In Search of a Stable, Short-Run M1 Demand Function

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Conventional[1] M1 demand equations went off track at least twice during the 1980s, failing to predict either the large decline in M1 velocity in 1982-83 or the explosive growth in M1 in 1985-86. A number of hypotheses were advanced to explain the prediction errors, but none of these were completely satisfactory.[2] As a result, several analysts have concluded that there has been a fundamental change in the character of M1 demand.

In recent years, some economists have sought to fix conventional M1 demand functions by focusing on specifications that pay adequate attention to the long-run nature and short-run dynamics of money demand. As is well known, conventional money demand functions have been estimated using data either in levels or in differences. Recent advances in time series analysis designed to deal with nonstationary data, however, have raised doubts about either specification. This has led several analysts to integrate these two specifications using cointegration[3] and error-correction techniques. In this approach, one first tests for the presence of a long-run, equilibrium (cointegrating) relationship between real money balances and its explanatory variables including real income and interest rates. If the test for cointegration indicates that such a relationship exists, an equilibrium regression is fit using the levels of the variables. The calculated residuals from the long-run money demand regression are then used in an error-correction model, which specifies the short-run behavior of money demand. This approach thus results in a money demand specification which could include both levels and differences of relevant explanatory variables.[4]

Those who have used cointegration techniques to test for the existence of a long-run, equilibrium M1 demand function, however, have found mixed results. For example, Baum and Furno (1990), Miller (1991), and Hafer and Jansen (1991) do not find a long-run equilibrium relationship between real M1, real income, and a short-term nominal interest rate. Other analysts including Hoffman and Rasche (1991), Dickey, Jansen and Thornton (1991), and Stock and Watson (1991), on the other hand, have presented evidence favorable to the presence of a long-run relationship among these variables.[5]

This study examines whether conventional M1 demand functions reformulated using error-correction techniques can explain the short-run behavior of M1. Much of the recent work on M1 demand has focused on the search for a long-run money demand function. In fact, those economists, who have found

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[1] The term *conventional* is meant to indicate those money demand specifications in which the demand for real M1 depends only on real income and short-term interest rates. [For examples, see specifications given in Rasche (1987), Mehra (1989) and Hetzel and Mehra (1989)].

[2] See Rasche (1987), Mehra (1989), and Hetzel and Mehra (1989) for a discussion of various hypotheses and reformulated M1 demand regressions.

[3] Let X_{1t}, X_{2t}, and X_{3t} be three time series. Assume that the levels of these time series are nonstationary but first differences are not. Then these series are said to be cointegrated if there exists a vector of constants $(\alpha_1, \alpha_2, \alpha_3)$ such that $Z_t = \alpha_1 X_{1t} + \alpha_2 X_{2t} + \alpha_3 X_{3t}$ is stationary. The intuition behind this definition is that even if each time series is nonstationary, there might exist linear combinations of such time series that are stationary. In that case, multiple time series are said to be cointegrated and share some common stochastic trends. We can interpret the presence of cointegration to imply that long-run movements in these multiple time series are related to each other.

[4] Miller (1991), Mehra (1992), and Baba, Hendry and Starr (1991), among others, have used this approach to estimate money demand functions.

[5] Sample periods, measures of income and interest rates, tests for cointegration, and estimators of cointegrating vectors used in these studies differ. These factors outwardly appear to explain part of different results found in these studies. However, as shown in Stock and Watson (1991), the main reason for the sensitivity to the sample period and estimator used is the presence of multicollinearity between real income and interest rate in the post-World War II period. The presence of this multicollinearity has made it difficult to get reliable estimates of the long-run money demand parameters. Stock and Watson (1991), however, note that the disappearance since 1982 of the trend in interest rates has reduced the extent of this multicollinearity. This may make it possible to get more reliable estimates of the long-run money demand function over the sample period that includes more of post-1982 observations.

a long-run cointegrating relationship between real M1 and its explanatory variables (like real income and interest rates), either have not constructed error-correction models of money demand or have constructed but failed to evaluate them for parameter stability and for explaining M1's short-run behavior.[6]

This study makes the basic assumption that there exists a long-run equilibrium relationship between real M1, real income, and an opportunity cost variable over the postwar period 1953Q1 to 1991Q2.[7] Under this assumption, error-correction models of M1 demand are constructed, tested for parameter stability, and evaluated for predictive ability. The empirical results indicate that these error-correction models do not depict parameter stability, nor do they adequately explain the short-run behavior of M1 in the 1970s and the 1980s. These results imply that the long-run M1 demand functions postulated here and in several recent M1 demand studies are misspecified. This has the policy implication that M1 remains unreliable as an indicator variable for monetary policy.

The plan of this study is as follows. Section I presents the basic error-correction model, reviews the Engle-Granger test of cointegration, and describes a simple procedure for estimating the error-correction model. Section II presents empirical results. Concluding observations are given in Section III.

I. THE MODEL AND THE METHOD

Specification of an M1 Demand Model

The general form of the error-correction money demand model estimated here is given below.

$$\ln(rM1)_t = \beta_0 + \beta_1 \ln(rY)_t$$
$$+ \beta_2 (R - RM1)_t + U_t \qquad (1)$$

[6] Only Hoffman and Rasche (1991) estimate the short-run error-correction model for M1, under the long-run specification that real M1 balances depend upon real income and a short-term interest rate. One important exception is the study by Baba, Hendry and Starr (1991), where the postulated long-run M1 demand function is complicated and differs substantially from that used by others. In particular, they assume that real M1 balances depend upon real income, one-month T-bill rate, the spread between long- and short-term rates, learning-adjusted yields on M1 and M2, and a moving standard deviation of holding period yields on long-term bonds. Given this long-run specification, they estimate an error-correction model for M1 and show that the model is stable over the sample period 1960Q3 to 1988Q3 studied there. The evaluation of this money demand model is outside the scope of the present study.

[7] I do, however, reproduce the mixed evidence found in recent studies on the existence of a long-run M1 demand function.

$$\Delta\ln(rM1)_t = \delta_0 + \sum_{s=1}^{n1} \delta_{1s} \Delta\ln(rM1)_{t-s}$$

$$+ \sum_{s=0}^{n2} \delta_{2s} \Delta\ln(rY)_{t-s}$$

$$+ \sum_{s=0}^{n3} \delta_{3s} \Delta(R - RM1)_{t-s}$$

$$+ \sum_{s=0}^{n4} \delta_{4s} \Delta^2\ln(p)_{t-s}$$

$$+ \delta_5 U_{t-1} + \epsilon_t, \qquad (2)$$

where rM1 is real M1 balances; rY real income; R a short-term nominal interest rate; RM1 the own rate of return on M1; p the price level; U and ϵ, random disturbance terms; ln the natural logarithm; Δ and Δ^2 the first- and the second-difference operators. Equation (1) is a long-run equilibrium M1 demand equation, which says that the long-run equilibrium demand for real M1 balances depends upon real income and an opportunity cost variable measured as the short-term nominal interest rate minus the own rate of return on M1. The parameter β_1 is the long-run real income elasticity and β_2 the long-run (semi-log) opportunity cost parameter. This equation is consistent with models of the transactions demand for money formulated in Baumol (1952) and Tobin (1956).

The presence of the disturbance term U_t in (1) implies that actual real M1 balances momentarily can differ from the long-run equilibrium value determined by factors specified in (1). Equation (2) describes the short-run behavior of M1 demand and is in a dynamic error-correction form, where δ_{is} (i = 2,3,4) measures the short-run responses of real M1 balances to changes in income, opportunity cost and inflation variables. The parameter δ_5 that appears on the disturbance term U_{t-1} is the error-correction coefficient and measures the extent to which actual real M1 balances adjust to clear disequilibrium in the public's long-term money demand holdings. This can be seen in (3), which is obtained by solving (1) for U_{t-1} and then substituting for U_{t-1} in (2).

$$\Delta\ln(rM1)_t = \delta_0 + \sum_{s=1}^{n1} \delta_{1s} \Delta\ln(rM1)_{t-s}$$

$$+ \sum_{s=0}^{n2} \delta_{2s} \Delta\ln(rY)_{t-s}$$

$$+ \sum_{s=0}^{n3} \delta_{3s} \Delta(R - RM1)_{t-s}$$

$$+ \sum_{s=0}^{n4} \delta_{4s} \, \Delta^2 \ln(p)_{t-s}$$

$$+ \delta_5 \, [\ln(rM1)_{t-1}$$

$$- \ln(rM1)^*_{t-1}] + \epsilon_t, \tag{3.1}$$

where

$$\ln(rM1)^*_{t-1} \equiv \beta_0 + \beta_1 \ln(rY)_{t-1}$$

$$+ \beta_2 \, (R - RM1)_{t-1}. \tag{3.2}$$

One can view $rM1^*$ as the long-term equilibrium real M1 balances, and $rM1$, of course, is actual real M1 balances. Thus, the term $[\ln(rM1) - \ln(rM1)^*]_{t-1}$ measures disequilibrium in the public's long-term real money balances. If the variables included in (1) are nonstationary but cointegrated, then the error-correction parameter is likely to be non-zero, i.e., $\delta_5 \neq 0$ in (3.1).

Another point to highlight is that equation (3.1) can be viewed as a generalization of the conventional partial-adjustment model, because the approach considered here allows separate reaction speeds to the different determinants of money demand (the coefficients δ_{2s}, δ_{3s}, δ_{4s} and δ_5 are different), yet via the error-correction mechanism ensures that actual real M1 balances converge to equilibrium levels in the long run.

The long-run money demand equation (1) is "conventional" in the sense that real M1 demand is assumed to depend only on real income and an opportunity cost variable. In particular, inflation is assumed to have no long-run effect on money demand. In this respect, the specification used here is similar to ones estimated recently in Dickey, Jansen and Thornton (1991), Hoffman and Rasche (1991), and Stock and Watson (1991). However, following Friedman (1959) the potential long-run influence of inflation on M1 demand is also examined (see footnote 11).

Even if inflation has no long-run effect on money demand, it could still influence real M1 balances in the short run because of the presence of adjustment lags.[8] Hence, the inflation variable appears in the short-run money demand equation (2) and is in first differences rather than in levels. This specification reflects the assumption that inflation is nonstationary.

However, the consequences of introducing inflation in levels or dropping it altogether from (2) are also examined (see footnote 18).

Estimation of the Error-Correction Model

If the disturbance term U_t is stationary, then the money demand model described above can be estimated in two alternative ways. The first is a two-step procedure given in Engle and Granger (1987). In the first step, the long-run money demand equation (1) is estimated by ordinary least squares and the residuals are calculated. In the second step, the short-run money demand equation (2) is estimated with U_{t-1} replaced by residuals in step one.

An alternative procedure is to estimate (1) and (2) jointly. This can be seen in (4), which is obtained by substituting (3.2) into (3.1).

$$\Delta \ln(rM1)_t = (\delta_0 - \delta_5 \beta_0) + \sum_{s=1}^{n1} \delta_{1s} \, \Delta \ln(rM1)_{t-s}$$

$$+ \sum_{s=0}^{n2} \delta_{2s} \, \Delta \ln(rY)_{t-s}$$

$$+ \sum_{s=0}^{n3} \delta_{3s} \, \Delta(R - RM1)_{t-s}$$

$$+ \sum_{s=1}^{n4} \delta_{4s} \, \Delta^2 \ln(p)_{t-s}$$

$$+ \delta_5 \, \ln(rM1)_{t-1}$$

$$- \delta_5 \beta_1 \, \ln(rY)_{t-1}$$

$$- \delta_5 \beta_2 \, (R - RM1)_{t-1} + \epsilon_t, \tag{4}$$

where all variables are defined as before. As can be seen, the long- and short-run parameters of the money demand model now appear in (4). All of the key parameters of (1) and (2)—such as those pertaining to income and opportunity cost variables—can be recovered from those of (4). The M1 demand equation here is estimated using the second procedure.[9]

Test for Cointegration: Engle-Granger Procedure

An assumption that is necessary to yield reliable estimates of the money demand parameters is that

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[8] The empirical work reported in Goldfeld and Sichel (1987) and Hetzel and Mehra (1989) is consistent with the presence of an inflation effect on money demand in the short run.

[9] The money demand model was also estimated using the first procedure, which generated qualitatively similar results on parameter stability and predictive ability.

the nonstationary variables included in (1) or in (4) are cointegrated as discussed in Engle and Granger (1987). Hence, one must first test for a cointegrating relationship between real M1 balances, real GNP and an opportunity cost variable, i.e., test whether U_t is stationary in (1).

Several tests for cointegration have been proposed in the literature [see, for example, Engle and Granger (1987) and Stock and Watson (1991)]. The test for cointegration used here is the one proposed in Engle and Granger (1987) and consists of two steps. The first tests whether each variable in (1) is nonstationary, which is done performing unit root tests on the variables. (The presence of a single unit root in a series implies that the series is nonstationary in levels but stationary in first differences.) The second step tests for the presence of a unit root in the residuals of the levels regressions estimated using the nonstationary variables. To explain further, assume that $\ln(rM1)_t$, $\ln(rY)_t$ and $(R - RM1)_t$ are nonstationary in levels. In order to test whether these variables are cointegrated, one needs to estimate the following regressions:

$$\ln(rM1)_t = \beta_0 + \beta_1 \ln(rY)_t$$
$$+ \beta_2 (R - RM1)_t + U_{1t}, \qquad (5.1)$$

$$\ln(rY)_t = \beta_3 + \beta_4 \ln(rM1)_t$$
$$+ \beta_5 (R - RM1)_t + U_{2t}, \qquad (5.2)$$

$$(R - RM1)_t = \beta_6 + \beta_7 \ln(rM1)_t$$
$$+ \beta_8 \ln(rY)_t + U_{3t}. \qquad (5.3)$$

If the residuals in any one of these regressions are stationary, then these variables are cointegrated.

Data, Definition of Variables, and Alternative Specifications

The money demand regression (4) is estimated using quarterly data over the period 1953Q1 to 1991Q2. Here rM1 is nominal M1 deflated by the implicit GNP deflator; rY real GNP; p the implicit GNP deflator; R the three-month Treasury bill rate; and RM1 the own rate of return on M1. The variable RM1 is defined as a weighted average of the explicit interest rates paid on the components of M1.[10]

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[10] The construction of the own rate on M1 is described in Hetzel (1989).

The opportunity cost variable in (1) is not in logarithms, whereas other variables are. This (semi-log) specification implies that the long-run opportunity cost elasticity varies positively with the level of the opportunity cost variable. I consider an alternative double-log specification in which the opportunity cost variable is also in logarithms. This specification implies that the long-term opportunity cost elasticity is constant. Furthermore, following Hoffman and Rasche (1991), the test for cointegration is also implemented including trend in the long-run part of the model (see the appendix in this paper).

II. EMPIRICAL RESULTS

Unit Root Test Results

The unit root tests are performed by estimating augmented Dickey-Fuller regressions of the form

$$X_t = a + \rho X_{t-1} + \sum_{s=1}^{k} b_s \Delta X_{t-s} + n_t, \qquad (6)$$

where X_t is the pertinent variable; n_t a random disturbance term; and k the number of lagged changes in X_t necessary to make n_t serially uncorrelated. If ρ equals one, then X_t has a unit root and is nonstationary. Two statistics are calculated to test the null hypothesis $\rho = 1$. The first is the t-statistic, $t_{\hat{\rho}}$, and the second is the normalized bias statistic, $T(\hat{\rho} - 1)$, where T is the number of observations. If these statistics have small values, then the null hypothesis is accepted.

Table 1 reports the unit root test results for the logarithm of real M1, the logarithm of real GNP, the level and the logarithm of the opportunity cost variable $(R - RM1)_t$, and the logarithm of the price level. These results indicate that real M1, real GNP and the opportunity cost variable are nonstationary in levels, but stationary in first differences. (The tests indicate the presence of a single unit root in these variables.) The test results for first differences of the logarithm of the price level, however, are mixed. The t-statistic, $t_{\hat{\rho}}$, indicates that the inflation variable is nonstationary, whereas the other statistic, $T(\hat{\rho} - 1)$, indicates that it is stationary.

Cointegration Test Results

Given the unit root test results, the logarithm of real M1, the logarithm of real GNP, and the logarithm (or the level) of opportunity cost are included in the cointegration tests. The inflation rate is not included because unit root test results are ambiguous

Table 1

Unit Root Test Results; 1953Q1–1991Q2

Augmented Dickey-Fuller Statistics

X_t	ρ	$t_{\hat\rho}$	$T(\hat\rho - 1)$	k	$\chi^2(1)$	$\chi^2(2)$	Q(36)
ln(rM1)$_t$	.99	− 1.3	− 1.6	5	.1	5.2	23.2
ln(rY)$_t$	.99	− .6	− .2	3	.6	1.1	25.6
(R − RM1)$_t$	.95	− 2.2	− 7.8	6			
ln(R − RM1)$_t$	.96	− 1.9	− 5.8	6	1.1	1.1	19.2
ln(p)$_t$	1.0	− .1	0.0	5	.6	1.0	28.8
Δln(rM1)$_t$	.65	− 3.6*	− 53.1*	6	.9	1.6	18.3
Δln(rY)$_t$	.32	− 6.5*	− 104.6*	2	.6	1.1	25.5
Δ(R − RM1)$_t$	.03	− 5.7*	− 158.8*	6	.5	.9	27.3
Δln(R − RM1)$_t$	.00	− 6.3*	− 153.6*	5	1.4	1.4	19.4
Δln(p)$_t$	.89	− 1.8	− 16.3*	4	.5	1.0	28.8

Notes: rM1 is real M1 balances; rY real GNP; R − RM1 the difference between the three-month Treasury bill rate (R) and the own rate on M1(RM1); and p the implicit GNP deflator. RM1 is a weighted average of the explicit rates paid on the components of M1. ln is the natural logarithm and Δ the first-difference operator.

Augmented Dickey-Fuller statistics are from the regression

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

where X_t is the pertinent variable; k the number of lagged first differences of X_t included to remove serial correlation in the residuals. $t_{\hat\rho}$ is the t-statistic and $T(\hat\rho - 1)$ the normalized bias statistic. Both are used in the test of the null hypothesis that $\rho = 1$. T is the number of observations used in the regression. k is chosen by the final prediction error criterion given in Akaike (1969). $x^2(1)$ and $x^2(2)$ are Godfrey statistics, which test for the presence of first- and second-order serial correlation in the residuals. Q(36) is the Ljung-Box statistic, which tests for the presence of higher-order serial correlation and is based on 36 autocorrelations.

"*" indicates significant at the 5 percent level. The 5 percent critical values for $t_{\hat\rho}$ and $T(\hat\rho - 1)$ statistics are − 2.89 and − 13.7, respectively. [See Tables 8.5.1 and 8.5.2 of Fuller (1976).]

about its nonstationarity.[11] Table 2 presents cointegration test results using the Engle-Granger procedure. As can be seen, these test results are mixed. For the semi-log specification, the test results indicate that real M1 balances are cointegrated with real income and interest rates, and this conclusion is not sensitive to the particular normalization chosen, i.e., the choice of the dependent variable in the cointegrating regression (compare results in rows 1

[11] Is the inflation variable, when treated as nonstationary and included in the cointegration regression, statistically significant? In order to answer this question, I estimated, following Stock and Watson (1991), the dynamic version of (1) by ordinary least squares. That is, the cointegrating regression (1) was estimated including, in addition, current, past, and future values of first differences of real income, opportunity cost and inflation variables and the current value of the inflation variable. The estimated coefficient on the current value of the (level) inflation variable is small and not statistically significant. This result indicates that the inflation variable does not enter the cointegrating regression (1). (In contrast, real income and opportunity cost variables were statistically significant.)

through 3 of Table 2). For the double-log specification, the test results indicate cointegration only if the cointegrating regression is normalized on the interest rate variable (compare results in rows 4 through 6 of Table 2).[12] Despite these mixed results, I proceed under the assumption that real M1 is cointegrated with real income and interest rates over the period studied here.

The Engle-Granger procedure also generates point-estimates of the long-run income and opportunity cost coefficients. For the semi-log specification, the point-estimates of the long-run income elasticity range from .31 to .44 and those for the opportunity cost parameter range from − .03 to − .04. For the

[12] This explains why Baum and Furno (1990) and Miller (1991) conclude that real M1 is not cointegrated with real income and interest rates. These authors implement the test for cointegration by estimating the cointegration regression normalized on the M1 variable.

Table 2

Cointegration Test Results: Engle-Granger Procedure

Row #	Dependent Variable	Cointegrating Vector			Augmented Dickey-Fuller Statistics			
		ln(rY)	(R − RM1)	ln(R − RM1)	$t_{\hat{d}}$	k	$x^2(1)$	$x^2(2)$
1	ln(rM1)	.31	−.03		−3.58*	5	.6	3.9
2	ln(rY)	.45	−.04		−3.90*	5	.9	4.5
3	(R − RM1)	.44	−.05		−4.83*	5	.3	2.4
4	ln(rM1)	.36		−.15	−2.57	6	.3	1.2
5	ln(rY)	.52		−.22	−2.89	6	.6	1.1
6	ln(R − RM1)	.53		−.29	−4.98*	5	1.6	1.6

Notes: The left part of the table reports estimates of the long-run income and interest rate coefficients from the cointegrating regressions estimated using alternative dependent variables [see equation (6) in the text]. The right part of the table presents statistics from the augmented Dickey-Fuller (ADF) regression that is used to test for the presence of a unit root in the residuals of the relevant cointegrating regression. The ADF regression is of the form

$$\Delta \hat{U}_t = d \, \hat{U}_t + \sum_{s=1}^{k} b_s \, \Delta \hat{U}_{t-s},$$

where $\hat{U}_t$ is the residual from the relevant cointegrating regression. $t_{\hat{d}}$ is the t-statistic that tests the null hypothesis that $\hat{d}=0$. k is the number of lagged differences of $\hat{U}_t$ in the regression and is chosen by the final prediction error criterion. $x^2(1)$ and $x^2(2)$ are Godfrey statistics, which test for the presence of first- and second-order serial correlation in the residuals of the relevant ADF regression.

"*" indicates significant at the 5 percent level. The 5 percent critical value for $t_{\hat{d}}$ is 3.62 [see Table 3 in Engle and Yoo (1987)].

double-log specification, the ranges for income and opportunity cost elasticities are .36 to .53 and −.15 to −.29, respectively.[13]

Figure 1 shows actual and fitted values from the long-run, semi-log money demand function ($\hat{\beta}_1$ = .44, $\hat{\beta}_2$ = −.05, $\hat{\beta}_0$ = −1.5), whereas Figure 2 shows the same for the double-log version ($\hat{\beta}_1$ = .53, $\hat{\beta}_2$ = −.29, $\hat{\beta}_0$ = −2.11). As can be seen, actual and predicted real money balances do not permanently drift away from each other in the long run. However, over several fairly long intervals actual real money balances persistently differ from the levels predicted by these cointegrating regressions. In order to examine whether such misses can be explained by short-run dynamics, error-correction models are estimated.

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[13] The point-estimates of the long-run income and interest rate coefficients are sensitive to the normalization chosen. To explain further, consider the cointegration regression (1). One can re-write (1) as

$$\ln(rY)_t = -\beta_0/\beta_1 + (1/\beta_1) \ln(rM1)_t - (\beta_2/\beta_1) (R - RM1)_t,$$

which is the cointegrating regression normalized on the income variable. From this regression, one can recover estimates of the long-run income elasticity β_1 [which is the inverse of the estimated coefficient on ln(rM1)$_t$] and the long-term interest rate coefficient β_2 [which is the coefficient on (R − RM1)$_t$ divided by the coefficient on ln(rM1)$_t$]. Another set of point-estimates can be recovered from the cointegration regression normalized on the interest rate variable.

Error-Correction M1 Demand Regressions

The results of estimating (4) are reported in Table 3. The opportunity cost variable, (R − RM1), is in levels in Equation A and in logarithms in Equation B. Equations A and B include levels, first differences, and second differences of the pertinent variables and are estimated by ordinary least squares. The estimated regressions look reasonable: all estimated coefficients possess theoretically correct signs and are generally statistically significant. The point-estimates of the long-run GNP elasticity range from .48 to .54. The point-estimate of the long-run opportunity cost elasticity is −.23 in Equation B and −.21 in Equation A; the latter elasticity is calculated as the product of the estimated semi-log opportunity cost parameter (−.04) and the sample mean value of the opportunity cost variable (5.19). These point-estimates of the long-run income and opportunity cost elasticities are close to the estimates generated by the (two-step) Engle-Granger procedure (see Table 2). The hypothesis that the long-run income elasticity is .5 could not be rejected.[14]

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[14] The test of this hypothesis is that the estimated coefficient on ln(rY)$_{t-1}$ and one-half of the estimated coefficient on ln(rM1)$_{t-1}$ add up to zero, i.e., ½ δ₅ − δ₅ β₁ = ½ δ₅ − ½ δ₅ = 0 in (3). The F-statistic (1,143) that tests the above hypothesis is .09 for Equation A and .08 for Equation B. These F-values are small and indicate that the long-run income elasticity is not different from .5.

Figure 1

ACTUAL AND PREDICTED VALUES BY THE COINTEGRATING REGRESSION

Predicted Value

Actual Real Money Balances

Cointegrating Regression: ln(rM1) = −1.5 + .44 ln(rY) − .05 (R−RM1)
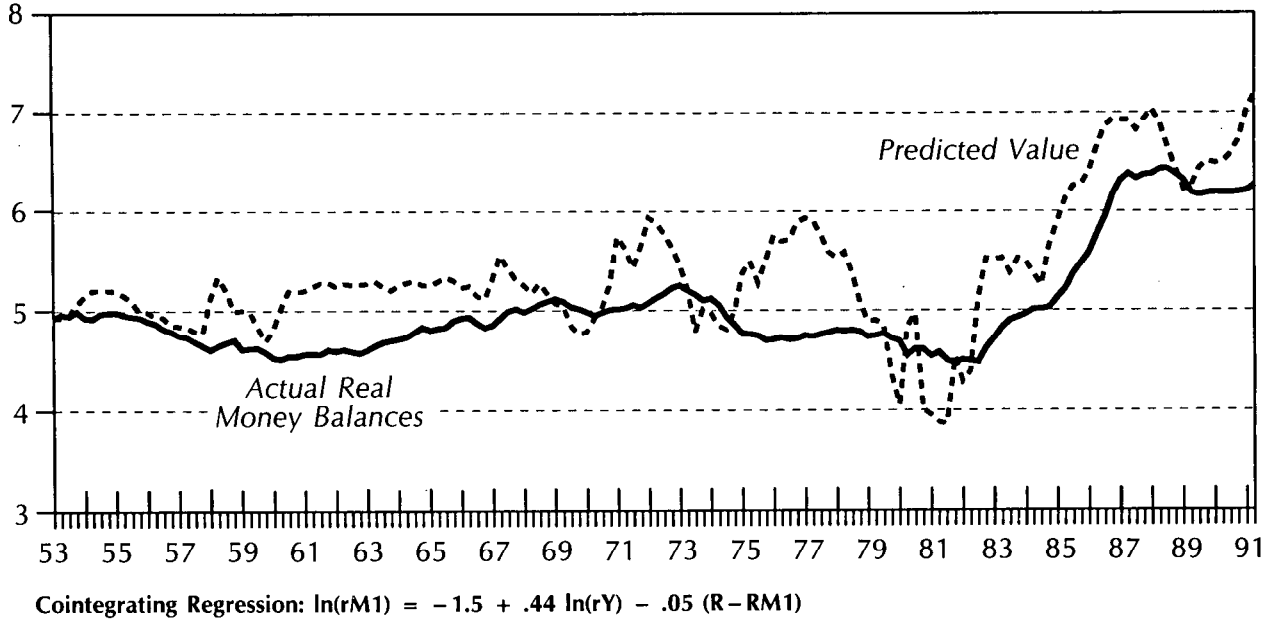
Figure 2

ACTUAL AND PREDICTED VALUES BY THE COINTEGRATING REGRESSION

Predicted Value

Actual Real Money Balances

Cointegrating Regression: ln(rM1) = −2.11 + .53 ln(rY) − .29 ln(R−RM1)

Table 3

Error-Correction M1 Demand Regressions; 1953Q1–1991Q2

A. Semi-Log Specification

$$\Delta\ln(rM1)_t = -.04 -.023 \ln(rM1)_{t-1} + .011 \ln(rY)_{t-1} - .0009 (R-RM1)_{t-1} + .11 \Delta\ln(rY)_t + .39 \Delta\ln(rM1)_{t-1}$$
$$\quad (2.2) \quad (2.2) \qquad\qquad (2.5) \qquad\qquad (2.1) \qquad\qquad\qquad (1.8) \qquad\qquad (5.7)$$

$$+ .25 \Delta\ln(rM1)_{t-2} - .000 \Delta(R-RM1)_t - .005 \Delta(R-RM1)_{t-1} - .71 \Delta^2\ln(p)_t - .26 \Delta^2\ln(p)_{t-1}$$
$$\quad (3.7) \qquad\qquad (0.0) \qquad\qquad (7.9) \qquad\qquad\qquad (6.5) \qquad\qquad (2.1)$$
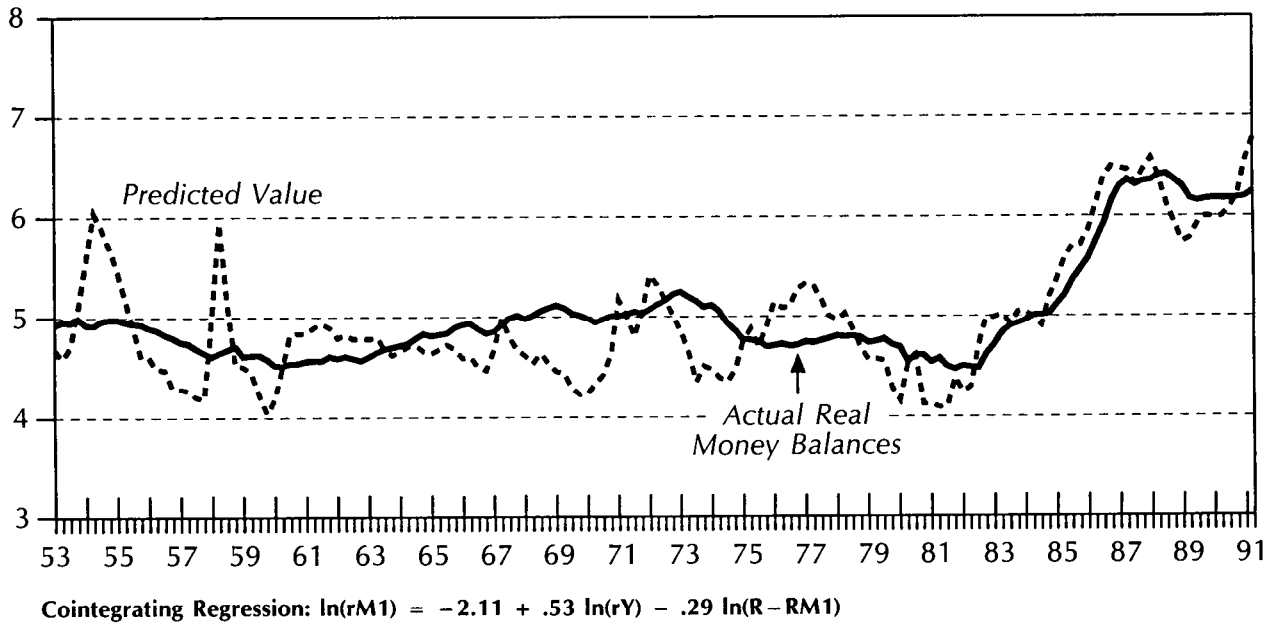
CRSQ = .68   SER = .00598   DW = 1.96   Q(5) = 3.4   Q(10) = 13.5 N_{ry} = .48 $N_{(R-RM1)}$ = −.04

B. Double-Log Specification

$$\Delta\ln(rM1)_t = -.06 - .026 \ln(rM1)_{t-1} + .014 \ln(rY)_{t-1} - .006 \ln(R-RM1)_{t-1} + .11 \Delta\ln(rY)_t + .39 \Delta\ln(rM1)_{t-1}$$
$$\quad (2.3) \quad (2.3) \qquad\qquad (2.5) \qquad\qquad (2.2) \qquad\qquad\qquad (1.7) \qquad\qquad (5.2)$$

$$+ .24 \Delta\ln(rM1)_{t-2} -.001 \Delta\ln(R-RM1)_t - .023 \Delta\ln(R-RM1)_{t-1} - .72 \Delta^2\ln(p)_t - .29 \Delta^2\ln(p)_{t-1}$$
$$\quad (3.1) \qquad\qquad (.6) \qquad\qquad (5.1) \qquad\qquad\qquad (6.0) \qquad\qquad (2.2)$$

CRSQ = .61   SER = .00659   DW = 2.0   Q(5) = 8.5   Q(10) = 16.5 N_{ry} = .54 $N_{\ln(R-RM1)}$ = −.23

Notes: Error-correction regressions are estimated by ordinary least squares. Parentheses contain the absolute value of t-statistics. CRSQ is the corrected R^2; DW the Durbin-Watson statistic; and SER the standard error of regression. Q(5) and Q(10) are Ljung-Box Q-statistics and are based, respectively, on five and ten autocorrelations of the residuals. N_{ry} is the long-term real GNP elasticity and is given by the estimated coefficient on $\ln(rY)_{t-1}$ divided by the estimated coefficient on $\ln(rM1)_{t-1}$. The relevant long-term interest rate coefficient N_{R-RM1} (or $N_{\ln(R-RM1)}$) is given by the coefficient on $(R-RM1)_{t-1}$ [or $\ln(R-RM1)_{t-1}$] divided by the coefficient on $\ln(rM1)_{t-1}$.

Another result to highlight is that the error-correction money demand regressions reported here yield estimates of the long-term opportunity cost $(R-RM1)$ elasticity substantially greater than those given by existing money demand regressions.[15] Hoffman and Rasche (1991), who also use error-correction techniques, report estimates (absolute values) of equilibrium interest elasticities that are of the order .4 to .5 for real M1, versus .21 to .23 reported here.[16]

Evaluating Money Demand Regressions

The money demand regressions reported in Table 3 are now evaluated by examining their structural stability and out-of-sample forecast performance.

The structural stability of these regressions is examined by means of a Chow test, with alternative

breakpoints which begin in 1971Q4 and end in 1983Q4 (the start and end dates include periods over which conventional M1 demand functions show instability). The Chow test is implemented using slope dummies on the variables. The restriction that the long-run real GNP elasticity is .5 is imposed. In addition, the stability of the regressions estimated allowing more lags on the explanatory variables than are used in the regressions given in Table 3 is also examined.

Table 4 presents results of the Chow test. F is the F-statistic that tests whether all of the slope dummies plus the one on the constant term are zero. F-statistics for Equations A and B of Table 3 are reported under the columns labeled "Specific." The columns labeled "General" contain results for regressions estimated with more lags on the explanatory variables. As can be seen, the F-values reported there are generally large and thus consistent with the hypothesis that the money demand regressions reported in Table 3 are not stable over the sample period studied.

Equation A of Table 3, which permits varying opportunity cost elasticity, is stable relative to Equation B (compare F-values for Equations A and B under the columns "Specific" in Table 4). This money demand regression depicts parameter stability during the 1970s, but then it breaks down during

[15] For example, a conventional M1 demand equation given in Hetzel and Mehra (1989) was reestimated using data in differences over the period 1953Q1 to 1980Q4. The income elasticity was estimated to be .52 and the opportunity cost elasticity −.04. The estimated income elasticity is close to the value generated using the error-correction model of M1 demand; in contrast, the opportunity cost elasticity is low, i.e., .04 versus .23 given by the error-correction model.

[16] Hoffman and Rasche (1991) do not include the own rate on M1 in defining the opportunity cost variable. This omission could bias upward the coefficient estimated on the interest rate variable and could explain relatively higher estimates of equilibrium interest elasticities reported in their study.

Table 4

Stability Tests; 1953Q1–1991Q2

Breakpoint	Equation A		Equation B	
	General	Specific	General	Specific
	F (26,102)	F (10,134)	F (26,102)	F (10,134)
1971Q4	1.01	1.22	1.91*	4.24*
1972Q4	1.04	1.24	2.09*	4.99*
1973Q4	1.37	1.50	2.75*	5.75*
1974Q4	1.38	1.61	2.46*	5.44*
1975Q4	1.26	.84	2.37*	5.02*
1976Q4	1.53	.76	2.57*	5.17*
1977Q4	1.64*	.88	2.89*	5.84*
1978Q4	1.52	1.09	2.97*	6.05*
1979Q4	1.87*	1.33	2.78*	6.16*
1980Q4	1.89*	1.22	2.51*	3.86*
1981Q4	1.97*	1.74	2.78*	3.19*
1982Q4	1.55	2.05*	1.53*	2.17*
1983Q4	2.00*	2.14*	1.89*	2.24*

Notes: The reported values are the F-statistics that test whether slope dummies when added to Equations A and B are jointly significant. The values reported under the column "Specific" are for Equations A and B reported in Table 3. The values reported under the column "General" are for versions of Equations A and B that are estimated including five lags of first-differenced variables. The breakpoint refers to the point at which the sample is split in order to define the dummies. The dummies take values one for observations greater than the breakpoint and zero otherwise. Parentheses contain degrees of freedom for the F-statistics.

"*" indicates significant at the 5 percent level.

the 1980s. In order to provide a different insight into the timing of predictive failure, I generate out-of-sample predictions of M1 growth conditional on actual values of income and interest rate variables. The predicted values are generated using Equation A of Table 3 and are for forecast horizons one to three years in the future.[17]

The results are reported in Table 5, which contains actual M1 growth as well as prediction errors (with summary statistics) for various forecast horizons. The results presented there suggest two observations. The first is that this regression cannot account for the "missing M1" in 1974-76 and "too much M1" in 1985-86. The explosion in M1 that occurred in 1982-83 is, however, well predicted. The

second is that prediction errors do not decline much as the forecast horizon is extended. The root mean squared error (RMSE), which is 2.7 percentage point for one-year horizon, declines slightly to 2.3 percentage point for three-year horizon. This result suggests that short-term misses in M1 are not reversed soon and can persist over periods longer than three years in the future.[18]

The out-of-sample predictions given in Table 5 are further evaluated in Table 6, which presents regressions of the form

$$A_{t+s} = c_0 + c_1 P_{t+s}, \quad s = 1,2,3, \qquad (7)$$

[17] The forecasts and errors were generated as follows. The money demand model was first estimated over an initial estimation period 1953Q1 to 1970Q4 and then simulated out-of-sample over one to three years in the future. For each of the forecast horizons, the difference between actual and predicted growth was computed, thus generating one observation on the forecast error. The end of the initial estimation period was then advanced four quarters and the money demand function was re-estimated, forecasts generated, and errors calculated as above. This procedure was repeated until it used the available data through the end of 1990.

[18] The short-run M1 demand equations were also estimated excluding inflation or including inflation in levels as opposed to first differences. Such regressions were then examined for their parameter stability and forecast performance. The results were qualitatively similar to those presented in the text. In particular, such M1 demand equations continue to depict parameter instability and fail to explain the weak M1 growth in 1974-76 and the subsequent explosion in 1985-86. The M1 demand equation estimated excluding inflation cannot even explain the explosive growth in 1982-83.

Standard M1 demand equations reported in Hetzel and Mehra (1989) were also estimated and simulated over the updated sample period 1981Q1 to 1991Q2. Such M1 demand regressions continue to underpredict M1 growth in the 1980s.

Table 5

Rolling-Horizon Forecasts of M1 Growth; 1971–1990

Year	1 Year Ahead			2 Years Ahead			3 Years Ahead		
	Actual	Predicted	Error	Actual	Predicted	Error	Actual	Predicted	Error
1971	6.4	9.3	−2.8	—	—	—	—	—	—
1972	8.0	7.3	.7	7.2	8.5	−1.3	—	—	—
1973	5.5	5.4	.1	6.8	5.9	.8	6.7	6.9	−.2
1974	4.7	7.0	−2.3	5.1	6.1	−1.0	6.1	6.3	−.2
1975	4.7	10.5	−5.8	4.7	8.9	−4.3	4.9	7.7	−2.7
1976	5.9	7.7	−1.8	5.3	9.9	−4.5	5.1	8.9	−3.7
1977	7.9	8.7	−.8	6.9	8.4	−1.5	6.2	9.6	−3.4
1978	7.9	7.7	.1	7.9	8.1	−.2	6.7	8.2	−.9
1979	7.0	5.2	1.8	7.4	6.5	.9	7.6	7.1	.5
1980	7.2	4.9	2.2	7.1	5.1	1.9	7.4	6.1	1.3
1981	5.2	3.0	2.2	6.2	3.8	2.4	6.5	4.1	2.3
1982	8.4	7.5	.9	6.8	4.9	1.9	6.9	4.7	2.3
1983	9.9	9.5	.4	9.1	7.7	1.5	7.8	5.8	2.0
1984	5.3	6.0	−.7	7.6	7.7	−.1	7.9	6.9	.9
1985	11.3	7.2	4.1	8.3	6.6	1.7	8.8	7.5	1.3
1986	14.4	8.9	5.4	12.9	6.9	5.9	10.3	6.6	3.7
1987	6.1	11.8	−5.6	10.2	8.1	2.1	10.6	6.9	3.7
1988	4.2	3.9	.3	5.1	9.3	−4.2	8.2	7.3	.9
1989	.6	1.8	−1.2	2.4	3.1	−.7	3.6	6.9	−3.3
1990	4.11	4.7	−.6	2.3	2.9	−.6	2.9	3.5	−.5

Mean Error			−.18			.03			.21
RMSE			2.7			2.5			2.3

Notes: Actual and predicted values are annualized rates of growth of M1 over 4Q to 4Q periods ending in the years shown. The predicted values are generated using money demand Equation A of Table 3 (see footnote 17 in the text for a description of the forecast procedure used). The predicted values are generated under the constraint that the long-run real GNP elasticity is .5.

Table 6

Out-of-Sample Forecast Performance

Error-Correction Equation	1 Year Ahead		2 Years Ahead		3 Years Ahead	
	c_0	c_1	c_0	c_1	c_0	c_1
Semi-Log	2.8	.57	3.9	.43	5.6	.19
(Equation A, Table 3)	(1.7)	(.23)	(1.9)	(.27)	(2.5)	(.32)
Double-Log	2.8	.57	4.1	.39	6.1	.12
(Equation B, Table 3)	(2.1)	(.28)	(2.3)	(.31)	(2.5)	(.34)

Notes: The table reports coefficients (standard errors in parentheses) from regressions of the form $A_{t+s} = c_0 + c_1 P_{t+s}$, where A is actual M1 growth; P predicted M1 growth; and s (= 1,2,3) number of years in the forecast horizon. For Equation A, the values used for A and P are reported in Table 5. For Equation B, the predicted values used are not reported.

where A and P are the actual and predicted values of M1 growth. If these predictions are unbiased, then $c_0 = 0$ and $c_1 = 1$. As can be seen, estimated values of c_1 are less than one and those of c_0 different from zero.[19] These results suggest that the predictions of M1 growth generated by these error-correction models are biased.

III. CONCLUDING OBSERVATIONS

Recent advances in time series analysis designed to deal with nonstationary data have yielded new procedures for estimating long- and short-run econometric relationships. Several analysts have employed these techniques to study M1 demand, and some of them have concluded there exists a long-run equilibrium relationship between real M1, real income, and an opportunity cost variable.

This study also provides evidence consistent with the existence of a stationary linear relationship among these variables. Thus, actual real M1 balances do not drift permanently away from the levels predicted by such cointegrating regressions in the long run. However, in the short run, which can be fairly long,

actual real M1 balances differ persistently from the level predicted. The dynamic error-correction models estimated here generally fail the test of parameter stability and do not predict well the short-run changes in M1. In particular, the dynamic models estimated here fail to explain the well-known episodes of "missing M1" in 1974-76 and "too much M1" in 1985-86.[20]

The negative empirical results described above rather suggest that the character of M1 demand has changed in the 1980s. As recently shown in Hetzel and Mehra (1989) and Gauger (1992), the financial innovations of the 1980s caused M1 to become highly substitutable with the savings-type instruments included in M2. Conventional M1 demand equations reformulated here using error-correction techniques yield a high equilibrium interest rate elasticity and thereby capture somewhat better the increase in portfolio substitutions than do the standard (first-differenced) money demand equations. However, the results here suggest that they fail to capture all of the increase in portfolio substitutions. Until that is done, M1 remains unreliable as an indicator variable for monetary policy.

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[19] The Ljung-Box Q-statistics (not reported) that test for the presence of higher-order serial correlation in the residuals of (7) were generally small and not statistically significant. This result indicates that the estimated standard errors for coefficients (c_0 and c_1) reported in Table 6 are unbiased.

[20] Additional results presented in the appendix to this paper indicate that these conclusions are robust to some changes in specifications used in the text. In particular, the use of alternative measures of the scale variable and/or the inclusion of trend in money demand regression do not alter qualitatively the results summarized above.

APPENDIX

SENSITIVITY ANALYSIS

Introduction

M1 demand functions reported in the text used real GNP as a scale variable and are estimated without including a linear trend in the long-run part of the model. The results presented there suggested two major conclusions. The first is that the statistical evidence on the existence of a long-run cointegrating relationship among real M1, real income, and a short-term nominal rate is mixed. The second is that short-term M1 demand functions estimated using error-correction techniques depict parameter instability.

This appendix presents additional evidence suggesting that the conclusions stated above are not sensitive to the use of alternative scale measures (such as real personal income or real consumer spending)

in money demand equations. Nor do these conclusions change when a linear trend is included in the long-run part of the cointegrating regression. There, however, is one difference. When a linear trend is included in the cointegrating regression, the hypothesis that the long-run real GNP elasticity is unity, not .5, appears consistent with the data. Estimates of the long-run opportunity cost coefficient are, however, unchanged.

Cointegration Test Results: Alternative Scale Measures and Linear Trend

Table A.1 presents cointegration test results with alternative scale variables but with linear trend excluded from cointegrating regressions (as in the text), whereas Table A.2 presents results with linear trend

Table A.1

Cointegration Test Results; Linear Trend Excluded; Different Scale Measures

Row	Dependent Variable	Cointegrating Vector				Augmented Dickey-Fuller Statistics	
		ln(rPY)	ln(rC)	(R − RM1)	ln(R − RM1)$_t$	$t_{\hat{a}}$	k
1	ln(rM1)	.29		−.03		−3.6*	5
2	ln(rM1)		.29	−.03		−3.6*	5
3	ln(rPY)	.42		−.04		−3.8*	5
4	ln(rC)		.42	−.03		−3.8*	5
5	(R − RM1)	.41		−.05		−4.8*	5
6	(R − RM1)		.41	−.05		−4.8*	5
7	ln(rM1)	.33			−.14	−2.5	6
8	ln(rM1)		.33		−.13	−2.4	6
9	ln(rPY)	.49			−.21	−2.8	6
10	ln(rC)		.48		−.20	−2.6	6
11	ln(R − RM1)	.50			−.29	−4.9*	5
12	ln(R − RM1)		.49		−.27	−3.8*	6

Notes: See notes in Table 2 of the text. rPY is real personal income and rC real consumer spending.

Table A.2

Cointegration Test Results; Linear Trend Included

Row	Dependent Variable	Cointegrating Vector					Augmented Dickey-Fuller Statistics	
		ln(rY)	ln(rPY)	ln(rC)	(R − RM1)	ln(R − RM1)$_t$	$t_{\hat{a}}$	k
1	ln(rM1)	.61			−.03		−3.4	5
2	ln(rM1)		.85		−.03		−3.2	5
3	ln(rM1)			1.5	−.02		−3.0	5
4	ln(rY)	4.2			−.04		−1.9	3
5	ln(rPY)		4.2		−.04		−1.7	5
6	ln(rC)			3.9	−.02		−1.9	1
7	(R − RM1)	1.02			−.05		−4.6*	5
8	(R − RM1)		1.3		−.05		−4.4*	5
9	(R − RM1)			1.3	−.04		−4.8*	5
10	ln(rM1)	.96				−.17	−3.2	5
11	ln(rM1)		1.2			−.17	−2.6	6
12	ln(rM1)			1.8		−.13	−3.3	5
13	ln(rY)	3.3				−.27	−2.9	5
14	ln(rPY)		3.7			−.26	−2.3	3
15	ln(rC)			3.5		−.14	−3.1	5
16	ln(R − RM1)	1.6				−.29	−5.3*	5
17	ln(R − RM1)		1.9			−.29	−5.3*	5
18	ln(R − RM1)			1.9		−.22	−5.6*	5

Notes: See notes in Table 2 of the text. rY is real GNP; rPY real personal income; and rC real consumer spending.

included in such regressions. The results are presented for alternative scale measures such as real GNP, real personal income, and real consumer spending. As can be seen, these test results indicate cointegration if the test is implemented with cointegrating regressions normalized on the interest rate variable. Otherwise, cointegration test results are sensitive to the particular specification employed. In particular, with cointegrating regressions normalized on real M1, the test results indicate cointegration if linear trend is excluded and if the semi-log specification is employed.

If we focus on specifications which indicate cointegration among real M1, real income (or real consumer spending) and an opportunity cost variable, the resulting point-estimates of the long-run income elasticity are sensitive to the treatment of linear trend. When linear trend is included in cointegrating regressions, it is difficult to reject the hypothesis that the long-term income elasticity is unity. However, when linear trend is excluded, the results instead indicate that the long-term income elasticity is not different from .5. Estimates of the long-term opportunity cost parameter (or elasticity) are not sensitive. In sum, cointegration test results are sensitive to the treatment of linear trend in the nonstationary part of the model and thus provide mixed evidence on the presence of a cointegrating relationship between variables studied here.

Error-Correction M1 Demand Regressions: Tests of Parameter Stability

Despite the mixed evidence on cointegration, error-correction M1 demand regressions were estimated using alternative scale measures and including linear trend in the long-run part of the money demand model. Tables A.3 and A.4 present such regressions for selected measures of income. (In Table A.3, regressions are estimated without including trend and real personal income is used as the income variable. In Table A.4, regressions are estimated including linear trend and real GNP is used as the scale variable. Regressions using other alternative measures considered here are similar and not reported.) As can be seen, estimated regressions look reasonable. The point-estimates of the long-term income elasticity is between 1.04 and 1.09 when linear trend is included in regressions, but is between .44 and .48 if not. The point-estimate of the opportunity cost elasticity, however, is quite robust.

Table A.5 and A.6 present results of implementing the Chow test of stability (as explained in the text). As can be seen, reported regressions do not depict parameter stability over the sample period studied here.

Table A.3

**Error-Correction M1 Demand Regressions; Linear Trend Excluded;
Real Personal Income as a Scale Variable**

C. Semi-Log Specification

$$\Delta\ln(rM1)_t = .01 - .023 \ln(rM1)_{t-1} + .01 \ln(rPY)_{t-1} - .001 (R - RM1)_{t-1} + .19 \Delta\ln(rPY)_t + .40 \Delta\ln(rM1)_{t-1}$$
$$(1.1) \quad (2.2) \qquad\qquad (2.5) \qquad\qquad (2.1) \qquad\qquad (2.4) \qquad\qquad (5.9)$$

$$+ .24 \Delta\ln(rM1)_{t-2} - .0005 \Delta(R - RM1)_t - .005 \Delta(R - RM1)_{t-1} - .64 \Delta^2\ln(p)_t - .22 \Delta^2\ln(p)_{t-1}$$
$$(3.5) \qquad\qquad (.8) \qquad\qquad (7.8) \qquad\qquad (5.6) \qquad\qquad (1.8)$$

CRSQ = .69   SER = .00589   DW = 2.0   Q(5) = 3.8   Q(10) = 13.3 N_{rpy} = .44

$N_{(R - RM1)}$ = −.22 [evaluated at the sample mean value of (R − RM1)]

D. Double-Log Specification

$$\Delta\ln(rM1)_t = .01 - .027 \ln(rM1)_{t-1} + .013 \ln(rPY)_{t-1} - .006 \ln(R - RM1)_{t-1} + .26 \Delta\ln(rPY)_t + .40 \Delta\ln(rM1)_{t-1}$$
$$(1.1) \quad (2.5) \qquad\qquad (2.7) \qquad\qquad (2.4) \qquad\qquad (3.0) \qquad\qquad (5.4)$$

$$+ .21 \Delta\ln(rM1)_{t-2} - .005 \Delta\ln(R - RM1)_t - .02 \Delta\ln(R - RM1)_{t-1} - .62 \Delta^2\ln(p)_t - .23 \Delta^2\ln(p)_{t-1}$$
$$(2.8) \qquad\qquad (1.2) \qquad\qquad (5.1) \qquad\qquad (5.0) \qquad\qquad (1.7)$$

CRSQ = .62   SER = .00648   DW = 2.11   Q(5) = 9.8   Q(10) = 16.9 N_{rpy} = .48 $N_{(R - RM1)}$ = −.22

Notes: See notes in Table 3 of the text.

Error-Correction M1 Demand Regressions; Linear Trend Included;
Real GNP as the Scale Variable

E. Semi-Log Specification

$$\Delta\ln(rM1)_t = -.13 - .023 \ln(rM1)_{t-1} + .024 \ln(rY)_{t-1} - .001 \ln(R-RM1)_{t-1} - .0001 TR_{t-1} + .11 \Delta\ln(rY)_t$$
$$(1.3) \quad (2.2) \qquad\qquad (1.6) \qquad\qquad (2.2) \qquad\qquad\qquad (.9) \qquad\qquad (1.8)$$

$$+ .40 \Delta\ln(rM1)_{t-1} + .25 \Delta\ln(rM1)_{t-2} - .0005 \Delta\ln(R-RM1)_t - .006 \Delta\ln(R-RM1)_{t-1}$$
$$(5.8) \qquad\qquad (3.8) \qquad\qquad (.7) \qquad\qquad\qquad (7.9)$$

$$- .71 \Delta^2\ln(p)_t - .26 \Delta^2\ln(p)_{t-1}$$
$$(6.6) \qquad\qquad (2.2)$$

CRSQ = .68   SER = .00594   DW = 1.98   Q(5) = 3.62   Q(10) = 12.9 N_{ry} = 1.04

$N_{(R-RM1)}$ = −.22 [evaluated at the sample mean value of (R − RM1)]

F. Double-Log Specification

$$\Delta\ln(rM1)_t = -.19 - .031 \ln(rM1)_{t-1} + .034 \ln(rY)_{t-1} - .007 \ln(R-RM1)_{t-1} - .0001 TR_{t-1} + .11 \Delta\ln(rY)_t$$
$$(1.4) \quad (2.5) \qquad\qquad (1.6) \qquad\qquad (2.4) \qquad\qquad\qquad (.9) \qquad\qquad (1.7)$$

$$+ .39 \Delta\ln(rM1)_{t-1} + .24 \Delta\ln(rM1)_{t-2} - .004 \Delta\ln(R-RM1)_t - .022 \Delta\ln(R-RM1)_{t-1}$$
$$(5.2) \qquad\qquad (3.1) \qquad\qquad (.9) \qquad\qquad\qquad (5.0)$$

$$- .74 \Delta^2\ln(p)_t - .30 \Delta^2\ln(p)_{t-1}$$
$$(6.1) \qquad\qquad (2.3)$$

CRSQ = .61   SER = .00659   DW = 2.04   Q(5) = 8.7   Q(10) = 15.6 N_{ry} = 1.09 $N_{(R-RM1)}$ = −.22

Notes: TR is linear trend, and other variables are as defined before. See notes in Table 3 of the text.

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Table A.5

Stability Tests

Breakpoint	Equation C		Equation D	
	General	Specific	General	Specific
	F (26,102)	F (10,134)	F (26,102)	F (10,134)
1971Q4	1.20	1.17	2.63*	4.27*
1972Q4	1.31	1.30	3.03*	5.02*
1973Q4	1.44	1.79	3.21*	5.83*
1974Q4	1.55	1.52	3.12*	5.15*
1975Q4	1.49	1.25	2.99*	5.07*
1976Q4	1.86*	1.39	3.54*	5.33*
1977Q4	2.05*	1.70	3.92*	5.86*
1978Q4	2.15*	1.90*	4.32*	5.80*
1979Q4	1.89*	2.36*	3.64*	6.26*
1980Q4	2.00*	1.95*	2.88*	4.02*
1981Q4	1.92*	1.94*	3.33*	3.84*
1982Q4	1.62	2.16*	2.00*	2.56*
1983Q4	1.38*	2.27*	1.55*	2.60*

Notes: See notes in Table 4 of the text. Equations (specific) C and D are reported in Table A.3.

Table A.6

Stability Tests

Breakpoint	Equation E		Equation F	
	General	Specific	General	Specific
	F (26,102)	F (10,134)	F (26,102)	F (10,134)
1971Q4	1.13	1.75	1.87*	4.58*
1972Q4	1.29	1.98*	2.06*	5.51*
1973Q4	1.86*	3.02*	3.27*	8.15*
1974Q4	1.74*	2.88*	3.14*	8.39*
1975Q4	1.40	1.80	2.71*	7.40*
1976Q4	1.72*	1.46	3.02*	7.06*
1977Q4	1.65*	1.24	3.24*	7.60*
1978Q4	1.38	1.09	3.48*	7.75*
1979Q4	1.83*	1.94*	2.95*	7.51*
1980Q4	1.99*	2.30*	2.54*	4.92*
1981Q4	1.92*	2.63*	2.57*	3.90*
1982Q4	1.76*	2.73*	1.77*	2.01*
1983Q4	1.83*	2.25*	1.71*	2.20*

Notes: The reported values are the F-statistics that test whether slope dummies when added to Equations E and F are jointly significant. The statistics test stability of all coefficients except the one on the trend term. See also notes in Table 4 of the text. Equations (specific) E and F are reported in Table A.4.

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