemporaneous correlation. I call this ratio the "q-ratio" of a given security.

The q-ratios of Wayne-Gossard (WG) and Eastman Kodak (EK), for example, are derived as follows:

$$q_{WG} = \frac{\rho_{WG,m}^{-1} + \rho_{WG,m}^{+1}}{\rho_{WG,m}} = \frac{0.106 + 0.060}{0.143} = 1.161 .$$

$$q_{EK} = \frac{\rho_{EK,m}^{-1} + \rho_{EK,m}^{+1}}{\rho_{EK,m}} = \frac{0.094 + 0.189}{0.626} = 0.452 .$$

The S&P 500 itself has a measurable q-ratio, which is equal to:

$$q_m = \frac{\rho_{mm}^{-1} + \rho_{mm}^{+1}}{\rho_{mm}} = \frac{2\rho_m^{-1}}{1} = 2(0.285) = 0.570 ,$$

where $\rho_{mm}^{-1} = \rho_{mm}^{+1}$ = the autocorrelation coefficient with a lag of one day in the S&P 500, whose value was found to equal 0.285 over the period January 1970 to December 1973. Since the contemporaneous price movements of the S&P 500 are perfectly positively correlated with themselves, $\rho_{mm} = 1$, hence the q-ratio of the S&P 500 equals twice its autocorrelation coefficient.

Importance of the q-Ratio

The presence of daily intertemporal cross-

Table II Daily Correlation Coefficients and q-Ratios (January 1970–December 1973)

	ρ_{im}	ρ_{im}^{-1}	ρ_{im}^{+1}	q_{im}
Wayne-Gossard	0.143	0.106	0.060*	1.161
Mich. Seamless Tube	0.165	0.177	0.044*	1.321
Publicker Inds.	0.281	0.150	0.108	0.981
Great West. United	0.233	0.131	0.096	0.974
Family Finance	0.248	0.200	0.046*	0.992
Bobbie Brooks	0.339	0.108	0.100	0.614
Monogram Inds.	0.459	0.119	0.166	0.621
Faberge	0.337	0.087	0.134	0.664
Dillingham Corp.	0.192	0.119	0.047*	0.865
Vornado	0.391	0.225	0.124	0.893
Big Three Inds.	0.355	0.219	0.090	0.870
Cabot Corp.	0.324	0.175	0.037	0.654
General Development	0.328	0.101	0.132	0.710
Addresso-Multigraph	0.404	0.074	0.168	0.599
Great West. Financial	0.546	0.176	0.180	0.652
Colgate-Palmolive	0.334	0.164	0.047*	0.632
Aluminium Co. of America	0.437	0.148	0.112	0.569
Shell Oil	0.394	0.198	0.110	0.782
Kresge, S.S.	0.505	0.137	0.168	0.608
Eastman Kodak	0.626	0.094	0.189	0.452
S&P 500	1.000	0.285	0.285	0.570

*Statistically insignificant at the 5 per cent level.

correlations, whose relative strengths are measured by the q-ratio, causes beta estimates to be related to the length of the return interval. The beta coefficient of security i (β_i), estimated over return intervals of T-day length, is defined as the ratio of the T-day covariance ($\sigma_{im}(T)$ between the returns of i and those of the market index m to the T-day variance ($\sigma_m^2(T)$) of the returns of the market index, or:

$$\beta_i(T) = \frac{\sigma_{im}(T)}{\sigma_m^2(T)} \quad . \tag{1}$$

The appendix demonstrates that $\sigma_{im}(T)$ can be expressed as a function of the daily covariance $\sigma_{im}(1)$, the length T of the return interval in days, and the security's q-ratio (q_{im}) according to:

$$\sigma_{im}(T) = \sigma_{im}(1)[T + (T-1)q_{im}] \ . \tag{2}$$

An analogous expression can be derived to relate the T-day variance $\sigma_m^2(T)$ to the daily variance $\sigma_m^2(1)$ as well as T and q_m, such that:

$$\sigma_m^2(T) = \sigma_m^2(1)\,[T + (T-1)q_m] \ , \tag{3}$$

where

$$q_m = 2\rho_m^{-1} \ ,$$

that is, twice the first order daily autocorrelation coefficient.

Substituting Equations (2) and (3) in Equation (1) gives:

$$\beta_i(T) = \frac{\sigma_{im}(1)[T+(T-1)q_{im}]}{\sigma_m^2(1)[T+(T-1)q_m]} = \beta_i(1)\frac{T+(T-1)q_{im}}{T+(T-1)q_m} \ . \tag{4}$$

Equation (4) clearly shows how intertemporal cross-correlations affect beta as T varies. First, consider the simple case in which the security's q-ratio and the market's q-ratio both equal zero. In this case, the security's beta over the measurement interval will equal its daily beta, regardless of the length of the return interval. In this highly unlikely situation, beta will be invariant to the length of the return interval. Beta will also be invariant when the security's q-ratio is non-zero and equals that of the market index. This situation may possibly arise. In general, however, a security's q-ratio will differ from the market's, causing beta to change as the return interval varies.

Predicting Direction and Strength of a Beta Shift

For which securities will beta rise, remain constant or fall as the return interval changes? How strong will the variation in beta be? To answer these questions, we measure the change in the numerator of Equation (4) with respect to the denominator; that is, we measure the response of $\beta_i(T)$ to a small change in T. We have:

$$\frac{d\beta_i(T)}{dT} = \frac{\beta_i(1)[q_{im} - q_m]}{[T + (T-1)q_m]^2} \quad . \tag{5}$$

For beta to decrease as T is shortened, the change must be positive. This will occur whenever the security's q-ratio is *larger* than the market's. For beta to increase as the return interval shortens, the change must be negative. This will occur whenever the security's q-ratio is *smaller* than the market's. Furthermore, beta will decrease faster the larger is the security's q-ratio relative to the market's, and it will increase faster the smaller the security's q-ratio is relative to the market's.

Consider, for example, Wayne-Gossard and Eastman Kodak. The former has a q-ratio of 1.161, far in excess of that of the market index (0.570). As the return interval is shortened, its beta should decrease at a faster rate than the beta of other securities with q-ratios exceeding 0.570 but smaller than 1.161. This is indeed what we observe (see Table I). Eastman Kodak has a q-ratio of 0.452, below that of the market index. Its beta should increase as the return interval is shortened. This is again what we observe from Table I. The q-ratio of Aluminium Co. of America equals 0.569 (Table I), which is approximately the value of the q-ratio of the market index. We should expect its beta to remain constant as the return interval varies, and this is indeed the case.

Is there a faster way to tell if beta will shift upward or downward? In other words, is it possible to predict the direction of the shift in beta without knowing its q-ratio? In fact, a security's q-ratio is inversely related to that security's market value of shares outstanding (MVSO).[5] This means that securities with large MVSO (relative to the market average) will have small q-ratios in comparison to the market. Those with small MVSO (relative to the market average) will have high q-ratios in comparison to the market. Therefore, we may use MVSO (a proxy for a security's relative market thinness) to determine the direction of the shift in beta. Eastman Kodak, with a large MVSO (relatively small q-ratio), will have an increasing beta, and Wayne-Gossard, with a small MVSO (relatively high q-ratio), will have a decreasing beta, when the return interval is shortened. Table II, which lists companies in increasing order of MVSO, illustrates that q-ratios do indeed have a tendency to fall as MVSO rises.

Implications for Analysis

Because beta estimates depend upon the length of the interval used to measure securities' returns, any computation that incorporates beta may also be affected by the length of the return interval. This includes risk-adjusted measures of investment performance. Consider the performance index suggested by Treynor.[6] It is the ratio of a portfolio's excess return to its beta. Because beta depends upon the return interval used, so will the performance index. Furthermore, since betas do not usually shift in the same direction, the ranking of portfolios according to the Treynor index may differ depending upon the return interval used.

Another implication concerns the estimation of the Security Market Line of Sharpe and Lintner.[7] From an estimated SML, one gets the value of the market price of risk reduction (MPR)—that is, the estimated market reward associated with an increase in each unit of risk. If one obtains a monthly estimate of MPR, then it will usually be incorrect to convert this number to an annual basis by simply multiplying by 12.[8] Likewise, a monthly estimate of MPR cannot usually be converted to a weekly MPR by simply dividing by four.

Finally, because of intertemporal cross-correlations between securities' price movements, beta estimates will generally depend upon the length of the return interval, implying that betas measured over return intervals of arbitrary length are biased. In particular, securities with large MVSO (relative to the market average) will have estimated betas that are biased upward, whereas securities with small MVSO (relative to the market average) will have estimated betas that are biased downward.[9] Hence securities with relatively small market values may appear to be less risky than they truly are, whereas securities with relatively large market values may appear to be more risky than they truly are.[10] ■

Footnotes

1. See, for example, Bruce D. Fielitz and Myron T. Green, "Shortcomings in Portfolio Evaluation via MPT," *Journal of Portfolio Management*, Summer 1980, pp. 13–19; Gabriel A. Hawawini, "Intertemporal Cross Dependence in Securities Daily Returns and the Short Run Intervaling Effect on Systematic Risk," *Journal of Financial and Quantitative Analysis*, March 1980, pp. 139–149; Cheng F. Lee and K. Morimune, "Time Aggregation, Coefficient of Determination, and Systematic Risk of the Market Model," *The Financial Review*, Spring 1978, pp. 36–47; David Levhari and Haim Levy, "The Capital Asset Pricing Model and the Invest-

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2. The CRSP tapes compiled at the University of Chicago's Center for Research in Security Prices are typical.
3. Kalman Cohen, Gabriel Hawawini, Steven Maier, Robert Schwartz and David Whitcomb, "Implications of Microstructure Theory for Empirical Research on Stock Price Behavior," *Journal of Finance*, May 1980, pp. 249-257.
4. Intertemporal correlation coefficients of order higher than the first—not reported in Table II—are usually statistically insignificant. See Gabriel Hawawini, "The Intertemporal Cross Price Behavior of Common Stocks: Evidence and Implications," *Journal of Financial Research*, Fall 1980, pp. 153-167.
5. This is demonstrated in Hawawini, "The Intertemporal Cross Price Behavior of Common Stocks."
6. Jack L. Treynor, "How to Rate Management of Investment Funds," *Harvard Business Review*, January/February 1965, pp. 63-75.
7. William F. Sharpe, "Capital Asset Prices: A Theory of Market Equilibrium Under Conditions of Risk," *Journal of Finance*, September 1964, pp. 425–442 and John Lintner, "The Valuation of Risk Assets and the Selection of Risky Investments in Stock Portfolios and Capital Budgets," *The Review of Economics and Statistics*, February 1965, pp. 13–37.
8. See Gabriel A. Hawawini and Ashok Vora, "Temporal Aggregation and the Estimation of the Market Price of Risk," *Economics Letters*, 1980, pp. 165–170.
9. See Hawawini, "The Intertemporal Cross Price Behavior of Common Stocks."
10. Recent research has suggested some techniques that can possibly be used to correct for this return interval bias. See Kalman Cohen, Gabriel Hawawini, Steven Maier, Robert Schwartz and David Whitcomb, "Estimating and Adjusting for the Intervaling-Effect Bias in Beta," *Management Science*, forthcoming; Elroy Dimson, "Risk Measurement when Shares are Subject to Infrequent Trading," *The Journal of Financial Economics*, 1979, pp. 197–226; and Myron Scholes and Joseph Williams, "Estimating Beta from Non-Synchronous Data," *Journal of Financial Economics*, 1977, pp. 309–327.

Appendix

Derivation of Equation (2)

To derive Equation (2), write the T-day covariance of the *i*th security's returns with the returns on a market index (m) as a function of daily returns:

$$\sigma_{im,t}(T) = \sum_{k=o}^{T-1} \sum_{u=u}^{T-1} \mathrm{Cov}(r_{iT-k,i}, r_{iT-u,m}) \qquad \text{(A1)}$$

Equation (A1) is a TxT cross-covariance matrix. Assuming stationarity, its diagonal elements are all equal to the daily contemporaneous covariance denoted $\sigma_{im}(1)$, and there are T of those. The off-diagonal elements are intertemporal cross-covariances. Under stationarity, all intertemporal cross-covariances for which the difference $k-u = +1$ are equal to $\rho_{im}^{+1}\sigma_i(1)\sigma_m(1)$, and there are $(T-1)$ of those. We have assumed that for $k-u > +1$, intertemporal cross-covariances are zero. Likewise, all intertemporal cross-covariances for which the difference $k-u=-1$ are equal to $\rho_{im}^{-1}\sigma_i(1)\sigma_m(1)$, and there are $(T-1)$ of those. Again, for $k-u < -1$, intertemporal cross-covariances are zero. From this discussion and Equation (A1) it follows that:

$$\sigma_{im}(T) = T\sigma_{im}(1) + (T-1)\rho_{im}^{+1}\sigma_i(1)\sigma_m(1) + (T-1)\rho_{im}^{-1}\sigma_i(1)\sigma_m(1) \ ,$$

from which Equation (2) in the text follows directly, because

$$\sigma_i(1)\sigma_m(1) = \sigma_{im}(1)/\rho_{im} \ .$$

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MCC DATA REQUEST
DATED JULY 12, 2013
DOCKET NO. D2012.9.100

MCC-201 **Regarding:** **Beta**
Witness: **J. Stephen Gaske**

Please provide an electronic copy of Figure 2: Rolling Average Value Line Beta for Proxy Group Companies, as shown on page 27 of your rebuttal testimony. Please also include all data, formulas, worksheets and supporting documents used in calculating each of the points depicted in the chart.

Response:

Please see the Excel file on the enclosed CD labeled "MCC-201 Proxy Group Beta.xlxs".

MONTANA-DAKOTA UTILITIES CO.
MONTANA CONSUMER COUNSEL
MCC DATA REQUEST
DATED JULY 12, 2013
DOCKET NO. D2012.9.100

MCC-202 **Regarding: Flotation Cost Adjustment**
Witness: J. Stephen Gaske

Please provide a copy of the full chapter of the publication from which the quotation at page 30, footnote 35 of your rebuttal testimony was obtained: Myron J. Gordon, *The Cost of Capital to a Public Utility*, Michigan State University, 1974.

Response:

Please see Attachment A.

THE COST OF CAPITAL TO A PUBLIC UTILITY

Myron J. Gordon

1974
MSU Public Utilities Studies

Division of Research
Graduate School of Business Administration
Michigan State University
East Lansing, Michigan

The Institute of Public Utilities, Graduate School of Business Administration, Michigan State University, publishes books, monographs, and occasional papers as part of its program of promoting academic interest in the study of public utilities and regulation. The views and opinions expressed in these works are solely those of the authors, and acceptance for publication does not constitute endorsement by the Institute, its member companies, or Michigan State University.

Library of Congress Catalog Card Number: 73-620076
ISBN: 0-87744-122-7

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To Joe and David

7.6 Conclusion

This chapter has provided the basis for questioning the validity of the PCCM leverage theorem and the PCCM and traditional dividend rate and stock financing theorems. In the process, alternative hypotheses about the relation between share yield (and price) and these financing policy variables have been indicated. The following chapter undertakes the development of a stock valuation and cost of capital model which incorporates these alternative hypotheses.

8

The Cost of Capital in Imperfect Capital Markets

Recognition of the fact that capital markets are not perfectly competitive requires significant departure from the line of attack followed in the previous chapters in arriving at a utility's cost of capital. In perfectly competitive capital markets the cost of capital is independent of the level and financing of the firm's investment and is equal to the leverage-free yield on the firm's stock. Cost of capital determination under the theory involved little more than measuring share yield.

The traditional theory departs from the PCCM theory only in the leverage assumption; the cost of capital is a function of the utility's leverage rate. However, in implementing the theory the practice is to fix the leverage rate on the basis of exogenous considerations. Consequently, given the leverage rate, the cost of capital here also is independent of the level and financing of investment, and the only problem is measuring share yield.

We have seen that the yield at which a utility's stock sells is an increasing function of its retention and stock financing rates as well as of its leverage rate. Knowing a share's yield under the firm's existing investment and financing policies does not tell us what the yield would be under alternative policies. What is more important, knowing share yield, even under the desired investment

and financing policies, does not tell us the cost of capital. All we know is that, with share yield a function of investment rate, the marginal yield investors require at any investment rate is above the average yield at that investment rate. Our task then is to construct a share value model which correctly represents the influence on share price of a utility's investment and financing policies. Manipulation of the model will provide the yield at which the share sells for any combination of investment and financing policies.

Given this model of share valuation, how should it be used to determine a utility's cost of capital? There are two answers to this question depending on the policy adopted by the regulatory agency. Under one policy the agency sets the allowed rate of return under the constraint that the utility may take it as given in making its investment and financing decisions. The utility makes these decisions with the objective of maximizing share price. Under the other policy the agency recognizes the bilateral monopoly relation with the utility and adjusts the allowed rate of return in response to changes in the utility's investment and financing decisions to make the price of the stock independent of these decisions. It will be seen that the bilateral monopoly policy results in a lower cost of capital and also requires less information to implement than the alternative policy, which we call the constrained policy.

The pages that follow will develop the imperfect capital markets stock value model and arrive at the cost of capital from each source—retention, debt, and sale of stock—under the constrained and the bilateral monopoly regulatory policies. The following chapters will test the model, estimate its parameters, and provide illustrative estimates of the cost of capital from each source under both policies.

8.1 Retention Share Value and the Cost of Capital

Our fundamental and perfectly general stock value model is that the price of a share is the present value of the expected future dividends. With the current dividend $(1 - b)Y$ and with the assumption that the dividend is expected to grow forever at the rate $g = br$, our stock value model becomes

$$P = \frac{(1 - b)Y}{k - br}. \tag{8.1.1}$$

The evidence in chapter 6 established that rather than k being independent of g, it increases with g. If we assume that k is a linear function of g,

$$k = \alpha_0 + \alpha_1 br. \tag{8.1.2}$$

α_0 is the yield at which the share would sell in the absence of growth and the growth coefficient $\alpha_1 > 0$. Substituting Eq. (8.1.2) for k, Eq. (8.1.1) becomes

$$P = \frac{(1 - b)Y}{\alpha_0 + (\alpha_1 - 1)br}. \tag{8.1.3}$$

Taking the derivative $\partial P/\partial b$ we find

$$\frac{\partial P}{\partial b} = \frac{Y}{[\alpha_0 + (\alpha_1 - 1)br]^2}[(1 - \alpha_1)r - \alpha_0]. \tag{8.1.4}$$

This equation states that price increases with retention if $(1 - \alpha_1)r > \alpha_0$. This implies that a firm should retain all of its earnings if $r > \alpha_0/(1 - \alpha_1)$ and retain nothing if $r < \alpha_0/(1 - \alpha_1)$.

The alternative hypothesis to a linear relation between k and g is that k increases with g by increasing amounts. This assumption is plausibly and conveniently represented by the expression

$$k = br + 1/[\alpha_0 e^{\alpha_1 br}]. \tag{8.1.5}$$

$1/\alpha_0$ is the yield at which the share would sell when $br = 0$ and $\alpha_1 > 0$. As br increases k may fall at first, but beyond some value of br, k will increase with g in decreasing amounts.[1] An interesting property of this function is that $k - br$ falls as br increases, asymptotically approaching zero as $br \to \infty$. Figure 1 illustrates how k and $k - g$ vary with g. $k - g$ is the vertical distance between the curves marked k and g.

Eq. (8.1.5) is a plausible expression for k since $k - g$, the dividend yield, approaches but never reaches zero as g rises. By contrast, for small values of α_1, the Eq. (8.1.2) value of k may result in

[1] k is minimized at $br = (1/\alpha_1)\ln(\alpha_1/\alpha_0)$.

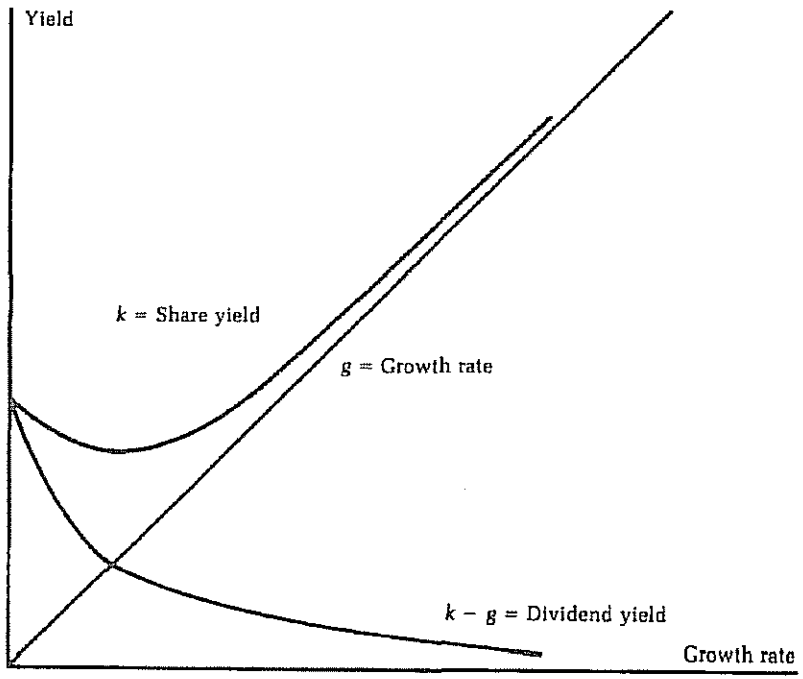

FIGURE 1. *Variation in Share Yield and in Dividend Yield with the Growth Rate*

$k - g = 0$ for some value of g. Eq (8.1.5) is convenient for a number of reasons, as will be seen in the course of what follows. Substituting the equation for k in Eq. (8.1.1) results in

$$P = \alpha_0(1 - b)\,Ye^{\alpha_1 br}. \tag{8.1.6}$$

This equation states that P is the multiple α_0 of the dividend when $br = 0$, and the multiple increases with $g = br$. The coefficient α_1 may be looked on as the price the market is willing to pay for growth.

Eq. (8.1.6) is linear in the logarithims; hence, linear regression analysis of sample data may be used to estimate the values of the coefficients α_0 and α_1. This equation and the related Eq. (8.1.5) for k have two clear advantages over the Eq. (8.1.2) expression for k. With k a linear function of g we are led to an estimating equation of the form

$$\frac{D}{P} = \alpha_0 + \alpha_1 g + \ldots. \tag{8.1.7}$$

In Eq. (8.1.7) the change in price with growth is an increasing percentage of the price, but if P/D is made the dependent variable, the change in price is a decreasing percentage of the price as g increases. By contrast, in Eq. (8.1.6) the change in price with growth is a constant percentage of the price, and it is the same with P, P/D, or D/P the dependent variable. This is the first advantage. The second is that in Eq. (8.1.6) the least squares estimates of the coefficients minimize the squared ratios of the actual to the estimated price or dividend yield. Eq. (8.1.7) minimizes the squared differences between the actual and estimated dividend yield, which will not necessarily make the error expressed as a percentage of the price independent of the price.[2]

Taking the derivate $\partial P/\partial b$ in Eq. (8.1.6) and setting it equal to zero, we find that share price is maximized when

$$b = 1 - 1/\alpha_1 r. \tag{8.1.8}$$

Hence, when k is given by Eq. (8.1.5) there is an optimal retention rate which increases both with the rate of return on investment, r, and with α_1, the index of the market's willingness to pay for growth in the dividend.

Let us assume that the regulatory agency allows the utility the freedom to make whatever retention decision it wishes given the rate of return decision of the regulatory agency. Since the objective of the utility is to maximize the share price, Eq. (8.1.8) provides the utility's retention rate as a function of the agency's return decision. Solving the equation for r, we find that

$$r = 1/\alpha_1(1 - b) \tag{8.1.9}$$

[2]It is also true that the least squares estimates of the coefficients of Eq. (8.1.7) with D/P the dependent variable differ from the coefficients with P/D the dependent variable so that given values of D and g produce different values of P depending on which set of coefficients are used. By contrast, when Eq. (8.1.6) is used it does not matter whether D/P or P/D is made the dependent variable.

is the return on common equity the utility must be allowed to earn to persuade it to retain the fraction b of its income. It is clear that r increases with b.

A retention rate of b and a return on common equity equal to r produce a rate of growth of br in the common equity and assets. Hence, the investment rate is br. To illustrate the model, assume that $\alpha_1 = 15$ and that the desired investment rate is .05. Multiply both sides of Eq. (8.1.9) by b, set $br = .05$, set $\alpha_1 = 15$, and solve for b; the result is $b = .43$. With $br = .05$, the solution for $r = .117$. If the demand for service requires a rate of growth in assets of $br = .04$, the combination of $b = .375$ and $r = .106$ maximizes share price at the required investment rate.

Recall that the above solution for the rate of return allows the utility to take the rate of return on common as given in making its retention decision. Under this assumption, with the debt-equity ratio fixed and with retention the only source of equity funds, the utility's cost of capital may be determined as follows. With no debt in the capital structure, $r = x$, and Eq. (8.1.9) provides a firm's cost of capital. With a leverage rate of B/E and the coupon rate of interest equal to c, we have $r = x + (x - c)\,B/E$. Substituting this expression for r, the cost of capital is

$$z = x = \frac{1}{\alpha_1(1 - b)(1 + B/E)} + \frac{cB/E}{1 + B/E}. \qquad (8.1.10)$$

Since the investment rate is an increasing function of the retention rate when there is no stock financing and the leverage rate is fixed, it is clear from Eq. (8.1.10) that the cost of capital is an increasing function of the investment rate. It also can be shown that the cost of retention capital for any value of b falls as the given leverage rate is raised as long as $1/\alpha_1(1 - b)$ is greater than the coupon rate of interest.

8.2 *The Bilateral Monopoly Cost of Retention Capital*

The previous section established a utility's cost of retention capital if the utility may take the allowed rate of return as given in deciding its retention rate. However, there is a bilateral monopoly relation

between the utility and the regulatory agency. The latter may adjust the allowed rate of return in response to the retention rate selected by the utility. The consequences of each of these regulatory policies for the cost of capital, its relation to share yield, and the market-book value relation now are examined.

Figure 2 presents share price as a function of the retention rate for alternative values of the allowed rate of return. The retention rate b is on the horizontal axis, and share price P is on the vertical axis. The horizontal line E is the book value per share. If the allowed rate of return is set at r_2, share price is maximized at a retention rate of b_2, and the investment rate is $b_2 r_2$. Similarly, b_1 and b_3 are the optimal retention rates when r_2 and r_3 are the allowed rates of return. It is clear that the price of the stock and the welfare of the stockholders are increasing functions of the investment rate needed to satisfy the demand for service. Stimulating the demand

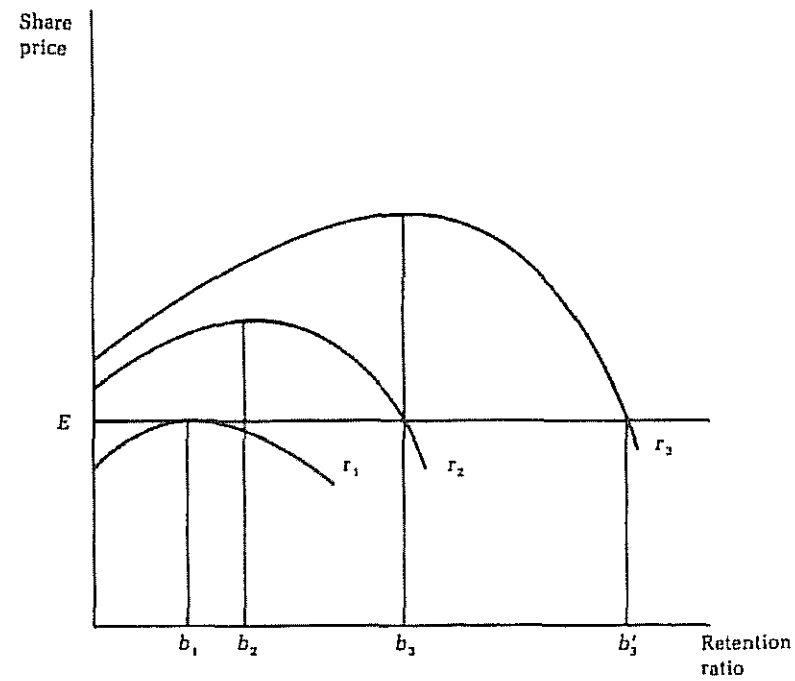

FIGURE 2. ***Comparison of Constrained and Bilateral Monopoly Cost of Retention Capital***

for service and/or increasing the capital requirements to meet the demand for service benefit the stockholders.

Assume now that the regulatory agency decides to adjust the allowed rate of return on assets (and common, given the debt-equity ratio) to keep the price per share of the common stock equal to its book value. Furthermore, let the investment rate that satisfies the public's demand for service be $g = b_3' r_3$. If the regulatory agency knows this and if it knows the coefficients of Eq. (8.1.6), it will set the return on assets so that the return on common $r = r_3$. If the utility sets $b = b_3'$, the result will be an investment rate that satisfies the demand for service and $P = E$.

The utility well may recognize that with $r = r_3$ setting $b = b_3 < b_3'$ will result in a higher price per share. However, the management knows that if it sets $b = b_3$ the regulatory agency will reduce r from r_3 to r_2, and the combination $b_3 r_2$ results in $P = E$ once again. The management might just as well set $b = b_3'$ and provide the investment rate that satisfies the demand for service. The stockholder, of course, is indifferent to the retention rate since the price of the stock is independent of the retention rate. Furthermore, the dividend plus the growth in price during the coming period divided by the current price will be exactly equal to the yield he requires on the stock regardless of the retention rate.

What if the utility management refuses to accept the regulatory agency's policy of maintaining $P = E$? With $r = r_3$, the management sets $b = b_3$, and when the agency reduces r to r_2, the utility counters with $b = b_2 < b_3$. The downward spiral continues until $b = 0$ or a positive b maximizes share price at $P = E$. The stockholders have gained nothing, but the investment rate is inadequate to meet the demand for service. If the public's dissatisfaction with inadequate service can be turned against the agency, it may be forced to give up the policy that maintains $P = E$. Although this possibility cannot be denied, it would seem that the relative power position of a regulatory agency committed to the public interest would enable it to enforce its policy. However, it must be acknowledged that where a bilateral monopoly exists considerations outside the scope of economic theory, narrowly defined, enter into the solution that will obtain.

Under a bilateral monopoly policy the agency need not burden itself with the difficult tasks of estimating the demand for service and the coefficients of Eq. (8.1.6). The agency need only concern

itself with keeping the market value of the stock equal to its book value. With stockholders indifferent to the investment rate, the management may just as well keep consumers happy by selecting the investment rate that satisfies the demand for service. Since the utility management is likely to be better informed on the demand for service than the agency, it would seem advisable for the agency to follow this simple policy. The only argument against leaving the responsibility to the management is that the management may inject its own preference into the decision. The bias then would be in the direction of a higher investment rate since growth provides various benefits to a management. However, carrying this policy too far may destroy the utility's monopoly position. Raising the investment rate increases the price of the product due to the higher capital base and the rise in the required rate of return. At some point the demand for service may not be equal to the supply at a price which covers the cost of the required capital, and the utility is in the position of a competitive firm that has overinvested.

To see the implications of this policy for the cost of capital and allowed rate of return, set $P = E$, substitute $Y = Er$ in Eq. (8.1.6), and rewrite it as follows:

$$r = 1/\alpha_0(1 - b)e^{\alpha_1 br}. \tag{8.2.1}$$

Comparison with Eq. (8.1.9) reveals that when $b = 0$ the two values of r are the same if $\alpha_0 = \alpha_1$. If $\alpha_0 > \alpha_1$ the bilateral monopoly solution results in a lower r at $b = 0$. As b rises both the Eq. (8.2.1) and the Eq. (8.1.9) values of r rise, but the bilateral monopoly value rises less rapidly due to the presence of the exponential term.

All the relevant considerations support the bilateral monopoly approach to regulatory policy. The cost of capital is lower, and the utility has no inducement to overinvest or underinvest. Furthermore, the regulatory agency has two alternative and complementary courses of action in implementing the policy. The agency can estimate the investment rate that satisfies the demand for service, set br equal to that value in Eq. (8.2.1), and solve for r. Alternatively, the agency can rely on the possibly superior information available to the utility on the investment rate that satisfies the demand for service and simply adjust the allowed rate of return to keep the market-book value ratio equal to one.

8.3 Leverage Share Value and the Cost of Capital

The presence of leverage in the capital structure is recognized by adding a leverage rate variable to Eq. (8.1.6) as follows:

$$P = \alpha_0(1 - b)\,Ye^{\alpha_1 br}e^{\alpha_2 B/E}. \tag{8.3.1}$$

The leverage rate, B/E, is based on the book values of the debt and equity since the empirical work of the previous chapter provided no basis for believing that market values should be used. Given the dividend and its growth rate, an increase in leverage and the risk associated with it should cause the share price to fall. Hence, we would expect that $\alpha_2 < 0$. With this form of the function the share price goes down by α_2 percent for every 1 percent increase in the leverage. In consequence, share yield increases by increasing amounts as leverage rises. Ezra Solomon [43] and most other advocates of the traditional theory postulate a relationship of this kind between leverage and share yield. By contrast, the linear regression of share yield on leverage and other variables implies that the rise in share yield with leverage is independent of the leverage rate.

Leverage has a favorable impact on share price through its impact on earnings and return on investment since $Y = Er$ and $r = x + (x - c)B/E$. Assuming for the moment that $b = 0$, Eq. (8.3.1) can be written

$$P/E = \alpha_0[x + (x - c)B/E]\,e^{\alpha_2 B/E}. \tag{8.3.2}$$

Dividing both sides of Eq. (8.3.1) by book value so that the market-book value ratio is the dependent variable facilitates the analysis that follows. Taking the derivative of P/E with respect to leverage results in

$$\frac{\partial(P/E)}{\partial(B/E)} = \left[\alpha_0 e^{\alpha_2 B/E}\right] \times \left[(x - c)(1 + \alpha_2 B/E) - \frac{\partial c}{\partial(B/E)}\frac{B}{E} + \alpha_2 x\right]. \tag{8.3.3}$$

We are interested in whether P/E is maximized at some value of B/E and, if so, what that value is. The second order conditions necessary to answer the first question are extremely difficult to evaluate. However, we can proceed as follows. Eq. (8.3.3) is positive at $(B/E) = 0$ if $-\alpha_2 < (x - c)/x$. Since $\alpha_2 < 0$ and is very small in absolute amount relative to $(x - c)/c$, this condition is easily satisfied. At $(B/E) = 0$ it is possible that $\partial c/\partial(B/E)$ is zero or negative, but when B/E passes some value, $\partial c/\partial(B/E)$ becomes positive, and it rises by increasing amounts. Hence, P/E does reach a maximum at a finite value of B/E. Over this, the relevant range of B/E, we find its optimal value by setting Eq. (8.3.3) equal to zero and solving for B/E. The result is

$$\frac{B}{E} = \frac{x - c + \alpha_2 x}{-\alpha_2(x - c) + \partial c/\partial(B/E)}. \tag{8.3.4}$$

The optimal leverage rate increases with $x - c$, and it decreases as $-\alpha_2$ and $\partial c/\partial(B/E)$ are raised.

The cost of debt capital is obtained by setting Eq. (8.3.3) equal to zero and solving for x instead of B/E. The result is

$$z = x = \frac{c[1 + \alpha_2(B/E) + \partial c/\partial(B/E)]}{1 + \alpha_2 + \alpha_2(B/E)}. \tag{8.3.5}$$

With $\alpha_2 < 0$ it can be shown that the cost of debt capital increases with the interest rate and with the leverage rate. The cost of capital also increases with $-\alpha_2$, the leverage risk coefficient, and with the response of the interest rate to changes in leverage.

It should be noted that the interest rate in the above model is the coupon or imbedded rate of interest. If the current interest rate is above the coupon rate, the latter will move over time to equality with the current rate. Hence, for a given leverage rate, the appropriate value of the coupon rate is a weighted average of the coupon and current interest rates. The weights depend on the refunding and the new debt financing required to maintain the existing leverage rate during the year.

As the firm's leverage rate is raised, the interest rate at which it borrows is raised more or less regardless of the relation between the coupon and the current interest rates. It could be argued that

only the new debt during the year will be at the higher interest rate. In that event the definition of the coupon rate in the previous paragraph accounts for this change in the interest rate since $\partial c/\partial(B/E) > 0$ in Eq. (8.3.5). However, bond contracts subordinate or otherwise restrict additional debt, and the change in the coupon rate with the leverage rate should reflect some refinancing of the existing debt or a risk premium on the new debt commensurate with its subordination.

With a positive retention rate the optimum leverage rate is finite if $-\alpha_2 > (x - c)\alpha_1 b$. In that event the counterpart to Eq. (8.3.4) with $\partial c/\partial(B/E) = 0$ is

$$\frac{B}{E} = \frac{x - c + \alpha_2 x + (x - c)x\alpha_1 b}{-\alpha_2(x - c) - (x - c)^2\alpha_1 b}. \tag{8.3.6}$$

Since $\alpha_1 > 0$, the optimum leverage rate is an increasing function of the retention rate. The value of x that satisfies $\partial(P/E)/\partial(B/E) = 0$ when $b > 0$ is a quadratic function of x. The economically valid solution and hence the cost of debt capital is a decreasing function of the retention rate. The cost of capital remains an increasing function of the interest rate, the leverage rate, and $-\alpha_2$.

Under the analysis just presented the regulatory agency was presumed to follow a constrained policy. That is, it allows the rate of return that maximizes share price at the debt-equity ratio it desires the utility to adopt. Under the policy the cost of debt capital and share price are increasing functions of the desired debt-equity ratio. By contrast, under a policy that recognizes the bilateral monopoly relation between the agency and the utility, the allowed rate of return is adjusted to maintain the market-book ratio for the common stock equal to one. How does the allowed rate of return vary with the debt-equity ratio when the condition that the price of the stock remains unchanged is satisfied?

It is not easy to answer this question by analytic methods since Eq. (8.3.2) cannot be arranged to make x an explicit function of B/E with P/E set equal to one. It will be shown in the next chapter, however, that the values of the parameters cause x to fall initially and then rise as B/E rises. The leverage rate at which x is minimized, of course, minimizes the cost of debt capital.

The yield at which the common stock sells rises continuously

with the debt-equity ratio, but this only compensates shareholders for their risk. When the share price is independent of the debt-equity ratio, stockholders are indifferent to that ratio. Consumers are concerned with the return on assets, and their interest is served by the debt-equity ratio that minimizes the return on assets. It should be noted, however, that the utility's risk increases with the leverage rate, and management may consider its welfare to vary inversely with the risk to the firm. Management therefore may prefer a debt-equity ratio lower than the one that minimizes the cost of capital, particularly since the stockholders are indifferent to the ratio's value.

8.4 External Finance, Share Value, and the Cost of Capital

The third source of funds to finance investment, the sale of additional shares, is the most difficult to incorporate into our model. The consequences of stock financing may be classified as short term and long term. This section will develop an imperfect capital markets model that captures the long-term consequences of stock financing. The next section modifies the model to incorporate the short-term consequences.

Our perfectly general model with stock financing present is

$$\frac{P}{E} = \frac{(1 - b)r}{k - br - vs}, \tag{2.8.8}$$

where s is the stock financing rate and v is the equity accretion rate on stock financing. As shown in chapter 2, the equity accretion rate is

$$v = (r - k)/(r - rb - s). \tag{2.8.12}$$

Under the PCCM theory of stock valuation, k is independent of s. Hence, if $r = k$, $v = 0$ regardless of s, and share price is independent of the stock financing rate. The cost of external equity capital is ρ, the leverage-free value of k.

However, if $r > k$ the consequence of stock financing for the value of P poses some problems. First, with v given, the denominator

of Eq. (2.8.8) approaches zero as s rises and $P \to \infty$. But we cannot take v as given since Eq. (2.8.12) shows that with $r > k$ v rises with s. The unreasonable conclusion is that stock financing easily can generate an infinite share price for a profitable firm for which $r > k$.

Recognition that capital markets are not perfectly competitive is a means of avoiding the above ridiculous conclusion.[3] Our hypotheses are that the yield investors require on a share is an increasing function of both the dividend growth rate and the stock financing rate. The expression for share yield that incorporates these hypotheses and the leverage theorem is

$$k = br + vs + 1/[\alpha_0 e^{\alpha_1 br} e^{\alpha_2 B/E} e^{\alpha_3 vs} e^{\alpha_4 s^2}]. \tag{8.4.1}$$

With $\alpha_3 > 0$ and $\alpha_4 < 0$, k increases with vs, the dividend growth due to stock financing, and also with s, the stock financing rate. Substituting this expression for k in Eq. (2.8.8) and representing all the terms of no immediate interest by A_0, we have

$$P/E = A_0(1-b)re^{\alpha_3 vs}e^{\alpha_4 s^2}. \tag{8.4.2}$$

This expression simply states that, given the current dividend $(1-b)rE$, share price increases with the expected rate of dividend growth due to stock financing, vs, and it decreases with the stock financing rate s.

Eq. (8.4.2) poses two questions. Why are there two stock financing terms, and why does the second term have the squared value of s? On the first question, it is clear that yield should rise with vs, the stock financing growth rate. However, with vs the only stock financing variable, share yield is independent of the stock financing rate when $v = 0$. This is clearly not reasonable. Rather, as reflected in Eq. (8.4.1), share yield should increase with the stock financing rate and with the growth due to stock financing. Substituting this expression for k in Eq. (2.8.8) results in a stock price equation with two stock financing terms.

[3] Another way out of the dilemma is to assume r is a decreasing function of the investment rate. This is not a correct assumption for regulated companies. For unregulated companies the assumption is unlikely to deal with the finite price problem.

To answer why the squared value of s is used in Eq. (8.4.2), let us replace it with s and take the derivative of P/E with respect to s:

$$\frac{\partial(P/E)}{\partial s} = (P/E)\left[\alpha_3 v + \alpha_3 s\frac{\partial v}{\partial s} + \alpha_4\right]. \tag{8.4.3}$$

With $\partial(P/E)/\partial s > 0$ at $s = 0$ it will remain so as s increases. The reason is that with $v = 1 - E/P$, when $\partial(P/E)/\partial s > 0$, $\partial v/\partial s > 0$. Hence, the second term in Eq. (8.4.3) is positive and increases with s if $\partial(P/E)/\partial s > 0$ at $s = 0$, and P/E rises with s without end.

By contrast, with Eq. (8.4.2) as it stands, we obtain, using $v = 1 - E/P$,

$$\frac{\partial(P/E)}{\partial s} = \frac{(P/E)[\alpha_3 v + 2\alpha_4 s]}{\left[1 - \dfrac{\alpha_3 s}{P/E}\right]}. \tag{8.4.4}$$

As s is increased, P/E increases in decreasing amounts and is maximized at the optimal stock financing rate

$$s = -\alpha_3 v/2\alpha_4. \tag{8.4.5}$$

The optimal stock financing rate varies directly with α_3 and v and inversely with $-\alpha_4$. However, since $v = 1 - E/P$, it is an implicit function of all the other variables and parameters which determine P, and Eq. (8.4.5) cannot be used directly to find the optimal stock financing rate.

To establish the optimal stock financing rate we must determine v for any value of s and the other variables. To do so we may proceed as follows. Substitute the Eq. (8.4.1) value of k in Eq. (2.8.12). The result is

$$v = \frac{r - br - vs - 1/[A_0 e^{\alpha_3 vs} e^{\alpha_4 s^2}]}{r - br - s}$$

$$= 1 - \frac{1}{r(1-b)A_0 e^{\alpha_3 vs} e^{\alpha_4 s^2}}. \tag{8.4.6}$$

When $s = 0$, v is an explicit function of the variables on the right-hand side of Eq. (8.4.6), and we may obtain $v \gtreqless 0$, depending on the values of the variables and the coefficients.

When $v < 0$ at $s = 0$ we also have $P/E < 1$ since $v = 1 - E/P$. As s rises above zero, both v and P/E fall, and the rise in s makes a bad situation worse. Hence, when $v < 0$ at $s = 0$ the optimal course of action is $s = 0$.[4] When $v = 0$ and $P/E = 1$, a rise in s reduces v and P/E. Hence, in this case also the optimal course of action is $s = 0$.

When $v > 0$ and $P/E > 1$ at $s = 0$, a rise in s raises v and P/E. There is some positive value of v greater than the value of v at $s = 0$ which satisfies Eq. (8.4.6) at each value of s. Trial and error calculations may be used to find this value of v and the associated value of P/E at each value of s. For each value of s and the values of the variables other than v on the right-hand side of Eq. (8.4.6) there is also a negative value of v that satisfies the equation. However, this solution to the equation can be ruled out on economic grounds. As s rises the one admissible solution for v and P/E for each s either will reach a maximum at a finite s or rise continuously with s. In the former case we have an optimal s for the given values of the variables and coefficients. In either case we can find the value of r necessary to make a stock financing rate at least equal to some value optimal policy.

8.5 Further Consequences of Stock Financing

The previous section established how we measure v given s, r, the other variables, and the parameters of Eq. (8.4.2). It also established that the optimal stock financing rate is

$$s = \alpha_3 v / -2\alpha_4. \tag{8.4.5}$$

The interpretation of this expression is somewhat tricky because v depends on s, among other variables. Instead of attempting this, let us consider the equity accretion rate necessary to make some value of s optimal policy. Solving Eq. (8.4.5) for v we obtain

[4] There are some combinations of the coefficients and variables in Eq. (8.4.6) which result in two positive solutions for v with $v < 0$ at $s = 0$. However, these combinations of the variables and coefficients can be ruled out on economic grounds.

$$v = -2\alpha_4 s / \alpha_3. \tag{8.5.1}$$

If a stock financing rate of $s = s^*$ is desired, the value of $v = v^*$ needed to make it optimal policy is provided by Eq. (8.5.1). The problem that remains is to determine the return on common equity that will realize $v = v^*$ when $s = s^*$, and we proceed as follows. Assign a value to r in Eq. (8.4.6) and find the value of v that satisfies the equation with $s = s^*$ and the given values of all the other variables. If there is no solution for v or if $v < v^*$, raise r. There is some value of $r = r^*$ at which $v = v^*$ with $s = s^*$. It is the return on common equity needed to make s^* optimal policy. Given the firm's debt-equity ratio, the return on assets $x = x^*$ needed to realize $r = r^*$ can be determined.

The bilateral monopoly solution to the cost of new equity capital is considerably simpler than the above competitive solution. Under a policy of maintaining $P/E = 1$, $v = 0$ regardless of s. Market price is independent of the stock financing rate, the existing shareholders are indifferent to the stock financing rate, and the regulatory agency need only maintain $P/E = 1$ to generate whatever stock financing rate the demand for service requires.

To determine the consequences of such a policy for the rate of return on common equity, we set $P/E = 1$ and $v = 0$ in Eq. (8.4.2) and solve for r. The result is

$$r = 1/A_0(1 - b)e^{\alpha_4 s^2}. \tag{8.5.2}$$

Taking the derivative with respect to s we find that

$$\frac{\partial r}{\partial s} = \frac{-2\alpha_4 s}{A_0(1 - b)e^{\alpha_4 s^2}}. \tag{8.5.3}$$

Since $\alpha_4 < 0$, Eq. (8.5.3) is positive, and r rises with the stock financing rate. That is, maintaining $P/E = 1$ involves raising r with s. However, this is of no benefit to the shareholder since the yield he requires has gone up correspondingly. Also, it can be shown that the value of r that satisfies Eq. (8.5.2) is less than the value of r that satisfies Eq. (8.5.1) for any value of s.

The previous analysis recognized that if a corporation is expected to stock finance periodically at some average annual rate then the yield investors require on the stock is raised directly by the stock

financing rate and indirectly by the growth due to the stock financing. In addition to these long-run influences on share price, stock financing will have two short-run consequences. First, Eq. (2.8.12) should be modified to recognize that the firm does not receive the entire amount paid by the public for the issue. Second, the expression for v should recognize that the issue causes a slight temporary depression in the price. This is accomplished as follows. With Q_n the amount invested in the corporation by new shareholders during n, the equity in the corporation they receive is $(1 - v - \psi)Q_n$, where ψ is the fraction of Q_n that accrues to the investment banker. The dividend they can expect in $n + 1$ is Eq. (2.8.10) changed accordingly:

$$D^*_{n+1} = (1 - b)r(1 - v - \psi)Q_n. \tag{8.5.4}$$

Once in the corporation the new shares are identical with the old. Their dividends also are expected to grow at the rate $br + vs$. However, to bring the shares in required a higher yield or some depression of the price at the time of the issue. This is recognized by changing Eq. (2.8.11) to

$$Q_n = \sum_{t=n+1}^{\infty} \frac{(1 - b)r(1 - v - \psi)Q_n(1 + br + vs)^{t-n-1}}{(1 + \eta k)^{t-n}}$$

$$= [(1 - b)r(1 - v - \psi)Q_n]/[\eta k - br - vs], \tag{8.5.5}$$

where $\eta > 1$ is the ratio of the yield on a share at the time of a new issue to the yield at other times. Dividing both sides by Q_n and solving for v results in

$$v = \frac{r[1 - \psi(1 - b)] - \eta k}{r - br - s}. \tag{8.5.6}$$

Plausible values of ψ and η are $\psi = .05$ and $\eta = 1.05$. The consequence is that the numerator of this expression is less than the numerator of Eq. (2.8.12) for given values of r, b, s, and k, while the denominator is unchanged. In other words, introducing ψ and η reduces v.

Substituting the Eq. (8.4.1) value for k in this expression for v results in

$$v = \frac{r[1 - b(\eta - \psi) - \psi]}{r[1 - b] + s[\eta - 1]} - \frac{\eta}{[r(1 - b) + s(\eta - 1)]\, A_0 e^{\alpha_3 vs} e^{\alpha_4 s^2}}. \tag{8.5.7}$$

While this expression may appear considerably more involved than Eq. (8.4.6), the differences are not important. The first term is reduced from one to a number that is slightly less than one, and it falls somewhat as s rises. The numerator of the second term is raised slightly, and the denominator is reduced slightly and rises less rapidly with s. The characteristics of the solution to Eq. (8.5.7) are the same as the characteristics of the solution to Eq. (8.4.6). The only difference is that for any s the solution value of v is reduced somewhat. Eqs. (8.4.5) and (8.5.1) remain unchanged. However, the value of r needed to generate any value of v is raised so that the cost of stock financed capital under our constrained model of regulatory policy is an increasing function of ψ and η.

However, with $\eta > 1.0$ and/or $\psi > 0$, the bilateral monopoly solution to regulatory policy is somewhat more complicated than it is with $\eta = 1.0$ and $\psi = 0$. The proceeds to the corporation from a stock issue are less than the market price of the stock, and with price equal to book the issue reduces both the book and market value of the stock.[5] The equity of the existing shareholders is diluted by the stock issue. In other words, with $P = E$ and $r = k$, a stock issue or stock financing at some rate make the return the shareholders can expect to earn a decreasing function of the stock financing. Therefore, it is against the interest of the existing shareholders for the firm to engage in stock financing when the proceeds per share on the issue are less than the book value per share.

To make shareholders indifferent to the rate at which a utility stock finances, the regulatory agency must set $r > k$ so that $P > E$ by a margin that makes the proceeds per share on the new issue equal to the book value of the outstanding stock. With $r > k$, $P > E$, and the stock sells to earn k prior to a new issue. With

[5]The price (and book value) per share after the issue is a weighted average of the preissue price and the proceeds to the corporation per share issued. The weights are the relative amounts of old and new shares.

the new shares sold at a price of E per share, the buyers receive a return above k for coming into the company, but they do not receive r since part of the proceeds on the issue goes to the investment banker. Both the new and the old shares earn Er per share, and they sell at a yield of k and a price of P, the old price per share.

All this may seem very complicated, but it need not trouble the regulatory agency. The agency need only estimate the proportion that the proceeds per share on an issue bear to the price of the stock and adjust the allowed rate of return so that the price per share is the indicated ratio of the book value per share. If the proceeds on an issue are 91 percent of market price, the agency should maintain market price at about 110 percent of book value. The welfare of the stockholders is independent of the firm's stock financing rate, and the utility may be expected to set s to satisfy the demand for service.

9

Estimation of the Parameters and Implementation of the Model

This chapter begins with tests of the stock value model developed in the previous chapters and experiments with alternative forms of the model, using the sample of utility firms for the years 1958–1968 described in chapter 5. The objective is to find the form that is best suited for estimating the cost of capital from each source and in the aggregate. The parameter estimates and the sample data then are used to arrive at illustrative estimates of the cost of capital from each source: retention, debt, and the sale of stock. These results are obtained under the constrained and the bilateral monopoly policies.

The next chapter takes up the problem of the overall cost of capital. That is, it develops an algorithm for determining the allowed rate of return necessary to obtain a given investment rate with the choice between retention and stock financing being made so as to minimize the allowed rate of return. The debt-equity ratio is taken as fixed for reasons that will become clear in the course of the discussion. Chapter 10 continues with a comparative analysis of the cost of capital under the different theories and the allowed rate of return for certain of the sample years, and it concludes

Response No. MCC-20
Attachment
Page 13 of 1

MONTANA-DAKOTA UTILITIES CO.
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MCC DATA REQUEST
DATED JULY 12, 2013
DOCKET NO. D2012.9.100

MCC-203 Regarding: Relative Risk
Witness: J. Stephen Gaske

Please provide a copy of the document referenced in footnote 40 of your rebuttal testimony: Michael Annin, *Equity and the Small-Stock Effect*, Public Utilities Fortnightly, October 15, 1995.

Response:

Please see Attachment A.

Financial News

By Michael Annin

Equity and the Small-Stock Effect

The capital asset pricing model shows risk inherent in return on equity. But something goes wrong when it's used for small-sized companies.

Does the size of a company affect the rate of return it should earn? If smaller companies should earn a higher return than larger firms, then small utilities, because of their size, should be allowed to adjust the rates they charge to customers.

By far the most notable and well-documented apparent anomaly in the stock market is the effect of company size on equity returns. The first study focusing on the impact that company size exerts on security returns was performed by Rolf W. Banz. Banz sorted New York Stock Exchange (NYSE) stocks into quintiles based on their market capitalization (price per share times number of shares outstanding), and calculated total returns for a value-weighted portfolio of the stocks in each quintile. His results indicate that returns for companies from the smallest quintile surpassed all other quintiles, as well as the Standard & Poor's 500 and other large stock indices. A number of other researchers have replicated Banz's work in other countries; nevertheless, a consensus has not yet been formed on why small stocks behave as they do.

One explanation for the higher returns is the lack of information on small companies. Investors must search more diligently for data. For small utilities, investors face additional obstacles, such as a smaller customer base, limited financial resources, and a lack of diversification across customers, energy sources, and geography. These obstacles imply a higher investor return.

The Flaw in CAPM

One of the more common cost of equity models used in practice today is the capital asset pricing model (CAPM). The CAPM describes the expected return on any company's stock as proportional to the amount of systematic risk an investor assumes. The traditional CAPM formula can be stated as:

$$R_s = [\beta_s \times RP] + R_f$$

where:

R_s = expected return or cost of equity on the stock of company "s"

β = the *beta* of the stock of company "s"

RP = the expected equity risk premium

R_f = expected return on a riskless asset.

Table 1: The Size Premium in CAPM
(By Decile Portfolio in NYSE, 1926-94)

Decile	*Beta*	Arithmetic Mean Return	Actual Return in Excess of Riskless Rate**	CAPM Return in Excess of Riskless Rate**	Size Premium (Return in Excess CAPM)
1	0.90	11.01%	5.88%	6.33%	–0.44%
2	1.04	13.09	7.97	7.34	0.63
3	1.09	13.83	8.71	7.70	1.01
4	1.13	14.44	9.32	7.98	1.33
5	1.17	15.50	10.38	8.22	2.16
6	1.19	15.45	10.33	8.38	1.95
7	1.24	15.92	10.79	8.75	2.05
8	1.29	16.84	11.72	9.05	2.67
9	1.36	17.83	12.71	9.57	3.14
10	1.47	21.98	16.86	10.33	6.53

*Betas are estimated from monthly returns in excess of the 20-year government bond income return, January 1926-December 1994.
**Historical riskless rate measured by the 69-year arithmetic mean income return component of 20-year government bonds.
Source: SBBI 1995 Yearbook

Table 2: CAPM vs. CAPM w/ Size Premium

(By Percentile for Electric, Gas, and Sanitary Services Utilities)

	CAPM	CAPM with Size Premium
90th Percentile	16.42%	18.92%
75th Percentile	12.56%	14.72%
Median	10.89%	12.58%
25th Percentile	9.86%	11.39%
10th Percentile	8.63%	10.65%

(Weighted by Market Capitalization)

	CAPM	CAPM with Size Premium
Industry Composite	11.76%	12.33%
Large Company Composite	12.05%	12.07%
Small Company Composite	13.93%	17.95%

Source: Cost of Capital Quarterly '95 Yearbook by Ibbotson Associates
Note: Public utilities include electric, gas, and sanitary services companies.

Table 1 shows *beta* and risk premiums over the past 69 years for each decile of the NYSE. It shows that a hypothetical risk premium calculated under the CAPM fails to match the actual risk premium, shown by actual market returns. The shortfall in the CAPM return rises as company size decreases, suggesting a need to revise the CAPM.

The risk premium component in the actual returns (realized equity risk premium) is the return that compensates investors for taking on risk equal to the risk of the market as a whole (estimated by the 69-year arithmetic mean return on large company stocks, 12.2 percent, less the historical riskless rate). The risk premium in the CAPM returns is *beta* multiplied by the realized equity risk premium.

The smaller deciles show returns not fully explainable by the CAPM. The difference in risk premiums (realized versus CAPM) grows larger as one moves from the largest companies in decile 1 to the smallest in decile 10. The difference is especially pronounced for deciles 9 and 10, which contain the smallest companies.

Based on this analysis, we modify the CAPM formula to include a small-stock premium. The modified CAPM formula can be stated as follows:

$$R_s = [\beta_s \times RP] + R_f + SP$$

where:

SP = small-stock premium.

Because the small-stock premium can be identified by company size, the appropriate premium to add for any particular company will depend on its equity capitalization. For instance, a utility with a market capitalization of $1 billion would require a small capitalization adjustment of approximately 1.3 percent over the traditional CAPM; at $400 million, approximately 2.1 percent, and at only $100 million, approximately 4 percent.

Again, these additions to the traditional CAPM represent an adjustment over and above any increase already provided to these smaller companies by having higher *betas*.

Implications for Smaller Utilities

These findings carry important ramifications for relatively small public utilities. Boosting the traditional CAPM return by a full 400 basis points for small utilities translates into a substantial premium over larger utilities.

Table 2 shows the results of an analysis of 202 utility companies that calculated cost of equity figures. Composites (arithmetic means) weighted by equity capitalization were also calculated for the largest and smallest 20 companies. The results show the impact size has on cost of equity.

For the traditional CAPM, the large-company composite shows a cost of equity of 12.05 percent; the small company composite, 13.93 percent. However, once the respective small capitalization premium is added in, the spread increases dramatically, to 12.07 and 17.95 percent, respectively. Clearly, the smaller the utility (in terms of equity capitalization), the larger the impact that size exerts on the expected return of that security. ▼

Michael Annin, CFA, is a senior consultant with Ibbotson Associates, specializing in business valuation and cost of capital analysis. He oversees the Cost of Capital Quarterly, *a reference work on using cost of capital for company valuations.*

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MCC DATA REQUEST
DATED JULY 12, 2013
DOCKET NO. D2012.9.100

MCC-204 Regarding: Flotation Cost Adjustment
Witness: J. Stephen Gaske

At page 2 of JSG-03 you claim that a flotation cost adjustment is required whether or not the regulated company has incurred flotation costs in the past or whether it will incur such costs in the near future. Please provide a calculation showing the amount of test year revenue that MDU will receive from Montana ratepayers if the flotation cost adjustment that you propose is approved in this case.

Response:

The requested calculation is provided below:

			Notes
1	Montana Gas Rate Base	$43,759,831	
2	x Common Equity Ratio	53.387%	
3	Equity Portion of Rate Base	$23,362,061	
4	Proposed Primary Market ROE	10.5000%	
5	÷ Flotation Cost Adjustment	1.04	
6	Secondary Market ROE	10.0962%	
7	Flotation Cost Adjustment	0.4038%	line 4 - line 6
8	÷ Income Tax Factor	60.6125%	39.3875% tax rate
9	Pre-Tax Flotation Cost Adjustment	0.6663%	
10	$ Amount of Flotation Cost Adjustment	**$155,656**	line 3 x line 9

Thus, this cost element constitutes one-fourth of one percent (0.25%) of the revenue requirement of $62.8 million.

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MCC DATA REQUEST
DATED JULY 12, 2013
DOCKET NO. D2012.9.100

MCC-205 Regarding: Comparable Investment Standard
Witness: J. Stephen Gaske

At page 4 of Exhibit JSG-03 you state that setting the authorized return on common equity for Montana-Dakota's natural gas distribution operations at 9.0 percent would violate the comparable investment standard set forth in the Supreme Court's *Hope* decision. Please provide the specific language from the *Hope* decision that you say establishes such a standard.

Response:

Below is the specific language from the *Hope* decision:

> "The rate-making process under the Act, *i.e.*, the fixing of 'just and reasonable' rates, involves a balancing of the investor and the consumer interests. Thus, we stated in the *Natural Gas Pipeline Co.* case that 'regulation does not insure that the business shall produce net revenues.' 315 U.S. at page 590, 62 S.Ct. at page 745, 86 L.Ed. 1037. But such considerations aside, the investor interest has a legitimate concern with the financial integrity of the company whose rates are being regulated. From the investor or company point of view it is important that there be enough revenue not only for operating expenses but also for the capital costs of the business. These include service on the debt and dividends on the stock. *Cf. Chicago & Grand Trunk R. Co. v. Wellman*, 143 U.S. 339, 345, 346, 12 S.Ct. 400, 402, 36 L.Ed. 176. ***By that standard, the return to the equity owner should be commensurate with returns on investments in other enterprises having corresponding risks.*** That return, moreover, should be sufficient to assure confidence in the financial integrity of the enterprise, so as to maintain its credit and to attract capital. *See State of Missouri ex rel. Southwestern Bell Tel. Co. v. Public Service Commission*, 262 U.S. 276, 291, 43 S.Ct. 544, 547, 67 L.Ed. 981, 31 A.L.R. 807 (Mr. Justice Brandeis concurring). The conditions under which more or less might be allowed are not important here. Nor is it important to this case to determine the various permissible ways in which any rate base on which the return is computed might be arrived at. For we are of the view that the end result in this case cannot be condemned under the Act as unjust and unreasonable from the investor or company viewpoint."[1] [***Emphasis added***]

[1] *Federal Power Commission v. Hope Natural Gas Company*, 320 U.S. 591, 603 (1944).

MONTANA-DAKOTA UTILITIES CO.
MONTANA CONSUMER COUNSEL
MCC DATA REQUEST
DATED JULY 12, 2013
DOCKET NO. D2012.9.100

MCC-206 Regarding: Analysts' Forecasts
Witness: J. Stephen Gaske

At page 7 of Exhibit JSG-03 you state that "Although analysts' long term growth forecasts are typically expressed as five-year forecasts, these forecasts generally represent growth rate expectations for a longer period of time than the five years expressed in the forecasts." Please provide all documentation supporting your contention that the published five year forecasts generally represent growth rate expectations for a longer period of time than the five years expressed in the forecasts.

Response:

When analysts' are surveyed they typically are asked for their five-year growth rate estimates. However, the fact that long-term growth rate estimates are often expressed as five-year estimates does not mean that analysts' expect growth rates to change after five years. For example, see pages 7-8 of Exhibit No. ___(JSG-8) for a description of some of the economics literature that establishes analysts' growth rate estimates as the most accurate estimates for determining stock prices. Because stock prices are best explained by the analysts' five-year estimates, and not by multi-stage growth rate methods that change the growth rate after five years, the analysts' estimates are treated by investors as being valid for a period longer than five years.

MONTANA-DAKOTA UTILITIES CO.
MONTANA CONSUMER COUNSEL
MCC DATA REQUEST
DATED JULY 12, 2013
DOCKET NO. D2012.9.100

MCC-207 Regarding: Analysts' Forecasts
Witness: J. Stephen Gaske

At page 8 of Exhibit JSG-03 you state that there is no data supporting the observation that stock prices are typically somewhat lower than analysts' price forecasts. Please provide the current stock prices for each of your comparable gas utility companies and the current consensus price forecast from Yahoo Finance for each of these same companies.

Response:

Please see Attachment A for this data. As shown on Attachment A, the one year target price is lower than the current price for four proxy group companies, while the one year target price is higher than the current price for three proxy group companies.

Montana-Dakota Utilities Co.

Company	Ticker	Current Price	1-Year Target Price	Target Price Less Current Price
AGL Resources Inc.	GAS	$ 45.36	N/A	N/A
Atmos Energy Corp.	ATO	$ 43.43	$ 42.00	$ (1.43)
Laclede Group, Inc.	LG	$ 46.90	$ 48.50	$ 1.60
New Jersey Resources Corp.	NJR	$ 43.42	$ 46.00	$ 2.58
Northwest Natural Gas Co.	NWN	$ 44.45	$ 43.00	$ (1.45)
Piedmont Natural Gas Co., Inc.	PNY	$ 34.89	$ 32.60	$ (2.29)
South Jersey Industries, Inc.	SJI	$ 60.54	$ 63.75	$ 3.21
Southwest Gas Corp.	SWX	$ 49.83	$ 47.00	$ (2.83)
MDU Resources Group Inc.	MDU	$ 28.01	$ 27.78	$ (0.23)

Source: Yahoo! Finance. Closing prices as of July 15, 2013.

MONTANA-DAKOTA UTILITIES CO.
MONTANA CONSUMER COUNSEL
MCC DATA REQUEST
DATED JULY 12, 2013
DOCKET NO. D2012.9.100

MCC-208 Regarding: Advanced Research Center
Witness: J. Stephen Gaske

Please identify, by location and affiliation, the specific "Advanced Research Center" that you refer to in Footnote 9 and identify, by name, the author or authors of the "updated study" that you cite.

Response:

The "Advanced Research Center" is a division of State Street Global Advisors. State Street Global Advisors is the asset management business of State Street Corporation, one of the world's leading providers of financial services to institutional investors. State Street Global Advisors has offices in 16 countries around the world including Australia, Belgium, Canada, China, France, Germany, India, Ireland, Italy, Japan, Netherlands, Singapore, Switzerland, United Arab Emirates, United Kingdom, and United States. The authors of the "updated study" are Anita Xu and Ami Teruya who worked with Duke University professor Dr. James H. Vander Weide, one of the authors of the original study conducted in 1988.[2]

[2] Vander Weide, J. H., and W. T. Carleton. "Investor Growth Expectations: Analysts vs. History." *The Journal of Portfolio Management*, Spring 1988, pp. 78-82.

MONTANA-DAKOTA UTILITIES CO.
MONTANA CONSUMER COUNSEL
MCC DATA REQUEST
DATED JULY 12, 2013
DOCKET NO. D2012.9.100

MCC-209 **Regarding: Exhibit JSG-10 (MDU Common Stock)**
Witness: J. Stephen Gaske

At page 31 of Exhibit JSG-03 and in Exhibit JSG-10 you refer to the three most recent common equity issuances by MDU Resources. Please specify the book value of MDU Resources Group total common stock at 1/1/1998 and at 12/31/2012.

Response:

According to the SEC Form 10-K for the respective years, the book value per share of MDU Resources Group at January 1, 1998 was $10.39, and at December 31, 2012 was $13.95.

MONTANA-DAKOTA UTILITIES CO.
MONTANA CONSUMER COUNSEL
MCC DATA REQUEST
DATED JULY 12, 2013
DOCKET NO. D2012.9.100

MCC-210 Regarding: Local Economy
Witness: J. Stephen Gaske

At page 31 of Exhibit JSG-03 you state that MDU serves an undiversified local economy. Please list the ten largest business sectors in this local economy served by the MDU gas distribution utility in Montana and the percentage of the local economy accounted for by each of these sectors.

Response:

Data by industry for Montana-Dakota's specific service territory in eastern Montana are not available. According to the U.S. Department of Commerce's Bureau of Economic Analysis, the following ten industries had the largest contribution to Montana's Gross Domestic Product in 2012:

Table 1: Montana's Gross Domestic Product (2012)

Industry	$MM	% of Total
Government	6,342	15.69%
Real estate and rental and leasing	4,805	11.89%
Health care and social assistance	3,736	9.24%
Manufacturing	2,860	7.08%
Retail trade	2,698	6.67%
Wholesale trade	2,100	5.20%
Professional, scientific, and technical services	2,100	5.20%
Finance and insurance	2,088	5.17%
Construction	2,033	5.03%
Mining	1,949	4.82%
All industry total	40,422	

Please see the Response No PSC-079(a), PSC-134(a), and MCC-211(b).

MONTANA-DAKOTA UTILITIES CO.
MONTANA CONSUMER COUNSEL
MCC DATA REQUEST
DATED JULY 12, 2013
DOCKET NO. D2012.9.100

MCC-211 Regarding: Purpose of proceeding
Witness: J. Stephen Gaske

At page 32 of Exhibit JSG-3 you state that the purpose of this proceeding is to establish the authorized ROE for MDU's Montana natural gas distribution business as if it were a stand-alone entity that is independently going to capital markets to raise equity capital.

a. Should the cost of the debt component of the Company's capital likewise reflect debt costs as if the Montana natural gas distribution business was a stand-alone entity that is independently going to capital markets to raise debt capital? Please fully explain your answer.

b. Please provide documentation showing that the Commission's ROE determination in any previous MDU gas distribution rate case was made assuming that the Company's Montana natural gas distribution business was a stand-alone entity that independently went to capital markets to raise equity capital

c. Please provide documentation showing that any of MDU's equity capital was actually obtained as if the Company's Montana natural gas distribution business was a stand-alone entity going to capital markets independently to raise equity capital.

d. Please specify and fully explain the regulatory basis under which it would be just and reasonable to the Company's ratepayers for the Commission to establish an authorized ROE in this proceeding under the assumed pretense that MDU's Montana natural gas distribution business was a stand-alone entity that independently went to capital markets to raise equity capital.

Response:

a. Yes. Although Montana-Dakota's Montana natural gas distribution business does not issue its own debt, the stand-alone principle applies to the cost of debt for the same reasons that it applies to the cost of equity. Specifically, the cost of debt should reflect the stand-alone cost of debt as if MDU's Montana natural gas distribution business were issuing debt independently in capital markets. The primary reason, as discussed in more detail below, is that the stand-alone principle protects ratepayers from paying a higher cost of debt if the parent company is engaged in higher risk non-regulated activities. There is precedent for this approach among regulators. For example, in Quebec, Gazifere (a small gas distribution company) is owned by Enbridge. The Régie has imputed a cost of debt for Gazifere based on the cost of debt for Enbridge Gas

Response No. MCC-211 (cont.)

Distribution's operations in Ontario plus a risk premium to account for the small size of Gazifere.

b. The most recent Montana-Dakota gas distribution rate case decision was Order No. 5856b in Docket No. D95.7.90 issued by the Commission in 1996. As shown in the following excerpt, the Commission's precedent regarding risks is explicitly based on an assessment of the business risks of the Montana jurisdictional operations on a stand-alone basis:

> 25. Gaske argues that MDU is riskier than the companies in his sample group. His arguments are contained in his direct testimony, pp. 29-39. In his conclusions on risk he provides the following reasons why MDU is riskier than his sample group:
>
> > There are considerable risks associated with investments in gas distribution companies and these risks have increased in recent years. In my opinion, Montana-Dakota's overall risks are greater than those of any of the companies in the comparison group. **The considerably higher business risk is due primarily to the small size *of the Montana jurisdictional natural gas operations* relative to the size of the comparison companies** and the perceived risks of bypass due to the unusually large amount of direct competition with another gas utility. Montana Dakota's Montana operations also face regulatory risks that are above average relative to those of the comparison group. In addition, Montana-Dakota's financial risks are clearly greater than the average financial risks faced by the comparison companies.
>
> 26. **The Commission is persuaded that Gaske's conclusions regarding increased risk, in comparison to the sample group, appear to be valid for financial risk and, to a limited extent, for business risk.** MDU, in comparison to the sample group average, is financially more risky because it has a higher magnitude of debt in its capital structure and its bond rating is lower. **In the area of business risk MDU's smaller revenue base, compared to that of the sample group, increases its risk,**

Response No. MCC-211 (cont.)

> **because MDU has a smaller proportion of return available to absorb fixed costs during periods of economic downturns.** ...
> 27. ... The Commission's acceptance of the argument that MDU has greater financial and business risk than the sample group of companies, entitles MDU to an increase in its ROE above that found reasonable for the sample group. In addition to being riskier than the sample group the Commission finds that MDU should be rewarded for its good management practices such as diversifying its gas portfolio, minimizing rate filings, and minimizing its costs of doing business. The Commission determines that MDU should be authorized a 12.0 percent ROE. [Emphasis added].

c. The equity capital used to finance Montana-Dakota's regulated gas distribution operations in Montana is obtained from the parent company, MDU Resources Group. Each business in which MDU Resources invests on must provide an opportunity to earn a return that is appropriate for the risks of that particular business in order to justify the investment of investors' equity funds in that business. The assessment of risks and the adequacy of expected returns for each asset or business on a stand-alone basis is a fundamental principle of financial management.

d. Many regulators have adopted the stand-alone principle because it insulates ratepayers from high costs of capital associated with riskier operations and, as discussed in response to part (c) above, it ensures that the company can compete for funds that are to be invested in the stand-alone regulated business. These principles have been explained by many regulators.

For example, in Quebec, the Régie de l'Energie has explained that the return on equity for Hydro Quebec Distribution and Hydro Quebec TransEnergie should be set on a "stand-alone" basis as if the entities were independently seeking to attract capital in the equity markets.

Similarly, in a 2007 decision involving Potomac Electric Power, the Maryland Public Service Commission stated: "We reject People's Counsel's proposed capital structure [reflecting a double leverage adjustment] because it suffers from numerous flaws. First, it assumes that

Response No. MCC-211 (cont.)

the rate of return depends on the source of capital rather than the risks faced by the capital."[3]

The Alberta Energy and Utilities Board described the stand-alone principle as follows: "This first application of the stand-alone principle is designed to remove the effects of diversification by utilities into non-regulated activities. Using the stand-alone principle in this case, a utility is regulated as if the provision of the regulated service were the only activity in which the company was engaged. This application of the principle ensures that the revenue requirement of regulated utility operations is not influenced up or down by the operations of a parent or "sister" company. Thus, the cost (or revenue requirement) or providing utility service reflects only the expenses, capital costs, risks and required returns associated with the provision of regulated service." [4]

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[3] Maryland Public Service Commission, Order No. 81517, Case No. 9092, In the Matter of Potomac Electric Power Company for Authority to Revise its Rate and Charges for Electric Service and for Certain Rate Design Changes, July 19, 2007. Clarification added.

[4] Alberta Energy and Utilities Board, Decision 2001-92, December 12, 2001, at 25.

MONTANA-DAKOTA UTILITIES CO.
MONTANA CONSUMER COUNSEL
MCC DATA REQUEST
DATED JULY 12, 2013
DOCKET NO. D2012.9.100

MCC-212 Regarding: Small Utilities
Witness: J. Stephen Gaske

At page 33 of Exhibit JSG-03 you quote Ibbotson regarding investors' return requirements. Provide any documentation which you assert supports Ibbotson's credibility and expertise on the subject of investors' return requirements for equity capital

Response:

Ibbotson Associates is a recognized source of financial data that investors rely upon for historical data regarding returns achieved for various asset classes including returns for stocks with various market capitalizations. The financial data provided by Ibbotson Associates represents actual historical returns in the market, not the opinion of Ibbotson Associates about those returns.

Ibbotson Associates was founded in 1977 by Dr. Roger G. Ibbotson and Rex Sinquefield, both graduates from the University of Chicago. Dr. Ibbotson is a finance professor at the Yale School of Management, while Mr. Sinquefield is a retired financial executive who created the first S&P indexes in 1973, and co-founded Dimensional Fund Advisors in 1981, which has more than $200 billion in assets under management, and co-authored *Stocks, Bonds, Bills and Inflation* with Dr. Ibbotson. Ibbotson Associates was acquired by Morningstar, another established provider of financial data, in 2006. Ibbotson Associates is a registered investment advisor whose purpose is to deliver innovative asset allocation solutions, help investors reach their financial goals, and provide asset allocation thought leadership.

The reference to a small size premium appeared in an article by Michael Annin, an employee of Ibbotson Associates, which was published in Public Utilities Fortnightly. Mr. Annin received an M.B.A from the University of Texas and a B.B.A from the University of Wisconsin. He is currently a Senior Managing Director for Mesirow Financial in Chicago. Mr. Annin was previously a Managing Director at Ibbotson Associates, and held positions at IPC Group and Abbott Laboratories. He holds the Chartered Financial Analyst designation.