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New Hope for the Expectations Hypothesis of the Term Structure of Interest Rates

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ABSTRACT

Survey data on interest rate expectations permit separate testing of the two alternative hypotheses in traditional term structure tests: that the expectations hypothesis fails, and that expected future interest rates are ex post inefficient forecasts. We find that the source of the spread's poor predictions of future interest rates varies with maturity. At short maturities the expectations hypothesis fails. At long maturities, however, changes in the yield curve reflect changes in expected future rates one-for-one, an implication of the expectations hypothesis. This result confirms earlier findings that long rates underreact to short rates, but now it cannot be attributed to term premia.

IF THE ATTRACTIVENESS OF an economic hypothesis is measured by the number of papers which statistically reject it, the expectations theory of the term structure is a knockout. Most tests beginning with Macaulay (1938) find no evidence supporting the expectations hypothesis.[1] Many cannot even reject statistically the alternative hypothesis that the spread between long and short rates contains no information about future interest-rate changes.[2] To make matters worse, in U.S. postwar data, future long rates tend to rise when short rates are above long rates. Since the expectations hypothesis would predict that long rates tend to fall, the theory often does worse than even the näive model that future interest-rate changes are always zero.

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[1] Among the many studies which reject the expectations theory in U.S. postwar data are Campbell and Shiller (1984), Fama (1984a,b), Fama and Bliss (1987), Mankiw (1986), Mankiw and Summers (1984), Mankiw and Miron (1986), Shiller (1979), and Shiller, Campbell, and Schoenholtz (1983). See Shiller (1987) for a thorough survey of this literature.

[2] Several papers, however, report more supportive evidence. Fama (1984b) finds the spread has some positive preditive power for future short-term (one-month) interest-rate changes. Mankiw and Miron (1986) also discover evidence of the spread's predictive power, but only as recently as 1890–1915. Campbell and Shiller (1984) and Fama and Bliss (1987) find that medium- and long-term spreads have some positive predictive power for short-rate changes farther into the future. Nevertheless, all of these papers statistically reject restrictions imposed by the expectations hypothesis. Shiller (1981) presents the strongest evidence in support of the expectations hypothesis. He finds not only that the spread has statistically significant predictive power for excess returns on five-year bonds, but also that his data cannot reject the expectations theory.

Naturally, this literature does not test the expectations hypothesis in isolation. It examines instead a joint hypothesis: that investors' expectations conform with the expectations theory, and that those expectations are rational in the sense of Muth. Since the data cannot tell us how each hypothesis would fare individually, authors often attribute the failure of the joint hypothesis in accordance with their priors. Most authors argue that their rejections are a consequence of time-varying term premia. However, others suggest that the expectations hypothesis may be true and that over-or underreaction of expected future rates to short-rate changes is responsible for the results.[3]

By restricting ourselves to existing methods, there is simply no way of choosing between these alternative views. In this paper, we extract new information from surveys of interest-rate expectations to help resolve these basic issues. Survey data give us a unique opportunity to decompose the spread's biased predictions into a component attributable to expectational errors and a component attributable to term premia. The results of this decomposition indicate a striking difference in the importance of these competing explanations at opposite ends of the maturity spectrum.

This paper is structured as follows. Section I describes the data and our treatment of them. Section II reviews the linearized model of the term structure. We perform in Section III the standard test of whether the spread is an unbiased predictor of future interest-rate changes. Section IV separates the spread's bias into a component attributable to a failure of the expectations theory and a component attributable to systematic expectational errors. Section V then uses the surveys to test the first part of the joint hypothesis—that the survey expectations themselves conform to the expectations theory. Section VI tests for the significance of the second component—expectational errors. In Section VII, we present some simple but revealing statistics from the surveys to clarify the role of term premia in the pricing of bills and bonds. Section VIII concludes.

I. The Data

The interest-rate expectations come from a survey conducted by the *Goldsmith-Nagan Bond and Money Market Letter*, now published in the investor newsletter, *Reporting on Governments*.[4] At the end of each quarter from mid-1969 to the end of 1986, Goldsmith-Nagan (GN) surveyed financial-market participants on their expectations of interest rates on three-month Treasury bills, three-month Eurodollar deposits, twelve-month Treasury bills, the Bond Buyer index,[5] and the

[3] Shiller (1981), Shiller, Campbell, and Schoenholtz (1983), and Mankiw and Summers (1984) assume that the expectations hypothesis holds and interpret the spread's inability to predict future interest-rate changes as evidence that long rates underreact to short-rate changes. Under the expectations hypothesis, the long rate is a weighted average of expected future short rates. Thus, the underreaction of long rates can also be thought of as the underreaction of expected future short rates to current short-rate changes.

[4] Friedman (1979, 1980) used some of the data from these surveys.

[5] This is an index of twenty general obligation issues with twenty-year maturities. The index is designed to reflect the current yield-to-maturity on new issues.

thirty-year mortgate rate.[6] Each respondent is asked for his or her expectation of the level of these rates in three and six months. GN reports the median response.[7] The duration and consistency of this set of surveys—currently, seventy contiguous quarters of data for a variety of instruments—is highly unusual.

We wish to be skeptical about the survey data, because they are likely to measure the market's true expectation with error. In the tests below, we follow the literature in speaking of a single "market" expectation. However, different survey respondents have different beliefs, suggesting that if there is a single market expectation the median survey response measures it with error.[8] Thus, we assume the survey measure is equal to the sum of the market's (unobservable) expectation plus random measurement error:[9]

$$\mathbf{s}_{t+j}^{(k-j)} = \mathbf{e}_t^{(j,k)} + \epsilon_{t,j}, \tag{1}$$

where $\mathbf{e}_t^{(j,k)} = (\mathbf{i}_{t+j}^{(k-j)})^{\mathbf{e}} - \mathbf{i}_t^{(j)}$, the market's expectation at time t of the (net) yield to maturity on a $k-j$ period bond issued in period $t+j$ less the yield to maturity at time t on a j period bond, and $E(\epsilon_{t,j} \mid \mathbf{e}_t^{(j,k)}) = 0$. Our identifying assumption is that the measurement error, $\epsilon_{t,j}$, is conditionally independent of $\mathbf{e}_t^{(j,k)}$.[10]

There is a clear analogy between (1) and the standard rational-expectations methodology, in which the "measurement error" is the difference between the ex post future realization and the market's expectation, $(\mathbf{i}_{t+j}^{(k-j)} - \mathbf{i}_t^{(j)}) - \mathbf{e}_t^{(j,k)}$, usually attributed to unforecastable news. But the survey measurement error has important advantages over the ex post prediction error. First, the variance of the measurement error is approximately an order of magnitude smaller than the variance of the ex post prediction error. This reduces the standard errors in regression tests. Second, an increase in the forecast horizon, j, in (1) does not reduce the number of nonoverlapping observations in a given sample, since the survey expectations are recorded at time t. For any given time series of length T, surveys contain more independent information as j grows. Finally, the survey measurement error is immune to the effects of learning, irrationality, and to the presence of peso problems, all of which arise out of differences between the ex ante and sample ex post distributions, and which therefore infect only ex post prediction errors.

[6] Mortage instruments carry an implicit "put" option: mortages that were granted at rates higher than those prevailing are often refinanced. Their risk characteristics are therefore not a function of duration alone.

[7] We matched the survey medians with actual interest rates from Data Resources, Inc., recorded on the last day of each quarter.

[8] When it is possible to aggregate investors' demands into a single representative agent, expected future rates are a complicated weighted average of individual investors' expectations (see Rubinstein (1974)).

[9] Measurement error could also arise because of disparities between the times at which the survey responses and contemporaneous interest rates were recorded. To see whether this source of error affects the results, we experimented with different dating assumptions for contemporaneous interest rates. Specifically, we constructed information sets dated one week and two weeks before the end of the quarter. The results below remain the same under these alternative dating assumptions.

[10] The presence of measurement error implies that survey median need not reflect precisely the market's true expectation at each point in time. We require, however, that the median survey response does not differ *systematically* from the market's expectation.

II. The Model

We adopt the linearized model of the term structure of Shiller, Campbell, and Schoenholtz (1983). The model's approximations to actual forward rates and holding-period returns are useful here for two reasons.[11] First, the linearization resolves any ambiguities which arise when choosing among alternative definitions of the expectations hypothesis. Second, the linearized model allows us to test a single specification of the expectations hypothesis, and yet draw implications for many other specifications.

We denote the forward rate at time t on a $k - j$-period bill, j periods into the future by $\mathbf{f}_t^{(j,k)}$, and the corresponding forward premium by the forward rate less the j-period short rate, $\mathbf{fp}_t^{(j,k)} = \mathbf{f}_t^{(j,k)} - \mathbf{i}_t^{(j)}$. Under the linearized model, the forward premium is proportional to the spread between the long rate and short rate. Shiller, Campbell, and Schoenholtz (1983) show that:

$$\mathbf{fp}_t^{(j,k)} = \left(\frac{D_k}{D_k - D_j}\right)\left(\mathbf{i}_t^{(k)} - \mathbf{i}_t^{(j)}\right), \tag{2}$$

where D_m is Macaulay's (1938) definition of duration for an m-period bond when priced at par, $D_m = \dfrac{1 - (1 + \bar{i})^{-m}}{1 - (1 + \bar{i})^{-1}}$, and $\bar{i}$ is the coupon rate.[12] The term premium on a k-period bond held from period t to period $t + j$ is the difference between the forward premium and expected future interest-rate change:

$$\theta_t^{(j,k)} = \mathbf{fp}_t^{(j,k)} - \mathbf{e}_t^{(j,k)}. \tag{3}$$

Also, let $(\mathbf{h}_t^{(j,k)})^{\mathrm{e}}$ denote the expected holding-period yield obtained from purchasing a k-period bond at time t, holding it for j periods, and then selling it. Under the linearized model, the expected excess holding-period return is proportional to the term premium in (3):

$$\theta_t^{(j,k)} = \left(\frac{D_j}{D_k - D_j}\right)((\mathbf{h}_t^{(j,k)})^{\mathrm{e}} - \mathbf{i}_t^{(j)}). \tag{4}$$

The expectations hypothesis implies that $\theta_t^{(j,k)} = 0$: forward rates are equal to expected future spot rates and expected excess holding-period returns are zero.

III. A Standard Test of the Expectations Hypothesis

There are many ways to test the expectations hypothesis. Here we focus on a single specification. Consider a regression of the subsequent change in the interest rate on the forward premium:

$$\mathbf{i}_{t+j}^{(k-j)} - \mathbf{i}_t^{(j)} = \alpha + \beta\mathbf{fp}_t^{(j,k)} + \eta_{t+j}. \tag{5}$$

[11] The linearization appears to sacrifice little accuracy in comparison with nonlinear models, especially over the relatively short forecast horizons considered in this paper. See Shiller, Campbell, and Schoenholtz (1983) for evidence.

[12] Treasury bills have no coupon, so their duration is just their time to maturity. For longer-term bonds, we assume that the coupon rate is equal to the average return over the sample period.

The expectations hypothesis and rational expectations together imply that the forward premium is an efficient forecast of the future interest-rate change. Thus the null hypothesis in (5) is that $\alpha = 0$, $\beta = 1$, and the residual η_{t+j} is purely random.[13]

Most studies find that β is statistically less than one (once their results are expressed in terms of equation (5)). For shorter maturities β is frequently not statistically different from zero, so that the forward premium is of no help in forecasting future changes in the short rate. Indeed, it is not unusual to find a coefficient less than zero, which implies that interest rates on average move in the direction opposite to that predicted by movements in the slope of the yield curve.[14]

A finding of $\beta = 0$ in (5) has two polar interpretations. At one extreme, it is consistent with a model in which all movements in the slope of the yield curve reflect changes in risk. Under this view, which rejects the expectations hypothesis, expected interest-rate changes are uncorrelated with the slope of the yield curve. Indeed, if interest rates follow a random walk, expected interest rate changes would be zero.[15] At the opposite extreme, $\beta = 0$ can be consistent with a model (e.g., the expectations theory) in which all movements in the slope of the yield curve reflect changes in expected future rates. Under this view, an increase in the spread implies the expectation of future rate increases which on average are not realized ex post. Investors would forecast better if they reduced their expectations of future interest-rate changes toward zero.[16] Note that this latter view need not reject rational expectations. Learning about the interest-rate process, or "peso problems" generated by infrequent events could also be consistent in small samples with a repeated tendency to mispredict interest rates.

A. Results from the Standard Test

Table I presents estimates of (5) for the survey sample period. We report these estimates in part to show that the particularities of our sample do not lead to unusual conclusions about the forward premium's forecasting ability. The time to maturity of each instrument in Table I is $k - j$ (column 1), and the forecast horizon (either three or six months) is j (column 2). For each forecast horizon, instruments are ordered by maturity. Before discussing the parameter estimates, we pause briefly to clarify several econometric issues.

[13] Most studies do not test the precise formulation in (5). Often either the change in the long rate or the realized excess holding return is the dependent variable, and the spread or the forward premium above the long rate is the regressor. Under the linearized model in Section II, however, all of these tests are exact transformations of (5). See the discussion on pages 7 and 8, and in Table 1 of the NBER Working Paper version of this paper for more on the equivalence of different regression tests under the linearized model.

[14] The finding that forward premia contain little information about future spot rate changes is not limited to the term structure. See Fama and French (1986) for similar evidence in commodity markets, and Hodrick and Srivastava (1984) and Froot and Frankel (1989) for the evidence in foreign exchange markets.

[15] See Mankiw and Miron (1986) for a discussion of this view.

[16] See, for example, Shiller (1981), Shiller and Campbell, and Schoenholtz (1983) and Mankiw and Summers (1984).

Table I

Tests of the Term Structure of the U.S.

OLS regressions of

$$\mathbf{i}_{t+j}^{(k-j)} - \mathbf{i}_t^{(j)} = \alpha + \beta\, \mathbf{fp}_t^{(j,k)} + \eta_{t+j}{}^{a}$$

using quarterly data, 1969–1986.

Instrument $(k-j)$	Forecast Horizon (j)	β	t: $\beta = 1$	β_{cc}[b]	DW	$\bar{R}^2$	DF	F-test $\alpha = 0$ $\beta = 1$
3-Month T-Bill	3 Mns	0.0592	−3.616***	4.9737	2.06	0.35	67	14.37***
		(0.2602)	−3.925***	(0.8679)				14.64***
		(0.2397)		(1.1277)				
3-Month Euro Dollar	3 Mns	0.4267	−1.817*	NA	1.89	0.02	33	5.55***
		(0.3155)	−1.209					3.28*
		(0.4744)						
12-Month T-Bill	3 Mns	0.2909	−4.134***	2.0542	2.13	0.27	67	19.73***
		(0.1715)	−2.939***	(0.4861)				23.31***
		(0.2412)		(0.9818)				
Buyer Bond Index	3 Mns	0.8342	−3.497***	0.1576	2.37	0.87	67	4.09**
		(0.0474)	−3.988***	(0.0927)				5.31***
		(0.0416)		(0.2200)				
30-Year Mortgages	3 Mns	0.7568	−3.523***	0.4209	2.74	0.69	67	6.58***
		(0.0690)	−2.208**	(0.1275)				5.26***
		(0.1101)		(0.2193)				
3-Month T-Bill	6 Mns	−0.1813	−7.027***	0.4642	1.82	0.04	66	15.04***
		(0.1681)	−7.928***	(0.3767)				13.32***
		(0.1490)		(0.8363)				
12-Month T-Bill	6 Mns	−0.1219	−5.374***	2.3456	1.51	0.18	66	22.29***
		(0.2088)	−4.755***	(0.5963)				20.93***
		(0.2360)		(0.8378)				
Buyer Bond Index	6 Mns	0.6355	−5.127***	0.2946	1.29	0.71	66	8.85***
		(0.0711)	−4.875***	(0.1317)				9.03***
		(0.0748)		(0.0984)				
30-Year Mortgages	6 Mns	0.5680	−4.819***	0.4355	1.53	0.47	66	9.15***
		(0.0897)	−4.614***	(0.1552)				11.58***
		(0.0936)		(0.1261)				

[a] $\mathbf{i}_{t+j}^{(k-j)}$ is the yield on a $k-j$-period bill, purchased at time $t+j$, and $\mathbf{i}_t^{(j)}$ is the yield on a j-period bill, purchased at time t. $\mathbf{fp}_t^{(j,k)} = \mathbf{f}_t^{(j,k)} - \mathbf{i}_t^{(j)}$ is the forward premium: the forward rate at time t on a $k-j$-period bill, j periods into the future, less the yield on a j-period bill. Under the null hypothesis that the expectations hypothesis holds and that expectations are Muth rational, $\beta = 1$, and the error term η_{t+j} is a purely random innovation. Upper and lower standard errors (in parentheses) are computed using GMM under the assumption of homoskedasticity and also allowing for conditional heteroskedasticity, respectively.

[b] β_{cc} is the slope parameter dummy for the Carter Special Credit Restraint Program.

*, **, *** Significant at the ten, five, and one percent levels, respectively.

Unless otherwise noted, estimation in all tables below is by OLS, with standard errors calculated using Hansen's (1982) Generalized Method of Moments (GMM). In the regressions at six-month forecast horizons, we allow the residuals to follow an MA(1) process, to correct for the overlapping data problem. Using a technique due to Newey and West (1985), we compute the parameter covariance matrixes in such a way to guarantee they will be positive definite.[17]

[17] This estimator multiplies the lth order autocovariance by $1 - l/(m+1)$. To be conservative, we set $m = 2$ for the MA(1) process.

In addition, the null hypothesis does not imply that the residuals will be homoskedastic. Due to the downward finite-sample bias of the heteroskedasticity-consistent GMM covariance estimates, however, we report two sets of standard errors for the coefficients. The upper set assume the residuals are homoskedastic, and the lower set allow for conditional heteroskedasticity. If we wish to be on the safe side, we should weigh this downward bias more heavily than a loss in power, and therefore draw inferences based on the larger of the two reported standard errors.[18]

All of the regressions below include constant terms, which we do not report to save space. We also include a slope-parameter dummy, β_{cc}, in all the regressions during 1980 when the Fed changed operating procedures, and Jimmy Carter announced, put in place, and then dismantled his temporary Special Credit Restraint Program. Chow tests reject the hypothesis that β does not change during this period. There was, however, no evidence of a change in constant terms. In addition, we tried splitting the sample into pre- and post-1979 subsamples (based on the change in Fed operating procedures), but we could not reject the hypothesis that the coefficients were equal in the two subsamples.

The results in Table I are reminiscent of those from previous studies. All of the point estimates of β are less than one, and all but one are significantly less. The data reject the hypothesis that the spread is an efficient forecast of future interest rate changes. For the shorter-term three-month and twelve-month bills, the parameter estimates are statistically indistinguishable from zero, so we are unable to reject the hypothesis that the spread is of no use in forecasting future changes in short rates. At the six-month forecast horizon, the point estimates for these shorter-term instruments are actually negative.

For the two longer-term instruments in Table I—the Bond Buyer index and thirty-year mortgages—the coefficient estimates are significantly different from both zero and one.[19] At these longer maturities, it is not surprising to find that the spread has predictive power for differences between tomorrow's long rate and today's short rate. Nevertheless, this predictive power does not provide support for the expectations hypothesis.

To see this, consider the usual test of the expectations hypothesis at longer maturities, which asks the spread to predict the upcoming change in the long rate:

$$\mathbf{i}_{t+j}^{(j,k)} - \mathbf{i}_t^{(k)} = \alpha_1 + \beta_1(\mathbf{i}_t^{(k)} - \mathbf{i}_t^{(j)}) + \mu_{1,t,j}, \tag{6}$$

where it can be shown that the null hypothesis implies $\beta_1 = D_j/(D_k - D_j)$, a number slightly greater than zero. Even though β_1 may be close to one, β_1 can still have the wrong sign. For the case of the Bond Buyer index at the six-month forecast horizon,

$$\hat{\beta}_1 = \frac{\hat{\beta}D_k}{D_k - D_j} - 1 = \frac{0.6355D_k}{D_k - D_j} - 1 = (0.6355 \times 1.0426) - 1 = -0.3374$$

[18] See Froot (1989) for evidence of the downward bias in heteroskedasticity-consistent standard errors. The bias is present regardless of the presence of conditional heteroskedasticity.

[19] In calculating the spread for these instruments, we assume that the term structure is flat between durations of $k - j$ and k periods. For example, for a twenty-year instrument such as the Bond Buyer index, this says that the 240-month rate is the same as the 234- and 237-month rates.

with a standard error of $0.0748 \times 1.0426 = 0.0786$. The finding that β_1 is significantly less than zero shows that the slope of the yield curve systematically predicts in the wrong direction the future change in the long rate. This finding is almost universal in tests of long maturities on U.S. post-war data. Thus, estimates of β for long maturities that are statistically less than one, but qualitatively close to one, still provide an economically meaningful rejection of the expectations hypothesis.[20]

IV. Decomposition of the Standard Test

We now split the deviation from the null hypothesis into a component attributable to the term premium and a component attributable to systematic expectational errors. The ex post interest-rate change can be written as the market's expectation plus an ex post prediction error:

$$\mathbf{i}_{t+j}^{(k-j)} - \mathbf{i}_t^{(j)} = \mathbf{e}_t^{(j,k)} + \eta_{t+j}. \tag{7}$$

Using (7), the coefficient β in (5) converges in probability to:

$$\beta = \frac{\mathrm{cov}(\mathbf{e}_t^{(j,k)}, \mathbf{fp}_t^{(j,k)}) + \mathrm{cov}(\eta_{t+j}, \mathbf{fp}_t^{(j,k)})}{\mathrm{var}(\mathbf{fp}_t^{(j,k)})}. \tag{8}$$

With a little algebra, β can be written as one (the null hypothesis) plus a deviation attributable to the term premium plus a second deviation attributable to systematic expectational errors:

$$\beta = 1 + \beta_{\mathbf{tp}} + \beta_{\mathbf{ee}}, \tag{9}$$

where

$$\beta_{\mathbf{tp}} = \frac{-\mathrm{cov}(\theta_t^{(j,k)}, \mathbf{fp}_t^{(j,k)})}{\mathrm{var}(\mathbf{fp}_t^{(j,k)})}, \tag{10}$$

$$\beta_{\mathbf{ee}} = \frac{\mathrm{cov}(\eta_{t+j}, \mathbf{fp}_t^{(j,k)})}{\mathrm{var}(\mathbf{fp}_t^{(j,k)})}. \tag{11}$$

Clearly, $\beta_{\mathbf{tp}}$ is zero if the variance of the term premium is zero—that is, if expectations conform to the expectations hypothesis—and $\beta_{\mathbf{ee}}$ if zero if there are no systematic expectational errors. To obtain estimates of these two components, we use (1) to write the term premium and the ex post prediction error as

$$\theta_t^{(j,k)} = \mathbf{fp}_t^{(j,k)} - \mathbf{s}_{t+j}^{(k-j)} - \epsilon_{t,j}, \tag{12}$$

$$\eta_{t+j} = (\mathbf{i}_{t+j}^{(k-j)} - \mathbf{i}_t^{(j)}) - \mathbf{s}_{t+j}^{(k-j)} - \epsilon_{t,j}. \tag{13}$$

With the help of the survey data these two terms are observable, up to the random measurement error, $\epsilon_{t,j}$. Using (12) and (13) we then can obtain consistent estimates of $\beta_{\mathbf{tp}}$ and $\beta_{\mathbf{ee}}$.

[20] Interestingly, the credit controls dummy term, $\beta_{\mathbf{cc}}$, shows that the spread does have additional positive predictive power when a preannounced, temporary (and large) change in monetary policy takes place. This is reminiscent of the finding in Mankiw and Miron (1987) that the spread had predictive power for future short-rate changes around the turn of the century when there were large seasonal fluctuations in short rates.

The first two columns of Table II present point estimates of these two components to gain a sense for their economic significance. (We test their statistical significance in the following sections.) For completeness, column 3 reports the coefficient β from Table I.

There are several striking results in Table II. First, all of the estimates of β_{ee} are negative and relatively similar in sign and absolute value. This indicates that a positive portion of the deviation of β from its hypothesized value of 1 is attributable to expectational errors.

The second striking fact to emerge from Table II is the different behavior of short versus long maturities. The estimates of β_{tp} explain relatively less of the bias at longer maturities than they do at shorter maturities. In the case of the Bond Buyer index, for example, β_{tp} is relatively unimportant compared with β_{ee} (at the three-month horizon β_{tp} and β_{ee} are 0.010 and −0.176, respectively), and in any case β_{tp} is positive, raising β above one. For the shorter maturities, however, the qualitative contribution of β_{tp} to the overall deviation from the null hypothesis is larger. For example, for three-month Treasury bills at the three-month horizon, $\beta_{tp} = -0.602$ and $\beta_{ee} = -0.338$.

In sum, Table II thus suggests that: (i) systematic expectational errors contribute to the well-documented bias in the spread's predictions; and (ii) the qualitative importance of the term premium for the spread's bias appears greater for shorter-term instruments. The economic importance of risk in the pricing of bills and bonds is investigated further in Section VII below.

Table II

Components of the Failure of the Expectations Hypothesis

Quarterly data, 1969–1986.[a]

Instrument	Forecast Horizon	(1) Component Attributable to the Term Premium β_{tp}	(2) Component Attributable to Expectational Errors β_{ee}	(3) Regression Coefficient $\beta = 1 + \beta_{tp} + \beta_{ee}$
3-Month T-Bill	3 Mns	−0.602	−0.338	0.059
3-Month Euro Dollar	3 Mns	−0.557	−0.016	0.427
12-Month T-Bill	3 Mns	−0.373	−0.336	0.291
Buyer Bond Index	3 Mns	0.010	−0.176	0.834
30-Year Mortgages	3 Mns	−0.051	−0.192	0.757
3-Month T-Bill	6 Mns	−1.048	−0.137	−0.181
12-Month T-Bill	6 Mns	−0.503	−0.619	−0.122
Buyer Bond Index	6 Mns	0.032	−0.395	0.636
30-Year Mortgages	6 Mns	−0.042	−0.389	0.568

[a] β is the regression coefficient from the regression, $\mathbf{i}_{t+j}^{(k-j)} - \mathbf{i}_t^{(j)} = \alpha + \beta \mathbf{fp}_t^{(j,k)} + \eta_{t+j}$, where $\mathbf{i}_{t+j}^{(k-j)}$ is the yield on a $k-j$-period bill, purchased at time $t+j$, $\mathbf{i}_t^{(j)}$ is the yield on a j-period bill, purchased at time t, $\mathbf{fp}_t^{(j,k)} = \mathbf{f}_t^{(j,k)} - \mathbf{i}_t^{(j)}$ is the forward premium: the forward rate at time t on a $k-j$-period bill, j periods into the future, less the yield on a j-period bill. Under the null hypothesis that the expectations hypothesis holds and that expectations are Muth rational, $\beta = 1$ and the error term, η_{t+j} is a purely random innovation. $\beta_{tp} = \dfrac{-\text{cov}\,(\theta_t^{(j,k)}, \mathbf{fp}_t^{(j,k)})}{\text{var}(\mathbf{fp}_t^{(j,k)})}$, where $\theta_t^{(j,k)}$ is the term premium. $\beta_{ee} = \dfrac{\text{cov}(\eta_{t+j}, \mathbf{fp}_t^{(j,k)})}{\text{var}\,(\mathbf{fp}_t^{(j,k)})}$.

V. A Direct Test of the Expectations Hypothesis

We now test whether the biased predictions of the spread can be attributed statistically to a time-varying premium. That is, we test whether $\beta_{\mathbf{tp}} = 0$. To do this we regress the survey expected change on the forward premium:

$$\mathbf{s}_{t+j}^{(k-j)} = \alpha_2 + \beta_2 \mathbf{fp}_t^{(j,k)} + \epsilon_{t,j}. \tag{14}$$

Equation (14) uses the surveys to test the expectations hypothesis directly. Thus the null hypothesis is that $\alpha_2 = 0$ and $\beta_2 = 1$, and $\epsilon_{t,j}$ is purely random measurement error.

It is easy to show that the probability limit of β_2 is:

$$\beta_2 = 1 + \beta_{\mathbf{tp}}. \tag{15}$$

A finding that β_2 is statistically indistinguishable from one implies that we cannot reject $\beta_{\mathbf{tp}} = 0$. Put differently, $\beta_2 = 1$ implies we cannot reject the hypothesis that the variance of the term premium is zero (or, more precisely, that the covariance of the term premium with the forward premium is zero). Notice also that a test of $\beta_2 = 1$ may shed light on the presence of measurement errors in the returns on long-term bonds. Mankiw (1986) and Shiller (1979) discuss the possibility that such errors are responsible for the poor predictions of the spread. A finding of $\beta_2 = 1$ would suggest that the errors-in-variables problem is not important.

There are two alternative hypotheses of interest in (14). The first is the hypothesis that expected interest-rate changes are static, or at least unrelated to the level of the spread. For this we test $\beta_2 = 0$. The second concerns the relative variability of the term premium and expected interest-rate changes. For short-term instruments it is useful to write the coefficient as:

$$\beta_2 = \tfrac{1}{2} + \frac{\mathrm{var}(\mathbf{e}_t^{(j,k)}) - \mathrm{var}(\theta_t^{(j,k)})}{2\ \mathrm{var}(\mathbf{fp}_t^{(j,k)})}. \tag{16}$$

Equation (16) says that if β_2 is statistically less than ½, the variance of the term premium is greater than the variance of expected interest-rate changes. Alternatively, if β_2 is greater than ½, the variance of the term premium is less than the variance of expected interest-rate changes.[21]

A. Results

Table III reports the estimates of (14). To begin, note that the Durbin-Watson statistic in most of the regressions rejects the hypothesis that the residuals are

[21] For longer-term instruments, it is more appropriate to compare the variance of long-rate changes with the variance of the term premium. Some manipulation of (16) and the results in section II yield that:

$$\beta_2 = \frac{D_j}{2D_k}\left(\frac{\mathrm{var}((\mathbf{i}_{t+j}^{(k-j)})^{\mathbf{e}} - \mathbf{i}_t^{(k)}) - \mathrm{var}(\theta_t^{(j,k)})}{2\ \mathrm{var}(\mathbf{f}_t^{(j,k)} - \mathbf{i}_t^{(k)})} - 1\right) + 1.$$

Thus, if $\beta_2 > 1 - \dfrac{D_j}{2D_k}$, the variance of expected long-rate changes is greater than the variance of the term premium.

Table III

Direct Tests of the Expectations Hypothesis

OLS regressions of

$$\mathbf{s}_{t+j}^{(j,k)} = \alpha_2 + \beta_2 \mathbf{fp}_t^{(j,k)} + \epsilon_{t,j}{}^{a}$$

using quarterly data, 1969–1986.

Instrument $(k-j)$	Forecast Horizon (j)	β_2	t: $\beta_2 = 1$	$\beta_{cc}{}^{b}$	DW	$\bar{R}^2$	DF	F-test $\alpha_2 = 0$ $\beta_2 = 1$
3-Month T-Bill	3 Mns	0.3974	−5.958***	0.5715	0.66	0.31	67	23.29***
		(0.1011)	−3.302***	(0.2946)				13.83***
		(0.1825)		(0.3443)				
3-Month Euro Dollar	3 Mns	0.4428	−3.968***	NA	0.57	0.28	33	34.23***
		(0.1404)	−4.984***					18.70***
		(0.1118)						
12-Month T-Bill	3 Mns	0.6273	−3.203***	0.0773	0.47	0.44	67	20.01***
		(0.1164)	−2.603***	(0.2583)				16.12***
		(0.1432)		(0.3102)				
Buyer Bond Index	3 Mns	1.0100	0.364	−0.0024	1.01	0.98	67	0.76
		(0.0276)	0.350	(0.0463)				3.64**
		(0.0287)		(0.0327)				
30-Year Mortgages	3 Mns	0.9493	−1.541	−0.1288	0.64	0.95	67	8.88***
		(0.0329)	−1.636	(0.0565)				48.27***
		(0.0310)		(0.0346)				
3-Month T-Bill	6 Mns	−0.0483	−16.674***	−0.3604	1.00	0.13	67	72.83***
		(0.0651)	−13.460***	(0.1428)				97.12***
		(0.0789)		(0.1560)				
12-Month T-Bill	6 Mns	0.4974	−4.001***	−0.1243	0.56	0.26	67	24.12***
		(0.1256)	−3.499***	(0.3041)				22.97***
		(0.1437)		(0.2149)				
Buyer Bond Index	6 Mns	1.0316	0.895	0.0198	0.68	0.97	67	1.04*
		(0.0353)	1.096	(0.0575)				9.83***
		(0.0288)		(0.0303)				
30-Year Mortgages	6 Mns	0.9576	−0.874	−0.1925	0.60	0.90	67	8.69***
		(0.0485)	−0.869	(0.0807)				23.76***
		(0.0488)		(0.0597)				

[a] $\mathbf{s}_{t+j}^{(k-j)}$ is the survey expected yield on a $k-j$-period bill, purchased at time $t+j$ less the yield at time t on a j-period bill, purchased at time t. $\mathbf{fp}_t^{(j,k)} = \mathbf{f}_t^{(j,k)} - \mathbf{i}_t^{(j)}$ is the forward premium: the forward rate at time t on a $k-j$-period bill, j periods into the future, less the yield on a j-period bill. Under the null hypothesis that the expectations hypothesis holds, $\beta_2 = 1$ and the error term, $\epsilon_{t,j}$, is purely random measurement error. Upper and lower standard errors (in parentheses) are computed using GMM under the assumption of homoskedasticity and also allowing for conditional heteroskedasticity, respectively.

[b] β_{cc} is the slope parameter dummy or the Carter Special Credit Restraint Program.

*, **, ***, Significant at the ten, five, and one percent levels, respectively.

serially uncorrelated. To construct standard errors, we use the covariance matrix estimator suggested by Newey and West (1985) to handle serial correlation and heteroskedasticity of unknown form. We continue to use both homoskedastic and heteroskedasticity-consistent standard errors.

Notice that for instruments with a duration of one year or less, β_2 is statistically less than one. For these instruments, therefore, the estimates of $\beta_{\mathbf{tp}}$ given in Table II are statistically as well as qualitatively significant. In other words, the expectations hypothesis fails at the short end of the maturity spec-

trum.[22] Nevertheless, expectations conform more closely to the expectations hypothesis than the usual ex post regressions in Table I reveal: the estimates of β_2 are in all cases greater than the estimates of β. Also, the $\bar{R}^2$s in Table III are well above those in Table I.

The estimates of β_2 for the longer maturities are more supportive of the expectations hypothesis. All of the estimates for the Bond Buyer index and thirty-year mortage rate are statistically indistinguishable from one at the one-percent level. Indeed, the point estimates for the Bond Buyer index are actually greater than one. We therefore cannot reject the hypothesis that the corresponding estimates of $\beta_{\mathbf{tp}}$ in Table II are zero.[23]

A third feature of the estimates of β_2 is that we can reject the hypothesis that expectations are static. In all but one of the regressions, β_2 is statistically greater than zero.

Finally, while expectations do not appear to be static, there is little evidence that they vary consistently more than do term premia. For the shorter-term instruments (three- and twelve-month bills), the point estimates of β_2 indicate that the variance of the term premium is greater than the variance of expected interest-rate changes. Nevertheless, in only one of the regressions can we reject the hypothesis that the variance of the term premium is equal to the variance of expected short-rate changes (i.e., $\beta_2 = 1/2$). For the longer maturities, we cannot reject either the hypothesis that the variance of the term premium is equal to the variance of expected long-rate changes or the hypothesis that the variance of the term premium is zero.

VI. Tests of Rational Expectations

We now turn to the second explanation for biased predictions of the spread: expectational errors. The survey data give us a unique opportunity to test directly the under-or overreaction hypothesis. Others have tested this hypothesis, but to do so they need to impose the expectations theory as a maintained hypothesis.

A. Over- or Underreaction to the Short Rate

Suppose the market's expectation of the future $k - j$-period interest rate is a linear combination of the contemporaneous short rate and an arbitrary combination of other inputs:

$$(\mathbf{i}_{t+j}^{(k-j)})^{\mathrm{e}} = \omega_1 \mathbf{i}_t^{(j)} + (1 - \omega_1)\mathbf{x}_t, \tag{17}$$

where $0 \leq \omega_1 \leq 1$. Similarly, suppose the actual realized interest rate is a linear combination of the same factors, plus a stochastic news term:

$$\mathbf{i}_{t+j}^{(k-j)} = \omega_2 \mathbf{i}_t^{(j)} + (1 - \omega_2)\mathbf{x}_t + \mu_{t+j}. \tag{18}$$

[22] The F-tests in the last column of each table show the overall importance of the term premium in the survey data. In every case the size of the statistic is sufficient to permit rejection at the one percent level.

[23] The F-tests of $\alpha_2 = 0$, $\beta_2 = 1$ reject for thirty-year mortgages, but not for the Bond Buyer index. This difference is primarily due to differences in the constant terms.

Subtracting (17) from (18), and using (17) to substitute for $\mathbf{x}_t$, we have:

$$\mathbf{i}_{t+j}^{(k-j)} - (\mathbf{i}_{t+j}^{(k-j)}) = \alpha_3 + \beta_3 \mathbf{e}_t^{(j,k)} + \mu_{t+j}, \tag{19}$$

where $\beta_3 = \dfrac{\omega_1 - \omega_2}{1 - \omega_1}$. Under the null hypothesis that the market expectation is rational, $\alpha_3 = \beta_3 = 0$, and the residual, μ_{t+j}, is purely random.

The alternative hypothesis in (19) is that expected future rates over- or underreact to short-rate changes. If β_3 is positive ($\omega_1 > \omega_2$), expectations place a greater weight on the contemporaneous short rate than is rational: the expected future rate overreacts to changes in the short rate. If β_3 is negative, agents' expectations of future rates do not move enough in response to changes in the current short rate: expected rates underreact to short-rate changes.

To estimate (19), we use (1) to get:

$$(\mathbf{i}_{t+j}^{(k-j)} - \mathbf{i}_t^{(j)}) - \mathbf{s}_{t+j}^{(k-j)} = \alpha_3 + \beta_3 \mathbf{s}_{t+j}^{(k-j)} + \mu_{t+j} - (1 + \beta_3)\epsilon_{t,j}. \tag{20}$$

Table IV presents estimates of (20). Because the survey data appear on both sides of (20), OLS estimates of β_3 would be biased toward minus one. We use instrumental variables estimation to eliminate the errors-in-variables problem. Contemporaneous short and long rates are the instruments for $\mathbf{s}_{t+j}^{(k-j)}$. Most point estimates of β_3 in Table IV are less than zero, an indication that expected interest rates underreact to changes in short rates. The estimates for the three-month instruments are not statistically significant. At the long end of the maturity spectrum, however, we can reject the hypothesis that expected future rates respond optimally to sample changes in short rates, in favor of the alternative hypothesis of underreaction.[24]

B. Reaction to the Long-Rate

A second test of the expectational errors turns out to be useful. We regress the survey prediction error on the spread:

$$(\mathbf{i}_{tj}^{(k-j)} - \mathbf{i}_t^{(j)}) - \mathbf{s}_{tj}^{(k-j)} = \alpha_4 + \beta_4 \mathbf{fp}_t^{(j,k)} + \eta_{t+j} - \epsilon_{t,j}, \tag{21}$$

where the null hypothesis is again that $\alpha_4 = \beta_4 = 0$, and the composite error term is purely random.[25] Equation (21) allows us to test for statistical significance the portion of the spread's bias in Table II that is attributable to expectational errors. It is easy to see that the coefficient β_4 is precisely:

$$\beta_4 = \beta_{\mathbf{ee}}. \tag{22}$$

The alternative hypothesis in (21) is that expected future interest rates over- or underreact to changes in the long rate (for a given short rate). To see this, note that by replacing $\mathbf{x}_t$ with $\mathbf{f}_t^{(j,k)}$ in (18) and (19) and then taking the difference

[24] The coefficients in Table IV using instrumental variables are very similar to those obtained using OLS. This suggests that the variance of the survey measurement error is small in comparison with the variance of the market's unobservable expected change, $\mathbf{e}_t^{(j,k)}$.

[25] We use OLS to estimate (21). The results are equivalent to an instrumental variables estimate of (20), where the spread is the instrument. Since our null hypothesis is that $\beta_4 = 0$, either estimation method leads to the same t-test.

Table IV

Tests of Rational Expectations: Reaction to the Short Rate

Regressions of

$$(\mathbf{i}_{t+j}^{(k-j)} - \mathbf{i}_t^{(j)}) - \mathbf{s}_{t+j}^{(k-j)} = \alpha_3 + \beta_3 \mathbf{s}_{t+j}^{(k-j)} + \mu_{t+j} - \epsilon_{t,j}{}^{a}$$

using quarterly data, 1969–1986.

Instrument $(k-j)$	Forecast Horizon (j)	β_3	t: $\beta_3 = 0$	β_{cc} [b]	DW	$\bar{R}^2$	DF	F-test $\alpha_3 = 0$ $\beta_3 = 0$
3-Month T-Bill	3 Mns	−0.0815	−1.151*	−1.9200	2.18	.005	67	1.23
		(0.5410)	−0.163	(1.0829)				0.33
		(0.4999)		(2.6524)				
3-Month Euro Dollar	3 Mns	−0.0476	−0.034	NA	1.79	.000	33	0.56
		(0.4304)	−0.036					0.71
		(0.4106)						
12-Month T-Bill	3 Mns	−0.4225	−1.976***	1.8475	2.48	0.17	67	4.47***
		(0.2138)	−1.584**	(0.4321)				1.28
		(0.2667)		(1.0154)				
Buyer Bond Index	3 Mns	−0.1780	−3.390***	0.1633	2.14	0.15	67	3.98**
		(0.0525)	−3.437***	(0.0986)				3.87**
		(0.0518)		(0.1990)				
30-Year Mortgages	3 Mns	−0.2217	−2.913***	0.6489	2.41	0.26	67	8.40***
		(0.0761)	−1.939*	(0.1445)				4.44***
		(0.0495)		(0.2133)				
3-Month T-Bill	6 Mns	−0.1321	0.121	12.7769	1.62	0.25	66	6.95***
		(1.0902)	0.126	(2.8049)				17.92***
		(1.0465)		(1.7967)				
12-Month T-Bill	6 Mns	−0.6863	−2.682***	1.9525	1.34	0.19	66	4.44***
		(0.2559)	−3.270***	(0.5194)				9.33***
		(0.2099)		(0.4439)				
Buyer Bond Index	6 Mns	−0.3824	−5.150***	0.2633	1.28	0.33	66	8.87***
		(0.0743)	−4.766***	(0.1327)				10.59***
		(0.0802)		(0.1001)				
30-Year Mortgages	6 Mns	−0.4456	−4.392***	0.7529	1.46	.32	66	9.66***
		(0.1015)	−4.822***	(0.1890)				41.66***
		(0.0924)		(0.0756)				

[a] $\mathbf{i}_{t+j}^{(k-j)}$ is the yield on a $k-j$-period bill, purchased at time $t+j$, and $\mathbf{i}_t^{(j)}$ is the yield on a j-period bill, purchased at time t. $\mathbf{s}_{t+j}^{(k-j)}$ is the survey expected yield on a $k-j$-period bill, purchased at time $t+j$, less the yield on a j-period bill, purchased at time t. $\mu_{t+j} - (1+\beta_3)\epsilon_{t,j}$ is a composite error term which is purely random under the null hypothesis that expectations are Muth rational. Upper and lower standard errors (in parentheses) are computed using GMM under the assumption of homoskedasticity and also allowing for conditional heteroskedasticity, respectively.

[b] β_{cc} is the slope parameter dummy for the Carter Special Credit Restraint Program.

*, **, *** Significant at the ten, five, and one percent levels, respectively.

we have (21), with $\beta_4 = \omega_1 - \omega_2$. A finding that $\beta_4 < 0$ implies that ω_2 is "too" large: expectations place *excessive* weight on the contemporaneous long rate. The opposite holds if $\beta_4 > 0$.

Estimates of (21) are given in Table V. The results agree closely with those in Table IV. For the shorter maturities, expectational errors continue to appear unsystematic. We also reject $\beta_4 = 0$ for the longer-maturity instruments, which indicates that the corresponding estimates of β_{ee} are statistically significant. Expectations of long-maturity instruments thus appear to overreact to the long rate: agents would do better to place more weight on the contemporaneous short

Table V

Tests of Rational Expectations: Reaction to the Long Rate

OLS regressions of

$$(\mathbf{i}_{t+j}^{(k-j)} - \mathbf{i}_t^{(j)}) - \mathbf{s}_{t+j}^{(k-j)} = \alpha_4 + \beta_4 \mathbf{fp}_t^{(j,k)} + \eta_{t+j} - \epsilon_{t,j}{}^{a}$$

using quarterly data, 1969–1986.

Instrument $(k-j)$	Forecast Horizon (j)	β_4	t: $\beta_4 = 0$	$\beta_{cc}{}^{b}$	DW	$\bar{R}^2$	DF	F-test $\alpha_4 = 0$ $\beta_4 = 1$
3-Month T-Bill	3 Mns	−0.3382	−1.255	4.4022	2.07	0.24	67	8.09***
		(0.2696)	−1.374	(0.8993)				6.53***
		(0.2461)		(1.0817)				
3-Month Euro Dollar	3 Mns	−0.0161	−0.054	NA	1.83	0.00	33	0.59
		(0.2955)	−0.036					0.30
		(0.4417)						
12-Month T-Bill	3 Mns	−0.3363	−2.209**	1.9769	2.44	0.22	67	7.27***
		(0.1522)	−1.527	(0.4314)				6.06***
		(0.2203)		(0.7326)				
Buyer Bond Index	3 Mns	−0.1759	−3.345***	0.1600	2.11	0.12	67	4.05**
		(0.0526)	−3.356***	(0.1029)				3.87**
		(0.0524)		(0.2156)				
30-Year Mortgages	3 Mns	−0.1925	−2.647***	0.5500	2.42	0.21	67	7.50***
		(0.0727)	−1.789*	(0.1344)				3.63**
		(0.1076)		(0.1934)				
3-Month T-Bill	6 Mns	−0.1373	−0.864	0.8252	1.75	0.08	66	1.80
		(0.1596)	−0.938	(0.3580)				0.84
		(0.1431)		(0.6760)				
12-Month T-Bill	6 Mns	−0.6193	−3.631***	−2.4708	1.66	0.32	66	10.69***
		(0.1706)	−3.400***	(0.4884)				7.73***
		(0.1823)		(0.6194)				
Buyer Bond Index	6 Mns	−0.3948	−5.165***	0.2740	1.24	0.31	66	9.34***
		(0.0764)	−4.661***	(0.1412)				9.61***
		(0.0847)		(0.1063)				
30-Year Mortgages	6 Mns	−0.3886	−3.969***	0.6268	1.38	0.27	66	8.87***
		(0.0979)	−4.235***	(0.1696)				33.66***
		(0.0918)		(0.0795)				

[a] $\mathbf{i}_{t+j}^{(k-j)}$ is the yield on a $k-j$-period bill, purchased at time $t+j$, and $\mathbf{i}_t^{(j)}$ is the yield on a j-period bill, purchased at time t. $\mathbf{s}_{t+j}^{(k-j)}$ is the survey expected yield on a $k-j$-period bill, purchased at time $t+j$, less the yield on a j-period bill, purchased at time t. $\mathbf{fp}_t^{(j,k)} = \mathbf{f}_t^{(j,k)} - \mathbf{i}_t^{(j)}$ is the forward premium: the forward rate at time t on a $k-j$-period bill, j periods into the future, less the yield on a j-period bill. $\eta_{t+j} - \epsilon_{t,j}$ is a composite error term which is purely random under the null hypothesis that expectations are Muth rational. Upper and lower standard errors (in parentheses) are computed using GMM under the assumption of homoskedasticity and also allowing for conditional heteroskedasticity, respectively.

[b] β_{cc} is the slope parameter dummy for the Carter Special Credit Restraint Program.

*, **, ***, Significant at the ten, five, and one percent levels, respectively.

rate and less weight on the long rate in forming their expectations of future long rates.

VII. Variation in Term Premia

Our findings thus far could be summarized as documenting the importance of term premia for instruments of shorter duration and the unimportance of term

Table VI

Components in the Slope of the Term Structure

Quarterly data, 1969–1986.[a]

		Mean of:			
		(1)	(2)	(3)	(4)
Instrument $(k-j)$	Forecast Horizon (j)	$\mathbf{fp}_t^{(j,k)}$ Forward Premium	$\mathbf{s}_{t+j}^{(k-j)}$ Expected Change	$\hat{\theta}_t^{(j,k)}$ Term Premium	$\left(\frac{D_k - D_j}{D_j}\right)\hat{\theta}_t^{(j,k)}$ Holding Premium
3-Month T-Bill	3 Mns	0.406	−0.061	0.468	0.468
3-Month Euro Dollar	3 Mns	0.569	−0.279	0.850	0.850
12-Month T-Bill	3 Mns	0.997	0.175	0.820	3.318
Buyer Bond Index	3 Mns	0.082	−0.151	0.070	2.915
30-Year Mortgages	3 Mns	3.051	2.822	0.223	8.418
3-Month T-Bill	6 Mns	2.031	−0.241	2.277	1.133
12-Month T-Bill	6 Mns	1.022	0.018	1.004	2.018
Buyer Bond Index	6 Mns	−0.296	−0.362	0.066	1.339
30-Year Mortgages	6 Mns	2.920	2.574	0.337	6.100

[a] $\mathbf{fp}_t^{(j,k)} = \mathbf{f}_t^{(j,k)} - \mathbf{i}_t^{(j)}$ is the forward premium: the forward rate at time t on a $k-j$-period bill, j periods into the future, less the yield on a j-period bill. $\mathbf{s}_{t+j}^{(k-j)}$ is the survey expected yield on a $k-j$-period bill, purchased at time $t+j$ less the yield at time t on a j-period bill, purchased at time t. $\hat{\theta}_t^{(j,k)} = \mathbf{fp}_t^{(j,k)} - \mathbf{s}_{t+j}^{(k-j)}$. The term D_j is Macaulay's definition of duration on a bond with j periods to maturity. All figures are expressed in percent per annum.

premia for instruments of longer duration in the biased forecasts of the spread. Could it be that risk is more important in pricing short-term bills than long-term bonds? In this section we ignore the restrictions implied by the expectations hypothesis and investigate the survey term premia directly. To preview our findings, the answer to the above question is no. Term premia become increasingly important in pricing bonds as duration increases.

Table VI presents means of the data used in the foregoing tests, expressed in percent per annum. In the first column is the forward premium, $\mathbf{fp}_t^{(j,k)}$. Using the survey data, we separate $\mathbf{fp}_t^{(j,k)}$ into the survey expected change, $\mathbf{s}_t^{(k-j)}$, and term premium, $\hat{\theta}_j^{(j,k)}$. As long as the survey measurement error is random, these averages are consistent estimates of the true market values. Column (4) reports the survey estimate of the holding premium, $(\mathbf{h}_t^{(j,k)})^{\mathrm{e}} - \mathbf{i}_t^{(j)}$, the expected excess return from holding a k-period bond for j periods. Note that the holding premium generally increases with duration: the average expected excess return to holding six-month bills for three months is 0.47 percent per annum, while the average expected excess return to holding tax-exempt twenty-year Buyer bonds for three months is about 2.92 percent per annum.[26]

The fact that average holding premia rise with duration does not itself imply that time variation in term premia is important. We gain a sense of the relative variability of the premia in two ways. The most direct route is to plot the survey premia (although one must bear in mind that they are contaminated by meas-

[26] Kane (1983) reports similar findings in his analysis of a different survey source. See also Fama (1984a) who uses ex post data to measure term premia.

urement error). Figures 1 and 2 display the term structure of the premia for a six-month holding-period for several instruments. It is clear that as duration increases, both the mean and the variability of the survey holding premium increase. Even though the size of the holding premium for nine-month bills is

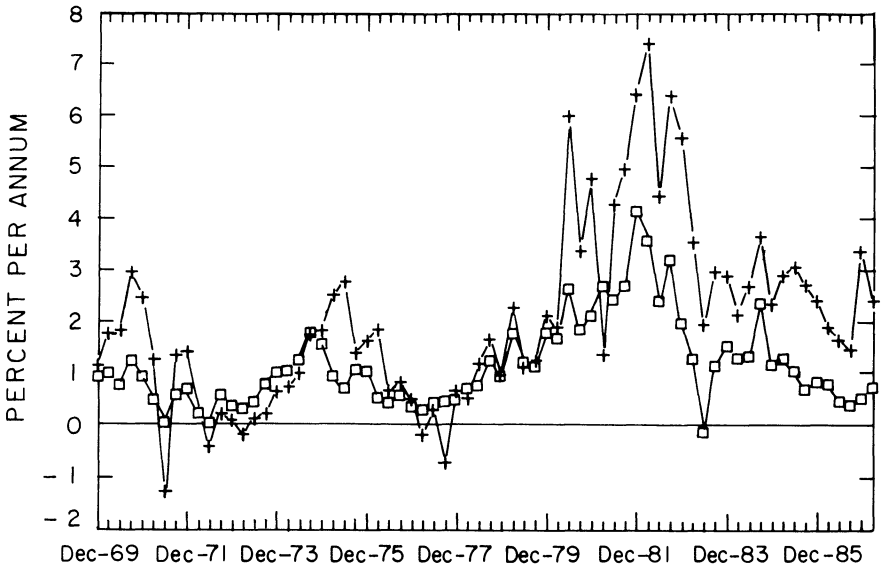

Figure 1. The term structure of the holding premium over a six-month holding period. □ nine-month Treasury bills, + eighteen-month Treasury Notes.

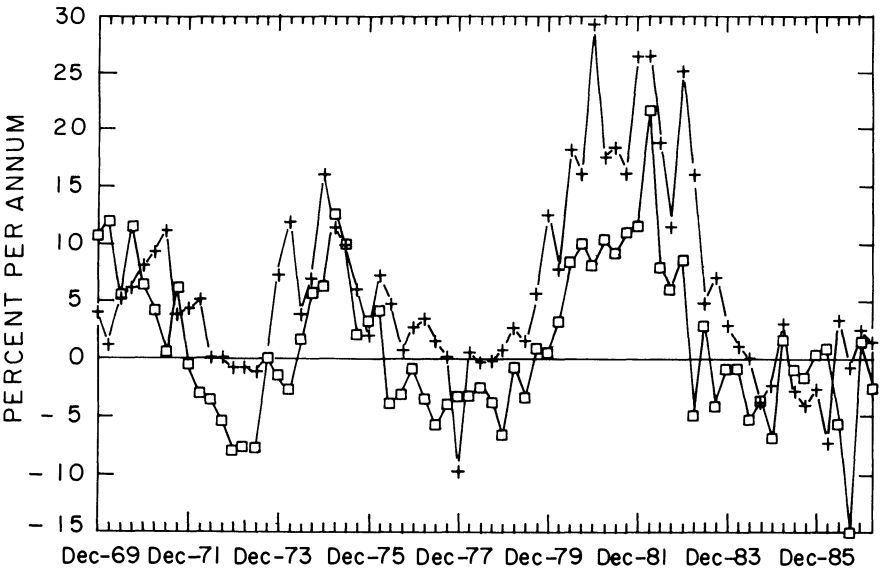

Figure 2. The term structure of the holding premium over a six-month holding period. □ Bond Buyer index, + thirty-year mortgages.

Table VII

Survey Risk Premium on Short and Long Rates

OLS regressions of

$$\hat{\theta}_t^{(j,k)} = \alpha_5 + \beta_l \mathbf{i}_t^{(k)} + \beta_s \mathbf{i}_t^{(j)} + \epsilon_{t,k}{}^{a}$$

using quarterly data, 1969–1986.

Instrument $(k-j)$	Forecast Horizon (j)	β_s	β_l	DW	$\bar{R}^2$	DF	F-test $\alpha = \beta_s = \beta_l = 0$
3-Month T-Bills	3 Mns	−1.3203	1.4272	0.95	0.56	67	44.20***
		(0.1488)	(0.1555)				
3-Month Euro Dollar	3 Mns	−1.3780	2.1337	1.13	0.71	67	85.40***
		(0.2586)	(0.2569)				
12-Month T-Bill	3 Mns	−1.5734	1.7968	1.37	0.76	32	53.29***
		(0.1618)	(0.1768)				
Buyer Bond Index	3 Mns	0.5273	1.9494	1.26	0.16	67	7.75***
		(0.6993)	(0.9228)				
30-Year Mortgages	3 Mns	−1.2102	3.7015	0.81	0.32	67	17.56***
		(0.8895)	(0.8914)				
3-Month T-Bill	6 Mns	−0.9291	1.0712	0.78	0.93	67	393.57***
		(0.0899)	(0.0670)				
12-Month T-Bill	6 Mns	−0.7000	1.3797	0.88	0.78	67	108.53***
		(0.1674)	(0.1597)				
Buyer Bond Index	6 Mns	0.8698	0.5381	0.80	0.22	67	7.58***
		(0.4259)	(0.5380)				
30-Year Mortgages	6 Mns	0.0397	1.816	0.63	0.36	67	14.40***
		(0.6575)	(0.5120)				

[a] $\hat{\theta}_t^{(j,k)}$ is the survey term premium on a k-period bond, above the j-period rate and the expected future rate on a $k-j$-period bill purchased at time $t+j$. $\mathbf{i}_t^{(k)}$ and $\mathbf{i}_t^{(j)}$ are the yields on a k-period and a j-period bill at time t, respectively.

small relative to longer maturities, a premium of 100 basis points (which is not unusual) on U.S. government Treasury bills still seems large in absolute terms. The surveys suggest that term premia rose substantially during periods of high interest-rate volatility.[27]

While these figures are of interest, it is possible that the survey premia vary over time primarily because of measurement error. To remove the measurement error we obtained the predicted values from a regression of the survey premium in (12) on a constant, the current short rate, and the current long rate, and then computed a "cleaned" holding premium. These regressions are reported in Table VII. As a benchmark measure, we estimated the predictable component of actual returns from a regression of the ex post holding returns on the same regressors. We graph in Figures 3 and 4 the cleaned premia measures for two of the instruments in Figures 1 and 2: six-month holding premia for both nine-month Treasury bills and the Bond Buyer index.[28]

[27] To clarify the maturities of the holding premia, note that the expected return on three-month (twelve-month) bills six months into the future allows us to compute the six-month expected holding premium on nine-month (eighteen-month) bills.

[28] The cleaning regression has little effect on the short-term premia, as can be seen by comparing Figures 1 and 3. By comparison, cleaning smooths out the survey premia for the Bond Buyer index in Figures 2 and 4.

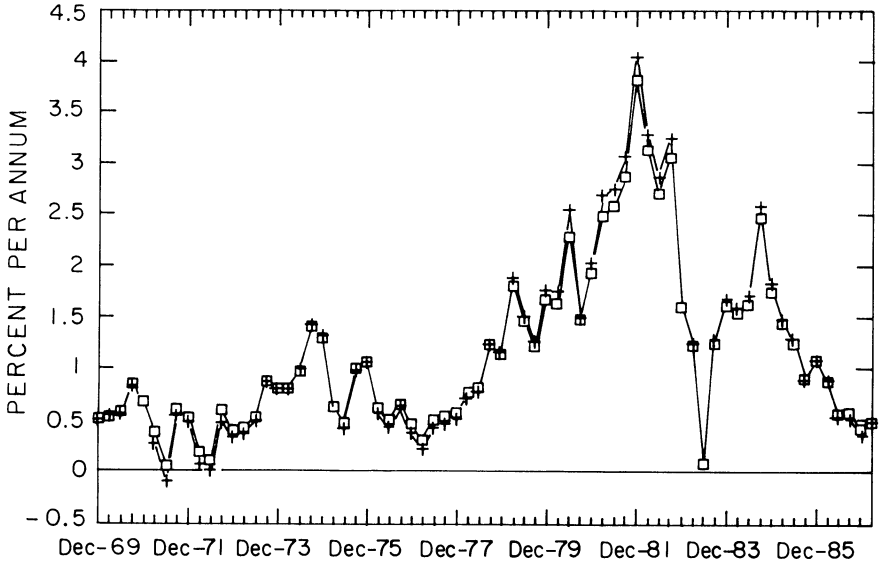

Figure 3. Expected excess holding-period returns over six months for nine-month Treasury bills. □ cleaned survey premia, + predictable excess holding-period returns.

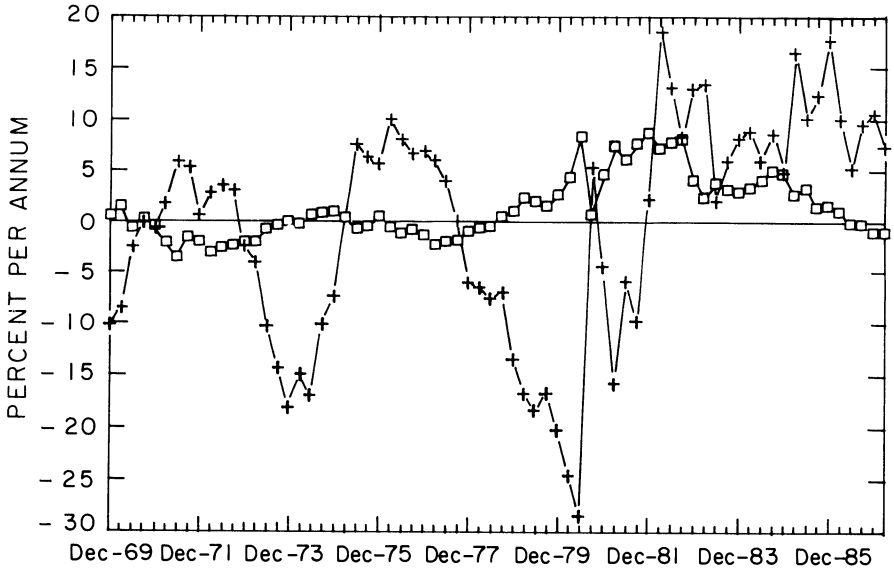

Figure 4. Expected excess holding-period returns over six months for the Bond Buyer index. □ cleaned survey premia, + predictable excess holding-period returns.

Two striking observations come out of Figures 3 and 4. First and foremost is the powerful visual impression created by how closely the survey premia (after being purged of any measurement error) move with predictable excess returns on short-maturity bills in Figure 3. It appears the short-maturity surveys do not tell us anything new about term premia that we could not have learned with ex post

Table VIII

Variation in Estimates of the Holding Premium

Quarterly data, 1969–1986.[a]

		Variance of:	
Instrument ($k-j$)	Forecast Horizon (j)	(1) Cleaned Survey Premium	(2) Predictable Excess Holding Returns
3-Month T-Bill	3 Mns	0.003	0.004
		(0.000)	(0.055)
3-Month Euro Dollar	3 Mns	0.005	0.005
		(0.000)	(0.046)
12-Month T-Bill	3 Mns	0.095	0.157
		(0.000)	(0.000)
Buyer Bond Index	3 Mns	0.426	1.870
		(0.006)	(0.034)
30-Year Mortgages	3 Mns	0.886	2.097
		(0.000)	(0.019)
3-Month T-Bill	6 Mns	0.011	0.013
		(0.000)	(0.000)
12-Month T-Bill	6 Mns	0.040	0.088
		(0.000)	(0.000)
Buyer Bond Index	6 Mns	0.163	1.984
		(0.001)	(0.004)
30-Year Mortgages	6 Mns	0.404	1.471
		(0.000)	(0.026)

[a] Figures above are estimated variances obtained from projections of the survey term premium and realized excess returns, respectively, onto a constant and the appropriate long and short rates. Estimates are annualized variances, expressed in percent. In parentheses are the probability values from joint F-tests that the variances are zero.

realizations. Contrast this with the markedly different behavior of the two series for longer maturities, graphed in Figure 4. The long-maturity survey premia differ radically from the predictable excess returns on bonds.

A second striking fact emerges in Figures 3 and 4: as duration increases, the cleaned survey premia are substantially less volatile relative to the predictable component of excess returns. In Figure 4 the survey premium is much smoother than are predictable excess returns, which exhibit enormous swings. Although the survey premia are smooth they will vary considerably. Changing perceptions of risk are clearly an important determinant of changes in bond prices.

Table VIII evaluates the statistical significance of fluctuations in these two measures of risk premia. For each asset we report the point estimate for the variance and the probability that the variance is equal to zero (in parentheses). The probabilities are from F-tests that the coefficients on the long and short rates in the "cleaning" regressions are jointly zero. Even though the estimated variance of predictable excess returns is relatively large, we frequently cannot reject the hypothesis that the actual variance is zero. By constrast, we strongly

reject the hypothesis that the cleaned survey premia have zero variance. Thus, while risk appears more variable when extracted from excess returns, it is measured less precisely than when extracted from the surveys.

Finally, the survey premia in Figures 3 and 4 are intuitively very reasonable: they are highly positively correlated with nominal interest rates and inflation. They are also smooth, suggesting that long-horizon perceptions of underlying economic fundamentals change slowly. By contrast, the predictable excess returns in Figure 4 are highly volatile and less easily understood.[29]

VIII. Conclusions

We used survey data on interest-rate expectations to investigate the reasons why the spread is such a poor predictor of future interest-rate changes. In evaluating the conclusions below, it is worth remembering that our statistical results rely on the identifying assumption that the survey data measure accurately the market's (unobservable) expectation, up to random measurement error.

Our major findings are summarized:

(1) We confirm earlier findings that predictions of future interest-rate changes by the spread contain bias of a similar nature for short and long maturities. The explanations for this bias, however, differ markedly at opposite ends of the maturity spectrum.

(2) *We use the survey data to test directly the expectations hypothesis on short-term instruments and we reject it.* The test does not require the auxiliary assumption of rational expectations. The surveys on short-maturity instruments reveal term premia which are both large and variable. For short maturities, the biased predictions of the spread are predominantly attributable to time variation in term premia. Changes in the slope of the yield curve therefore reflect changing perceptions of risk.

(3) *We find little evidence that expected future short rates underreact to current short-rate changes.* While our point estimates suggest underreaction, they are not statistically significant. We therefore cannot reject the hypothesis that the market's expectation of future short rates is rational. This is the same as saying that we cannot reject the hypothesis that the spread's bias in short-maturity instruments is entirely attributable to term premia.

(4) *The survey data show that expected long-rate changes conform with the expectations theory, in that changes in the spread are reflected one-for-one in changes in expected future long rates.* This fact suggests that the frequently-cited tendency of the spread to predict long-rate changes perversely cannot be explained by errors made in measuring long-term rates, or by variation in term premia. We cannot reject the hypothesis that the premium is uncorrelated with the spread, even though the survey premium exhibits significant variability after the measurement error is removed.

[29] Fama and Bliss (1987) use an autoregressive measure of expected future interest rates to extract a measure of the term premium. Their measure seems to move procyclically with the business cycle and appears negatively correlated with nominal short-term interest rates.

(5) *The inability of the spread to forecast future long-rate changes is attributable primarily to systematic expectational errors.* We cannot reject the hypothesis that the change in the long rate in excess of the spread is attributable exclusively to these expectational errors. Expected profits from trading on these prediction errors are small and highly variable (see Mankiw and Summers (1984)). The behavior of the expectational errors suggests that expected future rates *underreact* to changes in the short rate. Mankiw and Summers (1984) and Campbell and Shiller (1984) also interpret their results as evidence that long rates underreact; our evidence suggests that this result cannot be attributed to term premia.

(6) *Perceptions of risk become increasingly important in the pricing of bonds as duration increases.* We find that average expected excess holding-period returns increase with maturity. We also document large and statistically significant swings in term premia on long-term bonds, and substantially smaller (but nevertheless significant) swings on short-term bills. The survey premia on long-term bonds are much smoother than the predictable component of realized excess returns, but are measured much more precisely. In contrast to the behavior of predictable excess returns, these premia are large when inflation and nominal interest rates are high.

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