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FURTHER EVIDENCE ON THE STABILITY OF THE SHORT-RUN DEMAND FOR MONEY

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I. Introduction

In response to various financial innovations over the last decade, the Federal Reserve Board has recently announced redefinitions of monetary aggregates. Two new transaction-type aggregates have replaced the old M1 measure. The M1A measure is equal to M1 less demand deposits due to foreign commercial banks and official institutions. The M1B aggregate is a somewhat broader measure: It adds to M1A those financial items that have the dual characteristic of being held both for check-writing purposes and as interest bearing assets. The most notable assets satisfying this dual role are negotiable orders of withdrawal (NOW accounts) and automatic transfer systems (ATS accounts).[1/]

This paper analyzes money-demand relationships using these alternative transaction-type monetary aggregates. Such an investigation is especially meaningful in light of the continuing problems associated with estimating traditional money-demand functions. An important issue considered is whether the new measurements resolve such problems.

The paper proceeds as follows: The estimation and forecasting accuracy of conventional log-level money demand equations, using M1, M1A, and M1B as alternative money definitions, are examined in section II. Section III investigates the possibility of intercept shifts in the relationship. As a preliminary step, the traditional log-level specification is first-differenced prior to estimation. The estimation results for these transformed equations are quite different from those obtained for the original log-level specification and are fully consistent with the notion of once-and-for-all intercept shifts. The hypothesis of once-and-for-all intercept shifts receives further support from the use of selective dummy variables in the original level equations. The conclusion drawn from these empirical tests is that the short-run money demand relationship has been subject to intercept shifts, but that the association between money demand and real income and interest rates remains unchanged. The importance of this conclusion regarding the implementation of monetary policy is examined in section IV. Concluding remarks are presented in section V.

II. The Log-Level Money-Demand Relationship

Using pre-1974 data, Goldfeld found the following conventional money-demand relationship to be quite "sturdy":

$$(1) \quad \ln (M_t/P_t) = \alpha_0 + \alpha_1 \ln y_t + \alpha_2 \ln RCP_t + \alpha_3 \ln RCB_t + \alpha_4 \ln (M_{t-1}/P_t) + \epsilon_t,$$

where M is M1, P represents the implicit GNP deflator, y is real GNP, RCP is measured by the 4-6 month prime commercial paper rate, and RCB equals a weighted average of commercial bank passbook rates (data sources and definitions are presented in the Appendix).[2/]

When this relationship was estimated for the sample period III/1959 - IV/1973 using ordinary least squares, there was evidence of significant first-order autocorrelation in the error terms.[3/] Consequently, the equation was reestimated utilizing the Cochrane-Orcutt (CORC) iterative technique. The estimated coefficients and key summary statistics from this estimation are presented in the first row of table 1. All estimated coefficients have the anticipated sign and are significantly different from zero at the five percent level. While the sample period is slightly different from that used by Goldfeld, the regression results reported in table 1 are very similar to his.[4/]

The same relationship given by equation (1) was also estimated with the monetary variable defined by the new transaction-type aggregate definitions, M1A and M1B. These regression results, again estimated utilizing the CORC procedure (again because of first-order serial correlation in the disturbances in the original equation), are presented in table 1. Because there is actually little difference in these alternative aggregate measures during this period, the coefficient estimates are quite similar. Utilizing the new monetary aggregates, however, increases the standard error of the estimated equation by approximately 10 percent over that obtained using M1 as the dependent variable. Also, the Durbin-h statistics are lower when these new aggregates are employed.

In a subsequent study, Goldfeld found that the conventional money-demand function was not as sturdy as demonstrated previously.[5] Specifically, he noted that the estimated relationship consistently and significantly over-projected money-demand during the 1974-76 period. Unable to explain this outcome to his own satisfaction, Goldfeld concluded that "some sort of shift has occurred" (his emphasis).[6] Other studies that have dealt with alternative measurements of the variables in (1) also have failed to resolve this problem.[7]

To investigate whether the adoption of the new aggregates improves the predictive performance of the money-demand relationship, each estimated relationship in table 1 was used to statically simulate the level of the respective nominal money balances over the period I/1974 - IV/1979. Key summary statistics outlining the results of this exercise are reported in table 2.[8/]

Unfortunately, the out-of-sample performance for each of the estimated money-demand relationships is very poor. First, the root-mean-squared-error (RMSE) for each equation is distressingly large. More importantly, however, is the problem of one-sided simulation errors: money-demand is consistently <u>over-predicted</u> for the entire post-sample period. This problem is evidenced by the Theil decomposition statistics which indicate that, in each case, over 75 percent of the simulation error is due to bias.[9/] Indeed, the poor predictive performance of the money-demand equations, especially that of the M1B equation, suggests that recent problems associated with estimating money-demand are <u>not</u> due solely to the introduction of ATS and NOW accounts, as some have suggested.

The persistent one-sided simulation errors during the I/1974 - IV/1979 period are very discouraging and suggest a significant deterioration in the underlying economic relationships. This conclusion is supported by the full III/1959 - IV/1979 sample period regression results for each money demand equation presented in table 3. A comparison of these results with those obtained for the earlier sample period (III/1959 - IV/1973) indicates that, irrespective of the monetary aggregate employed, the estimated coefficient on the commercial paper rate is the only coefficient which is similar in magnitude for the two alternative sample periods. Further, the estimated coefficient on the lagged dependent variable exceeds unity in all full-sample regressions, making economic interpretation of the results difficult if not meaningless. Finally, the coefficient on the real income variable is much smaller for the full sample period, implying a dramatic decline in the short-run income elasticity.

Chow tests were employed to ascertain whether the change in the estimated relationship was statistically significant. The evidence from these tests is reported in table 4.[10/] Two possible breakpoints were investigated: IV/1969, a point near the middle of the full sample; and IV/1973, a point of concern with regard to out-of-sample simulations. The null hypothesis in each case is that all regression coefficients remained unchanged over the full sample period. The calculated F-statistics are sufficiently large to allow the rejection of the null hypothesis at the 5 percent significance level in each of the six different cases considered. Thus, the estimated equations are statistically unstable over the III/1959 - IV/1979 sample period.

The evidence offered in this section suggests that the new transaction-type aggregates aid little, if any, in solving the "missing money" problem. Estimated money-demand relationships exploiting these new measures (M1A and M1B) do not appear more stable, nor do they significantly improve upon the out-of-sample prediction of money-demand relative to old M1.

III. The Breakdown Reconsidered

One aspect of estimating the short-run money-demand equation that has not received much attention in previous studies is the possibility that the equation has undergone one or several intercept shifts during the estimation period. As noted earlier, interest has been focused primarily on the "correctness" of the conventional equation in terms of the explanatory variables employed. Investigation into the existence and effects of once-and-for-all types of intercept shifts in the equation appears warranted.

In the present case, a relatively straight forward procedure is available to test for the presence of intercept shifts in the log-level specification. That procedure is to first-difference the log-level equation and estimate the resulting specification using ordinary least squares. Thus, instead of estimating equation (1), the equation

$$(2)\quad \ln (M_t/P_t) - \ln (M_{t-1}/P_{t-1}) = (\alpha_0 - \alpha_0) + \alpha_1 \ (\ln y_t - \ln y_{t-1}) + \alpha_2 (\ln RCP_t - \ln RCP_{t-1}) + \alpha_3 (\ln RCB_t - \ln RCB_{t-1}) + \alpha_4 [\ln (M_{t-1}/P_t) - \ln (M_{t-2}/P_{t-1})] + \epsilon_t - \epsilon_{t-1}$$

is estimated. Comparing equations (1) and (2), it is clear that they differ in one important feature; namely, the intercept term in the log-level relationship has no direct counterpart in the first-difference equation.

This difference in the two relationships, while insignificant on the surface, is of great importance in the event of an intercept shift in the log-level relationship. For example, consider the simple relationship

$$(3)\quad y_t = \beta_0 + \beta_1 X_t + \epsilon_t \qquad t=1, 2, \ldots, K.$$

Suppose that at some point, say $t=T$ $(T < K)$, there is a once-and-for-all intercept (constant) shift in the relationship so that the new intercept is β_0' $(\beta_0' \neq \beta_0)$.

Because of the intercept shift, two separate regimes exist:

(4) $y_t = \beta_0 + \beta_1 X_t + \epsilon_t$ for $t=1, 2, \ldots, T-1$

and

(5) $y_t = \beta_0' + \beta_1 X_t + \epsilon_t$ for $t=T, T+1, \ldots, K$.

Suppose that we desire to estimate equation (3) utilizing all K observations. In such a situation we will have T-1 observations from regime I (equation 4) and K-T observations from regime II (equation 5). If all K observations were employed in the estimation of relationship (3), and it was assumed (incorrectly) that no intercept shift occurred, the resulting estimate of β_1 clearly will be biased.[12/]

However, using the first-difference of the relationship (and assuming any required correction for first-order serial correlation is made),

(6) $y_t - y_{t-1} = \beta_1 (X_t - X_{t-1}) + \epsilon_t - \epsilon_{t-1}$,

the estimate of β_1 should be very close to its true value since equation (6) would be valid for all t

<u>except t=T</u>: Only at the point when the once-and-for-all intercept shift occurred would (6) be incorrect. At time T, equations (4) and (5) indicate that the true first-difference relationship is

$$(7) \quad y_T - y_{T-1} = \beta_0' - \beta_0 + \beta_1 (X_T - X_{T-1}) + \epsilon_T - \epsilon_{T-1}.$$

For any other time $t \neq T$, equation (6) would be correct.

It is in this sense that the first-difference specification may be viewed as a "robust" estimation technique in the face of a once-and-for-all intercept shift in the original log-level equation. The essential difference between the level and first-difference estimation is that, in the case of the former, the equation to be estimated, specifically equation (3), will be <u>false</u> in K-T observations. On the other hand, if equation (6) is estimated, this relationship would be false for <u>only one</u> observation, the point of the supposed shift.

To determine if the log-level specification has been subject to shifts in the intercept term, our analysis proceeds as follows. First, the first-difference specification of the alternative money-demand relationships are estimated. Then, like the analysis of the previous section, the "sturdiness" of these relationships in terms of their predictive performance as well as more formal tests are considered. Finally, information from the pattern observed in the estimated residuals from the first-difference equations is utilized to test directly for the existence and impact of intercept shifts on the conventional log-level money-demand equation.

First-Difference Results[13/]

Preliminary estimation of equation (2) using the alternative monetary aggregates yielded an estimated coefficient on the commercial bank passbook rate that was never significantly different from zero. This finding held for many different sample periods and for all monetary aggregates employed.[14/] In light of the passbook rate variable's statistical insignificance, it was excluded from estimation of all three money-demand relationships.

First-difference estimation results of the resulting money-demand specification for the sample period III/1959 - IV/1973 are found in table 5. The coefficient estimates reported there compare favorably in sign and magnitude with their respective counterparts obtained earlier for this same sample period. Further comparison of the CORC and first-difference regression results indicates that the latter estimation technique yields a slightly larger standard error of the estimating equation in each of the three relationships. It is also important to recognize that the Durbin-h statistics are now much smaller than those obtained in the log-level estimation, even after the use of the CORC procedure. Employing the first-differenced relationships, therefore, yield no evidence that suggests first-order serial correlation in the disturbances.

To compare the out-of-sample performance of the first-difference results with those obtained for the log-level equations, the estimated relationships in table 5 were transformed to statically simulate the respective post-sample nominal balances.[15] These results, provided in table 6, indicate a marked improvement over those of the log-level equations. For example, the RMSE's in table 6 are 60 to 70 percent lower than those reported in table 2. In addition, the absolute value of the average out-of-sample simulation errors have been reduced from a range of $5-$7 billion for the log-level regressions to less than $1 billion.

The marked reduction in the average post-sample simulation errors is further demonstrated in figures 1, 2, and 3, which plot the simulation errors in billions of nominal dollars for the log-level and first-difference money demand equations. As indicated by these figures, there is a striking difference between the error pattern for the log-level and first-difference equations. The growing out-of-sample simulation errors in the log-level equation are not evident when the first-difference specification is employed.

The Theil decomposition statistics indicate that the first-difference estimation procedure also alleviates the important problem of bias in the out-of-sample money-demand simulations found earlier. Indeed, the first-difference specification yields simulation errors that are due primarily to unequal variation between simulated and realized values of the aggregates. The values of the Theil covariation statistic (U^c) in table 5 are all greater than 86 percent while the proportion of the simulation errors due to bias (U^m) never exceed 13 percent. In contrast, log-level simulation results yielded a covariation statistic that never exceeds 11 percent while the bias proportion of the error (U^m) is <u>always</u> greater than 75 percent. Clearly, the first-difference simulation results are statistically superior to those of the log-level specification.

As far as a comparison of the new aggregates M1A and M1B is concerned, the out-of-sample simulation results obtained using first-differences slightly favor M1B. The smaller RMSE and mean error for the M1B equation, however, is due solely to the simulation results antedating IV/1978: Prior to then, the simulation errors for M1A and M1B are indistinguishable. It is not surprising that M1B performs better in out-of-sample simulations postdating III/1978, since this period coincides with the legalization of negotiable order of withdrawal (NOW) accounts in New York state and automatic transfer system (ATS) accounts on a nationwide basis. To the extent that individuals substituted NOW and ATS accounts for conventional demand deposits, M1B, being a more inclusive measure, captures this substitution.

The simulation results suggest that the gain from using M1B is minor. Moreover, by II/1979 it is again impossible to distinguish between M1A and M1B simulation errors. In this regard, it appears that the transition period following the introduction of NOW accounts lasted less than three quarters, and that the broader monetary aggregates is preferable over this period.

The ability to accurately simulate money-demand over the turbulent post-1973 period suggests that each of the first-differenced money-demand relationships is likely to be structurally stable over the full sample period. This supposition is supported by the full-sample coefficient estimates for the alternative money-demand relationships. Unlike the log-level regression results, the estimated coefficients listed in table 7 are remarkably similar to the estimates found for the shorter period III/1959 - IV/1973 in table 5. Indeed, the marked change in the coefficient estimates on the income and lagged dependent variables previously observed (tables 1 and 3) is no longer evident.

Chow tests were used to examine the structural stability of the first-difference equations. The results of these tests--using IV/1969 and IV/1973 as possible break points--are reported in table 8. In sharp contrast to the test results reported for the log-level equations (table 4), the findings for the first-difference equations (table 8) indicate that each specification indeed has remained stable over the III/1959 - IV/1979 sample period.[16/] Consequently, the results presented here suggest that each of the alternative first-differenced money-demand equations has not undergone the statistical breakdown in the underlying relationship shown to have occurred for the log-level regressions. Rather, the evidence is fully consistent with the notion of once-and-for-all intercept shifts in the log-level form of these relationships.

Testing for Intercept Shifts

We now directly investigate the possibility that the log-level money-demand relationship has been susceptible to once-and-for-all intercept shifts. A comparison of equations (6) and (7) in the previous section indicates that at the time of an intercept shift in the level relationship an unusually large error (residual) will occur if equation (6) is estimated. That is, at the point of the once-and-for-all intercept shift, time T, the observed estimation residual will include not only the difference in the disturbances, $\epsilon_T - \epsilon_{T-1}$, but also the change in the intercept from the level equation, $\beta_0' - \beta_0$. This suggests that the residuals from the first-difference relationship can be examined for outliers signaling possible points at which once-and-for-all intercept shifts occurred in the original log-level equation.

Examination of both the M1 and M1A first-difference equations' residuals revealed the same points of possible intercept shift. Over the III/1959 - IV/1979 period there were five residuals larger than twice the estimated standard error of each equation: III/1966, I/1975, IV/1975, I/1979, and II/1979.[17/] With the exception of II/1979, all of these large residuals were negative, suggesting "downshifts" in the log-level money-demand relationship (i.e., an absolutely larger intercept term). On the other hand, the residual in II/1979 was a large positive one suggesting a reversal of the previous "downshifting."

The pattern of residuals for M1B was slightly different from that for M1 or M1A. The large residual outliers for M1B occurred in III/1966, I/1975, IV/1975, and II/1979. Thus, the estimation results for this aggregate do not suggest a shift in the log-level relationship in I/1979, as did the M1 and M1A equations. Similar to the analysis of post-sample simulations, this evidence suggests a sizable movement of assets from traditional commercial bank demand deposits to ATS accounts and New York NOW accounts at this point in time. A downshift in the level of money-demand is observed when these latter assets are ignored, but is not apparent when they are included in the monetary aggregate.

To test for intercept shifts in the alternative log-level specifications, (0, 1) variables were employed. Because each shift in the intercept term is hypothesized to be a "once-and-for-all" type of displacement, the dummy variables are constructed to represent the constant term for each period in question. Thus, six dummy variables were employed when re-estimating the M1 and M1A equations. They are

$$D_0 = \begin{cases} 1 & \text{III/1959 - II/1966} \\ 0 & \text{otherwise} \end{cases}$$

$$D_1 = \begin{cases} 1 & \text{III/1966 - IV/1974} \\ 0 & \text{otherwise} \end{cases}$$

$$D_2 = \begin{cases} 1 & \text{I/1975 - III/1975} \\ 0 & \text{otherwise} \end{cases}$$

$$D_3 = \begin{cases} 1 & \text{IV/1975 - IV/1978} \\ 0 & \text{otherwise} \end{cases}$$

$$D_4 = \begin{cases} 1 & \text{I/1979} \\ 0 & \text{otherwise} \end{cases}$$

and

$$D_5 = \begin{cases} 1 & \text{II/1979 - IV/1979} \\ 0 & \text{otherwise.} \end{cases}$$

When re-estimating the M1B equation, the D_0, D_1 and D_2 dummies are again employed. In place of D_3, D_4, and D_5, however, two other dummies,

$$D_6 = \begin{cases} 1 & \text{IV/1975 - I/1979} \\ 0 & \text{otherwise} \end{cases}$$

and

$$D_7 = \begin{cases} 1 & \text{II/1979 - IV/1979} \\ 0 & \text{otherwise} \end{cases}$$

are used.

Preliminary CORC estimation results for the log-level money-demand relationships that included the respective dummy variables (not reported) revealed that certain constant terms are not significantly different from others. Regression results for M1 and M1A indicate that the hypothesized intercept shift occuring in III/1966 was not statistically significant using a conventional t-test. In other words, the constant term D_0 was not significantly different from D_1. Therefore, in the regression results reported (table 9), the variable D_0' is used where

$$D_0' = \begin{cases} 1 & \text{III/1959 - IV/1974} \\ 0 & \text{otherwise.} \end{cases}$$

A slightly different story unfolds with regard to M1B, however. In testing the significance of the various constants relative to each other, it was found again that D_1 was not significantly different from D_0; consequently, the variable D_0' is employed. More importantly, in the case of this broader M1B aggregate D_6 was found not to be significantly different from D_7. As a result, only one intercept was estimated over the post-IV/1975 period;

$$D_6' = \begin{cases} 1 & \text{IV/1975 - IV/1979} \\ 0 & \text{otherwise.} \end{cases}$$

Based on these results, the log-level specification employing the M1B definition of money was more stable (i.e., subject to fewer intercept shifts) than either M1 and M1A. The M1B relationship was determined to have shifted only twice over the full sample period and then again in IV/1975.

The results of estimating the log-level equations employing the intercept dummies are presented in table 9.[18/] The evidence strongly supports the hypothesis that the respective log-level relationships have been subject to significant shifts in the intercept term over the sample period.[19/] More important, however, is the fact that accounting for changes in the intercepts produces equations that are quite consistent with pre-1974 empirical results. Indeed, the results indicate that <u>the</u>

underlying economic relationship between real money balances and income and interest rates has not undergone significant changes over the past twenty years. The short-run income and interest elasticities remain quite stable.

IV. Policy Implications

The implications for the conduct of monetary policy, which result from the empirical evidence presented in this study, are at odds with what appears to have become accepted belief. Previous evidence, based on the standard log-level estimation, has suggested a wholesale evaporation of the money-demand relationship.[20/] Not only was the relationship found to be subject to random disturbances, but key elasticities changed markedly. This view is consistent with the log-level results presented in section II. Because of the highly unpredictable nature of money demand, therefore, the claim has been made that a monetary policy aimed at controlling growth rates in the monetary aggregates will fail. The belief has been that there is essentially no information available for the policymaker to determine future money-demand. As a result, it is impossible for the policymaker to determine the impact of any given pattern of monetary growth upon the economy.

The first-difference estimation results, however, indicate that the supposed breakdown in the money-demand relationship clearly has been exaggerated. Specifically, the money-demand elasticities required for the effective implementation of monetary policy have been quite stable over the last twenty years. Although this relationship was found to be subject to random shocks, the issue as to the choice of a monetary policy instrument (i.e., monetary aggregates versus interest rates) remains crucially related to the size of these disturbances _vis-a-vis_ real sector disturbances.[21/] Thus, those who have suggested that not only are there large random disturbances in money demand, but also that the partial relation between real money balances and income and interest rates is irreparably uncertain, have overstated the case against aggregate targeting. Clearly, further research on this topic is needed.

Conclusion

The empirical evidence presented in this paper indicates that, when properly estimated, the economic relationship between the demand for real money balances and real income and interest rates has remained invariant during recent economic history.

This claim is based on the results obtained when the first-difference specification of the conventional log-level short-run money demand equation is estimated for the periods III/1959-IV/1973 and III/1959-IV/1979. The reason for this finding is that the first-difference form of the money demand equation is not sensitive to intercept shifts. In other words, the position that the short-run money demand relationship has deteriorated markedly during the post-1973 period is not substantiated by the evidence.

One important finding of this study lies in pinpointing the key estimation problem as being unforeseen shifts in the intercept term. Our econometric evidence points to the acceptance of the first-difference specification over the log-level form of the relationship because of these shifts. The question of "what caused these intercept shifts?" is partially answered by comparing the full sample results for the M1A and M1B equations.

The most notable difference is the relative stability of the M1A and M1B equations. The former equation, when estimated in log-level form, was found to have had four significant intercept shifts during the full sample period. In contrast, the log-level M1B specification was demonstrated to have shifted only twice during this period. In addition, the equation using M1B as the definition of money has been stable since IV/1975. This finding suggests that the previous estimation problem may exist in failure to capture the substitution of interest earning NOW and ATS accounts, both of which are included in M1B, for conventional demand deposits. Indeed, the relative stability finding of the M1B equation suggests that the action of redefining the transactions aggregates to include interest bearing checkable deposits has been worthwhile.

While we believe the relative stability of the M1B equation documented in this paper is a significant finding, we recognize that the shifts in this relationship that did transpire in 1975 are still left unexplained. Until that time when the cause of these unforeseen shifts is provided, the first-difference specification of the conventionally defined money demand relationship employing M1B is preferred. In the final analysis, however, the case of the missing money is not yet closed.

FOOTNOTES

1/ M1B incorporates NOW and ATS accounts held at thrift institutions, as well as those held at commercial banks. In addition, M1B includes demand deposits at non-bank thrift institutions and credit union share drafts. Because of data insufficiencies and technical difficulties the components added to M1A to calculate M1B are, at this time, not seasonally adjusted. For an expanded description of the new transactions aggregates, see R. W. Hafer, "The New Monetary Aggregates," Federal Reserve Bank of St. Louis Review (February 1980) pp. 25-32.

2/ The lagged adjustment form of equation (1) can be motivated by a lagged adjustment process or, alternatively, by an adaptive expectational adjustment process. With regard to the former, deflating the lagged money stock by the contemporaneous price level permits a lagged response by economic agents to the change in real money holdings resulting from a change in the price level. On this issue, see S. Goldfeld, "The Demand for Money Revisited" Brookings Papers on Economic Activity (3:1973) pp. 577-638; and W. White, "Improving the Demand-for-Money Function in Moderate Inflation," International Monetary Fund Staff Papers (September 1978) pp. 564-607.

3/ The beginning of the sample period, which does not coincide with Goldfeld's, was dictated by data limitations. Data for M1A and M1B are not available prior to 1959. Goldfeld also corrected for first-order serial correlation in his estimations.

4/ Although the Durbin-h statistic is not large enough to allow rejection of the hypothesis of zero first-order autocorrelation in the error terms at the 5 percent level, it is sufficiently close to the 5 percent critical level (1.645) to raise the issue of whether first-order autocorrelation still remains, even after the CORC correction procedure has been employed.

Alternatives to using the CORC estimation procedure were also investigated. One was the Hildreth-Lu (HILU) grid scanning technique. For each of the equations reported in tables 1 and 3, as well as the sub-period estimations required to perform the Chow tests in table 2, the $\hat{\rho}$ chosen by the HILU scanning procedure coincided with that selected by CORC.

Maximum-likelihood estimates were also obtained for each of the seperate equations. This was done by choosing the value of ρ over the (-1, 1) interval which minimized Z, where

$$Z = \frac{T}{2} \ln SSR(\rho) - 1/2 \ln (1-\rho^2)$$

(T equals the number of observations and SSR is the sum of the squared residuals) instead of simply the SSR (ρ). Again, the value of ρ which minimized Z was the same as that obtained by the CORC procedure.

On these various techniques and the problems associated with each, see Henri Theil, Principles of Econometrics (New York: John Wiley and Sons, 1971) pp. 413-14; J. Johnston, Econometric Methods, 2nd ed. (New York: McGraw-Hill, 1972) pp. 262-63; and G. S. Maddala, Econometrics (New York: McGraw-Hill, 1977) pp. 279-84.

5/ S. Goldfeld, "The Case of the Missing Money," Brookings Papers on Economic Activity (3:1976) pp. 683-730.

6/ S. Goldfeld, "The Case of the Missing Money," p. 727.

7/ For an examination of several such studies, see R. W. Hafer and Scott E. Hein, "Evidence on the Temporal Stability of the Demand for Money Relationship in the United States," Federal Reserve Bank of St. Louis, Review (December 1979) pp. 3-14.

8/ For those familiar with the recent money demand literature, it may be surprising to find static rather than dynamic forecasts employed. However, Scott E. Hein, "Dynamic Forecasting and the Demand for Money" (forthcoming, Federal Reserve Bank of St. Louis, Review) shows that this latter technique can provide a biased measure of the shift in a relationship. Consequently, the more widely understood static forecasting procedure is employed throughout this paper.

9/ On the importance of the U^m statistic, Theil notes that "If it [U^m] is large, this means that the average predicted change deviates substantially from the average realized change (relative to the RMS prediction error). This is clearly a serious error; on heuristic grounds we should expect that forecasters must be able to reduce such errors in the course of time." For a complete description of the derivation and interpretation of the Theil inequality coefficients, see H. Theil, Applied Economic Forecasting (Amsterdam: North-Holland Publishing Company, 1966) pp. 27-32.

10/ Application of the Chow test is complicated by the presence of serially dependent disturbances. The Chow test employed in forming the statistics reported in table 4 is based on estimating the respective equations in each subperiod using the CORC iterative technique. This allowed the serial correlation coefficient (rho) to assume different values in the alternative subperiods. An alternative test--one that constrains the serial correlation coefficient to be equal across subperiods--was also used. The findings of this latter approach did not alter the conclusions presented in the text. The likelihood ratio test, more appropriate given the presence of a lagged dependent variables and the serial correlation correction, also provided evidence to reject the null hypothesis of coefficient equality across time. For a discussion of these alternative tests, see Franklin M. Fisher, "Test of Equality Between Sets of Coefficients in Two Linear Regressions: An Expository Note," Econometrica (March 1970) pp. 361-66 and Richard E. Quandt, "The Estimation of the Parameters of a Linear Regression System Obeying Two Separate Regimes," American Statistical Association Journal (December 1958) pp. 873-80.

11/ Estimating the first-difference relationship may also yield desirable econometric properties. On this point, see C. W. J. Granger and P. Newbold,"Spurious Regressions in Econometrics," Journal of Econometrics (June 1974) pp. 111-20; Charles I. Plosser and G. William Schwert, "Money, Income, and Sunspots: Measuring Economic Relationships and the Effects of Differencing," Journal of Monetary Economics (4:1978) pp. 637-60; and D. Williams, "Estimating in Levels or First Differences: A Defense of the Method Used for Certain Demand-for-Money Equations," The Economic Journal (September 1978) pp. 564-68.

12/ The discussion in the text is, essentially, dealing with the exclusion of a relevant variable from the regression. Excluding a relevant variable, in this case the intercept shift term, biases not only the β_0 and β_1 estimates, but also the estimate of the residual variance. On this point see, J. Johnston, Econometric Methods, pp. 168-69.

13/ Equation (2) indicates that a constant term should not be included in the first-differenced money-demand relationship. Including a constant term may be thought of, however, as representing a time trend variable excluded in the original log-level specification. To ascertain whether a trend in money-demand, not explained by the other independent variables, is present, a constant term was included in preliminary regressions of the first-difference equations. The estimated coefficient on the constant term was never significantly different from zero at the 5 percent level, indicating that there is no trend in the original log-levels. Based on these results, the constant term is omitted from the regressions reported. This evidence is different from that of Charles Lieberman, "Structural and Technological change in Money Demand," American Economic Review Papers and Proceeding (May 1979) pp. 324-29. Lieberman found evidence of a significant time trend in the money-demand relationship he estimated.

14/ Interestingly enough, the coefficient on the passbook rate was generally similar in magnitude to that obtained employing the CORC estimation procedure. This suggests that even though nonstationarity in the disturbances has yielded unbiased coefficient estimates on the passbook rate, the coefficient's significance tests in the case of log-level estimation were biased towards rejecting (incorrectly) the null hypothesis. See, C. W. J. Granger and P. Newbold, "Spurious Regressions in Econometrics," Journal of Econometrics (2: 1974) pp. 111-20.

15/ This transformation was made to enable a direct comparison with that of the log-level estimation results detailed previously. The ability to do quite well in out-of-sample simulations that is subsequently detailed for the transformed equation also holds prior to the transformation. For example, the static RMSEs for the untransformed M1, M1A, and M1B relationships over the I/1974 - IV/1979 period are, respectively, 0.0071, 0.0065, 0.0057. These values are not significantly larger than the respective in-sample standard errors of the equation given in table 5.

16/ Since there was some evidence of heteroscedasticity in each of the relationships, we also used a test of structural change developed by S. Gupta ("Testing the Equality between Sets of Coefficients in Two Linear Regressions When the Disturbance Variances are Unequal," unpublished Ph. D. Thesis, 1978, Purdue University) which has been shown to have more power than the standard Chow test. Upon employing this test, however, the null hypothesis of coefficient stability could not be rejected. This conclusion is unaltered by the inclusion of a constant term and/or the commercial bank passbook rate in the relationship.

17/ The evidence of shifts in the relationship provided here is not consistent with the contention that shifts in money demand have been due to cash management innovations caused by high interest rates. This view suggests that the shifts in money demand should coincide with peaks in interest rates e.g., see Richard D. Porter, Thomas D. Simpson, and Eileen Mauskopf, "Financial Innovations and the Monetary Aggregates," Brookings Papers on Economic Activity (1:1979) pp. 213-29.). Market interest rates peaked IV/1969, III/1974, and III/1979, which are points quite different from those suggested by the above analysis. In addition, entering peak interest rate values of the commercial paper rate into the log-level estimations for the III/1959-IV/1979 sample period did not alleviate the satistical deterioration of the equations.

18/ The estimation results employing all of the various intercept dummies are available upon request of the authors.

19/ The HILU and maximum-likelihood techniques again were employed without altering the results reported in the text.

20/ For example, a recent Federal Reserve Board staff study stated that the causes of the deterioration in the function was due to "a shift in the relationship between money and GNP" and the fall in the elasticity of money-demand with respect to interest rates. These conclusions are invalid: properly estimated, the relationship between money, GNP and market interest rates has held up quite well throughout the post-1973 experience. See, David J. Bennett, Flint Brayton, Eileen Mauskopf, Edward K. Offenbacker, and Richard D. Porter, "Econometric Properties of the Redefined Monetary Aggregates," manuscript, Federal Reserve Board (February 1980) p. 16.

21/ On the problems of selecting the appropriate monetary policy instrument, see William Poole, "Optimal Choice of Monetary Policy Instruments in a Simple Stochastic Macro Model," Quarterly Journal of Economics (May 1970) pp. 197-216.

Appendix: Data Definitions and Sources

Money Stock (M1)—currency plus demand deposits at commercial banks (in billions of dollars), seasonally adjusted, quarterly average of monthly figures. Source: Federal Reserve Board

Money Stock (M1A)--currency plus demand deposits exclusive of deposits due to foreign commercial banks and official institutions (in billions of dollars), seasonally adjusted, quarterly average of monthly figures. Source: Federal Reserve Board

Money Stock (M1B)--M1A plus NOW accounts and ATS accounts at commercial banks and other thrift institutions and credit union share draft balances (in billions of dollars), seasonally adjusted, quarterly average of monthly figures. Source: Federal Reserve Board

Income (y)—gross national product in billions of 1972 dollars, seasonally adjusted. Source: U.S. Department of Commerce, Bureau of Economic Analysis

Commercial paper rate (RCP)--4-6 month prime commercial paper rate. Prior to III/1974, average of most representative daily offering . After III/1974 average of midpoint of range of daily dealer closing rates. After November 1, 1979, rate is for 4-month commercial paper. Source: Federal Reserve Bank of New York

Commercial bank passbook rate (RCB)--weighted average of commercial bank passbook rates. Source: Federal Reserve Board of Governors

Price level (P)--implicit gross national product price deflator (1972 = 100). Source: U.S. Department of Commerce, Bureau of Economic Analysis

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TABLE 1
Short-Run Money Demand Regression Results
Log-Level Estimates: III/1959 - IV/1973

Dependent Variable	Coefficient[a]					Summary Statistics[b]	
	Constant	y_t	RCP_t	RCB_t	M_{t-1}/P_t	$\bar{R}^2/SEE\times10^{-3}$	Durbin-h statistic/ rho
OLD M1	-0.767	0.150	-0.015	-0.029	0.761	0.9844	1.63
	(4.49)	(4.07)	(3.52)	(2.29)	(7.89)	3.91	0.464
M1A	-0.784	0.155	-0.017	-0.028	0.739	0.9820	1.36
	(4.54)	(4.15)	(3.74)	(2.15)	(7.35)	4.21	0.445
M1B	-0.781	0.154	-0.017	-0.028	0.742	0.9823	1.36
	(4.51)	(4.12)	(3.72)	(2.14)	(7.37)	4.22	0.442

[a] All variables enter logarithmically and all equations are estimated using the Cochrane-Orcutt iterative technique. The numbers in parentheses are absolute values of t-ratios.

[b] $\bar{R}^2$ is the coefficient of determination corrected for degrees of freedom, SEE is the standard error of the estimated equation, and rho is the Cochrane-Orcutt estimate of the autocorrelation coefficient.

TABLE 2
Post-Sample Static Simulation
Results: Log-Level Equations*
I/1974 - IV/1979

Variable	RMSE	Mean Error	Theil Statistics			
			U	U^m	U^s	U^c
OLD M1	6.22	-5.44	0.0096	0.765	0.129	0.106
M1A	7.36	-6.64	0.0116	0.814	0.128	0.059
M1B	6.23	-5.74	0.0097	0.850	0.075	0.077

*Notes: RMSE is the root-mean-squared error in terms of nominal money balances; the mean error is also in nominal money balances. Both RMSE and mean error are in billions of dollars. U is the Theil inequality coefficient; U^m the bias coefficient, U^s the variance coefficient, and U^c the covariance coefficient. See Henri Theil, Applied Economic Forecasting (Amsterdam; North Holland Publishing Co., 1971), pp. 26-32.

TABLE 3
Short-Run Money Demand Regression Results
Log-Level Estimates: III/1959 - IV/1979

Dependent Variable	Coefficient[a] Constant	y_t	RCP_t	RCB_t	M_{t-1}/P_t	Summary Statistics[b] $\bar{R}^2$/SEEx10^{-3}	Durbin-h statistic/ rho
OLD M1	-0.238	0.035	-0.013	-0.004	1.042	0.9823	0.65
	(3.76)	(3.08)	(4.02)	(0.39)	(45.91)	4.86	0.264
M1A	-0.246	0.037	-0.014	-0.004	1.040	0.9834	0.46
	(4.08)	(3.45)	(4.51)	(0.46)	(51.63)	4.79	0.225
M1B	-0.283	0.044	-0.013	-0.008	1.026	0.9825	0.42
	(4.63)	(4.05)	(4.34)	(0.84)	(42.92)	4.60	0.279

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a All variables enter logarithmically and all equations are estimated using the Cochrane-Orcutt iterative technique. The numbers in parentheses are absolute values of t-ratios.

b $\bar{R}^2$ is the coefficient of determination corrected for degrees of freedom, SEE is the standard error of the estimated equation, and rho is the Cochrane-Orcutt estimate of the autocorrelation coefficient.

TABLE 4
Stability Test Results
Log-Level Equations

Variable	Break-point Tested and Calculated F-values* IV/1969	IV/1973
OLD M1	3.43	2.52
M1A	3.29	2.83
M1B	4.20	3.86

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* Critical F-values for the calculated F-statistics, F (5, 73), are 2.35 at the 5 percent level and 3.28 at the 1 percent level of significance.

TABLE 5
Short-Run Money-Demand Regression Results
First-Difference Estimation: III/1959 - IV/1973

Dependent Variable	Coefficient[a]				Summary Statistic[b]	
	$\dot{y}_t$	$\dot{RCP}_t$	$(\dot{M}_{t-1}/P_t)$	$\bar{R}^2$	SEEx10^{-3}	Durbin-h statistic
OLD M1	0.130 (2.47)	-0.018 (2.97)	0.701 (6.83)	0.51	4.48	0.24
M1A	0.143 (2.55)	-0.020 (3.13)	0.652 (6.20)	0.47	4.80	-0.19
M1B	0.142 (2.53)	-0.020 (3.09)	0.653 (6.18)	0.47	4.81	-0.19

[a] The (·) above each variable is used to represent the fact that all variables enter as log first-differences. The equations are estimated using ordinary least squares. The numbers in parentheses are absolute values of t-ratios.

[b] $\bar{R}^2$ is the coefficient of determination corrected for degrees of freedom and SEE is the standard error of the estimated equation.

TABLE 6
Post-Sample Static Simulation
Results: First-Difference Equations*
I/1974 - IV/1979

Variable	RMSE	Mean Error	Theil Statistics			
			U	U^m	U^s	U^c
OLD M1	2.40	-0.54	0.0037	0.050	0.005	0.930
M1A	2.07	-0.63	0.0033	0.009	0.015	0.903
M1B	1.80	-0.41	0.0028	0.052	0.052	0.890

* Notes: RMSE is the root-mean-squared error in terms of nominal money balances; the mean error is also in nominal money balances. Both RMSE and mean error are in billions of dollars. U is the Theil inequality coefficient; U^m the bias coefficient, U^s the variance coefficient, and U^c the covariance coefficient. See Henri Theil, Applied Economic Forecasting (Amsterdam; North Holland Publishing Co., 1971), pp. 26-32.

TABLE 7
Short-Run Money Demand Regression Results
First-Difference Estimation: III/1959 - IV/1979

Dependent Variable	Coefficient[a] $\dot{y}_t$	$\dot{RCP}_t$	$(\dot{M_{t-1}/P_t})$	Summary Statistic[b] $\bar{R}^2$	$SEEx10^{-3}$	Durbin-h statistic
OLD M1	0.142 (2.77)	-0.016 (2.85)	0.699 (7.99)	0.53	5.43	-0.45
M1A	0.144 (2.86)	-0.019 (3.43)	0.694 (8.10)	0.54	5.35	-1.16
M1B	0.153 (3.15)	-0.018 (3.46)	0.686 (7.94)	0.55	5.06	-0.96

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[a] The (•) above each variable is used to represent the fact that all variables enter as log first-differences. The equations are estimated using ordinary least squares. The numbers in parentheses are absolute values of t-ratios.

[b] $\bar{R}^2$ is the coefficient of determination corrected for degrees of freedom and SEE is the standard error of the estimated equation.

TABLE 8
Stability Test Results
First-Difference Equations

Variable	Break-point Tested and Calculated F-values* IV/1969	IV/1973
OLD M1	1.64	0.10
M1A	1.33	0.22
M1B	1.70	0.33

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* Critical F-values for the calculated F-statistics, F (3, 76), are 2.73 at the 5 percent level and 4.06 at the 1 percent level of significance.

TABLE 9
Short-Run Money-Demand Regression Results
Log-Level Estimates Including Intercept Shift Variables: III/1959 - IV/1979

Dependent Variable	Coefficients[a]										Summary Statistics[b]	
	Y_t	RCP_t	RCB_t	M_{t-1}/P_t	D0'	D2	D3	D4	D5	D6'	$\bar{R}^2$/SEEx10^{-3}	Durbin-h statistic/ rho
OLD M1	0.128	-0.019	-0.026	0.843	-0.674	-0.687	-0.698	-0.716	-0.697		0.9842	0.62
	(5.96)	(5.99)	(2.63)	(16.41)	(6.53)	(6.51)	(6.50)	(6.62)	(6.39)		3.76	0.405
M1A	0.124	-0.020	-0.024	0.848	-0.653	-0.666	-0.678	-0.693	-0.677		0.9818	0.66
	(5.93)	(6.21)	(2.41)	(17.07)	(6.42)	(6.40)	(6.39)	(6.48)	(6.30)		3.95	0.397
M1B	0.136	-0.019	-0.028	0.809	-0.707	-0.721				-0.732	0.9797	0.67
	(6.10)	(5.72)	(2.62)	(15.43)	(6.46)	(6.45)				(6.43)	4.15	0.402

a All variables (except D_i (i=0', ...,6') enter logarithmically and all equations are estimated using the Cochrane-Orcutt iterative technique. The numbers in parentheses are absolute values of t-ratios.

b $\bar{R}^2$ is the coefficient of determination corrected for degrees of freedom, SEE is the standard error of the estimated equation, and rho is the Cochrane-Orcutt estimate of the autocorrelation coefficient.

FIGURE 1

PLOT OF M1 (NOMINAL) STATIC

POST-SAMPLE SIMULATION ERRORS

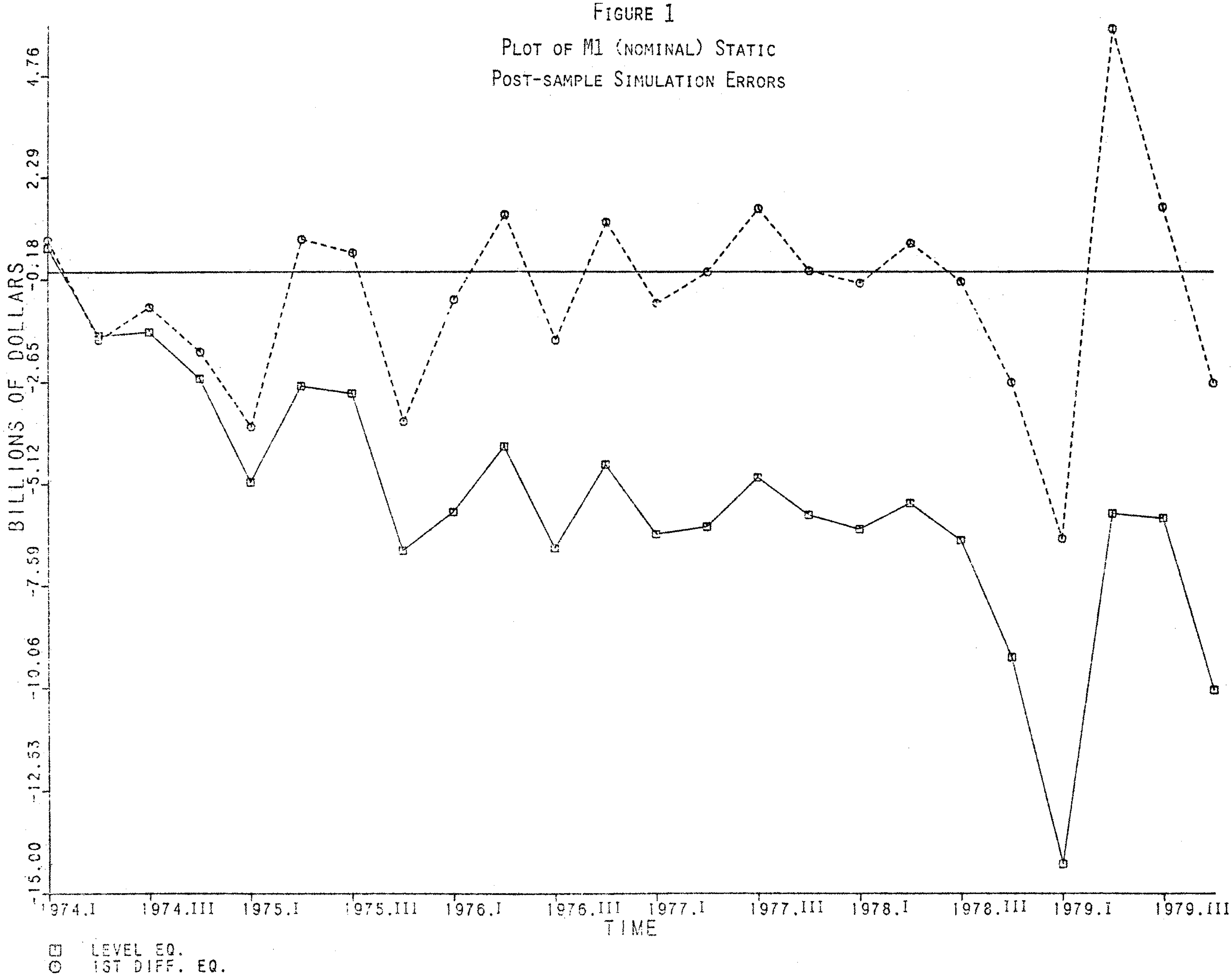

Figure 2

Plot of M1A (nominal) Static

Post-sample Simulation Errors

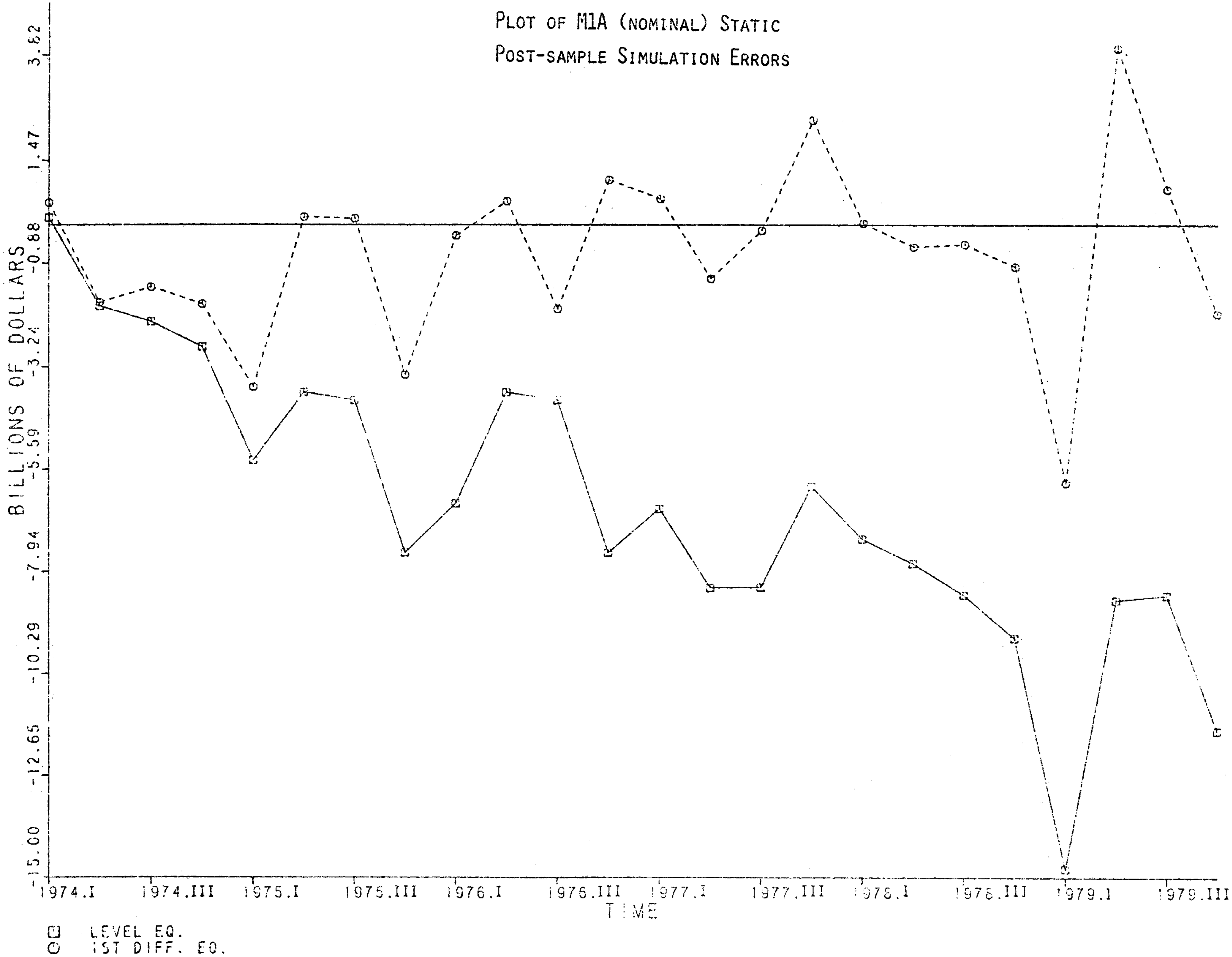

FIGURE 3

PLOT OF M1B (NOMINAL) STATIC

POST-SAMPLE SIMULATION ERRORS