Inflation Forecast Errors and Time Variation in Term Premia

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Abstract

The expectations theory of the term structure is well known to give wrong signals as to the future course of long-term interest rates. One explanation involves rational time-varying term premia. However, the "anomaly" may also be due to inflation forecast errors. We study survey forecasts of inflation. It seems that the respondents' forecasts are insufficiently adaptive. Interest rates reflect expectations similar to the inflation forecasts. As a result, past survey forecast errors reliably predict premia on U.S. Government Bonds.

I. Introduction

The puzzling behavior of nominal and real interest rates, and the role of inflationary expectations, have long been studied in finance. In theory, nominal rates (at all maturities) should vary one-for-one with movements in expected inflation. Yet, according to the economist most closely associated with this theory, Irving Fisher ((1930), pp. 493–494, 415),

> ... the money rate of interest, while it does change somewhat according to the theory, ... does not usually change enough to fully compensate for the appreciation or depreciation.... Men are unable or unwilling to adjust at all accurately and promptly the money interest rates to changed price levels.

Evidently, Fisher (1928), (1930) believed that the public's forecasts of consumer prices are subject to systematic bias, too low in periods of rising inflation and too high in periods of falling inflation.[1]

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[1]Fisher went on to argue that "negative real interest rates could scarcely occur if contracts were made in a composite commodity standard. The erratic behavior of real interest rates is evidently a trick played on the money market by ... money illusion." He relied on this concept to explain the Gibson paradox. Friedman and Schwartz (1982) and Summers (1983) largely support this view, but Barsky (1987) argues against it. Yet, there is wide agreement that, before 1930, little of the variation in interest rates is explained by inflation, actual or expected.

The slow adjustment of expectations and market interest rates to past inflation, however, may also be consistent with rationality and Bayesian learning. This depends on investors' prior beliefs, the predictability of inflation, and the uncertainty associated with possible regime changes. A study by Barsky and De Long (1988) provides a good example. These authors examine gold discoveries, inflation, and interest rates during the 1890s. Standard orthogonality conditions fail. Investors did not expect a change in trend inflation, which now, after the fact, looks predictable. But, this alleged lack of foresight reflected the belief, quite reasonable at the time, that increased gold production could not be sustained.

In this paper, we study the 1953–1987 period and document the same inertia in inflationary expectations and interest rates that other authors have discussed for earlier times. Our more novel evidence is that these inflation forecast errors are related to predictable time-variation in term premia on U.S. Government Bonds. We leave it to the reader to decide whether the findings are evidence of naive information processing by investors.

Empirically, it is quite difficult to capture the public's beliefs about future changes in the price level. The past literature follows two tracks. One approach is to start from survey expectations data. But, of course, surveys do not always describe the forecasts inherent in market behavior. A second method is to infer the unobservable market expectation from security prices. In that case, one has to correct for movements in expected real returns—say, a Tobin-Mundell effect or supply shocks (Peek and Wilcox (1983))—and taxes (Darby (1975)).

This paper relies on both approaches. After a description of the data in Section II, Section III examines the forecasts of the Consumer Price Index (CPI) gathered by Joseph Livingston of the *Philadelphia Enquirer*. It appears that the respondents' forecasts are insufficiently adaptive. The survey participants are reluctant to predict inflation levels that greatly differ from past experience.

Based on new tests that combine survey with price data, Sections IV and V examine whether market interest rates reflect similar expectations. Section IV finds that the inflation forecast errors that are implicit in ex post real rates parallel those of the Livingston surveys.

Section V relates the evidence to the term structure literature. As early as 1938, Macaulay observed that interest rates do not move as the expectations theory predicts. In fact, "experience is more nearly the opposite" ((1938), p. 33). Under the joint hypothesis of the expectations theory and rationality, term premia should be unforecastable. Yet, in actuality, they are positively related to the slope of the term structure (for a review, see Shiller and McCulloch (1990)). These findings support either a form of less-than-fully rational expectations, or time-varying risk premia, or both. Ex post, inflation forecast errors appear as part of the premia. The surveys and the results for ex post real rates suggest that, after periods of low (high) inflation, long rates are slow to rise (fall) relative to short rates. This inertia therefore means that, compared to money market instruments, bonds perform poorly when inflation accelerates, but that they perform well when inflation subsides.

By combining survey with market data, we can distinguish changing expectations from changing premia. In agreement with the expectations theory, variation in the term structure slope clearly reflects changes in survey expected inflation.

However, movements in the slope and in term premia are also related to *past* survey forecast errors. If recent past inflation numbers are higher than expected, the survey expected change in inflation, the term structure slope, and term premia all predictably decline. When recent inflation is lower than expected, the opposite occurs. These results help to explain the predictive power of the slope for term premia.

II. Data

We investigate the period between June 1953 and June 1987. The annualized inflation rate $\pi_{t,k}$ is defined as the change from the end of month t to month $t+k$ in the natural log of the U.S. seasonally unadjusted CPI. The CPI forecasts are taken from the Livingston surveys. Twice a year, in early May and November, professional economists predict several macroeconomic variables including the CPI. We work with consensus forecasts, i.e., the means of (on average) 52 individual predictions. Carlson (1977) argues that, since the May (November) questionnaires are returned in early June (December), the latest CPI information available to forecasters is the price level for April (October). Therefore, the 8- and 14-month forecasts solicited in May are found by comparing the CPI for the previous April with the predicted levels for December or June of the subsequent year. Likewise, the forecasts solicited in November are computed on the basis of the October CPI.[2]

The interest rate data are from two sources: 1) the "Fama Term Structure Files" produced by the Center for Research of Security Prices (CRSP) at the University of Chicago; and 2) the data collected by McCulloch (Shiller and McCulloch (1990)). CRSP finds spot yields from averages of end-of-month bid and asked prices of Treasury Bills and Government Bonds. Most CRSP data refer to the period between 1959 and 1986. McCulloch provides a zero coupon yield curve series (computed on an annual percentage continuously compounded basis) between December 1946 and February 1987. The yields are available for maturities ranging from 0 through 6 months, 9 months, 1 through 5 years, as well as 10, 15, and 20 years. The data sets include most marketable U.S. Government Treasury Bills, Notes, and Bonds.[3]

III. Survey Forecasts of Inflation

A. Testing for Survey Rationality

We use three approaches to study the rationality of the consensus forecasts. Our initial tests are for unconditional bias. Next, we compare the forecast errors of the Livingston surveys with those of time-series models. Finally, and most importantly, we model how the survey expectations were formed and we examine their rationality from this point of view.

[2] All calculations were repeated under the assumption that forecasts were for 6 (7) and 12 (13) months. We also defined the consensus forecast as the median of the individual predictions. None of these robustness checks had a big enough impact on the estimates to affect their interpretation.

[3] See Shiller and McCulloch (1990) for details. Below, we sometimes estimated the yield on 8- or 14-month instruments by linear interpolation using, respectively, the yields on 6- and 9-month bills, and 1- and 2-year zero-coupon bonds.

To test for unconditional bias, the regression equation is

$$\pi_{t,k} = \alpha + \beta\pi^e_{t,k} + u_{t,k}, \tag{1}$$

where $\pi^e_{t,k}$ is the forecast at t of the annualized inflation rate between months t and $t+k$. We choose our notation so that t corresponds to the last trading day of April (October) and k equals either 8 or 14.

Tests of rationality based upon OLS are inappropriate because the usual assumptions concerning the residuals are not met. In particular, since the forecast interval is longer than the sampling interval (which is 6 months), the prediction errors for future periods are serially correlated. Correct standard errors are found using Hansen and Hodrick's (1980) method of moments with two adjustments: 1) White's (1980) correction for conditional heteroskedasticity; and 2) the correction of Newey and West (1987), which insures that the variance-covariance matrix is positive definite by discounting the mth order autocovariance.[4] We estimate Equation (1), but we use the corrected covariance matrix. Unbiased forecasts are characterized by $(\alpha, \beta) = (0, 1)$ or $B = B_0$. The test statistic for the joint null hypothesis is $q = (\hat{B} - B_0)'X'X(X'QX)^{-1}X'X(\hat{B} - B_0)$. It is distributed χ^2 with $k = 2$ degrees of freedom.

Our second approach follows Theil (1966) and provides some insight into the causes of survey forecast error. Theil decomposes the mean squared error (MSE) into U^M (the proportion of MSE due to mean level bias), U^R (the proportion due to regression bias), and U^D (the proportion due to random error). The three terms on the right-hand side of Equation (2) correspond to the sources of error,

$$\text{MSE} = 1/N\sum(\pi^e_{t,k} - \pi_{t,k})^2 = (P - A)^2 + (s_p - rs_a)^2 + (1 - r^2)s_a^2, \tag{2}$$

where P and A are the mean values of $\pi^e_{t,k}$ and $\pi_{t,k}$, s_p and s_a are the corresponding standard deviations, and r is the correlation coefficient.

To provide an additional perspective, we use Equation (2) to evaluate two other sets of forecasts: 1) a random walk model (RW); and 2) an interest rate model (FG). The forecast dates match those of the Livingston surveys (LS). RW assumes that inflation remains unchanged from one 8- or 14-month period to the next. Thus, all changes in inflation are permanent. FG is as in Fama and Gibbons (1984). Isolating the inflation forecast implicit in interest rates requires a model of the ex ante real return. Fama and Gibbons find that month-to-month changes in the ex post real rate follow an MA(1)-process. Real returns are well approximated by an equally-weighted average of real rates for the last 12 months. We subtract this average from the 1-month Treasury Bill yield observed at the end of April and October. The resulting 1-month expected inflation number is multiplied by either 8 or 14.

A third approach to characterize the Livingston forecasts—and the most relevant for our purposes—is to model how inflation expectations are formed. As

[4]The asymptotically consistent covariance matrix V equals $(X'X)^{-1}X'QX(X'X)^{-1}$, where X is the matrix of regressors (N observations; $k = 2$ predictors) and Q is the adjusted covariance matrix of the residuals. The (i,j)th element of Q, denoted with $w(i,j)$, equals zero except when $m \leq n$, where $m = |i - j|$ and n is the order of the moving average process of the residuals. If $m \leq n, w(i,j) = [1 - m/(n+1)]u_iu_j$. For the 8-month forecasts, we postulate a first order moving average process; for the 14-month forecasts, a second order moving average process.

observed by Frankel and Froot (1987), four simple models of expectation formation are described by

$$(3) \qquad \pi^e_{t,k} = \theta Z + (1-\theta)\pi_{t-k,k},$$

where $0 \leq \theta \leq 1$, and Z represents either the twice-lagged inflation rate $\pi_{t-2k,k}$ (extrapolative expectations), the lagged forecast $\pi^e_{t-k,k}$ (adaptive expectations), or a long-run constant equilibrium level of inflation π (regressive expectations). Also, for a random walk process (static expectations), $\theta = 0$. The regression models that we estimate are obtained by subtracting $\pi_{t-k,k}$, the most recently observed inflation rate, from both sides of Equation (3). Then,

$$(4) \qquad \pi^e_{t,k} - \pi_{t-k,k} = \lambda(\pi_{t-2k,k} - \pi_{t-k,k}),$$

$$(5) \qquad \pi^e_{t,k} - \pi_{t-k,k} = \gamma(\pi^e_{t-k+2,k} - \pi_{t-k+2,k-2}),$$

$$(6) \qquad \pi^e_{t,k} - \pi_{t-k,k} = \delta(\pi - \pi_{t-k,k}).$$

In each case, the dependent variable is the expected change in inflation.[5] The regressor for the adaptive model is adjusted to insure that it is based only on information known at t, the time of the forecast. For example, for 8-month forecasts, $\pi^e_{t-6,8}$ is compared with $\pi_{t-6,6}$, the annualized inflation rate for the first 6 months of the forecast period. Since π is assumed constant over the sample period, the regressor for the regressive model is $-\pi_{t-k,k}$.

To test for rationality, we compare the parameters of Equations (4)–(6) with the imaginary "true" parameters (λ^*, γ^*, and δ^*) that describe the inflation process. Starting from Equation (3), we subtract actual inflation, $\pi_{t,k}$, from both sides and obtain

$$(4') \qquad \pi^e_{t,k} - \pi_{t,k} = (\lambda - \lambda^*)(\pi_{t-2k,k} - \pi_{t-k,k}),$$

$$(5') \qquad \pi^e_{t,k} - \pi_{t,k} = (\gamma - \gamma^*)(\pi^e_{t-k+2,k} - \pi_{t-k+2,k-2}),$$

$$(6') \qquad \pi^e_{t,k} - \pi_{t,k} = (\delta - \delta^*)(\pi - \pi_{t-k,k}).$$

In (4′)–(6′), the dependent variable is the inflation survey forecast error. As earlier for Equation (1), we face the econometric difficulties arising from overlapping observations. We again rely on the methods of Hansen and Hodrick, White, and Newey and West. Under rational expectations, the estimated coefficients should not be significantly different from zero. The advantage of running Equations (4′)–(6′) as well as (4)–(6) is that, if rationality fails, we have ready-made descriptions of the nature of the bias.

[5] Since the regressions are run with consensus forecasts, an aggregation theorem is implicit in their formulation. For example, Bierwag and Grove (1966) show that if expectations are formed adaptively, the mean forecast is not adaptive unless the adjustment coefficient is the same for all individuals. See also Keane and Runkle (1990).

Even though the forecast periods are longer than the sampling period, this should not cause serial correlation in the error terms. The reason is that the forecasts are based on past information. However, as it turns out, the OLS regressions do show serial correlation. Therefore, we use GLS Prais-Winsten estimates. We retain White's correction for heteroskedasticity.

B. Results

Table 1 reports the test results for unconditional bias. For the 1953–1987 period, the joint hypothesis that $(\alpha, \beta) = (0, 1)$ cannot be rejected. However, the positive intercept indicates that the forecasts are systematically too low. But, during the 1980s, predictions are too high. The Durbin-Watson coefficient from the OLS-regression shows that a correction for serial correlation is indeed necessary. Also, for 1953–1987, White's (1980) heteroskedasticity test (not shown) yields a χ^2 of 13.4 (12.1) for the 8- (14-) month horizon. Both numbers are significant at the 1-percent level.

TABLE 1

Regressions of Actual Inflation Rates on Livingston Consensus (Mean) Forecasts of Inflation

Periods	α	t_α	β	t_β	Q-Test	D-W	Adj. R^2	N
Eight-Month Forecasts								
4/53–4/87	0.856	2 86	1.055	0.51	2.08	0.77	0.711	69
4/53–4/79	0.812	2 83	1.268	2 17	4.91*	1.08	0.789	53
10/79–4/87	−2.072	−2.11	1.269	1.46	1.70	1 60	0.813	16
Fourteen-Month Forecasts								
4/53–10/86	0.924	2.77	1.003	0.02	1.45	0.42	0.657	68
4/53–4/79	0.706	2.20	1.295	1.97	3.50	0.58	0.792	53
10/79–10/86	−2.355	−2 08	1.205	1 10	2.68	1.18	0.778	15

The regressions are OLS with the t-statistics corrected for serial correlation and heteroskedasticity with the methods of Hansen and Hodrick, White, and Newey and West. The t-statistic for the slope coefficient tests whether it equals one. D-W is the Durbin-Watson coefficient. Q tests the joint hypothesis that $(\alpha, \beta) = (0, 1)$. Q-entries marked with "*" are significant at the 1-percent level. N is the number of observations

Previous studies (Huizinga and Mishkin (1984), (1986) and Clarida and Friedman (1984)) conclude that the relationship between inflation and nominal interest rates shifts with the October 1979 change in Federal Reserve policy. We find evidence of a 1979 structural break in the survey data using Wald tests. The test yields a χ^2 of 23.7 (39.3) for a shift in β for the 8- (14-) month horizon. Tests for shifts in both α and β produce even stronger rejections of parameter stability.

More evidence consistent with mean level bias and a break in the survey forecasts around 1979 is found in Table 2. The mean absolute, average, and root mean square errors (RMSE) indicate that, at least until 1979, the random walk and interest rate models perform well relative to the surveys and that they show little systematic bias. In contrast, for the 8-month forecasts, about one quarter of the survey errors is attributable to a tendency to underpredict inflation. However, after 1979, the surveys beat RW and FG. See Hafer and Hein (1985) for similar results based on ASA-NBER survey data. A reasonable conjecture is that the improved relative performance of the surveys merely follows from the 1979 shift in Federal Reserve policy, a turning point that the other models miss by construction. Yet, if we remove the October 1979 and April 1980 observations, the result stands. For instance, for the 8- (14-) month forecasts, the RMSE for the surveys is 1.32 (1.81). It is 2.33 (2.62) for the Fama-Gibbons interest rate model.

Table 3 presents our estimates of Equations (4)–(6) and (4′)–(6′). To save space, we only present the 1953–1987 results. There is considerable agreement

TABLE 2

Theil-Analysis for Three Sets of Inflation Forecasts: 1) Livingston Surveys; 2) Random Walk Model; 3) Fama-Gibbons Interest Rate Model

Periods	Model	MAE	RMSE	U^M	U^R	U^D	AVE	s_a	s_p	r^2	N
Eight-Month Forecasts											
4/53–10/86	LS	1.54	2.16	0.25	0.00	0.75	−1 06	3 56	2.85	0.72	68
	RW	1.46	1.92	0 00	0.08	0.92	−0.01	3 57	3.60	0.73	67
	FG	1.21	1.57	0.01	0 07	0.92	−0.12	3.55	3 63	0.82	66
4/53–4/79	LS	1.60	2.26	0.43	0.08	0.48	−1 50	3.42	2.40	0 79	53
	RW	1.41	1.73	0.02	0.00	0.98	−0.25	3.47	3 12	0.75	52
	FG	1 06	1 32	0.01	0.00	0.98	−0 16	3 44	3.19	0.85	51
10/79–10/86	LS	1.31	1.79	0.08	0.17	0.76	0.50	3.83	2.72	0.82	15
	RW	1.64	2.46	0.10	0.26	0.64	0.79	3.83	4.54	0.72	15
	FG	1.73	2 23	0.00	0.39	0.61	0.01	3.83	4 81	0 78	15
Fourteen-Month Forecasts											
4/53–10/86	LS	1 58	2.20	0.18	0.00	0.82	−0 93	3.43	2.79	0 66	68
	RW	1.78	2.34	0.00	0.13	0.87	−0 10	3 42	3.49	0.59	66
	FG	1.30	1.76	0.01	0.13	0.86	−0.15	3.42	3.63	0.77	66
4/53–4/79	LS	1.59	2.30	0.43	0.10	0.48	−1.51	3.50	2.41	0.80	53
	RW	1.11	2.19	0.07	0.01	0.92	−0.57	3.50	2.99	0.63	51
	FG	1 17	1.51	0.04	0.00	0.96	−0 32	3 50	3 19	0 82	51
10/79–10/86	LS	1.55	1 84	0.35	0.07	0.58	1.09	3.20	2.37	0.79	15
	RW	2.29	2.78	0.29	0 31	0 40	1 50	3.20	4.25	0.68	15
	FG	1 73	2 45	0.03	0.62	0.35	0.41	3 20	4.81	0.78	15

LS is the Livingston consensus forecast RW is the forecast of a random walk model. FG is the forecast of the Fama-Gibbons interest rate model. Three error measures are given· the average error (AVE), the mean absolute error (MAE), and the root mean square error (RMSE). The standard errors of actual and predicted inflation are denoted s_a and s_p. r^2 is the squared correlation between actual and predicted inflation. N is the number of observations. For U^M, U^R, and U^D, see Equation (2) in the text.

among the three models. With respect to the rationality tests, four of six slopes are significantly positive. For the extrapolative model in rows A.1 and A.2, this suggests that large initial movements, up or down, are expected to partly reverse themselves. However, since $\lambda > \lambda^*$, the expected reversal is excessive. The surveys give too much weight to inflation in the distant past, relative to recent past inflation.

Equations B.1 and B.2 send a similar message. In effect, the tests check for serial correlation between adjacent forecast errors. Expectations are insufficiently adaptive: if the economists paid more attention to recent inflation, and interpreted the prevailing rate as less of a surprise, they would not make the same error repeatedly. For example, for the 8-month forecasts, γ—the weight given to the previous forecast—is 0.60. Yet, from the rationality tests, $(\gamma - \gamma^*)$ equals 0.42, which implies that the "correct" weight is only 0.18.[6]

The results for the regressive model in rows C.1 and C.2 are much weaker. However, for the 1953–1979 period (not reported), $(\delta - \delta^*)$ for 8-month forecasts

[6]A convenient way to interpret the findings is in the context of Muth (1960). Let actual inflation $\pi_{t,k}$ be subject to permanent (η) as well as temporary shocks (μ). Then, the forecast that minimizes the mean square error follows an adaptive model, $\pi^e_{t,k} = \gamma\pi^e_{t-k,k} + (1-\gamma)\pi_{t-k,k}$, where γ is a function of $\sigma^2_\eta/\sigma^2_\mu$. When all shocks are permanent, $\gamma = 0$ and $\pi^e_{t,k} = \pi_{t-k,k}$. When all shocks are temporary, $\gamma = 1$ and $\pi^e_{t,k} = \pi^e_{t-k,k}$. Since $\gamma > \gamma^*$, Table 3 indicates that the experts impute a larger temporary component to CPI movements than they should. Indeed, as Table 2 suggests, much of the time, it would be better to assume that all inflation shocks are permanent and none temporary!

TABLE 3

Predicting Survey Expected Changes in Inflation and Survey Forecast Errors

	Expectation Formation			Rationality Tests				
Regressors	Slope	t_{slope}	Adj. R^2	Intercept	$t_{\text{int.}}$	Slope	t_{slope}	Adj. R^2
Eight-Month Forecasts								
A.1 $\pi_{t-16,8} - \pi_{t-8,8}$	0.470	7.61	0.279	−1.001	−4.42	0.361	3.02	0.152
B.1 $\pi^e_{t-6,8} - \pi_{t-6,6}$	0.604	13.28	0.717	−0.601	−3.20	0.418	3.28	0.206
C.1 $-\pi_{t-8,8}$	0.492	13.99	0.410	−0.472	−1.48	0.131	1.47	0.048
Fourteen-Month Forecasts								
A.2 $\pi_{t-28,14} - \pi_{t-14,14}$	0.308	5.83	0.203	−0.928	−3.04	0.382	3.84	0.197
B.2 $\pi^e_{t-12,14} - \pi_{t-12,12}$	0.516	11.27	0.616	−0.507	−1.67	0.443	3.09	0.179
C.2 $-\pi_{t-14,14}$	0.482	17.74	0.390	−0.799	−2.07	0.031	0.30	−0.012

We study the 1953–1987 period. The regressions with the expected change in inflation as the dependent variable (left-hand side of the table) are estimated using GLS and White's correction for heteroskedasticity. The adjusted R-square coefficients refer to the OLS-regression. We do not report the intercept. The regressions with Livingston survey forecast errors as the dependent variable (right-hand side of the table) are estimated using OLS, with Hansen and Hodrick's correction for serial correlation, White's adjustment for heteroskedasticity, and the Newey and West weighting of covariances. All regressions have 68 observations.

is 0.25 (t-statistic: 2.47). Thus, prior to 1980, the surveys did place too much weight on the past.

C. Discussion

How do our results compare to previous research? Prior studies of the Livingston inflation forecasts yield little agreement on whether the predictions are rational. Two papers argue in favor (Mullineaux (1978) and Caskey (1985)) and five argue against (Pesando (1975), Pearce (1979), Jacobs and Jones (1980), Figlewski and Wachtel (1981), and Sterman (1987)). Brown and Maital (1981) conclude that the forecasts are unbiased but inefficient with respect to monetary growth.

Most often, the authors' strategy is to test for rationality in Muth's sense, i.e., are the expectations equal to mathematical expectations, conditional on the relevant information? However, this concept of rationality may be too narrow. The information necessary to form expectations on the basis of the true economic model may be too costly to obtain (Feige and Pearce (1976)). Also, in a nonstationary environment, rational Bayesian learners may not appear rational in Muth's sense (Friedman (1979)). An important study that takes these shortcomings as a starting point is Caskey (1985). Caskey's empirical work continues to assume that the Livingston economists consider the parameters of the inflation process to be stable. But, since a credible set of initial beliefs combined with Bayesian updating can account for the forecasts, they may be the product of a rational learning process.

Our own orthogonality tests—which offer mild evidence against forecast unbiasedness and stronger evidence against forecast efficiency—have little to say about whether, in a deeper philosophical sense, the predictions are rational. We look for a specific type of "bias." The inertia that we observe is consistent with Fisher's explanation, but it is equally consistent with Bayesian learning. This view gains plausibility from the literature on dynamic inconsistency in monetary policy (for a review, see Alesina (1988)). The Federal Reserve may want to look like a "strong" inflation fighter, yet turn "weak" when the economy slows. Thus, the public can never be certain about the Fed's intentions. Some models (e.g., Backus and Driffill (1985))—which rely on reputation to solve the credibility problem—predict the public's skepticism toward each new attempt to fight inflation. Rational expectations resemble adaptive expectations. This theory seems particularly relevant to the experience of the 1970s and early 1980s.

IV. Interest Rate Forecasts of Inflation

Are the inflation forecasts implicit in interest rates similar to the Livingston survey forecasts? We follow two approaches to address this question. First, we check whether movements in ex post real rates—which contain an inflation forecast error—are predicted by the same variables that predict inflation survey errors. Second, we test whether time-variation in nominal rates mirrors survey forecasts more closely than time-series forecasts.

Our tests start from the Fisher equation, $R_{t,k} = r^e_{t,k} + \pi^e_{t,k}$, which equates the k-month nominal interest rate to the expected real rate plus expected inflation. Subtracting later observed inflation from both sides,

$$r_{t,k} = r^e_{t,k} + (\pi^e_{t,k} - \pi_{t,k}). \tag{7}$$

Equation (7) expresses the ex post real rate as the sum of the ex ante real rate and an inflation forecast error that, under rational expectations, is unpredictable.

If the ex ante real rate were constant (except for white noise), a regression of $r_{t,k}$ on any variable known at time t should yield a coefficient of zero. Since there is virtually no end to the list of variables that one can try, we restrict ourselves to those that, from Section III, we already know to predict survey forecasts errors, i.e., the recent change in inflation, past survey errors, and past inflation. The variables are of interest because they allow for a ready-made interpretation. The regressions remain interesting even if ex ante real rates are not constant. Whereas, in principle, the right-hand side variables may proxy for fundamental determinants of real rates (such as the covariance of consumption with bond returns, as in Benninga and Protopapadakis (1983)), it seems easier to interpret their predictive power as evidence of forecast bias.

The results in Table 4 are based on 6- and 12-month Treasury Bill yields at the end of each June and December since 1953.[7] As before, the t-statistics are corrected for serial correlation and heteroskedasticity. The data suggest that interest rates are biased predictors of inflation. Thus, when inflation is accelerating, the prices of Treasury Bills seem predictably too high while, with falling inflation, prices seem too low.[8] Just as in Table 3, all slope coefficients are positive and many are significant. Chow tests that compare the estimated slopes in Tables 3 and 4 do not allow one to reject that the movements in real rates exactly parallel the patterns

[7]We report the findings for 6- and 12-month yields rather than for 8- and 14-month yields (measured at the end of April and October) because the test is more conservative and because the data are more precise (see footnote 3). However, the results are virtually identical.

[8]Compared to Mishkin (1981), the R^2s are surprisingly high. Mishkin finds that real rates since the 1930s show little cyclical variation. Except for a negative correlation with past inflation, he cannot find a link with other macroeconomic variables (such as growth in M1).

in the survey forecast errors.[9] But if, for example, a third set of variables—other than inflation forecasts—were responsible for the findings in Table 4, there is little reason to expect this parallel movement.

TABLE 4

Predicting Ex Post Real Interest Rates

Regressors	α	t_α	β	t_β	D-W	Adj. R^2
Six-Month Real Rate						
A.1 $\pi_{t-16,8} - \pi_{t-8,8}$	1.522	4.00	0.664	3.10	0 95	0.219
B.1 $\pi^e_{t-6,8} - \pi_{t-6,6}$	2.299	5.17	0.778	4.72	1 22	0.315
C.1 $-\pi_{t-8,8}$	2.254	4.85	0.177	1 45	0.68	0.034
Twelve-Month Real Rate						
A.2 $\pi_{t-28,14} - \pi_{t-14,14}$	1.741	3.58	0.549	2.57	0.48	0.190
B.2 $\pi^e_{t-12,14} - \pi_{t-12,12}$	2.404	4.12	0.698	3.48	0.51	0.211
C.2 $-\pi_{t-14,14}$	1.812	3.28	0.018	0.11	0 34	−0.015

We study the 1953–1987 period. The ex post real rates are computed from Treasury Bill yields observed on the last trading day of the months of June and December. Time t corresponds to April or October. The regressors are the same as in Table 3. The regressions are OLS with the t-statistics corrected for serial correlation and heteroskedasticity (using the methods of Hansen and Hodrick, White, and Newey and West). D-W is the Durbin-Watson coefficient. All regressions have 68 observations.

A different method permitting us to judge whether the market expectation behaves like the survey data is to regress nominal interest rates on various inflation forecasts. We compare the Livingston expectations (LS) with the "superior" random walk (RW) and Fama-Gibbons (FG) forecasts studied earlier. In bivariate OLS regressions with, e.g., 8-month yields and inflation forecasts between 1953 and 1986, the adjusted R^2 is larger for LS (0.76) than for RW (0.59) or FG (0.53). In a regression with all three sets of forecasts, the coefficient for LS is 1.16 (t-statistic: 6.71). The t-statistics for the other forecasts are 0.05 (RW) and −1.03 (FG). The R^2 remains at 0.76. Our results differ from those of Pearce (1979), who finds that, for the less volatile 1959–1975 period, interest rates are more explained by constructed "rational" forecasts of inflation than by LS.

V. Term Structure Forecasts of Inflation

Past studies have repeatedly rejected the joint hypothesis of the expectations theory of the term structure and rational expectations. There are two competing but not mutually exclusive explanations for this rejection. The first is that term premia rationally vary through time. While intertemporal asset pricing theories easily accommodate time-varying risk premia, it remains a challenge to account for actual bond price movements, especially negative expected returns. The second explanation involves some form of nonrational expectations. For example, Campbell and Shiller (1984) and Mankiw and Summers (1984) test and reject the

[9]The Chow tests are done for the June 1953–June 1979 as well as the December 1979–December 1986 periods. We compare the rationality tests in Table 3 to the equivalent regressions in Table 4 (that share the same predictor variable). However, the regressions with the real rate as the dependent variable are for 8- and 14-month yields. For none of the six sets of two equations do the Chow tests reject the equality of the slopes. We also compute Wald statistics that do not assume equal error variances in the two samples. Of 12 test statistics, none are significant at the 5-percent level.

hypothesis that long rates "overreact" to short rates. Our empirical tests below use survey as well as market data. As a result, they allow us to distinguish between changing forecasts and changing risk premia.

A. Inflation and the Slope of the Term Structure

Let the term premium, ϕ_t, be defined as $E_t(H_t - h_t)$, where E_tH_t and E_th_t represent, respectively, the expected holding returns (at t) on long- and short-term bonds (or bills). Then, the actually realized excess holding return equals the sum of the premium and a forecast error ν_t so that $H_t - h_t = \phi_t + \nu_t$. The expectations theory states that ϕ_t is constant through time. Under rational expectations, ex post variation in $(H_t - h_t)$ is unforecastable on the basis of information known at t.

Starting with Shiller (1979), various studies have shown that $(H_t - h_t)$ is erratic but not totally unpredictable. Excess returns on long bonds are positively related to the yield spread between long and short rates (also called "the slope"). In Panel A of Table 5, we confirm these earlier findings for 1953–1986, using only the semi-annual Livingston survey dates. The yield spread is defined as the difference between the yields on a 12- and a 6-month bill. The (ex post) term premia are the annualized 6-month returns on bonds (with maturities of 1 to 5 and 10 years) *over and above* the yield on a 6-month bill. To calculate the bond returns, we follow previous work (e.g., Froot (1989)) and use the linearized approximations of Shiller et al. (1983).[10]

The expectations theory and investor rationality imply that, when long-term rates are above short-term rates, long rates ought to rise. Correcting for the term premia, the resulting capital loss equates expected holding period returns across assets. Similarly, when long rates are below short rates, long rates ought to fall. However, in practice, the opposite tends to happen. As seen in Table 5 (Panel A), the more the term structure is upward-sloping, the more long-duration instruments outperform bills.

Why does the slope have predictive power? Following our evidence on inflation forecast errors, the predictive power of the yield spread may reflect a delayed reaction of the bond market to news about the CPI. For example, after the double-digit inflation of the late 1970s, the public may well have been skeptical that the much lower inflation of the early 1980s would continue. Perhaps, for too much time, long rates remained high relative to short rates and, only when the lower inflation continued, did investors drive long-term interest rates down. Exactly the reverse story could be told for the mid-1970s. Thus, the observed "underreaction" of long rates to short rates may reflect the relative sluggishness with which long-term inflation expectations adjust to news about the CPI, compared to short-term expectations.[11]

[10]Since, under the null hypothesis, the slope and the forecast error are uncorrelated, the regressions are properly estimated using OLS. The t-statistics are corrected for heteroskedasticity. The flavor of the results does not change if the yield spread equals the difference between the yields on a 10-year bond and a six-month bill.

[11]Remember, however, that the public's skepticism may be rational. Our theory is consistent with work by Froot (1989), which is based on survey expectations of interest rates (rather than inflation) collected from the investment newsletter, *Reporting on Governments*. Froot argues that the "perverse

TABLE 5

Term Premia as Predicted by the Slope of the Term Structure, Survey Expected Changes in Inflation, and Past Survey Forecast Errors

Maturity (Years)	α	t_α	β	t_β	SEE	Adj. R^2
Panel A Regressor: $R_{t,12} - R_{t,6}$						
1	−0.138	−0.64	1.438	2.13	1.45	0.088
2	−0.538	−1.04	3.903	2.28	3.66	0.102
3	−0.994	−1.33	6.126	2.49	5.78	0.101
4	−1.527	−1.69	8.754	2.95	7.03	0.139
5	−2.019	−1.92	10.947	3.24	8.42	0.151
10	−4.345	−2.54	20.142	3.87	14.90	0.162
Panel B. Regressor: $\pi^e_{t,14} - \pi^e_{t,8}$						
1	−0.022	−0.09	1 143	1 11	1.51	0.019
2	−0.260	−0.47	3.360	1.42	3.81	0.031
3	−0.587	−0.73	5.485	1 52	5.99	0.034
4	−0.963	−0 95	7.969	1.82	7.38	0.052
5	−1.359	−1.15	10.294	1.98	8.85	0.062
10	−3.196	−1.71	19.404	2.35	15.68	0.072
Panel C. Regressor: $\pi^e_{t-6,8} - \pi_{t-6,6}$						
1	0.311	1.38	0.164	1.26	1.49	0.038
2	0.743	1 25	0.504	1.62	3.74	0.063
3	1 133	1.16	0.902	1.79	5.83	0.086
4	1.394	1 19	1.176	1.96	7.20	0.096
5	1.653	1.16	1 489	2.07	8 63	0 108
10	2.427	0.91	2.755	2 35	15 29	0.118

The dependent variable is the excess holding return on an instrument of stated maturity over-and-above the yield on a 6-month bill. The yield spread is the difference between the yields on a 12- and 6-month bill. It is observed on the last trading day of every Livingston survey month between April 1953 and April 1986. The regressor in panel B is the Livingston expected 14-month inflation minus the expected 8-month inflation. The regressor in Panel C is the most recent Livingston survey 8-month forecast error. The regressions are OLS with the t-statistics corrected for heteroskedasticity (White (1980)). SEE refers to the standard error of the estimate. All regressions have 67 observations

In contrast, under rational expectations, the behavior of the yield spread implies that this variable moves importantly with changing risk premia (which are part of ex ante real rates) and possibly more so than with changing forecasts of inflation. In other words, if we decompose the yield spread $(k > j), R_{t,k} - R_{t,j} = (r^e_{t,k} - r^e_{t,j}) + (\pi^e_{t,k} - \pi^e_{t,j})$, variation in the first term on the right-hand side obscures variation in the second.

The survey data argue against this presumption in a variety of ways. For example, $(R_{t,14} - R_{t,8})$ clearly moves with the Livingston expectations of future inflation, $(\pi^e_{t,14} - \pi^e_{t,8})$. For 1953–1986, the correlation between these series is 0.60. Table 6 (Panel A) makes the same point. After 1979, one cannot reject the hypothesis that the variables move in perfect parallel, i.e., $(\alpha, \beta) = (0, 1)$.[12] In addition, Table 7 shows that $(R_{t,14} - R_{t,8})$ and $(\pi^e_{t,14} - \pi^e_{t,8})$ respond similarly to the variables that are of special interest to us, i.e., the recent change in inflation, past

predictions of the spread reflect investors' failure to raise sufficiently their expectations of future long rates when the short rate rises."

[12]In an earlier version of this paper, we obtain a similar result for long-term survey forecasts of inflation collected by *Drexel Burnham Lambert*. For 1980–1987, we compare the monthly difference between 10- and 5-year inflation forecasts with the comparable slope of the term structure, i.e., 10-year yields minus 5-year yields (taken from Shiller and McCulloch (1990)). The correlation between both series is 0.78.

survey errors, and past inflation. Notice the surprising closeness of the estimates. Chow tests based on the equivalent regressions in Panels A and B never allow us to reject their equality.

TABLE 6

Survey Expected Changes in Inflation, Actual Inflation, and Survey Forecast Errors as Predicted by the Slope of the Term Structure

Periods	Dependent Variable	α	t_α	β	t_β	D-W	Adj. R^2
A: 4/53–10/86	$\pi^e_{t,14} - \pi^e_{t,8}$	0.045	1.16	0.689	4 84	1.53	0 357
10/79–10/86		−0.016	−0.29	1.052	8.27	2.23	0.778
B: 4/53–10/86	$\pi_{t,14} - \pi_{t,8}$	−0.140	−1.11	1.084	2.14	2.15	0.056
10/79–10/86		−0.732	−3.49	1.659	3.02	2.32	0.213
C: 4/53–10/86	$\pi^e_{t+8,6} - \pi_{t+8,6}$	−1.285	−2.48	3.725	2.15	0.90	0.087
10/79–10/86		1.427	2.44	1.962	2.09	1.65	0.042

The regressions are OLS with the t-statistics corrected as in Table 1. D-W is the Durbin-Watson coefficient. The regressor is the slope of the term structure, defined as the difference between the yields on a 14-month zero-coupon instrument and on an 8-month bill. Both yields are observed on the last trading day of each April and October between 1953 and 1986. The number of observations is 68 for the April 1953–October 1986 period and 15 for the period after 1979.

In sum, we find that the yield spread varies as survey forecasts of inflation vary. This is consistent with the expectations theory. However, the results do not speak to the rationality of the forecasts. If, as suggested by Section III, $(\pi^e_{t,14} - \pi^e_{t,8})$ is biased, so may be the slope of the term structure.

At least three empirical predictions follow from this hypothesis. First, the ex post forecast error in $\pi^e_{t+8,6}$ is predictable from the yield spread, even as the forecast is being made. When the spread is positive, economists' forecasts of inflation tend to be too high. When it is negative, they are too low. See Table 6 (Panel C). On the one hand, this finding questions the rationality of the surveys. On the other, it suggests that the yield spread contains the same error.[13]

Second, $(\pi^e_{t,14} - \pi^e_{t,8})$ predicts ex post term premia. See Table 5 (Panel B). If the slope of the term structure predicts bond excess returns because it proxies for time-varying risk, there is no reason why, at time t, the survey expected change in inflation between $t + 8$ and $t + 14$ would also predict excess returns. Presumably, bond prices rationally incorporate such expected changes at t and no further predictable price movements should occur between t and $t + 6$. Nevertheless, the parameter estimates in Panels A and B of Table 5 are very similar. This suggests that the predictive power of the yield spread derives, at least in part, from $(\pi^e_{t,14} - \pi^e_{t,8})$. Apparently, when inflation accelerates, $\pi^e_{t,14}$ is low relative to $\pi^e_{t,8}$, long rates are low relative to short rates, and long-maturity instruments are poor investments. Conversely, long bonds are excellent investments when inflation slows down.

The third and perhaps most noteworthy implication is that past survey errors—which get repeated and predict the spread—also predict ex post term premia. Table 5 (Panel C) indeed confirms that bonds earn higher returns when the economists

[13]The yield spread may contain systematic error and yet predict the path of future inflation. See Table 6 (Panel B). This is consistent with Fama (1988) and Mishkin (1988).

overpredicted inflation in the prior survey, i.e., when $\pi^e_{t-6,8} > \pi_{t-6,6}$.[14] This result suggests potentially profitable bond trading strategies. A thought experiment, inspired by Mankiw (1986), makes this clear. Suppose that an investor decides to buy 10-year bonds worth \$1 million with funds that are borrowed at the 6-month rate when the past (annualized) forecast error is two standard deviations above its sample mean for the 1953–1986 period (−1.06 percent). This implies that, on an annual basis, the Livingston economists overestimated inflation by 3.21 percent during the last 6 months. Then, from Table 5 (Panel C), the expected return on the strategy is 11.28 percent before transaction costs (or \$56,400 for the next 6 months) with a standard error of 15.29 percent (\$76,450). If returns are normally distributed, there is a 77 percent chance that the strategy is profitable.

TABLE 7

Predicting the Slope of the Term Structure and Survey Expected Changes in Inflation

Regressors	α	t_α	β	t_β	D-W	Adj. R^2
Panel A. Dependent Variable: $R_{t,14} - R_{t,8}$						
$\pi_{t-28,14} - \pi_{t-14,14}$	0.142	7.28	0.037	3.87	1.20	0 152
$\pi^e_{t-12,14} - \pi_{t-12,12}$	0.201	10.76	0.061	5.45	1 21	0.257
$-\pi_{t-14,14}$	0.228	9.73	0.020	2.90	0.81	0.042
Panel B. Dependent Variable: $\pi^e_{t,14} - \pi^e_{t,8}$						
$\pi_{t-28,14} - \pi_{t-14,14}$	0.142	5.59	0.035	2.48	1.24	0.094
$\pi^e_{t-12,14} - \pi_{t-12,12}$	0.205	8.40	0.066	5.30	1.48	0.239
$-\pi_{t-14,14}$	0.244	5.36	0.024	2.64	1.03	0.057

The regressions are GLS with the t-statistics corrected for heteroskedasticity (White (1980)). The R-squares measure the fit of the structural part of the model after transforming for the autocorrelation. D-W indicates the Durbin-Watson coefficient for the OLS regression. The slope of the term structure is the difference between the yields on a 14- and an 8-month bill. Both yields are observed on the last trading day of each April and October between 1953 and 1986. All regressions have 68 observations.

From the above discussion, we conclude that movements in term premia are partly driven by inflation forecast errors. Interestingly, however, the yield spread does not lose *all* its predictive power if past inflation forecast errors are taken into account. For example, for 1953–1986, a regression with $(R_{t,14} - R_{t,8})$ and $(\pi^e_{t-6,8} - \pi_{t-6,6})$ on the right-hand side and with excess returns on 10-year bonds as the dependent variable has an R^2 of 0.22, considerably higher than do the bivariate regressions in Table 5. The estimated coefficients (beta weights) are 16.87 (0.35) for the spread (t-statistic: 3.12) and 2.10 (0.28) for the past forecast error (t-statistic: 2.46). Both variables retain about the same level of statistical significance if the regression includes survey and other conventional measures of inflation and output uncertainty, factors that we consider next.

B. Term Premia, Risk, and the Business Cycle: Survey Evidence

Even though term premia systematically vary through time, few researchers have found any ex ante observable economic variables (besides the term structure

[14]The results are similar for other measures of past forecast errors, e.g., $(\pi^e_{t-12,14} - \pi_{t-12,12})$. However, once $(\pi^e_{t-6,8} - \pi_{t-6,8})$ is included in the regressions, other forecast errors or further lagged errors add no explanatory power.

spread) that reliably predict their movements.[15] We now discuss the relationship between term premia, the yield spread, and survey measures of business cycle risk.

Fama (1976a), (1976b) suggests that the premia reflect uncertainty about future inflation. In the spirit of this earlier work, we consider five measures: 1) the standard deviation of the 8-month Livingston forecasts; 2) the standard deviation of the 14-month forecasts; 3) the standard deviation of the 6-month forecasts, starting 8 months after the survey date; 4) the mean absolute monthly change in inflation for the 24 months centered around the Livingston survey month; and 5) the measure proposed by Fama, i.e., the mean absolute change in the yield of a 1-month Treasury Bill for 24 months (centered around the Livingston survey month). As Fama concedes, there is an obvious problem with all of these measures. Contrary to modern portfolio theory, they assume that risk and return uncertainty are equivalent.[16]

Panel A (B) in Table 8 reports regressions with the excess returns on a 12-month bill (10-year bond)—i.e., the (annualized) 6-month return minus the yield on a 6-month bill—as the dependent variable. With risk aversion, the premia should be positively related to inflation uncertainty. However, none of the coefficient estimates (models #1–#5) are significantly different from zero. The results stand in contrast to Fama's (1976a), (1976b) findings at the short end of the maturity spectrum and to the work of Engle, Lilien, and Robins (1987).

Possibly, our estimates are biased toward zero because of measurement error. Therefore, we also consider regressions with the yield spread—the yield on a 12-month bill (or a 10-year bond) minus the yield on a 6-month bill—as the predictor variable and the five inflation uncertainty proxies as regressands. If the measurement error is uncorrelated with the yield spread, there should be no bias. All the estimated slopes (not reported) are negative; five out of ten are significant but the R^2s are very low. Thus, contrary to intuition, if the yield spread is a risk proxy, it would appear that long-term instruments are less risky when inflation uncertainty is high.

The Livingston survey respondents also provide 8- and 14-month forecasts of industrial production (IP). When they are compared with base numbers collected from the *Federal Reserve Bulletin,* we can use the cross-sectional standard deviations of the forecasted percentage change in IP as measures of output uncertainty. Are there links between output uncertainty and term premia? Once again, the results in Table 7 (models #6 and #7) are unfavorable. Regressions with the yield spread as the predictor variable (not reported) are equally disappointing. However, consistent with Hasbrouck (1984), the output uncertainty measures show a positive sign when related to a stock market risk premium, e.g., the (annualized) 6-month

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[15] For example, Friedman (1980) finds no evidence that term premia vary with indicators of economic policy. Mankiw (1986) checks whether bond price volatility, consumption covariability, or changes in asset supply explain the movements in the yield spread. None do. Consistent with consumption-based asset pricing models, Harvey (1989) finds that, since 1953, the real term structure has predictive power for per capita growth in real consumption. However, the R^2s are small and they depend a lot on the results of one subperiod, 1972–1987.

[16] Also implicit in measures 1–3 is the assumption that the standard deviation of a set of predictions made by different individuals (predictive dissent) is an acceptable proxy for the uncertainty felt by the representative investor. Zarnowitz and Lambros (1987) study NBER-ASA probabilistic inflation forecasts and find that lack of consensus and uncertainty are usually positively correlated.

TABLE 8

Term Premia and Uncertainty about Inflation and Industrial Production

Model	α	t_α	β	t_β	D-W	Adj. R^2
Panel A. Dependent Variable: Excess Return on 12-Month Bill						
1	−0.022	−0.05	0.119	0.34	2.21	−0.014
2	0.053	0.10	0.068	0.15	2.22	−0.015
3	0.654	1.41	−0.316	−1.24	2.17	0.008
4	0.210	0.65	−0.032	−0.13	2.22	−0.015
5	0.301	0.99	−0.428	−0.73	2.22	−0.007
6	0.282	0.44	−0.044	−0.24	2.19	−0.015
7	−0.100	−0.18	0.085	0.19	2.22	−0.012
Panel B. Dependent Variable: Excess Return on 10-Year Bond						
1	5.157	1.01	−4.514	−1.22	1.77	0.007
2	3.815	0.67	−3.949	−0.82	1.75	−0.005
3	6.336	1.29	−4.152	−1.54	1.73	0.021
4	2.138	0.31	−1.078	−0.41	1.76	−0.013
5	2.840	0.88	−8.458	−1.36	1.79	0.013
6	3.039	0.45	−1.080	−0.54	1.74	−0.011
7	−3.310	−0.56	1.010	0.50	1.80	−0.011

The regressions are OLS. D-W refers to the Durbin-Watson statistic. All regressions have 68 observations. Each corresponds to the last trading day of every Livingston survey month between June 1953 and June 1986. The returns, over and above the yield on a 6-month bill, are measured between the last trading day of June (December) and the following December (June). Inflation uncertainty is measured by: the standard deviation of Livingston 8-month inflation forecasts (model 1); the standard deviation of 14-month forecasts (model 2); the standard deviation of 6-month forecasts, 8 months after the survey date (model 3); the mean absolute monthly change in inflation for the 24 months centered around the Livingston survey month (model 4); the mean absolute monthly change in the 1-month Treasury bill yield during 24 months centered around the survey month (model 5). Output uncertainty is measured by: the standard deviation of Livingston 8-month expected percentage changes in industrial production (model 6); the standard deviation of 14-month expected percentage changes in industrial production (model 7). Returns as well as inflation and output forecasts are annualized.

excess returns on the Standard & Poor's Index for Industrial Companies. As with bonds, the stock market risk premium is not explained by inflation uncertainty.

VI. Conclusion

The approach of this study is to combine inflation survey with bond market data for the 1953–1987 period. This research strategy allows us to distinguish changing inflation expectations from other factors that influence interest rates, for example, liquidity, risk, and tax effects. Apparently, past inflation forecast errors predict future forecast errors in surveys, predict future movements in real rates, and predict term premia on U.S. Government Bonds. Even though the inflation forecasts fail standard rationality tests, movements in the yield spread strongly reflect their variation through time.[17]

The evidence is consistent with a Fisher effect and with the expectations theory of the term structure. It also suggests that past experience weighs too heavily in how investors intuitively judge future levels of inflation and interest rates. However, the inertia in expectations may be rational if we consider the costs and benefits of more accurate forecasts and/or possible regime changes (with the

[17]It would be interesting to test the robustness of this result with data for foreign countries. Studies of inflation survey data for Britain (Batchelor and Dua (1987)) and Sweden (Jonung and Laidler (1988)) find, as we do, that the forecasts seem insufficiently adaptive. However, the link with interest rates and the term structure has not been investigated.

implication that rational expectations resemble adaptive expectations). From this perspective, it seems that investors need considerable time to appreciate the real and monetary shocks underlying movements in consumer price levels.

References

Alesina, A. "Macroeconomics and Politics." In *Macroeconomics Annual 1988,* S. Fischer, ed. Cambridge, MA: MIT Press (1988).

Backus, D., and J. Driffill. "Inflation and Reputation." *American Economic Review,* 75 (June 1985), 530–538.

Barsky, R. B. "The Fisher Hypothesis and the Forecastability and Persistence of Inflation." *Journal of Monetary Economics,* 19 (Jan. 1987), 3–24.

Barsky, R. B., and J. B. De Long. "Forecasting Pre-World War I Inflation: The Fisher Effect Revisited." Working Paper, NBER #2784 (1988).

Batchelor, R. A., and P. Dua. "The Accuracy and Rationality of UK Inflation Expectations: Some Quantitative Evidence." *Applied Economics,* 19 (June 1987), 819–828.

Benninga, S., and A. Protopapadakis. "Real and Nominal Interest Rates under Uncertainty: The Fisher Theorem and the Term Structure." *Journal of Political Economy,* 41 (Oct. 1983), 856–867.

Bierwag, G. O., and M. A. Grove. "Aggregate Koyck Functions." *Econometrica,* 34 (Oct. 1966), 828–832.

Brown, B. W., and S. Maital. "What Do Economists Know? An Empirical Study of Expert Expectations." *Econometrica,* 49 (March 1981), 491–504.

Campbell, J. Y., and R. J. Shiller. "A Simple Account of the Behavior of Long-Term Interest Rates." *American Economic Review,* 74 (May 1984), 44–48.

Carlson, J. G. "A Study of Price Forecasts." *Journal of Economic and Social Measurement,* 6 (1977), 27–56.

Caskey, J. "Modeling the Formation of Price Expectations: A Bayesian Approach." *American Economic Review,* 75 (Sept. 1985), 768–776.

Clarida, R. H., and B. M. Friedman. "The Behavior of U.S. Short-Term Interest Rates since October 1979." *Journal of Finance,* 39 (July 1984), 671–682.

Cox, J. C.; J. E. Ingersoll, Jr.; and S. A. Ross, "A Theory of the Term Structure of Interest Rates." *Econometrica,* 35 (March 1985), 385–408.

Darby, M. R. "The Financial and Tax Effects of Monetary Policy on Interest Rates." *Economic Inquiry,* 14 (June 1975), 259–269.

Engle, R. F.; D. M. Lilien; and R. P. Robins. "Estimating Time-Varying Risk Premia in the Term Structure: The ARCH-M Model." *Econometrica,* 55 (May 1987), 391–407.

Fama, E. F. "Forward Rates as Predictors of Future Spot Rates." *Journal of Financial Economics,* 3 (Oct. 1976a), 361–377.

__________. "Inflation Uncertainty and Expected Returns on Treasury Bills." *Journal of Political Economy,* 84 (June 1976b), 427–448.

__________. "Term Structure Forecasts of Interest Rates, Inflation, and Real Returns." Working Paper, Graduate School of Business, Univ. of Chicago (1988).

Fama, E. F., and M. R. Gibbons. "A Comparison of Inflation Forecasts." *Journal of Monetary Economics,* 13 (May 1984), 327–348.

Feige, E. L., and D. K. Pearce. "Economically Rational Price Expectations: Are Innovations in the Rate of Inflation Independent of Innovations of Monetary and Fiscal Policy?" *Journal of Political Economy,* 84 (June 1976), 499–522.

Figlewski, S., and P. Wachtel. "The Formation of Inflationary Expectations." *Review of Economics and Statistics,* 43 (Feb. 1981), 1–10.

Fisher, I. *The Money Illusion.* New York: Adelphi Co. (1928).

__________.*The Theory of Interest.* Clifton, NJ: A. M. Kelley (1965).

Frankel, J. A., and K. A. Froot. "Using Survey Data to Test Standard Propositions Regarding Exchange Rate Expectations." *American Economic Review,* 77 (March 1987), 134–153.

Friedman, B. M. "Optimal Expectations and the Extreme Information Assumptions of Rational Expectations Macromodels." *Journal of Monetary Economics,* 5 (Jan. 1979), 23–41.

__________. "Interest Rate Expectations versus Forward Rates: Evidence from an Expectations Survey." *Journal of Finance,* 34 (Sept. 1979), 965–973.

Friedman, M., and A. J. Schwartz. *Monetary Trends in the United States and the United Kingdom: Their Relation to Income, Prices, and Interest Rates,* 1867–1975. Chicago: National Bureau of Economic Research (1982).

Froot, K. A. "New Hope for the Expectations Hypothesis of the Term Structure of Interest Rates." *Journal of Finance,* 44 (June 1989), 283–305.

Gibson, W. E. "Interest Rates and Inflationary Expectations: New Evidence." *American Economic Review,* 62 (Dec. 1972), 854–865.

Hafer, R. W., and S. E. Hein. "On the Accuracy of Time-Series, Interest Rate and Survey Forecasts of Inflation." *Journal of Business,* 58 (Oct. 1985), 377–398.

Hansen, L. P., and R. J. Hodrick. "Forward Exchange Rates and Optimal Predictors of Future Spot Rates: An Econometric Analysis." *Journal of Political Economy,* 88 (Oct. 1980), 829–853.

Hasbrouck, J. "Stock Returns, Inflation, and Economic Activity: The Survey Evidence." *Journal of Finance,* 39 (Dec. 1984), 1293–1310.

Harvey, C. "The Real Term Structure and Consumption Growth." *Journal of Financial Economics,* 22 (Dec. 1988), 305–333.

Huizinga, J., and F. S. Mishkin. "Inflation and Real Interest Rates on Assets with Different Risk Characteristics." *Journal of Finance,* 39 (July 1984), 699–714.

__________. "Monetary Policy Regime Shifts and the Unusual Behavior of Real Interest Rates." Working Paper, NBER #1678 (1984).

Jacobs, R. L., and R. A. Jones. "Price Expectations in the United States: 1947–1975." *American Economic Review,* 70 (June 1980), 269–277.

Jonung, L., and D. Laidler, "Are Perceptions of Inflation Rational? Some Evidence from Sweden." *American Economic Review,* 78 (Dec. 1988), 1080–1087.

Keane, M. P., and D. E. Runkle. "Testing the Rationality of Price Forecasts: New Evidence from Panel Data." *American Economic Review,* 80 (Sept. 1990), 714–735.

Macaulay, F. A. *The Movements of Interest Rates, Bond Yields and Stock Prices in the United States Since 1865.* New York: National Bureau of Economic Research (1938).

Mankiw, N. G. "The Term Structure of Interest Rates Revisited." *Brookings Papers on Economic Activity* (1, 1986), 61–96.

Mankiw, N. G., and L. H. Summers. "Do Long-Term Interest Rates Overreact to Short-Term Interest Rates?" *Brookings Papers on Economic Activity* (1, 1984), 223–242.

Mishkin, F. S. "Are Market Forecasts Rational?" *American Economic Review,* 71 (June 1981), 295–306.

__________. "The Real Interest Rate: An Empirical Investigation." Working Paper, NBER #622 (1981).

__________. "What Does the Term Structure Tell Us about Future Inflation?" Working Paper, NBER #2626 (1988).

Mullineaux, D. J. "On Testing for Rationality: Another Look at the Livingston Price Expectations Data." *Journal of Political Economy,* 86 (April 1978), 329–336.

Muth, J. F. "Optimal Properties of Exponentially Weighted Forecasts." *Journal of the American Statistical Association,* 55 (June 1960), 299–306.

Newey, W. K., and K. D. West. "A Simple, Positive Semi-Definite, Heteroscedasticity and Autocorrelation Consistent Covariance Matrix." *Econometrica,* 55 (May 1987), 703–708.

Pearce, D. K. "Comparing Survey and Rational Measures of Expected Inflation." *Journal of Money, Credit, and Banking,* 11 (Nov. 1979), 447–456.

Peek, J., and J. A. Wilcox. "The Postwar Stability of the Fisher Effect." *Journal of Finance,* 38 (Sept 1983), 1111–1124.

Pesando, J. E. "A Note on the Rationality of the Livingston Price Expectations." *Journal of Political Economy,* 83 (Aug. 1975), 849–858.

Shiller, R. J. "The Volatility of Long Term Interest Rates and Expectations Models of the Term Structure." *Journal of Political Economy,* 87 (Dec. 1979), 1190–1219.

Shiller, R. J.; J. Y. Campbell; and K. L. Schoenholtz. "Forward Rates and Future Policy: Interpreting the Term Structure of Interest Rates." *Brookings Papers on Economic Activity* (1, 1983), 173–217.

Shiller, R. J., and J. H. McCulloch. "The Term Structure of Interest Rates." In *Handbook of Monetary Economics,* B. M. Friedman and F. H. Hahn, eds. Amsterdam: North-Holland Publ. (1990).

Sterman, J. D. "Expectation Formation in Behavioral Simulation Models." *Behavioral Science,* 32 (July 1987), 190–211.

Summers, L. H. "The Nonadjustment of Nominal Interest Rates: A Study of the Fisher Effect." In *Macroeconomics, Prices, and Quantities,* J. Tobin, ed. Washington, D.C.: The Brookings Institution (1983).

Theil, H. *Applied Economic Forecasting.* Amsterdam: North-Holland Publ. (1966).

White, H. "A Heteroscedasticity-Consistent Matrix Estimator and Direct Test for Heteroscedasticity." *Econometrica,* 48 (May 1980), 817–838.

Zarnowitz, V., and L. A. Lambros. "Consensus and Uncertainty in Economic Prediction." *Journal of Political Economy,* 95 (June 1987), 591–621.