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The Term Structure of Inflationary Expectations and Market Efficiency

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FISHER'S [15] ORIGINAL FORMULATION OF an equilibrium relationship in which a nominal interest rate is equal to a real rate of interest plus the expected rate of inflation has stimulated considerable interest and research. Although many empirical studies have utilized a single equation distributed lag specification[1] of past price changes (which might be induced from a rational expectations argument) as a proxy for "expectations," T. F. Cargill and R. A. Meyer [5] have emphasized the necessity of a simultaneous equation structure if such distributed lags were to be employed. An alternative approach using Livingston's direct observations on expectations and allowing for the possibility of incomplete adjustment to expectations was employed by Cargill [2], W. Gibson [16] and D. Pyle [22]. These evolving improvements in conception of the empirical problem, data and econometric approach have led to a particularly interesting paper by K. Lahiri [18] in which S. Turnovsky's [25] expectations formation equations are coupled to the original Fisherian equation and then appropriate simultaneous equation techniques are applied to the resulting econometric relationships.

From a slightly different perspective, E. Fama [13] has emphasized the implication for market efficiency embedded in the interest rate and expected inflation relationship. All analyses to date rest on the assumption that the real rate of interest is constant over time. Thus, as Fama noted, empirical implications for market efficiency tests really constitute a test of the joint hypothesis of efficiency and a constant rate.[2]

Despite the advances in conception of the empirical problem, data improvements, and improvements in econometric technique, the question of variation in the real rate of interest and the possibility of incomplete adjustment to changes in the real rate as well as to expected inflation leaves the extent and stability of the Fisherian relationship far from a settled matter. In addition, a much richer class of hypotheses emerges when Fama's emphasis on the market equilibrium aspects is carried further. No prior work has utilized any type of data to estimate the term structure of expectations which is the key to Fisher's hypothesis when applied across maturities. In addition, the contributions of W. Sharpe [24] and J. Lintner [20] in providing an equilibrium capital asset pricing theory can be

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[1] The use of OLS yields biased and inconsistent estimates.

[2] Numerous researchers have questioned the constancy assumption.

utilized to extend the study of the Fisherian relationship to risky as well as riskless instruments.

In the following discussion, Section I presents the underlying theory and econometric approach. Section II discusses the data employed. Sections III and IV contain the empirical results and discussion for the term structure of riskless and risky yields, respectively. A brief conclusion summarizes the main results.

I. Theoretical Model Formulation

With a view to analyzing both riskless and risky yields, it is particularly useful to begin with the basic capital market equilibrium pricing relationship derived by Sharpe and Lintner. The equilibrium price of any asset can be expressed as

$$V_t = \frac{E[r_t] - \beta\rho\sigma}{r_t^f} \tag{1}$$

where V_t is the market price per unit, r_t is the return at time t, β is a measure of nondiversifiable risk, ρ is the correlation coefficient between the return on a portfolio consisting of all market assets and the return on the asset under consideration, σ is the standard deviation of r_t, and r_t^f is the risk free discount factor. Alternatively, (1) may be expressed as

$$\frac{E[r_t]}{V_t} = r_t^f + \frac{\beta\rho\sigma}{V_t} \tag{2}$$

i.e., as a relationship between an expected nominal return and a risk free return plus a "risk premium." Expression (2) could also be expressed as

$$R_t = r_t^* + \dot{\pi}_t + d_t + l_t \tag{3}$$

The nominal rate, R_t, is expressed as the sum of a real rate, r_t^*, an expected inflation rate, $\dot{\pi}_t$, and a risk premium decomposed into default risk, d_t, and a liquidity premium, l_t. In this form, the traditional econometric expression for the Fisherian relationship is seen as a special case ($d_t = 0$, $l_t = 0$), since most researches implicitly assume that in equilibrium one would find $r_t^f = r_t^* + \dot{\pi}_t$. In addition, the need for explicit recognition of the nonconstancy of the real rate of interest is apparent when time series data is to be employed as the vehicle for isolating the relationship and/or testing market efficiency hypotheses.

Expression (3), however, does not fully reveal the fact that tests across different maturities will not in general be correct if the *same* numerical value for the expected inflation rate is used for all maturities. The total nominal return of an asset in period t may be expressed as

$$R_t = \frac{V_t - V_{t-1}}{V_{t-1}}$$

or

$$V_{t-1} = \frac{V_t}{1 + R_t} \tag{4}$$

If the asset is sold (or matures) at t for \$1, then V_{t-1} depends solely on the one period nominal yield (hence, only a one period inflationary expectation), but recursive substitutions using (4) yields an expression for an asset with k periods to maturity as

$$V_{t-1,k} = \pi_{j=0}^{k} \frac{1}{1 + R_{t+j}} \tag{5}$$

Thus the relevant inflationary expectation in (3) is necessarily a geometric mean of the one period rates from $t - 1$ to $t - 1 + k$. The same argument applies to estimates of r_r^*, d_t, and l_t for maturities beyond one year. In essence, an estimate of the term structure of expectations is needed.

Direct observations on inflationary expectations are available in the Livingston data[3] for 6 and 12 months ahead. With such a paucity of data three approaches might be considered for estimating a term structure of expectations.

First, the most direct approach is to posit a smooth relationship among expectations of different maturities and analytically fit a curve to the sample points. The frequently observed yield curve shape

$$\dot{\pi}_k = \alpha_0 + \alpha_1 e^{-\alpha_2 k} \tag{6}$$

and its special case

$$\dot{\pi}_k = \alpha_0 (1 - e^{-\alpha_2 k}) \tag{7}$$

provide a useful starting point. Table I reports nonlinear least squares estimates for equation (7) over several periods from 1954 to 1975.

The most striking impression from these results is the almost total absence of stability in expectations from quarter to quarter.[4] Only when observations were grouped on a purely empirical basis of relative stability did R^2's even begin to edge above zero—and even this improvement was judged too slight to form a reliable basis for estimating the unobserved points on the term structure of expectations curve.

As a second approach, since we seek an estimate of $\dot{\pi}_{t,k}$, the k period mean

[3] The Livingston data is a biannual series of expected rates of inflation based on a survey of business, government, labor, and academic economists on their expectations of the future value of the CPI 6 and 12 months ahead. The survey has been conducted by Joseph Livingston, a finnancial columnist, on a biannual basis since 1946. The data represent a widely used and useful measure of inflationary expectations based on survey data; however, the reader should be aware of the inherent problems of using the Livingston data. Gibson [16] has provided some discussion of the problems of using survey data to represent broad based inflationary expectations, but J. Carlson [9] has provided the most useful discussion of the problems associated with using the Livingston data. Early applications of the data contain errors in the interpretation of the survey results published by Livingston. Carlson has detailed these issues and developed a revised series of expected inflation that forms the data base used in this paper. Despite the inherent difficulties of using the Livingston data, they represent reasonably good direct estimates of inflationary expectations that can be used to generate a term structure of inflationary expectations.

[4] The estimates reported in table 1 are based on a quarterly version of the Livingston data. As published, these data are available twice each year, for June and December. Quarterly values are based on a linear interpolation of the biannual measures of expected inflation. Quarterly data were used to obtain a reasonable number of degrees of freedom to estimate expression (7) over a variety of short periods of time. The quarterly values were disregarded at this point and the subsequent analysis is based solely on the biannual measures of expected inflation as published by Carlson.

Table I

Nonlinear Least Squares Estimates of Yield Curve

$\dot{\pi}_k = \alpha_0(1 - e^{-\alpha_2 k})$

Time Period	$\hat{\alpha}_0$	$\hat{\alpha}_2$	R^2
Whole Period			
1954I - 1975IV	2.23	6.01	0.001
Large Subperiods Partitionary			
1954I - 1959IV	0.55	3.87	0.005
1960I - 1969IV	1.74	5.06	0.005
1970I - 1975IV	4.74	7.57	0.001
Smaller Subperiod Partitionary			
1954I - 1956II	0.18	148.03	0.000
1956III - 1957II	1.14	152.62	0.000
1957III - 1960IV	0.94	1.35	0.159
1961I - 1965IV	1.11	5.99	0.012
1966I - 1973I	3.15	4.32	0.038
1973II - 1975IV	6.00	52.53	0.000

expectation which satisfies the consistency relation

$$(1 + \dot{\pi}_{t,k})^k = \pi_{j=1}^{k} (1 + \dot{\pi}_{t+j,1}), \tag{8}$$

one could estimate an expectations formation equation such as Turnovsky analyzed but with coefficients which vary by maturity, i.e.,

$$\dot{\pi}_{t,k} = \beta_0(k) + \beta_1(k)\dot{p}_t + \beta_2(k)(\dot{p}_t - \dot{p}_{t-1}) + \cdots \tag{9}$$

where $\dot{p}_t$ is the observed percentage change in prices at time t. The coefficients in (9) can be approximated by Taylor series and truncated to permit estimation.

A third approach is suggested by the apparently sharp divergences between 1950s and 1960s samples (e.g., Gibson [16] and Cargill and Meyer [6] [7]) and the purely statistical aberrations introduced by trends in economic data. Although prior to the 1960s no consistent trend in inflation existed, this is not true in the 1960s. There is some suspicion the marked increases in explanatory power for later periods noted in Fisherian studies and in expectations formation model tests may be due to the presence of trends rather than an intrinsic improvement in the quality of the description by various models presented [17]. Such an observation mitigates against estimation based on relationships like (8) and (9). Since it appears that quarter to quarter expectation changes contain information relevant to subsequent tests and the more traditional route of estimation of expectation curve parameters yields insufficient information (table 1), direct analytic fits of various functional forms provide a reasonable method for extracting the implicit information. This procedure maintains the independence of quarter to quarter expectation estimates which would not be true for forecasts based on a regression procedure. Using the two observations at each point in time successive experiments included the following:

$$\dot{\pi}_k = \beta_0 + \beta_1 k \tag{10}$$

$$\dot{\pi}_k = \alpha_0(1 - e^{-\alpha_2 k}) \tag{11}$$

$$\dot{\pi}_k = \alpha_1 k^{\alpha_2} \tag{12}$$

$$\dot{\pi}_k = \alpha_0 + \alpha_1 k + \alpha_2 k^2 \tag{13}$$

where the quadratic emerges from a truncated Taylor series approximation of $e^{-\alpha_2 k}$ in (11). While more observations naturally would admit a richer set of candidates, equation (12) yielded such well-behaved estimates that these were used. At each point in time separate estimates were made for α_1 and α_2. These were then used to estimate unobserved expectations at 2, 3, 4, 5, 10, and 20 years.[5] Combined with the estimates for 6 and 12 month forecast horizons, the 2, 3, 4, 5, 10, and 20 year estimates yield 8 points on the term structure of expected inflation at each point in time.[6]

Estimation now requires a general expression with the provision for market efficiency tests with respect to both a real rate and inflationary expectations.

$$R_t = \beta_0 + \beta_1 r_t^* + \beta_2 \dot{\pi}_t + u_t \tag{14}$$

Fama and Schwert [14] have suggested that observed market rates may also pick up a contemporaneous effect from the difference between the expected inflation and the actual inflation rate, hence (14) is augmented for this potential effect.[7]

$$R_t = \beta_0 + \beta_1 r_t^* + \beta_2 \dot{\pi}_t + \beta_3(\dot{p}_t - \dot{\pi}_t) + u_t \tag{15}$$

[5] The small number of degrees of freedom available to estimate the tail end of the expectations curve (beyond one year maturity) is a matter of concern; however, three considerations lead us to regard this as a reasonable procedure. First, there is a considerable amount of data available for estimating the term structure if one is willing to use a pooled-time series sample. Unfortunately, this would imply homogeneity in the way expectations are formed at each point in time when the majority of evidence indicates that there is considerable heterogeneity in the underlying expectation formation process. Thus it seems more reasonable to base the estimation on the data available at each point in time. Second, the lack of data for inflation expectations beyond one year certainly leads to a high variance of the estimates for two years and beyond; that is, there is a wide confidence interval around the tail end of the term structure of expected inflation. More direct observations would have been desirable, but would this significantly improve the estimates? Additional direct observations would not necessarily reduce the variance of the estimates since the direct observations themselves would be subject to considerable statistical variation, especially for expectations beyond a few years. Thus, whether we have additional data or use the procedure outlined in this paper, there will be considerable statistical variation in the tail end of the term structure curve. Third, additional observations would have been desirable in any event; however, they are simply not available. There is a series of 18 month expected inflation rates based on the Livingston survey data, but these are available only once a year and Carlson's estimates end in 1971.

[6] Previous work [10] and [18] has suggested that the Livingston data may contain errors of measurement which produce a downward bias in the estimated coefficient on the expected inflation term in Fisher-type equations for the nominal interest rate. We have not employed a correction for measurement error since regressions of expected inflation on lagged actual rates of inflation (treated as instruments) did not yield reliable predicted values that could be used in a second stage estimation procedure. The periods 1960–1969, 1970–1975, and 1960–1975, yielded adjusted R^2 values of .73, .57, and .81 for a three-period distributed lag equation of actual changes in the GNP deflator to predict the 6 months expected inflation rate.

[7] The data used in this study permit estimates of expression (15) for maturities one year or less since the only available measure of unexpected inflation from the Livingston data base is for 6 months and 12 months. Estimates for the unexpected inflation variable were neither consistent in sign nor always significant. Apart from a lagged adjustment phenomena, only expected inflation should be relevant to yields which are also determined ex ante, as opposed to unexpected inflation which is only

where β_0 yields an estimate of the combined default and liquidity components.[8]

It is inappropriate to interpret the constant term in a regression of R_t on $\dot{\pi}_t$ as an estimate of "the real rate of interest" since the estimated parameter is really equal to $d_t + l_t + \beta_1 r_t^*$. Even if the real rate is constant and the liquidity premium is constant and the default premium is constant, such an interpretation requires the implicit assumption that the nominal rate adjusts completely to all three components for each sample observation. Given the reasonable possiblity of incomplete adjustment to expected inflation, the asymmetry of assumptions appears somewhat incongruous.

Since (14) will be estimated for both riskless and risky term structures, the difference between the two $\hat{\beta}_0$'s will yield an estimate of the default risk for risky bonds plus any differential in liquidity premiums between riskless and risky instruments. Previous work by J. McCulloch [21] suggests the liquidity premiums on riskless instruments are quite small, thus barring an inexplicably different result for risky bonds, the principal component of the difference between the $\hat{\beta}_0$'s will be the default variation across the term structure.

II. Time Interval and Data

In order to investigate the issues raised in the previous section, expression (14) is estimated for a variety of riskless and risky yields using estimates of the term structure of expected inflation. The constancy of the $(\beta_0 + \beta_1 r^*)$ term is tested by comparing constant coefficient estimates of $(\beta_0 + \beta_1 r^*)$ with estimates that allow $(\beta_0 + \beta_1 r^*)$ to vary along a polynominal in time. This type of stability test has been used by the authors in previous work [8]. In addition, expression (14) is estimated in an expanded form to allow for variables other than exact time patterns which may influence $\beta_0 + \beta_1 r^*$ over time.

Direct observations on expected inflation are available for the period from June 1947 through December 1975 on a biannual basis; however, the estimates to be presented below are not based on the entire period. The bond pegging program initiated by the Federal Reserve System in 1941 and continued until March 1951,

known ex post. M. Levi and J. Makin [19] have also experimented with measures of risk such as the variance of the inflation rate.

While at first glance it would appear useful to estimate a real rate and include it directly at the estimation stage, unfortunately this is not possible with existing data. Equation (3) emphasizes that subtracting a measure of price expectations does not yield an estimate of the real rate. In fact, such a procedure assumes complete adjustment to price expectations (one of the questions to be asked, rather than assumed) and it begs the question of the degree of adjustment to "the" real rate. Even if the coefficient of adjustment to price expectations were known with certainty, a nominal rate minus the adjustment coefficient times the expectation term leaves liquidity and default premia and an adjustment coefficient times the real rate effects scrambled.

[8] Viewed as a set of relations drawn from a simultaneous equation system, interpretation of estimates (and by implication tests of market efficiency) poses severe difficulties for Fisherian phenomena. As an example, consider a set of three equations, one for price expectations $\dot{\pi}$, one for the Fisherian relationship, R_t, and one for a real rate, r_t^*. Since direct observations on r_t^* are not available, estimation of a relationship for R_t implicitly assumes a partial reduced form in which the relationship for r_t^* is substituted into the one for R_t. This leaves coefficients in the R_t relationship theoretically unidentifiable without further structural information. Individual coefficients may be positive, negative, and greater or less than unity without corresponding inference for market efficiency.

requires that we eliminate observations during the pegging period. While the pegging program ended officially in March 1951, it has been suggested that data prior to 1954 are sufficiently influenced by the pegging program to warrant exclusion from the sample.

Estimates are based on the following periods: December 1954–December 1959; June 1960–December 1969; June 1970–December 1975; and June 1960–December 1975. The variety of time periods are employed to highlight significant structural change over the period from 1954 through 1975 found in previous work.

Risky returns are represented by the 4–6 month prime commercial paper rate and the Aaa corporate bond rate [1]. Riskless returns are represented by the Salomon Brothers [23] estimates of yields for government securities at three month, 1 year, 2 year, 3 year, 4 year, 5 year, 10 year, and 20 year maturities. The Livingston data are biannual observations and in order to use observed interest rates as close to the point at which the forecast was made lead to the use of first of June and December observations for the Salomon data (starting in 1959) and average values for May and November for the commercial paper and corporate bond rate.

The 6 month expected inflation rate is related to the 4–6 commercial paper rate and the 3-month Treasury bill rate. The Aaa bond rate represents a variety of long-term maturities and given the lack of an exact maturity represented by the Aaa rate it is related to both the 10 and 20 year expected inflation rates. The form of the data make it impossible to match the maturities of the interest rate and expected inflation exactly for these three interest rates just mentioned; however, the maturities should be close enough to approximate the theoretical relationship between expected inflation and nominal interest rates. In the case of the Salomon data for 1, 2, 3, 4, 5, 10, and 20 year interest rate maturities, exact corresponding maturity of expected inflation can be used thus providing a close matching of dependent and explanatory variables.

III. Simple Relationships Between Nominal Yields and Expected Inflation

Table II presents estimates of expression (14) for three time periods;[9] December 1954–December 1959, June 1960–December 1969, and June 1970–December 1975. There are several implications worth noting. First, the coefficient for the expected inflation term is almost always positive and significant at conventional test levels. Second, coefficients decline with increases in the maturity of the nominal interest rate and expected rate of inflation. Third, many of the coefficients on the expected

[9] The estimates presented in these tables are based on the Cochrane-Orcutt iterative procedure to adjust for first order serial correlation in the residual process. While the Cochrane-Orcutt procedure theoretically improves the asymptotic efficiency of the regression results, the resulting Durbin-Watson statistics indicated that serial correlation still remained a problem in some regressions. In several cases, the Cochrane-Orcutt procedure did not achieve a stable solution. The procedure would iterate to a value very close to -1 or $+1$. In those cases, the rho value was set to $-.99$ or $+.99$. While not a completely satisfactory solution, it was only necessary in a few of the reported regressions. Specifically, those estimates with very large constant terms are the ones where this procedure was followed.

The estimates for the period June 1960–December 1975 are not reported to conserve space; however, they conform to the implications drawn from the results for the three other periods.

Table II

Estimates of $R_t = (\beta_0 + \beta_1 r^*) + \beta_2 \dot{\pi}_{t,k}$ over the Periods 1954-1959, 1960-1969 and 1970-1975[a]

	1954-1959					1960-1969					1970-1975				
	Coefficient					Coefficient					Coefficient				
R_t	$\beta_0 + \beta_k r^*$	β_2	k	R^2	D.W.	$\beta_0 + \beta_1 r^*$	β_2	k	R^2	D.W.	$\beta_0 + \beta_1 r^*$	β_2	k	R^2	D.W.
4-6 month	2.96*	.58	6 mo.	.00	2.12	2.04*	1.53*	6 mo.	.74	1.77	1.22	1.04*	6 mo.	.49	1.23
Commercial	(5.24)	(.91)				(4.64)	(7.14)				(.69)	(3.25)			
Paper Rate															
Aaa Bond	3.45*	.75*	10 yr.	.76	1.86	2.58*	1.24*	10 yr.	.92	1.66	5.25*	.62*	10 yr.	.37	2.43
Rate	(2.32)	(5.45)				(12.17)	(14.62)				(4.17)	(2.63)			
Aaa Bond	3.57*	.57*	20 yr.	.67	1.85	2.56*	1.18*	20 yr.	.88	1.59	4.67*	.72*	20 yr.	.38	2.45
Rate	(22.07)	(4.41)				(9.67)	(11.78)				(3.24)	(2.68)			
3 month	1.90*	1.33	6 mo.	.24	1.79	1.69*	1.46*	6 mo.	.75	1.84	2.02	.74*	6 mo.	.45	1.26
Treasury	(3.56)	(1.97)				(4.18)	(7.36)				(1.56)	(3.05)			
Bill Rate															
1 year	1.27*	2.40*	1 yr.	.78	2.12	1.72*	1.46*	1 yr.	.80	1.72	2.38	.77*	1 yr.	.51	1.28
Treasury	(4.13)	(5.67)				(4.65)	(8.50)				(2.00)	(3.34)			
Bond Rate															
2 year	1.87*	1.70*	2 yr.	.68	1.62	1.98*	1.29*	2 yr.	.81	1.56	2.66*	.74*	2 yr.	.76	1.80
Treasury	(6.18)	(4.52)				(6.05)	(8.82)				(4.02)	(5.71)			
Bond Rate															
3 year	2.30*	1.26*	3 yr.	.64	1.64	2.20*	1.17*	3 yr.	.84	1.50	2.88*	.72*	3 yr.	.82	2.06
Treasury	(8.63)	(4.11)				(7.94)	(9.67)				(5.27)	(6.73)			
Bond Rate															
4 year	2.53*	1.01*	4 yr.	.55	1.68	2.36*	1.08*	4 yr.	.85	1.47	3.08*	.70*	4 yr.	.82	2.29
Treasury	(9.35)	(3.46)				(9.60)	(10.30)				(5.80)	(6.78)			
Bond Rate															
5 year	2.67*	.85*	5 yr.	.50	1.70	2.46*	1.02*	5 yr.	.86	1.46	3.15*	.70*	5 yr.	.84	2.45
Treasury	(10.34)	(3.18)				(10.82)	(10.66)				(6.30)	(7.23)			
Bond Rate															
10 year	2.90*	.58*	10 yr.	.60	1.75	2.90*	.83*	10 yr.	.78	1.32	3.96*	.54*	10 yr.	.60	1.88
Treasury	(17.33)	(3.84)				(10.69)	(8.08)				(5.64)	(4.03)			
Bond Rate															
20 year	3.00*	.44*	20 yr.	.71	1.82	16.69*	.11	20 yr.	.01	1.47	.76	1.17*	20 yr.	.78	2.42
Treasury	(26.08)	(4.79)				(2.56)	(1.08)				(.75)	(6.12)			
Bond Rate															

[a] R^2 is adjusted for degrees of freedom and the absolute value of the "t-statistic" is enclosed by parentheses. An asterisk indicates that the coefficient is significant at the .05 level. Durbin-Watson and R^2 values are based on transformed residuals.

inflation term are either close to or significantly above unity. Complete adjustment of the nominal rate to expected inflation requires a coefficient above unity to take into account the presence of a progressive income tax system where interest income in subject to taxation [12]. Fourth, the constant terms which measure the combined influence of default and other factors are generally larger for the risky yields than for the riskless government security yields. Fifth, the size and statistical significance of the coefficient for the expectation term is directly related to the maturity of the nominal yield. The coefficients decline as the length of the term structure increases. Sixth, there is a considerable variation in the results across the three time periods. An interesting and important result is seen by examining the estimates for the 1950s which shows some evidence of a relationship between interest rates and expected inflation. Previous work has concluded that the relationship is very weak or not statistically significant; however, this evidence has been based on including several of the earlier years of the 1950s and not comparing the same maturity lengths of the interest rate and expected inflation for yields greater than one year. It would appear that data before 1954 should be omitted from the analysis as well as only comparing interest rates and estimates of expected inflation of the appropriate corresponding maturity; however, evidence for the 1950s still presents perplexing features. For example, the price expectation coefficients are not significant for the 4–6 month commercial paper rate and the 3-month Treasury bill rate while they are significant for maturities equal to or greater than one year.

As mentioned earlier, the constant term in regressions like expression (14) does not measure the real rate, but rather the combined influence of a real rate, a liquidity premium, and in the case of risky assets, a default premium. Considerable simplicity is obtained by regressing nominal rates on a constant term and price expectation variable; however, the assumed constancy of the term $\beta_0 + \beta_1 r^*$ is a rather weak reed. A direct test of this assumption can be made by assuming that the constant term can be approximated by a polynominal in time. Since the first question deals with stability per se, a time polynomial is sufficient for a myriad of specific causes while choice of a specific variable(s) would provide only a more restricted stability test. Table III presents the F-statistics of the test between the restricted and unrestricted interest rate equations. In a number of cases the constancy of the term $\beta_0 + \beta_1 r^*$ assumed in the estimates of table 2 is rejected at the .05 level. An interesting exception occurs when the 1960s and 1970s data are pooled, when a majority of F-statistics fail to reject the constancy assumption. A number of investigators have expressed the opinion that simple regressions of the nominal rate on expected inflation are too restrictive; however, to our knowledge, the results in table III are the first direct test of the assumption of constancy in the context of a term structure of price expectations.

IV. An Expanded Relationship Between Nominal Yields and Expected Inflation

The results presented in Table III suggest that the basic interest rate model should be expanded to take into account factors that cannot realistically be held

Table III

A Test of the Constancy of the Term $\beta_0 + \beta_1 r^*$ over the period from 1954 through 1975[a]

R_t	F-Statistics			
	1954–1959	1960–1969	1970–1975	1960–1975
4-6 month Commercial Paper Rate	2.65	4.35*	2..00	2.41
Aaa Corporate Bond Rate[b]	2.51	24.51*	15.00*	1.90
3 month Treasury Bill Rate	3.53	3.29	1.78	1.93
1 year Treasury Rate	1.66	2.58	1.40	1.58
2 year Treasury Rate	4.43	3.40*	2.46	1.64
3 year Treasury Rate	4.22	4.62*	3.55	1.83
4 year Treasury Rate	2.92	6.10*	3.99	2.17
5 year Treasury Rate	2.03	7.77*	4.24	2.44
10 year Treasury Rate	2.48	23.92*	15.97*	6.45*
20 year Treasury Rate	4.12	17.18*	6.87*	6.90*

[a] An asterisk indicates that the F-statistic is significant at the .05 level.
[b] F-statistic based on 20 year maturity for expected inflation.

constant. There are many possible models of the nominal rate of interest, based on liquidity preference, loanable funds, or some combination of the two; however, J. A. Carlson [10] has recently suggested an approach to an expanded model of the Fisherian mechanism that suggests a measurable proxy for the real rate.

The real rate might be regarded as a positive function of the expected return to capital and Carlson agues that the percentage capacity utilization rate for manufacturing (CUM) is a reasonable proxy for this return. The capacity utilization rate is estimated and published by the Board of Governors [26]. Liquidity elements can be incorporated by including the percentage change in the money supply defined as currency plus demand deposits.[10]

Table IV presents estimates of

$$R_t = \alpha_0 + \alpha_1 \dot{\pi}_{t,k} + \alpha_2\ (\%\Delta M1) + \alpha_3\ \mathrm{CUM} \tag{16}$$

for the periods 1954–1959, 1960–1969, and 1970–1975. The importance of the price expectation variable is still apparent and behavior of this varible across price expectations and yield maturities discussed above is also present in the expanded model results.

The capacity index generally has the expected positive sign, but is seldom significant. Only at the short end of the maturity spectrum is the capacity variable positive with t-statistics above 1.5. This is consistent with the argument made by Carlson that the capacity variable may only be a satisfactory measure of short-run expected real returns. The liquidity variable is negative in the majority of regressions and statistically significant in a large number. These results suggest that improvements can be obtained by modeling the liquidity and real interest rate components separately.

[10] The capacity and liquidity variables used in this paper differ from those employed by Carlson. Carlson used deviations of the capacity index from the mean value and a moving average measure of the liquidity variable, while we use the direct estimates of the capacity index and the percentage change in the level of the money supply.

V. The Question of Stability and Adjustment of the Relationship

Recent literature on the relationship between nominal interest rates and expected inflation has emphasized two considerations. First, is the relationship stable over time? A prime objective of the study by Cargill and Meyer [7] was to investigate the stability issue over a long period of time. Other studies have also reached findings of intertemporal instability in the relationship. Second, is the adjustment of the nominal ratè to expected inflation complete in the presence of a progressive income tax structure? Empirical work by Cargill [3], J. Carr, J. Pesando, and L. Smith [11], and Carlson [10] suggest that the adjustment is not complete for the 1960s, though Cargill and Carlson found a larger adjustment of the nominal rate of expected inflation than Carr, et al. We now turn to a consideration of these two issues.

Stability of the Relationship: A detailed study such as presented in [7] is not possible with the type of data and the time interval under consideration in this paper. The biannual nature of the Livingston data effectively rules out monthly and even quarterly analysis and the time interval from 1960 to 1975 limits the number of subperiods that can be reasonably analyzed.

Given the limitations of the data, separate estimates for subperiods from 1954 through 1975 were obtained and compared (tables II and III). The estimates indicate that the coefficient of the expected price variable is generally positive and statistically significant; however, the magnitude of the relationship between nominal rates and expected inflation varies significantly. For example, in table II the estimated price expectation term for the 4-6 month commercial paper rate is 1.53 for the period 1960-1969 while declining to 1.04 for the period 1970-1975. In addition, the R^2 values of the estimated equations are consistently lower for the later period.

Another significant difference between the two periods is illustrated by the stability test of the constancy of the term $\beta_0 + \beta_1 r^*$ via an F-statistic. The constancy of the term $\beta_0 + \beta_1 r^*$ cannot generally be rejected at the .05 level for the pooled period 1960-1975, while it is rejected for the 1960-1969 period.[11]

Complete Versus Incomplete Adjustment: The adjustment mechanism can be most easily examined in the relationship for short term yields. In this case, the use of yields to maturity is a close approximation to holding period yields. In the case of longer term yields, there is, of course, a significant difference between yields to maturity and holding period yields. The adjustment coefficients for the short term yields (summarized in table IV) are significantly greater than unity for the 1960s, thus at least minimally consistent with complete adjustment (depending on marginal tax rates). By including data from the 1970s the adjustment coefficients decline markedly below unity. The adjustment coefficients for the 1950s are often not statistically significant.

VI. Concluding Comments

The major results of this study can be briefly summarized. First, explicit recognition of the term structure of price expectations should be considered in the

[11] The pooled period of 1960-1975 combines periods of such diverse inflationary behavior that any results for this period should be regarded with caution.

Table IV

Estimates of $R_t = \alpha_0 + \alpha_1 \dot{\pi}_{t,k} + \alpha_2\ (\%\Delta M1) + \alpha_3$ CUM over the Periods 1954–1959, 1960–1969 and 1970–1975[a]

	1954–1959						1960–1969						1970–1975					
	Coefficients						Coefficients						Coefficients					
R_t	α_0	α_1	α_2	α_3	R^2	D.W.	α_0	α_1	α_2	α_3	R^2	D.W.	α_0	α_1	α_2	α_3	R^2	D.W.
4-6 month	−3.20	.02	−.26*	.10	.78	2.39	−3.30	1.47*	−.18*	.07	.80	1.77	−11.75	.90*	−.07	.18*	.71	2.15
Commercial Paper Rate	(1.06)	(.07)	(4.89)	(2.83)			(.90)	(6.98)	(3.49)	(1.67)			(2.24)	(3.62)	(.68)	(2.78)		
Aaa Bond	2.42	.83*	−.09	.01	.84	2.24	3.24	1.28*	−.02	−.78	.91	1.68	11.69*	.60	−.03	−.08*	.72	2.79
Rate (10 year)	(1.09)	(5.85)	(1.77)	(.54)			(1.63)	(11.90)	(.30)	(.32)			(4.47)	(3.27)	(.46)	(2.71)		
Aaa Bond	1.88	.68*	−.10	.02	.79	2.26	2.33	1.17*	.01	.00	.87	1.59	10.95*	.66*	−.05	−.07*	.70	2.72
Rate (20 year)	(.72)	(5.04)	(1.67)	(.71)			(.96)	(9.51)	(.08)	(.09)			(3.69)	(2.89)	(.74)	(2.33)		
3 month	−1.64	.77	−.28*	.06	.69	2.01	−3.60	1.41*	−.16*	.07	.80	1.66	−8.44*	.63*	−.10	.14*	.75	2.01
Treasury Bill Rate	(.41)	(1.86)	(3.84)	(1.20)			(1.03)	(7.08)	(3.19)	(1.70)			(2.26)	(3.59)	(1.22)	(3.21)		
1 year	−3.91	1.07	−.24*	.09	.82	2.23	−2.60	1.50*	−.15*	.06	.83	1.52	−7.68*	.75*	−.05	.13*	.80	2.49
Treasury Bond Rate	(1.09)	(1.93)	(3.61)	(1.95)			(.85)	(8.24)	(3.55)	(1.58)			(2.45)	(4.52)	(.57)	(3.47)		
2 year	−3.52*	1.88*	−.06	.06*	.95	1.37	−2.33	1.32*	−.14*	.06	.82	1.54	−.78	.71*	−.04	−.05	.78	2.29
Treasury Bond Rate	(2.77)	(12.39)	(1.51)	(4.41)			(.83)	(8.10)	(2.77)	(1.66)			(.30)	(4.66)	(.49)	(1.56)		
3 year	−3.50*	1.55*	−.04	.07*	.94	2.14	−1.51	1.20*	−.12*	.05	.84	1.54	1.11	.69*	−.04	.03	.80	2.31
Treasury Bond Rate	(2.70)	(12.09)	(1.07)	(4.57)			(.62)	(8.66)	(2.28)	(1.63)			(.50)	(4.90)	(.48)	(1.04)		
4 year	−3.95*	1.36*	−.03	.08*	.92	2.23	−.87	1.11*	−.10*	.04	.85	1.51	2.39	.65*	−.05	.02	.79	2.36
Treasury Bond Rate	(2.57)	(10.08)	(.62)	(4.26)			(.40)	(9.01)	(1.98)	(1.56)			(1.13)	(4.63)	(.60)	(.61)		
5 year	−3.91*	1.20*	−.01	.08*	.89	2.21	−.38	1.04*	−.09	.04	.85	1.49	3.66	.63*	−.06	.00	.82	2.41
Treasury Bond Rate	(2.27)	(8.72)	(.29)	(3.86)			(.18)	(9.04)	(1.73)	(1.45)			(1.93)	(4.76)	(.80)	(.10)		
10 year	−1.56	.79*	−.00	.05*	.82	2.28	19.87*	.12	−.08*	−.02	.21	1.52	5.04*	.46*	−.06	−.00	.55	1.89
Treasury Bond Rate	(.78)	(6.27)	(.08)	(2.23)			(2.39)	(.84)	(2.57)	(.05)			(2.13)	(2.67)	(.79)	(.17)		
20 year	−.39	.57*	.02	.04	.83	2.42	17.51*	.11	−.05	.01	.12	1.48	1.91	.92*	−.18*	.01	.86	2.14
Treasury Bond Rate	(.22)	(6.16)	(.53)	(1.86)			(2.48)	(1.20)	(1.95)	(.02)			(.78)	(4.52)	(2.10)	(.63)		

[a] R^2 is adjusted for degrees of freedom and the absolute value of the "t-statistic" is enclosed by parentheses. An asterisk indicates that the coefficient is

Table V

Adjustment Coefficient for Yields Equal to or Less than One Year[a]

R_t	1954–1959		1960–1969		1970–1975		1960–1975	
	$\hat{\alpha}_1$	$\hat{\beta}_2$	$\hat{\alpha}_1$	$\hat{\beta}_2$	$\hat{\alpha}_1$	$\hat{\beta}_2$	$\hat{\alpha}_1$	$\hat{\beta}_2$
4–6 month Commerical Paper Rate	[b]	[b]	1.47*	1.53*	.90	1.04	.80	.86
3 month Treasury Bill Rate	[b]	[b]	1.41*	1.46*	.63	.74	.59	.68
1 year Treasury Bond Rate	[b]	2.40*	1.50*	1.46*	.75	.77	.78	.72

[a] The estimates $\hat{\alpha}_1$'s come from equation (16) while the $\hat{\beta}_2$'s come from equation (14). An asterisk indicates that the adjustment coefficient is significantly greater than unity at the .05 level based on a one-tailed t-test.

[b] Adjustment coefficient not statistically significant at the .05 level.

relationship between nominal yields and expected inflation. Estimates of the term structure of price expectations are significantly related to nominal yields and provide results different from those in which only a single "price expectations" is used for all maturities. Second, simple regressions of nominal yields on expected inflation measures are subject to specification error because liquidity default factors and the real rate cannot be held constant; certainly it seems implausible to do so over the decade of the 1960s. An expanded model of the nominal yield is required to estimate adequately the magnitude of the relationship between expected inflation and nominal interest rates. Third, evidence of instability and/or sample sensitivity in the relationship is apparent. Many of the results obtained for one period fail to hold for estimates obtained for other periods. Fourth, evidence of complete or incomplete adjustment (hence market efficiency or inefficiency) appears to be heavily influenced by the particular time period of estimation. Complete adjustment appears to have been achieved only in the 1960s at the short end of the maturity structure.

A final comment on the framework of the study is appropriate. Both the work reported here and by prior researchers is based on equilibrium model interpretations of sample observations. It is natural to ask if despite the fluidity of financial markets a disequilibrium/partial adjustment formulation may yield superior results. To our knowledge no prior work has employed lagged adjustment models, yet the presence of marked serial correlation in equilibrium based formulations makes it seem potentially fruitful to estimate a lagged dependent variable version of (14). Durbin's h-statistic was calculated and found to be significant in only one period for OLS estimates; however, despite the weak h-statistic evidence, equation (14) was re-estimated with a lagged dependent variable via Sargan's instrumental variables technique. Rarely is the coefficient on the lagged dependent variable significant and the coefficient on the price expectations term varies so widely that we feel the equilibrium assumption more empirically robust (except for the Aaa bond rate during the 1960s and 1970s). There is little empirical support for a disequilibrium approach based on a partial adjustment model.

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