Some Key Empirical Determinants of Short-Term Nominal Interest Rates

Yash P. Mehra

Many previous studies cannot account for the high level of interest rates in the early 1980s. For example, Clarida and Friedman (1983, 1984) demonstrate that relative to the predictions of either a structural or an astructural model, interest rates in the early 1980s were too high. This article suggests that the customary empirical measures that gauge the short-run impact of monetary policy on interest rates have become increasingly noisy in the 1980s and that this factor may have been partly responsible for deterioration in the predictive ability of these interest rate equations.

In most previous studies, the short-run impact of monetary policy on rates has been captured by detrended measures of the real money supply defined either as real M1 or real M2. Such empirical measures of the money supply do not provide a consistent basis of comparison over time when deposit rate ceilings are removed and new liquid financial claims are introduced. These and other financial developments of the 1980s also have altered the underlying relationships between the public's demand for these assets and their traditional economic determinants, including the nominal rate (Hetzel and Mehra 1989; Feinman and Porter 1992). As a result, many of the published reduced forms for nominal rates that gauge the short-run impact of monetary policy by the

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real money supply perform poorly in predicting the actual behavior of rates in the 1980s.[1]

In this article, I present an interest rate equation in which the short-run impact of monetary policy on the real component of rates is captured by changes in the real federal funds rate rather than in the real money supply.[2] In addition, I distinguish between the short- and long-run empirical determinants of rates using cointegration and error-correction methodology. In previous short-rate studies, stationarity properties of the data have largely been ignored, thereby muddling the important distinction between the short- and long-run determinants of rates.[3]

The empirical work presented here focuses on the behavior of one-year Treasury bill rates and finds that inflation is the main long-run economic determinant of the level of the nominal rate. Conversely, several customary empirical measures of fiscal policy and a measure capturing foreign capital inflows are not significant when included in the long-run (cointegrating) interest-inflation regression. Also, in the short run, changes in the nominal rate depend largely upon changes in inflation, real output, and the real federal funds rate. Changes in fiscal policy measures and foreign capital inflows do not affect the nominal rate even in the short run.

The short-run interest rate equation estimated here does not exhibit any simultaneous equation bias. The real funds rate is therefore exogenous in the

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[1] Other studies that have addressed this issue are those of Peek and Wilcox (1987) and Hendershott and Peek (1992). Peek and Wilcox continue to employ the real money supply measure and attribute the high level of rates in the early 1980s to a less accommodative monetary policy. They, however, use dummy variables to capture differences in the average tightness of policy during different Fed chairman regimes. Hendershott and Peek, on the other hand, abandon the money supply measure and use instead innovations in the slope of the term structure (the ratio of the 6- to 60-month Treasury rates) to measure the short-run impact of monetary policy on the real component of short-term rates. Given this proxy, monetary policy is highly relevant in explaining the behavior of rates in the early 1980s.

[2] Unlike the slope of the term structure used in Hendershott and Peek (1992), the nominal federal funds rate has been the main instrument of monetary policy during most of the sample period studied here. I use the real funds rate because in the short run the Fed influences the nominal rate mainly by affecting its real component. Moreover, the real funds rate is correlated with the nominal funds rate in the short run. The simple correlation between these two variables is 0.46. Recently, Goodfriend (1993) has used the federal funds rate to measure the short-run impact of monetary policy on the real component of the long rate.

[3] This distinction is important in describing the effect of monetary policy on the nominal rate. Many analysts believe that monetary policy can control the level of the nominal rate in the long run only through its control over inflation, even though in the short run it could have substantial effects. This distinction between short- and long-run effects can be made more precise using cointegration and error-correction methodology. Thus, the statement that monetary policy determines the level of the nominal rate in the long run through its control over inflation can be interpreted to mean that measures of monetary policy are not at the source of long-run, stochastic movements in the level of the nominal rate, but inflation is. That is, the nominal rate is cointegrated with inflation, but not with measures of monetary policy. However, short-run stationary movements in the nominal rate could still be correlated with measures of monetary policy, indicating that in the short run, monetary policy also influences the nominal rate.

short-run equation, which means that the real funds rate is not correlated with contemporaneous shocks to the short rate. Thus, the Federal Reserve does influence the market rate in the short run. In addition, the nominal rate equation that captures the impact of monetary policy by the real funds rate explains reasonably well the nominal rate in the 1980s, although it does not completely solve the puzzle of high rates during the early 1980s. The equation, for example, significantly underpredicts the level of the short rate in 1981. Finally, the results with other measures of short- and medium-term nominal rates indicate that in the short run, nominal yields on the short end of the U.S. Treasury term structure are determined largely by movements in inflation, the real funds rate, and real GDP. The short-run influences of these economic factors, however, decline with the term to maturity, suggesting that the yield curve at the short end is affected most by the outlook for Fed policy and the state of the economy.

The plan of this article is as follows. Section 1 briefly describes the model and the method used in estimating the nominal rate equation. Section 2 presents empirical results, and Section 3 contains concluding observations.

1. THE MODEL AND THE METHOD

An Economic Specification of the Interest Rate Equation

The nominal interest rate equation that underlies the empirical work presented here is based on a variant of a loanable funds model employed by Sargent (1969), among others. According to that model, the nominal rate depends upon anticipated inflation, changes in the real money supply and income, the deficit, and the level of income.[4] The version examined here measures the impact of monetary policy on the nominal rate by changes in the real federal funds rate rather than in the real money supply. In addition, I examine the role of other fiscal policy measures such as government purchases, net taxes, and foreign capital inflows in determining the nominal rate. Since this model has already been described in an earlier paper (Mehra 1994), I report below just the relevant econometric specifications that are used to investigate the behavior of the short rate.

The short-run interest rate equation is estimated using cointegration and error-correction modeling. If the nominal rate and empirical measures of its potential economic determinants are nonstationary, then tests for cointegration provide inferences about the existence of a long-run, equilibrium relationship between the nominal rate and its potential determinants. The error-correction equation then explains short-run changes in the nominal rate.

[4] The nominal rate responds positively to anticipated inflation, the deficit, and real growth and responds negatively to increases in the real money supply. A rise in the level of real income, however, generates a larger volume of savings and hence depresses the equilibrium real rate (Sargent 1969).

The nominal rate equation estimated here has two parts: a long-run part and a short-run part. The long-run part that specifies the potential long-run determinants of the level of the nominal rate is given in (1):

$$R_t = a_0 + a_1\dot{p}_t^e + a_2 RFR_t + a_3 FP_t - a_4 \ln ry_t + a_5 \Delta \ln ry_t + U_t, \tag{1}$$

where R is the nominal interest rate, $\dot{p}^e$ is anticipated inflation, RFR is the real federal funds rate, FP is a fiscal policy variable, $\ln ry$ is the logarithm of real income, and U is the disturbance term. Equation (1) describes the long-run response of the nominal rate to anticipated inflation, the real funds rate, a fiscal policy variable, changes in real income, and the level of real income. The coefficients $a_i, i = 1$ to 5, measure the long-run responses in the sense that they are the sums of coefficients that appear on current and past values of the relevant economic determinants. The term $a_1\dot{p}_t^e$ in (1) captures the inflation premium in the nominal rate, whereas the remaining terms capture the influence of other variables on the equilibrium real component of the short rate. If the nominal rate and anticipated inflation variables are nonstationary but cointegrated as in Engle and Granger (1987), then the other remaining long-run impact coefficients in (1) may all be zero.

Equation (1) may not do well in explaining short-run movements in the nominal rate for a number of reasons. First, it ignores the short-run effects of economic factors. Some economic factors, including those measuring monetary policy actions, may be important in explaining short-run changes in the nominal rate, even though they may have no long-run effects. Second, it completely ignores short-run dynamics. Hence, in order to explain short-run changes in the nominal rate, consider the following error-correction model of the nominal rate:

$$\begin{aligned}\Delta R_t = c_0 + c_1 \Delta \dot{p}_t^e + c_2 \Delta RFR_t + c_3 \Delta FP_t + c_4 \Delta \ln ry_t \\ + c_5 \Delta^2 \ln ry_t + \sum_{s=1}^{n} c_{6s} \Delta R_{t-s} + c_7 U_{t-1} + \epsilon_{1t},\end{aligned} \tag{2}$$

where U_{t-1} is the lagged residual from the long-run nominal equation (1), Δ^2 is the second-difference operator, and other variables are as defined earlier. Equation (2) is the short-run interest rate equation, and the coefficients c_i, $i = 1$ to 5, capture the short-run responses of the interest rate to economic determinants suggested here.

Estimation Issues: OLS Works If Empirical Measures of Economic Determinants Are Nonstationary or Exogenous

Both the long- and short-run equations (1) and (2) contain contemporaneous values of economic fundamentals. Those values are likely to be correlated with contemporary shocks to the nominal rate. In particular, if the Federal Reserve contemporaneously adjusts its short-run funds objective with respect to movements in short rates, then the real federal funds rate is likely to be correlated with the disturbance term in the regression.[5] Hence, these equations cannot be consistently estimated by ordinary least squares unless some special assumptions hold.

The long-run equation (1) can be consistently estimated by ordinary least squares if empirical measures of the economic factors included in (1) are nonstationary but cointegrated as in Engle and Granger (1987). Tests of hypotheses on coefficients that appear in (1) can then be carried out by estimating Stock and Watson's (1993) dynamic OLS regressions. I, therefore, test first for nonstationarity and cointegration. The empirical work here examines the stationarity properties of the data using unit root and mean stationarity tests. The test for cointegration used is the one proposed in Johansen and Juselius (1990).[6]

The economic variables that appear in the short-run equation (2) are stationary. This equation can be consistently estimated by ordinary least squares if contemporaneous right-hand side explanatory variables are uncorrelated with the disturbance term. That condition can be tested by performing the test for exogeneity given in Hausman (1978). To implement the test, consider the following VAR representation of these contemporaneous right-hand side explanatory variables. (For simplicity, I am ignoring fiscal policy and some other variables.)

$$\begin{aligned}\Delta\dot{p}^e_t = d_0 + \sum_{s=1}^{k} d_{1s}\Delta\dot{p}^e_{t-s} + \sum_{s=1}^{k} d_{2s}\Delta RFR_{t-s} \\ + \sum_{s=1}^{k} d_{3s}\Delta \ln ry_{t-s} + \sum_{s=1}^{k} d_{4s}\Delta R_{t-s} + \epsilon_{2t}\end{aligned} \tag{3}$$

$$\begin{aligned}\Delta \ln ry_t = e_0 + \sum_{s=1}^{k} e_{1s}\Delta\dot{p}^e_{t-s} + \sum_{s=1}^{k} e_{2s}\Delta RFR_{t-s} \\ + \sum_{s=1}^{k} e_{3s}\Delta \ln ry_{t-s} + \sum_{s=1}^{k} e_{4s}\Delta R_{t-s} + \epsilon_{3t}\end{aligned} \tag{4}$$

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[5] To illustrate, consider a scenario in which the incoming new data indicates that in the current quarter real growth or inflation is going to be higher than what the market expected based on the past information. If the market believes such information, short rates could rise because accelerations in real growth or inflation are generally associated with higher rates. If the Fed also reacts contemporaneously to such new information and the resulting rise in short rates, then changes in the funds rate would be correlated with the disturbance term. Such correlation will be absent, however, if the Fed does not react or reacts with a lag.

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

$$\Delta RFR_t = f_0 + \sum_{s=1}^{k} f_{1s} \Delta \dot{p}^e_{t-s} + \sum_{s=1}^{k} f_{2s} \Delta RFR_{t-s} + \sum_{s=1}^{k} f_{3s} \Delta \ln ry_{t-s} + \sum_{s=1}^{k} f_{4s} \Delta R_{t-s} + \epsilon_{4t} \tag{5}$$

This VAR includes only past values of the economic factors that appear in the economic model used here and hence can be consistently estimated by ordinary least squares. Then consider an expanded version of equation (2) given below:

$$\Delta R_t = d_0 + d_1 \Delta \dot{p}^e_t + d_2 \Delta RFR_t + d_3 \Delta \ln ry_t + \sum_{s=1}^{n} d_4 \Delta R_{t-s} + d_5 U_{t-1} + d_6 \hat{\epsilon}_{2t} + d_7 \hat{\epsilon}_{3t} + d_8 \hat{\epsilon}_{4t} + \epsilon_{1t}, \tag{6}$$

where $\hat{\epsilon}_2, \hat{\epsilon}_3$, and $\hat{\epsilon}_4$ are residuals from the VAR. If $d_6 = d_7 = d_8 = 0$, then $\Delta \dot{p}^e_t, \Delta RFR_t$, and $\Delta \ln ry$ are uncorrelated with the disturbance term and hence exogenous in this equation. The hypothesis $d_6 = d_7 = d_8 = 0$ can be tested using the F-test.[7]

Data and Definition of Variables

The empirical work uses quarterly data from 1955:1 to 1994:3. The short-term nominal rate, $R1$, is the nominal yield on one-year U.S. Treasury bills. I consider two proxies for anticipated inflation. The first one uses actual inflation as measured by the behavior of the consumer price index ($\dot{p}$). The second one uses one-year-ahead inflation rates from the Livingston survey ($\dot{p}^e$). The real federal funds rate, RFR, is the nominal federal funds rate minus the actual, annualized quarterly inflation rate. Nominal interest rate data are observations from the last month of the quarter, and inflation ($\dot{p}$) is calculated as the change in the log form from the last month of the previous quarter price level to that of the current. In some specifications in which the real money supply is used to measure the impact of monetary policy actions on the real component of the nominal rate, the measure of money used is M2 scaled by the real GDP deflator. Real income, ry, is real GDP. The real deficit scaled by real GDP (DEF/y),[8] real government purchases (fg), and real government tax (net of transfers) receipts (tx) are alternatively used to measure the impact of fiscal policy on the real component of the short rate. I also consider the impact of

[7] The hypothesis that the right-hand side contemporary regressors in (2) are independent of the disturbance term also can be tested by comparing ordinary least squares and instrumental variables estimates of the equation. Under the null hypothesis that there is no simultaneity equation bias, OLS estimates ($\hat{\beta}_{OLS}$) should not be statistically different from IV estimates ($\hat{\beta}_{IV}$). Hausman (1978) shows that the statistic that tests the null hypothesis $\hat{\beta}_{OLS} = \hat{\beta}_{IV}$ is distributed Chi-squared with a degree of freedom parameter equal to the number of parameters estimated in the equation. See Maddala (1988) for a simple description of these test procedures.

[8] This specification reflects the assumption that in a growing economy higher deficits result in higher rates only if the deficit rises relative to GDP.

foreign capital inflows measured as the ratio of U.S. Treasury securities held by foreigners to the total of U.S. securities held by domestic and foreign residents (fh).[9]

2. ESTIMATION RESULTS

On the Long-Run Determinants of the Nominal Rate

I first present test results that help determine which economic determinants suggested in the long-run equation (1) are relevant.

Table 1 presents test results for determining whether empirical measures of potential determinants such as $R1, \dot{p}, \dot{p}^e, DEF/y, \ln fg, \ln tx, fh, \ln rM2, \ln ry$, and RFR 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 all these series. 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 with the exception of RFR. These results indicate that $R1, \dot{p}, \dot{p}^e, DEF/y, \ln fg, fh, \ln tx, \ln rM2_t$, and $\ln ry$ have a unit root and thus are nonstationary in levels.[10] The results are inconclusive for the real funds rate RFR. Together, these results indicate that most empirical measures of the potential determinants suggested here are nonstationary and hence could be the source of long-run stochastic movements in the nominal rate.

Table 2 presents test statistics for determining whether the short-term nominal rate ($R1$) is cointegrated with any of these nonstationary measures of inflation, fiscal and monetary policies, and foreign capital inflows. Trace and maximum eigenvalue statistics, which test the null hypothesis that there is no cointegrating vector, are large for systems ($R1, \dot{p}$), ($R1, \dot{p}^e$), ($R1, \ln ry$), ($R1, DEF/y$), and ($R1, \ln rM2$), but are very small for systems ($R1, RFR$), ($R1, \ln fg$), ($R1, \ln tx$), and ($R1, fh$). These results indicate that the nominal rate is cointegrated with inflation (actual or anticipated), the real money supply, the level of income, and the deficit, but not with the real federal funds rate, government purchases, net taxes, and foreign capital inflows. The evidence continues to favor the presence of at least one cointegrating vector even in expanded systems that include inflation and fiscal and monetary policy variables

[9] The data on the Livingston survey are provided by the Philadelphia Fed. The data used in measuring capital inflows are from the Federal Reserve Board's flow of funds data. All other data series are from the Citibank database.

[10] The t-statistic that tests the null hypothesis that first differences of a series have a unit root takes values −5.2, −4.9, −6.4, −15.9, −4.2, −6.2, −5.2, −5.6, −4.7, and −7.4 for $\Delta R1$, $\Delta \dot{p}$, $\Delta \dot{p}^e$, ΔRFR, $\Delta \ln rM2$, $\Delta DEF/y$, $\Delta \ln fg$, $\Delta \ln tx$, Δfh, and $\Delta \ln ry$, respectively. These t-values are large, indicating that first differences of these series are stationary. The 5 percent critical value taken from Fuller (1976) is −2.9.

Table 1 Tests for Unit Roots and Mean Stationarity

Series X	Panel A Tests for Unit Tools			Panel B Tests for Mean Stationarity
	$\hat{\rho}$	$t_{\hat{\rho}}$	k	$\hat{n}_u$
$R1$	0.94	−1.93	2	0.84*
$\dot{p}$	0.84	−2.83	7	0.49*
$\dot{p}^e$	0.98	−1.80	2	0.98*
RFR	0.85	−2.51	2	0.37
$\ln rM2$	0.99	−1.87	1	1.80*
$\ln ry$	0.97	−1.80	1	0.33
DEF/y	0.92	−2.54	1	1.42*
$\ln fg$	0.98	−1.63	3	0.91*
$\ln tx$	0.96	−1.54	1	1.54*
fh	0.98	−1.58	5	1.19*

* Significant at the 5 percent level.

Notes: $R1$ is the one-year Treasury bill rate; $\dot{p}$ is the annualized quarterly inflation rate measured by the consumer price index; $\dot{p}^e$ is the Livingston survey measure of one-year-ahead expected inflation; RFR is the real federal funds rate; $rM2$ is the real money supply; ry is real GDP; DEF/y is the ratio of federal government deficits to nominal GDP; fg is real federal government purchases; tx is real federal government tax (net of transfers) receipts; and fh is the ratio of U.S. Treasury securities held by foreigners to total of U.S. Treasury securities held by domestic and foreign residents. The sample period studied is 1955:1 to 1994:3. 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 $k\,\mathrm{max}$, chosen a priori (eight quarters here). Estimate (a) above with k set at $k\,\mathrm{max}$. If the last included lag is significant, select $k = k\,\mathrm{max}$. 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.

[see systems $(R1, \dot{p}, DEF/y, \ln rM2, \ln ry)$, $(R1, \dot{p}^e, DEF/y, \ln rM2, \ln ry)$, $(R1, \dot{p}, DEF/y, \ln ry)$, and $(R1, \dot{p}^e, DEF/y, \ln ry)$].

Panels A and B in Table 3 help determine which variables included in the cointegrating regression[11] are statistically significant. It presents the dynamic OLS estimates of the potential cointegrating regressions with and without the real money supply. As can be seen, inflation (actual or anticipated) is the only variable that enters significantly in these cointegrating regressions. Other

[11] In this article, I focus on a single cointegrating regression that is normalized on the short-term nominal rate. The analysis here thus ignores the possibility that in larger systems there may be multiple cointegrating vectors.

Table 2 Cointegration Test Results

System	Trace Test	Maximum Eigenvalue Test	k
$(R1,\ \dot{p})$	22.2*	17.9*	8
$(R1,\ \dot{p}^e)$	22.1*	18.9*	2
$(R1,\ RFR)$	17.6	14.3	8
$(R1,\ \ln rM2)$	20.3*	16.9*	8
$(R1,\ \ln fg)$	10.1	5.3	2
$(R1,\ \ln tx)$	12.6	8.5	4
$(R1,\ DEF/y)$	30.7*	27.5*	8
$(R1,\ fh)$	14.7	10.7	8
$(R1,\ \ln ry)$	48.6*	43.4*	2
$(R1,\ \dot{p},\ DEF/y,\ \ln ry)$	97.1*	42.9*	4
$(R1,\ \dot{p}^e,\ DEF/y,\ \ln ry)$	86.1*	34.9*	4
$(R1,\ \dot{p},\ DEF/y,\ \ln rM2,\ \ln ry)$	148.8*	65.2*	2
$(R1,\ \dot{p}^e,\ DEF/y,\ \ln rM2,\ \ln ry)$	146.4*	64.9*	2

* Significant at the 5 percent level.

Notes: Trace and maximum eigenvalue tests are tests of the null hypothesis that there is no cointegrating vector in the system. 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 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 is the one that results in the rejection of the restricted model.

nonstationary variables such as the real deficit and the real money supply are not significant. Real GDP is significant in some regressions and not in others (see Panels A and B, Table 3). These results thus indicate that inflation is the main long-run economic determinant of the short-term nominal rate.

Panel C in Table 3 presents the cointegrating regressions that include only the inflation variable. The cointegrating regression is estimated with and without the restriction that the nominal rate adjusts one for one with inflation in the long run. The χ^2 statistic that tests the validity of the full Fisher-effect restriction is not large, indicating that this restriction is consistent with data. These results also indicate that the real rate of interest on one-year Treasury bills is mean stationary. The estimate of this mean falls in a 2.3 to 3.2 percent range (see the constant term in Panel C regressions).

On the Short-Run Determinants of the Nominal Rate

The short-run equation (2) is estimated here jointly with its long-run part and hence includes levels as well as first differences of the relevant economic

Table 3 Cointegrating Regressions; Dynamic OLS

(Leads, Lags)	
	Panel A: With Real Money Supply
(−4, 4)	$R1_t = 0.7\dot{p}_t - 17.6 \ln rM2_t + 2.05 \ln ry_t + 0.22(DEF/y)_t$ (3.9) (0.9) (1.0) (0.5)
(−4, 4)	$R1_t = 1.1\dot{p}^e_t + 6.1 \ln rM2_t + 6.8 \ln ry_t + 0.02(DEF/y)_t$ (5.9) (0.4) (0.4) (0.1)
	Panel B: Without Real Money Supply
(−4, 4)	$R1_t = 0.9\dot{p}_t - 1.0 \ln ry_t + 0.18(DEF/y)$ (21.8) (2.2) (1.5)
(−4, 4)	$R1_t = 1.1\dot{p}^e_t - 1.5 \ln ry_t - 0.10(DEF/y)_t$ (23.1) (12.9) (0.7)
	Panel C: With Inflation Only
(−4, 4)	$R1 = 3.2 + 0.8\dot{p}_t; R1 = 2.3 + 1.0\dot{p}_t; \chi^2(1) = 2.8(0.10)$
(−4, 4)	$R1 = 2.5 + 1.1\dot{p}^e_t; R1 = 2.6 + 1.0\dot{p}^e_t; \chi^2(1) = 0.32(0.57)$

Notes: All regressions 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. Parentheses contain t-values corrected for the presence of moving average serial correlation. $\chi^2(1)$ is the χ^2 statistic with one degree of freedom (significance levels in parentheses); it tests the hypothesis that the coefficient on $\dot{p}$ or $\dot{p}^e$ is unity.

determinants.[12] A preliminary specification search indicated that in the short run, changes in the nominal rate depend largely upon contemporaneous changes in inflation, the real funds rate, and real GDP. The lagged level of the funds rate is also significant. The empirical measures of fiscal policy and foreign capital inflows, however, did not enter the short-run equation. (I formally test these restrictions later on.)

Table 4 presents ordinary least squares as well as instrumental variables estimates of the pertinent short-run equation. The instruments chosen are basically the lagged values of the right-hand side explanatory variables that appear

[12] The short-run equation (2) of the text includes a one-period lagged value of the residual from the long-run equilibrium equation. In joint estimation, the lagged residual is replaced by lagged levels of the variables that enter the long-run equilibrium equation. To see it, assume for the sake of explanation that the long-run cointegrating regression is given in (b) below:

$$R1_t = a_0 + a_1\dot{p}^e_t + U_t. \tag{b}$$

If we solve (b) for U_{t-1} and substitute for U_{t-1} into (2), then the short-run equation (2) will include the lagged level of the nominal rate ($R1$) and the inflation rate ($\dot{p}^e$).

Table 4 Short-Run Nominal Interest Rate Equations, 1957:1 to 1994:3

Explanatory Variables	Instrumental Variables (A.1)	(A.2)	Ordinary Least Squares (B.1)	(B.2)
constant	−0.22(1.4)	1.0 (2.9)	0.06(0.8)	0.6 (3.8)
$\Delta \dot{p}_t$	0.71(8.2)		0.72(12.8)	
$\Delta \dot{p}^e_t$		0.9 (3.2)		0.8 (3.3)
ΔRFR_t	0.61(6.5)	0.11(1.2)	0.64(11.3)	0.17(3.5)
$\Delta \ln ry_t$	0.12(2.6)	−0.02(0.3)	0.06(3.9)	0.06(2.1)
$R1_{t-1}$	−0.37(3.4)	−0.55(4.9)	−0.33(4.1)	−0.53(6.1)
$\dot{p}_{t-1}$	0.37(3.4)		0.33(4.1)	
$\dot{p}^e_{t-1}$		0.55(4.9)		0.53(6.1)
RFR_{t-1}	0.34(2.7)	0.28(3.3)	0.30(3.3)	0.30(4.7)
$\Delta R1_{t-1}$	−0.23(6.2)		−0.21(5.0)	
$\Delta R1_{t-2}$	0.07(1.2)		0.04(0.9)	
$\Delta R1_{t-3}$	−0.04(0.5)		−0.06(0.9)	
$\Delta R1_{t-4}$	0.15(2.9)		0.14(2.9)	
SER	0.547	0.98	0.52	0.94
Q(36)	33.8	27.7	37.1	38.9
Q(8)	6.9	5.9	9.2	11.9
Q(4)	2.8	3.9	2.2	8.5
Sargan's χ^2	12.5(0.19)	6.5(0.36)		
F	1.4(0.24)	1.0(0.37)		

* Significant at the 5 percent level.

Notes: See notes in Table 1 for definition of variables. Parentheses following coefficients contain t-values. SER is the standard error of estimate, and Q(36), Q(8), and Q(4) are the Ljung-Box Q-statistics based on 36, 8, and 4 autocorrelations of the residuals.

Sargan's χ^2 tests the independence of the instruments and the disturbance term (parentheses contain the significance level of the test). F is the F-statistic that tests the null hypothesis that the residuals from the reduced-form regressions of real growth, inflation, and the real federal funds rate are not jointly significant when included in the interest rate equation (parentheses contain the significance level of the test). See footnote 13 of the text for a description of instruments used.

in (2).[13,14] The null hypothesis that contemporary values of changes in inflation, the real funds rate, and real growth are jointly exogenous with respect to

[13] When actual inflation is the proxy for anticipated inflation, the instruments used for estimating the short-run equation are a constant, one-period lagged value of the short rate, inflation, and the real funds rate, two-period lagged values of the change in the short rate, and four-period lagged values of changes in inflation, real income, and the real funds rate. When the Livingston survey is the proxy for anticipated inflation, I use similar instruments, but I treat the Livingston survey as exogenous in the short equation. Hence, first differences and one-period lagged values of the Livingston survey ($\Delta \dot{p}^e_t$, $\dot{p}^e_{t-1}$) are used as instruments.

[14] Two considerations are important in the choice of instruments. First, the instruments chosen should be uncorrelated with the disturbance term. Second, they should be highly correlated

short-run impact coefficients is not rejected (the relevant Hausman F-statistics reported in Table 4 are small).[15] Sargan's (1964) χ^2 specification test statistic presented in Table 4 is also small, indicating that the instruments chosen are independent of the disturbance term. A casual look at the estimates reported in Table 4 indicates that instrumental variables estimates of short-run impact coefficients are not strikingly different from OLS estimates.[16] These results suggest that in the short-run equation, changes in inflation, the funds rate, and real income are not correlated with contemporary shocks to the short rate. Hence, the estimates of short-run impact coefficients reported here are consistent.[17]

The short-run equation is reported for the full sample 1955:1 to 1994:3.[18] As can be seen, the relevant explanatory variables have coefficients that are

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with the contemporaneous values of endogenous variables. In the short-run equation estimated here, lagged endogenous variables are valid as instruments if the equation does not exhibit any serial correlation. The Ljung-Box Q-statistics reported in Table 4 indicate that serial correlation is not a problem in the regressions reported there. As regards the second point, lagged endogenous variables are good instruments because they are likely to be highly correlated with the contemporaneous endogenous variables. However, these variables may not be strictly exogenous in the sense that they are completely independent of past shocks to economic variables, including the nominal rate. Thus, changes in the real funds rate may be uncorrelated with the disturbance term in the regression but may not be independent of the past behavior of economic fundamentals, including the short rate.

[15] The individual t-statistics that appear on the residuals from the reduced-form regressions of inflation, the funds rate, and real growth (the t-statistics for $d_6 = 0, d_7 = 0$, or $d_8 = 0$ in (6)) are not large in the short-run interest rate equation either. In contrast, the null hypothesis that in the real funds rate equation contemporaneous values of changes in inflation, real GDP, and the nominal rate are jointly exogenous is usually rejected by the F-test.

[16] The null hypothesis that instrumental variables estimates of the short-run equation (2) of the text are not jointly different from OLS estimates is not rejected by the Hausman test described in footnote 7. For equations A.1 and A.2 of Table 4, the relevant χ^2 statistics are 2.1 and 2.6, respectively. Both these statistics are small.

[17] If one begins with a structural equation in which the nominal rate depends upon contemporaneous anticipated values of economic variables, then lagged values are valid as instruments for unobservables if the order of lag in instruments chosen exceed the order of lag in the serial correlation of the disturbance term. The empirical work here does not begin with any particular structural equation. It does, however, assume that the disturbance term in the short-run equation does not have any serial correlation. Nevertheless, in order to check the robustness of results to the order of lag in instruments, I also reestimated the short-run equations using successively more than one-period lagged values of the right-hand side explanatory variables as instruments, going as far back as five- through eight-period lags. The point-estimates of the short-run impact coefficients that appear on inflation, real GDP, and the funds rate move around somewhat, but they continue to have expected signs and are generally significant. As expected, standard errors of the estimated coefficients increase as the order of lag in instruments chosen increases.

[18] The instrumental variables estimates of the short-run equations for the subsample 1955:1 to 1979:3 are given below, and those estimates look very similar to the ones for the whole period.

$$\Delta R1_t = 0.06 + \underset{(8.7)}{0.76}\Delta \dot{p}_t + \underset{(6.8)}{0.67}\Delta RFR_t + \underset{(0.0)}{0.00}\Delta \ln ry_t - \underset{(2.2)}{0.16}\Delta R1_{t-1} - \underset{(2.2)}{0.15}\Delta R1_{t-2}$$
$$- \underset{(1.7)}{0.13}R1_{t-1} + \underset{(1.7)}{0.13}\dot{p}_{t-1} + \underset{(1.5)}{0.12}RFR_{t-1}$$

of expected signs and statistically significant. Thus, the nominal rate responds positively to short-run increases in inflation, real GDP, and the real funds rate. The coefficients that appear on contemporary values of these variables range from 0.7 to 0.9 for inflation, 0.1 to 0.6 for the real funds rate, and 0.0 to 0.12 for real growth. Thus, a one percentage point rise in inflation raises the short rate between 70 and 90 basis points, whereas a similar increase in the real funds rate raises it by 10 to 60 basis points in the short run. A one percentage point rise in the growth rate of real GDP raises the short rate by about 12 basis points in the regression that uses actual inflation data.[19]

The short-run equations reported in Table 4 embody the long-run relationship among the levels of economic determinants. The null hypothesis that coefficients appearing on one-period lagged levels of inflation and the nominal rate sum to zero is not rejected, indicating that the nominal rate adjusts one for one with inflation in the long run. The coefficient that appears on the lagged level of the real funds rate is large and remains statistically significant, indicating that (stationary) movements in the real funds rate have substantial short-run effects on the real component of the short rate.

The short-run equations reported in Table 4 do not include any fiscal policy measure. Nor do they allow for the effect of foreign capital inflows. At this point I formally test the hypothesis that these variables have no significant effects on the nominal rate. Table 5 presents Lagrange multiplier tests for omitted variables.[20] Those test results indicate that the real deficit, real government

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$$\Delta R1_t = \underset{}{1.0} + \underset{(3.4)}{1.0\Delta \dot{p}^e_t} + \underset{(2.0)}{0.29\Delta RFR_t} - \underset{(0.1)}{0.06\Delta \ln ry_t} + \underset{(1.4)}{0.17\Delta R1_{t-1}}$$

$$- \underset{(3.3)}{0.54R1_{t-1}} + \underset{(3.3)}{0.54\dot{p}^e_{t-1}} + \underset{(2.5)}{0.38RFR_{t-1}}$$

Parentheses contain t-values.

[19] The empirical work here uses actual values of the real funds rate, in contrast with previous studies in which the short-run impact of monetary policy on rates is captured by employing innovations in the pertinent money supply or term-structure measure. The latter approach reflects two assumptions. First, anticipated values of these variables affect the short rate by altering its expected-inflation component. In the short run, only unanticipated changes have an effect on the real component of the short rate. Second, it is possible to decompose the pertinent monetary policy measure into its anticipated and unanticipated components. Some of these assumptions are questionable. Nevertheless, I also estimate the short-run equation using residuals from the short-run real funds rate equation like (5) as a measure of unanticipated monetary policy actions. The coefficient that appears on this measure of policy remains positive and is statistically significant. The estimated coefficient is 0.46 (t-value = 8.6) when the actual inflation data is used, whereas it is 0.20 (t-value = 3.2) when the expected inflation measure is used. These results indicate that monetary policy effects when using the actual funds rate are not spurious.

[20] A Lagrange multiplier test for a set of p-omitted variables is constructed by regressing the model's residuals on both the set of original regressors and on the set of omitted variables. If the omitted variables do not belong in the equation, then multiplying the R^2 statistic from this regression by the number of observations will produce a statistic distributed as χ^2 with p degrees of freedom (Engle 1984; Breusch and Pagan 1980).

Table 5 χ^2 Tests for Omitted Variables

Candidate Variable X	Lags (0 to k)	Equation B.1	Equation B.2
$\Delta(DEF/y)$	(0,0)	0.0 (0.96)	1.5 (0.21)
	(0,1)	0.7 (0.68)	1.5 (0.47)
	(0,4)	6.1 (0.30)	4.6 (0.47)
$\Delta \ln fg$	(0,0)	0.0 (0.80)	0.9 (0.33)
	(0,1)	1.8 (0.40)	1.2 (0.55)
	(0,4)	3.9 (0.54)	1.8 (0.87)
$\Delta \ln tx$	(0,0)	0.2 (0.28)	2.2 (0.14)
	(0,1)	2.4 (0.30)	4.7 (0.10)
	(0,4)	4.2 (0.51)	7.3 (0.20)
Δfh	(0,0)	1.2 (0.27)	0.1 (0.71)
	(0,1)	3.2 (0.20)	0.4 (0.81)
	(0,4)	6.1 (0.30)	1.6 (0.89)

Notes: See notes in Table 1 for definition of variables. The statistics reported are the Wald test of the null hypothesis that the pertinent variable is not an omitted variable from the relevant regression. If this statistic is large for some variable, then the variable should be included in the regression. Parentheses contain significance levels of the test.

purchases, net taxes, and foreign capital inflows do not enter the short-run interest equation. Overall, these results assign no significant role to fiscal policy measures and foreign capital inflows in explaining short-run movements in the nominal rate.

Predicting the Behavior of the Nominal Rate in the 1980s

I now examine whether the short-run equation estimated here can predict the actual behavior of the nominal rate during the 1980s. The predicted values used are the out-of-sample, one-year-ahead dynamic forecasts that cover the subperiod from 1979 to 1993. I focus on the equation that uses actual inflation.

Table 6 presents predicted values generated using the interest rate regression presented in Table 4. Actual values and prediction errors are also reported there. As can be seen, this equation predicts reasonably well the actual behavior of the nominal rate during this period. The mean error is small, only 3 basis points, and the root mean squared error is 0.4 percentage points. This regression outperforms a purely eight-order autoregressive model of the short rate. For the time series model, the mean prediction error is 51 basis points and the root mean squared error is 1.3 percentage points.[21]

[21] The interest rate regression with the real funds rate also outperforms the version in which changes in the real funds rate are replaced by changes in the real money supply. The mean error is 7 basis points and the root mean squared error is 1.22 percentage points.

Table 6 Predictive Performance

Year	**Actual (*A*)**	**Predicted (*P*)**	**Error**
Panel A: Actual and Predicted One-Year Treasury Bill Rate, 1979 to 1993			
1979	10.7	11.2	−0.5
1980	12.6	12.2	0.4
1981	14.5	13.5	1.0*
1982	11.9	12.2	−0.3
1983	9.7	9.9	−0.2
1984	10.9	10.8	0.1
1985	8.3	8.9	−0.6
1986	6.3	6.7	−0.3
1987	6.9	6.4	0.5
1988	7.8	7.9	−0.1
1989	8.5	9.0	−0.6
1990	7.8	7.5	0.3
1991	5.7	5.4	0.3
1992	3.9	−0.1	0.3
1993	3.5	3.2	0.2
Mean Error			0.03[−0.051]
RMSE			0.43[1.31]
Panel B: $A_t = d_0 + d_1 P_t + e_t$			
Model	d_0	d_1	F
Interest Rate Equation (A.1, Table 4)	0.05	1.0	0.1
AR(8)	0.48	0.89	3.1

* The predictor error is twice the root mean squared error.

Notes: The predicted values reported above are generated using the regression A.1 of Table 4. AR(8) is an eight-order autoregressive process for explaining changes in the one-year Treasury bill rate. RMSE is the root mean squared error. Brackets contain the mean error and the RMSE value generated using the time series model. F is the F-statistic that tests the null hypothesis that $d_0 = 0$ and $d_1 = 1$.

I evaluate further the predictive performance of the interest rate equation with the real funds rate from 1979:1 to 1993:4 by estimating regressions of the form

$$A_t = d_0 + d_1 P_t,$$

where A is the quarterly value of the short rate and P is the value predicted by the short-rate regression. If $d_0 = 0$ and $d_1 = 1$, then regression forecasts are unbiased. As can be seen from Table 6, the coefficients d_0 and d_1 take values

0.05 and 1.0, respectively, for the interest rate equation (A.1 of Table 4) and 0.48 and 0.89, respectively, for the time series model. The hypothesis $d_0 = 0$ and $d_1 = 1$ is rejected for the time series model, but not for the interest rate equation.

Despite this overall good predictive performance, the interest rate equation with the real funds rate does not completely solve the puzzle of high rates during the early 1980s. The equation significantly underpredicts the level of the nominal rate in 1981. It predicts very well, however, the declines in nominal rates that have occurred since 1990.[22]

The Short End of the Term Structure Is Dominated by the Outlook for Inflation, Fed Policy, and the State of the Economy

The empirical analysis of one-year Treasury bills summarized in Tables 4 and 5 indicate that in the short run, changes in the nominal rate depend largely upon changes in inflation, real GDP, and the real funds rate. I now argue that the same economic factors pretty much determine the behavior of the short end of the U.S. Treasury term structure. Table 7 presents short-run coefficients that appear on these economic determinants when alternative measures of short- to medium-term interest rates are used in the regression (2). As can be seen, those estimates indicate that changes in inflation, real GDP, and the real funds rate influence most the short end of the term structure. The short-run impact coefficients that appear on these variables steadily decline in size with the term to maturity. These results indicate that the yield curve at the short end of the term structure is dominated by the outlook for inflation, Fed policy, and the state of the economy.

3. CONCLUDING OBSERVATIONS

It is a widely held view both in financial press and academic circles that the Federal Reserve influences short-term nominal interest rates. The empirical work presented here provides one perspective on the potential role of the Federal Reserve in determining short-term rates. The results indicate that it is inflation, not the real federal funds rate, that is at the source of long-run stochastic movements in the level of the nominal rate. Therefore, only through its control over inflation can the Federal Reserve exercise control over the level of the nominal rate in the long run. In other words, the Federal Reserve cannot permanently lower the nominal rate by affecting its real component.

[22] In contrast, the interest rate regression with the real money supply performs very poorly in predicting the behavior of the nominal rate during the early 1980s. Nor does it predict well the declines in nominal rates that have occurred since 1990.

Table 7 Term Structure Effects

Dependent Variable	Coefficients (t-values) on Contemporary Values of $\Delta\dot{p}_t$	ΔRFR_t	$\Delta \ln ry_t$
R3M	0.73 (8.9)	0.57 (7.5)	0.14 (3.5)
R1	0.71 (8.2)	0.62 (6.4)	0.11 (2.6)
R3	0.51 (6.9)	0.43 (4.5)	0.07 (1.4)
R5	0.41 (5.9)	0.33 (3.6)	0.05 (1.2)
R10	0.29 (5.0)	0.22 (2.8)	0.03 (0.9)

Notes: R3M is the three-month Treasury bill rate; R1, R3, R5, and R10 are the nominal yields on one-year, three-year, five-year, and ten-year Treasury bills. The regression equation used is equation A.1 reported in Table 4. In all of these regressions, the long-run coefficient that appears on the level of the inflation rate ($\dot{p}$) is constrained to be unity, so that in the long run the nominal rate adjusts one for one with inflation. The regressions are estimated by instrumental variables over the sample period 1957:1 to 1994:3. Parentheses above contain t-values.

The results also indicate that in the short run, changes in the real funds rate have considerable effects on the nominal rate. Moreover, the real funds rate variable is exogenous in the short-run interest rate equations, indicating that short-run changes in the real federal funds rate do not respond to contemporaneous movements in the nominal rate. This finding, however, is quite consistent with the possibility that in reduced-form regressions like (5), changes in the real funds rate are highly correlated with lagged values of economic variables, including the nominal rate. Together, these results indicate that in the short run, the Federal Reserve influences the market as well as may be influenced by it.

Fiscal policy measures such as the deficit, government purchases, net taxes, and foreign capital inflows do not affect the short rate, once one controls for the effects of inflation, the real funds rate, and real growth. Overall, the results indicate that in the short run, the behavior of short-term nominal rates is dominated by the outlook for inflation, the funds rate, and the state of the economy.

The interest rate equation predicts reasonably well the nominal rate during the 1980s. It does not, however, completely solve the puzzle of high rates during the early 1980s, particularly in 1981. The analysis here indicates that a significant part of the rise in rates in the early 1980s can be attributed to the behavior of inflation and the real federal funds rate. Since over long periods these variables are endogenously determined, a complete explanation of the behavior of short rates must include an explanation of the behavior of these two variables.

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