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The Pricing of Interest-Rate Risk: Evidence from the Stock Market

RICHARD J. SWEENEY and ARTHUR D. WARGA*

ABSTRACT

This paper addresses the issue of whether firms are required to pay an *ex ante* premium to investors for bearing the risk of interest-rate changes. A two-factor APT model with the market and changes in the yield on long-term government bonds as factors is employed. The paper shows that, empirically, most of the interest-sensitive stocks are in the utility industries, and that there is reasonable evidence that the interest factor is priced in the sense of the APT. Several sources for the interest sensitivity are considered, and regulatory lags are focused on as a likely candidate.

SOME FIRMS EXHIBIT *ex post* sensitivity in their stock returns to unanticipated changes in interest rates. Are these firms required to pay investors *ex ante* for bearing this risk of interest-rate changes? The framework for examining pricing of unanticipated interest-rate changes is a two-factor APT model with the market and changes in the yield on long-term government bonds as factors. The issue is whether the interest-rate change premium is significant. To study this, the paper uses full information maximum likelihood (FIML) estimation on groups of twenty-five individual firms, with both cross-equation constraints and within-equation nonlinear constraints on the parameters as mandated by the APT model.

Whether the risk of unanticipated changes in interest rates is priced in the market is of particular interest with regard to regulated firms. Regulated firms (especially electric utilities) turn out to comprise the big majority of firms that are sensitive to unanticipated interest-rate changes. After some preliminary investigation, the analysis focuses on electric utilities.

Much empirical work on the APT relies on factor analysis, e.g., Roll and Ross [30]. Bower, Bower, and Logue [5] use factor analysis in comparing the APT and the CAPM to explain utilities' returns. The present paper, however, uses observable variables at the macro level. Chan, Chen, and Hsieh (CCH; [7]), and Chen, Roll, and Ross (CRR; [8]) also examine APT models with specific macroeconomic factors. CCH and CRR use a traditional Fama-MacBeth type of cross-sectional estimation approach in the premia calculations rather than a FIML approach. Gibbons and Stambaugh have applied FIML techniques to models using only the market as an explanatory variable. Gibbons [20] used a noniterative algorithm on portfolios of firms, equivalent to a two-step cross-sectional approach. Stam-

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baugh [33] used an iterative FIML approach (as we do here) and a variant of the likelihood ratio test employed by Gibbons, on portfolios formed by industry. Both Gibbons and Stambaugh were concerned with testing the constraints that the CAPM imposes, whereas this paper asks whether unanticipated changes in interest rates are priced in a manner consistent with the APT. It is also of at least tangential interest that evidence for or against the pricing of the nonmarket factor constitutes evidence for or against a CAPM of the Sharpe-Lintner-Mossin or Black type.

Section I of the paper shows that, empirically, most of the interest-sensitive stocks are in utilities industries, particularly electric utilities. Section II contains the main results of the paper, focusing on electric utilities. It argues that there is reasonable evidence that the interest factor is priced in the sense of the APT. A brief discussion of the APT and the specific formulation employed here is also provided. Section III looks at data beyond utilities for further evidence of the pricing of interest-rate risk. A main difficulty is reliably detecting interest-rate sensitive stocks. Section IV considers several sources for the interest sensitivity found among utilities and focuses on regulatory lags as a likely candidate. It also discusses whether the detected interest sensitivity is to changes in expected inflation, changes in the expected real interest rate, or both, and presents some further empirical results to support our conjectures. Section V concludes the paper.

I. "Market Model" Results

This section uses industry portfolios, similar to Stambaugh [32], to show that regulated industries, particularly electric utilities, display interest sensitivity. The investigation uses simple changes in interest rates to measure unanticipated interest-rate changes.

Three measures of changes in "the" interest rate were used, with all giving very similar results.[1] The calculations below use changes in the yield on twenty-year government bonds, FYGT20 from the NBER Database. Over the sample period investigated, this variable also equals the residuals of an ARIMA model, since the first difference of FYGT20 appears to be white noise.

If expectations are formed rationally, the shape of the yield curve contains

[1] The first measure was an index of yields on U.S. government bonds with twenty years to maturity; the second measure was three-month U.S. T-bill rates (FYGT20 and FYGM3, respectively, from the NBER Database). A third measure arose because one problem with using a particular index is that some of its movements may be idiosyncratic to that index, rather than due to changes in the overall level of rates. To reduce this source of error, one could use a cross-section relationship for the yield curve in any month, explaining the interest rate (Int) for any time to maturity (τ) as a quadratic relationship

$$\text{Int} = \gamma_0 + \gamma_1\tau + \gamma_2\tau^2,$$

where the equation is viewed as a different cross-sectional relationship for each of the months in a sample. The three time series on γ_0, γ_1, and γ_2 then give a picture of the level of the yield curve over time (see Heller and Khan [23] and also Friedman and Schwartz [19] for similar approaches). Experiments with all three variables—the two indices of yields and the computed $\Delta\hat{\gamma}_i$ series—yielded similar results.

optimal forecasts of how yields will change over time. Hence, if the yield curve is approximately flat in the longer maturities (as it very often is), then the implication is that, at any time, the expected change in long-term yields is (approximately) zero. Any change over time, then, in the yield on long-term bonds with a given maturity reflects changes in the height of the yield curve rather than movements along a yield curve. As a result, changes in long-term yields can be interpreted as unexpected. Calling the long-term yield I_t, the hypothesis that all changes in it are unexpected is $E(\Delta I_t) = 0$, $E(\Delta I_t \cdot \Delta I_{t-j}) = 0$ for all $j \neq 0$. ARIMA models of FYGT20 for the period 1960–1979 essentially conform to these restrictions.

Preliminary work considered a simple market model with the addition of an interest rate change variable,

$$R_{it} = \beta_{i0} + \beta_{i1} R_{Mt} + \beta_{i2} \Delta I_t + \varepsilon_{it} \tag{1}$$

where R_{it} is the return to the i^{th} firm at time t, R_{Mt} is the CRSP value-weighted return to the market, ΔI_t is the change in the interest rate, and ε_{it} is an error term. The above regression (1) was run on equally weighted portfolios formed from all two-digit SIC code industries for the period 1960–1979, using monthly CRSP data. Twenty-one portfolios were formed, similar to Stambaugh [33], save that utilities (code 49) were broken into three portfolios.

The average beta on ΔI for the three utilities portfolios was larger than −70.0 with an average t-statistic above 4.50. Note that firms in these portfolios are subject to substantial regulation. The only other portfolios with significant second betas were "Banking, Finance, and Real Estate" (codes 60–67), where some of the firms included are subject to regulation, with a beta of −28.41 and a t-statistic of 2.28, and "Stone, Clay, and Glass" (code 32) with a beta of −47.98 and a t-statistic of 2.88.[2] However, the betas on ΔI for the latter portfolios were not very stable over subperiods. The average beta on ΔI across portfolios was only about −8.0, not significantly different from zero. The value-weighted sum of these betas is implied by the model to be equal to zero since one of the included factors is the value-weighted return on the market (as Section III discusses[3]).

Some feel for the instability is provided by Table I. For four five-year periods, it shows the squares of t-statistics on the estimated betas for ΔI for the five portfolios with significant second betas for the overall period, 1960–1979. The sum of these for any industry is distributed (asymptotically) χ^2 with four degrees of freedom under the null hypothesis that the coefficients are all zero. Despite the overall period's results for "Stone, Clay and Glass," the β_2 do not appear very significant or stable over subperiods, and the null hypothesis of zero betas cannot be rejected. While the null can be rejected for "Banking, Finance and Real Estate," this is only because of the very large t-statistic (above 4.0) in 1965–1969, with no t-statistic in any other period above 1.5. Based on these results, we concentrate on industry 49, Electric and Gas Utilities, which is dominated by electric utilities.

Table II presents results of regression (1) run over various periods from 1960–

[2] Complete results are available from the authors upon request.

[3] See Sweeney and Warga [36].

Table I

t^2 for Subperiods

Portfolio	1960–1964	1965–1969	1970–1974	1975–1979	Sum
7	0.4877	3.7183	0.7603	2.3159	7.2822
16	6.5149	7.7255	7.8956	2.5878	24.7238
17	1.00322	7.7083	10.1162	3.24	22.0677
18	7.9705	8.6068	7.9741	4.2305	28.7814
21	0.0898	17.0765	1.9666	2.2456	21.3785

χ^2 (4): 95%, 9.48773; 99%, 13.2767; 99.5%, 14.8602

Portfolios (with codes):

1. Mining (10–14)
2. Food and Beverage(20)
3. Textile and Apparel (22, 23)
4. Paper Products (26)
5. Chemical (28)
6. Petroleum (29)
7. Stone, Clay, and Glass (32)
8. Primary Metals (33)
9. Fabricated Metal (34)
10. Machinery (35)
11. Appliances, Electrical Equipment (36)
12. Transportation Equipment (37)
13. Misc. Manufacturing (38, 39)
14. Railroads (40)
15. Other Transportation (41, 42, 44, 45, 47)
16. Utilities (49)
17. Utilities (49)
18. Utilities (49)
19. Department Stores (53)
20. Other Retail Trade (50–52, 54–59)
21. Banking, Finance, and Real Estate (60–67)

Table II

Monthly Portfolio Regression Results for

$R_{pt} = b_0 + b_1 R_{Mt} + b_2 \Delta I_t + e_{it}$

	Constant	b_1	b_2	$\bar{R}^2$
1960–1979	0.0040	0.62	−70.9	0.59
	(2.3)	(14.6)	(5.1)	
1960–1969	0.0036	0.61	−105	0.61
	(18)	(11.2)	(4.4)	
1970–1979	0.0047	0.63	−62	0.58
	(17)	(9.8)	(3.3)	
1960–1964	0.0063	0.70	−108	0.72
	(2.9)	(12.0)	(2.6)	
1965–1969	0.0003	0.53	−100	0.52
	(0.07)	(5.6)	(3.1)	
1970–1974	0.0013	0.55	−73	0.58
	(0.3)	(6.5)	(3.0)	
1975–1979	0.0052	0.77	−54	0.59
	(1.2)	(8.4)	(1.9)	

Note: R_p = rate of return on equally weighted portfolio of 75 utilities; t-statistics are in parentheses.

1979, for a sample of 75 utilities (the majority of which are electrical utilities) formed into an equally weighted portfolio.[4] While ΔI_t is significant, these regressions do not tell us whether this factor is "priced" in the sense that the CAPM tells us the market is, i.e., that an *ex ante* premium is paid proportional to the loading or coefficient β_2 regardless of whether ΔI_t realizes values equal to its mean. The issue is traditionally expressed in the market model literature as one regarding the diversifiability of the second factor.

The following section examines whether the ΔI_t factor is priced in the sense of the APT.

II. Evidence of the Existence of a Premium for Unanticipated Interest-Rate Changes

The APT starts with the assumption that n assets returns, R_{it}, have generating relationships[5]

$$R_{it} = E_i + \beta_{i1}\delta_{1t} + \cdots + \beta_{ik}\delta_{kt} + e_{it} \quad (i = 1, n) \tag{2}$$

where the k factors, δ_j, and the specific errors, e_i, have zero mean and are serially uncorrelated, and where the e_i are orthogonal to each other.[6] The expected return on asset i is then E_i, and a major interest is to infer the structure of the vector of expected returns when there are no arbitrage profit opportunities. Ross [31, 32] shows that, if no arbitrage profits can be made, then (2) implies that approximately[7]

$$E_i = a_0 + \beta_{i1}a_1 + \cdots + \beta_{ik}a_k \quad (i = 1, n) \tag{3}$$

where the $(k + 1)a_j$ are constants; the E_i depend linearly on the betas and $k + 1$ constants. Ross [31] also shows that the a_i's may be viewed as premia $(E^j - E_0)$, where E^j is the expected return from a portfolio with unit beta on factor j and zero betas on all other factors, and $E_0 = a_0$ is either the riskless or zero-beta return.

The empirical model for the APT may be expressed as:

$$R_{it} = E_0 + \beta_{i1}[F_{1t} - E(F_1) + a_1] + \cdots + \beta_{ik}[F_{kt} - E(F_k) + a_k] + e_{it}$$
$$i = 1, \cdots, n; t = 1, \cdots, T \tag{4}$$

where δ_{jt} is set equal to $F_{jt} = E(F_j)$, with $E\delta_{jt} = E[F_{jt} - E(F_j)] = 0$, in preparation for the point made below that we observe the factors F_j, but do not necessarily

[4] Similar results are found in Sweeney and Warga [34].

[5] The discussion is in terms of nominal returns solely for convenience, and all analytical points hold *mutatis mutandis* for real returns models.

[6] In fact, the e_i need display only enough cross-sectional independence to allow the Law of Large Numbers to work to eliminate (approximately) the influence of the e_i in a well-diversified portfolio. The δ_i and e_i need not be orthogonal in principle; however, they must be orthogonal for valid use of OLS.

[7] Huberman [24] provides an alternative proof which does not require an assumption of risk aversion for mathematical rigor (as does Ross). Chen and Ingersoll [7] provide an elegant proof that the Ross pricing equation holds *exactly* in an infinite asset economy with a risk-averse agent. Bounds on an APT for a finite economy are provided in Grinblatt and Titman [21] and Dybvig [9].

know their population means, $E(F_j)$. Separating out constants and variables, rewrite (4) as:

$$R_{it} = E_0 + \sum_{j=1}^{K} \beta_{ij}[a_j - E(F_j)] + \sum_{j=1}^{K} \beta_{ij} F_{jt} + \varepsilon_{it} \quad i = 1, \cdots, n;\ t = 1, \cdots, T. \tag{5}$$

Equation (5) takes advantage of pooled cross-sectional information, in that it can simultaneously yield estimates of betas and (potentially) premia by use of FIML estimation techniques which can constrain the a_j, $E(F_j)$, and E_0 to be identical across firms.

It should be emphasized that *a priori* information on $E(F_j)$ is required to identify the premium on the j^{th} factor rather than just $a_j - E(F_j)$ (see Sweeney and Warga [36]). For the specific implementation here, we have $F_1 = R_M$ and $F_2 = \Delta I$. While we have $E(F_2) = E(\Delta I) = 0$ as Section I argues, we do not know that $E(F_1) = E(R_M)$. However, the market return is itself a portfolio, and thus is naturally scaled so that

$$a_1 = E^1 - E_0 = E(R_M) - E_0$$

since the market beta is equal to one. Notice that $E(R_M)$ thus cancels out of (5), giving:

$$R_{it} = E_0 + \beta_{i1}(-E_0) + \beta_{i2}(E^{\Delta} - E_0) + \beta_{i1} R_{Mt} + \beta_{i2} \Delta I_t + \varepsilon_{it}. \tag{6}$$

Equation (6) highlights the fact that the premium $(E^1 - E_0) = (ER_M - E_0)$ is not estimable. The analogue of this point is that empirical work using the Fama-MacBeth, Black-Jensen-Scholes methodology provides an estimate of $(\bar{R}_M - E_0)$, where $\bar{R}_M$ is the sample mean and $\bar{R}_M - E_0$ is the sample (not *ex ante*) market premium.

While the particular two-factor version of the APT employed here was not derived from a prior model of decision-making, we hoped that the market return would summarize the net effect of most pertinent variables and believed that there are strong reasons for considering the ΔI factor (see the discussion in Section IV).[8] In any case, evidence supporting the pricing of the ΔI factor casts doubt on use of the traditional market model for the purpose of producing "the" risk measure for a firm. Under the CAPM, the β_{i1} risk measure must be the only systematic explanation of cross-sectional differences in mean return (assuming ΔI and R_M are orthogonal). The pricing of ΔI, i.e., the cross-sectional significance of the coefficient on β_{i2}, would be contrary to the CAPM, where CAPM of course refers to the mean variance efficiency of the particular market proxy employed.

However, if ΔI and R_M are not orthogonal, any measured pricing of ΔI may

[8] The model can be more formally justified if we assert that our market proxy is a well-diversified portfolio with zero idiosyncratic risk. (It would be difficult to think of another portfolio that better satisfies these conditions.) In this case, R_M can serve as a factor by just transforming the original factor basis. Our tests for the pricing of the interest-rate variable are correct if the true model is a two-factor model or if the loadings on omitted factors are near zero for the assets (electric utilities) examined here. Flannery and James [17] examine a model similar to ours for the purpose of testing the nominal contracting hypothesis (they employ returns to an index of default free bonds rather than changes in yields). They do not examine whether the interest rate factor is priced.

Table III

Parameter Estimates of a_0 and a_2

$R_{it} = a_0 - a_0\,\beta_{i1} + a_2\,\beta_{i12} + \beta_{i1}R_{Mt} + \beta_{i12}\Delta I + e_{it}$

	A		B		C	
	a_0	a_2	a_0	a_2	a_0	a_2
1960–1979	0.105	−0.599*	0.506	−0.332**	−0.028	−0.492*
	(0.2)	(2.1)	(1.5)	(1.8)	(0.07)	(2.4)
1960–1969	0.228	−0.235*	0.259	−0.203**	−0.08	−0.277**
	(0.8)	(2.0)	(0.7)	(1.9)	(0.21)	(1.7)
1970–1979	0.060	−0.795*	0.348	−0.357	−0.169	−0.702*
	(0.12)	(2.2)	(0.9)	(1.5)	(0.3)	(2.6)
1960–1964	1.27*	−0.0571	1.138*	−0.081	NC	NC
	(3.7)	(0.9)	(3.4)	(1.2)		
1965–1969	−0.293	−0.343*	−0.31	−0.265*	−0.592**	−0.535*
	(1.4)	(2.1)	(0.9)	(2.3)	(1.8)	(3.12)
1970–1974	0.295	−0.349	−0.203	−0.489*	−0.503	−0.707*
	(0.6)	(1.3)	(0.5)	(2.3)	(0.9)	(2.3)
1975–1979	−0.519	−0.784*	−0.654	−0.761*	NC	NC
	(1.4)	(3.3)	(1.4)	(2.7)		

Note: Estimates of a_2 have been multiplied by 10^4. Estimates of a_0 have been multiplied by 10^2. NC denotes nonconvergence of the algorithm. t-statistics are in parentheses.

* Significant at the 95% confidence level.

** Significant at the 90% level.

simply reflect the relation of ΔI to R_M and the fact that R_M is priced (Sweeney and Warga [35]; see also Elton, Gruber, and Rentzler [11]). Thus, discovering a priced second factor need not imply the CAPM is falsified if the market and the second factor are correlated. FIML estimation performed below will be done on both ΔI and an orthogonalized version of ΔI. We proceed first to the pricing of ΔI without concern for whether it is at odds with the CAPM, and without attention to the orthogonality issue.

FIML Estimation. A FIML approach was used to estimate E_0 and the premium on ΔI over various periods.[9] Based on (5) and (6), the form of the equation estimated was:

$$R_{it} = a_0 - a_0\beta_{i1} + a_2\beta_{i2} + \beta_{i1}R_{Mt} + \beta_{i2}\Delta I_t + e_{it}. \tag{7}$$

The investigation focused on a sample of seventy-five NYSE listed utilities with complete returns over the period 1960–1979.

Table III shows estimates of a_0 and a_2 from Equation (7) for four nonoverlapping five-year periods, the two nonoverlapping ten-year periods, and the full twenty-year period. The seventy-five utilities were placed alphabetically in three groups of twenty-five stocks each labeled A, B, and C. Another experiment, with the firms randomly assigned across three groups of twenty-five, gave quite similar results. The estimates of a_2, which under the APT hypothesis is the risk premium on the interest-rate factor, has significant negative values in all five-year periods except for 1960–1964. Both ten-year periods and the twenty-year period provide

[9] The particular algorithm employed was LSQ in the TSP software package (see Hall and Hall [21]).

Table IV

Implied Annual Premia from ΔI (in percent per year)*

	A	B	C	Average
1960–1979	5.2	2.7	4.3	4.07
1960–1969	3.3	2.3	3.5	3.03
1970–1979	5.8	2.6	5.4	4.6
1960–1964	0.9	1.2	NC	1.05
1965–1969	4.5	2.6	6.9	4.7
1970–1974	3.3	4.1	6.2	4.53
1975–1979	4.5	5.0	NC	4.75

* For equally weighted portfolios of firms in groups A to C. The implied annual premium is calculated as the product of β_2 and a_2, where β_2 is the coefficient of ΔI for the portfolio. NC denotes nonconvergence of the algorithm.

significant estimates for virtually every group of stocks save group B in 1970–1979 with a t-statistic for a_2 of -1.49. All of the significant estimates are negative. Table IV lists the implied annual premia for equally weighted portfolios of groups A to C. These are computed by multiplying the portfolio coefficient (always negative), found from regressions such as in Table II, times the premia (also negative) from Table III. Over the various periods and groups, implied annualized premia vary from just under 1% to just over 6%.

Orthogonality of ΔI and R_M. Neither the theoretical validity of the APT nor its empirical implementation requires that the factors be orthogonal to each other. However, with the first factor given to be the market, it could happen that a second factor is priced as found above simply through the market being priced and the second factor being correlated with the market. Sweeney and Warga [35] show that *if* the CAPM holds, pricing gives

$$ER_{it} = E_0 + \beta_{i1}(ER_M - E_0) + \beta_{i2}(ER_M - E_0)\gamma_1$$

where γ_1 is found from the purging relationship

$$\Delta I_t = \gamma_0 + \gamma_1 R_{Mt} + u_t,$$

with u_t an error term. Hence, the premium on ΔI found above may simply reflect the correlation of ΔI and R_M, and the estimates of a_2 may simply be estimates of $(ER_M - E_0)\gamma_1$. This suggests examining the APT for values of ΔI purged of the influence of R_M. First, this helps in discriminating between the APT and CAPM. The CAPM implies the premium on the purged ΔI is zero, while it may be nonzero under the APT (Sweeney and Warga [35]). Second, it offers insight into whether some risk premium is missed in using the standard market model instead of also including ΔI; if the only role of ΔI is through its correlation with R_M, either approach will give the same results asymptotically (although efficiency differs).

Table V shows that $\hat{\gamma}_1$ is sample dependent but still significant over most of our sampling period. One approach is to give full weight to the data by assuming that the sample correlation between R_M and ΔI is the true correlation, and to orthogonalize ΔI with respect to R_M by replacing ΔI with the residuals of a regression of ΔI on R_M. These residuals are, by construction, orthogonal to R_M, and the resulting β_{i2} coefficients on ΔI in the generating relationship will be unchanged on the residual variable.

Table VI presents the estimates of a_2 for this orthogonalized version of the model and the implied annual premia. These estimates are, of course, a deter-

Table V

$$\Delta I = \gamma_0 + \gamma_I R_M + e$$

	$\hat{\gamma}_1$	t-Statistic	Standard Error
1960–1979	−0.11817	−6.4	1.846×10^{-4}
1960–1969	−0.06407	−3.12	2.054×10^{-4}
1970–1979	−0.1485	−5.2	2.856×10^{-4}
1960–1964	−0.0102	−0.55	1.855×10^{-4}
1965–1969	−0.11496	−3.3	3.484×10^{-4}
1970–1974	−0.1541	−3.83	4.024×10^{-4}
1975–1979	−0.1577	−3.75	3.087×10^{-4}

Note: ΔI = first difference of FYGT20 from the NBER database. R_M is the value-weighted return on the market (including dividends) from the CRSP monthly returns file. All figures for γ are multiplied by 100.

Table VI

Estimates of a_2 for Orthogonalized Version of Model* (t-statistics are in parentheses)

	A	B	C
1960–1979	−0.329	−0.110	−0.206
	(1.4)	(0.6)	(1.2)
1960–1969	−0.024	0.005	−0.047
	(0.2)	(0.05)	(0.3)
1970–1979	−0.485	−0.09	−0.358
	(1.6)	(0.4)	(1.6)
1960–1964	−0.0818	−0.105	NC
	(1.32)	(1.5)	
1965–1969	0.133	0.212	−0.024
	(0.9)	(1.8)	(0.17)
1970–1974	−0.280	−0.344	−0.515
	(1.2)	(1.6)	(1.9)
1975–1979	−0.168	−0.124	−0.39
	(0.9)	(0.5)	(2.1)

Implied Annual Premia for Orthogonalized Version of ΔI** (in percent per year)

	A	B	C	Average
1960–1979	2.9	0.9	1.8	1.87
1960–1969	0.3	−0.056	0.6	0.3
1970–1979	3.6	0.7	2.8	2.4
1960–1964	1.3	1.5	NC	1.4
1965–1969	−1.7	−2.1	.31	−1.16
1970–1974	2.7	2.9	4.5	3.4
1975–1979	1.0	0.8	2.5	1.4

Note: NC denotes nonconvergence of the algorithm.

* a_0 is identical in the orthogonalized and nonorthogonalized versions.

** Calculated as in Table III.

Table VII

Premia Implied by the CAPM*

$\hat{a}_2/\hat{\gamma}_1 = (ER_M - E_0)$

	A	B	C	Average
1960–1979	60.79/–	33.71/–	49.9/–	48.14/–
1960–1969	44.02/23.87	38.06/20.64	51.88/28.14	44.65/24.22
1970–1979	64.21/80.65	28.82/36.20	56.72/71.24	49.92/62.70
1960–1964	67.18/5.80	95.6/8.25	NC	81.4/7.0
1965–1969	35.8/34.8	27.77/26.94	55.8/54.3	39.8/38.7
1970–1974	27.16/35.41	38.11/49.68	55.06/71.79	40.11/52.29
1975–1979	59.63/79.56	57.9/77.3	NC	58.8/78.4

Note: Figures are in percentage per year. NC denotes nonconvergence of the algorithm.

* Top figure uses $\hat{\gamma}$ calculated in the same period as $\hat{a}_2$. Bottom figure uses $\hat{\gamma}$ calculated from 1960–1979.

ministic function of the previous estimates and the purging equations. Results are much more mixed now, and at first glance, it is unclear whether one would want to argue that the variable is priced in any of the periods from 1960–1969. Note, however, that potentially, the measurement error introduced by use of purged values of ΔI can account for the decline in the t-statistics.

Further, to assume that the premia in the nonorthogonalized case arose from correlation of ΔI and R_M is to assume the true values of the premia estimated in Table III are $(ER_M - E_0)\gamma_1$, and this causes two serious problems. First, reasonable values of $(ER_M - E_0)$, taken with the estimates of γ_1 in Table V, will suffice to explain only a relatively small portion of the estimates of a_2 in Table III. For example, if the expected excess return on the market is 10% per year, then $(ER_M - E_0)\gamma_1$ explains only 16.4% of the estimated a_2 for group A for 1960–1979, leaving an unexplained portion equal to 1.76 standard errors.[10]

Similarly, the values of $(ER_M - E_0)$ required to account fully for the estimated a_2 are absurdly large. For group A, $(ER_M - E_0)$ would have to have the implausible value of 60.79% per year. Table VII shows implied values of $(ER_M - E_0)$ for the various groups and periods.

Finally, note that any decrease in the assumed absolute value of γ_1 moves the results closer to those in Tables III and IV.

The premia in Table III, then, have some fraction that arises from correlation of ΔI with the priced market. However, a borderline significant and perhaps substantial share is due to ΔI being priced in addition to the market. This means that estimating the required return on electric utilities with the market model, omitting the ΔI factor, will lead to error. The market beta so estimated, when combined with the market premium, will pick up some (small) fraction of the premium due to ΔI. Indeed, Table VII implies that, for the average firm in the

[10] For example, suppose the annual risk premium on the market is 0.1 (in decimal form), so that per month $(ER_M - E_0) = 0.8333 \times 10^{-2}$. Further, suppose the true value of γ_1 is that estimated for 1960–1979 in Table V, or -0.118×10^{-2}. Then, the true $(ER_M - E_0)\gamma_1$ is 0.098333×10^{-4}. The estimated a_2 for 1960–1979 for group A is 0.599×10^{-4}. This means that -0.5007×10^{-4} is unaccounted for by $(ER_M - E_0)\gamma_1$; since the standard error is 0.285×10^{-4}, the unexplained part of a_2 is 1.76 standard errors away from zero.

1960–1979 sample, the analyst would have to assume an annual $(ER_M - E_0)$ of 48% for the usual procedure to capture fully the second premium.

III. Further Tests of the Pricing of Interest-Rate Risk

This section discusses using other data sets to test the pricing of the interest sensitivity found above. Further results are not as straightforward as with utilities. Such tests should look at stocks that: (a) are interest sensitive; (b) display some stability in this sensitivity (to offer some assurance both that parameters are stationary and that the measured sensitivity in any one period is not simply measurement error); (c) show sufficient spread in the β_{i2} to give some precision to estimates; and (d) exist in large enough numbers to give adequate degrees of freedom.

For any particular industry, its sensitivity to ΔI may have any sign, $\beta_2 \gtreqless 0$. To see this, take the generating function

$$R_{it} = E_0 + \beta_{i1}(-E_0) + \beta_{i2}(E^{\Delta} - E_0) + \beta_{i1}R_{Mt} + \beta_{i2}\Delta I_t + \varepsilon_{it} \tag{6}$$

and find the value-weighted sum for t,

$$R_{Mt} = E_0\Sigma w_i + (-E_0)\Sigma w_i\beta_{i1} + (E^{\Delta} - E_0)\Sigma w_i\beta_{i2} + R_{Mt}\Sigma w_i\beta_{i1} + \Delta I_t\Sigma w_i\beta_{i2} + \Sigma w_i\varepsilon_{it}, \tag{8}$$

where Σ is over all i. By definition, $\Sigma w_i = 1$. The only way for (8) to hold for all realizations of R_M, ΔI and ε is for $\Sigma w_i\beta_{i1} = 1$, $\Sigma w_i\beta_{i2} = 0$, and $\Sigma w_i\varepsilon_{it} = 0$. These restrictions hold at the theoretical level if (6) is true, and also as an empirical matter if (6) is estimated simultaneously over all N securities in R_M.

The interpretation of $\Sigma w_i\beta_{i2} = 0$ is that the (value-weighted) average of β_{i2} is zero. In other words, the β_{i2} in (6) measures the interest sensitivity of any industry *relative* to the market's sensitivty. The negative β_2 for utilities means they are substantially more harmed by unanticipated increases in interest rates than is the average NYSE firm.

While the pricing of interest-rate risk in the utilities industry means it is priced for *all* interest-sensitive stocks, empirically it may be difficult finding reliably interest-sensitive stocks. In Table II, the average β_{i2} for electric utilities was about −70.0. Thus, because the value-weighted sum of the β_{i2} must equal zero, nonutilities β_{i2} must be slightly positive on average. This relationship can be satisfied by having many nonutilities β_{i2} that are not statistically significant, and this is what is observed. In turn, it is difficult to select nonutilities industries with interest sensitivity for further testing.

For example, as Section I discussed, there were only two nonutilities portfolios, of the twenty-one examined, that showed significant interest sensitivity. Unfortunately, these two portfolios showed substantial instability over subperiods in Table I.[11]

[11] Searching for industries to examine implicitly assumes that statistically significant, stable sensitivity at the industry level is accompanied by the same characteristics for firms in the industry, and that there is sufficient spread among firms' β_{i2}. Neither assumption necessarily holds, of course, but the procedure seems to have worked for utilities. We can, and did, check to see if there is spread

Table VIII

Parameter Estimates of a_0 and a_2

$R_{it} = a_0 - a_0\beta_{i1} + a_2\beta_{i2} + \beta_{i1}R_{Mt} + \beta_{i2}\Delta I_t + \varepsilon_{it}$

	a_0	a_2
1960–1979	−1.38*	−1.166*
	(3.04)	(4.41)
1960–1969	−8.293	0.019
	(0.26)	(0.17)
1970–1979	−0.180	−0.989*
	(0.51)	(3.8)
1960–1964	NC	NC
1965–1969	NC	NC
1970–1974	2.326**	3.048*
	(1.92)	(2.0)
1975–1979	−0.256	−0.246
	(0.60)	(1.2)

Note: Estimates of a_2 have been multiplied by 10^4. Estimates of a_0 have been multiplied by 10^2. NC denotes nonconvergence of the algorithm.

* Significant at 95% level.

** Significant at 90% level.

Because of these difficulties, it seemed reasonable to focus on a large part of the entire NYSE rather than on other single industries. Hence, we used the twenty-one portfolios discussed in Section I. Table VIII shows results for FIML estimation. The a_2 estimates for the entire period and for the 1970–1979 period are somewhat higher than those in Table III but seem to be broadly consistent. The a_2 estimate for 1960–1969 is quite insignificant, and the algorithm did not converge for the two 1960's subperiods. The 1970's subperiods are mixed, with one negative but insignificant, the other significant but positive.

Clearly, the results are not as impressive as for the utilities. We conjecture that this is due to two reasons. First, using industry portfolios has masked some spread of β_{i2} within an industry. This is a general problem with portfolio formation; one must take care not to average out in the portfolio the phenomenon that is being investigated. Second, too large a portion of the spread in β_{i2} across industries is due to measurement error; the average nonutilities β_{i2} is insignificant at conventional levels. In other words, Table VIII may well be showing the results of trying to measure the pricing of interest sensitivity when only one-seventh of the portfolios (the three utilities portfolios) are in fact reliably interest sensitive.

These considerations suggest that further testing may require searching for individual firms, not industries, that have some on-going interest sensitivity. This is not simple since stock returns are notoriously noisy; it is thus hard to tell if a single security that shows interest sensitivity in some periods should be

of the β_{i2} across firms; based on the types of tests described in Warga [37], greater spread in the second betas for utilities would have been highly desirable, but evidently the spread was adequate for acceptably precise estimates. Checking on stability and significance at the firm level is difficult because of the noise for individual firms. In effect, we used the industry characteristics in this regard as indicators of these characteristics of firms.

judged interest sensitive or merely to have substantial measurement error. It was in fact the clear interest sensitivity of portfolios of utilities, as in Table II, that led to the focus on them above.

IV. Reasons for Utilities' Interest Sensitivity

The β_{i2} are measures of a stock's interest sensitivity relative to the market's. A natural question is why utilities appear so much more sensitive than average. Changes in interest rates are by definition the sum of changes in real rates and changes in expected inflation. Up to the late 1970's (and the results in this paper go through 1979), many economists viewed a constant real rate as a very reasonable first approximation (e.g., Fama [12]). In this view, the interest sensitivity of utilities would have to be explained by their sensitivity to unanticipated changes in expected inflation. Now there is substantial evidence that ex ante real rates vary (e.g., Fama and Gibbons [13]). Thus, explaining utilities' interest sensitivity must also take into account real interest rate changes. We start by looking at explanations that focus on inflation, holding constant ex ante real rates. Note that use of long-term government bond yields implies that the expected inflation rate considered is also long term; we look at short-term rates below.

Explanations Based on Inflation. The negative sign for β_{i2} reported above implies that, if expected inflation rises from the end of one month to the next, this will make the stock price change smaller than otherwise, or *ceteris paribus*, reduce price. This is interpreted as the net effect of four separate influences. First, higher inflation implies a lower real present value of interest payments implied by already outstanding debt and would thus raise the stock's price (Alchian and Kessel's [1] effect of unanticipated inflation). Second, the "regulatory lag" implies the utility will face (perhaps ongoing) lags between rises in its costs before being allowed to raise it output's price; this will reduce the stock's price. Third, current debt will have to be rolled over at higher interest rates due to the higher inflation; since the higher interest will be paid with dollars with lower real value, this may be a "wash." Fourth, Logue and Sweeney [27] provide cross-country evidence that increases in inflation bring greater real economic instability, and argue that this is costly and reduces the real return on capital. On balance, then, regulatory lags plus the generalized phenomenon of more erratic growth at higher inflation rates seem to cause the negative coefficient on ΔI. However, the majority of firms with significant loadings on ΔI come from regulated industries. Thus, regulatory lags seem to be mainly responsible here, although the inferential support for this view has to be regarded as quite tentative.

There is some additional support for the regulatory lag hypothesis in Table II. As noted in Clarke [9], the utility industry as a whole experienced a decreasing reliance on the regulation of prices over the sample period in this paper. Automatic fuel adjustment clauses (FAC's) came into more widespread use in the early 1970's for rates charged to residential customers (many commercial users were subject to FAC's before 1970). This effect should reveal itself by having coefficients on an inflation factor that decrease over the sampling period. Table

II reveals that the β_{i2} coefficients decreased monotonically in magnitude over the four sampling periods.

The negative β_2's found above are all the more striking in light of the first interpretation, where unanticipated inflation should reduce the real value of interest payments on outstanding debt and thus raise stock prices, since electric utilities have higher than average leverage. However, empirical evidence seems to have a hard time substantiating the debt-unanticipated inflation relationship (see French, Ruback and Schwert [18]). Bernard [2] has found support for the nominal contracting hypothesis on a small sample of firms, but in doing so has revealed that the effect of unanticipated inflation on real cash flows dominates the nominal contracting (or debt-unanticipated inflation) relationship. Bernard did not isolate utilities in his study, but the approach he takes does offer the potential for discerning the relative importance of various effects—nominal contracting or regulatory lag—in the observed utility sensitivity to unanticipated changes in interest rates.

Explanations Based on Real Interest-Rate Changes. There are many channels through which changes in real interest rates can affect firms, and utilities more than average. However, to illustrate we focus on one, the present value of dividends. Many researchers have considered how dividend policy can *ceteris paribus* affect firm values, through tax effects, clientele effects, and so forth; the evidence on dividend effects is mixed and controversial (Black and Scholes [4], Litzenberger and Ramaswamy [26], Miller and Scholes [29]). However, utilities pay higher than average dividends, so it is possible that their stockholders pay a premium for this. If this premium is based on the present value of dividends, a rise in the real interest rate will reduce the present value of utilities' dividends more than low dividend industries', and thus will reduce the relative attractiveness of utilities. Thus, the interest-rate sensitivity discovered above may simply be revealing a type of dividend effect. Note that, in the absence of regulatory lag and a nominal contracting effect, increases in interest rates due to inflation with a constant real rate should leave the present value of real dividends unaffected.

Regulators' behavior may also play a role in this explanation based on changes in real rates. If interest rates rise, either due to real rates or expected inflation, and regulators do not allow a complete pass-through, the real present value of utilities' dividends is reduced.

It is a major project to try to measure the relative importance of changes in expected inflation and in real rates. Nevertheless, some simple tentative evidence is available. We consider two types of evidence that bear on the issue.

(A) Alternative Interest Rates. If the second factor is truly a real rate factor, it would be expected that an even better variable would be an index of yields on public utility bonds, i_p, where i_p is FYPUT from the NBER Database. In fact, F-tests reveal that Δi_p adds explanatory power to the single-factor model in only a relatively few cases, and then mostly at the 10% level. A plausible interpretation is that ΔFYGT20 (ΔI) is measuring the change in expected long-term inflation rate. Δi_p measures the change in interest costs to utilities, and this is shown to add little to the single-factor market model.

When changes in the three-month T-bill rate are used in place of ΔI, the results are weaker, but quite similar. This is not surprising since such changes show correlation above 0.95 with ΔI. This seems to argue against necessarily interpreting the second factor as unanticipated changes in expected *long*-term inflation. However, the point to emphasize is that daily observations on short-term interest rates—say one month—can be much more volatile than monthly observations on daily averages of rates on either three-month or 20-year instruments. In periods of sharply negatively sloped yield curves in, say, 1974, 1980, and 1981, very short rates exaggerated the slope as compared to even three-month rates, and monthly averages usually made the yield curve smoother. Further, such very short rates may be more heavily influenced by liquidity-shortage pressures instead of expected inflation than are longer rates.

(*B*) *Models of Changes in Expected Inflation and Real Interest Rates.* Explicit models for expected inflation and expected real rates, based on short-term interest rates and consumer price indices (CPI), have been investigated by many researchers. Two recent examples appear in Fama and Gibbons [14] and Huizinga and Mishkin [25]. Fama-Gibbons present a simple model where expected inflation equals the difference between the T-bill rate and a proxy for the expected real rate (comprised of an equally weighted moving average of the past twelve months of ex post real rates). Ex post real rates are always measured as the difference between the treasury bill and an inflation rate measured by the CPI. Huizinga and Mishkin [25] produce a measure of expected inflation based on a projection of ex post inflation rates on a constant, the one-month nominal interest rate (T-bill), the one-month inflation rate lagged one and two months, and a supply shock variable lagged one month (log of the relative price of fuel and related products in the producer price index).

A major difference in these two studies is their use of different measures of inflation. Fama and Gibbons examine measures based on the more commonly used CPI-U index, which employs a mortgage interest subcomponent in the measure of housing costs. Huizinga and Mishkin employ a revised measure (now used by the Bureau of Labor Statistics as the primary CPI measure) that is less dependent on fluctuations in mortgage rates. The CPI employed by Huizinga and Mishkin measures homeowners' costs as the homeowners' equivalent rent and household insurance (the owner's equivalent rent is defined to be the income the homeowner foregoes by choosing not to rent his house). In replicating the Fama-Gibbons work, we replaced the CPI index they used with that employed by Huizinga-Mishkin; we obtained the Huizinga-Mishkin data from the authors.

Tables IX and X show the results of regressions of the return to an equally weighted portfolio of 54 electric utilities on the value-weighted return to the market and on changes in the expected inflation rate ($\Delta EInf$) and on changes in the expected real rate of interest (ΔER), measured as described above. Both data sets support the view that changes in expected inflation negatively affect utilities' returns. The Huizinga-Mishkin measures provide support for the view that changes in real interest also have an important negative effect, while the Fama-Gibbons measures give less support to the view that ΔER is important.

Mismeasurement in the Market Index. A final possibility considered was that electric utilities might be very sensitive to the lack of an interest rate component in R_M, which is after all a proxy for the true market return in a CAPM view. The

Table IX

$R_{pt} = \beta_0 + \beta_1 R_{Mt} + \beta_2 \Delta EInf_t + \beta_3 \Delta ER_t + e_{pt}$ for Huizinga-Mishkin Measures

Period	β_0	β_1	β_2	β_3	$\bar{R}^2$	DW
1960–1979	0.0017	0.683	−12.02	−12.72	0.51	1.80
	(0.90)	(15.0)	(2.50)	(2.40)		
1960–1969	0.0015	0.668	−8.54	−9.26	0.50	1.66
	(0.70)	(10.60)	(1.10)	(1.10)		
1970–1979	0.0019	0.690	−13.23	−13.8	0.51	1.86
	(0.60)	(10.40)	(2.00)	(1.90)		
1960–1964	0.0065	0.726	−0.18	−6.65	0.67	1.40
	(2.70)	(10.90)	(0.02)	(0.70)		
1965–1969	−0.0030	0.573	−17.50	−9.72	0.38	1.98
	(0.80)	(5.40)	(1.40)	(0.70)		
1970–1974	−0.0009	0.686	−10.36	−23.28	0.53	1.65
	(0.20)	(8.10)	(1.30)	(2.50)		
1975–1979	0.0049	0.682	−18.19	−5.52	0.49	1.64
	(0.70)	(5.70)	(1.40)	(0.40)		

Note: R_p = rate of return on equally weighted portfolio of 54 electric utilities; $\Delta EInf$ = change in expected rate of inflation; ΔER = change in expected real interest rate; t-statistics are in parentheses.

Table X

$R_{pt} = \beta_0 + \beta_1 R_{Mt} + \beta_2 \Delta EInf_t + \beta_3 \Delta ER_t + e_{pt}$ for Fama-Gibbons Measures

Period	β_0	β_1	β_2	β_3	$\bar{R}^2$	DW
1960–1979	0.0017	0.682	−12.29	−13.35	0.51	1.80
	(0.9)	(15.0)	(2.6)	(1.3)		
1960–1969	0.0016	0.662	−8.14	1.48	0.50	1.66
	(0.7)	(10.7)	(1.1)	(0.1)		
1970–1979	0.00175	0.696	−13.48	−21.91	0.51	1.86
	(0.6)	(10.5)	(2.1)	(1.4)		
1960–1964	0.0067	0.714	−1.40	−7.20	0.66	1.42
	(2.8)	(10.9)	(0.2)	(0.46)		
1965–1969	−0.0028	0.571	−13.08	7.63	0.39	1.90
	(0.8)	(5.5)	(1.0)	(0.3)		
1970–1974	−0.0019	0.671	−13.28	−36.07	0.53	1.65
	(0.4)	(7.9)	(1.76)	(1.9)		
1975–1979	0.0037	0.712	−10.54	−3.61	0.47	1.57
	(0.75)	(5.9)	(0.81)	(0.12)		

Note: R_p = rate of return on equally weighted portfolio of 54 electric utilities; $\Delta EInf$ = change in expected rate of inflation; ΔER = change in expected real interest rate; t-statistics are in parentheses.

regressions reported in Table II were rerun with the market return R_M having both long- and short-term government rates added, according to their market valuation. The weights were drawn from Stambaugh [33], and as could be expected from that study, virtually no changes were observed in either parameter estimates or their statistical significance.

V. Conclusions

This paper provides evidence that changes in government bond yields clearly affect ex post returns to electric utilities, and that this phenomenon is concentrated to a much larger extent in this particular industry than in NYSE firms as a whole (Table I). Evidence of the pricing is provided within the framework of the APT. Empirical work employed a pooled cross-section approach using a nonlinear constrained version of FIML estimation.

The interest rate factor can be thought of as equal to unanticipated changes in expected inflation plus changes in the real rate of interest. This interpretation allows us to consider several mechanisms, in particular regulatory lags, as responsible for the effect. The identification of regulatory lags would indicate that the cause of the interest rate loading is a recognized source of risk in the market, and hence help rationalize the evidence that this factor is priced by market participants. Furthermore, this sensitivity to changes in interest rates seems to be priced, in the sense that ex ante returns incorporate a premium based on the risk of interest-rate changes, with the premium proportional to the stocks sensitivity to these changes.

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