The Bond Rate and Actual Future Inflation

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It is widely believed that bond yields contain useful information about expected inflation. Many have empirically investigated this issue by examining whether the slope of the term structure has any predictive content in forecasting future inflation. That research, however, has produced disparate results. In a series of papers, Mishkin (1990a, 1990b, 1991) and Jorian and Mishkin (1991) report evidence that indicates the slope has predictive content at long horizons but not at short horizons.[1] In contrast, Engsted (1995) investigates whether the spread between the long-term interest rate and the one-period inflation rate predicts future one-period inflation. While this spread *does* help predict future inflation for a number of countries, it does not for the United States.[2]

In this article, I provide new evidence on the predictive content of the bond rate for future inflation using cointegration and error-correction modeling. The empirical work here corrects for two possible shortcomings of the previous research that may account for the disparate results described above. First, I relax the assumption made in previous studies that the ex ante real interest rate is constant. If the ex ante real rate is variable, then short-run movements

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[1] In this research the horizon forecasts of inflation match that of the slope of the term structure. Hence, the result that the slope has predictive content at long horizons but not at short horizons should be interpreted to mean long-term bonds help predict inflation at long horizons, but short-term bonds do not help predict inflation at short horizons. In this article, by contrast, there is no such matching. In fact, I explore the predictive content of the long bond rate at short and long inflation horizons.

[2] Though this spread does Granger-cause the U.S. inflation rate, the sum of coefficients that appear on lagged values of the spread in the inflation equation is small in magnitude. Engsted, however, does not test whether the sum of these coefficients is different from zero.

in the bond rate do not necessarily reflect movements in its expected inflation component. In that environment the predictive content of the bond rate for future inflation should be investigated controlling for the influences of variables that capture movements in the real rate of interest. The empirical results indicate that inferences regarding the predictive content of the bond rate for future inflation are sensitive to such conditioning. Second, the empirical work in this article examines whether the predictive content has changed over time, in particular between pre- and post-1979 periods. Recent research reported in McCallum (1994) and Rudebusch (1995) indicates that the term structure's ability to predict future economic variables may be influenced by the way the Fed conducts its monetary policy.[3] Most economists would agree that since 1979, the Fed has made repeated attempts to bring down the trend rate of inflation and contain inflationary expectations. In that environment an increase in the current bond rate, even if it correctly signals an increase in long-term expected inflation, may not necessarily translate into higher actual future inflation.

The results here focus on the behavior of the nominal yield on ten-year U.S. Treasury bonds over the period 1959Q1 to 1996Q4. The economic variables that appear in the cointegration and error-correction modeling are the bond rate, the actual inflation rate, the nominal federal funds rate, and the output gap. The last two variables control for variations in the real component of the bond rate that are due to funds rate policy actions and the state of the economy. The test results indicate that the bond rate is cointegrated with the actual inflation rate during the full sample period, implying that the bond rate and the inflation rate move together in the long run. Permanent movements in the inflation rate are associated with permanent movements in the bond rate. The estimated error-correction model, however, indicates that a change has occurred in the way these two variables have adjusted in the short run. In the pre-1979 period, when the bond rate rose above the current inflation rate, actual future inflation accelerated. In the post-1979 period, however, the rise in the bond rate was reversed, and actual future inflation did not accelerate. Thus the bond rate signaled an acceleration in future inflation in the period before 1979 but not thereafter.

The results indicate that the above-noted change in the predictive content of the spread for future inflation may be due to change in Fed policy since 1979. In the post-1979 period, future inflation is inversely related to the current stance of Fed policy measured by the real funds rate, indicating Fed policy was geared towards reducing inflation. No such effect is found prior to 1979. Together these results are consistent with the hypothesis that after 1979, Fed policy prevented any increase in inflationary expectations (evidenced by

[3] In the context of rational expectations hypothesis tests, McCallum (1994) shows how the reduced-form regression coefficients depend upon the Fed's policy rule when the Fed smooths interest rates and responds to movements in the long-short spread.

the rise in the bond rate spread) that would have become embodied in higher actual future rates of inflation. As markets understand and believe in such Fed behavior, increases in inflationary expectations will be less common. The bond rate will then increasingly reflect phenomena other than expected inflation, thereby undermining its usefulness as a precursor of actual future inflation.

1. THE MODEL AND THE METHOD

The Fisher Relation, the Bond Rate, and Future Inflation

In order to motivate the empirical work, I discuss what the Fisher relation implies about the predictive content of the bond rate for future inflation. The Fisher relation for the m-period bond rate is

$$BR_t^{(m)} = rr_t^{(m)} + \dot{p}_t^{e(m)}, \tag{1}$$

where $BR^{(m)}$ is the m-period bond rate, $\dot{p}^{e(m)}$ is the m-period expected inflation rate, and $rr^{(m)}$ is the m-period expected real rate of interest. The Fisher relation (1) relates the bond rate to expectations of inflation and the real rate over the maturity (m) of the bond.

If the expected real interest rate is constant and if expectations of inflation are rational, then the Fisher relation above can be expressed as in (2) or (3):

$$BR_t^{(m)} = rr + \dot{p}_{t+m} - \epsilon_{t+m} \tag{2}$$

or

$$BR_t^{(m)} - \dot{p}_t = rr + (\dot{p}_{t+m} - \dot{p}_t) - \epsilon_{t+m}, \tag{3}$$

where rr is the constant real rate, $\dot{p}_{t+m}$ is the m-period future inflation rate, $\dot{p}_t$ is the one-period current inflation rate, and ϵ_{t+m} is the m-period future forecast error that is uncorrelated with past information. Equation (2) indicates that the bond rate contains information about the (m-period) future inflation rate, and equation (3) similarly shows that the spread between the bond rate and the current inflation rate has information about a change in the future inflation rate.

Testing the Predictive Content of the Bond Rate for Future Inflation

Previous Studies

Equations (2) and (3) above form the basis of empirical work in most previous studies of the predictive content of the bond rate for future inflation. Previous researchers have investigated the term structure's ability to predict future inflation by running regressions that are of the form

$$(\dot{p}_{t+m} - \dot{p}_{t+n}) = a + b\ (BR_t^{(m)} - BR_t^{(n)}) + \epsilon_{1t} \tag{4}$$

and

$$(\dot{p}_{t+m} - \dot{p}_t) = c + d\ (BR_t^{(m)} - \dot{p}_t) + \epsilon_{2t}, \tag{5}$$

where $BR^{(n)}$ is the n-period bond rate, $\dot{p}_{t+n}$ is the n-period future inflation rate, and other variables are as defined before. As can be seen, these regressions are merely rearranged versions of Fisher relations (2) and (3).[4] In (4) the spread between the m-period and n-period nominal interest rates is used to predict the difference between the m-period and n-period inflation rates, and in (5) the spread between the m-period bond rate and the (one-period) inflation rate is used to predict change in future inflation. Regressions like (4) appear in Mishkin (1990a, 1990b, 1991) and those like (5) appear in Engsted (1995). If $b \neq 0$ in (4) or $d \neq 0$ in (5), then that result indicates that the slope of the term structure does help predict future inflation.

But, as noted before, equations (2) and (3) (or regressions [4] and [5]) embody the assumption that the expected real interest rate is constant. This is a questionable assumption. Plosser and Rouwenhorst (1994) in fact present evidence that indicates that the long end of the term structure does seem to contain information about the real economic activity and therefore about the real rate of interest. If the expected real rate is not constant, then the disturbance term in these regressions ([4] or [5]) may be correlated with the spread.[5] In that case ordinary least squares may provide inconsistent estimates of parameters b and d, biasing inferences concerning the predictive content of the term structure for future inflation. Hence, in order to examine robustness to change in the assumption about the real rate, the predictive content should be investigated conditioning on variables that may control for potential short-term movements in the real rate.

Another issue not investigated fully in previous research is that slope parameters in (4) and (5) are likely to be influenced by the way the Fed conducts its monetary policy (McCallum 1994). For example, if the Fed has in place a disinflationary policy, then higher actual inflation may not follow a current increase in the bond rate spread (as in [5]). This could happen if current widening in the bond rate spread causes the Fed to raise the funds rate, leading to slower real growth and lower actual inflation in the future. In this scenario a current increase in the bond rate spread still reflects expectations of rising future

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[4] That is, we get (2) and (3) if we impose the restrictions $a = 0$, $b = 1$, $c = -rr$, and $d = 1$.

[5] To explain it further, assume the real rate c in (5) is not constant but in fact moves systematically with certain economic factors as follows:

$$c_t = c_0 + c_1 Z_t + u_t,$$

where Z is a set containing determinants of the real rate. If we replace c in (5) by c_t as above, then the disturbance term in (5) contains terms like $c_1 Z_t + u_t$. If the spread variable in (5) is correlated with the determinants Z_t, then the spread will be correlated with the disturbance term.

inflation. However, the ensuing Fed behavior prevents those expectations that would have been embodied in higher actual inflation. Therefore, in regressions like (5), the estimate of the slope parameter (d) may be small in periods during which the Fed has been vigilant. Those considerations suggest that parameters that measure the predictive content of the term structure for future inflation may not be stable during the sample period.

Cointegration and Error-Correction Modeling

The empirical work here examines the predictive content of the bond rate using cointegration and error-correction modeling. This empirical procedure, as I illustrate below, yields regressions that are similar in spirit to those employed in some previous research but differ in that it includes additional economic variables that control for potential movements in the real rate of interest.

As indicated before, the Fisher relation (1) for interest rates relates the bond rate to expectations of future inflation and the real interest rate. If one assumes that those expectations can be proxied by distributed lags on current and past values of actual inflation and other fundamental economic determinants, then the Fisher relation implies the following regression (6):

$$BR_t = a + \sum_{s=0}^{k} b_s \dot{p}_{t-s} + \sum_{s=0}^{k} c_s X_{t-s} + U_t, \tag{6}$$

where $\dot{p}_t$ is the actual inflation rate, X_t is the vector containing other economic determinants of the real rate, and U_t is the disturbance term. The presence of the disturbance term in (6) reflects the assumption that distributed lags on actual values of economic determinants may be good proxies for their anticipated values in the long run but not necessarily in the short run.[6]

If levels of the empirical measures of these economic determinants, including the bond rate, are unit root nonstationary, then the bond rate may be cointegrated with these variables as in Engle and Granger (1987). Under those assumptions, regression (6) can be reformulated as in (7):

$$BR_t = d_0 + d_1 \dot{p}_t + d_2 X_t + e_t. \tag{7}$$

Equation (7) is the cointegrating regression. The coefficients that appear on $\dot{p}_t$ and X_t in (7) then measure the long-run responses of the bond rate to inflation and its other real rate determinants. I investigate the question whether the bond rate incorporates expectations of future inflation by testing whether the bond rate is cointegrated with the actual inflation rate. My analysis thus views the positive relationship between the bond rate and actual inflation as a long-run phenomenon.

[6] The only assumption I make about the random disturbance term in (2) is that it has a zero mean.

The cointegrating bond rate regression defines the long-run equilibrium value of the bond rate. Should the bond rate rise above its long-run equilibrium value, then either the bond rate should fall, the economic determinants including inflation should adjust in the direction needed to correct the disequilibrium, or both. I examine such short-run dynamic adjustments by building a vector error-correction model that consists of short-run inflation and bond rate equations. The behavior of the error-correction variable, defined below, then provides information about ways the bond rate and inflation adjust in the short run. Therefore, if the error-correction term is positive and statistically significant in the short-run inflation equation, then that evidence can be interpreted to mean that the bond rate signals future inflation.[7]

To illustrate, assume that the bond rate depends only on the inflation rate in the long run and that the expected real rate is mean stationary. The cointegrating regression is then defined by the relation

$$BR_t = a + b\,\dot{p}_t + U_t, \tag{8}$$

where U_t is the short-term error. This variable, defined as the error-correction variable, measures the extent to which the bond rate differs from its long-run equilibrium value in the short run. The presence of cointegration implies the following error-correction model in ΔBR and $\Delta\dot{p}$:

$$\Delta BR_t = c_0 + \sum_{s=1}^{k} c_{1s}\Delta BR_{t-s} + \sum_{s=1}^{k} c_{2s}\Delta\dot{p}_{t-s} + \lambda_1 U_{t-1} + \epsilon_{1t} \tag{9a}$$

and

$$\Delta\dot{p}_t = d_0 + \sum_{s=1}^{k} d_{1s}\Delta BR_{t-s} + \sum_{s=1}^{k} d_{2s}\Delta\dot{p}_{t-s} + \lambda_2 U_{t-1} + \epsilon_{2t}, \tag{9b}$$

where U_{t-1} is the lagged value of the error-correction variable from (8) and where all other variables are as defined above. The presence of cointegration between BR_t and $\dot{p}_t$ implies that in (9) either $\lambda_1 \neq 0$, $\lambda_2 \neq 0$, or both. Thus, if λ_2 is positive and statistically significant, then a rise in the spread ($U_t = BR_t - a - b\,\dot{p}_t$) signals higher actual future inflation. Since the real interest rate is assumed to be mean stationary, not constant, the error-correction equations should be estimated including other (stationary) short-run determinants of the real interest rate.[8]

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[7] Miller (1991) has used this methodology to investigate short-run monetary dynamics.

[8] It is worth pointing out that Engsted (1995) uses an equation like (9b) to investigate whether the spread between the bond rate and the actual inflation rate (U_{t-1} in [8] here) helps predict future inflation. He, however, derives this implication of the Fisher hypothesis under the assumptions that expectations of inflation are rational and forward-looking and that the expected real interest rate is constant. To explain it further, consider the following version of the Fisher

Data and Definition of Economic Determinants in the Multivariable Analysis

The empirical work here examines the dynamic interactions between the bond rate and the inflation rate within a framework that allows for movements in the real component of the bond rate. The descriptive analysis of monetary policy in Goodfriend (1993) and the error-correction model of the bond rate estimated in Mehra (1994) indicate that the real component of the bond rate is significantly influenced by monetary policy actions and the state of the economy. Therefore, the economic variables that enter the analysis here are the bond rate, the actual inflation rate, the nominal federal funds rate, and the output gap that measures the state of the economy.

The empirical work uses quarterly data that spans the period 1959Q1 to 1996Q4. The estimation period, however, begins in 1961Q2, the earlier observations being used for lags. In addition, the sample is broken in 1979Q3, and results are provided for subperiods 1961Q2 to 1979Q3 and 1979Q4 to 1996Q4. The bond rate is the nominal yield on ten-year U.S. Treasury bonds (*BR*). Inflation as measured by the behavior of the consumer price index (excluding food and energy) is the actual, annualized quarterly inflation rate ($\dot{p}$). The measure of monetary policy used is the nominal federal funds rate (*NFR*), and the output gap (*gap*) is the natural lag of real GDP minus the natural log of potential GDP; the latter is generated using the Hodrick-Prescott filter (Hodrick and Prescott 1997).[9] The interest rate data are for the last month of the quarter.

Tests for Unit Roots and Cointegration

Cointegration and error-correction modeling involves four steps. First, determine the stationarity properties of the empirical measures of economic

hypothesis (1) for the long-term bond rate:

$$BR_{(t)} = rr + (1 - b) \sum_{j=1}^{\infty} b^j \; E_t \; \dot{p}_{t+j}, \tag{a}$$

where rr is the constant real rate and $b = \bar{e}^{\,i} \approx (1 + rr)$ is the discount factor (Engsted 1995). That is, the long bond rate is given as the constant real rate plus a weighted average of expected future one-period inflation rates ($E_t \; \dot{p}_{t+j},\ j \geq 1$). If BR_t and $\dot{p}_t$ are nonstationary and expectations are rational, then the above equation can be reformulated as

$$BR_t - b \; \dot{p}_t \equiv \; S_t = rr + \sum_{j=1}^{\infty} b^j \; E_t \Delta \dot{p}_{t+j}. \tag{b}$$

Equation (b) implies that BR_t and $b \; \dot{p}_t$ are cointegrated and that the spread $S_t = BR_t - b \; \dot{p}_t$ is an optimal predictor of future changes in inflation. Engsted (1995) examines the second implication by estimating a VAR in S and $\Delta\dot{p}$ and then testing whether S Granger-causes $\Delta\dot{p}$.

[9] I have examined the sensitivity of results to some changes in specification. For example, alternatively defining output gap relative to a linear trend produces qualitatively similar results.

determinants suggested above. Second, test for the presence of cointegrating relationships in the system. Third, estimate the cointegrating regression and calculate the residuals. Fourth, construct the short-run error-correction equations.

In order to determine whether the variables have unit roots or are mean stationary, I perform both unit root and mean stationarity tests. The unit root test used is the augmented Dickey-Fuller test, and the test for mean stationarity is the one advocated by Kwiatkowski, Phillips, Schmidt, and Shin (1992). Thus a variable X_t is considered unit root nonstationary if the hypothesis that X_t has a unit root is not rejected by the augmented Dickey-Fuller test and the hypothesis that it is mean stationary is rejected by the mean stationarity test.

The test used for cointegration is the one proposed in Johansen and Juselius (1990), and the cointegrating relations are identified by imposing restrictions as in Johansen and Juselius (1994). Also, the cointegrating relations are estimated using an alternative estimation methodology, Stock and Watson's (1993) dynamic OLS procedure.

2. EMPIRICAL RESULTS

Unit Root and Mean Stationarity Test Results

As indicated before, the economic variables that enter the analysis are the bond rate (*BR*), the inflation rate ($\dot{p}$), the nominal funds rate (*NFR*), and the output gap (*gap*). The output gap variable by construction is stationary. Table 1 reports test results for determining whether other variables have a unit root or are mean stationary. As can be seen, the t-statistic ($t_{\hat{\rho}}$) that tests the null hypothesis that a particular variable has a unit root is small for *BR*, $\dot{p}$, and *NFR*. On the other hand, the test statistic ($\hat{n}_u$) that tests the null hypothesis that a particular variable is mean stationary is large for all these variables. These results indicate that *BR*, $\dot{p}$, and *NFR* have a unit root and are therefore nonstationary in levels.

Cointegration Test Results

Table 2 presents test statistics for determining the number of cointegrating relations in the system (*BR*, $\dot{p}$, *NFR*, *gap*). Trace and maximum eigenvalue statistics presented in the table indicate that there are three cointegrating relations in the system.[10] This result holds in both the sample periods 1961Q2 to 1996Q4 and 1961Q2 to 1979Q3.

[10] The lag length parameter (k) for the VAR model was chosen using the likelihood ratio test described in Sims (1980). In particular, the VAR model initially was estimated with k set equal to a maximum number of eight quarters. This unrestricted model was then tested against a restricted model, where k is reduced by one, using the likelihood ratio test. The lag length finally selected in performing the Johansen-Juselius procedure is the one that results in the rejection of the restricted model.

Table 1 Tests for Unit Roots and Mean Stationarity

Series	Panel A Test for Unit Roots					Panel B Test for Mean Stationarity
X	ρ	$t_{\hat{\rho}}$	k	$x^2(2)$	$x^2(4)$	$\hat{n}_u$
BR	0.96	−1.7	5	1.6	1.3	0.80*
$\dot{p}$	0.89	−2.4	2	2.1	1.8	0.39**
NFR	0.89	−2.8	5	1.1	0.40	0.46*

*Significant at the 5 percent level.

**Significant at the 10 percent level.

Notes: BR is the bond rate; $\dot{p}$ is the annualized quarterly inflation rate measured by the behavior of the consumer price index excluding food and energy; and NFR is the nominal federal funds rate. The sample period studied is 1961Q2 to 1996Q4. ρ and t-statistics ($t_{\hat{\rho}}$) for $\rho = 1$ in panel A above are from the augmented Dickey-Fuller regressions of the form

$$X_t = a_0 + \rho\, X_{t-1} + \sum_{s-1}^{k} a_s\, \Delta X_{t-s},$$

where X is the pertinent series. The series has a unit root if $\rho = 1$. The 5 percent critical value is 2.9. The lag length k is chosen using the procedure given in Hall (1990), with maximum lag set at eight quarters. $x^2(2)$ and $x^2(4)$ are Chi-squared statistics that test for the presence of second-order and fourth-order serial correlation in the residual of the augmented Dickey-Fuller regression, respectively. The test statistics $\hat{n}_u$ in panel B is the statistic that tests the null hypothesis that the pertinent series is mean stationary. The 5 percent critical value for $\hat{n}_u$ given in Kwiatkowski et al. (1992) is 0.463 (10 percent critical value is 0.347).

Table 3 presents estimates of the cointegrating relations found in the system. I first test the hypothesis that the three-dimensional cointegration space contains cointegrating relations that are of the form (10) through (12):

$$BR_t = a_0 + a_1\, \dot{p}_t + u_{1t};\ a_1 = 1, \tag{10}$$

$$NFR_t = b_0 + b_1\, \dot{p}_t + u_{2t};\ b_1 = 1, \tag{11}$$

and

$$gap_t = c_0 + u_{3t}. \tag{12}$$

Equation (10) can be interpreted as the Fisher relation for the bond rate and equation (11) as the Fed reaction function. Equation (12) simply states that the output gap variable is stationary. As shown in Johansen and Juselius (1994), these cointegrating relations can be identified imposing restrictions on long-run parameters in the cointegrating space.

In the full sample period, the hypotheses that cointegrating relations are of the form (10) through (12) and that $a_1 = b_1 = 1$ are consistent with data (the x_1^2 statistic that tests those restrictions is small; see Table 3, panel A). However, in the subsample 1961Q2 to 1979Q3, the restrictions that $a_1 = b_1 = 1$ are

Table 2 Cointegration Test Results

Panel A: 1961Q2 to 1996Q4

System	**Trace H0**		**Maximum Eigenvalue H0 vs H1**	***k***
(BR, $\dot{p}$, NFR, gap)	$r = 0$	8.9*	$r = 0$ vs $r \leq 1$: 28.6*	8
	$r \leq 1$	40.3*	$r = 1$ vs $r \leq 2$: 23.9*	
	$r \leq 2$	16.3*	$r = 2$ vs $r \leq 3$: 11.7*	
	$r \leq 3$	4.6	$r = 3$ vs $r \leq 4$: 4.6	

Panel B: 1961Q2 to 1979Q3

System	**Trace H0**		**Maximum Eigenvalue H0 vs H1**	***k***
(BR, $\dot{p}$, NFR, gap)	$r = 0$	66.2*	$r = 0$ vs $r \leq 1$: 33.4*	5
	$r \leq 1$	32.8*	$r = 1$ vs $r \leq 2$: 19.8*	
	$r \leq 2$	13.0*	$r = 2$ vs $r \leq 3$: 10.6*	
	$r \leq 3$	2.5	$r = 3$ vs $r \leq 4$: 2.5	

*Significant at the 10 percent level.

Notes: Trace tests the null hypothesis that the number of cointegrating vectors (r) is less than or equal to a chosen value, and maximum eigenvalue tests the null hypothesis that the number of cointegrating vectors is r, given the alternative of $r + 1$ vectors. The VAR lag length (k) was chosen using the likelihood ratio test in Sims (1980).

rejected by data, and the cointegrating relations are thus estimated without such restrictions.[11] As can be seen, estimates indicate that the bond rate is cointegrated with the inflation rate, but the bond rate does not adjust one-for-one with inflation. Therefore, inflation is the main source of the stochastic trend in the bond rate.

The estimation procedure in Johansen and Juselius (1990, 1994) is a system estimation method. In order to check the robustness of estimates, I also present estimates of the cointegrating relations (10) and (11) using a single equation estimation method. Panel B in Table 3 presents results using the dynamic OLS procedure given in Stock and Watson (1993). As shown in the table, this procedure yields estimates that are remarkably close to those reported above.

Results on the Error-Correction Coefficient in the Error-Correction Model

The cointegration test results described in the previous section are consistent with the presence of cointegrating relations that are of the form

[11] In estimating error-correction equations for the pre-1979 period, I consider cointegrating regressions with $a = b = 1$. This restriction implies that the bond rate does adjust one-for-one with inflation. The basic results do not change if instead this restriction is not imposed (see footnote 15).

Table 3 Estimates of Restricted Cointegrating Vectors

	Panel A: Johansen-Juselius Procedure	
	Sample Period 1961Q2 to 1996Q4	**Sample Period 1961Q2 to 1979Q3**
A1	$BR_t = 3.1 + \dot{p}_t + U_{1t}$	$BR_t = 3.2 + 0.67\ \dot{p}_t + U_{1t}$
A2	$NFR_t = 2.3 + \dot{p}_t + U_{2t}$	$NFR_t = 2.7 + 0.66\ \dot{p}_t + U_{2t}$
	$x_1^2(3) = 0.92\ (0.82)$	$x_2^2(1) = 0.01\ (0.91)$
	Panel B: Dynamic OLS	
	1961Q2 to 1996Q4	**1961Q2 to 1979Q3**
A1	$BR_t = 2.9 + 1.0\ \dot{p}_t + U_{1t}$	$BR_t = 3.2 + 0.66\ \dot{p}_t + U_{1t}$
A2	$NFR_t = 2.2 + 1.0\ \dot{p}_t + U_{2t}$	$NFR_t = 2.5 + 0.67\ \dot{p}_t + U_{2t}$

Notes: Panel A above reports two of the three cointegrating vectors that lie in the cointegration space spanned by the four-variable VAR ($BR, \dot{p}, NFR, gap$). The cointegrating vectors A1 and A2 are the Fisher relation for the bond rate and the funds rate. $x_1^2(3)$ and $x_2^2(1)$ are Chi-squared statistics (degrees of freedom in parentheses) that test the null hypothesis that the identifying restrictions imposed are consistent with data (Johansen and Juselius 1994).

Panel B above reports the same cointegrating vectors estimated using the dynamic OLS procedure (eight leads and lags are used).

$$BR_t = a_0 + a_1\ \dot{p}_t + U_{1t} \tag{13}$$

and

$$NFR_t = b_0 + b_1\ \dot{p}_t + U_{2t}, \tag{14}$$

where U_1 and U_2 are stationary disturbance terms. I now examine the behavior of the error-correction term $U_{1t} = BR_t - a_0 - a_1\dot{p}_t$ in short-run equations of the form

$$\Delta BR = b_0 + \sum_{s=1}^{k1} b_{1s}\ \Delta BR_{t-s} + \sum_{s=1}^{k2} b_{2s}\Delta\dot{p}_{t-s} + \sum_{s=1}^{k3} b_{3s}\Delta NFR_{t-s} + \sum_{s=1}^{k4} b_{4s}\ gap_{t-s} + \lambda_1\ U_{1t-1} + \delta_1\ U_{2t-1} \tag{15a}$$

and

$$\Delta\dot{p}_t = c_0 + \sum_{s=1}^{k1} c_{1s}\ \Delta BR_{t-s} + \sum_{s=1}^{k2} c_{2s}\Delta\dot{p}_{t-s} + \sum_{s=1}^{k3} c_{3s}\Delta NFR_{t-s} + \sum_{s=1}^{k4} c_{4s}\ gap_{t-s} + \lambda_2\ U_{1t-1} + \delta_2\ U_{2t-1}, \tag{15b}$$

where all variables are as defined before. The short-run equations include first differences of the bond rate, inflation, and the funds rate and level of the output gap, even though the last two variables do not enter the long-run bond equation (13). These variables capture the short-run impacts of monetary policy and the state of the economy on the bond rate and other variables. As indicated before, the parameters of interest are λ_1, λ_2 and the sums of coefficients that appear on the bond rate in equation (15b). The expected signs of the error-correction term U_{1t-1} are positive for $\Delta\dot{p}$ and negative for ΔBR.

Following Campbell and Perron (1991), the lag lengths used in the error-correction model are chosen using the procedure given in Hall (1990). This procedure starts with some upper bound on lags, chosen a priori for each variable (eight quarters here) and then drops all lags beyond the lag with a significant coefficient. I do present tests of the hypothesis that excluded lags are not significant, however.

Table 4 reports the error-correction coefficients (t-values in parentheses) when the long-run bond equation is (13). In addition, it also reports the sums of coefficients that appear on (first differences of) the bond rate in the inflation equation. Parentheses that follow contain t-statistics for the sum of coefficients, whereas brackets contain Chi-squared statistics for exclusion restrictions. Panel A reports results for the full sample 1961Q2 to 1996Q4 and panels B and C for the subperiods 1961Q2 to 1979Q3 and 1979Q4 to 1996Q4.[12] In full sample regressions the error-correction coefficient is negative and statistically significant in the bond equation (ΔBR), but in inflation equations ($\Delta\dot{p}$), it is generally small and not statistically different from zero.[13] Furthermore, individual coefficients that appear on two lagged values of the bond rate in the inflation equation are 0.50 and -0.33. These coefficients are individually significant, but their sum is not statistically different from zero, indicating that ultimately, increases in the bond rate have not been associated with accelerations in actual inflation.[14] Together, these results indicate that the short-run positive deviations of the bond rate from its long-run equilibrium values were corrected mainly through reversals in the bond rate. Actual inflation did not accelerate.

The results for the first subperiod 1961Q2 to 1979Q3 reported in panel B of Table 4 are, however, strikingly different. As can be seen, the error-correction coefficient is negative and significant in the bond rate equation but is positive and significant in the inflation equation. These results suggest that

[12] Inflation equations include dummies for President Nixon's price and wage controls.

[13] The error-correction coefficients are in fact negative in the inflation equation that includes other determinants of the real rate. In the inflation equation that includes only lagged values of inflation, the coefficient that appears on the error-correction term is positive, small in magnitude, and not statistically different from zero. The latter result is similar in spirit to the one in Engsted (1995).

[14] This result, of course, means that the bond rate Granger-causes inflation.

Table 4 Granger-Causality Results from Error-Correction Equations: General to Specific, Using Hall Approach

Panel A: Cointegrating Regressions, 1961Q2 to 1996Q4

$$BR_t = 2.9 + \dot{p}_t + U_{1t};\ NFR_t = 2.2 + \dot{p}_t + U_{2t}$$

Equation	U_{1t-1}	$\sum_{s=1}^{k1} \Delta BR_{t-s}$	(k_1, k_2, k_3, k_4)	$x^2(sl)$
ΔBR_t	−0.20 (3.5)		(7,7,8,1)	9.5 (0.39)
$\Delta \dot{p}_t$	−0.13 (1.3)	0.17 (0.6) [10.2]*	(2,8,8,8)	5.3 (0.51)

Panel B: Cointegrating Regressions, 1961Q2 to 1979Q3

$$BR_t = 1.7 + \dot{p}_t + U_{1t};\ NFR_t = 1.0 + \dot{p}_t + U_{2t}$$

Equation	U_{1t-1}	$\sum_{s=1}^{k1} \Delta BR_{t-s}$	(k_1, k_2, k_3, k_4)	$x^2(sl)$
ΔBR_t	−0.24 (3.5)		(8,7,6,1)	8.9 (0.54)
$\Delta \dot{p}_t$	0.32 (3.2)		(0,0,0,0)	38.4 (0.24)

Panel C: Cointegrating Regressions, 1979Q4 to 1996Q4

$$BR_t = 4.2 + \dot{p}_t + U_{1t};\ NFR_t = 2.5 + \dot{p}_t + U_{2t}$$

Equation	U_{1t-1}	$\sum_{s=1}^{k1} \Delta BR_{t-s}$	(k_1, k_2, k_3, k_4)	$x^2(sl)$
ΔBR_t	−0.39 (2.9)		(7,0,6,8)	14.5 (0.21)
$\Delta \dot{p}_t$	−0.01 (0.4)		(0,6,8,8)	11.4 (0.33)

Notes: The coefficients reported are from error-correction regressions that include the bond rate (*BR*), the inflation rate ($\dot{p}$), the nominal federal funds rate (*NFR*), and the output gap (*gap*) (see equation [15] of the text). In addition, the model has two error-correction variables (U_{1t} and U_{2t}). (k_1, k_2, k_3, k_4) refers to lag lengths that are chosen for $BR, \dot{p}, NFR$, and *gap*. Parentheses contain t-statistics for the error-correction variable (U_{1t-1}) or for the sum of coefficients that appear on the bond rate $\left(\sum_{s=1}^{k1} \Delta BR_{t-s}\right)$. For the latter, brackets contain the Chi-squared statistic for the null hypothesis that every coefficient in this sum is zero. $x^2(sl)$ tests the null hypothesis that remaining lags are not significant (significance levels follow in parentheses).

positive deviations of the bond rate from its long-run equilibrium value were eliminated partly through declines in the bond rate and partly through increases in actual inflation. Consequently, in the pre-1979 period, actual inflation did accelerate when the spread between the bond rate and the one-period inflation rate rose.[15]

[15] I get similar results if cointegrating regressions (13) and (14) are estimated without restrictions $b_1 = a_1 = 1$. In particular, over the sample period 1961Q2 to 1979Q3, the error-correction variable U_{1t-1} has a positive coefficient in the inflation equation, indicating that actual inflation did accelerate following an increase in the bond rate spread.

In the aforementioned result, the fact that the spread signaled an increase in inflation in the pre-1979 period but not in the full sample period implies that the spread must have lost its predictive content in the post-1979 period. This implication is consistent with the subperiod results reported in panel C of Table 4: the error-correction term is no longer significant in the inflation equation.

Comparison with Previous Studies

The full sample results discussed in the previous section indicate that the spread between the bond rate and the one-period inflation rate does not help predict one-quarter-ahead changes in the rate of inflation. Since inflation is a unit root process, the results above also imply that the spread has no predictive content for long-horizon forecasts of future inflation. The latter implication is in contrast with the finding in Mishkin (1990a, 1990b, 1991) that at long horizons the long end of the slope of the term structure does help predict future inflation.

As indicated before, an important assumption implicit in the regressions used by Mishkin is that the ex ante real rate of interest is constant. This assumption may not be valid. Therefore, the predictive content of the spread for future inflation should also be investigated conditioning on variables that capture changes in the short-run determinants of the real rate.

In order to illustrate whether results are sensitive to such conditioning, I also investigate the predictive content of the spread between the bond rate and the (one-period) inflation rate for future inflation by estimating regressions of the form

$$(\ln[P_{t+m}/P_t]/m) - \ln(P_t/P_{t-1}) = a_0 + \lambda_c\ U_{1t} + V_{1t}, \tag{16}$$

$$\begin{aligned}(\ln[P_{t+m}/P_t]/m) - \ln(P_t/P_{t-1}) &= a_0 + \lambda_d\ U_{1t} + \sum_{s=1}^{k1} a_{1s}\ \Delta\dot{p}_{t-s} \\ &+ \sum_{s=1}^{k2} a_{2s}\ \Delta NFR_{t-s} + \sum_{s=1}^{k3} a_{3s}\ \Delta BR_{t-s} + \sum_{s=1}^{k4} a_{4s}\ gap_{t-s} + V_{2t},\end{aligned} \tag{17}$$

and

$$\begin{aligned}(\ln[P_{t+m}/P_t]/m - \ln(P_t/P_{t-1}) &= a_0 + \lambda_e\ U_{1t} + \delta U_{2t} + \sum_{s=1}^{k1} a_{1s}\ \Delta\dot{p}_{t-s} \\ &+ \sum_{s=1}^{k2} a_{2s}\ \Delta NFR_{t-s} + \sum_{s=1}^{k3} a_{3s}\ \Delta BR_{t-s} + \sum_{s=1}^{k4} a_{4s}\ gap_{t-s} + V_{2t},\end{aligned} \tag{18}$$

where

$$U_{1t} = BR_t - a_0 - a_1\ \dot{p}_t,$$

$$U_{2t} = NFR - b_0 - b_1\ \dot{p}_t,$$

and where m is the number of quarters, and other variables are as defined.[16] U_1 measures the spread between the bond rate and the (one-period) inflation rate and U_2 the spread between the nominal funds rate and the inflation rate. Regression (16) examines the predictive content of the spread for long-horizon forecasts of future inflation without controlling for variations in the spread due to real growth, monetary policy actions, and inflation. Regressions (17) and (18), however, control for such variations. Regression (18) is similar to regression (17) except in that it also includes the current stance of short-run monetary policy measured by the funds rate spread (U_{2t}). The regressions are estimated over the full sample period as well as over subperiods 1961Q2 to 1979Q3 and 1979Q4 to 1996Q4 and for horizons up to four years in the future. In addition, I consider the subperiod 1983Q1 to 1996Q4, during which inflation has remained relatively low.

In Tables 5 and 6, I present estimates of the coefficient (t-values in parentheses) that appears on the bond rate spread variable (λ_c in [16]), λ_d in [17], and λ_e in [18]).[17,18] I also report the coefficient on the funds rate spread variable (δ in [18]). In Table 5 the results are for the full sample period and the first subperiod and in Table 6 for two post-1979 subperiods. If we focus on pre-1979 regression estimates, we will see that they indicate that the bond rate spread does help predict future inflation (see t-values on λ_c, λ_d, and λ_e in Table 5, panel B). This result holds at all forecast horizons and is not sensitive to the inclusion of other variables in regressions. Furthermore, the funds rate spread variable that controls for policy-induced movements in the real component of the bond rate is never significant in those regressions, indicating that at the time the current stance of monetary policy had no predictive content for future inflation. Therefore, the widened bond rate spread was followed by higher actual future inflation during this subperiod.

[16] These regressions differ from those reported in Mishkin (1990a, 1990b, 1991). Mishkin uses zero-coupon bond data, derived from coupon-bearing bonds that have actually been traded. So, he is able to match the horizon of the inflation forecast with that of the term spread. The empirical work here instead uses yield-to-maturity data on coupon bonds and the inflation forecast horizon does not match with that of the term spread. These differences, however, do not reduce the importance of examining the potential role of additional variables that may provide information about movements in the real rate of interest.

[17] The t-values were corrected for the presence of moving-average serial correlation generated due to overlap in forecast horizon. The degree of correction in the moving-average serial correlation was determined by examining the autocorrelation function of the residuals. This procedure generated the order of serial correlation correction close to the value given by $(m-1)$, where m is the number of quarters in the forecast horizon. Furthermore, the use of realized multi-period inflation in these regressions led to the loss of observations at the end of the sample, so that the effective sample sizes are 1961Q2 to 1996Q4-m and 1961Q2 to 1979Q3-m.

[18] All regressions are estimated including four lagged values of other information variables. Furthermore, those lagged values are always statistically significant as a group in regressions (17) and (18).

Table 5 Long-Horizon Inflation Equations

Horizons in Quarters (m)	Equation (C) λ_c (t-value)		Equation (D[c]) λ_d (t-value)		Equation (E) λ_e (t-value)		δ (t-value)	
Panel A: 1961Q2 to 1996Q4								
Cointegrating Regressions: $BR_t = 2.9 + \dot{p}_t + U_{1t}$; $NFR_t = 2.2 + \dot{p}_t + U_{2t}$								
4	0.16	(1.5)	0.09	(1.0)	0.02	(0.2)	0.07	(0.6)[a]
8	0.20	(1.5)	0.04	(0.4)	0.14	(0.9)	−0.12	(1.0)[a]
12	0.24	(1.6)	0.01	(0.1)	0.21	(1.2)	−0.26	(1.4)[a]
16	0.25	(1.7)	−0.07	(0.6)	0.22	(1.2)	−0.33	(1.7)[a]
Panel B: 1961Q2 to 1979Q3								
Cointegrating Regressions: $BR_t = 1.7 + \dot{p}_t + U_{1t}$; $NFR_t = 1.0 + \dot{p}_t + U_{2t}$								
4	0.56	(5.4)	0.61	(7.4)	0.62	(3.3)	−0.00	(0.0)[b]
8	0.81	(7.9)	0.80	(6.5)	1.0	(4.4)	−0.25	(0.8)[b]
12	0.96	(7.9)	0.89	(12.6)	0.94	(3.4)	−0.10	(0.2)[b]
16	0.99	(9.1)	0.99	(13.2)	0.77	(2.8)	0.33	(0.9)[b]

[a]The restriction $\lambda_e + \delta = 0$ is consistent with data.

[b]The restriction $\lambda_e + \delta = 0$ is not consistent with data.

[c]Additional variables included in equations (D) and (E) are always statistically significant as a group.

Notes: The coefficients reported are from regressions of the form

$$p(t,m) = f_0 + \lambda_c\ U_{1t}, \tag{C}$$

$$p(t,m) = g_0 + \lambda_d\ U_{1t} + \sum_{s=1}^{k1} g_{1s}\ \Delta BR_{t-s} + \sum_{s=1}^{k2} g_{2s}\ \Delta\dot{p}_{t-s} + \sum_{s=1}^{k3} g_{3s}\ \Delta NFR_{t-s} + \sum_{s=1}^{k4} g_4\ gap_{t-s} \tag{D}$$

and

$$p(t,m) = \lambda_e\ U_{1t} + \delta\ U_{2t} + \text{ other variables as in (D)}, \tag{E}$$

where $p(t,m)$ is $(log[P_{t+m}/P_t])/m - log(P_t/P_{t-1})$, m is the number of quarters in the forecast horizon, and the rest of the variables are as defined before. All regressions are estimated setting $k_1 = k_2 = k_3 = k_4 = 4$.

The full sample regression estimates, however, suggest strikingly different results. The coefficient that appears on the bond rate spread variable is now about one-third the size estimated in subsample regressions.[19] For forecast

[19] Mishkin (1990a) also finds that in full sample regressions the coefficients that appear on term spreads are generally smaller in size than those in pre-1979 regressions. Nonetheless,

Table 6 Long-Horizon Inflation Equations

Cointegrating Regressions: $BR_t = 4.2 + \dot{p}_t + U_{1t}$; $NFR_t = 2.5 + \dot{p}_t + U_{2t}$

Panel A: 1979Q4 to 1996Q4

Horizons in Quarters (m)	Equation (C) λ_c (t-value)		Equation (D[a]) λ_d (t-value)		Equation (E) λ_e (t-value)		δ (t-value)	
4	0.21	(2.0)	0.18	(3.0)	0.29	(2.7)	−0.10	(1.1)[b]
8	0.31	(2.1)	0.28	(3.6)	0.58	(3.7)	−0.26	(2.2)[b]
12	0.35	(2.0)	0.31	(4.1)	0.61	(5.9)	−0.28	(2.9)[b]
16	0.42	(2.2)	0.37	(6.1)	0.73	(6.3)	−0.35	(3.5)[b]

Panel B: 1983Q1 to 1996Q4

Horizons in Quarters (m)	Equation (C) λ_c (t-value)		Equation (D[a]) λ_d (t-value)		Equation (E) λ_e (t-value)		δ (t-value)	
4	0.08	(0.4)	0.14	(2.0)	0.08	(0.3)	0.07	(0.07)[b]
8	0.08	(0.9)	0.17	(2.7)	0.10	(4.4)	0.10	(1.1)[b]
12	0.11	(0.9)	0.21	(3.9)	0.30	(3.6)	−0.11	(1.7)[b]
16	0.14	(0.7)	0.28	(4.4)	0.42	(6.2)	−0.22	(2.9)[b]

[a]Additional variables included in equations (D) and (E) are always statistically significant as a group.

[b]The restriction $\lambda_e + \delta = 0$ is not consistent with data.

Notes: The cointegrating regressions are estimated over the period 1979Q4 to 1996Q4. The coefficients reported above are from regressions like those given in Table 5. See notes in Table 5.

horizons up to three years in the future, this coefficient is not statistically significant, and for somewhat longer horizons, it is marginally significant at the 10 percent level (see t-values on λ_c, λ_d, or λ_e in Table 5, panel A). Those estimates suggest there had been a significant deterioration in the predictive content of the bond rate spread for future inflation in the period since 1979. Furthermore, results are now sensitive to variables included in the conditioning set. If we ignore the current stance of Fed policy measured by the funds rate spread, then the bond rate spread has no predictive content for actual future inflation at any forecast horizon (see λ_d in Table 5, panel A). However, when the funds rate spread variable is included in the conditioning set, then in long-horizon inflation regressions, the bond rate spread variable appears with a positive coefficient. Yet in those same regressions the coefficient that appears on the funds rate spread is negative though barely statistically significant (see δ in Table 5, panel

his regressions pass the conventional test of parameter stability. The regressions estimated here, however, do not depict such parameter constancy.

A). This result is consistent with the presence of policy-induced movements in the real component of the funds rate and their subsequent negative effects on future inflation rates. In fact, the coefficients that appear on the bond rate and the funds rate spreads are equal in size but opposite in signs. Those estimates suggest that increases in the bond rate spread accompanied by equivalent increases in the funds rate spread have had no effect on actual future inflation rates.[20]

The first subperiod and full sample results discussed above suggest the presence of considerable subperiod instability. In order to gain more insight into subperiod differences, Table 6 presents estimates from long-horizon inflation equations for two post-1979 subperiods, 1979Q4 to 1996Q4 and 1983Q1 to 1996Q4. Those estimates permit the following inferences about the predictive content of the spread for future inflation in the post-1979 period. First, the predictive content of the spread for future inflation has deteriorated in the post-1979 period. The size of the coefficient that appears on the bond rate spread variable is not only small relative to its value found in the pre-1979 period, but its size declines further during the relative low inflation period of the late 1980s and the 1990s (compare values of λ_c in Tables 5 and 6). Second, the marginal predictive content of the spread for future inflation is now sensitive to variables included in regressions (compare values of λ_c and λ_e in panels A and B of Table 6). Third, the current stance of monetary policy measured by the funds rate spread correlates negatively with future inflation, indicating Fed policy was geared towards reducing inflation in the post-1979 period. In those regressions the bond rate spread variable remains significant, indicating the bond rate spread does contain information about future inflation. However, the results also indicate that actual future inflation may not accelerate following the rise in the bond rate spread if the Fed reacts aggressively by raising the funds rate (see estimates of λ_e and δ in Table 6).

The descriptive analysis of monetary policy in Goodfriend (1993) in fact indicates that since 1979 the Fed has had a disinflationary policy in force to reduce the trend rate of inflation and contain inflationary expectations. Accordingly, this Fed behavior may be at the source of deterioration in the predictive content of the bond rate for actual future inflation. To the extent that rising long-run inflationary expectations evidenced by the rise in the bond rate were triggered in part by news of strong actual or anticipated real growth, the Fed may have calmed those expectations by raising the funds rate. The induced tightening of monetary policy may have reduced inflationary expectations by reducing actual or anticipated real growth, thereby preventing any increase in actual inflation. Given such Fed behavior, observed increases in the bond rate

[20] This result is similar in spirit to the finding reported using cointegration and error-correction methodology.

do not necessarily indicate that actual inflation is going to accelerate in the near term.

3. CONCLUDING OBSERVATIONS

This article views the Fisher hypothesis as a long-run relationship with short-run variation in the real interest rate. The findings show that the bond rate is cointegrated with the inflation rate over the 1962Q2 to 1996Q4 period, which indicates that in the long run, permanent movements in actual inflation have been associated with permanent movements in the bond rate.

The short-run error-correction equations help identify ways in which the bond rate and inflation adjust in the short run. In the pre-1979 period, increases in the bond rate were followed by an acceleration in actual inflation, whereas that did not happen in the post-1979 period. In the latter period, short-run increases in the bond rate have usually been reversed, with no follow-up in actual inflation.

In the period since 1979, the Fed has made serious attempts to reduce the trend rate of inflation and contain inflationary expectations. Such Fed behavior may have prevented the short-run increases in inflationary expectations, as evidenced by increases in the bond rate, from finally producing higher actual inflation. These results imply that if the Fed retains its hard-won credibility for inflation stability, then the bond rate may reflect phenomena other than expected inflation, thereby undermining its usefulness as a precursor of actual future inflation.

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