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THE PRICING OF UNANTICIPATED CHANGES IN EXPECTED INFLATION: EVIDENCE FROM THE STOCK MARKET

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1. Introduction

This paper examines whether firms that exhibit *ex post* sensitivity in their stock returns to proxies for unanticipated changes in expected inflation are required to pay investors *ex ante* for this risk. The framework for examining pricing of the inflation variable 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 inflation premium is significant. To study this, the paper uses full information maximum likelihood (FIML) estimation on systems 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 expected inflation is priced in the market is of particular interest with regard to regulated firms. The well known regulatory lag problem implies that regulated firms may be harder hit by unexpected inflation than firms that can can easily adjust their prices. Regulated firms (especially electric utilities) turn out to comprise the vast majority of firms that are sensitive to the inflation factor. After some preliminary investigation, the analysis focuses on electric utilities.

Chan, Chen, and Hsieh (CCH, 1983), and Chen, Roll, and Ross (CRR, 1983), also examine APT models with specific macroeconomic factors. These factors include (among others) both the market and proxies for unanticipated inflation, as we do. CCH and CRR use a traditional Fama-MacBeth type of cross-sectional estimation approach in their premia calculations rather than a FIML approach. The FIML approach taken here has appeared before, but applied to the simple CAPM using only the market. Gibbons (1981) used a non-iterative algorithm on portfolios of firms equivalent to a two-step cross-sectional approach. Stambaugh (1982) 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 expected inflation are priced in a manner consistent with the APT. The size of the premium is also of interest; thus efficient estimates of this parameter take on added importance and use of FIML is appropriate. It is also of at least tangential interest that evidence for or against the pricing of the non-market factor constitutes evidence for or against a CAPM (of the Sharpe-Lintner-Mossin or Black type).

Section 2 of the paper argues that the various interest-rate-change variables employed here may be proxying for unanticipated change in expected inflation. It also shows that empirically most of the interest-sensitive stocks are in regulated industries, particularly electric utilities. Section 3 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. In section 4 several sources for the pricing mechanism are considered, and the role of regulatory lags is focused on as a likely candidate. Some further empirical results are presented to support the claims made. Section 5 concludes the paper.

2. INTERPRETATION OF THE INTEREST RATE FACTOR AND "MARKET MODEL" RESULTS

This section (i) discusses using interest rates as forecasts of expected inflation, and changes in interest rates as changes in such forecasts, and (ii) uses industry portfolios, similar to Stambaugh (1982), to show that regulated industries, particularly electric utilities, are those that display interest sensitivity.

Interest Rate-Changes as Changes in Expected Inflation. A number of authors have investigated the hypothesis that the expected real rate of interest is constant, and hence changes in nominal yields equal changes in the expected rate of inflation (Fama (1975), Fama and Schwert (1977), Schwert (1981)). While there is evidence that expected real rates are not constant (Nelson (1976), Nelson and Schwert (1977), Hess and Bicksler (1975), Fama and Gibbons (1981)), some have argued that movements in nominal interest rates are dominated by changes in expected inflation rather than in expected real rates (Fama and Schwert (1977)), and by the same token changes in interest rates can serve adequately as a proxy for changes in expected inflation (and this seems plausible at least for many of the periods examined here). For example, Schwert's (1981) investigation of returns on the market uses the one-month T-bill rate to proxy the one-month expected inflation rate, but reports that results using an ARIMA model of inflation gives no substantial differences. The empirical results presented below are, of course, independent of whether the interest rate-change factor is interpreted as unanticipated changes in expected inflation.

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

Experiments using changes in long term bond yields are implicitly using forecasts of longer term inflation rates. Nominal yields for bonds with maturity n can be interpreted equal to the expected average inflation rate, less the (assumed constant) expected average real rate over n periods. If expectations are formed rationally, the shape of the yield curve contains optimum forecasts of how expected inflation 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 run inflation 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 <u>unexpected</u> changes in longer-run expected inflation rates. 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.

We turn now to how such changes in expected inflation affect returns on electric utilities.

<u>"Market Model" Evidence on the Interest Factor</u>. Preliminary work considered a simple market model with the addition of an interest rate change variable,

$$R_{it} = \beta_{io} + \beta_{i1}R_{Mt} + \beta_{i2}\Delta I_t + \varepsilon_{it} ,$$

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, and ΔI_t is a change in interest rate variable and ε_{it} is an error term. The above regression was run on equally weighted portfolios of all two-digit SIC code industries. Regulated industries appeared most sensitive to the interest change factor, and electric utilities were consistently the most sensitive of all firms, displaying significant loadings on ΔI_t in all sample periods.[2]
Table 1 presents results of this regression run over various periods from 1960-1979, for a sample of 75 electric utilities formed into an equally weighted portfolio. 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 <u>ex ante</u> 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.

3. EVIDENCE OF THE EXISTENCE OF A PREMIUM FOR UNANTICIPATED CHANGES IN EXPECTED INFLATION

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

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

where the k factors δ_j and the specific errors e_i are zero mean and serially uncorrelated, and the e_i are orthogonal to each other.[3] 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 (1976, 1977) shows that if no arbitrage profits can be made then (1) implies that approximately[4]

$$(2) \quad E_i = a_o + \beta_{i1}a_1 + \ldots + \beta_{ik}a_k \qquad (i=1,n)$$

where the (k+1) a_j are constants; the E_i depend linearly on the β's and k+1 constants. Ross (1977) also shows that the a_i's may be viewed as premia $(E^j - E_o)$, where E^j is the expected return from a portfolio with unit beta on factor j and zero betas on all other factors, and $E_o = a_o$ is either the riskless or zero beta return.

The empirical model for the APT may be expressed as:

$$(3) \quad R_{it} = E_o + \beta_{1i}[F_{1t}-E(F_1)+a_1] + \ldots + \beta_{ki}[F_{kt}-E(F_k)+a_k] + e_{it} \qquad i=1,\ldots,n;\ t=1,\ldots,T$$

where δ_{jt} is set equal to $F_{jt} - E(F_j)$ in preparation for the assumption made below that we observe the factors F_j, but do not necessarily know their population means $E(F_j)$. Separating out constants and variables, rewrite (3) as:

$$(4) \quad R_{it} = E_o + \sum_{j=1}^{k} \beta_{ij}\ [(a_j) - E(F_j)] + \sum_{j=1}^{k} \beta_{ij} F_{jt} + e_{it}$$

$$i=1,\ldots,n; \quad t=1,\ldots,T.$$

Equation (4) takes advantage of pooled-cross sectional information in that it can simultaneously yield estimates of β's and (potentially) premia by use of FIML estimation techniques which can constrain the a_j, $E(F_j)$, and E_o 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 $a_j - E(F_j)$. For the specific implementation here we have $F_1 = R_M$ and $F_2 = \Delta I$. While we have $E(F_2) = E(\Delta I) = 0$, we do not have knowledge of $E(F_1) = E(R_M)$. However, the market return is itself a portfolio, and thus is naturally scaled so that

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

since the market beta is equal to one. Notice that $E(R_M)$ thus cancels out of (4), giving;

$$(5) \quad R_{it} = E_o + \beta_{i1}(-E_o) + \beta_{i2}(E^{\Delta I}-E_o) + \beta_{i1}R_{Mt} + \beta_{i2}\Delta I_t + e_{it}$$

(5) highlights the fact that the premium $(E^1-E_o) = (ER_M-E_o)$ is not estimable. The analogue of this point is that empirical work using the Fama-MacBeth, Black-Jensen-Scholes methodology can at best provide an estimate of $(\bar{R}_M-E_o)$,[5] where $\bar{R}_M$ is the sample mean and $\bar{R}_M - E_o$ 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 below on the regulatory lag problem).[6] 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 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 simply reflect the relation of ΔI to R_M and the fact that R_M is priced (Sweeney and Warga (1984); see also Elton, Gruber, and Rentzler (1983)). Thus, discovering a priced second factor need not imply the CAPM is falsified if the market and the second factor are correlated. FIML estimation carried out 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_o and the premium on ΔI over various periods. The particular algorithm employed was LSQ in the TSP software package (see Hall and Hall (1981)). The form of the equation estimated was:

$$(6) \quad R_{it} = a_o - a_o \beta_{i1} + a_2 \beta_{i2} + \beta_{i1} R_{Mt} + \beta_{i2} \Delta I_t + e_{it}$$

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

Table 2 shows estimates of a_o and a_2 from equation (6) for four non-overlapping five year periods the two non-overlapping 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. The estimates of a_2, which under the APT hypothesis is the risk premium on the inflation factor, has significant values in all five year periods except for 1960-64. Both ten year periods and the twenty year period provide significant estimates for virtually every group of stocks save the group B in 1970-79 with a t-statistic for a_2 of -1.49. All of the significant estimates are negative. Table 3 lists the implied annual premia for equally weighted portfolios of groups A-C. These are computed by multiplying the portfolio coefficient (always negative), found from regressions such as in Table 1, times the premia (also negative) from Table 2. Over the various periods and groups, implied annualized premia vary from just under one percent to just over six percent.

Orthogonality of ΔI and R_M. Neither the theoretical validity of the APT nor its empirical implementation require that the factors be orthogonal to each other. However, with the first factor 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 (1984) show that if the CAPM holds, pricing gives

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

where γ_1 is found from the purging relationship

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

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 $(ER_M - E_o)\gamma_1$. 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 non-zero under the APT (Sweeney and Warga (1984)). 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 (though efficiency differs).

Table 4 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 5 presents the estimates of a_2 for this orthogonalized version of the model and the implied annual premia. 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 1960-69. 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 non-orthogonalized case arose from correlation of ΔI and R_M is to assume the true values of the premia estimated in Table 2 are $(ER_M - E_O)\gamma_1$, and this causes two serious problems. First, reasonable values of $(ER_m - E_O)$, taken with the estimates of γ_1 in Table 4, will suffice to explain only a relatively small portion of the estimates of a_2 in Table 2. Second, and related, the values of $(ER_m - E_O)$ required to account fully for the estimated a_2 are absurdly large. For example, suppose the annual risk premium on the market is .1 (in decimal form), so that per month $(ER_M - E_O) = .8333 \times 10^{-2}$. Further, suppose the true value of γ_1 is that estimated for 1960-79 in Table 4, or $-.118 \times 10^{-2}$. Then, the true $(ER_M - E_O)\gamma_1$ is $--.098333 \times 10^{-4}$. The estimated a_2 for 1960-79 for group A is $-.599 \times 10^{-4}$. This means that $-.5007 \times 10^{-4}$ is unaccounted for by $(ER_M - E_O)\gamma_1$; since the standard error is $.285 \times 10^{-4}$, the unexplained part of a_2 is 1.76 standard errors away from zero.

Another way of looking at the same data is to ask what values of $(ER_M - E_O)$ would give the estimated annualized premia in Table 3. For group A, $(ER_M - E_O)$ would have to have the implausible value of 60.79% per year. Table 6 shows implied values of $(ER_M - E_O)$ for the various groups and periods.

Finally, note that any decrease in the assumed value of γ_1 moves the results closer to those in Tables 2 and 3.

The premia in Table 2, 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 required rates of returns 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 5 implies that for the average firm in the 1960-1979 sample, the analyst would have to assume an annual $(ER_M - E_O)$ of 48% for the usual procedure to capture fully the second premium.

4. INTERPRETATION OF THE EFFECT OF UNANTICIPATED CHANGES IN THE EXPECTED INFLATION RATE

The effect of unanticipated changes in expected long run inflation rates could, in principle, have either sign or even be zero. The negative sign 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 (1959) 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 its 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 (1981) provide cross-country evidence that increases in inflation bring greater real economic instability, and argue 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 a significant loading on ΔI come from regulated industries. Thus, regulatory lags seem to be mainly responsible here, though the inferential support for this view has to be regarded as tentative.

It might be argued that the first interpretation, as a result of unanticipated inflation on the real value of interest payments on outstanding debt explains the results, 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 (FRS, 1983)). Bernard (1983) 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 which effect--nominal contracting or regulatory lag--is the primary cause of the observed utility sensitivity to unanticipated changes in expected inflation.

<u>Changes in Expected Inflation v. An Interest Rate Factor</u>. It might be argued that the ΔI coefficient in Table 1 is really measuring the effect of changes in interest rates rather than

changes in expected long term inflation rates. However, if the second factor is truly an interest 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 3 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 .95 with ΔI. This seems to argue against necessarily interpreting the second factor as unanticipated changes in expected <u>long</u> term inflation. However, the point to emphasize is that daily observations on short-term interest rates--say 1 month--can be much more volatile than monthly observations on daily averages of rates on either 3-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 3-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.

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 basic regressions (i.e., those reported in Table 1) where rerun with the market return R_M having both long and short term government rates added in according to their market valuation. The weights were drawn from Stambaugh (1982), and as could be expected from that study, virtually no changes were observed in either parameter estimates or their statistical significance.

5. 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. A reasonable interpretation of the interest rate factor is as a proxy for unanticipated changes in expected inflation. 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 this factor might be priced by market participants. Evidence of the pricing is provided within the framework of the Arbitrage Pricing Theory (APT). Empirical work employed a pooled cross-section approach using a non-linear constrained version of Full Information Maximum Likelihood estimation.

FOOTNOTES

[1]Three measures of changes in "the" interest rate were used, with all giving very similar results. The first measure was an index of yields on U.S. government bonds with 20 years to maturity; the second measure was 3-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 ( )as a quadratic relationship

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

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 (1979) and also Friedman and Schwartz (1963) for similar approaches.) Experiments with all three variables -- the two indices of yields and the computed $\Delta\hat{\gamma}_i$ series -- yielded similar results.

[2]Complete results are available from the authors upon request. The betas on ΔI for twenty-one industry-based portfolios were sometimes negative, sometimes positive. Utilities were broken into three portfolios, with the average beta on ΔI being larger than -60.0 with an average t-statistic above 4.50. The only other portfolios with significant second betas were "Banking, Finance, and Real Estate," with a beta of -28.41 and a t-statistic of 2.28, and "Stone, Clay, and Glass" with a beta of -44.98 and a t-statistic of 2.88. However, the beta on ΔI for the latter portfolio was not very stable over sub-periods.

The average beta on DFY was only about -7.0, not significantly different from zero. The value-weighted sum of these betas is implied by the model to be equal to zero.

[3]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.

[4]Huberman (1981) provides an alternative proof which does not require an assumption of risk aversion for the mathematical rigor (as does Ross). Chen and Ingersoll (1982) 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 (1983), and Dybvig (1983).

[5]The purpose of this footnote is to show that the existence of a parameter representing the premium λ_i (corresponding to factor i) in an estimable version of the APT relies critically on whether or not the factor used is observed in a mean zero form or else has a known mean. Without loss of generality we will assume that the observed proxy factors are exactly the portfolios of assets perfectly correlated with the true factors, so that issues of scale can be ignored. In the infinite asset APT's of Grinblatt and Titman (1983) and Dybvig (1983), it is always the case that portfolios exist which are perfectly correlated with the underlying factors. It is important to emphasize that the degree of indeterminacy concerning the estimability of premia discussed here is independent of errors in variables or any misspecification issues. The only "wrinkle" to be introduced here is the possibility that the factor's mean is unknown and/or that the factor is not available measured mean zero.

The APT generating relationship is

$$(1)\quad R_t - E = \beta\delta_t + \varepsilon_t = \beta(F_t - E(F)) + \varepsilon_t$$

where R_t an n-vector of asset returns at time t, E is the n vector of (time independent) mean asset returns, β is an n by k matrix of factor loadings, $\delta_t = F_t - E(F)$ is the k by 1 vector of orthogonal common factor realizations at time t, and ε_t is the n by 1 vector of time t idiosyncratic errors. The APT result is that

$$(2)\quad E = E_0 + \beta\lambda$$

where E_0 is the constant vector representing a risk free rate or zero-beta return and λ is the k-vector of factor premia. Combining (1) and (2) yields

$$(3)\quad R_t = E_0 + \beta(F_t - E(F) + \lambda) + \varepsilon_t$$

$$(4)\quad = E_0 + \beta(\lambda - E(F)) + \beta F_t + \varepsilon_t$$

We now make two points. The first is the assertion that the risk premium on the market cannot be estimated by usual cross section-time series approaches. Suppose (for expositional purpose only) we have a single factor model with F_t already measured in units such that $F_t = R_{pt}$, where R_{pt} is a portfolio of assets. Then we have

(5) $R_{pt} = E_o + \beta_p(\lambda - E(F)) + \beta_p R_{pt} + \varepsilon_{pt}$

and this can be true if and only if

$$\lambda = E(F) - E_o = E(R_p) - E_o$$

and

$$\beta_p = 1, \quad \varepsilon_{pt} = 0$$

Thus, (4) can be rewritten

(6) $R_t = E_o + \beta(-E_o) + \beta R_{pt} + \varepsilon_t$

Equation (6) makes clear the fact that $\lambda = (E(F) - E_o) = (E(R_p) - E_o)$ no longer appears parametrically in the model. The motivation behind this exposition is the consideration of the market return as a factor. In this case the market premium is clearly not estimable. Any inference concerning the market premium using (6) must be based on some <u>other</u> model relating R_{mt} to its mean. This was the case in the cross sectional work in Black, Jensen and Scholes (1972) and Fama and MacBeth (1973), where estimates of $\bar{R}_M - E_o$ (where $\bar{R}_M$ is the sample mean) were estimated. To the extent that the sample mean is undesireable as an estimator of the expected market return (it can be negative), then so must estimates of it be. In any event, the usefulness of equation (6) for the purpose of investigating the premium to a portfolio that is also a factor is clearly non-existent. This is not to say that the APT never contains information about a premium to a factor that is observed in a similar manner to the market premium (i.e., not in mean zero form and with an unknown mean). Relationships between premia exist (our example above had only one), and if the premium on one factor may be estimated (as we show below), then inference about premia on other factors may be carried out (see Sweeney and Warga (1984) for a discussion of this).

For our second point, suppose we observe a factor in a mean zero form, although we may not know the mean of the factor F_t (think of unanticipated changes in expected inflation). Again consider for expositional purposes a single factor model with $\delta_t = F_t - E(F)$ observed. Then (3) yields

(7) $R_t = E_o + \beta\lambda + \beta\delta_t + \varepsilon_t$

and we see that a constrained pooled cross-sectional estimation technique can provide us with estimates of E_o, β, and λ. This also makes clear that knowledge of the mean of a factor, either by observing the variable in a mean zero form or by actually knowing E(F), allows λ to appear parametrically in the model because the model can always be written in a form like (7).

[6]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 inflation 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.

TABLE 1
MONTHLY PORTFOLIO REGRESSION RESULTS FOR

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

(R_p = rate of return on equally weighted portfolio of 75 Electric Utilities; t-statistics in parentheses)

	Const.	b_1	b_2	$\bar{R}^2$
1960-1979	.0040 (2.3)	.62 (14.6)	-70.9 (5.1)	.59
1960-1969	.0036 (18)	.61 (11.2)	-105 (4.4)	.61
1970-1979	.0047 (17)	.63 (9.8)	- 62 (3.3)	.58
1960-1964	.0063 (2.9)	.70 (12.0)	-108 (2.6)	.72
1965-1969	.0003 (.07)	.53 (5.6)	-100 (3.1)	.52
1970-1974	.0013 (.3)	.55 (6.5)	- 73 (3.0)	.58
1975-1979	.0052 (1.2)	.77 (8.4)	- 54 (1.9)	.59

TABLE 2

PARAMETER ESTIMATES OF a_o and a_2

$$R_{it} = a_o - a_o\beta_{i1} + a_2\beta_{i2} + \beta_{i1}R_{mt} + \beta_{i2}\Delta I + e_{it}$$

(t-statistics in parentheses)

	A		B		C	
	a_o	a_2	a_o	a_2	a_o	a_2
1960-79	.105 (.2)	-.599* (2.1)	.506 (1.5)	-.332 (1.8)	-.028 (.07)	-.492* (2.4)
1960-69	.228 (.8)	-.235* (2.0)	.259 (.7)	-.203 (1.9)	-.08 (.21)	-.277 (1.7)
1970-79	.060 (.12)	-.795* (2.2)	.348 (.9)	-.357 (1.5)	-.169 (.3)	-.702* (2.6)
1960-64	1.27* (3.7)	-.0571 (.9)	1.138* (3.4)	-.081 (1.2)	NC	NC
1965-69	-.293 (1.4)	-.343* (2.1)	-.31 (.9)	-.265* (2.3)	-.592 (1.8)	-.535* (3.12)
1970-74	.295 (.6)	-.349 (1.3)	-.203 (.5)	-.489* (2.3)	-.503 (.9)	-.707* (2.3)
1975-79	-.519 (1.4)	-.784* (3.3)	-.654 (1.4)	-.761* (2.7)	NC	NC

*Significant at the 95% confidence level.

Note: Estimates of a_2 have been multiplied by 10^4. Estimates of a_o have been multiplied by 10^2. NC denotes non-convergence of the algorithm.

TABLE 3

IMPLIED ANNUAL PREMIA FROM ΔI

(in percent per year)*

	A	B	C	Average
1960-79	5.2	2.7	4.3	4.07
1960-69	3.3	2.3	3.5	3.03
1970-79	5.8	2.6	5.4	4.6
1960-64	0.9	1.2	NC	1.05
1965-69	4.5	2.6	6.9	4.7
1970-74	3.3	4.1	6.2	4.53
1975-79	4.5	5.0	NC	4.75

*For equally weighted portfolios of firms in groups A-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 non-convergence of the algorithm.

TABLE 4

$$DFY = \gamma_0 + \gamma_1 R_M + e$$

	$\hat{\gamma}_1$	t-stat	Standard Error
1960-79	-.11817	-6.4	1.846x10-4
1960-69	-.06407	-3.12	2.054×10^{-4}
1970-79	-.1485	-5.2	2.856×10^{-4}
1960-64	-.0102	-.55	1.855×10^{-4}
1965-69	-.11496	-3.3	3.484×10^{-4}
1970-74	-.1541	-3.83	4.024×10^{-4}
1975-79	-.1577	-3.75	3.087×10^{-4}

*DFY = 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 $\hat{\gamma}$ are multiplied by 100.

TABLE 5

ESTIMATES OF a_2 FOR ORTHOGONALIZED VERSION OF MODEL

(t-statistics in parentheses)

	A	B	C
1960-79	-.329 (1.4)	-.110 (.6)	-.206 (1.2)
1960-69	-.024 (.2)	.005 (.05)	-.047 (.3)
1970-79	-.485 (1.6)	-.09 (.4)	-.358 (1.6)
1960-64	-.0818 (1.32)	-.105 (1.5)	NC
1965-69	.133 (.9)	.212 (1.8)	-.024 (.17)
1970-74	-.280 (1.2)	-.344 (1.6)	-.515 (1.9)
1975-79	-.168 (.9)	-.124 (.5)	-.39 (2.1)

IMPLIED ANNUAL PREMIA FOR ORTHGONALIZED VERSION OF ΔI

(in percent per year)

	A	B	C	Average
1960-79	2.9	0.9	1.8	1.87
1960-69	0.3	-.056	0.6	.3
1970-79	3.6	0.7	2.8	2.4
1960-64	1.3	1.5	NC	1.4
1965-69	-1.7	-2.1	.31	-1.16
1970-74	2.7	2.9	4.5	3.4
1975-79	1.0	0.8	2.5	1.4

*Calculated as in Table 3.
NC denotes non-convergence of the algorithm.

TABLE 6

PREMIA IMPLIED BY THE CAPM

$$P\Delta'/\hat{\gamma}_1 = (\hat{ER}_M - ER_Z)$$

(figures are in percentage per year)

	A	B	C	Average
1960-79	60.79/ -	33.71/ -	49.9/ -	48.14/-
1960-69	44.02/23.87	38.06/20.64	51.88/28.14	44.65/24.22
1970-79	64.21/80.65	28.82/36.20	56.72/71.24	49.92/62.70
1960-64	67.18/5.80	95.6 /8.25	NC	81.4/7.0
1965-69	35.8 /34.8	27.77/26.94	55.8 /54.3	39.8/38.7
1970-74	27.16/35.41	38.11/49.68	55.06/71.79	40.11/52.29
1975-79	59.63/79.56	57.9 /77.3	NC	58.8/78.4

*Top figure uses $\hat{\gamma}$ calculated in the same period as PΔ'. Bottom figure uses $\hat{\gamma}$ calculated from 1960-1979.

NC denotes non-convergence of the algorithm.

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