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WP/90/53

INTERNATIONAL MONETARY FUND

Research Department

Long-Run Money Demand in Large Industrial Countries

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June 1990

Abstract

The reputation of the aggregate demand function for money balances has plummeted since the mid-1970s, owing to the destabilizing effects of financial innovation and deregulation. There is, nonetheless, a renewed effort among economists to uncover stable relationships, a revival that reflects in part the development of new econometric approaches, especially those related to cointegration and error correction models. This paper examines the long-run properties of money demand functions in the large industrial countries, under the hypothesis that the long-run functions have been stable but that the dynamic adjustment processes are more complex than those represented in most earlier models. The results do broadly support this hypothesis, but for certain aggregates they also call into question some basic hypotheses about the nature of the demand function, including notably that of homogeneity with respect to the price level.

JEL Classification Numbers:

311, 211, 431

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\* I would like to thank Charles Adams, Michael Bordo, Matt Canzoneri, Bob Flood, Peter Garber, Ben McCallum, Jerome Stein, George Tavlas, Mark Taylor, and participants at seminars at the IMF and at Brown and Rutgers Universities for many helpful comments and suggestions, without implicating them in any remaining errors. The views in this paper are my own and should not be attributed to the International Monetary Fund.

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## I. Introduction

The reputation of the aggregate demand function for money balances has plummeted since the mid-1970s. Once viewed as a pillar of macroeconomic models, it is now widely regarded as one of the weakest stones in the foundation. The origins of this fall from grace are not hard to find: the past two decades have witnessed a large number of financial innovations and deregulatory measures in many countries, which have dismembered traditional payments patterns and have rendered the identification of the line between money and other liquid assets all but impossible. What is remarkable is not that the estimation of functions relating money (conventionally defined) to other macroeconomic variables has become much more difficult, but rather that it remains possible at all.

In spite of these difficulties, a renaissance seems to be in progress, or at least a renewed effort among economists to uncover stable relationships. Part of the basis for this revival has been the development of new econometric approaches, especially those related to cointegration and error correction models. A rapidly growing literature--exemplified by the work of David Hendry and his co-authors--has raised the possibility that models that combine a conventional steady-state function with a complex set of dynamics may be reasonably stable even over periods of substantial institutional change. In a sense allied with this econometric army, though scarcely on speaking terms with it, Milton Friedman and Anna Schwartz (among others) have emphasized that one need not throw out the long run demand function simply because it cannot always predict shorter-run developments. With appropriate allowances, they argue, the qualitative characteristics of the underlying relationships have changed but little over very long stretches of time.

The object of this paper is to examine the nature of the long-run demand for money in the large industrial countries. The maintained hypothesis is that money demand in these countries has remained stable but that the dynamic adjustment processes are more complex than those represented in most earlier models. By "nature of the long run" is meant the parameterization of the steady state associated with time series estimates of equations relating the stock of money to aggregate income, prices, interest rates, and perhaps other variables. The paper begins (Section II) with a selective review of the recent empirical literature, in order to assess what seems to be known or not known on the subject. Section III then examines various methods for estimating the long-run component of money demand equations, and Section IV applies the preferred methodology to data for the five largest industrial countries. Conclusions are summarized in Section V.

## II. The State of the Art

What is the starting point: What do we think we know about the long-run demand for money? There is, of course, a vast literature on this

subject, and no attempt will be made here to review it systematically. <sup>1/</sup> What is of primary interest is the literature that has focused on the effects of financial innovation and deregulation on the nature and stability of money demand. In order to narrow the field a little further, this discussion will focus on two aggregates that have received close attention in the literature: M1 in the United States and in the United Kingdom. <sup>2/</sup>

## 1. United States

The conventional point of departure for looking at M1 in the United States is the classic study by Goldfeld (1976). In that era, the conventional way to estimate the demand for money was to specify a dynamic adjustment comprising a Koyck transformation of a stock adjustment process and a first-order serially correlated error term. The steady-state solution of Goldfeld's "basic" equation of this type, estimated over the 1952-73 period, was

$$(1) \quad m - p = 0.629y - 0.197r^{td} - 0.056r^s,$$

where  $m$ ,  $p$ , and  $y$  are the logarithms of money, real GNP, and the GNP deflator; and  $r^{td}$  and  $r^s$  are the interest rates on time deposits and short-term securities (Treasury bills in this case). <sup>3/</sup> Price homogeneity was imposed on theoretical grounds. The long-run real income elasticity (0.629) is consistent with the Baumol-Tobin proposition of economies of scale in holdings of cash. Both interest rates are intended to represent substitute prices, so the negative coefficients are as expected.

In estimating this equation, Goldfeld was interested in analyzing its inability to predict the sharp rise in velocity that began in 1974. Although it was then too early to judge whether the 1974-76 period constituted an unusually large blip or a more permanent shift, subsequent research generally has confirmed the notion that the demand for money has become more difficult to predict, and that the problems have become much more serious since the early 1980s. What has never been satisfactorily determined is whether the problem relates primarily to the short-run dynamics or whether

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<sup>1/</sup> See Laidler (1985) for a general review.

<sup>2/</sup> For a detailed analysis of the properties of the broader aggregates in these two countries, see Friedman and Schwartz (1982). For cross-country comparisons covering all or most of the major industrial countries, see, inter alia, Atkinson et al. (1984), Boughton (1981), Boughton and Tavlas (1990), Domowitz and Hakkio (1990), Fair (1987), and Leventakis (1990).

<sup>3/</sup> This equation, like all of those presented below, also includes a constant term, which is omitted for simplicity. Except as noted, all equations were estimated with quarterly data.

one must re-examine the shape and stability of the long-run demand function as well. 1/

Interestingly, re-estimation of Goldfeld's equation with an updated sample (1963-88) produces a rather similar steady state: 2/

$$(1') \quad m - p = 0.758y - 0.053r^S.$$

The time deposit rate used by Goldfeld is excluded here, owing to the multiplicity of deposit rates in the latter part of the period. Nonetheless, both the income elasticity and the semi-elasticity of the interest rate are close to earlier values. To some extent, the problems with the equation seem to reside in the dynamics; the adjustment rate--one minus the coefficient on the lagged dependent variable--falls from 18 percent per quarter in Goldfeld's sample to 8 percent in the update. More fundamentally, however, the literature on this issue has called attention to possible specification errors in the conventional approach, including both the Koyck lag and the "correction" for first-order serial correlation.

Jumping ahead to the "modern" era, Michael Darby's buffer stock approach (Darby (1972), Carr and Darby (1981), Carr, Darby, and Thornton (1985)) provides a useful point of departure. 3/ This approach was developed to explain short-run phenomena, but it is of interest to determine whether the nature of the steady state is affected when the adjustment process is modeled more extensively. The U.S. equation that Carr and Darby (1972) estimate for the 1957-76 period (their Table 3) has a very slow adjustment rate (1 percent per quarter, insignificantly different from zero), and a rather high long-run real income elasticity (1.87). The simple

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1/ Major contributions on this issue have been made recently by Bordo and Jonung (1987, 1990) and by Lucas (1988); that literature suggests that the long-run properties of money demand have been stable, but the authors get to that conclusion by different paths. Bordo and Jonung argue that conventional formulations of money demand have undergone secular shifts that are explained in large part by institutional developments. Lucas examined Meltzer's (1963) estimates that were derived from annual data on U.S. M1 over the 1900-57 period, and argued that Meltzer's elasticities do not change when data are added through 1985. That conclusion was based, however, on constrained estimation; Lucas' unconstrained OLS estimates for the 1958-85 period produced radically different elasticities. Whether the constraints are an appropriate way to deal with the different behavior of the data in the two periods remains controversial.

2/ All of the estimates presented in this paper have been made using PC-GIVE; see Hendry (1989).

3/ Another important vein of research (and source of controversy; see Hendry and Ericsson (1989)) has been the work of Friedman and Schwartz (1982 and forthcoming), using annual data from 1870 to 1975 for the United States and the United Kingdom. Their tests, however, used M2 rather than M1 and so are not included in this review.

buffer-stock model thus calls into question the existence of a long-run solution. <sup>1/</sup>

A more general examination of the short-run dynamics was made by Judd and Scadding (1982). Comparing conventional approaches with several buffer-stock models, they found that the best results--not only over the 1959-74 estimation period but also over the 1974-80 post-sample simulation period--were obtained for Coats' (1982) model, which postulates that prices adjust to the excess of monetary growth over the previously expected rate of inflation. They estimated the steady state of that model to have a real income elasticity very close to 1/2, with the usual imposition of unitary long-run price elasticity and with estimated negative responses to short-term interest rates.

The very long adjustment rates in all of these partial-adjustment models raises the issue of whether the relationship between the level of the stock of money and the levels of other macroeconomic variables is characterized by unit roots. Consequently, a number of researchers have examined the cointegration properties of money demand relationships. In this vein, Baba, Hendry, and Starr (1987) estimated a general dynamic equation for U.S. M1 that has a steady state of the following form (slightly simplified):

$$(2) \quad m - p = 0.5y + 0.046r^s - 0.072r^l$$

where  $r^l$  is a long-term interest rate. <sup>2/</sup> As with most earlier models, this equation was estimated subject to the prior constraint--in this case based on initial estimation of a simplified equation--that the long-run price elasticity was unity. The real income coefficient was initially estimated to be close to 0.5, and that value was then imposed as being consistent with the simple version of the Baumol-Tobin model.

The main departure that equation (2) makes from the conventional approach is the inclusion of both short- and long-term interest rates. When the demand function is estimated with all arguments entering contemporaneously, it is very difficult to sort out the effects of different

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<sup>1/</sup> More recent studies of the Carr-Darby approach have confirmed the long-run problems with the model. Estimates by Carr, Darby, and Thornton (1985) using the same 1957-76 data period but a slightly modified equation yielded a negative adjustment coefficient, implying that a steady state did not exist. Boughton and Tavlas (1989) estimated the original Carr-Darby equation over the 1973-85 period, using a modified estimate of unanticipated changes in money, and found an adjustment rate of 4 percent per quarter (still insignificantly different from zero), with a long-run income elasticity around 2.8.

<sup>2/</sup> This equation was estimated by the Engle-Granger two-stage procedure discussed below. The imposed error-correction term was  $m - p - 0.5y$ , and the steady-state interest rate effects emerged in the second stage. Baba, Hendry, and Starr measured interest rates in decimal rather than percentage form; the interest rate coefficients reported in their paper have been divided by 100 for consistency with the other equations reported here.

interest rates, owing to collinearity. In a more general specification, in which short and long rates might, for example, enter with different lags, disentangling their effects could be enhanced. 1/ This appears to be the case with equation (2), where the short rate acts as an own rate with a positive coefficient, and the long rate captures substitution effects. That finding is puzzling, however, since the equation estimated by Baba, Hendry, and Starr also included the yield on NOW accounts as a proxy for the own rate.

Porter, Spindt, and Lindsey (1987) experimented with an error correction model that incorporated a trend, intended to capture various exogenous innovations that help to economize of money holdings. They found the trend to be significant, with a value close to -1 percent per year over the 1961-86 period (their Table 3.3); allowing for this trend, the real income elasticity was estimated to be approximately unity. 2/

## 2. United Kingdom

A convenient place to begin reviewing the modern study of the demand for money in the United Kingdom is the study by Hacche (1974) for the Bank of England, which determined that the narrow aggregate (M1) appeared to be more stable than the broad aggregate (M3). 3/ Hacche used a conventional function similar to that of Goldfeld; the steady state of his estimated M1 equation (1963-72) was approximately

$$(3) \quad m - p = 0.697y - 0.010r^S - 0.029r^L.$$

This equation is completely standard except for the inclusion of both long- and short-term interest rates in collinear form. 4/

Artis and Lewis (1976) presented a number of estimates over almost the same data period (1963-73), experimenting with different functional forms, which suggested that the long-run income elasticity had risen quite sharply in the early 1970s and probably exceeded unity. For example, when they specified the equation in real per capita terms and included a measure of interest rate volatility as an additional argument, they calculated the

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1/ For the United States over the period 1963-88, the correlation coefficient between the quarterly first differences of short- and long-term interest rates (the 6-month commercial paper rate and the yield on long-term government bonds) is 0.73. When one rate is lagged by one quarter, the correlation drops to 0.15.

2/ These estimates were derived from disaggregated equations for currency and demand deposits.

3/ For reviews of earlier work, see Artis and Lewis (1976) and Coghlan (1978). For a criticism of Hacche's methodology and further tests relating to M3, see Hendry and Mizon (1978).

4/ Hacche used  $\log(1 + r)$  instead of the level of the interest rate; the two are quite close ( $\times 100$ ) for low levels of  $r$ .

long-run income elasticity to be 1.35. The same equation truncated at 1971:4 yielded an elasticity of 0.75, which is close to Hacche's estimate.

Artis and Lewis also reported equations in nominal form, which allows the price elasticity to differ from unity but constrains it to equal the real income elasticity. Those equations also produced elasticities that shifted from less than unity to greater than unity with the lengthening of the sample. But when Coghlan (1978) allowed both elasticities to range freely and used unrestricted lags, he estimated the price elasticity to be around 0.7 and the real income elasticity to be quite close to unity. Perhaps the extreme low estimate for the income elasticity was obtained by Cuthbertson and Taylor (1987). They estimated a Carr-Darby buffer stock model for U.K. M1 (1964-81), and found an elasticity of 0.32. In contrast, Hall, Henry, and Wilcox (1989) estimated an error correction model (1963-87), allowed both price and income elasticities to vary, and found both to be close to unity.

Muscatelli (1989) tested an error correction model against forward-looking buffer stock models, and found that the error correction model consistently outperformed the alternatives. His preferred equation (estimated over the 1963-82 period) had a steady state of the form

$$(4) \quad m - p = 0.602y - 0.035r^S,$$

which is very close to Hacche's much earlier estimate (equation (3), above), with the short-term interest rate here capturing the combined effects of short and long rates in Hacche's equation.

### 3. Implications

This sketchy review of recent developments suggests a few conclusions and some open issues to be explored. First, there are important interactions between the specification of the short-run dynamics and the long-run properties of the demand for money. The conventional partial-adjustment model (including the Carr-Darby buffer-stock variant) is an inadequate representation of the former and probably distorts the estimation of the latter. Second, there is near-unanimous agreement that the function is homogeneous in prices in the long run; most researchers have simply imposed that condition, but most of those who have tested for it have failed to reject the hypothesis. Third, there is no consensus on the nature of the long-run income elasticity; estimates range from less than 1/2 to well above unity, both in the United States and the United Kingdom. Fourth, while there is a clear negative relationship between money demand and the level of market interest rates, the relative roles of short and long rates are less evident. Fifth, there is at least preliminary evidence suggesting that the demand for money in the United States has been characterized by a negative trend over the past three decades.



### III. Cointegration Analysis

One of the main sources of confusion regarding the parameterization of the long-run demand function has been the variety of methods employed for obtaining approximately white-noise residuals in time series estimates. The most common method employed for many years was simply to "correct" for first-order serial correlation through a Cochrane-Orcutt or similar transformation. It is now known, however, that this procedure covered up serious specification errors and led to the acceptance of excessive restrictions (see Hendry and Mizon (1978)). During the past ten years, there has emerged a growing acceptance of more complex dynamic structures, combined with prior analysis of static relationships through analysis of the cointegrating properties linking the various arguments in the model; i.e., tests of whether the levels of variables are linked by a stable long-run relationship.

This section examines the implications for long-run money demand of several approaches to cointegration analysis. In each case, the initial step is to examine the stationarity properties of the data; only if the data are (or can be transformed so as to be) stationary in differences is it appropriate to apply these procedures for estimating cointegrating relationships. All of the data discussed here meet the usual tests for difference stationarity. <sup>1/</sup>

#### 1. Two-stage estimation: the Engle-Granger procedure

Engle and Granger (1987) propose a two-step procedure under which a cointegrating vector is estimated first, with the dynamics estimated subject to that steady state. They demonstrate the general consistency of the procedure, but they also emphasize the non-uniqueness of the solution. That is, there is a variety of methods for obtaining the first stage equation--and, in the multivariate case, there may be as many as  $n-1$  valid cointegrating vectors--and there is no clear basis for choosing among them. As alternatives to simple static estimation (as originally proposed by Engle and Granger), one could directly estimate the steady state of a VAR, or one could arbitrarily select one of the vectors estimated by the methodology discussed in the following subsection, or one could impose prior values on the coefficients. Choosing among the results raises issues of both efficiency and identification. In any case, the lagged residuals from the cointegrating vector would be introduced as an argument (the error correction term) in the dynamic equation.

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<sup>1/</sup> One problem is that these tests have low power for rejecting stationarity and may be misleading if the data are not pure autoregressive processes. For a discussion and application, see Schwert (1987). By some tests, price indexes are stationary only in second differences; see, for example, Boughton, Branson, and Muttardy (1989). For this study, however, they are treated as difference-stationary.

Regardless of how the first stage is obtained, the second stage is to estimate

$$(5) \Delta m_t = \alpha + \beta'_1 L(p) + \beta'_2 L(y) + \beta'_3 L(r_s) + \beta'_4 L(r_l) - \gamma EC_{t-1} + \epsilon_t$$

where  $L$  is the lag operator and  $EC_t$  is the time series of residuals from the cointegrating vector. Equation (5) can then be reduced to a parsimonious equation through the elimination of insignificant terms and the imposition of constraints that hold to a reasonable approximation. 1/ If and only if the levels of all variables vanish in this reduction process will the initial estimate of the cointegrating vector be accepted as the steady state (i.e., the long run demand function). Hence the specification of the dynamics cannot be treated as recursive to the specification of the steady state. 2/

The difficulty of applying the Engle-Granger procedure may be illustrated by reference to the M1 equation for the United States. When the cointegrating vector is estimated through a least-squares regression relating the levels of the variables in the model (with  $m$  as the dependent variable), one obtains (1963:1 -1988:4):

$$(6) m = 0.960p + 0.629y + 0.017r^s - 0.006r^l.$$

This static equation has coefficients that, for the most part, approximate their a priori values: the price elasticity is close to unity, the real income elasticity is less than unity, and the coefficient on the long-term

1/ The algorithm used here for this specification process is as follows. First, for each variable, drop the lag (or the current value) with the lowest t-ratio, as long as the ratio is less than unity. Repeat this operation as necessary. Second, eliminate further lags if t-ratios are below the 5 percent level, taking due account of interactions and of the effects on the final specification. If eliminating a variable has a noticeable effect on other coefficients or would qualitatively alter the steady state solution, further testing (including F tests) may be needed to determine if the variable should be retained. Third, test to determine whether two or more lags can be combined to form simple or compound differences. For example, if  $x_t$  has a positive coefficient and  $x_{t-1}$  has a negative coefficient of similar magnitude, test to see if the two can be replaced by  $\Delta_1 x_t$  without significantly raising the standard error of the estimate. At each step, account is taken of the shape of the equation residuals; problems such as instability, autocorrelation, heteroskedasticity, or skewness may reflect a misspecification.

2/ If the constraints in the cointegrating vector are valid, then  $EC_{t-1}$  should capture all of the relationships among the levels of the variables; see Engle and Granger (1987). But if the researcher imposes restrictions that seem to hold approximately (such as price homogeneity), or omits key variables (such as interest rates), or if there are different lags linking the levels, then  $EC$  may not be a sufficient characterization of the steady state. These complications are discussed further, below.

interest rate is quite small but negative. The only oddity is that the coefficient on the short rate is not only positive (which is acceptable, given that a portion of M1 pays interest) but larger than the coefficient on the long rate.

Because of the high degree of autocorrelation in the errors of equation (6), the standard errors of this regression are not meaningful. What is clear, however, is that the coefficient estimates are highly sensitive to the postulated structure of lags linking the variables. Suppose, alternatively, that one starts by estimating a VAR with 4 lags on each detrended variable:

$$(7) \quad \beta_0 L(m) + \beta_1 L(p) + \beta_2 L(y) + \beta_3 L(r_s) + \beta_4 L(r_l) - \alpha = \mu_t.$$

Solving for the static long-run solution gives

$$(8) \quad m = 0.615p + 1.424y + 0.022r^s - 0.030r^l \\ (0.171) \quad (0.336) \quad (0.019) \quad (0.021)$$

Now the interest rate coefficients are in the "right" range, but the price and income elasticities are not. It is noteworthy, however, that the standard errors in equation (8)--which are meaningful because the relevant dynamics have been included in the estimating regression--suggest that one or both of the interest rates should be excluded from the cointegrating equation.

The "best fit" for this type of VAR occurs when the short rate is excluded:

$$(9) \quad m = 0.702p + 1.225y - 0.014r^l \\ (0.165) \quad (0.317) \quad (0.011)$$

Here, the coefficients move somewhat closer to their expected values, but it still seems doubtful that equation (9) is a valid characterization of a long-run money demand relationship, since the interest rate is barely significant and the long-run price elasticity is well under unity.

In addition to this ambiguity, two-step estimation raises the difficulty that the final dynamic process may be restricted relative to what would emerge from a more general regression strategy. 1/ To illustrate this problem for a simple bivariate model, let the cointegrating vector be

$$(10) \quad y_t = \alpha + \beta x_t + \mu_t$$

and the dynamic error correction process be

$$(11) \quad D(y) = D(x) + \delta \mu_{t-1} + \epsilon_t,$$

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1/ For evidence on the difficulties that can arise with application of the two-step methodology, see Banerjee et al. (1986).

where  $D$  denotes a general differencing operator. In principle, these differences can be made general enough to incorporate any adjustment process. Nonetheless, because the specification of the dynamics is necessarily data-based rather than a priori, the specification of equation (11) may well be affected by the ordering of the estimation. In contrast, a single-step approach could be written as

$$(11') \quad D(y) = D(x) + \delta L(y, x) + \epsilon'_t,$$

where  $L$  is a general lag operator on the levels of  $y$  and  $x$ . Depending on the exact specification of the lag distributions in (11) and (11'), the two estimates may well be non-nested, but in any event the final specifications can be compared through encompassing tests, as discussed below.

## 2. Multiple cointegrating vectors: the Johansen procedure

The initial development of cointegration analysis was aimed at bivariate models. In that context, the question is whether a single relationship exists and, if so, how it is specified. In the case of money demand functions, the model is multivariate and there may exist multiple cointegrating vectors linking some or all of the included variables. Johansen (1988) has devised a general procedure for the multivariate case, and the test statistics have been generalized in Johansen and Juselius (1989). For this test, consider a VAR on the detrended logarithms of money, the price level, and real income, plus short- and long-term interest rates (equation (7), above). The null hypothesis is that there are 5 unit roots in this system (no cointegrating vectors). If that hypothesis is rejected, one tests sequentially for additional cointegrating vectors. For the present problem, it is also interesting to examine the coefficients of any significant vectors to determine if they have the signs and order of magnitude that are expected for a long-run money demand relationship.

A complication that arises is that the steady-state demand function, if it exists, may exclude one or both interest rates. In that case, the 5-variable VAR may still be characterized by one or more cointegrating vectors, none of which might have the desired characteristics. Furthermore, there is a strong likelihood of multiple cointegrating vectors, because--in addition to the presumed long-run money demand relationship--the two interest rates are normally linked with each other and possibly with other included variables through related demand functions. The procedure followed here is therefore also to examine sub-systems that exclude, first, one interest rate (generally the short rate, which is less likely to enter the long-run function 1/ and then both interest rates.

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1/ In most cases, bivariate tests indicate that short- and long-term interest rates are cointegrated; the apparent exception is the United Kingdom, where the Dickey-Fuller test just fails at the 90 percent level to reject the null hypothesis of no cointegration. The long-term rates, however, have stationarity properties that come closer to those of the other variables in the system; specifically, the levels of short-term rates have somewhat higher Durbin-Watson statistics.

The results of this exercise, with four lags included in each VAR, are summarized in Table 1. The interpretation of this table may be illustrated by taking the first two rows (M1 for the United States) as an example. For the full 5-variable VAR, there are 3 significant cointegrating vectors (at the 90 percent level or higher), of which the third (i.e., the least highly significant of those 3) comes closest to matching the prior values on the coefficients. This vector is not very satisfying, however, because the price elasticity is low. Consequently, a second test has been conducted without the short-term interest rate. In the second row, 2 of the 4 cointegrating vectors are significant, of which the second looks reasonably as expected. Overall, the results shown in Table 1 do support the hypothesis that these data sets are characterized by error correction representations, with steady states that could be interpreted as conventional money demand relationships. There are, however, a few exceptions. First, for both M2 in Japan and M3 in the United Kingdom the estimated long-run price elasticities are well below unity (0.6 and 0.5, respectively). Second, for M1 in Germany the matrix may be full rank; 1/ in this case, although there is a steady state, there may not be a stable dynamic adjustment process.

One way to employ these results would be to take the vectors in Table 1 as point estimates of the steady state, and incorporate the lagged residuals from those equations as arguments in a dynamic Engle-Granger equation linking changes in money to changes in the other variables. There are, however, a number of difficulties with that procedure. First, as already noted, the key parameters are not always consistent with conventional priors regarding the shape of the long-run demand function. Second, compounding this first problem, it is not always obvious which of perhaps several candidates should be selected as the most relevant cointegrating vector. Third, the estimated steady state may change in the context of a more fully specified model, especially when constraints have been imposed at the initial stage.

To illustrate the problems selecting among Johansen vectors and integrating them as steady states, consider the 4-variable vector listed as the second row of Table 1:

$$(12) \quad m_t = 0.885p_t + 0.895y_t - 0.045r_t^l + \mu_t$$

While these are consistent estimates subject to the constraints under which they were estimated, there is no particular reason to expect them to be consistent estimates of a more general model. For example, application of

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1/ If significance were restricted to the 95 percent rather than the 90 percent level, then only 4 of the 5 vectors would pass the test.

Table 1. Tests for the Existence of Cointegrating Vectors

	Number of Significant Vectors <u>2/</u>	Preferred Vector <u>3/</u>	Price Level	Elasticities			<u>1/</u>
				Real Income	Short Rate	Long Rate	
United States							
M1	3 of 5	3 *	0.760	1.084	0.110	-0.027	
M1	2 of 4	2 *	0.885	0.895		-0.045	
M2	2 of 5	1 ***	0.791	2.013	0.140	-0.153	
M2	2 of 4	2 *	1.091	0.917		-0.016	
Japan							
M1	4 of 5	2 ***	1.198	0.962	0.027	-0.028	
M2	1 of 5	1 ***	0.625	1.423	0.038	-0.079	
Germany							
M1	5 of 5	1 ***	0.976	1.061	0.041	-0.036	
M3	3 of 5	1 ***	0.533	2.123	0.033	-0.027	
M3	2 of 3	1 ***	0.876	1.630			
France							
M1	3 of 5	2 ***	0.862	0.816	0.030	-0.050	
M2	3 of 5	1 ***	1.272	0.821	0.024	-0.060	
United Kingdom							
M1	2 of 5	2 *	0.753	1.709	0.125	-0.154	
M3	2 of 5	1 ***	0.534	3.280	0.090	-0.079	

1/ Semi-elasticities for interest rates.

2/ The first number is the number of vectors for which the trace of the eigenmatrix exceeds the 90 percent significance level, as reported in Table A2 in Johansen and Juselius (1989). The second number is the maximum number (when the matrix is full rank); i.e., the number of variables included in the VAR.

3/ The first vector is the one with the highest eigenvalue (and significance level), and so forth. \*, \*\*, and \*\*\* indicate significance levels of 90, 95, and 99 percent, respectively.

equation (12) as the first stage of an error correction model leads to the following dynamic equation:

$$\begin{aligned}
 (13) \quad \Delta m_t = & - \frac{0.113}{(0.018)} EC_{t-1} + \frac{0.025}{(0.009)} p_{t-1} + \frac{0.045}{(0.017)} y \\
 & + \frac{0.0016}{(0.0008)} r^s_{t-3} + \frac{0.0018}{(0.0010)} r^l_{t-3} - \frac{0.0032}{(0.0006)} \Delta r^s_t - \frac{0.0029}{(0.0013)} \Delta r^l_{t-3} \\
 & - \frac{0.292}{(0.058)} (p_{t-1} - p_{t-4}) - \frac{0.029}{(0.002)} dum80
 \end{aligned}$$

1964:1 - 1988:4  
 $R^2 = 0.62$   
 $DW = 1.80$

where dum80 is a dummy variable representing the temporary imposition of credit controls in the United States in 1980. 1/ The error correction term ( $EC_t = \mu_t$  from equation (12)) is significant, but the levels of  $y$ ,  $p$ ,  $r^s$ , and  $r^l$  are also all significant. Solving for the steady state of equation (13) gives

$$(13') \quad m = 0.669p + 1.293y + 0.014r^s - 0.029r^l.$$

This more general estimate implies price and income elasticities that are rather different from those in equation (12), and it restores the short-term interest rate. This solution is similar to the 5-variable vector at the top of Table 1, which suggests that the additional constraint in the second row of the table is invalid. Thus the estimation of the error correction model may help to choose among multiple cointegrating vectors.

### 3. Single-stage estimation

The final alternative is to estimate the steady state implicitly, through a general-to-specific regression strategy along the lines advanced by David Hendry (e.g., 1987). That is, starting from a VAR like equation (7), one can eliminate the most insignificant lagged elements and reduce the system to a more parsimonious equation in levels and differences, and then solve directly for the steady state of that equation. There are two main differences between this approach and the two-step procedure. First, the error-correction term need not be limited to contemporaneous observations. That is, a cointegrating relation like equation (10) could in effect be

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1/ Dum80 = 1 in 1980:1, -1 in 1980:3, and 0 otherwise.

estimated with non-contemporaneous (lagged or led) data on  $x$  and  $y$ . 1/ Second, because the extraneous elements in the original regression are eliminated and restrictions may be imposed on the coefficients, there may be a substantial gain in degrees of freedom. While neither of these differences is likely to be consequential asymptotically, both may be important in small samples. The extension to non-contemporaneous observations turns out to be especially important for the problem at hand.

Applying the general-to-specific methodology (see footnote 1, page 8) to U.S. M1 and ignoring all changes yields the following relation among levels:

$$(14) \quad m_t = 0.652p_{t+1} + 1.338y_{t+2} + 0.015r_{t-1}^s - 0.030r_{t+1}^l.$$

With this lag structure, the short-term interest rate reappears with a positive coefficient. If the model reduction is valid, equation (14) is preferred to (8) or (9) because of the rise in degrees of freedom. The inference is that the short rate was inappropriately excluded from equation (9) because of extraneous regressors, and from the Johansen estimates because of the imposition of invalid constraints. 2/

One additional contrast between the Johansen procedure and this general VAR methodology is that the latter may give different results depending on the choice of normalization. Equation (14) was estimated with the nominal money stock as the dependent variable. Normalizing on real rather than nominal balances would make little difference as long as the price elasticity was not constrained a priori. Normalizing on prices rather than

1/ Any equation in levels and differences can be transformed into one in which the levels all appear with the same lag, simply by adding the appropriate difference terms subject to constraints on the parameters. To take a simple example, the equation

$$(a) \quad \Delta y = -\beta_1 y_{-2} + \beta_2 x_{-1}$$

may be transformed as

$$(b) \quad \Delta y = -\beta_1 y_{-1} - \beta_1 \Delta y_{-1} + \beta_2 x_{-1}, \text{ or as}$$

$$(b') \quad \Delta y = -\beta_1 EC_{-1} - \beta_1 \Delta y_{-1}, \text{ where } EC = y - \beta_2 x.$$

Imposition of a single lag structure is equivalent to estimating (b) or (b') without the constraint. That procedure is inefficient and may lead the researcher to drop terms that would be (implicitly) significant in (a). In this example, one would be likely to find the coefficient on  $\Delta y_{-1}$  to be insignificant; dropping it, however, would produce an equation that would be encompassed by (a).

2/ It is tempting to try to draw causal inferences from the lag structure in equation (14); it should be kept in mind, however, that the dynamic relationships in the full equation are much more complex.



money, however, could make a more substantial difference; the dynamic adjustment process would be affected, which could in turn lead to a different estimate of the steady state. For U.S. M1, the renormalized levels portion of the equation for the parsimonious VAR with prices as the dependent variable is as follows:

$$(14') \quad m_t = 0.522p_{t+2} + 1.544y_t + 0.021r_t^S - 0.040r_t^L.$$

There are some minor differences between (14) and (14'), notably in that the real income elasticity is higher and the price elasticity lower. Qualitatively, however, a similar picture emerges. <sup>1/</sup>

Yet another variant of the two-stage procedure is to impose some or all of the coefficients of the error correction term a priori. In particular, since all theoretical models of money demand hypothesize long-run homogeneity with respect to prices, it is natural to consider imposing unitary price elasticity. In addition, a number of theoretical models imply constraints on the range of acceptable values for the real income elasticity. Commonly considered possibilities would include a simple velocity model, used by, among others, Hendry and Ericsson (1989) for M2 in the United Kingdom and by Hall, Henry, and Wilcox (1989) for M1 in the United Kingdom:

$$(15a) \quad m - p - y = \mu_t$$

and the Baumol-Tobin model of economies of scale, used by Baba, Hendry, and Starr (1987) for M1 in the United States:

$$(15b) \quad m - p - .5y = \mu_t.$$

It has already been seen that the restrictions in the Baumol-Tobin model are rejected in a general VAR for U.S. M1, notably in that the income elasticity appears to be at least unity. The use of (15a) is therefore certain to dominate (15b) in any comparison on this data set. When the residuals from equation (15a) are used as the error correction term in a VAR like equation (7), the price level is a significant additional determinant with a negative sign, indicating again that the long-run price elasticity is less than unity. For U.S. M1, the steady state of this model is:

$$\begin{array}{cccc} 0.095(m-p-y) = & -0.017p & + 0.002r^S & - 0.003r^L \\ (0.020) & (0.006) & (0.001) & (0.001) \end{array}$$

or 
$$(16) \quad m = 0.825p + y + 0.018r^S - 0.037r^L$$

In addition to these various models and approaches, there are many other functional forms and estimation methods that could be explored.

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<sup>1/</sup> The equation was also estimated with the short-term interest rate as the dependent variable; that formulation did not produce results that could be interpreted as a money demand function.

including the addition of other arguments. Without attempting a comprehensive survey, it may be worth trying to incorporate the effects of the financial innovations that have occurred during the past two or three decades through the addition of a trend. As noted in Section II.1, there is some evidence that money demand in the United States may have been subjected to a negative trend over the postwar period, possibly as a result of innovation or deregulation. When a linear trend ( $t$ ) is added to equation (7), the trend is significantly negative, and the steady state of the final parsimonious specification is as follows:

$$(17) \quad m = 1.761p + 5.004y - 0.041t.$$

The difficulty here is that the inclusion of the trend sharply raises the estimated elasticities on income and prices to implausible levels, and it eliminates the significance of the levels of the interest rates, which enter only in the dynamic process.

#### 4. Evaluation of the various approaches

In view of the variety of methods available for estimating cointegrating vectors, it is necessary to find a means of selecting among them. The various steady-state relationships discussed above for M1 in the United States are collected in Table 2, where it is immediately apparent that the choice of estimation strategy makes a substantial difference for the qualitative interpretation of the steady state relationship. Most estimates that allow for non-homogeneity with respect to the price level produce an elasticity that is significantly less than unity, although that result holds neither in the simple static equation (6) nor in the equation (17) in which a trend is included. Similarly, most but not all estimates generate a real income elasticity that exceeds unity and a positive coefficient on the short-term interest rate. Even the finding of a negative steady-state relationship with the long-term interest rate, which is supported by almost all models, is overturned when the trend is included.

These various models--at least in the final parsimonious form--are almost all nonnested, so an appropriate test for dominance is the procedure developed by Davidson and MacKinnon (1981), in which the predictions from one model are added to another: if the predictions from model 1 are significant in model 2, but those of model 2 are insignificant in model 1, then model 1 is preferred over model 2. <sup>1/</sup>

Tests comparing several models for M1 in the United States and the United Kingdom, and the two basic approaches for the other aggregates, are summarized in Table 3. It would have been a pleasant surprise to find that these tests conclusively favored one approach over the others; more realistically, they do broadly support unconstrained single-stage estimation over

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<sup>1/</sup> See Hendry (1989) for a summary of several other encompassing tests. Those tests are programmed into PC-GIVE, but only for smaller data sets than those used here.

Table 2. Summary of Steady-State Estimates for United States M1 1/

- (1) Goldfeld [1952-73]  
 $m = p + 0.629y - 0.197r^{td} - 0.056r^s,$
- (1') Equation (1) updated [1963-88]  
 $m = p + 0.758y - 0.053r^s.$
- (2) Baba, Hendry, and Starr [1960-84]  
 $m = p + 0.5y + 0.046r^s - 0.072r^l$
- (6) simple static equation  
 $m = 0.960p + 0.629y + 0.017r^s - 0.006r^l$
- (8) static solution to 4-lag VAR  
 $m = 0.615p + 1.424y + 0.022r^s - 0.030r^l$
- (9) steady state of two-stage Engle-Granger estimation (model EG2)  
 $m = 0.702p + 1.225y - 0.014r^l$
- (12) 4-variable Johansen vector  
 $m = 0.885p + 0.895y - 0.045r^l$
- (13') steady state of model JV2 (2-stage estimation with eq. 11 as error-correction term)  
 $m = 0.669p + 1.293y + 0.014r^s - 0.029r^l$
- (14) steady state of model EC1 (single-stage estimation)  
 $m = 0.652p + 1.338y + 0.015r^s - 0.030r^l$
- (14') model EC1 estimated with p as dependent variable  
 $m = 0.522p + 1.544y + 0.021r^s - 0.040r^l$
- (14'') model EC1 with m-p as dependent variable (model EC1c)  
 $m = 0.648p + 1.364y + 0.013r^s - 0.030r^l$
- (16) steady state of model PV2 (2-stage estimation with m-p-y as error-correction term)  
 $m = 0.825p + y + 0.018r^s - 0.037r^l$
- (17) model EC1 with trend included (model EC1b)  
 $m = 1.761p + 5.004y - 0.041t.$
- 

1/ Sample period is 1964-88 except as noted. Equation numbers refer to those in the text, except for (14''), which was not shown explicitly.

Table 3. Encompassing Tests for Non-Nested Models, 1964-88

(F statistics) 1/

Additional Model								2/
EG2	PV2	JV2	EC1	EC1a	EC1b	EC1c		
A. United States (M1)								
Basic Model								
EG2	--	0.55	2.27	2.89*	2.08	2.85*	0.25	
PV2	2.87*	--	3.49*	4.41**	1.47	4.82**	0.00	
JV2	5.26**	4.86**	--	4.20**	4.18**	5.34**	0.55	
EC1	2.09	1.63	0.01	--	-- 3/	2.15	0.03	
EC1a	4.44**	1.48	1.30	3.00*	--	5.04**	0.02	
EC1b	4.65**	4.03**	2.15	4.27**	3.99**	--	0.00	
EC1c	5.45**	4.34**	1.26	3.59*	3.62*	4.79**	--	
B. United Kingdom (M1)								
Basic Model								
EG2	--	0.90	0.51	4.91**				
PV2	3.21	--	2.34	5.75**				
JV2	2.58	3.45*	--	6.27**				
EC1	6.02**	4.72**	4.86**	--				
C. Other Aggregates								
	<u>EG2 vs. EC1</u>			<u>EC1 vs. EG2</u>				
United States								
M2	20.36***			5.45**				
Japan								
M1 4/	10.43***			2.33				
M2	62.73***			0.01				
Germany								
M2	19.76***			6.46**				
M3 5/	12.63***			1.80				
France								
M1	0.84			6.26**				
M2	0.24			7.64***				
United Kingdom								
M3	17.01***			2.59				

1/ Test of the significance of adding the predictions from the "additional" model to the regression of the basic model; see Davidson and MacKinnon (1981). The null hypothesis is that the information from the additional model is already in the basic model. The symbols \*, \*\*, and \*\*\* indicate rejection of the null hypothesis at the .10, .05, and .01 levels, respectively.

2/ EG2 = Engle-Granger 2-stage estimation, with first stage estimated as a 4-lag VAR.

PV2 = Same procedure as EG2 except first stage imposed, with unitary elasticities on prices and real income and both interest rates omitted.

JV2 = Same procedure as PV2 except that long-term interest rate included, and coefficients taken from Table 1.

EC1 = Single-stage general-to-specific modeling, starting from the 4-lag VAR.

EC1a = real income elasticity constrained to unity.

EC1b = trend included.

EC1c = real balances as dependent variable.

3/ For this aggregate, EC1a is nested in EC1.

4/ Estimated in first differences.

5/ Estimated with real balances as the dependent variable.

most alternatives. For M1 in the United States, the conclusion is unambiguous: model EC1 is preferred over all other estimates. For M1 in the United Kingdom, the tests imply that a more general model may be required, since the predictions of model EC1 are significant when added to the other three estimates, but the estimates from the other models are also significant when added to model EC1. Comparisons of single-stage and two-stage estimation for the other eight aggregates indicate that the single-stage estimates are preferred in four cases; the two-stage estimates in two cases (both of the French aggregates); and a more general model in the remaining two. Thus with the exception of France, relatively little seems to be gained--and a possibly heavy cost could be incurred--by two-stage estimation. The sub-models included for the United States in Table 3 are also of interest. Constraining the real income elasticity to unity (model EC1a) has little effect one way or another, <sup>1/</sup> but estimating the model with real rather than nominal money balances as the dependent variable (model EC1c) makes matters decidedly worse. Inclusion of a trend (model EC1b) is either neutral or superior to all other models except for the unconstrained single-stage model.

The clearest conclusion to be drawn from this exercise is that there does exist at least one cointegrating vector linking the variables that are expected to enter the long-run demand function. This conclusion holds for all ten data sets (two aggregates in each of five countries) being examined. There is thus a *prima facie* case for estimating error correction models of the demand for these aggregates. It also is clear that the imposition of prior constraints leads, in most cases, to inferior performance. In spite of the very strong theoretical prior for price homogeneity, the imposition of that constraint can be quite misleading. Whether a single- or two-stage estimation strategy is to be preferred is, however, much less clear.

#### IV. Final Error Correction Estimates

The procedure adopted here for estimating error correction models for each of these aggregates is to pursue at least two approaches--the general single-stage and two-stage strategies discussed above--and then to select between them on the basis of both encompassing tests and plausibility of the parameters. If both models have serious problems, then related models are also estimated and compared.

The results of this approach are summarized in Table 4; the complete dynamic regressions are listed in the Appendix. With two exceptions, one or the other of the two basic models has been successfully estimated for each aggregate. For M3 in Germany, single-stage estimation failed to confirm the existence of an error correction model; the final equation contained only

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<sup>1/</sup> Unconstrained estimation (equation 14) yields a long-run real income elasticity of 1.338, which is close to that in most other estimates (see Table 2). However, an F-test fails to reject (at 95 percent) the hypothesis that the true coefficient is unity.

Table 4. Long-Run Elasticities: Final Estimates

Aggregate	Model <u>1</u> /	Elasticity		Semi-elasticity	
		Price Level	Real Income	Short-term Interest Rate	Long-term Interest Rate
United States					
M1	EC1	0.652	1.338	1.487	-3.013
M2	EC1	1.0	1.0	-0.779	--
Japan					
M1	EC1d	0.831	0.831	--	-2.628
M2	EC1	0.387	1.604	1.777	-5.572
Germany					
M1	EG2	1.0	1.199	-2.447	-1.482
M3	EC1c	0.418 <u>2</u> /	2.240	2.180	-3.183
France					
M1	EG2	1.0	0.434	-1.226	0.960
M2	EG2	0.685	1.965	-1.280	--
United Kingdom					
M1	EC1	1.0	3.332	-10.343	--
M3	EC1	0.699	3.281	--	-3.441

1/ EG2 = Engle-Granger 2-stage estimation, with first stage estimated as a 4-lag VAR.

EC1 = Single-stage general-to-specific modeling, starting from the 4-lag VAR.

EC1c = real balances as dependent variable.

EC1d = estimated in first differences.

2/ Insignificantly different from zero.

differences in the variables, except for the level of the long-term interest rate. Two stage estimation generated an acceptable steady state, but the error correction term was nonstationary (i.e., the hypothesis that the data are not cointegrated could not be rejected), and the dynamic equation was encompassed by other models. The preferred estimate was taken to be that of model EC1c: single-stage estimation starting with real balances as the dependent variable. For M1 in Japan, none of the models produced an acceptable error correction equation, so the basic models were rerun in first-difference form.

The most striking finding from this exercise is that six of the ten aggregates have long-run price elasticities that are significantly less than unity. The reasons for this anomaly are not obvious, but there are two explanations that probably account for at least part of it. The first is the possibility of aggregation bias. For four of the five countries, the estimated price elasticity is much closer to unity for narrow money than for the broad aggregate; for the United States, the opposite holds. It may be, therefore, that excessive (or, in the U.S. case, insufficient) aggregation is introducing errors in the parameter estimates. If, for example, the output elasticities differ across components of the aggregates, constraining them to be equal could bias the estimates.

The second likely explanation for the low price elasticities is the shortness of the data sample (25 years). If the true underlying price elasticity is unity but adjustment of the stock of money to repeated inflationary shocks has been incomplete over the data sample, then the estimated elasticity would be less than unity. To the extent that the central bank has responded to increases in the price level by reducing the money stock or that the price level has responded slowly in response to monetary policies, lags would be increased and the downward bias would be aggravated. In this case, the discrepancy would be greatest for the aggregates that were the primary focus of monetary policy; this story would thus suggest that U.S. monetary policy had been concerned more with M1 during this period, while other countries had focused more on M2. Sorting out the importance of these various factors would require further research. <sup>1/</sup>

There is virtually no evidence here in support of economies of scale in cash holdings. That is, the microeconomic theories pioneered by Baumol and Tobin do not seem to apply to the aggregate data: for eight of the ten aggregates, the real income elasticities are unity or higher.

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<sup>1/</sup> Other factors may also help to explain the lack of homogeneity in the regression estimates. First, the aggregate price indexes may not measure the true prices on which asset-holders decisions are based. Second, there may be inflationary-expectations effects that have not been modeled, such that a rise in the price level would cause the desired stock of real balances to fall. These factors, however, would not explain why the irregularity is present much more for one aggregate than for the other in each country.

In all but one case (M1 in France), the total effect of interest rates (the effect of a combined change in short and long rates) is negative, as expected. In several cases, however, there is a significant term structure relationship. In three cases, short-term interest rates have a positive effect on money holdings in the long run, offset by a somewhat larger negative effect from long-term rates.

Table 5 presents tests of the stability of the regression estimates over the period 1972-1988, which indicate that most of the equations have been stable since at least the mid-1970s. These are N-step Chow tests, which test for sustained or permanent shocks to the relationships. In six of the ten cases, there is a significant shift at some point. In four of those cases, however, the shifts come in the early 1970s, around the time of the first oil shock and the switch from fixed to more flexible exchange rates.

In only two cases--M1 in the United States and M2 in France--is there any evidence of sustained parameter instability after 1973. The U.S. M1 equation has a very large residual in 1986:4, a period when M1 was growing quite rapidly against all expectation, apparently reflecting shifts in yields on included accounts relative to those on accounts excluded from M1. That residual is large enough to destabilize the parameter estimates around that time. The instability in the French M2 equation reflects a sharp increase in the volatility of the data after the end of 1984, which in turn may be attributable to the increased openness and flexibility of the French financial system around that time. <sup>1/</sup> When the equation is estimated through the end of 1984 and is then used to forecast through 1988, the forecasts stay on track but miss some unusually large swings.

## V. Conclusions

This paper has examined the long-run properties of money demand equations for five large industrial countries and has compared the performance of equations specified under various approaches. Examination of earlier estimates supported the hypothesis that the long-run elasticities should be reasonably robust with respect to changes in the estimation period or in the specification of the estimating equation, while the short-run properties tend to be much less stable. The estimates presented here, however, raise questions about the robustness of the long-run properties as well, and they suggest that some of the most commonly accepted restrictions employed in the money demand literature may be inconsistent with the data. These questionable properties include homogeneity with respect to the price level, unitary or less-than-unitary elasticities with respect to real

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<sup>1/</sup> From 1964 through 1984, the mean percentage quarterly change in M2 was 3.01 percent, with a standard deviation of 1.36 percent. For the period 1985-1988, the mean increase was 2.08 percent, with a standard deviation of 2.06 percent. For a review of financial innovation and deregulation in France during the 1980s, see Patat (1987).



Table 5. Stability Tests, 1972-1988 1/

	Date of Most Likely Instability <u>2/</u>	F <u>3/</u>
United States		
M1 <u>4/</u>	1986:4	* 2.55 (1.91)
M2	1972:3	* 2.24 (1.79)
Japan		
M1	1971:2	* 2.34 (1.87)
M2	1985:4	0.63 (1.85)
Germany, Fed. Rep. of		
M1	1988:1	1.88 (2.47)
M3	1974:2	0.69 (1.67)
France		
M1	1973:1	* 2.15 (1.71)
M2	1985:1	* 2.05 (1.77)
United Kingdom		
M1	1983:2	1.47 (1.68)
M3	1972:1	* 2.28 (1.83)

1/ N-step Chow tests for stability of recursive regression estimates. Each regression is first estimated using data only through 1971:4, and an F test is computed against the null hypothesis that the same equation fits the remaining observations (through 1988:4). The test is then repeated sequentially through the end of the sample, with one observation added each time. The test statistic has (N-T, T-k) degrees of freedom, where N is the total number of observations (100), T is the number of observations in the truncated sample, and k is the number of regressors.

2/ Date (T+1) at which the value of the F statistic is maximized.

3/ The 5 percent significance level is given in parentheses; asterisks flag cases where the null hypothesis of stability is rejected at that level.

4/ For this aggregate, the presence of the dummy variable (Dum80) necessitates the use of restricted least squares for the recursive regressions. The dependent variable was redefined as  $\Delta m - \beta \text{Dum80}$ , where  $\beta$  is the full-sample coefficient estimate.

income, and restriction of the set of included interest rates to either short- or long-term rates, to the exclusion of the other.

The results discussed here are robust with respect to a variety of estimation strategies. Notably, whether the dependent variable is nominal money, real balances, or prices is of little consequence. What does matter is whether one imposes prior constraints on the dynamic process or allows it to be driven by the data. Two-stage error-correction modeling, in which the errors from a cointegrating equation are used as an argument in a dynamic adjustment equation, is generally outperformed by a less restricted general-to-specific specification process. Here again, however, the choice has little effect on the long-run elasticities.

### Dynamic Regressions

The following are the regression estimates that underly the steady-state elasticities listed in Table 4. The sample period is 1964:1-1988:4. Each equation includes a constant term (not shown), and the variables are defined as in the text.  $\Delta_i x_{t-j} = x_{t-j} - x_{t-i-j}$ . Coefficients on interest rates are multiplied by 100 for clarity. Heteroskedasticity-corrected standard errors are given in parentheses. The test statistics, in addition to  $R^2$  and Durbin-Watson, are as follows:

$\sigma$  = standard error of the estimate (x100), followed by the standard deviation of the dependent variable in parentheses

$AR_4$  = F-test for 4th-order serial correlation in the residuals

$AR_{1-8}$  = F-test for serial correlation, jointly for orders 1 through 8

$\chi^2$  = Chi-square tests (with 2 degrees of freedom) for non-normality in the residuals.

Tests that reject the null hypothesis at the .05 level are marked with an asterisk. Five of the ten equations show significant departures from the normal distribution of the residuals (either leptokurtosis or skewness), and one of those five also shows significant higher-order residual autocorrelation. All of these cases also show significant parameter shifts (see Table 5), and the residual problems probably reflect those shifts.

### United States

M1 1/

$$\Delta m = -.101m_{-2} + .066p_{-1} + .135y + .150i_{-3} - .304r_{-1}$$

(.016)      (.016)      (.025)      (.059)      (.122)

$$-.028dum80 - .240\Delta_3 p_{-1} - .369\Delta i - .207\Delta_3 r_{-1}$$

(.003)      (.055)      (.053)      (.080)

$$\begin{aligned} R^2 &= 0.63 \\ \sigma &= 0.66 \text{ (1.03)} \\ DW &= 1.99 \\ AR_4 &= 0.25 \\ AR_{1-8} &= 0.82 \\ \chi^2 &= 22.38 * \end{aligned}$$

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1/ Dum80 = 1 in 1980:1, -1 in 1980:3, and 0 otherwise.

M2

$$\begin{aligned}\Delta m = & -.168(m_{-4} - p_{-2} - y) - .131i + .266\Delta p_{-2} \\ & (.023) \quad \quad \quad (.027) \quad (.091) \\ & -.340\Delta i - .165\Delta i_{-2} + .285\Delta r - .543\Delta r_{-1} \\ & (.069) \quad (.065) \quad (.153) \quad (.158)\end{aligned}$$

$$\begin{aligned}R^2 &= 0.66 \\ \sigma &= 0.51 \quad (0.85) \\ DW &= 1.69 \\ AR_4 &= 0.21 \\ AR_{1-8} &= 0.54 \\ \chi^2 &= 0.35\end{aligned}$$

Japan

M1

$$\begin{aligned}\Delta \Delta m = & -.638\Delta m_{-1} + .206\Delta m_{-3} + .360(\Delta p_{-2} + \Delta y_{-4}) \\ & (.068) \quad (.089) \quad (.084) \\ & - 1.137\Delta r_{-3} - .353\Delta_2 \Delta i_{-2} \\ & (.357) \quad (.130)\end{aligned}$$

$$\begin{aligned}R^2 &= 0.42 \\ \sigma &= 1.42 \quad (1.82) \\ DW &= 2.27 \\ AR_4 &= 3.86 \\ AR_{1-8} &= 3.20 \\ \chi^2 &= 94.96 \quad *\end{aligned}$$

M2

$$\begin{aligned}\Delta m = & -.050m_{-4} + .019p_{-4} + .079y_{-3} + .088i_{-4} \\ & (.009) \quad (.009) \quad (.013) \quad (.024) \\ & -.276r_{-1} + .415\Delta m_{-1} - .235\Delta y_{-3} \\ & (.045) \quad (.096) \quad (.054)\end{aligned}$$

$$\begin{aligned}R^2 &= 0.88 \\ \sigma &= 0.42 \quad (1.16) \\ DW &= 2.20 \\ AR_4 &= 0.51 \\ AR_{1-8} &= 0.96 \\ \chi^2 &= 1.10\end{aligned}$$



M2

$$\begin{aligned}\Delta m = & -.132(m - .685p - 1.965y + 1.280i)_{-1} \\ & (.027) \\ & -.187\Delta_2 y + .965(.3r_{-1} - r_{-3} + .7r_{-4}) \\ & (.088) \quad (.563)\end{aligned}$$

$$\begin{aligned}R^2 &= 0.25 \\ \sigma &= 1.34 \quad (1.52) \\ DW &= 1.93 \\ AR_4 &= 8.87 * \\ AR_{1-8} &= 2.92 * \\ \chi^2 &= 29.08 *\end{aligned}$$

United Kingdom

M1

$$\begin{aligned}\Delta m = & -.035(m - p)_{-1} + .116y_{-1} - .359i_{-4} - .560\Delta_4 i \\ & (.015) \quad (.024) \quad (.106) \quad (.072) \\ & + .754\Delta r_{-3} + .161\Delta m_{-2} \\ & (.269) \quad (.093)\end{aligned}$$

$$\begin{aligned}R^2 &= 0.58 \\ \sigma &= 1.43 \quad (2.14) \\ DW &= 2.04 \\ AR_4 &= 0.80 \\ AR_{1-8} &= 1.79 \\ \chi^2 &= 0.53\end{aligned}$$

M3

$$\begin{aligned}\Delta m = & -.072m_{-1} + .051p_{-3} + .237y_{-1} - .249r_{-3} \\ & (.020) \quad (.020) \quad (.057) \quad (.066) \\ & - .485\Delta_2 i_{-2} + .938\Delta r_{-3} + .173\Delta_3 m_{-1} \\ & (.153) \quad (.371) \quad (.052)\end{aligned}$$

$$\begin{aligned}R^2 &= 0.42 \\ \sigma &= 1.63 \quad (2.07) \\ DW &= 1.94 \\ AR_4 &= 1.90 \\ AR_{1-8} &= 0.65 \\ \chi^2 &= 56.16 *\end{aligned}$$

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