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

Behzad T. Diba
Federal Reserve Bank of Philadelphia

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. Our findings, however, are consistent with a more general bivariate model--suggested by what Barro (1987, Chapter 5) refers to as "the basic market-clearing model"--in which the real rate depends on its own lagged values and on lagged output. Taking this bivariate model as our null hypothesis, we find no evidence that money-stock changes have a significant liquidity effect on the expected real interest rate.

Address correspondence to:

Behzad Diba
Research Department
Federal Reserve Bank of Philadelphia
Ten Independence Mall
Philadelphia, PA 19106

INTRODUCTION

This paper examines the relationship, in postwar U.S. data, between changes in monetary aggregates and changes in the expected real interest rate--henceforth, referred to as the real rate. 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 real 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] Since RBC models suggest (implicitly or explicitly) that money-stock changes do not have a significant liquidity effect, our empirical analysis is relevant for assessing the importance of the RBC challenge to conventional macroeconomic models.

We consider two specifications of the process generating the real rate. In the first specification, under the null hypothesis, the real rate follows a univariate autoregression. In the second specification, under the null hypothesis, the real 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 real 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 real 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 real-rate changes [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 real rate. To test their model, they use quarterly postwar US 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 imperfect-information model) imply that monetary aggregates Granger-cause the real 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 real rate poses a challenge to these standard models of the business cycle. L&W's empirical analysis, however, jointly tests the exclusion of lagged values of output, the rate of inflation, and the money stock from their AR(1) model for the real rate. The present paper reports some evidence on the bivariate relations of monetary aggregates with the real rate, and also considers higher order autoregressive specifications.

(The cross-equation restrictions tested by L&W's maximum likelihood procedure become very complicated for higher than first-order processes; we use an instrumental variables procedure, instead of maximum likelihood.)

L&W's failure to reject the hypothesis that the real 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 findings of our earlier paper supported a more general bivariate model--suggested by what Barro (1987, Chapter 5) refers to as "the basic market-clearing model"--in which the real rate depends on its own lagged values and on lagged output. The present paper reports tests of this bivariate model, against the alternative hypothesis that money-stock changes have a significant liquidity effect on the real rate.

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 real rate, $r_t^e = E(r_{t+1} \mid \phi_t)$, 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 estimation of 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 real 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 real rate.

Although the finding that monetary aggregates Granger-cause the real 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 real 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 real 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, <u>in principle</u> constitute evidence against conventional macro models. In practice, however, even if these conventional models are true, a specification search would quite likely yield a list of real variables whose inclusion in Z_{t-1} would mask the explanatory power of monetary aggregates, in a given sample.

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 real rate, aggregate demand is a decreasing function of the real rate, and the real 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 real 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 real rate, we should at least allow for the effect of lagged output on the real rate.

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

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

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

ESTIMATION PROCEDURE

If lagged values of the real rate did not appear on the right hand side of equation (4), we could easily test the exclusion of monetary aggregates from Z_{t-1}, under the maintained hypothesis of Rational Expectations, using the procedure of Mishkin (1981). Mishkin's procedure exploits the fact that, according to equation (3) above, the measurement error resulting from replacing the dependent variable r_t^e by r_{t+1} is not correlated with the lagged variables included in Z_{t-1}--and, therefore, the OLS regression of r_{t+1} on Z_{t-1} yields valid test statistics. (Note that, because r_{t-i}^e in equation (4) can be correlated with lagged nominal variables, the finding that such variables are correlated with r_t^e--as reported, for example, by Mishkin (1981) and Barsky (1987)--does not contradict the exclusion of nominal variables from this equation.) We modify Mishkin's procedure below to accommodate the inclusion of lagged real rates on the right hand side of equation (4).

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

$$(8) \qquad r_{t+1} = Z_{t-1}\gamma + \sum_{j=0}^{n-1} r_{t-j}\alpha_{j+1} + \eta_t - v_{t+1} + \sum_{j=0}^{n-1} v_{t-j}\alpha_{j+1}.$$

(Note that the lagged values of the inflation forecast error, v_{t-j}, which enter equation (8), may be correlated with lagged nominal variables. Accordingly, the finding that a nominal variable Granger-causes the ex post real interest rate--as, for example, Shiller (1980) finds, using money supply data--does not necessarily imply that this nominal variable Granger-causes the real rate.)

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 (8),

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

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+1 or earlier qualify as valid instruments for estimation of equation (8), using a simple (linear) version of the estimation procedure suggested by Hansen (1982) and by Cumby, Huizinga, and Obstfeld (1983).[5]

EMPIRICAL RESULTS

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. 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.

Dickey-Fuller tests failed to reject the hypothesis that the time series of the ex post real interest rate in our sample has a unit root. Accordingly, to avoid the statistical problems, reviewed and illustrated by Stock and Watson (1987), that arise in causality tests based on integrated time series, we estimated variants of equation (8) in first differences. The moving average error term in the differenced regression equations is of order n+1, where n is the order of the autoregressive polynomial specified for the real rate. The instruments are a constant series and four lags each of first differences of nominal interest rates, first differences of rates of inflation, growth rates of output, and growth rates of the monetary aggregate included among the regressors.

Tables 1 and 2 report tests of univariate autoregressive specifications of the real-rate process. The estimated 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. Each regression includes four lags of money growth, beginning with lag m, among the regressors.

The middle two columns of Tables 1 and 2 report the p-values (marginal significance levels) of the Wald statistic for testing the exclusion of the four lags of money growth. The p-values are below the 0.05 level for several specifications of lag-lengths. Although these p-values indicate several rejections of the hypothesis that monetary aggregates do not Granger-cause the real rate, they are quite sensitive to minor changes in lag-lengths--suggesting that the apparent Granger-causality may be spurious.

The Wald tests reported in the last two columns of Tables 1 and 2 in all cases fail to reject the hypothesis that the sum of the coefficients on lagged monetary aggregates equals zero. These results suggest, in particular (with m set equal to one), that a monetary expansion sustained for four quarters either has no liquidity effect on the real rate, or else has a liquidity effect that dissipates within four quarters. Our failure to reject the hypothesis that the sum of coefficients on lagged money growth equals zero strengthens the possibility that the apparent Granger-causality running from the monetary aggregates to the real rate, rather than reflecting a structural relation, is spurious.

Since monetary aggregates and the real rate fluctuate considerably over the business cycle, the exclusion of lagged output from the regressions reported in Tables 1 and 2 is a potential source of spurious causality running from monetary aggregates to the real rate. To investigate this possibility, we reestimated the regressions of Tables 1 and 2 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.

The estimated regressions are of the form:

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

Table 3 reports the results.

Overall the p-values of the Wald statistics for testing the exclusion of the lagged monetary aggregates, reported in Table 3, are considerably higher than the corresponding p-values in Tables 1 and 2. In fact, out of the 56 regression equations (with various specifications of lag-lengths, the choice of the monetary aggregate, and sample period) that we started out with, we are left with only 8 equations in Table 3 for which we can reject the exclusion of lagged monetary aggregates at the 0.05 level. Interestingly, all but one of these rejections occur when we take M2, rather than the monetary base, as the monetary aggregate. (In our earlier paper, we failed to reject the exclusion of lagged values of M1 from univariate and bivariate autoregressive models for the real rate, in several cases.)

King and Plosser (1984) observe that much of the correlation of output with money is with inside money. Bernanke (1986) and Boschen Mills (1987) also find that measures of outside money do not predict output. Our failure to reject the exclusion of lags of the monetary base from the bivariate real rate model is consistent with these findings. As King and Plosser point out, the failure of measures of outside money to predict the fluctuations of real variables strengthens the claim that the correlations of real variables with measures of inside money do not reflect any structural dependence of these real variables on monetary aggregates.

Turning to the signs of the estimated coefficients, the last two columns of Table 3 report the point estimates and asymptotic standard errors for the sum of the coefficients on lagged monetary aggregates in the real rate equations. The point estimates are in most cases positive and in no case significantly different from zero. Table 4 reports tests of the hypothesis that money-stock changes have a contemporaneous (i.e., within the quarter) liquidity effect on the real rate. Again, we fail to find significant negative coefficients for the monetary aggregates.

We also experimented with several variations on the regressions reported in Table 4, without detecting any evidence of a significant liquidity effect. Our experiments included using a single lag of money growth, instead of contemporaneous money growth, among the regressors; using next quarter's money growth figures to test for a liquidity effect of anticipated changes in the money stock; and testing for liquidity effects that were only significant between 1979:3 and 1982:2.

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 US data changes of the expected real returns on Treasury bills are negatively related to past, contemporaneous, or expected future 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 real rate. Most of these rejections, however, occurred when we restricted the generating process of the real rate to a univariate autoregression. Using a bivariate model, which allows for the effect of lagged output on the real

rate, we failed to reject the exclusion of lagged monetary aggregates, especially the monetary base, in most cases.

Our findings obviously do not necessarily imply that monetary policy has no effect on the real rate. As McCallum (1986) emphasizes, money-stock changes may be a poor proxy for changes in the stance of monetary policy in the US. Our findings, however, lead us to conclude that a simple market-clearing model that ignores any structural dependence of the real rate on monetary aggregates provides an accurate characterization of the postwar US. data.

TABLE 1

TESTS OF UNIVARIATE AUTOREGRESSIVE MODELS FOR THE REAL 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	0.34	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	0.35	0.19	0.08	0.32
4	4	0.40	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 a monetary aggregate from autoregressive models for the real rate, and for testing the hypothesis that the sum of the coefficients on lags of the monetary aggregate 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 2

TESTS OF UNIVARIATE AUTOREGRESSIVE MODELS FOR THE REAL 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	0.35	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 a monetary aggregate from autoregressive models for the real rate, and for testing the hypothesis that the sum of the coefficients on lags of the monetary aggregate 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 3

TESTS OF BIVARIATE AUTOREGRESSIVE MODELS FOR THE REAL 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	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	-0.049	0.115
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	0.83	0.025	0.114

Table reports the p-values of the Wald statistics for testing the exclusion of four lags of monetary aggregates and the point estimates and asymptotic standard errors of the sum of the coefficients on lagged monetary aggregates 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 4

TESTS OF CONTEMPORANEOUS LIQUIDITY EFFECTS

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.

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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