Yield Curve Forecasts of Inflation: A Cautionary Tale

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A commonplace of financial market analysis is the dependence of interest rates on inflation and, further, the dependence of the term structure of interest rates (also called the yield curve) on the expected future path of inflation. Thus, for example, long-term interest rates that are unusually high relative to short-term interest rates are seen to indicate that the market expects increasing inflation in the future. News of strong economic growth may tend to increase long-term rates more than short-term rates, and we read that the market fears such growth will lead to increasing inflation in the future.

Given that the term structure reacts to inflation expectations, does it do so in a reasonable manner? Does the term structure embody inflation forecasts that bear a sensible relationship to the inflation that in fact occurs? This article will review the theoretical link between the term structure and inflation expectations, and then provide empirical evidence on the link in light of the theory.

The theory of the term structure of interest rates has received intensive scrutiny (Shiller 1990 provides a survey), as has the relationship between interest rates and inflation. However, relatively little work has been done linking the term structure to future changes in inflation, which is the focus of this article. The most closely related recent work is that by Fama (1990) and by Mishkin (1990), who run simple regressions of inflation changes on yield spreads. Regressions such as theirs will be shown to arise as a special case in the model considered here.

The link between the term structure and inflation is indirect. The theory of the term structure says only that the term structure should reflect expectations of future *interest rates*. The term structure should be useful in forecasting changes in *inflation* only if it is useful in forecasting changes in interest rates *and* changes in interest rates are, to a substantial extent, driven by changes in inflation. As shown below, proper accounting for the indirect nature of this link has important implications for interpreting the data. In particular, to the extent that the yield curve has no power to forecast changes in interest rates, it *cannot* have power to forecast changes in inflation within the standard theoretical mechanism.

Rather than give results for many different maturities, this study looks at only one pair. The relationship examined is that between one- and two-year interest rates and the one-year-ahead change in the one-year inflation rate. For example, the one- and two-year interest rates at the end of 1993 should embody expectations of the change in inflation from 1994 to 1995.

This horizon is chosen as a balance between practical and econometric considerations. Horizons

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shorter than one year are of limited relevance to policymakers, but examination of longer horizons is handicapped by data limitations. Reliable data are available only since World War II. That seemingly lengthy span contains only nine independent five-year periods, however, preventing reliable inference about the relationship between five-year interest rates and five-year inflation rates. Longer data series of lesser quality are available, and prewar data will be used in this study for comparing results across periods, an exercise that asks less of the data than using them to obtain results for long maturities.

This study finds little evidence of a link between the term structure and future inflation at this horizon. Regressions that control for expected changes in the real rate of interest find no statistically significant evidence of such a link for any time period examined. This evidence is consistent with previous studies, which find that the term structure does not predict changes in interest rates at this horizon.

I. Theory of the Term Structure and Inflation

The theoretical relationship between the term structure and inflation presented here combines the Fisher equation, relating nominal interest rates, real interest rates, and inflation, with the expectations theory of the term structure. The latter gives a relation between long-maturity interest rates and the expected path of short rates.

The box summarizes the relationships between the yield curve, spot rates, and forward rates, and the implications of the expectations theory of the term structure for those relationships. In the notation used there, let i_t refer to the nominal interest rate (spot rate) at the end of year t of the one-year security maturing at the end of year $t + 1$. (Throughout this article, time subscripts will refer to the *date of observation* of the given variable.) Let f_t be the one-year-ahead, one-year forward rate implicit in the term structure observed at the end of year t. As explained in the box, the forward rate is the interest rate that can be locked in, in advance, by appropriate purchase and sale of securities of different maturities. The expectations theory maintains that the spread (difference) between the forward and the spot rates should equal the expected change in the spot rate plus a term premium. Mathematically,

$$(f_t - i_t) = E_t(\Delta i_{i+1}) + \theta \qquad (1)$$

where $E_t(.)$ represents expectations as of the end of year t and θ is a term premium, which the expectations theory assumes is constant over time.[1]

Equation (1) is often called the "forward unbiasedness condition," because it implies that the forward rate provides an unbiased forecast of the future spot rate. As expressed by equation (1), the expectations theory directly implies that the term structure (as reflected in the forward–spot spread $f - i$) forecasts changes in *interest rates*, not changes in *inflation*. To bring inflation into the analysis, let π_{t+1} be the rate of price inflation from the end of year t to the end of year $t + 1$ (recalling that the subscripts denote the date of observation). Let r_{t+1} denote the ex post real rate of interest for the same period. The ex post real rate is simply the rate of interest earned over a period in excess of actual inflation over that period,[2] and therefore:

$$r_{t+1} \equiv i_t - \pi_{t+1} \qquad (2)$$

which implies:

$$E_t(\Delta i_{t+1}) = E_t(\Delta \pi_{t+2}) + E_t(\Delta r_{t+2}). \qquad (3)$$

[1] The term premium is permitted to vary across maturities, but not over time. Alternative theories that drop this assumption are discussed in Section IV below.

[2] Tax considerations are ignored. However, some of the data used below have been adjusted for differential tax treatment of the underlying securities.

The Expectations Theory of the Term Structure of Interest Rates

The expectations theory of the term structure is the benchmark model in economics and finance of the relationship between interest rates of differing maturities. An abbreviated development is presented here, examining only securities with maturities of one and two periods and confining the analysis to pure discount securities. For a thorough analysis that includes the extension to coupon bonds, see Shiller (1990).

Let p_t and $p2_t$ be the prices at time t of securities that will pay \$1 at times $t+1$ and $t+2$, respectively. The securities are assumed to have zero risk of default. Let i_t and $i2_t$ be the continuously compounded, per-period yields to maturity of the two securities. Then by definition:

$$p_t \cdot e^{i_t} = p2_t \cdot e^{2 \cdot i2_t} = 1$$

so that:

$$i_t = -\ln p_t \text{ and } i2_t = -(\ln p2_t)/2.$$

Now consider an agent who has no net borrowing needs at time t, but will need to borrow money at time $t+1$ to be repaid at time $t+2$. Without loss of generality, assume the borrowing need can be expressed as a need to repay \$1 at $t+2$. The agent could simply wait until $t+1$, and borrow p_{t+1} dollars at an interest rate of i_{t+1}.

An alternative transaction would lock in an interest rate at time t. The agent could issue a two-year bond, and invest the proceeds $p2_t$ in $p2_t/p_t$ one-year bonds. Then the agent would carry a zero balance from t to $t+1$. At $t+1$ the one-year bonds mature and the agent would receive $p2_t/p_t$ dollars, with a requirement to repay \$1 at $t+2$. The implied interest rate from $t+1$ to $t+2$ on this transaction is:

$$f_t = -\ln(p2_t/p_t) = 2 \cdot i2_t - i_t$$

which is called (if a period is a year) the one-year-ahead, one-year forward rate: it is the one-year rate of interest that can be locked in, one year ahead of time. Note that this definition implies that the two-year rate is the average of the one-year rate and the forward rate.

The agent faced with a choice between locking in the forward rate at time t and waiting to borrow at time $t+1$ is likely to compare the forward rate to the one-year rate expected to prevail at $t+1$. The expectations hypothesis of the term structure supposes that market forces will drive the forward rate to equal the expected one-year spot rate plus a "term premium," which is supposed to be constant over time for each maturity but might differ across maturities. Hence (since only one maturity is considered here), the expectations hypothesis can be stated:

$$f_t = E_t(i_{t+1}) + \theta.$$

The term premium θ is commonly understood to reflect the differing risk of the two strategies; alternatively it could reflect maturity-specific forces of supply and demand for funds (compare Culbertson 1957). In either case, the premium could in principle be positive, negative, or zero.

It is important to note that there is no theoretical reason for the expectations hypothesis to hold except as an approximation. If $\theta = 0$, as might be suggested by risk neutrality, the hypothesis in fact *cannot* hold for all maturities simultaneously. When non-zero term premia are allowed, there are no compelling reasons why they should be constant over time. Some recent attempt to explain failures of the expectations hypothesis concentrate on modelling changes in the term premia—for example, Engle and Ng (1993). See Shiller (1990) and the references therein for further detail on these matters.

Substituting (3) into (1) and rearranging, we obtain:

$$E_t(\Delta\pi_{t+2}) = (f_t - i_t) - E_t(\Delta r_{t+2}) - \theta. \qquad (4)$$

According to the expectations theory, the forward–spot spread reflects the expected change in the spot rate, which in turn reflects both expected changes in inflation and expected changes in the real rate of interest. As expressed mathematically in Equation (4), *the term structure cannot be linked to expected inflation without consideration of the real rate.* The forward–spot spread will directly measure expected changes in inflation only if the real rate is expected to remain unchanged.

Equation (4) forms the basis for the empirical investigation. Given an assumption about the expected change in the real rate, the right-hand side of (4) expresses the "market" expectations of the change in inflation. If those expectations are rational, they should predict the actual change in inflation. Therefore, regressions of the actual change in the inflation rate on the forward–spot spread two years previous and the expected change in the real rate can be used to assess the consistency of historical data with the predictions of the expectations hypothesis.

Two types of checks will be used. First, equation (4) predicts that, in a regression of the actual change of inflation on the lagged spread and the expected change in the real rate, the coefficients on those two variables will be 1 and −1, respectively. If the estimated coefficients are consistent with (not significantly different from) those values, the data are consistent with the expectations hypothesis. However, such a result alone does not measure the importance of the yield curve in predicting inflation. Results will also be reported for tests of the hypothesis that the coefficient on the forward–spot spread is 0. Only if this hypothesis is rejected can the spread be said to have a significant relationship to future inflation.

This study finds little evidence of a link between the term structure of interest rates and future inflation, at the horizon explored here.

II. Data and Econometric Method

Three different sets of data for interest rates on U.S. risk-free (or low-risk) securities are used. These are described fully in Appendix 1, "Data Sources." The data sets jointly cover the period 1919 to 1990. Because of the varying quality of the data and the possibility of temporal instability, results will be reported for subperiods corresponding to the individual data sets and to the first and second halves of the postwar sample, as well as for the full period.

The variables used in the empirical analysis are:

i_t : Yield on one-year-maturity securities, December observations.

f_t : One-year-ahead, one-year-forward interest rate, computed from December observations of one-year and two-year yields. (See the box.) Date subscript refers to time of observation, so f_t is the forward rate applicable to the period $t + 1$ to $t + 2$.

π_t : December-to-December percentage change in the Consumer Price Index. The dating convention means that π_t is the rate for the period $t - 1$ to t.

r_t : Ex post real rate of interest, as defined by equation (2).

$\Delta r^e_{t-2,t}$: Change in the real rate from $t - 1$ to t, "expected" at $t - 2$. See Appendix 4, "The 'Expected' Real Rate."

While monthly observations of all variables are available, only one annual observation is used, for several reasons. First, use of monthly data gives a misleading indication of the amount of information present in the data set. The number of fully independent observations is equal to the span of the data divided by the longest maturity used; thus, 20 years of data contain only 10 fully independent observations of two-year horizons, even if all 240 monthly observations are used. Monthly data do in principle contain more information than annual data, and in principle reported standard errors can be adjusted for the observation overlap. However, these adjustment methods are known to work poorly when the overlap is substantial (see, for example, Richardson and Stock 1989).

The use of December observations for the variables may be questioned. December financial data may be contaminated by "end-of-year effects" due to tax and accounting influences on portfolios. Any such effects should be mitigated here by the fact that *only* December data are used; changes are December-to-December rather than December to some other month. Nevertheless, December data might be especially noisy. For this reason, all equations were reestimated using June data for the periods 1950 to 1990, 1950 to 1970, and 1971 to 1990. The results using June data were broadly similar to those using December data and are not reported in detail; any significant differences are noted where appropriate.

Econometric details for the results presented below are given in Appendix 2, "Theoretical Structure of Error Terms" and Appendix 3, "Econometric Method."

Table 1
Estimates of Equation (5)
$\Delta\pi_t = \beta_0 + \beta_1(f_{t-2} - i_{t-2}) + u_t$

	Estimation Period						
Variable	1923–90	1950–90	1950–70	1971–90	1922–49	1932–49	1921–30
constant	−.561	−.351	−.108	−.782	−.060	−1.697	−.847
	(.442)	(.416)	(.600)	(.585)	(1.466)	(2.948)	(1.793)
(f − i)(t − 2)	1.741	1.764	2.475	1.838	.768	2.676	−2.024
	(.618)	(.693)	(1.501)	(.663)	(1.902)	(2.530)	(9.064)
$\bar{R}^2$	.050	.168	.018	.266	−.034	−.036	−.120
D.W.	2.217	1.519	1.389	1.656	2.024	2.272	2.018
s.e.	3.940	2.410	2.333	2.541	5.821	6.477	6.697

Newey-West standard errors in parentheses. See Appendix 3, "Econometric Method," for details of computations.

III. Results

Suppose that the expected change in the real rate Δr^e is always zero.[3] After imposing this condition, the theoretical relationship (4) implies a coefficient of 1 on the forward–spot spread $(f - i)$ in the regression:

$$\Delta\pi_t = \beta_0 + \beta_1(f_{t-2} - i_{t-2}) + u_t. \qquad (5)$$

This regression is closely related to those reported by Fama (1990) and Mishkin (1990).[4]

Results of estimating (5) are reported in Table 1. The coefficient on the forward–spot spread is nowhere significantly different from 1; the data are in this sense consistent with the expectations hypothesis with a zero expected change in the real rate. The spread also appears to have some explanatory value for future inflation: the hypothesis that the coefficient on the spread is 0 can be rejected for the full 1923–90 sample and for the 1950–90 and 1971–90 subsamples.

However, the results also show that the forward–spot spread forecasts very little of the subsequent change in inflation. The $\bar{R}^2$ is 0.05 for the full sample. Furthermore, the correlation between the spread and inflation arises almost entirely in the most recent 20-year period; $\bar{R}^2$ is 0.27 for 1971–90 but 0.02 for 1950–70 and negative for all pre-war samples.

If the expected real rate of interest is not constant, equation (5) is not a valid representation of the expectations hypothesis. As discussed above, that hypothesis says that the forward–spot spread predicts the sum of the expected change in inflation and the expected change in the real rate. If the expected change in the real rate is not zero, the hypothesis no longer implies that the coefficient on the forward–spot spread in (5) is 1; in fact, in this case the expectations hypothesis has no testable implications at all for equation (5).[5]

Direct measures of the expected change in the real rate of interest are not available. However, such changes have been in part predictable in the data used in this study. Figure 1 shows actual values and the fitted values of a regression of the change in the ex post real rate on observations of variables in our data set dated $t - 2$ or earlier, over the whole sample. Under rational expectations, this information would have been incorporated in market expectations. Therefore, the fitted values from this regression[6] can be used as proxies for the expected changes in the real rate. Details and further justification of this procedure are given in Appendix 4.

The next set of results incorporates this measure of the expected change in the real rate into the inflation change regression. Table 2 provides estimates of the regression:

$$\Delta\pi_t = \beta_0 + \beta_1(f_{t-2} - i_{t-2}) + \beta_2\Delta r^e_{t-2,t} + u_t. \qquad (6)$$

The expectations hypothesis as given by (4) predicts

[3] This hypothesis is implied by the hypothesis that the ex ante real rate is constant. An expectation of no change also allows the real rate to follow a random walk. Note that this discussion refers to the change from $t + 1$ to $t + 2$ expected at t.

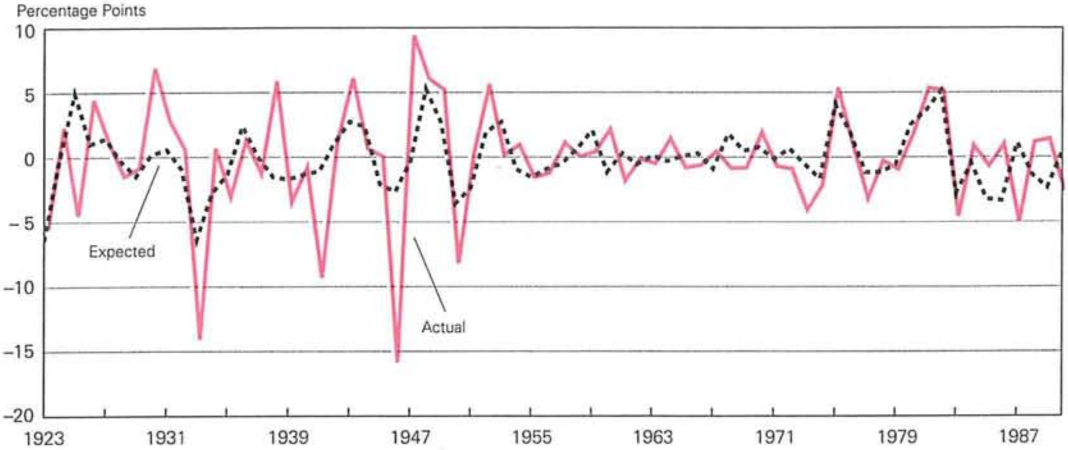
[4] The regressions are not directly comparable, because both Fama and Mishkin use the difference between long-term average and short-term average inflation rates instead of inflation rate changes, and yield spreads between long-term and short-term securities rather than forward–spot spreads.

[5] Mishkin (1990) discusses how the coefficients in a regression similar to (5) can be interpreted when the real rate varies, under the assumption that the expectations hypothesis holds exactly.

[6] The estimation is done separately for each sample period.

Figure 1

Actual and "Expected" Change in the Ex Post Real Rate of Interest



values of 1 and -1 for the coefficients β_1 and β_2, respectively. The last line of Table 2 gives p-values for tests of this restriction so that, for example, a p-value of 0.10 or less means that the restriction is rejected at the 10 percent level.

Table 2 contains good news and bad news for the expectations hypothesis. The data are consistent with the coefficient values predicted by the theory: the hypothesis that the coefficients are 1 and -1 is nowhere rejected at the 10 percent level or below. Furthermore, these regressions explain substantially more of the variation in inflation than those reported

Table 2

Estimates of Equation (6)

$\Delta\pi_t = \beta_0 + \beta_1(f_{t-2} - i_{t-2}) + \beta_2 \Delta r^e_{t-2,t} + u_t$

	Estimation Period						
Variable	1923–90	1950–90	1950–70	1971–90	1922–49	1932–49	1921–30
constant	.036	.175	.056	.066	−.181	−.114	−.278
	(.389)	(.429)	(.401)	(.693)	(1.196)	(2.203)	(.953)
(f − i)(−2)	.005	−.050	1.396	−.171	.116	.192	.905
	(.559)	(.852)	(1.300)	(1.022)	(1.553)	(2.090)	(6.417)
Δr^e	−.905	−.957	−.900	−.987	−.847	−.828	−.916
	(.222)	(.284)	(.240)	(.380)	(.186)	(.130)	(.351)
$\bar{R}^2$	.241	.417	.553	.419	.145	.184	.298
DW	2.596	1.882	1.565	1.996	2.428	2.839	2.934
s.e.	3.521	2.018	1.574	2.261	5.291	5.749	5.303
p-value[a]	.104	.157	.891	.101	.558	.307	.967

[a]p-value for the null hypothesis that the coefficients on (f − i) and Δr^e are 1 and −1, respectively. A value less than 0.10 means the hypothesis is rejected at the 10 percent level, and so on.

Newey-West standard errors in parentheses. See Appendix 3, "Econometric Method," for details of computations.

Table 3

Estimates of Equation (7)

$\Delta i_t = \beta_0 + \beta_1(f_{t-1} - i_{t-1}) + u_t$

Variable	Estimation Period 1922–89	1949–89	1949–69	1970–89	1921–48	1931–48	1920–29
constant	.100	.199	.074	.077	−.203	.215	−.226
	(.162)	(.227)	(.307)	(.420)	(.189)	(.445)	(.208)
(f − i)(−1)	−.178	−.132	1.276	−.196	−.002	−.324	1.174
	(.240)	(.274)	(.921)	(.275)	(.478)	(.775)	(.985)
$\bar{R}^2$	−.008	−.022	.067	−.047	−.038	−.047	.026
DW	2.077	1.981	2.022	1.606	2.542	3.240	1.924
s.e.	1.264	1.498	.902	1.954	.822	1.048	.687

Newey-West standard errors in parentheses. See Appendix 3, "Econometric Method," for details of computations.

in Table 1, which excluded the expected change in the real rate. However, little credit for these results is due to the forward–spot spread. While the estimates of β_1 are nowhere significantly different from 1, they are also nowhere significantly different from 0. The forward–spot spread is not making a statistically meaningful contribution to explaining the change in inflation in these regressions. To the extent that movements in inflation are predictable with these data, they are associated with movements in the expected real rate of interest and *not* with the shape of the yield curve.

Note in particular the results for the 1971–90 period, which were the most favorable for equation (5). When the expected real rate is included in the regression, the coefficient on the forward–spot spread has a negative sign and a large standard error. The simple correlation between the spread and the future change in inflation during this period, as reported in Table 1, appears to be an artifact of the omission of the expected real rate, and cannot be attributed to the expectations theory mechanism.[7]

According to the expectations hypothesis, any ability of the yield curve to forecast changes in inflation must be a byproduct of the ability of the yield curve to forecast changes in short-term interest rates. The results in Table 2 suggest an examination of the extent to which the yield curve performs this function. Table 3 presents results of the regression:

$$\Delta i_t = \beta_0 + \beta_1(f_{t-1} - i_{t-1}) + u_t \qquad (7)$$

where the forward unbiasedness condition (1) predicts a slope coefficient of 1.

The equation performs poorly. The slope coefficient is significantly less than 1 for the full sample and for the subperiods 1949–89 and 1970–89; estimated values are negative for all periods except 1949–69 and 1920–29. The coefficient is not significantly different from 0 for any period. Thus, the forward–spot spread has essentially *no* ability to forecast changes in the spot rate.[8] This result is consistent with previous findings in the term structure literature. (Shiller 1990 includes a summary of empirical work.)

The results in Table 3 do *not* mean that the expectations hypothesis is badly misguided as a description of the yield curve. Rather, they are consistent with a view that changes in the one-year rate have a negligible forecastable component. The expected change in the spot rate will be nearly zero throughout the sample. The strict expectations hypothesis would then require that the forward–spot spread be a constant (the term premium). If the expected change in the spot rate is zero, any variations in the forward–spot spread are necessarily variations in the term premium. Such deviations from the expectations hypothesis could be small while being consistent with the results in Table 3.[9]

[7] This particular result is sensitive to the use of December observations. When June observations are used for this time period, the coefficient on the forward–spot spread is 0.85 with a standard error of 0.42.

[8] When June data are used, the coefficient on the spread is positive and not significantly different from 1 for the post-war samples. The coefficient is not significantly different from 0 for any of these samples, however, so the conclusion that the spread has no ability to forecast changes in interest rates is robust to the change of dates.

[9] Mankiw and Miron (1986) give this interpretation, and suggest that the unforecastability of changes in short rates is a result of Federal Reserve behavior: they find that forward unbiasedness regressions have substantially more explanatory power for interest rates changes prior to the founding of the Fed.

Table 4
Estimates of Equation (8)*
$\Delta\pi_t = \beta_0 + \beta_1(f_{t-2} - i_{t-2}) + \beta_2\Delta r_t + u_t$

	Estimation Period						
Variable	1923–90	1950–90	1950–70	1971–90	1922–49	1932–49	1921–30
constant	.036	.175	.056	.066	−.181	−.114	−.278
	(.184)	(.245)	(.305)	(.616)	(.168)	(.578)	(.103)
(f − i)(−2)	.005	−.050	1.396	−.171	.116	.192	.905
	(.262)	(.379)	(.957)	(.742)	(.273)	(.663)	(.607)
Δr	−.905	−.957	−.900	−.987	−.847	−.828	−.916
	(.066)	(.103)	(.064)	(.309)	(.038)	(.058)	(.008)
$\bar{R}^2$	.911	.685	.865	.546	.976	.966	.994
DW	1.930	1.915	1.888	1.587	2.393	2.352	2.320
s.e.	1.209	1.484	.866	1.998	.882	1.172	.491
p-value[a]	.000	.001	.261	.003	.000	.004	.000

*Instruments: See Table A2.
[a]p-value for the null hypothesis that the coefficients on (f − i) and Δr are 1 and −1, respectively.
Newey-West standard errors in parentheses. See Appendix 3, "Econometric Method," for details of computations.

For purposes of this study, the point is that predictability of changes in the spot rate is essential for the yield curve to predict changes in inflation, under the expectations hypothesis. The forward–spot spread will predict changes in inflation *only* if changes in short rates are predictable *and* such changes reflect changes in expected inflation. If changes in short rates are not predictable, the yield curve has no role to play, and expected changes in inflation will be completely absorbed by expected changes in the real rate.

This point can be emphasized by examining a hybrid equation that replaces the expected change in the real rate with its actual (ex post) value. Table 4 presents estimates of the regression:

$$\Delta\pi_t = \beta_0 + \beta_1(f_{t-2} - i_{t-2}) + \beta_2\Delta r_t + u_t. \qquad (8)$$

Since the ex post real rate is likely to be correlated with the error term, the equation is estimated using instrumental variables. Again the coefficients on the spread and on the change in the real rate should be 1 and −1 under the expectations hypothesis.

The interest of equation (8) is that it can be derived by substracting the ex post real rate from both sides of the forward unbiasedness condition (equation 1). It therefore provides a direct link between that equation and the inflation forecasting equation (6). Equation (8) is not a forecasting equation because it includes the ex post real rate, which is not known in advance; on the other hand, it can be shown to have a smaller error variance under the expectations hypothesis than (6) and therefore permits more precise statistical inference. (See Appendix 2 for details.)

The results in Table 4 indicate that the expectations hypothesis is rejected at well below the 1 percent level for all periods except 1950–70. Even though the smaller error variances give smaller standard errors, the coefficient on the spread is nowhere significantly different from 0.[10] *For this data set, the yield curve has no ability to forecast changes in inflation within the expectations theory framework.*

IV. Alternative Hypotheses

The analysis to this point has been based entirely on the expectations theory of the yield curve, with rational expectations assumed. This section will briefly consider alternative theories.

As an empirical background for this discussion, first consider whether the yield curve has any ability to forecast changes in inflation when the constraints of the expectations hypothesis are dropped. Table 5 presents results of regressions of the form:

[10] When June observations are used, the coefficient on the forward–spot spread for the 1971–90 period is positive and marginally significantly different from 0 at the 5 percent level. Otherwise the results are very similar.

Table 5
Estimates of Equation (9)

$$\Delta\pi_t = \beta_0 + \sum_{i=1}^{n}\beta_i(f_{t-1-i} - i_{t-1-i}) + \sum_{i=1}^{n}\gamma_i\pi_{t-1-i} + \sum_{i=1}^{n}\delta_i i_{t-1-i} + u_t$$

	Estimation Period						
Variable	1923–90	1950–90	1950–70	1971–90	1922–49	1932–49	1921–30
constant	−.532	.249	−3.551	3.018	−.210	3.094	−23.012
	(.902)	(.724)	(1.273)	(3.263)	(3.681)	(3.245)	(16.799)
(f − i)(−2)	2.013	.756	1.512	.674	5.547	−.347	17.854
	(1.315)	(.881)	(1.161)	(1.021)	(3.434)	(5.781)	(19.003)
(f − i)(−3)	−.803	1.455	5.000	.855	−3.933	−.108	—
	(.989)	(.648)	(2.226)	(1.162)	(2.709)	(2.988)	
(f −i)(−4)	1.071	—	—	—	—	—	—
	(.670)						
$\pi(-2)$	−.329	−.313	−.031	−.498	−.458	−.515	−.681
	(.083)	(.142)	(.203)	(.287)	(.119)	(.149)	(.260)
$\pi(-3)$	−.036	.535	.659	.596	.198	−.091	—
	(.124)	(.167)	(.211)	(.219)	(.171)	(.190)	
$\pi(-4)$	.296	—	—	—	—	—	—
	(.149)						
i(−2)	.413	.062	−.511	.135	1.370	1.334	4.602
	(.375)	(.266)	(.289)	(.318)	(.993)	(.760)	(3.416)
i(−3)	−.426	−.386	.946	−.679	−1.326	−2.269	—
	(.307)	(.262)	(.376)	(.328)	(.996)	(1.302)	
i(−4)	−.004	—	—	—	—	—	—
	(.413)						
$\bar{R}^2$	.173	.412	.444	.427	−.012	−.095	.181
DW	2.609	1.995	1.477	2.584	2.509	2.931	2.932
s.e.	3.676	2.026	1.756	2.245	5.757	6.659	5.727
p-value[a]	.256	.011	.116	.591	.269	.988	.347

[a]p-value for the hypothesis that the coefficients on the (f −i) terms are jointly zero.
Newey-West standard errors in parentheses. See Appendix 3, "Econometric Method," for details of computations.

$$\Delta\pi_t = \beta_0 + \sum_{i=1}^{n}\beta_i(f_{t-1-i} - i_{t-1-i}) + \sum_{i=1}^{n}\gamma_i\pi_{t-1-i} + \sum_{i=1}^{n}\delta_i i_{t-1-i} + u_t \quad (9)$$

where the lag length n varies with the sample length.

The last row of Table 5 gives p-values for the hypothesis that the coefficients on the forward–spot spread are jointly 0. This hypothesis is accepted for the full period, but is rejected for the postwar sample 1950–90. That rejection appears to arise largely in the 1950–70 period.[11] This result provides some evidence that the spread has an association with future changes in inflation. Because the spread was not significant in the equations based on the expectations hypothesis, however, any such association must arise through some other mechanism.

Time-Varying Term Premia

If the term premium is not constant but rather reflects a time-varying risk premium, expected changes in inflation could be associated with changes in risk and therefore could affect the term premium.

[11] When June observations are used, the hypothesis that the coefficients on the spread are jointly 0 is not rejected for any postwar period.

This would lead to an association between the forward–spot spread and future changes in inflation that is not captured by the expectations hypothesis. This possibility deserves further investigation, which would require linking expected changes in inflation with the volatility of interest rates (possibly including a link between the level and the variance of inflation). Note, however, that in the Table 5 regressions the sum of coefficients on the forward–spot spread is positive, so increases in the spread predict increases in inflation. Since no change in the spot rate is predicted, this implies a predicted *decline* in the real rate (see also Fama 1990). A risk-based model would therefore have to explain why an increase in the risk premium embedded in the term structure anticipates a decline in the real rate of interest.

The results presented here undermine a particularly strong form of the view that interest rates respond to inflation expectations, the view that interprets a steep yield curve as a reliable forecast of accelerating inflation.

Dropping Rational Expectations

The discussion in this paper has presumed that market prices reflect rational expectations of future inflation and interest rates. Another avenue for exploration would examine the expectations hypothesis without assuming rational expectations. If the market expects a constant real rate, ignoring the forecastability found here, the real rate can be ignored in making the link between the yield curve and inflation. Equation (5) above takes that approach, but inconsistently presumes that the market rationally anticipates changes in the inflation rate while ignoring the evidence that such changes are likely to be offset by changes in real rates.

An alternative uses direct measurement of expectations. Froot (1989) finds some correspondence between the forward–spot spread and survey expectations of changes in spot rates. Combining those results with survey measures of inflation expectations could illuminate the relationship between inflation expectations and the yield curve without insisting on a further connection to actual inflation.

The Modigliani-Sutch Equation

Modigliani and Sutch (1966) and a number of successors model long rates as a distributed lag of short rates. Shiller (1987) argues that equations of this type have proved quite robust as a description of the yield curve. This approach can be interpreted within the expectations theory as assuming adaptive expectations of the future path of short rates.

This theory leaves little room for the forward–spot spread to predict inflation. If long rates are a distributed lag of short rates, so is the forward rate and so is the forward–spot spread. Therefore, in this view, the spread is an artifact of the recent history of spot rates, and any association between the spread and future inflation would be an indirect result of the impact of spot rates on economic activity.

V. Conclusion

The findings of this study may be summarized as follows. First, the expectations theory of the term structure implies that the forward–spot spread forecasts the *sum* of the expected change in inflation and the expected change in the real rate of interest. Second, changes in the real rate of interest are in part predictable, so that such expected changes should be taken into account in linking the term structure to expected changes in inflation. Third, after such account is taken, the forward–spot spread has essentially no power to explain one-year-ahead changes in one-year inflation.

The results presented here consider only one maturity. However, they provide some guidance for other horizons. Under the expectations hypothesis, the term structure can forecast inflation only if it forecasts changes in interest rates. Horizons for which other work has found little predictability in interest rate changes are unlikely to give results different from those in this paper.

The results presented here do not contradict the view that interest rates respond to inflation expectations. Rather they undermine a particularly strong form of that view, which interprets a steep yield curve as a reliable forecast of accelerating inflation.

Appendix 1: Data Sources

π_t: For all sample periods, the inflation rate is defined as the percentage change from December of year $t - 1$ to December of year t of the Consumer Price Index.

i_t, f_t: December observations of one-year and two-year interest rates were taken from the data sets described below. The one-year rate was used as i_t; the forward rate was constructed so that the two-year rate was the average of the forward rate and the one-year rate.

1946–90 Data provided by J. Huston McCulloch, constructed as described in McCulloch (1975) and summarized in Appendix B of Shiller (1990). Data are pure discount yields implied by observed end-of-month data on U.S. Treasury securities, adjusting for tax effects and using a cubic spline to fit a yield curve. While in principle these yields are subject to measurement error, for the maturities examined here the errors are surely trivial.[12]

1929–49 Data from Cecchetti (1988). These data also were constructed by fitting a yield curve to end-of-month U.S. Treasury data; Cecchetti corrected the data for distortions caused by an exchange privilege carried by many Treasury bonds during this period. The Cecchetti data are coupon bond yields, unlike the theoretically preferable discount yields provided by McCulloch, but comparison of the McCulloch pure discount series with coupon yields suggests the differences are small for these maturities. For combined data sets, Cecchetti data are used for the 1929–46 period.

1919–30 Data from Baum and Thies (1992). These data were constructed using curve-fitting methods, but using railroad bonds rather than Treasury securities. Like the Cecchetti data, these are coupon bond yields. The Baum and Thies data have been mean-adjusted so that the 1929 observations equal those from Cecchetti. This amounts to assuming that the railroad bonds carried a constant risk premium over Treasuries.

Table A-1 gives means and standard deviations of the various data series for the sample periods covered by each data set, including the constructed ex post real rate, forward–spot spread, and differenced inflation, real rate, and spot rate series.

Table A1
Sample Statistics

Variable		1919–30	1929–49	1946–90
π	mean	−.038	1.736	4.606
	std. dev.	6.087	6.403	4.005
i	mean	4.684	1.105	5.433
	std. dev.	1.099	.975	3.250
f	mean	4.712	1.825	5.750
	std. dev.	.926	1.148	3.200
r	mean	6.095	−.689	1.102
	std. dev.	4.735	7.238	3.274
f − i	mean	.028	.720	.317
	std. dev.	.199	.387	.625
$\Delta\pi$	mean	−1.904	−.133	−.273
	std. dev.	6.859	6.276	3.139
Δr	mean	.711	−.343	.227
	std.dev.	6.824	6.639	2.946
Δi	mean	−.188	−.116	.139
	std dev.	.661	1.022	1.439

Appendix 2: Theoretical Structure of Error Terms

The hypothesis of rational expectations has implications for the error terms of the equations estimated. Certain econometric points require understanding of these properties. First, define the expectational errors for the change in the spot rate and the change in the real rate:

$$\varepsilon_t \equiv \Delta i_t - E_{t-1}(\Delta i_t). \tag{A1}$$

$$\nu_t \equiv \Delta r_t - E_{t-2}(\Delta r_t). \tag{A2}$$

Under the hypothesis of rational expectations, ε_t is uncorrelated with any information available at time $t - 1$ or earlier, and ν_t is uncorrelated with any information available at time $t - 2$ or earlier (since the expectations are formed at $t - 2$). However, ν_t is likely to be correlated with information available at $t - 1$, and in particular with ν_{t-1}. Therefore, it is likely to have to have a first-order moving average (MA(1)) structure.

Then equations (1), (2), and (3) in the text imply that:

$$\Delta\pi_{t+2} = -\theta + (f_t - i_t) - E_t(\Delta r_{t+2}) - \nu_{t+2} + \varepsilon_{t+1} \tag{A3}$$

where the error terms are theoretically uncorrelated with the right-hand-side variables, justifying least squares estimation of the empirical relationship (5). However, ν_t is MA(1) and further may be correlated with ε_t, introducing an additional MA(1) effect to the total error in (A3). This serial correlation requires use of a correction in calculating standard errors and test statistics in the results, as described in Appendix 3 below.

Equation (A1) immediately gives the error term for the forward unbiasedness regression (7):

$$\Delta i_{t+1} = -\theta + (f_t - i_t) + \varepsilon_{t+1}. \tag{A4}$$

Under rational expectations ε_t is serially uncorrelated so no correction is needed for that equation. Finally, the following equation is obtained by subtracting the ex post real rate from both sides of (A4) :

[12] The discount bond yields used are almost indistinguishable from the constant maturity coupon bond yield series maintained by the Federal Reserve Board. This conclusion is further reinforced by comparison of the McCulloch data for recent periods to market yields for stripped Treasury bonds.

$$\Delta\pi_{t+2} = -\theta + (f_t - i_t) - \Delta r_{t+2} + \varepsilon_{t+1}. \quad \text{(A5)}$$

This equation has the same structure as (A3) but omits the expectational error for the change in the real rate. The smaller error variance explains the smaller standard errors obtained when the hybrid equation (8) is estimated instead of the forecasting equation (6).

Appendix 3: Econometric Method

All coefficient estimates were generated by ordinary least squares except for those in Table 4, which are instrumental variables estimates. Instruments for each time period are the independent variables for the real rate regressions in Table A-2. As noted above, theory suggests that the error terms of all but the forward unbiasedness regression are serially correlated. Therefore, all standard errors and hypothesis tests were computed using covariance matrices robust to heteroskedasticity and first-order serial correlation as per Newey and West (1987) except for those in Table 2, which omit the serial correlation adjustment.

The tables report Durbin-Watson statistics for first-order serial correlation. Again, the expectations hypothesis predicts MA(1) errors in all equations except those in Table 2; note also that except in Table 2 standard errors incorporate an (asymptotically) appropriate adjustment. Q-statistics for serial correlation of first and higher order (depending on sample length) were computed but are not reported in the tables. P-values for these tests fell below 0.15 only as follows: Table 1, 1923–90 (0.003); Table 2, 1932–49 (0.005); Table A-2, 1921–30 (0.109) and 1971–90 (0.027).

P-values for the hypothesis tests in Tables 3, 4, and 5 were computed from Wald test statistics generated by RATS version 4.01 (in which all computations were performed). The Wald statistics are asymptotically chi-square with degrees of freedom equal to the number of restrictions. In view of the limited number of observations in many of the regressions here, the reported p-values incorporate a small sample adjustment: the Wald statistic is divided by the number of restrictions and the result is compared to an F-distribution with degrees of freedom equal to the number of restrictions (numerator) and degrees of freedom of the regression (denominator). The adjustment slightly increases the p-values; those closer to 0 are increased proportionately more. A small number of p-values are changed from being slightly less than 0.10 to being slightly more than 0.10, but the adjustment does not affect significance of any hypothesis at the 5 percent or 1 percent level.

Table A-2

Real Rate Regression

$$\Delta r_t = \beta_0 + \sum_{i=1}^{n} \beta_i (f_{t-1-i} - i_{t-1-i}) + \sum_{i=1}^{n} \gamma_i \pi_{t-1-i} + \sum_{i=1}^{n} \delta_i i_{t-1-i} + u_t$$

	Estimation Period						
Variable	1923–90	1950–90	1950–70	1971–90	1922–49	1932–49	1921–30
constant	.930	.478	3.584	−1.170	.788	−2.564	25.881
	(.885)	(.618)	(1.464)	(4.162)	(3.866)	(3.313)	(17.277)
(f − i)(−2)	−3.509	−1.486	.007	−1.784	−6.910	1.276	−19.622
	(1.522)	(1.125)	(1.345)	(1.135)	(3.892)	(6.027)	(19.505)
(f − i)(−3)	1.547	−.572	−5.906	.215	4.548	−1.357	—
	(1.329)	(1.049)	(2.499)	(1.821)	(3.202)	(3.075)	
(f − i)(−4)	−.907	—	—	—	—	—	—
	(.810)						
π(−2)	.383	.424	.048	.558	.494	.522	.739
	(.085)	(.121)	(.214)	(.341)	(.133)	(.150)	(.286)
π(−3)	.028	−.579	−.701	−.549	−.218	.114	—
	(.119)	(.168)	(.203)	(.189)	(.179)	(.197)	
π(−4)	−.270	—	—	—	—	—	—
	(.160)						
i(−2)	−.868	−.373	.481	−.496	−2.101	−2.238	−5.240
	(.450)	(.359)	(.399)	(.405)	(1.116)	(.814)	(3.507)
i(−3)	.457	.541	−.828	.728	1.839	2.896	—
	(.365)	(.359)	(.455)	(.440)	(1.161)	(1.373)	
i(−4)	.352	—	—	—	—	—	—
	(.456)						
R^2	.288	.464	.613	.517	.254	.351	.471
DW	2.747	2.450	2.065	2.740	2.461	2.920	2.934
s.e.	3.872	2.176	1.831	2.580	5.987	6.825	6.079

Newey-West standard errors in parentheses. See Appendix 3, "Econometric Method," for details of computations.

Appendix 4: The "Expected" Real Rate

Estimation of equation (6) requires construction of a measure of the change in the real rate of interest from year $t + 1$ to $t + 2$ that is expected in year t. This variable $\Delta r^e_{t,t+2}$ is constructed as the fitted values of a regression of the ex post change in the real rate Δr_{t+2} on the variables $(f_t - i_t)$, π_t, and i_t, and lags of these variables. The regression is performed separately for each sample period. Table A-2 indicates the lags included for each subsample, gives coefficient estimates and Newey-West standard errors, and provides summary statistics.

Given the importance of the real rate in the empirical results, some discussion of this procedure is warranted. Most importantly, note that *these results are not intended as a structural estimate of the expected change in the real rate*. This constructed measure is not asserted to represent the expectations of "the market" or of any participant(s). Rather, a proxy is sought for the expected change that has desirable *econometric* properties under the hypothesis of rational expectations.

The advantage of the procedure used here is that the expectations errors (ν_t in Appendix 2) implied by this procedure are by construction uncorrelated with the independent variables in all regressions estimated, since those variables are used as explanatory variables in the real rate regressions. This implies that deviations of the coefficient estimates from their theoretical values cannot be due to mismeasurement of expectations: the procedure by construction cannot generate coefficient bias. (Deviations could be due to failure of the rational expectations assumption, a possibility discussed in the text.) Another way to understand this argument is to note that the coefficient estimates (but not the standard errors or measure of fit) produced by this procedure are identical to those given by instrumental variables estimation of (8), since the instruments are identical to the independent variables in the real rate regressions. Validity of instrumental variables estimation requires that instruments be correlated with the variables instrumented (the change in the ex post real rate) and be independent of the error term, but it does not require any structural relationship between the instruments and the instrumented variable.

Under rational expectations, "the market" uses all available information to form expectations, while here only variables in the data set are used. Thus the "market" expectation could be more accurate than that used here. This difference causes no estimation bias, although the estimates would be more precise if a more accurate forecast were used. On the other hand, the expected change used here is formed using sample information not available to market participants. No formal stability tests were performed on the results in Table A-2, but the estimates show no obvious sign of instability. In any case, this would not affect the validity of the instrumental variables interpretation.

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