Monetary Policy and Long-Term Interest Rates

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The standard view of the transmission mechanism of monetary policy assigns a key role to long-term interest rates. According to this view, a monetary policy tightening pushes up both short and long interest rates, leading to less spending by interest-sensitive sectors of the economy and therefore to lower real growth. Conversely, a monetary easing results in lower interest rates that stimulate real growth. An open question in discussions of this view is whether monetary policy has significant empirical effects on long-term interest rates.[1]

In this article, I provide new evidence on the quantitative effect of monetary policy on the long-term interest rate. The federal funds rate is used as a measure of monetary policy.[2] The work extends the previous research in

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[1] In most previous studies using an interest-rate-based measure of monetary policy, a short-term money market rate (the three-month or one-year Treasury bill rate) is used. Most of those studies surveyed recently in Akhtar (1995) find significant and large effects of short rates on long rates. In those studies the long-run response of nominal long rates ranges from about 22 to 66 basis points for every one percentage point change in nominal short rates. However, there is considerable skepticism about the reliability and interpretation of those effects. One main reason for such skepticism is even though monetary policy has its strongest effect on a short-term money market rate, the latter is also influenced by nonmonetary forces. Hence changes in short rates do not necessarily reflect changes in the stance of monetary policy.

There are a few other empirical studies that use the federal funds rate as a measure of monetary policy. But most of those studies examine the effect of policy on the long rate in a bivariable framework. In such studies the estimated impact of policy on the long rate is quantitatively modest and temporally unstable (see Akhtar 1995, Table 3). An exception is the recent work in Mehra (1994) which uses a multivariable framework and finds a significant effect of the real funds rate on the long rate. However, Mehra (1994) does not investigate the robustness of those results to alternative specifications or to different sample periods.

[2] Recent research has shown that the federal funds rate is a good indicator of the stance of monetary policy (Bernanke and Blinder 1992; Bernanke and Mihov 1995).

two main directions. First, following Goodfriend's (1993) description of funds rate policy actions, it distinguishes empirically between the long- and short-run sources of interaction between the funds rate and the long rate; this distinction is absent in previous studies. Second, the analysis in Goodfriend also suggests that the near-term effects of funds rate policy actions on the long rate may be quite variable. Hence the work examines the temporal stability of such effects, an issue also virtually ignored in previous research.

The empirical work focuses on the behavior of the nominal yield on ten-year U.S. Treasury bonds during the period 1957Q1 to 1995Q2. The results here indicate that the bond rate moves positively with the funds rate in the long run. However, this comovement arises because the bond rate automatically moves with trend inflation (the Fisher relation) and the Federal Reserve (Fed) keeps the level of the funds rate in line with the going trend rate of inflation (the long-run Fed reaction function). Apart from the correlation that occurs through the inflation channel, I find that empirically there is no other source of long-run interaction between the bond rate and the funds rate. This result arises because the bond rate's other component—the expected long real rate—is mean stationary and therefore unrelated to the level of the funds rate in the long run. These results have the policy implication that monetary policy can permanently lower the bond rate only by lowering the trend rate of inflation.

The short-run stance of monetary policy is measured by the spread between the funds rate and the trend rate of inflation (the funds rate spread). The results indicate that movements in the funds rate spread have a statistically significant effect on the bond rate and that the magnitude of its near-term effect on the bond rate has increased significantly since 1979.[3] In the pre-1979 period, the bond rate rose anywhere from 14 to 29 basis points whenever the funds rate spread widened by one percentage point. In the post-1979 period, however, the estimate of its near-term response ranges from 26 to 50 basis points.

The short-run results thus suggest that, ceteris paribus, a monetary policy tightening measured by an increase in the funds rate spread does result in higher bond rates in the short run, in line with the traditional view of the transmission mechanism. However, this increase in the short-run sensitivity of the bond rate to policy actions may itself be due to the way the Fed has conducted its monetary policy since 1979. The Fed's post-1979 efforts to bring down the trend rate of inflation, coupled with lack of credibility, may have amplified the near-term effects of funds rate changes on the bond rate.

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[3] In this article near-term effects refer to responses of the bond rate to recent past values of the funds rate spread. The immediate effect is the response to the one-period lagged value of the spread and the near-term effect is the cumulative response to all such past values. What I call the near-term effect is sometimes referred to as the long-run effect in previous studies. As indicated later, I use the long-run effect to measure the effect that arises from the existence of equilibrium or trending relationships among nonstationary variables.

The plan of this article is as follows. Section 1 presents the methodology that underlies the empirical work, Section 2 contains empirical results, and concluding observations are in Section 3.

1. THE MODEL AND THE METHOD

This section describes the methodology that underlies the empirical work in this article.

The Fisher Relation, the Bond Rate, and the Federal Funds Rate

In order to motivate the empirical work, I first describe how monetary policy may affect the bond rate in the short run and the long run.[4] As indicated before, the federal funds rate is used as a measure of the stance of monetary policy. Thus a monetary policy action is defined as a change in the funds rate, and a monetary policy strategy is defined as the reaction function that would lead to policy actions. The Fisher relation for interest rates provides a convenient framework within which effects of policy actions can be described. The Fisher relation is

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

where BR_t is the bond rate, rr_t^e is the expected long real rate, and $\dot{p}_t^e$ is the expected long-term inflation rate. Equation (1) relates the bond rate to expectations of inflation and the real rate.

The Fisher relation indicates that policy actions could conceivably affect the bond rate by altering expectations of inflation, the real rate, or both. Since policy actions may not always move the real rate and inflation components in the same direction, the near-term responses of the bond rate to such actions cannot be determined a priori. Much may depend upon the nature of the strategy being pursued by the Fed. Goodfriend (1993) discusses three different strategies that may lie at the source of interaction between the bond rate and the funds rate. Consider, first, pure cyclical strategies in which the Fed routinely raises (lowers) the funds rate in response to cyclical expansions (downturns) without attempting to affect the current trend rate of inflation expected by the public. Under that strategy, a funds rate increase will tend to raise the bond rate by raising current and expected future short real rates (i.e., by raising the rr_t^e component in [1]). This cyclical comovement is short run in nature.

The second strategy discussed by Goodfriend considers the response of the Fed to an exogenous change in the trend rate of inflation. If the trend rate of inflation increases, the bond rate automatically moves with inflation (equation

[4] The discussion in this section draws heavily from Goodfriend (1993).

[1]). The Fed may choose to keep the short real rate steady and will therefore move the funds rate in line with the rising or falling inflation rate. In this case the bond rate comoves with the funds rate because the Fed is responding to changing inflation in a neutral fashion. I refer to this source of comovement as long run in nature.

Finally, consider an aggressive strategy that could be taken either to promote real growth or to reduce the going trend rate of inflation. Under that strategy, the net impact of policy actions on the bond rate is complex because they can move the real rate (rr_t^e) and the inflation expectations ($\dot{p}_t^e$) in opposite directions. The real rate effect moves the bond rate in the same direction as the funds rate, while the inflation effect moves the bond rate in the opposite direction. Thus the net effect of policy actions on the long rate cannot be determined a priori.

To illustrate, consider an aggressive increase in the funds rate intended to reduce the trend rate of inflation. Such a tightening can shift both components of the bond rate. If the Fed's disinflation policy is credible, then short rates rise and expected inflation falls. The fall in expected inflation may thus offset somewhat the immediate response of the bond rate to the funds rate. If the decline in expected inflation persists, then the Fed may soon bring down the funds rate consistent with the lower trend rate of inflation. However, if the public does not yet have full confidence in the Fed's disinflation, then the Fed may have to persist with a sufficiently high funds rate until real growth slows and inflation actually declines. In this case, the immediate and near-term effects of the funds rate on the bond rate may be large relative to the previous case. These policy actions generate correlations between the bond rate and the funds rate which are both short and long run in nature.

Empirical Specifications of the Bond Rate Regressions: Short- and Long-Run Effects

The discussion in the previous section suggests the following observations. First, the bond rate may be correlated with current and past values of the funds rate, but the strength and duration of that correlation is a matter for empirical analysis.[5] Second, the correlation between the funds rate and the bond rate induced by pure cyclical and/or aggressive policy actions is likely to be short run, appearing over business cycle periods. In contrast, the correlation induced by the Fed's reaction to shifts in the trend rate of inflation may be long run. A rise in the trend rate of inflation that permanently raises the bond rate will

[5] This lag arises not because financial markets adjust slowly but rather because funds rate strategy puts considerable persistence in the funds rate. Such a lag can also arise if the bond rate depends upon anticipated policy moves which in turn are influenced partly by current and past values of the policy variable.

also result in a higher funds rate if the Fed is trying not to induce any cyclical or aggressive element into its policy. Third, other economic factors such as inflation, the deficit, or the state of the economy may also influence the bond rate. Such correlations are apart from the one induced by monetary policy actions.

Given the above-noted considerations, the empirical work here examines the relationship between the bond rate and the funds rate in a multivariable framework. The other economic determinants included in the analysis are the actual inflation rate and the output gap that measures the cyclical state of the economy.[6] The work identifies the short- and long-run sources of correlation between the funds rate and the bond rate using cointegration and error-correction methodology. In particular, I proceed under the assumption, whose validity I do examine, that levels of the empirical measures of the long rate, the inflation rate, the funds rate, and other economic determinants each have unit roots (stochastic trends) and that there may exist cointegrating relationships among these variables. I interpret cointegrating regressions as measuring the long-run equilibrium correlations and error-correction regressions as measuring short-run correlations among variables.

To illustrate, assume that we are examining the interaction between the bond rate and the funds rate in a system that also includes inflation. Assume further that tests for cointegration indicate the presence of the following two cointegrating regressions in the system:

$$BR_t = a_0 + a_1 \dot{p}_t + U_{1t}, a_1 = 1, \text{ and} \tag{2}$$

$$NFR_t = b_0 + b_1 \dot{p}_t + U_{2t}, b_1 = 1, \tag{3}$$

where BR_t is the bond rate, $\dot{p}_t$ is actual inflation, NFR is the nominal funds rate, and U_{1t} and U_{2t} are two stationary disturbances. Equation (2) indicates that the bond rate moves one-for-one with inflation and that the long real rate is mean stationary. Equation (3) indicates that the funds rate moves one-for-one with inflation and that the real funds rate is mean stationary. These two cointegrating regressions are consistent with the presence of long-run equilibrium correlations between variables. If in the cointegrating regression, say, equation (2),

[6] This framework differs somewhat from the ones used in Goodfriend (1993) and Mehra (1994). Goodfriend describes interactions between the bond rate and the funds rate, taking into account the behavior of actual inflation and real growth, whereas in Mehra (1994) the deficit also is included. That work, however, indicates that the deficit variable is not a significant determinant of the bond rate once we control for the influences of inflation and real growth. Hence the deficit variable is excluded from the work here. I use the output gap rather than real growth as a measure of the state of the economy because the bond rate appears more strongly correlated with the former than with the latter. The qualitative nature of results, however, is the same whether the output gap or real growth is used as a measure of the state of the economy. Moreover, I do examine the sensitivity of results to some changes in specification in the subsection entitled "Additional Empirical Results."

$\dot{p}_t$ is weakly exogenous, then the long-run correlation can be given a causal interpretation, implying that the bond rate is determined by the (trend) rate of inflation.[7] The hypothesis that inflation is weakly exogenous in (2) can be tested by examining whether in regressions (4) and (5)

$$\Delta BR_t = a_2 + \delta_1(BR - \dot{p})_{t-1} + \sum_{s=1}^{n1} a_{3s}\Delta BR_{t-s} + \sum_{s=1}^{n2} a_{4s}\Delta\dot{p}_{t-s} \tag{4}$$

$$\Delta\dot{p}_t = b_2 + \delta_2(BR - \dot{p})_{t-1} + \sum_{s=1}^{n1} b_{3s}\Delta BR_{t-s} + \sum_{s=1}^{n2} b_{4s}\Delta\dot{p}_{t-s}, \tag{5}$$

where $\delta_1 \neq 0$ but $\delta_2 = 0$.[8] That result indicates that it is the bond rate, not inflation, that adjusts in response to deviations in the long-run relationship.

The cointegrating regressions discussed above identify the long-run comovements among variables. In order to estimate the short-run responses of the bond rate to the funds rate, the empirical work uses the following error-correction model of the bond rate:

$$\begin{aligned}\Delta BR_t = d_0 + \lambda_1 U_{1t-1} + \lambda_2 U_{2t-1} + \sum_{s=1}^{n} d_{1s}\Delta BR_{t-s} \\ + \sum_{s=0}^{n} d_{2s}\Delta NFR_{t-s} + \sum_{s=0}^{n} d_3\Delta\dot{P}_{t-s} + \epsilon_t,\end{aligned} \tag{6}$$

where all variables are as defined before and where Δ is the first difference operator. U_{1t-1} and U_{2t-1} are one-period lagged values of the residuals from the cointegrating regressions. If we substitute for U_{1t} and U_{2t} from (2) and (3) into (6), we can rewrite (6) as in (7):

$$\begin{aligned}\Delta BR_t = \tilde{d} + \lambda_1(BR_{t-1} - a_1\dot{p}_{t-1}) + \lambda_2(NFR_{t-1} - b_1\dot{p}_{t-1}) \\ + \sum_{s=1}^{n} d_{1s}\Delta BR_{t-s} + \sum_{s=0}^{n} d_{2s}\Delta NFR_{t-s} + \sum_{s=0}^{n} d_{3s}\Delta\dot{p}_{t-s} + \epsilon_t,\end{aligned} \tag{7}$$

where $\tilde{d} = d_0 - \lambda_1 a_0 - \lambda_2 b_0$. The short-run regression (7) includes levels as well as differences of variables. The empirical effects of changes in the inflation rate and the funds rate on the bond rate may occur through two distinct channels. First, those changes may affect the bond rate directly by altering future expectations of the inflation rate and the real rate of interest. The parameters $d_{is},\ i = 2,\ 3,\ s = 0,\ n$, measure near-term responses of the bond rate to changes

[7] The concept of weak exogeneity is introduced by Engle et al. (1983). The hypothesis that inflation is weakly exogenous with respect to the parameters of the cointegrating vector simply means that inferences on such parameters can be efficiently carried out without specifying the marginal distribution of inflation. More intuitively, inflation in equation (2) could be considered predetermined in analyzing the response of the bond rate to inflation.

[8] This test is proposed in Johansen (1992).

in its economic determinants. But, as noted before, signs and magnitudes of those parameters are a matter for empirical analysis because they depend upon factors such as the strategy of policy actions, the credibility of the Fed, and the nature of persistence in data. Lagged values of changes in the bond rate are included in order to capture better its own short-run dynamics.

The second focuses on disequilibrium in the long-run relations which may be caused by changes in the inflation rate and the funds rate. For example, aggressive funds rate changes taken to affect real growth or inflation may result in the level of the funds rate that is out of line with its value determined by the long-run equilibrium relation ($NFR_t \lesseqgtr b_0 + \dot{p}_t$ in [3]). Such short-run disequilibrium can also occur if the Fed adjusts the funds rate with lags in response to rising or falling inflation. Similarly, even though the bond rate moves automatically with inflation, short-run influences from other economic factors may result in the level of the bond rate that is out of line with its long-run equilibrium value ($R_t \lesseqgtr a_0 + \dot{p}_t$ in [2]). Such transitory perturbations in long-run equilibrium relations may have consequences for short-run changes in the bond rate. The parameters λ_1 and λ_2 in (7) thus measure the responses of the bond rate to such disequilibrium. The expected sign for λ_1 is negative, because the presence of the error-correction mechanism implies that the bond rate should decline (increase) if it is above (below) its long-run equilibrium value. In contrast, the sign of λ_2 is expected to be positive. But note all these disequilibrium effects are short-run (cyclical) in nature because in the long run (defined here in the equilibrium sense) they disappear and the bond rate is at its long-run equilibrium value determined by (2), i.e., $a_0 + \dot{p}_t$.

Data and Estimation Procedures

The empirical work in this article focuses on the behavior of the long rate during the sample period from 1957Q1 to 1995Q2. The long rate is measured by the nominal yield on ten-year U.S. Treasury bonds (*BR*). In most previous studies a distributed lag on the actual inflation rate is used as proxy for the long-run anticipated inflation, and actual inflation is generally measured by the behavior of the consumer price index. I also use actual inflation as proxy for anticipated inflation. I, however, measure inflation as the average of change in the consumer price index, excluding food and energy, over the past three years ($\dot{p}$).[9] The output gap (gaph) is the natural log of real GDP minus the

[9] I get similar results if instead the consumer price index or the GDP deflator is used to measure actual inflation (see the subsection entitled "Additional Empirical Results").

In a couple of recent studies (Hoelscher 1986; Mehra 1994) the Livingston survey data on one-year-ahead inflationary expectations are used to measure long-run anticipated inflation. The results in Mehra (1994), however, indicate that the near-term impact of the funds rate on the bond rate remains significant if one-year-ahead expected inflation (Livingston) data are substituted for actual inflation in the empirical work (see Mehra 1994, Table 4). That result continues to hold in this article also (see the subsection entitled "Additional Empirical Results").

log of potential GDP, which is generated using the Hodrick-Prescott filter. The interest rates are monthly averages for the last month of the quarter.

The stationarity properties of the data are examined using tests for unit root and mean stationarity. The unit root test used is the augmented Dickey-Fuller test and the test for mean stationarity is the one in Kwiatkowski et al. (1992). The test used for cointegration is the one proposed in Johansen and Juselius (1990).[10]

2. EMPIRICAL FINDINGS

In this section I describe cointegration test results for a system that includes the bond rate (*BR*), the inflation rate ($\dot{p}$), and the nominal funds rate (*NFR*). I also discuss short-run results from error-correction regressions for the full sample period as well as for several subperiods. The section concludes with an explanation of different pre- and post-1979 sample results.

Cointegration Test Results

Test results for unit roots and mean stationarity are summarized in the appendix. They indicate that the bond rate (*BR*), the inflation rate ($\dot{p}$), and the nominal funds rate (*NFR*) each have a unit root and thus contain stochastic trends. The output gap variable by construction is stationary.

Test results for cointegration are also summarized in the appendix. I first focus on the bivariable systems $(BR, \dot{p})$, $(NFR, \dot{p})$, and (BR, NFR). Test results are consistent with the presence of cointegration between variables in each system, indicating that the bond rate is cointegrated with the inflation rate and the funds rate. The funds rate is also cointegrated with the inflation rate. Thus the bond rate comoves with each of these nonstationary variables, including the funds rate.

The presence of cointegration between two variables simply means that there exists a long-run stochastic correlation between them. In order to help determine whether such correlation can be given a causal interpretation, Table 1 presents test results for weak exogeneity of the long-run parameters. In the system $(BR, \dot{p})$ inflation is weakly exogenous but the bond rate is not, indicating that it is the bond rate that adjusts in response to deviations in the long-run relationship. Thus the long-run equilibrium relationship between the bond rate and the inflation rate can be interpreted as a Fisher relation in which the bond rate is determined by the (trend) rate of inflation. In the system $(NFR, \dot{p})$ inflation is again weakly exogenous but the funds rate is not. Here again the long-run relation can be interpreted as one in which the inflation rate drives the interest rate: in this case the short-term rate. Hence, I interpret the long-run

[10] These tests are described in Mehra (1994).

Table 1 Cointegrating Regressions and Test Results for Weak Exogeneity

Equation Number	Panel A Cointegrating Regressions	Panel B Error-Correction Coefficients (t-value) in Equations for		
		$\triangle BR$	$\triangle \dot{p}$	$\triangle NFR$
1	$BR_t = 3.3 + 0.93\ \dot{p}_t + U_{1t}$ (10.0)	−0.18 (3.4)	−0.01 (0.6)	
2	$BR_t = 1.2 + 0.91\ NFR_t + U_{2t}$ (50.7)	−0.17 (2.5)		0.21 (1.5)
3	$NFR_t = 1.8 + 1.1\ \dot{p}_t + U_{3t}$ (6.2)		0.00 (0.4)	−0.27 (3.2)
4	$(BR - \dot{p})_t = 2.2 + 0.09\ NFR_t$ (1.3) $\chi^2_1 = 1.7$			

Notes: Cointegrating regressions given in panel A above are estimated by the dynamic OLS procedure given in Stock and Watson (1993), using leads and lags of first differences of the relevant right-hand side explanatory variables. Eight leads and lags are included. Parentheses that appear below coefficients in cointegrating regressions contain t-values corrected for the presence of moving average serial correlation. The order of serial correlation was determined by examining the autocorrelation function of the residuals. χ^2_1 is the Chi-square statistic that tests the hypothesis that the coefficient that appears on *NFR* in equation 4 is zero.

Panel B above contains error-correction coefficients from regressions of the form

$$\Delta X_{1t} = \delta_1\ U_{t-1} + \sum_{s=1}^{4}\ a_s \Delta X_{1t-s} + \sum_{s=1}^{4}\ b_s\ \Delta X_{2t-s}$$

$$\Delta X_{2t} = \delta_2\ U_{t-1} + \sum_{s=1}^{4}\ c_s \Delta X_{1t-s} + \sum_{s=1}^{4}\ d_s\ \Delta X_{2t-s},$$

where U_{t-1} is the lagged value of the residual from the cointegrating regression that is of the form

$$X_{1t} = d_0 + d_1 X_{2t} + U_t,$$

and where X_{1t} and X_{2t} are the pertinent nonstationary variables. The relevant cointegrating regressions are given in panel A above. Parentheses that appear below error-correction coefficients contain t-values.

equilibrium relationship between the funds rate and the inflation rate as a kind of reaction function.

The test results for weak exogeneity discussed above for systems $(BR, \dot{p})$ and $(NFR, \dot{p})$ also imply that inflation causes the comovement between the bond rate and funds rate. The bond rate comoves with the funds rate because the bond rate moves automatically with inflation and the Fed keeps the funds rate in line with the trend rate of inflation in the long run.

The analysis above, based on bivariable systems, thus suggests that in the Fisher relation the funds rate should not be correlated with the bond rate once we control for the correlation that is due to inflation. I test this implication by examining whether the ex post real rate ($BR - \dot{p}$) is correlated with the funds rate. I do so by expanding the Fisher relation to include the funds rate while maintaining the Fisher restriction that the bond rate adjusts one-for-one with inflation. In that regression the funds rate is not significant (see Table 1, equation 4). The Chi-square statistic for the null hypothesis that the ex post real rate is not correlated with the funds rate is small, consistent with no correlation.

The result that the bond rate is cointegrated with the actual inflation rate and thus the ex post real rate ($R_t - \dot{p}_t$) is stationary also implies that the expected long real rate is stationary. This can be seen if we express the Fisher relation (1) as

$$BR_t - \dot{p}_t = rr_t^e + (\dot{p}_t^e - \dot{p}_t) = rr_t^e + U_t,$$

where all variables are defined as before and where U_t is the disturbance term. This disturbance arises because the long-term expected inflation rate may differ from the three-year inflation rate. As is clear, the stationarity of ($BR_t - \dot{p}_t$) implies the stationarity of the expected long real rate.

Error-Correction Regressions

Since the ex ante long real rate is mean stationary, not constant, cyclical and aggressive funds rate changes discussed before may still affect the bond rate by altering expectations of its real rate and inflation components. I now explore those short-run effects by estimating the error-correction equation.

The empirical results discussed in the previous section are consistent with the following two cointegrating relationships:

$$BR_t = a_0 + a_1\dot{p}_t + U_{1t} \quad \text{and} \tag{8}$$

$$NFR_t = b_0 + b_1\dot{p}_t + U_{2t}. \tag{9}$$

Equation (8) is the Fisher relation, and equation (9) the Fed reaction function. The latter captures that component of the funds rate that comoves with trend inflation. The residual U_{2t} is then the component that captures the stance of cyclical and aggressive funds rate policy actions. Consider then the following error-correction equation:

$$\Delta BR_t = d_0 + \lambda_1 U_{1t-1} + \lambda_2 U_{2t-1} + \sum_{s=1}^{n} d_{1s}\Delta BR_{t-s} + \sum_{s=1}^{n} d_{2s}\Delta\dot{p}_{t-s} + \sum_{s=1}^{n} d_{3s} NFR_{t-s} + \sum_{s=1}^{n} d_{4s} gaph_{t-s} + \epsilon_t. \tag{10}$$

The parameter λ_2 in (10) measures the response of the bond rate to the lagged value of the funds rate spread $(NFR - b_0 - b_1\dot{p})$. Since the short-run equation is in first differences, the near-term response of the bond rate to the funds rate spread can be calculated as $-\lambda_2/\lambda_1$.[11] Also, equation (10) includes only lagged values of economic determinants and hence can be estimated by ordinary least squares. In order to examine subsample variability, I estimate equations (8) through (10) over several sample periods, all of which begin in 1961Q2 but end in different years from 1972 through 1995.

Table 2 presents some key coefficients $(\lambda_1, \lambda_2, -\lambda_2/\lambda_1, a_1, b_1)$ from these regressions. If we focus on full sample results, then it can be seen that all these key coefficients appear with expected signs and are statistically significant. In cointegrating regressions the bond rate adjusts one-for-one with inflation and so does the funds rate. The hypotheses that $a_1 = 1$ and $b_1 = 1$ cannot be rejected and thus are consistent with our priors about the interpretation of (8) as the Fisher relation and of (9) as the Fed reaction function. In the error-correction regression λ_2 is positive and its estimated value indicates a one percentage point rise in the funds rate spread raises the bond rate by 13 basis points in the following period. The net increase totals 42 basis points. The mean lag $(-1/\lambda_1)$ in the short-run effect of the funds rate spread on the bond rate is approximately 3.2 quarters, indicating that these near-term responses dissipate quite rapidly.[12]

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[11] This can be shown as follows. Assume, for example, the level of the bond rate is related to inflation and the funds rate spread as in

$$BR_t = a_0 + a_1\dot{p}_t + a_2(NFR - b_1\dot{p})_t + V_t, \tag{a}$$

where the stationary component is $a_0 + a_2(NFR - b_1\dot{p})_t$ and the nonstationary component is $a_1\dot{p}_t$. The error-correction regression is

$$\Delta BR_t = \lambda_1 V_{t-1} + \sum_{s=1}^{n} \text{(other lagged differences of variables)} + \epsilon_t, \tag{b}$$

where λ_1 is negative. Substituting for V_{t-1} from (a) into (b) yields (c):

$$\Delta BR_t = \lambda_1 BR_{t-1} - \lambda_1 a_1 \dot{p}_{t-1} - \lambda_1 a_2 (NFR - b_1\dot{p})_{t-1} + \textit{ other terms}. \tag{c}$$

Equation (c) can be estimated and a_2 can be recovered as $\lambda_1 a_2/\lambda_1$, which is the minus of the coefficient on $(NFR - b_1\dot{p})_{t-1}$ divided by the coefficient on BR_{t-1}. The coefficient a_2 then measures the near-term response of the bond rate to the funds rate spread. I do not label a_2 as measuring the long-run effect because the spread is stationary. The long run is defined as the period over which trend relationships emerge. In the long run the funds rate spread $(NFR - \dot{p})$ is constant.

[12] In estimated short-run regressions the coefficients that appear on lagged differences of the bond rate are very small. If we ignore those coefficients, then the short-run equation (c) given in footnote 11 can be expressed as

$$BR_t = \frac{-\lambda_1 a_1}{1 - (1 + \lambda_1)L}\,\dot{p}_{t-1} + \frac{\lambda_2}{1 - (1 + \lambda_1)L}\,(NFR - \dot{p})_{t-1} \;+\; \textit{other terms},$$

where L is the lag operator and where λ_2 is $-\lambda_1 a_2$. The coefficients (w_i) that appear on lagged levels of $NFR - \dot{p}$ are then of the form $\lambda_2, \lambda_2(1 + \lambda_1), \lambda_2(1 + \lambda_1)^2$, etc. The mean lag then can be calculated as follows:

If we focus on subsample results, it can be seen that all key coefficients still appear with expected signs and are statistically different from zero. However, there are some major differences between pre- and post-1979 regression estimates. In pre-1979 cointegrating regressions the hypotheses that $a_1 = 1$ and $b_1 = 1$ are generally rejected. In contrast, that is not the case in most post-1979 regressions. The rejection, however, is more common in the Fisher relation than it is in the Fed reaction function.

In pre-1979 error-correction regressions the bond rate does respond to the funds rate spread—the one-period response (λ_2) ranges from 11 to 23 basis points. But the one-period response is generally quite close to the near-term net response ($-\lambda_2/\lambda_1$), indicating that the effect of policy actions on the bond rate did not persist too long. In post-1979 error-correction regressions, however, the near-term response of the bond rate to the funds rate spread is larger than the one-period lagged response. In particular, the immediate response of the bond rate to the funds rate ranges from 13 to 16 basis points and the near-term response from 36 to 48 basis points. Together these estimates imply that the near-term response of the bond rate to the funds rate spread has increased since 1979.[13] I argue below that these different results may be due in part to the way the Fed has conducted its monetary policy since 1979. In particular, I focus on the Fed's disinflation policy.

Before 1979 the Fed did not aggressively attempt to bring down the trend rate of inflation. In the long-run Fed reaction function estimated over 1961Q2 to 1979Q3, the parameter b_1 is less than unity, indicating that the Fed did not adjust the funds rate one-for-one with inflation (see Table 2). Moreover, the short-run reaction functions estimated in Mehra (1996) also indicate that in the pre-1979 period the Fed did not respond to accelerations in actual inflation. Hence, during this early period a monetary policy tightening measured by a widening in the funds rate spread may have affected the bond rate primarily by altering its expected real rate component. Because the funds rate increase alters only near-term expectations of future short real rates, its impact on the bond rate is likely to be modest, as confirmed by low estimates of λ_2 in Table 2. When the bond rate rises above the current inflation rate, both the bond rate and actual inflation rises, speeding up the adjustment, as confirmed by high estimates of λ_1 in Table 2. As a result, the immediate effect of policy on

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$$\textit{Mean Lag} = \sum_{i=1}^{\infty} w_i\, i \Big/ \sum_{i=1}^{\infty} w_i = \frac{\lambda_2[1 + 2(1+\lambda_1) + 3(1+\lambda_1)^2 + 4(1+\lambda_1)^3 + \ \dots \infty]}{\lambda_2[1 + (1+\lambda_1)^2 + (1+\lambda_1)^3 + \ \dots \infty]}$$

$$= \lambda_2 \frac{1}{(1-1-\lambda_1)^2} \times \frac{1}{\lambda_2 \frac{1}{1-1-\lambda_1}} = -\frac{\lambda_1}{\lambda_1^2} = -\frac{1}{\lambda_1}.$$

[13] This is confirmed by results of the formal Chow test that is discussed in the next subsection.

Table 2 Short-Run Error-Correction Regressions Using Residuals

Estimation Ends in Year	Panel A: Error-Correction Coefficients				Panel B: Coefficients from Cointegrating Regressions	
	$\lambda_2(\tilde{t})$	$\lambda_1(\tilde{t})$	$-\lambda_2/\lambda_1$	Mean Lag in Quarters	$a_1(\tilde{t})$	$b_1(\tilde{t})$
1995Q2	0.13(2.7)	−0.31(4.8)	0.42	3.2	0.93(10.0)	1.10 (6.0)
1994	0.12(2.4)	−0.29(4.5)	0.41	3.4	0.93(10.1)	1.10 (6.2)
1992	0.14(2.6)	−0.30(4.5)	0.46	3.3	0.92 (9.7)	1.10 (6.5)
1990	0.15(2.6)	−0.31(4.3)	0.47	3.3	0.92 (9.8)	1.10 (6.7)
1988	0.14(2.4)	−0.30(4.1)	0.48	3.3	0.94(10.9)	1.00 (6.6)
1986	0.16(2.7)	−0.34(2.7)	0.48	2.9	0.94(10.8)	1.10 (6.9)
1984	0.13(1.9)	−0.30(2.6)	0.46	3.3	0.93(11.1)	1.10 (6.7)
1982	0.19(2.5)	−0.54(3.6)	0.36	1.8	0.76(17.1)^c	0.84 (7.6)
1980	0.18(2.1)	−0.71(2.2)	0.26	1.4	0.68(30.8)^c	0.60 (5.6)^c
1979Q3	0.17(3.1)	−0.67(3.3)	0.25	1.5	0.69(31.1)^c	0.77 (8.9)^c
1978	0.16(2.8)	−0.64(3.1)	0.26	1.5	0.71(28.4)^c	0.78 (9.9)^c
1976	0.17(3.2)	−0.65(3.0)	0.26	1.5	0.71(21.4)^c	1.00(15.9)
1974	0.23(2.9)	−0.98(3.6)	0.23	1.0	0.78(17.9)^c	0.97(14.6)
1972	0.11(1.2)	−0.75(2.6)	0.14	1.3	0.73 (6.9)^c	0.68 (2.5)

^{c} indicates the relevant coefficient (a_1 or b_1) is significantly different from unity.

Notes: The coefficients reported above are from the following regressions:

$$BR_t = a_0 + a_1\dot{p}_t + U_{1t}, \quad \text{(a)}$$

$$NFR_t = b_0 + b_1\dot{p}_t + U_{2t}, \text{ and} \quad \text{(b)}$$

$$\Delta BR_t = d_0 + \lambda_1 U_{1t-1} + \lambda_2 U_{2t-1} + \sum_{s=1}^{n} d_{1s}\Delta BR_{t-s} + \sum_{s=1}^{2} d_{2s}\Delta\dot{p}_{t-s} + \sum_{s=1}^{2} d_{3s}\Delta NFR_{t-s} + \sum_{s=1}^{2} d_{4s}\, gaph_{t-s} + \epsilon_t, \quad \text{(c)}$$

where gaph is the output gap and other variables are defined as in Table A1. Equations (a) and (b) are estimated by the dynamic OLS procedure given in Stock and Watson (1993) and equation (c) by ordinary least squares. The estimation period begins in 1961Q2 and ends as shown in the first column. $\tilde{t}$ is the t-statistic. The mean lag is calculated as $-1/\lambda_1$.

the bond rate dissipates quickly and hence the near-term effect is close to the immediate impact.

In the post-1979 period, however, the Fed made a serious attempt to bring down the trend rate of inflation and to contain inflationary expectations. The descriptive analysis of monetary policy in Goodfriend (1993) and the short-run reaction function estimated in Mehra (1996) are consistent with this observation. Hence short-run increases in the funds rate spread may also have affected the bond rate by altering the long-term expected inflation. If the Fed's disinflation policy had been credible, then increases in the funds rate spread that raise the bond rate's real component may also lower expectations of the long-term expected inflation rate, thereby offsetting somewhat the immediate or the very-near-term response of the bond rate to the funds rate spread. The evidence reported in Table 2, however, indicates that the estimated coefficient $(-\lambda_2/\lambda_1)$ that measures the near-term response shows no tendency to fall in the post-1979 period. On the other hand, if the public does not have full confidence in the Fed's disinflation policy and if the Fed has to persist with sufficiently high short real rates to reduce the trend rate of inflation or contain inflationary expectations, then the estimated effect of a policy action on the bond rate would last longer. In this case, the near-term effect of the funds rate spread on the bond rate would be larger and the mean lag in the effect of such policy on the bond rate would also be higher. Both these implications are consistent with results in Table 2, where both the estimated short-run effect $(-\lambda_2/\lambda_1)$ and the mean lag $(-1/\lambda_1)$ are higher in the post-1979 period than they were in the period before.

Additional Empirical Results

In this section I discuss and report some additional test results which confirm the robustness of conclusions reached in the previous section. I consider several changes in the specification of the short-run equation (10). First, I reproduce the empirical work in Table 2 under the alternative specification that the short-run stance of monetary policy is measured by the real funds rate. I then consider additional changes in specification to address concerns raised by the potential endogeneity of monetary policy actions, alternative measures of inflation, and the potential stationarity of data. For these latter experiments I focus on results that pertain to the short-run effect of the funds rate on the bond rate over two sample periods only, 1961Q1 to 1995Q2 and 1961Q2 to 1979Q3. Hence I report two key coefficients, λ_2 and $(-\lambda_2/\lambda_1)$.

The results in Table 2 discussed in the previous section use the residual from the long-run Fed reaction function as a measure of the short-run stance of monetary policy. I now examine results if the short-run stance of policy is measured instead by the real funds rate $(NFR - \dot{p})$. Furthermore, I now estimate the Fisher relation (8) jointly with the short-run error-correction equation. This procedure allows for richer short-run dynamics in estimating the long-run effect of inflation on the bond rate. The short-run equation that incorporates these two new changes can be derived by replacing the residuals U_{1t-1} and U_{2t-1} in (10) by lagged levels of the variables and then by setting $b_1 = 1$. The resulting short-run equation is

$$\Delta BR_t = \tilde{d} + \lambda_1 BR_{t-1} + \lambda_3 \dot{P}_{t-1} + \lambda_2 (NFR - \dot{p})_{t-1} + \sum_{s=1}^{n} d_{1s} \Delta BR_{t-s} + \sum_{s=1}^{n} d_{2s} \Delta \dot{p}_{t-s} + \sum_{s=1}^{n} d_{3s} \Delta (NFR - \dot{p})_{t-s} + \sum_{s=1}^{n} d_{4s}\ gaph_{t-s} + \epsilon_t, \quad (11)$$

where $\lambda_3 = -\lambda_1 a_1$. In (11), if $a_1 = 1$, then the coefficients on BR_{t-1} and $\dot{p}_{t-1}$ sum to zero ($\lambda_1 + \lambda_3 = 0$ in (11)). I impose this restriction only if it is not rejected. Table 3 reports some key coefficients ($\lambda_1, \lambda_2, -\lambda_2/\lambda_1, \lambda_3$). These estimated coefficients confirm the qualitative nature of results in Table 2. First, the real funds rate is a significant predictor of the bond rate and this result

Table 3 Short-Run Error-Correction Regressions Using the Level of the Funds Rate Spread

Estimation Ends in Year	$\lambda_2(\tilde{t})$	$\lambda_1(\tilde{t})$	$-\lambda_2/\lambda_1$	Mean Lag in Quarters	$\lambda_3(\tilde{t})$	$-\lambda_3/\lambda_1$	F1
1995Q2	0.12(2.6)	−0.29(4.7)	0.40	3.5	0.29(4.7)	1.00	0.89
1994	0.10(2.2)	−0.26(4.3)	0.35	3.8	0.26(4.3)	1.00	0.89
1992	0.11(2.4)	−0.28(4.4)	0.39	3.6	0.28(4.4)	1.00	0.64
1990	0.11(2.2)	−0.28(4.1)	0.39	3.6	0.28(4.1)	1.00	0.68
1988	0.11(2.2)	−0.27(3.9)	0.41	3.7	0.27(3.9)	1.00	0.63
1986	0.13(2.5)	−0.30(4.2)	0.41	3.2	0.30(4.2)	1.00	0.55
1984	0.10(1.7)	−0.22(2.3)	0.41	4.5	0.22(2.3)	1.00	0.11
1982	0.21(3.4)	−0.53(3.5)	0.40	1.9	0.48(3.5)	0.90	3.70*
1980	0.12(1.9)	−0.24(2.0)	0.50	4.2	0.24(2.0)	1.00	1.00
1979Q3	0.13(2.5)	−0.41(2.4)	0.31	2.4	0.32(2.5)	0.80	3.60*
1978	0.12(2.4)	−0.41(2.3)	0.29	2.4	0.32(2.3)	0.78	3.40*
1976	0.14(2.6)	−0.53(2.6)	0.26	1.9	0.39(2.5)	0.74	6.60**
1974	0.22(3.1)	−0.92(3.8)	0.23	1.1	0.67(3.7)	0.73	14.20**
1972	0.21(2.3)	−0.97(3.4)	0.22	1.0	0.69(3.3)	0.71	10.10**

*Significant at the 10 percent level.

**Significant at the 5 percent level.

Notes: The coefficients reported are from regressions of the form

$$\Delta BR_t = d_0 + \lambda_1\ BR_{t-1} + \lambda_3\ \dot{P}_{t-1} + \lambda_2\ (NFR - \dot{p})_{t-1} + \sum_{s=1}^{2}\ d_{1s}\ \Delta BR_{t-s} + \sum_{s=1}^{2}\ d_{2s}\ \Delta \dot{p}_{t-s} + \sum_{s=1}^{2}\ d_{3s}\ \Delta (NFR - \dot{p})_{t-s} + \sum_{s=1}^{2}\ d_{4s}\ gaph_{t-s},$$

where gaph is the output gap and other variables are as defined in Table A1. All regressions are estimated by ordinary least squares. The estimation period begins in 1961Q2 and ends in the year shown. The mean lag is calculated as $-1/\lambda_1$. $\tilde{t}$ is the t-statistic. F1 tests the null hypotheses that λ_1 and λ_3 sum to zero, indicating that the Fisher restriction is consistent with data.

is fairly robust over several subsamples. Second, the Chow test indicates that two key parameters, λ_1 and λ_2, are unstable only between pre- and post-1979 periods (the date of the break is 1980Q2). This result is consistent with an increase in the near-term effect of policy on the bond rate in the post-1979 period.

The short-run error-correction equation (11) was alternatively re-estimated using the consumer price index and the GDP deflator to measure inflation, that is, the average inflation rate over the past three years. As in a few previous studies, I also used the Livingston survey data on one-year-ahead inflationary expectations. The results continue to indicate that the funds rate spread generally does help predict the bond rate and that the near-term response of the bond rate to the funds rate spread has increased since 1979 (see rows 1 through 3 in Table 4).[14]

The funds rate spread here measures the short-run stance of monetary policy because it is that component of the funds rate that does not comove with inflation. This spread, however, is still endogenous because, as noted before, the Fed routinely raises the funds rate during cyclical expansions and lowers it during cyclical downturns. The potential problem created by such endogeneity is that if the bond rate is directly influenced by variables that reflect the cyclical state of the economy and if those variables are omitted from short-run regressions, then the funds rate spread may be picking up the influence of those variables on the bond rate rather than the influence of monetary policy on the real component of the bond rate.

The short-run regressions reported in Tables 2 and 3 already include many of those variables such as the output gap that measures the cyclical state of the economy and changes in inflation, the bond rate, and the funds rate spread itself. While it is difficult to know all the information that the Fed may be using in setting its short-run funds rate policy, I re-estimated (11) alternatively including additional variables such as nonfarm payroll employment, sensitive materials prices, the deficit, and real growth. Those additional variables, when included in (11), are generally not significant and therefore do not change the qualitative nature of results in Table 3 (see rows 4a through 4d in Table 4).

It is sometimes argued that unit root tests used here have low power in distinguishing whether the variables are stationary or integrated. Hence the cointegration and error-correction methodology used here to distinguish between long- and short-run sources of comovement between the bond rate and the funds rate is suspect. However, the evidence presented above that the bond rate and the funds rate adjust one-for-one with inflation in the long run is confirmed even if I treat the bond rate, the funds rate, and inflation as stationary

[14] When inflation is measured by the behavior of the consumer price index or the Livingston survey, I get some mixed results. The statistical significance of the coefficient that appears on the funds rate spread is not robust over different sample periods.

Table 4 Sensitivity to Changes in Specification

Changes in Specification	Panel A: 1961Q2–1995Q2				Panel B: 1961Q2–1979Q3			
	$\lambda_2(\tilde{t})$	$(-\lambda_2/\lambda_1)$	f_2	$\tilde{f}_2$	$\lambda_2(\tilde{t})$	$(-\lambda_2/\lambda_1)$	f_2	$\tilde{f}_2$
1. CPI	0.08(1.7)	0.42			0.09(1.6)	0.16		
2. GDP Deflator	0.11(2.2)	0.49			0.10(2.0)	0.19		
3. Livingston Survey	0.08(1.7)	0.35			0.07(1.6)	0.23		
4. CPIEXFE								
Additional Variables								
a. $\Delta \ln PEM$	0.14(2.9)	0.48			0.15(2.8)	0.35		
b. d_t	0.10(2.3)	0.37			0.13(2.0)	0.28		
c. $\Delta \ln SMP$	0.10(2.4)	0.38			0.13(2.5)	0.29		
d. $\Delta \ln ry_t$	0.12(2.5)	0.43			0.10(1.5)	0.25		
5. Stationary: Level Regressions			0.07(1.9)	0.42			0.09(2.1)	0.29

Notes: The coefficients reported are from regressions of the form given in Table 3. The Fisher restriction is imposed in regressions estimated over 1961Q2–1995Q2 but not in those estimated over 1961Q2 to 1979Q3. CPI is the consumer price index; CPIEXFE is the consumer price index excluding food and energy; PEM is the nonfarm payroll employment; $\Delta \ln ry$ is real GDP growth; d is federal government deficits scaled by nominal GDP, and SMP is the sensitive materials prices. The coefficients reported in the row labeled 5 are from the regression (14) of the text. f_2 is the coefficient that measures the contemporary response of the bond rate to the funds rate spread and $\tilde{f}_2$ the near-term.

variables. The stationary versions of the long-run regressions (8) and (9) can be expressed as in (12) and (13):

$$BR_t = a_0 + \sum_{s=0}^{n} a_{1s}\ \dot{p}_{t-s} + \sum_{s=1}^{n} a_{2s}\ BR_{t-s} + U_{1t} \tag{12}$$

$$NFR_t = b_0 + \sum_{s=0}^{n} b_{1s}\ \dot{p}_{t-s} + \sum_{s=1}^{n} b_{2s}\ NFR_{t-s} + U_{2t}. \tag{13}$$

The net response of the bond rate to inflation is $\left(\sum_{s=0}^{n} a_{1s} \Big/ 1 - \sum_{s=1}^{n} a_{2s}\right)$ and to the funds rate is $\left(\sum_{s=0}^{n} b_{1s} \Big/ 1 - \sum_{s=1}^{n} b_{2s}\right)$. One cannot reject the hypotheses that these net responses each are unity. As for short-run correlations, consider the following stationary version of the short-run equation:

$$BR_t = f_0 + f_1\dot{p}_t + f_2\ (NFR - \dot{p})_t + f_3\ gaph_t + \sum_{s=1}^{n}\ f_{4s}\ BR_{t-s} + \epsilon_t, \tag{14}$$

where all variables are as defined before. Equation (14) already incorporates the long-run restriction that short-run funds rate policy actions affect the bond rate by altering the spread between the funds rate and the inflation rate. The other restriction can be imposed by the requirement that coefficients f_1 and $\sum_{s=1}^{n} f_{4s}$ in (14) sum to unity. The parameter f_2 in (14) measures the contemporaneous response of the bond rate to the funds rate spread and its net response can be calculated as $\left(f_2 \big/ 1 - \sum_{s=1}^{n} f_{4s}\right)$.

I estimate equation (14) by instrumental variables that are just lagged values of the right-hand side explanatory variables. In such regressions the funds rate spread still helps predict the bond rate (see row 5 in Table 4).

A Comparison with Some Previous Studies

In this section I discuss some previous studies that use entirely different methodologies but reach conclusions regarding the short-run impact of policy on long-term rates which are qualitatively similar to those reported here.

The first set consists of studies by Cook and Hahn (1989), Radecki and Reinhart (1994), and Roley and Sellon (1995). All three of these studies examine the response of long-term interest rates to changes in a measure of the federal funds rate target. They differ, however, with respect to the sample period studied and the length of the interval over which the interest rate response is measured. In Cook and Hahn the sample period studied is September 1974 to September 1979 and the interest rate response is measured on the day of the target change. In the other two studies the sample periods examined fall within the post-1979 period: 1989 to 1993 in Radecki and Reinhart (1994) and October 1987 to July 1995 in Roley and Sellon (1995). The measurement interval in Radecki and Reinhart spans the first ten days following the policy change, whereas in Roley and Sellon the time interval spans the period from the day after the previous policy action to the day after the current policy action. The economic rationale for the use of a wider measurement interval as in Roley and Sellon is that many times monetary policy actions are already anticipated by the markets, so that long-term interest rates move ahead of the announced change in the funds rate target. The relative magnitudes of interest rate responses before, during, and after the policy action, however, depend upon whether policy actions are anticipated or unanticipated and upon the degree of persistence in anticipated policy actions.

The measurement interval is the narrowest in Cook and Hahn and Radecki and Reinhart; the results there indicate that a one percentage point increase in the funds rate target induces 12 to 13 basis points movement in the ten-year bond rate (increases in longer maturity bond rates are somewhat smaller). The size of the interest rate response during and after the change in policy action as measured by Roley and Sellon is also modest; the 30-year Treasury bond yield rises by 10 basis points following one percentage point increase in the effective federal funds rate target. However, when the measurement interval includes days before the change in policy action, the measured interest rate response rises to 38 basis points. Thus a significant part of the response occurs before policy action is announced, indicating the presence of anticipated effect. What needs to be noted is that the magnitude of the total interest rate response measured by Roley and Sellon is quite close to the near-term response that I have estimated using an entirely different estimation methodology. Recall that for the complete sample period 1961Q2 to 1995Q2 the estimated near-term response of the ten-year bond rate to the funds rate spread is 42 basis points (see Table 2). The estimated near-term response is 36 basis points if instead I use the 30-year bond yield in my empirical work.

The other recent study showing that in the short run the long real rate comoves with the short nominal rate is by Fuhrer and Moore (1995). According to the expectational theory of the term structure of interest rates, the ex ante long-term real rate can be viewed as a weighted moving average of future short real rates. Fuhrer and Moore define the short real rate as the nominal yield on three-month Treasury bills minus the actual quarterly inflation rate. They then use a vector autoregression to construct long-horizon forecasts of the time paths of the three-month Treasury bill rate and the inflation rate. Given those forecasts they compute the 40-quarter-duration long-term real rate, since the average duration (maturity) of Moody's BAA corporate bond rate is 40 quarters. What they find is that over the period 1965 to 1992 the sample path of the ex ante long real rate tracks closely that of the short-term nominal rate (Fuhrer and Moore 1995, Figure 1, p. 224). The ex ante long real rate is still relatively stable, however, and only about one-fourth of the increase in the short nominal rate is reflected in the long real rate.

3. CONCLUDING OBSERVATIONS

This article has investigated empirically the immediate, near-term, and long-run effects of monetary policy on the bond rate. The federal funds rate is used as a measure of monetary policy, and the long run is viewed as the period during which trend relationships emerge. The results indicate that the long-run effect of monetary policy on the bond rate occurs primarily through the inflation channel.

In the short run, however, monetary policy also affects the bond rate by altering its expected real rate component. The short-run stance of monetary policy is measured by the spread between the funds rate and the ongoing trend rate of inflation. The results indicate that the near-term effect of the funds rate spread on the bond rate has increased considerably since 1979. In the pre-1979 period, the bond rate rose anywhere from 14 to 29 basis points whenever the funds rate spread widened by one percentage point. In the post-1979 period, however, the estimate of its near-term response ranges from 26 to 50 basis points.

This increase in the short-run sensitivity of the bond rate to monetary policy actions is consistent with the way the Fed has conducted its monetary policy since 1979. Since then the Fed has made a serious attempt to bring down the trend rate of inflation and contain inflationary expectations. If the public does not have full confidence in the Fed's disinflation policy, and if the Fed has to persist with sufficiently high short real rates to reduce the trend rate of inflation or contain inflationary expectations, then the estimated effect of a policy action on the bond rate would last longer. As a result, the near-term effect of policy on the bond rate would be stronger than would be the case if the disinflation policy were fully credible.

APPENDIX A

The stationarity properties of data are investigated using both unit roots and mean stationarity tests. The test for unit roots used is the augmented Dickey-Fuller test and the one for mean stationarity is the one in Kwiatkowski et al. (1992). Both these tests are described in Mehra (1994).

Table A1 presents test results for determining whether the variables BR, $\dot{p}$, NFR, $BR - \dot{p}$ and $NFR - \dot{p}$ have a unit root or are mean stationary. As can be

Table A1 Tests for Unit Roots and Mean Stationarity

Series	Panel A Tests for Unit Roots			Panel B Tests for Mean Stationarity
X	$\hat{\rho}$	$t_{\hat{\rho}}$	k	$\hat{n}_u$
BR	0.96	−1.7	5	0.83*
$\dot{p}$	0.99	−2.0	1	0.56*
NFR	0.89	−2.9[a]	5	0.46[a]
$BR - \dot{p}$	0.87	−3.2*	3	0.04
$NFR - \dot{p}$	0.70	−5.0*	5	0.03
ΔBR	−0.10	−5.6*	4	0.19
$\Delta\dot{p}$	0.60	−4.3*	7	0.18
ΔNFR	−0.30	−4.9*	7	0.07

[a]The test statistic is close to the relevant 5 percent critical value.

*Significant at the 5 percent level.

Notes: BR is the ten-year Treasury bond rate; $\dot{p}$ is the average inflation rate over the past three years; and NFR is the nominal funds rate. Δ is the first difference operator. Inflation is measured by the behavior of the consumer price index excluding food and energy. This price series begins in 1957; therefore the effective sample period studied is 1960Q1 to 1995Q2. The values for ρ and t-statistics ($t_{\hat{\rho}}$) for $\rho = 1$ in panel A above are from the augmented Dickey-Fuller regressions of the form

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

where X is the pertinent series. The number of lagged first differences (k) included in these regressions are chosen using the procedure given in Hall (1990). The procedure starts with some upper bound on k, say k max chosen a priori (eight quarters here). Estimate (a) above with k set at k max. If the last included lag is significant, select $k = k$ max. If not, reduce the order of the autoregression by one until the coefficient on the last included lag is significant. The test statistic $\hat{n}_u$ in panel B above 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 and for $t_{\hat{\rho}}$ given in Fuller (1976) is -2.89.

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, but large for $BR - \dot{p}$ and $NFR - \dot{p}$. On the other hand, the test statistics ($\hat{n}_u$) that tests the null hypothesis that a particular variable is mean stationary is large for BR, $\dot{p}$, and NFR, but small for $BR - \dot{p}$ and $NFR - \dot{p}$. These results indicate that BR, $\dot{p}$, and NFR have a unit root and are thus nonstationary in levels. In contrast $BR - \dot{p}$ and $NFR - \dot{p}$ do not have a unit root and are thus stationary in levels. Table A1 also presents unit roots and mean stationary tests using first differences of BR, $\dot{p}$, and NFR. As can be seen, the test results indicate that first differences of these variables are stationary.

The test for cointegration used is the one proposed in Johansen and Juselius (1990). The test procedure is described in Mehra (1994). Two test statistics—the trace test and the maximum eigenvalue test—are used to evaluate the number of cointegrating relationships. Table A2 presents these two test statistics for determining whether in bivariable systems like $(BR, \dot{p})$, (BR, NFR) and $(NFR, \dot{p})$ there exist a cointegrating vector. Those test results are consistent with the presence of cointegration between variables in each system.

Table A2 Cointegration Test Results

System	Trace Test	Maximum Eigenvalue Test	*k*
$(BR, \dot{p})$	16.2*	11.3	8
(BR, NFR)	36.4*	32.5*	6
$(NFR, \dot{p})$	22.3*	17.1*	8

*Significant at the 5 percent level.

Notes: Trace and maximum eigenvalue tests are tests of the null hypothesis that there is no cointegrating relation in the system. The test used for cointegration is the one proposed on Johansen and Juselius (1990). The lag length in the relevant VAR system is *k* and is chosen using the likelihood ratio test given in Sims (1980). In particular, the VAR model initially was estimated *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 is the one that results in the rejection of the restricted model.

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