Monetary Stabilization Policy: Evidence from Money Demand Forecasts

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ESTIMATED money demand relationships are a key ingredient in the formulation of monetary policy. Recently, some analysts have argued that financial innovations have rendered the money demand relationship unstable. Because of this, intermediate monetary targeting — a policy that is based on the predictability of money demand — has been viewed as a dubious policy procedure to follow.[1]

In this article, we investigate the stability of two commonly estimated money demand functions. Specifically, we examine whether there has been a statistically significant change in the estimated relationships between those found for the period 1960–79 and those for the period 1960–84.

We also examine the forecasting ability of the two models. To do this, the equations are estimated over the 1960–79 sample and are used to generate quarterly forecasts for the 1980–84 period. By observing the forecast errors in conjunction with the stability test results, we can better assess the validity of the recent arguments against monetary targeting.

ESTIMATING SHORT-RUN MONEY DEMAND

An extensive literature exists on the appropriate form of the short-run money demand function.[2] To investigate the issue of money demand stability, we have chosen two common specifications. These are

$$(1)\quad \ln(M/P)_t = \alpha_0 + \alpha_1 \ln y_t + \alpha_2 \ln R_t + \alpha_3 \ln(M/P)_{t-1} + \varepsilon_t ,$$

and

$$(2)\quad \ln(M/P)_t = \alpha_0' + \alpha_1' \ln y_t + \alpha_2' \ln R_t + \alpha_3' \ln(M_{t-1}/P_t) + \varepsilon_t' ,$$

where M = nominal M1,

P = the price level measured by the GNP deflator (1972 = 100),

y = a scale variable represented by real GNP ($1972),

R = a nominal market rate of interest, measured by the commercial paper rate, and

ε = a random error term.

Equations 1 and 2 are the so-called real and nominal adjustment specifications, respectively. These two equations differ in that the real adjustment specification assumes that individuals adjust their actual real money balances to their desired level. The nominal adjustment specification, on the other hand, assumes that individuals adjust their nominal money balances to their desired level. Although the two equations appear equivalent except for the adjustment variable, the dependent variable in equation 2 actually is the logarithm of nominal money.[3] Because there is no consensus on which of these two specifications is correct, both are used.

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[1]For example, Higgins and Faust (1981), p. 17, note that financial innovations create an atmosphere in which "it may be necessary to reevaluate the desirability of using monetary targets to achieve ultimate policy objectives." In this vein, Davis (1981), p. 24, suggests that "perhaps more subtle and pervasive questions about the desirability of pursuing rigorously monetary growth targets are raised by questions about the stability of the 'demand for money'."

[2]For a survey of the literature, see Laidler (1977).

[3]Thornton (1985) discusses this point and provides a more complete discussion of the derivation of the two money demand specifications.

Table 1
Money Demand Estimates

	I/1960–IV/1979		I/1960–IV/1984	
Variables	**Real**	**Nominal**	**Real**	**Nominal**
$\Delta\ln y_t$	0.168	0.151	0.239	0.187
	(3.10)	(3.20)	(3.32)	(2.82)
$\Delta\ln R_t$	−0.015	−0.015	−0.013	−0.014
	(2.60)	(3.06)	(1.77)	(2.22)
$\Delta\ln(M/P)_{t-1}$	0.533	—	0.309	—
	(5.37)		(3.04)	
$\Delta\ln(M_{t-1}/P_t)$	—	0.679	—	0.526
		(7.70)		(5.27)
$\bar{R}^2$	0.432	0.559	0.241	0.353
SE	2.219	1.956	3.318	3.063
Dh	−0.30	−0.70	N.A.[1]	N.A.[1]

NOTE: $\bar{R}^2$ is the coefficient of determination adjusted for degrees of freedom; SE is the regression standard error; and Dh is the Durbin h-statistic for autocorrelation.

[1]The h-statistic could not be calculated. The relevant Durbin-Watson statistics are 2.15 for the real adjustment specification and 2.43 for the nominal.

A number of studies have found that the estimated coefficients in equations 1 and 2 are statistically unstable when estimated across the mid-1970s. This instability has been ascribed to a variety of causes, including large changes in the price level, a wealth loss due to OPEC oil shocks, changes in financial management techniques and more.[4] It has been shown, however, that this instability of the *level* version is reduced greatly when the equation is estimated in first-difference form, at least up to 1980.[5] The general use of differencing has been suggested by Granger and Newbold (1974) and Plosser and Schwert (1978) to achieve stationarity and to reduce the possibility of a spurious regression result. On this point, a recent study by Layson and Seaks (1984) presents evidence indicating that the first-difference version of the money demand specification is statistically preferable to its level form.

Based on these findings, therefore, we use the first-difference versions of equations 1 and 2 in this study. Thus, the equations estimated and analyzed in this article are:

(3) $\Delta\ln(M/P)_t = \beta_1\,\Delta\ln y_t + \beta_2\,\Delta\ln R_t + \beta_3\,\Delta\ln(M/P)_{t-1} + \phi_t$

(4) $\Delta\ln(M/P)_t = \beta_1'\,\Delta\ln y_t + \beta_2'\,\Delta\ln R_t + \beta_3'\,\Delta\ln(M_{t-1}/P_t) + \phi_t'$.[6]

Estimation Results

Equations 3 and 4 are estimated for two sample periods: I/1960–IV/1979 and I/1960–IV/1984. The split at 1980 is used to determine the stability of the model during the past five years, a period of substantial financial market change. The question addressed is whether the results from the earlier period are statistically different from those of the latter.[7]

The results of estimating equations 3 and 4 are presented in table 1. Looking at the 1960–79 results, the estimated short-run income and interest rate elasticities are similar across specifications. The estimated coefficient on lagged money balances in equation 3 is

[4]A survey of the relevant literature is presented in Judd and Scadding (1982) and Roley (1985).

[5]This finding initially was reported in Hafer and Hein (1982).

[6]Note that the constant term does not appear in the first-difference equations. This is due to the algebraic manipulation of the level equation to generate the first-difference model. It should be noted, however, that incorporating a constant term into the first-difference equation represents a time trend variable from the level equation. Estimating the first-difference equations with the constant term found it to be insignificantly different from zero.

[7]The estimation properties of the (real) equation for the 1960–79 period are presented in Hafer and Hein (1982).

Table 2
Stability Test Results

	Test statistics for:	
Specification	Coefficient stability	Homoskedasticity
Real adjustment	4.18	37.72
Nominal adjustment	2.68	45.86

NOTE: The test statistic is distributed as a χ^2 with three degrees of freedom. The 5 percent critical value for these tests is 7.82.

0.533, implying an adjustment speed of 47 percent per quarter. For the nominal adjustment model, the estimated coefficient is 0.679, which yields an adjustment coefficient of 32 percent per quarter.[8]

The differences in the estimated adjustment speeds produce different long-run income and interest rate elasticities. The long-run income elasticity from the real specification is 0.36; from the nominal model it is 0.47. Each estimate is slightly less than values reported in previous studies.[9] The differences are especially noticeable in the long-run interest elasticities: the long-run interest elasticity from the real model is -0.032, while that from the nominal model is -0.047.

When the equations are estimated for the full 1960–84 period, some notable changes occur in the coefficient estimates. In each equation, the estimated short-run income elasticity increases in value, while the estimated coefficient on the lag term declines. Interestingly, the estimated short-run interest elasticities are little changed by the increased sample data.

ARE THE MODELS STABLE?

Comparing the two equations across the two sample periods indicates a substantial increase in the regression standard error. This increase suggests that the equations may not be statistically stable; that is, the estimated statistical relationship may have changed significantly across the sample.

To examine this issue, each equation was tested for stability of the estimated coefficients and for stability of the error structure. This dichotomy is important, because tests for coefficient stability in the presence of heteroskedasticity can be misleading.[10] Consequently, two test statistics are reported for each specification. One tests for coefficient stability, allowing the variance to change; the other tests for constant variance, with the coefficients allowed to change. The relevant test statistics are reported in table 2.

The results for each specification indicate that we cannot reject the hypothesis that the estimated coefficients are statistically constant across the 1979 break. Each of the calculated chi-squared statistics is well below the 5 percent critical value. The results of testing for homoskedastic errors indicates, however, that we can easily reject the hypothesis of constant variance over the two periods. This outcome suggests that the exogenous influences affecting the error term have changed between the two periods.

The stability evidence indicates that, contrary to some recent findings, the estimated coefficients of the real and nominal adjustment models of money demand have not changed significantly during the past five years when compared with those from the 1960–79 sample.[11] The question to which we now turn is, why has the variance of the estimate relationships changed? To do this, we examine the models' forecast errors for the post-1979 period.

[8]For a critical interpretation of such results, see Goodfriend (1985).

[9]For a comparison with previous results, see Judd and Scadding (1982).

[10]See Thornton for a related discussion on this point and the likelihood ratio tests used here.

[11]Thornton recently has reported that there is some evidence of instability for the real and nominal adjustment models. It should be noted, however, that his tests are based on the level specification. Also, his estimated equations include the passbook savings rate as an additional explanatory variable. Even with these differences, however, his parameter stability test results for the nominal adjustment model without the passbook rate estimated over the 1962–84 period indicate that stability cannot be rejected at the 5 percent level of significance.

Gordon (1984), on the other hand, reports the first-difference model to be unstable, based on a simple F-test. The reported test results examine the overall fit of the model. It is possible to test the stability of each coefficient over the 1979 break through the use of dummy variables. Let D1 = 1 for I/1960–IV/1979 and zero elsewhere, and D2 = 1 for I/1980–IV/1984 and zero elsewhere. Forming interaction terms with the right-hand-side variables, we may test the difference between coefficients estimated for each subsample. Testing the null hypothesis of coefficient equality, the absolute value of the calculated t-statistics for the real adjustment model variables are: income — 1.50; commercial paper rate — 0.14; and lagged term — 2.22. The t-statistics from the nominal adjustment model are: income — 0.92; commercial paper rate — 0.02; and lagged term — 1.68.

This evidence suggests that the lagged term in the real-adjustment model has changed. In contrast, none of the coefficients in the nominal adjustment model have changed, providing some basis for the preference of this version. It should be noted, however, that this test procedure does not account for changes in the error variance.

FORECASTING MONEY DEMAND: 1980–84

A computationally convenient procedure to examine the post-1979 forecast results for each specification is suggested by Dufour (1980, 1982). This technique uses separate (0, 1) dummy variables entered for each individual observation beyond a selected break point. In the present example, a dummy variable D1 was entered as 1.0 for I/1980 and zero elsewhere; D2 was entered as 1.0 for II/1980 and zero elsewhere; and so on through IV/1984. When added to equations 3 and 4 and estimated over the full 1960–84 sample period, the estimated coefficients on the dummy variables represent post-sample static forecast errors. Moreover, the t-statistic for each dummy variable provides information about which forecast error significantly departs from the 1960–79 regression model. Thus, by examining the estimated coefficients on the dummy variables for the I/1980–IV/1984 period, we can determine the magnitude of the forecast error and determine the sign pattern of the errors.[12]

On this last point, we especially are interested in whether there are transitory errors — errors that alternate in sign — or whether the errors are generally one-sided. Significant transitory errors suggest that the model is subject to random shocks that are larger during the forecast period than the average squared error experienced during the estimation sample. A forecast error pattern that has consistently significant, one-sided errors, however, suggests that the relationship embodied in the estimated model has changed from that in the estimation period.

To statistically investigate the nature of the forecast errors, it is informative to test whether the sum value of the forecast errors is statistically different from zero. If this hypothesis is rejected, the evidence would indicate that the forecast errors are offsetting in sign and magnitude.

The estimated dummy variable coefficients and t-statistics for both the real and nominal adjustment models are reported in table 3.[13] The evidence for the real adjustment model indicates that there have been several statistically significant departures from the regression model during the past five years. The first two are in II/1980 and III/1980, when special credit controls were initiated by the Carter administration. These errors are by far the largest; more important, however, is the fact that they are offsetting in sign and magnitude. This result is consistent with the notion that the credit control program had only a temporary effect on the money demand forecast errors.[14]

The remaining significant forecast errors are found mostly in 1981 and 1982. The errors in 1981 occur during the first three quarters, a period associated with the nationwide legalization of NOW accounts. More important is the result that the errors alternate in sign and are of approximately equal magnitudes. This also holds true for the errors found in the first two quarters of 1982. The forecast errors found in 1981 and 1982 corroborate previous findings about the increased variability of velocity growth during this period. The evidence here suggests that these errors were transitory.[15]

The forecast errors from the nominal adjustment specification follow a pattern similar to those from the real adjustment model. The sign pattern generally holds between the two error series, and the significant errors are located in the same periods, except for II/1983. In that quarter, the nominal adjustment model's forecast error (2.357), unlike that of the real adjustment model, is not statistically significant at the 5 percent level.

The F-statistic reported below each forecast series tests the hypothesis that the sum of the forecast errors is zero. The reported F-statistics are quite low and, as indicated by the significance levels reported in parentheses, do not permit rejection of the null hypothesis at any reasonable level of significance. Thus, finding that the sum of the money demand forecast errors from the real and nominal adjustment specifications are not different from zero corroborates the previous

[12]This procedure also is used by Hafer (1985) to investigate the stability of money demand during the 1920–39 period.

[13]The estimated coefficients on the other variables are not reported since, by construction of the test, they are identical to the 1960–79 estimates found in table 1.

[14]For relevant discussions of this finding, see Judd and Scadding (1981) and Hein (1982). Indeed, our evidence suggests that large fluctuations in the nominal money stock, such as those associated with the credit control period, may explain observed errors in the money demand model. Such a theory is suggested by Carr and Darby (1981).

[15]See Tatom (1983), Judd and Motley (1984), Hafer (1984a, b) and Gordon (1984) for discussions of this period. Interestingly, the signs of the forecast errors during this period do not conform with those predicted by some financial innovation arguments.

For a discussion of how financial innovations have influenced money demand estimates in Japan, see Suzuki (1984).

Table 3
Estimates of Post-1979 Forecast Errors

	Real		Nominal	
Variables	Estimates	t-statistic	Estimates	t-statistic
I/1980	−1.532	−0.68	0.357	0.18
II	−12.314	−5.31*	−11.211	−5.49*
III	13.613	5.25*	14.358	6.48*
IV	−1.520	−0.62	−0.923	−0.43
I/1981	−8.922	−3.92*	−8.991	−4.49*
II	6.657	2.84*	4.924	2.48*
III	−8.013	−3.60*	−6.098	−3.10*
IV	0.055	0.02	−0.185	−0.09
I/1982	6.908	3.07*	4.891	2.48*
II	−4.678	−2.07*	−4.684	−2.36*
III	2.915	1.29	1.857	0.93
IV	9.689	4.16*	9.251	4.51*
I/1983	−1.358	−0.53	−2.072	−0.95
II	4.696	2.04*	2.357	1.14
III	1.482	0.63	0.558	0.27
IV	−2.883	−1.24	−2.983	−1.48
I/1984	−0.617	−0.27	−0.704	−0.35
II	1.751	0.78	0.908	0.46
III	−1.112	−0.50	−1.137	−0.58
IV	−1.658	−0.73	−2.487	−1.23
F	0.094		0.048	
	(0.76)		(0.83)	

*Denotes significance at 5 percent level.

result of stable coefficient estimates.[16]

POLICY IMPLICATIONS

The empirical evidence suggests that the relationship between the growth of money balances and its economic determinants is more stable than some have argued. Although there is evidence of large post-1979 forecast errors, these errors are transitory and the sum of the forecast errors is not statistically different from zero. This evidence suggests that monetary policies relying on quarter-to-quarter forecasts of money demand growth may not fare well because of the random, unpredictable component inherent in the estimated relationship. It also suggests, however, that the secular relationships embodied in the money demand function may be exploited successfully by emphasizing long-run money growth and GNP growth objectives.[17]

[16]In determining the significance of the individual dummy variables, it should be noted that they are being compared with the regression model estimated through IV/1979. In this way, the large forecast errors do not influence the two-standard-error interval used to locate the significant forecast errors.

It may be argued that the evidence on the sum of the forecast errors holds only over the long period forecasted and that the use of selected subperiods would show the average error not to be zero. This argument misses the point: because there always are short-term forecast errors, some of which can be "large," policies that attempt to exploit such quarterly deviations from forecasts may fail to achieve desired longer-term monetary policy goals. Because the longer-term results indicate that the errors average to zero over time, a longer-view policy may better achieve desired longer-term goals, such as price stability and income growth.

[17]This conclusion also is reached by Hein and Veugelers (1983) in their study of velocity. In that article, the predictability of the quarter-to-quarter growth of M1 velocity was examined. Their evidence indicated that, on a quarterly basis, velocity growth fluctuated randomly about a fixed mean. As the forecast horizon was extended, the accuracy of the forecasts improved. Thus, in the context of a simple quantity theory model, given some desired growth for nominal income, determining the correct growth for money based on a forecast of velocity (or money demand) will be successful only for horizons longer than one or two quarters.

CONCLUSION

In this article, we have presented evidence indicating that the estimated coefficients from two common short-run money demand specifications are statistically stable across the 1960–84 period. Using IV/1979 as the hypothesized break point, we could not reject the hypothesis of stable coefficients. We also presented evidence showing that the estimated residuals have not remained constant over this time period. Further testing indicated that the reason for this heteroskedasticity stems from the large errors experienced by each equation primarily during the turbulent 1980–82 period.

Although the evidence reveals large quarterly forecast errors during the past five years, the results also show that these errors are offsetting in sign and magnitude. In fact, the sum of the forecast errors from each model is not statistically different from zero. This result substantiates previous findings from studies of velocity growth in which the forecast accuracy improved as the forecast horizon was lengthened. In this vein, arguments that monetary targeting to achieve the long-term goals of stable income growth and price stability has become useless because of purported money demand instability are not supported by the evidence.

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