HAVE MONEY-STOCK FLUCTUATIONS HAD A LIQUIDITY EFFECT ON EXPECTED REAL INTEREST RATES?

Behzad T. Diba
Georgetown University

Seonghwan Oh
UCLA

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ABSTRACT

Have Money-Stock Fluctuations Had a Liquidity
Effect on Expected Real Interest Rates?

This paper reports some empirical evidence on the relation between the expected real interest rate and monetary aggregates in postwar U.S. data. We find some evidence against the hypothesis, implied by the Real Business Cycle model of Litterman and Weiss (1985), that the expected real interest rate follows a univariate autoregressive process, not Granger-caused by monetary aggregates. But we find less evidence against a more general bivariate model--suggested by what Barro (1987) refers to as "the basic market-clearing model." The evidence against both models becomes much weaker when we take the monetary base, rather than M2, as our measure of the money stock. Using the bivariate model as our null hypothesis, we find no evidence that money-stock fluctuations have had a short-run liquidity effect on the real interest rate.

1. INTRODUCTION

Conventional macroeconomic models (such as the standard IS-LM model) imply that an increase in the money supply has a "liquidity effect" that reduces the expected real interest rate and, thereby, stimulates investment and output. Recent contributions to the "real business cycle"--henceforth, RBC--literature, however, contend that monetary disturbances have not played an important role in postwar U.S. business cycles.[1] As King and Plosser (1984) emphasize, the observed correlations of real variables with monetary aggregates may reflect the endogenous response of inside money to changes in the real variables, rather than the structural effect of changes in outside money on the real variables.

This paper examines the relationship between monetary aggregates and the expected real interest rate, in postwar U.S. data. We consider two specifications of the process generating the expected real interest rate. In the first specification, under the null hypothesis, the expected real interest rate follows a univariate autoregression. In the second specification, under the null hypothesis, the expected real interest rate depends not only on its own lagged values, but also on lagged values of output.

The motivation for specifying the generating process of the expected real interest rate as a univariate autoregression is the empirical evidence reported by Litterman and Weiss (1985)--henceforth, L&W. They begin their empirical analysis by replicating the finding of Sims (1980, 1982) that the inclusion of nominal interest rates in a vector autoregression (VAR) estimated on postwar US data eliminates much of the explanatory power of money-stock innovations for output fluctuations.[2] This finding, as L&W

show, is consistent with an RBC model in which real shocks, not observed by the econometrician, affect the expected real interest rate first, and the money stock and output subsequently.[3] In L&W's model, although money is neutral, the money stock Granger-causes output spuriously when lagged interest rates are not included among the regressors. Moreover, L&W's model is not contradicted by empirical evidence that money-stock changes are correlated with subsequent changes in expected real interest rates [e.g., Mishkin (1981)] and Granger-cause the ex post real interest rate [e.g., Shiller (1980)].

The key testable implication of L&W's model is that neither nominal nor real variables Granger-cause the expected real interest rate. To test their model, they use quarterly postwar U.S. data to estimate a VAR involving measures of output, the rate of inflation, the money stock, and the nominal interest rate. Using the VAR forecast as a proxy for the expected rate of inflation, L&W derive the overidentifying restrictions imposed on the VAR by the null hypothesis that the expected real interest rate follows an uncaused (in Granger's sense) first order univariate autoregression. They fail to reject these overidentifying restrictions.

L&W show that standard macroeconomic models (including the IS-LM model with price rigidities as well as the Lucas-Barro imperfect-information model) imply that monetary aggregates Granger-cause the expected real interest rate. They conclude--and, in his critique of RBC theories, McCallum (1986) concedes--that the failure of the variables included in their VAR to Granger-cause the expected real interest rate poses a challenge to these standard models of the business cycle.

L&W's failure to reject the hypothesis that the expected real interest rate follows an uncaused autoregression seems to be sensitive to their estimation procedure; alternative empirical strategies lead to strong rejections of this hypothesis [Frydman (1986), Diba and Oh (1987)]. These rejections, however, pertain to the particular RBC model formulated by L&W; they do not constitute tests of any general class of RBC models or of the importance of the liquidity effect of money-stock changes.

The present paper extends L&W's empirical analysis in two directions. First, instead of L&W's univariate AR(1) specification, we consider a more general bivariate model--suggested by what Barro (1987, Chapter 5) refers to as "the basic market-clearing model"--in which the expected real interest rate depends on its own lagged values and on lagged output. We report tests of this bivariate model, against the alternative hypothesis that money-stock changes have a significant liquidity effect on the expected real interest rate. Second, motivated by the findings of King and Plosser, we contrast the relationship of the expected real interest rate with M2 to its relationship with the monetary base.

In what follows, Section 2 sets up our notation and presents the market-clearing models to be tested. Section 3 describes our estimation procedure and data. Section 4 discusses our empirical results. Section 5 contains a brief summary and some concluding remarks.

2. THE MODEL

Let i_t denote the one-period nominal interest rate at date t, q_{t+1} the rate of inflation from date t to date t+1, and r_{t+1} the ex post real interest rate satisfying

$$i_t = q_{t+1} + r_{t+1}. \tag{1}$$

Let ϕ_t denote the information set of market participants at date t, and define $q_t^e = E(q_{t+1} | \phi_t)$ as the conditional expectation of q_{t+1} given ϕ_t. The market participants' forecast error, v_{t+1}, of the rate of inflation from date t to date t+1 is

$$v_{t+1} = q_{t+1} - q_t^e. \tag{2}$$

Because market participants observe the nominal interest rate at date t, the expected real interest rate, $r_t^e = E(r_{t+1} | \phi_t)$, satisfies

$$r_t^e = i_t - q_t^e = r_{t+1} + v_{t+1}. \tag{3}$$

Note that our dating convention uses the time subscript to denote the date at which market participants observe the relevant random variable. Thus, we denote the expected real interest rate from date t to date t+1, (which is known to market participants at date t) by r_t^e, and the corresponding ex post real interest rate (observed at date t+1) by r_{t+1}.

The hypotheses we would like to test involve estimating equations of the form:

$$r_t^e = Z_{t-1}\gamma + \sum_{i=1}^{n} r_{t-i}^e \alpha_i + \eta_t, \tag{4}$$

where Z_{t-1} is a vector of lagged variables observed by the econometrician, γ and $\alpha = (\alpha_1, \alpha_2, \ldots, \alpha_n)$ are coefficient vectors, and η_t is a white noise disturbance.

L&W's null hypothesis, that the expected real interest rate follows a univariate autoregression, corresponds to including only a constant series in Z_{t-1}, in equation (4). (L&W also set n=1 in their empirical work, but their theoretical model does not restrict the order of the autoregression.) Thus, to test L&W's RBC model, we can include lagged values of some monetary aggregate among the regressors in Z_{t-1} and test their exclusion--that is, we can test the hypothesis that the monetary aggregate does not Granger-cause the expected real interest rate.

Although the finding that monetary aggregates Granger-cause the expected real interest rate would provide evidence against L&W's RBC model, it would not constitute compelling evidence against any general class of RBC models. Such a finding can arise spuriously in a world governed by some RBC model because in estimating equation (4) we may have left out some relevant real variables.[4] In general then, with a given list of real variables included in Z_{t-1}, the finding that we cannot exclude lagged monetary aggregates from equation (4) would not be very informative.

In contrast, as L&W (pp. 137-138) demonstrate, failure to detect Granger-causality running from monetary aggregates to the expected real interest rate cannot, except for an extreme coincidence, be spurious. Accordingly, such an empirical finding would constitute unambiguous evidence against the conventional macro models that imply Granger-causality running from money to the expected real interest rate. More generally, given any list of real variables included in Z_{t-1}, failure to reject the exclusion of lagged monetary aggregates from equation (4) would, *in principle* constitute evidence against conventional macro models. In practice, of course, even if these conventional models are true, it is likely that some specifications of the real variables included in Z_{t-1} would mask the explanatory power of monetary aggregates, in a given sample.

Accordingly, the choice of the real variables to be included in Z_{t-1} should not be based on an extensive specification search.

In sum, the choice of the real variables to be included in Z_{t-1} necessarily involves a compromise to minimize the effects of spurious causality and of the specification search. To guide our choice we consider the basic market-clearing model [as presented, for example, by Barro (1987), Chapter 5] in which aggregate supply is an increasing function of the expected real interest rate, aggregate demand is a decreasing function of the expected real interest rate, and the expected real interest rate adjusts every period to clear the commodity market.

A log-linear version of this model (ignoring the intercept terms) consists of the following three equations:

$$y_t^s = a\, r_t^e + u_t^s \,, \tag{5}$$

$$y_t^d = -\, b\, r_t^e + u_t^d \,, \tag{6}$$

$$y_t^s = y_t^d = y_t \,, \tag{7}$$

where y_t^s is the logarithm of aggregate supply, y_t^d is the logarithm of aggregate demand, a and b are nonnegative constants, and u_t^s and u_t^d are supply and demand disturbances.

Assuming that u_t^s and u_t^d are first-difference stationary, and that the Wold representations of their first differences are invertible, we can solve equations (5) to (7) to express the change in the expected real interest rate at date t in terms of its own lags and lagged growth rates of output. This basic market-clearing model then implies that, to avoid the possibility of spurious causality running from monetary aggregates to the expected real interest rate, we should at least allow for the effect of lagged output on the expected real interest rate.

The bivariate autoregressive models for the expected real interest rate suggested by this market-clearing model are of the form:

$$\Delta r_t^e = c + \sum_{i=1}^{n} \Delta r_{t-i}^e \alpha_i + \sum_{j=1}^{m} \Delta y_{t-j} \beta_j + \eta_t \; . \tag{8}$$

(Taken literally, equations (5) to (7) imply some restrictions on the coefficients and lag-lengths of the bivariate autoregressive model, which we ignore.)

The specification of equations (5) to (7) assumes that the expected real interest rate is first-difference stationary. Our empirical analysis will also consider a specification in which the expected real interest rate and the growth rate of output are stationary. This specification amounts to modifying equations (5) and (6) to:

$$y_t^s = y_{t-1} + a\, r_t^e + u_t^s \, , \tag{5'}$$

$$y_t^d = y_{t-1} - b\, r_t^e + u_t^d \, , \tag{6'}$$

Assuming now that u_t^s and u_t^d are stationary in levels and have invertible Wold representations, we can solve equations (5'), (6'), and (7) to obtain:

$$(8') \qquad r_t^e = c + \sum_{i=1}^{n} r_{t-i}^e \alpha_i + \sum_{j=1}^{m} \Delta y_{t-j} \beta_j + \eta_t .$$

The market clearing models presented above, which involve only two fundamental shocks (u_t^s and u_t^d), imply the exclusion of lagged monetary aggregates from the bivariate representation of the process generating the expected real interest rate--equations like (8) or (8'). These models, however, do not imply that the expected real interest rate is uncorrelated with lagged monetary aggregates, lagged rates of inflation, etc.; such correlations may arise spuriously because lagged expected real interest rates, which are not directly observable, are excluded from the list of regressors. Accordingly, empirical studies that have detected such correlations [e.g., Mishkin (1981), Barsky (1987)] do not constitute rejections of these market clearing models.

Nor would empirical evidence that a nominal variable Granger-causes the *ex post* real interest rate--as, for example, Shiller (1980) finds, in the case of the money stock--necessarily contradict the market clearing models presented above. As equation (9) below illustrates, replacing the unobservable lagged expected real interest rates on the right hand side of the regression equation by their ex post realizations would introduce the lagged inflation forecast errors, v_{t-j}, into the error term. The error term would thus become correlated with lagged rates of inflation and (potentially) with lagged money growth.

3. ESTIMATION PROCEDURE AND DATA

The cross-equation restrictions tested by L&W's maximum likelihood procedure become quite complicated for higher orders than their AR(1) specification of the process generating the expected real interest rate. We use an instrumental variables procedure, instead of maximum likelihood, to estimate variants of equation (4). Frydman (1986) criticizes L&W's estimation procedure on the grounds that the overidentifying restrictions they test are valid only if the information set of market participants is restricted to the variables included in L&W's VAR. The instrumental variables procedure we employ assumes that market participants observe all the variables we use as instruments but does not preclude the possibility that they have additional information.

Equations (3) and (4) imply that the ex post real interest rate satisfies:

$$r_{t+1} = Z_{t-1}\gamma + \sum_{j=0}^{n-1} r_{t-j}\alpha_{j+1} + \eta_t - v_{t+1} + \sum_{j=0}^{n-1} v_{t-j}\alpha_{j+1}. \tag{9}$$

Under the null hypotheses that we will test, the white noise disturbance η_t in equation (4) is independent of the inflation forecast error process. Therefore, the moving average error term of equation (9),

$$\eta_t - v_{t+1} + \sum_{j=0}^{n-1} v_{t-j}\alpha_{j+1} ,$$

is at most of order n [see, for example, Granger and Newbold (1986), pp. 28-29]. Accordingly, under the maintained hypothesis of rational expectations, variables observed at date t-n or earlier qualify as valid instruments for estimation of equation (9), using a simple (linear) version of the estimation procedure suggested by Hansen (1982) and by Cumby, Huizinga, and Obstfeld (1983).[5]

Our data are quarterly from 1954:1 to 1987:1. We report our results for the entire sample period as well as the subperiod 1954:1 to 1979:3, which ends before the October 1979 change in the Fed's operating procedure. Our measure of the nominal interest rate is the annualized continuously compounded yield on 13-week Treasury bills. We use the continuously compounded growth rate of seasonally adjusted CPI less shelter as our measure of the rate of inflation. Following L&W, to match the periods over which we measure inflation and the nominal interest rate, we have used the price index for the last month of the quarter (which is based on a sample taken approximately during the middle week of the month) and the weekly average of interest rates during the second or third week of the month. The measure of output is the seasonally adjusted Index of Industrial Production for the last month of the quarter. We use seasonally adjusted M2 and the Federal Reserve Bank of St. Louis monetary base, for the last month of the quarter, as measures of the money stock.

4. EMPIRICAL RESULTS

Table 1 reports stationarity tests for the relevant time series. The test statistics reported in the first two columns of the Table yield mixed results about the stationarity properties of the (ex post) real interest rate and the rate of inflation. Specifically, the tests reject the presence of a unit root in both time series when the null hypothesis is a random walk (i.e., when the parameter k is set equal to zero in the Dickey-Fuller regressions reported in Table 1), but not when the null hypotheses are generalized to ARIMA(4,1,0) processes (i.e., with the parameter k set equal to four). For the first differences of both the real interest rate and the rate of inflation and for the growth rates of output, M2, and the monetary base, the tests reject the presence of a unit root, for tests of size 0.10.

As several studies (e.g., Evans and Savin, 1984) have emphasized, the unit root test reported in Table 1 has low power against borderline stationary alternatives, even for a test of size 0.10. Accordingly, because we cannot infer that the time series of the real interest rate and the rate of inflation are nonstationary, we report our results both assuming that these time series are stationary in levels and assuming that they are stationary in first differences.

When the expected real interest rate in equation (4) is assumed stationary in levels, we estimate equation (9) for the ex post real interest rate in levels, with an MA(n) error term. The instruments for these regressions are a constant series, lags n-1 to n+2 of the nominal interest rate, and lags n to n+3 of the rate of inflation, output growth, and the growth rate of the monetary aggregate included among the regressors. When the expected real interest rate is assumed stationary in first differences, we estimate equation (9) in first-differenced form, with an MA(n+1) error term. The set of instruments is lagged an additional period and includes the first differences, instead of the levels, of nominal interest rates, and rates of inflation.

Tables 2 to 5 report tests of univariate autoregressive models for the expected real interest rate. For Tables 2 and 3, the estimated regressions are of the form:

$$(10) \qquad r_{t+1} = c + \sum_{i=0}^{n-1} r_{t-i}\alpha_{i+1} + \sum_{j=m}^{m+3} \Delta m_{t-j}\gamma_j + \text{error} ,$$

where Δm denotes the growth rate of M2 or of the monetary base. For Tables

4 and 5, the estimated regressions are of the form:

$$(10') \qquad \Delta r_{t+1} = c + \sum_{i=0}^{n-1} \Delta r_{t-i}\alpha_{i+1} + \sum_{j=m}^{m+3} \Delta m_{t-j}\gamma_j + \text{error} .$$

Note that each regression includes four lags of money growth, beginning with lag m, among the regressors.

The middle two columns of Tables 2 to 5 report the p-values (marginal significance levels) of the Wald statistic for testing the exclusion of the four lags of money growth from these regressions. For some specifications of lag-lengths, the p-values reported in these Tables are below the 0.05 level, rejecting the hypothesis that monetary aggregates do not Granger-cause the expected real interest rate. The rejections, however, are more frequent for M2 than for the monetary base; in fact, for the undifferenced specification, there is very little evidence that the monetary base Granger-causes the expected real interest rate. The p-values reported in Tables 2 to 5 also seem rather sensitive to our choice of sample period and stationarity assumption, and to minor changes in lag-lengths.

The Wald tests reported in the last two columns of Tables 2 to 5 in all cases fail to reject the hypothesis that the sum of the coefficients on lagged monetary aggregates equals zero. In particular, with m set equal to one, these tests suggest that a monetary expansion sustained for four quarters either has no liquidity effect on the expected real interest rate, or else has a liquidity effect that dissipates within four quarters. The tests for values of m greater than one constitute evidence against the hypothesis that a monetary expansion sustained for four quarters has a significant liquidity effect that operates with a long time lag (of up to nine quarters after the expansion began).

Overall, the tests reported in Tables 2 to 5 seem consistent, on three grounds, with the view that money-stock changes have not had a significant liquidity effect on expected real interest rates. First, the weakness of the evidence indicating that the monetary base Granger-causes the expected real interest rate is consistent with the view (e.g., King and Plosser, 1984) that the observed correlations of real variables with monetary aggregates reflect the endogenous response of inside money to changes in the real variables, rather than the structural effect of changes in outside money on the real variables. Second, the sensitivity of the causality tests to minor changes in specification suggests that the Granger-causality we detect may be spurious. Third, our failure to reject the hypothesis that the sum of coefficients on lagged money growth equals zero suggests that sustained changes in money growth have no liquidity effect.

Before turning to tests for the existence of a liquidity effect within shorter periods (than the four-quarter horizon assumed above), we consider the bivariate models of equations (8) and (8') for the expected real interest rate. Since monetary aggregates and the expected real interest rate fluctuate considerably over the business cycle, the exclusion of lagged output from the regressions reported in Tables 2 to 5 is a potential source of spurious causality running from money growth to the expected real interest rate. To investigate this possibility, we reestimated those regressions of Tables 2 to 5 for which the exclusion of lagged monetary aggregates was rejected at the 0.05 level, with four lags of output growth added to the regressors of equations (10) and (10').

Tables 6 and 7 report tests of these bivariate models. In most cases, the p-values of the Wald statistics for testing the exclusion of lagged money growth reported in Tables 6 and 7 are higher than the corresponding p-values in Tables 2 to 5. In Tables 6 and 7, rejections of the hypothesis that money growth does not Granger-cause the expected real interest rate are mainly confined to specifications involving M2 and our shorter sample period. For the monetary base, out of the 56 regression equations (with various specifications of lag-lengths, the choice of sample period, and stationarity assumption) that we started out with, we are left with only 3 specifications in Tables 6 and 7 for which we can reject the exclusion of lagged money growth at the 0.05 level.

Turning to the signs of the estimated coefficients, the last two columns of Tables 6 and 7 report the point estimates and asymptotic standard errors for the sum of the coefficients on lagged money growth in our bivariate models for the expected real interest rate. The estimates are in no case significantly different from zero and, contrary to what the existence of a liquidity effect would imply, have a positive sign in most cases (especially for the monetary base). These findings are consistent with the ones reported in Tables 2 to 5, which failed to detect a significant liquidity effect resulting from changes in money growth that were sustained for four quarters.

Have the gyrations of money growth had a significant liquidity effect within the four-quarter horizon? Table 8 reports the coefficients on money growth--lagged zero, one, two, or three quarters--in regressions using the bivariate null models, given by equations (8) and (8'), for the expected real interest rate. The only two coefficients that are significantly different from zero in Table 8 have positive signs, contrary to what a significant liquidity effect would imply. To save space, Table 8 does not report the coefficients on the other regressors in our estimated equations. Tables 9 and 10 report two of these equations in full.

Finally, Table 9 reports tests of the hypothesis that money-stock changes had a significant liquidity effect between the fourth quarter of 1979 and the second quarter of 1982, when the Fed was paying closer attention to money-stock fluctuations than during the preceding and subsequent subsample periods. The regressions reported in Table 11 are similar to those of Table 8 except for using a dummy variable to test for a negative coefficient on money growth during the 1979:Q4-1982:Q2 period. Once again, the only two coefficients that are significantly different from zero, in Table 9, have positive signs.

SUMMARY AND CONCLUDING REMARKS

This paper reported tests of the hypothesis that an increase in the money-stock has a liquidity effect that reduces the expected real interest rate. We found no evidence that in postwar quarterly U.S. data changes of the expected real returns on Treasury bills are negatively related to contemporaneous or past changes of M2 or of the monetary base. For some specifications of lag-lengths, our tests rejected the exclusion of lagged monetary aggregates from estimated equations for the expected real interest rate. Most of these rejections, however, occurred when we used M2 rather than the monetary base as our measure of the money stock. Moreover, the rejections were less frequent when, instead of restricting the generating process of the expected real interest rate to a univariate autoregression, we used a bivariate model that also allows for the effect of lagged output on the expected real interest rate.

Our findings are consistent with the view that policy-induced changes in money growth have not had a significant liquidity effect on the expected real interest rate, and with the more general view that the observed correlations of monetary aggregates with real variables largely reflect the effects of changes in the real variables on inside money. Obviously, our findings do not necessarily imply that monetary policy cannot affect the expected real interest rate. As McCallum (1986) emphasizes, money-stock changes may be a poor proxy for changes in the stance of monetary policy in the U.S. Notably, however, our tests failed to detect a significant liquidity effect even during the period when the Fed was directly targeting nonborrowed reserves and was paying more attention to its money growth targets.

TABLE 1

UNIT-ROOT TESTS

End of Sample: **Number of Lags (k):**	**1987** **0**	**1987** **4**	**1979** **0**	**1979** **4**
Real Interest Rate	-6.27	-2.20	-7.75	-2.63
Rate of Inflation	-5.27	-2.53	-5.70	-3.11
Output Growth	-8.02	-6.18	-7.13	-5.34
Growth Rate of Of M2	-6.65	-3.72	-4.95	-3.79
Growth Rate of the Monetary Base	-9.30	-3.27	-10.10	-4.21
Differenced Real Interest Rate	-20.21	-6.60	-17.62	-3.21
Differenced Rate of Inflation	-20.69	-4.90	-17.10	-4.78

Table reports values of the Dickey-Fuller τ_τ statistic. For each time series (x_t) indicated in the first column, regressions of the form:

$$\Delta x_t = \mu + t\gamma + x_{t-1}\rho + \sum_{j=1}^{k} \Delta x_{t-j}\beta_j + \text{error} ,$$

were estimated for k=0 and k=4. The statistic τ_τ was calculated like the t-ratio for ρ. The rejection region, tabulated for a sample of 100 observations in Fuller (1976, p 373), is the set of vales of τ_τ below -3.15 (-3.45) for a test of size 0.10 (0.05).

TABLE 2

TESTS OF UNIVARIATE AUTOREGRESSIVE
MODELS FOR THE EXPECTED REAL INTEREST RATE
1954:1 - 1987:1 SAMPLE PERIOD

		Wald Test for $\gamma = 0$ (p-values)		Wald Test for $\Sigma \gamma_j = 0$ (p-values)	
n	m	M2	Base	M2	Base
1	1	0.83	0.89	0.64	0.86
1	2	0.46	0.76	0.71	0.86
2	1	0.16	0.09	0.83	0.33
2	2	0.08	0.99	0.58	0.94
2	3	0.14	0.99	0.83	0.99
3	1	0.38	0.15	0.30	0.62
3	2	0.03	0.26	0.35	0.21
3	3	0.05	0.80	0.75	0.52
3	4	0.03	0.89	0.61	0.58
4	1	0.62	0.23	0.36	0.69
4	2	0.16	0.20	0.48	0.28
4	3	0.23	0.87	0.34	0.71
4	4	0.22	0.78	0.54	0.95
4	5	0.26	0.02	0.30	0.31

Table reports the p-values of the Wald statistics for testing the exclusion of four lags of money growth from autoregressive models for the expected real interest rate, and for testing the hypothesis that the sum of the coefficients on lagged money growth equals zero. The regressions are of the form:

$$r_{t+1} = c + \sum_{i=0}^{n-1} r_{t-i}\alpha_{i+1} + \sum_{j=m}^{m+3} \Delta m_{t-j}\gamma_j + \text{error} ,$$

where Δm denotes the growth rate of M2 or of the monetary base.

TABLE 3

TESTS OF UNIVARIATE AUTOREGRESSIVE MODELS FOR THE EXPECTED REAL INTEREST RATE 1954:1 - 1979:3 SAMPLE PERIOD

		Wald Test for $\gamma = 0$ (p-values)		Wald Test for $\Sigma \gamma_j = 0$ (p-values)	
n	m	M2	Base	M2	Base
1	1	0.07	0.70	0.76	0.33
1	2	0.21	0.91	0.73	0.33
2	1	0.18	0.18	0.81	0.08
2	2	0.00	0.30	0.73	0.28
2	3	0.00	0.51	0.85	0.28
3	1	0.17	0.57	0.46	0.91
3	2	0.00	0.24	0.87	0.42
3	3	0.00	0.71	0.80	0.86
3	4	0.00	0.48	0.82	0.48
4	1	0.38	0.29	0.09	0.34
4	2	0.00	0.86	0.54	0.75
4	3	0.00	0.87	0.81	0.72
4	4	0.00	0.03	0.75	0.43
4	5	0.04	0.00	0.11	0.33

Table reports the p-values of the Wald statistics for testing the exclusion of four lags of money growth from autoregressive models for the expected real interest rate, and for testing the hypothesis that the sum of the coefficients on lagged money growth equals zero. The regressions are of the form:

$$r_{t+1} = c + \sum_{i=0}^{n-1} r_{t-i}\alpha_{i+1} + \sum_{j=m}^{m+3} \Delta m_{t-j}\gamma_j + \text{error} ,$$

where Δm denotes the growth rate of M2 or of the monetary base.

TABLE 4

TESTS OF UNIVARIATE AUTOREGRESSIVE MODELS FOR FIRST DIFFERENCES OF THE EXPECTED REAL INTEREST RATE 1954:1 - 1987:1 SAMPLE PERIOD

		Wald Test for $\gamma = 0$ (p-values)		Wald Test for $\Sigma \gamma_j = 0$ (p-values)	
n	m	M2	Base	M2	Base
1	1	0.16	0.02	0.59	0.21
1	2	0.13	0.97	0.84	0.57
2	1	0.68	0.05	0.21	0.93
2	2	0.05	0.51	0.32	0.15
2	3	0.04	0.85	0.95	0.29
3	1	0.36	0.00	0.33	0.80
3	2	0.05	0.00	0.84	0.43
3	3	0.10	0.21	0.28	0.62
3	4	0.10	0.07	0.16	0.74
4	1	0.01	0.03	0.98	0.54
4	2	0.01	0.01	0.87	0.46
4	3	0.36	0.19	0.08	0.32
4	4	0.41	0.25	0.12	0.65
4	5	0.06	0.14	0.13	0.63

Table reports the p-values of the Wald statistics for testing the exclusion of four lags of money growth from autoregressive models for the first differences of the expected real interest rate, and for testing the hypothesis that the sum of the coefficients on lags of money growth equals zero. The regressions are of the form:

$$\Delta r_{t+1} = c + \sum_{i=0}^{n-1} \Delta r_{t-i}\alpha_{i+1} + \sum_{j=m}^{m+3} \Delta m_{t-j}\gamma_j + \text{error} ,$$

where Δm denotes the growth rate of M2 or of the monetary base.

TABLE 5

TESTS OF UNIVARIATE AUTOREGRESSIVE MODELS FOR
FIRST DIFFERENCES OF THE EXPECTED REAL INTEREST RATE
1954:1 - 1979:3 SAMPLE PERIOD

		Wald Test for $\gamma = 0$ (p-values)		Wald Test for $\Sigma \gamma_j = 0$ (p-values)	
n	m	M2	Base	M2	Base
1	1	0.05	0.47	0.60	0.48
1	2	0.00	0.56	0.95	0.99
2	1	0.08	0.24	0.25	0.60
2	2	0.00	0.11	0.34	0.10
2	3	0.00	0.66	0.42	0.33
3	1	0.11	0.35	0.10	0.29
3	2	0.00	0.56	0.32	0.17
3	3	0.00	0.60	0.54	0.10
3	4	0.00	0.00	0.75	0.60
4	1	0.38	0.31	0.65	0.97
4	2	0.02	0.70	0.37	0.89
4	3	0.20	0.80	0.82	0.92
4	4	0.16	0.00	0.99	0.56
4	5	0.42	0.00	0.43	0.99

Table reports the p-values of the Wald statistics for testing the exclusion of four lags of money growth from autoregressive models for the first differences of the expected real interest rate, and for testing the hypothesis that the sum of the coefficients on lags of money growth equals zero. The regressions are of the form:

$$\Delta r_{t+1} = c + \sum_{i=0}^{n-1} \Delta r_{t-i} \alpha_{i+1} + \sum_{j=m}^{m+3} \Delta m_{t-j} \gamma_j + \text{error} ,$$

where Δm denotes the growth rate of M2 or of the monetary base.

TABLE 6

TESTS OF BIVARIATE AUTOREGRESSIVE MODELS FOR THE EXPECTED REAL INTEREST RATE

End of Sample	Aggregate	n	m	Wald Test for $\gamma = 0$ (p-values)	Point Estimate of $\Sigma \gamma_j$	Standard Error for $\Sigma \gamma_j$
1987	M2	3	2	0.44	0.077	0.096
1987	M2	3	3	0.21	0.109	0.101
1987	M2	3	4	0.34	0.087	0.087
1987	Base	4	5	0.04	0.094	0.082
1979	M2	2	2	0.00	-0.085	0.086
1979	M2	2	3	0.00	-0.076	0.086
1979	M2	3	2	0.00	-0.002	0.100
1979	M2	3	3	0.01	0.003	0.111
1979	M2	3	4	0.00	-0.095	0.099
1979	M2	4	2	0.00	0.053	0.129
1979	M2	4	3	0.00	-0.032	0.121
1979	M2	4	4	0.00	0.078	0.098
1979	M2	4	5	0.54	-0.058	0.100
1979	Base	4	4	0.10	0.025	0.089
1979	Base	4	5	0.01	0.165	0.118

Table reports the p-values of the Wald statistics for testing the exclusion of four lags of money growth and the point estimates and asymptotic standard errors of the sum of the coefficients on lagged money growth from regressions of the form:

$$r_{t+1} = c + \sum_{i=0}^{n-1} r_{t-i}\alpha_{i+1} + \sum_{j=1}^{4} \Delta y_{t-j}\beta_j + \sum_{j=m}^{m+3} \Delta m_{t-j}\gamma_j + \text{error} ,$$

where Δy denotes the growth rate of output, and Δm the growth rate of M2 or of the monetary base.

TABLE 7

TESTS OF BIVARIATE AUTOREGRESSIVE MODELS FOR FIRST DIFFERENCES OF THE EXPECTED REAL INTEREST RATE

End of Sample	Aggregate	n	m	Wald Test for $\gamma = 0$ (p-values)	Point Estimate of $\Sigma \gamma_j$	Standard Error for $\Sigma \gamma_j$
1987	M2	2	2	0.65	0.071	0.095
1987	M2	2	3	0.32	0.131	0.115
1987	M2	3	2	0.47	-0.033	0.123
1987	M2	4	1	0.04	0.148	0.126
1987	M2	4	2	0.04	0.171	0.135
1987	M2	4	5	0.45	0.166	0.149
1987	Base	1	1	0.39	-0.056	0.066
1987	Base	2	1	0.41	0.011	0.083
1987	Base	3	1	0.17	0.032	0.101
1987	Base	3	2	0.02	0.053	0.103
1987	Base	3	4	0.08	0.054	0.091
1987	Base	4	1	0.67	-0.293	0.314
1987	Base	4	2	0.23	0.091	0.130
1979	M2	1	1	0.08	0.001	0.061
1979	M2	1	2	0.00	0.048	0.065
1979	M2	2	2	0.19	0.004	0.065
1979	M2	2	3	0.00	0.080	0.056
1979	M2	3	2	0.00	0.029	0.111
1979	M2	3	3	0.02	0.010	0.120
1979	M2	3	4	0.02	-0.049	0.115
1979	M2	4	2	0.79	0.090	0.171
1979	Base	3	4	0.75	0.070	0.108
1979	Base	4	4	0.12	0.044	0.111
1979	Base	4	5	0.25	0.025	0.114

Table reports the p-values of the Wald statistics for testing the exclusion of four lags of money growth and the point estimates and asymptotic standard errors of the sum of the coefficients on lagged money growth from regressions of the form:

$$\Delta r_{t+1} = c + \sum_{i=0}^{n-1} \Delta r_{t-i}\alpha_{i+1} + \sum_{j=1}^{4} \Delta y_{t-j}\beta_j + \sum_{j=m}^{m+3} \Delta m_{t-j}\gamma_j + \text{error} ,$$

where Δy denotes the growth rate of output, and Δm the growth rate of M2 or of the monetary base.

TABLE 8

TESTS OF LIQUIDITY EFFECTS WITH SHORTER LAGS

Aggregate: **Sample:**		**M2** **1987**	**Base** **1987**	**M2** **1979**	**Base** **1979**
x	**j**				
r	0	-0.055	0.130	0.129	0.082
		(0.106)	(0.094)	(0.082)	(0.084)
r	1	0.045	0.014	0.104	0.006
		(0.079)	(0.087)	(0.073)	(0.073)
r	2	0.061	0.117	0.108	0.077
		(0.078)	(0.082)	(0.074)	(0.085)
r	3	0.057	0.073	0.079	0.036
		(0.070)	(0.082)	(0.068)	(0.072)
Δr	0	-0.009	0.164	-0.191	0.028
		(0.112)	(0.099)	(0.234)	(0.153)
Δr	1	0.112	0.200	0.032	0.061
		(0.115)	(0.113)	(0.161)	(0.129)
Δr	2	0.262*	0.234*	0.083	0.003
		(0.115)	(0.105)	(0.183)	(0.125)
Δr	3	-0.018	0.155	-0.069	0.093
		(0.114)	(0.089)	(0.188)	(0.131)

Table reports the point estimates and asymptotic standard errors (in parentheses, below coefficients) of the parameter γ from regressions of the form:

$$x_{t+1} = c + \sum_{i=0}^{3} x_{t-i}\alpha_{i+1} + \sum_{j=1}^{4} \Delta y_{t-j}\beta_j + \Delta m_{t-k}\gamma + \text{error} ,$$

where the dependent variable x_{t+1} is equated either to r_{t+1} or to Δr_{t+1}, and the parameter k is set equal to 0, 1, 2, or 3.
* indicates significant t-ratio at the 0.05 level.

TABLE 9

TESTS OF CONTEMPORANEOUS LIQUIDITY EFFECTS FOR THE UNDIFFERENCED SPECIFICATION

Aggregate: Sample:	M2 1987	Base 1987	M2 1979	Base 1979
c	0.764	-0.087	-1.008	-0.360
	(0.759)	(0.450)	(0.655)	(0.432)
r_t	0.106	-0.264	0.688	0.101
	(0.330)	(0.408)	(0.408)	(0.600)
r_{t-1}	1.227*	0.628*	0.144	0.705
	(0.319)	(0.231)	(0.309)	(0.374)
r_{t-2}	0.432*	0.542*	-0.248	0.266
	(0.201)	(0.171)	(0.306)	(0.437)
r_{t-3}	-0.776	-0.053	0.517	-0.147
	(0.409)	(0.417)	(0.267)	(0.527)
Δy_{t-1}	-0.082	0.058	0.029	0.008
	(0.081)	(0.063)	(0.035)	(0.038)
Δy_{t-2}	-0.032	-0.110	-0.011	0.005
	(0.077)	(0.061)	(0.072)	(0.076)
Δy_{t-3}	0.101	0.018	0.040	-0.009
	(0.071)	(0.057)	(0.059)	(0.066)
Δy_{t-4}	-0.076	-0.043	0.007	-0.018
	(0.045)	(0.028)	(0.021)	(0.026)
Δm_t	-0.055	0.130	0.129	0.082
	(0.106)	(0.094)	(0.082)	(0.084)
J	8.66	11.34	9.29	10.53
(p-value)	(0.28)	(0.13)	(0.23)	(0.16)

Table reports regressions of the form:

$$r_{t+1} = c + \sum_{i=0}^{3} r_{t-i}\alpha_{i+1} + \sum_{j=1}^{4} \Delta y_{t-j}\beta_j + \Delta m_t \gamma + \text{error} .$$

Asymptotic standard errors are below coefficients in parentheses.
* indicates significant t-ratio at the 0.05 level.
The statistic J is the minimized value of the GMM objective function, used for a specification test [see, Hansen (1982)].
The p-values reported below this statistic are from the $\chi^2(7)$ distribution.

TABLE 10

TESTS OF CONTEMPORANEOUS LIQUIDITY EFFECTS FOR THE DIFFERENCED SPECIFICATION

Aggregate: Sample:	M2 1987	Base 1987	M2 1979	Base 1979
c	0.513	-0.379	1.029	-0.558
	(0.743)	(0.591)	(1.453)	(0.946)
Δr_t	-0.909*	-0.903*	-1.495*	-1.724*
	(0.192)	(0.288)	(0.302)	(0.345)
Δr_{t-1}	-0.460	-0.974	-1.604*	-2.245*
	(0.344)	(0.594)	(0.463)	(0.696)
Δr_{t-2}	-0.192	-1.020*	-0.982*	-1.873*
	(0.350)	(0.503)	(0.338)	(0.696)
Δr_{t-3}	-0.236	-0.817*	-0.878*	-1.327*
	(0.218)	(0.295)	(0.388)	(0.530)
Δy_{t-1}	-0.064	-0.024	0.011	-0.009
	(0.056)	(0.052)	(0.077)	(0.074)
Δy_{t-2}	0.024	0.073	0.076	-0.138*
	(0.055)	(0.070)	(0.046)	(0.067)
Δy_{t-3}	-0.004	-0.061	0.148	-0.128*
	(0.048)	(0.062)	(0.081)	(0.063)
Δy_{t-4}	-0.087*	-0.103*	-0.176*	-0.209*
	(0.029)	(0.029)	(0.075)	(0.072)
Δm_t	-0.009	0.164	-0.191	0.028
	(0.112)	(0.099)	(0.234)	(0.153)
J	10.24	9.30	2.87	3.65
(p-value)	(0.18)	(0.23)	(0.90)	(0.82)

Table reports regressions of the form:

$$\Delta r_{t+1} = c + \sum_{i=0}^{3} \Delta r_{t-i}\alpha_{i+1} + \sum_{j=1}^{4} \Delta y_{t-j}\beta_j + \Delta m_t \gamma + \text{error} .$$

Asymptotic standard errors are below coefficients in parentheses.
* indicates significant t-ratio at the 0.05 level.
The statistic J is the minimized value of the GMM objective function, used for a specification test [see, Hansen (1982)].
The p-values reported below this statistic are from the $\chi^2(7)$ distribution.

TABLE 11

TESTS FOR EXISTENCE OF LIQUIDITY EFFECTS BETWEEN 1979:Q4 AND 1982:Q2

Aggregate:		M2		Base	
		γ_1	γ_2	γ_1	γ_2
x	j				
r	0	-0.132 (0.142)	0.142 (0.159)	0.029 (0.100)	0.202 (0.259)
r	1	-0.216 (0.139)	0.267 (0.248)	0.075 (0.123)	0.185 (0.297)
r	2	-0.187 (0.129)	0.253 (0.225)	-0.054 (0.137)	0.471 (0.360)
r	3	0.089 (0.176)	0.022 (0.216)	-0.018 (0.139)	0.828 (0.521)
Δr	0	0.083 (0.199)	0.157 (0.299)	-0.002 (0.268)	1.255 (0.737)
Δr	1	-0.203 (0.138)	0.466 (0.246)	0.040 (0.148)	0.762 (0.586)
Δr	2	-0.266 (0.198)	0.844* (0.342)	0.030 (0.155)	0.980* (0.444)
Δr	3	0.407 (0.342)	-0.389 (0.540)	0.144 (0.132)	0.606 (0.472)

Table reports the point estimates and asymptotic standard errors (in parentheses, below coefficients) of the parameters γ_1 and γ_2 from regressions of the form:

$$x_{t+1} = c + \sum_{i=0}^{3} x_{t-i}\alpha_{i+1} + \sum_{j=1}^{4} \Delta y_{t-j}\beta_j + \Delta m_{t-k}\gamma_1 + (D_{t-k}\Delta m_{t-k})\gamma_2 + \text{error} ,$$

where the dependent variable x_{t+1} is equated either to r_{t+1} or to Δr_{t+1}, and the parameter k is set equal to 0, 1, 2, or 3; D is a dummy variable that equals one between the fourth quarter of 1979 and the second quarter of 1982 and equals zero for the other observations.
* indicates significant t-ratio at the 0.05 level.

NOTES

1. There are two interpretations of the central claim of RBC theories. One interpretation is that real variables do not structurally depend on nominal variables and, therefore, variations in the time paths of monetary aggregates can never affect the time paths of real variables. We prefer the second interpretation, suggested by Eichenbaum and Singleton (1986), that the nature of monetary institutions and policy in the sample period under study was such that a model ignoring any structural dependence of real variables on nominal variables provides an accurate characterization of the data. See also McCallum (1986) for a critical discussion of RBC models and their interpretations.

2. This finding appears to be sensitive to assumptions about the stationarity properties of the relevant time series--see Stock and Watson (1987) and the studies they cite.

3. Boschen and Mills (1987) empirically substantiate the claim that the observed money-output correlations arise spuriously because researchers' models omit information, available to agents, about future values of real variables.

4. See King (1986) for an explicit example of how failure to include a relevant real variable among the regressors of the output equation can lead to spurious Granger-causality running from money to output.

5. For our estimation, we used a GMM program written by Greg Leonard and Dave Runkle, with the Newey and West (1987) covariance matrix.

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