Buffer Stock Models of the Demand for Money and The Conduct of Monetary Policy

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Popular wisdom informs us that one of the prime victims of the American financial deregulation of the last decade has been the demand for money function. While there is some evidence that the demise of standard monetary relations has been oversold (see Rasche 1987, 1988, and Darby, Mascaro, and Marlow, 1989), it is with some trepidation that we agreed to reexamine functional forms which we helped develop in 1980.

We use updated U.S. data to replicate the quarterly price equations estimated by Gandolfi and Lothian and by Darby and Stockman in <u>The International Transmission of Inflation</u>. These equations were based on buffer-stock money-demand functions and originally estimated for the U.S. and seven other countries. In this paper we confine ourselves to an examination of the American case. In our new estimations we use two alternative money definitions -- M1A and (new) M2. Estimates not reported here using the current M1 definition (inclusive of Other Checkable Deposits) were uniformly less successful.

Our results can be summarized simply. The buffer-stock emphasis on differential effects of expected and unexpected changes in nominal money are strongly supported by the new data. The simple partial-adjustment model first proposed by Chow (1966) and used by Darby and Stockman can be rejected, but a less restrictive adjustment or error-correction process used by Gandolfi and Lothian does survive the decade unscathed. We tentatively conclude that non-stationary disturbances to the long-run demand for money are present and that demand for money functions must be estimated -- explicitly or implicitly -- in growth-rate form to induce stationarity. However, cointegration tests leave us uncomfortable with this pat answer: There is borderline evidence that residuals of the long-run demand for money do exhibit stationarity although there are very persistent components.

The implications for monetary policy include the continuing value of monetary aggregates as indicators, potential cumulative effects of base drift in monetary targeting, and the long period over which monetary policy takes effect.

<u>I. Empirical Framework</u>

David Laidler (1984) has already admirably summarized most of the theoretical framework underlying work which we have previously done with coauthors -- Carr and Darby (1981), Darby and Stockman (1983) and Gandolfi and Lothian (1983).[1] Here we first outline some difference of emphasis and expression which we believe provide significant background, and then present the basic estimating equations that two of us set out at the beginning of the decade.

The Adjustment Process

The core of the buffer-stock approach is that the money balances of individual economic agents are conceived of as inventories which follow a stochastic path based on the net outcome of transactions generally conducted without explicit consideration of the level of cash balances. Overall plans --or rather strategies -- are formulated so that the expected net change is approximately zero over the planning period. Further, the length of the planning period is rationally chosen along lines pioneered by Baumol (1952) and Tobin (1956) to balance the additional carrying costs for transacting, payroll, billing, and the like. These plans incorporate expectations about levels of income and expenditure, about prearranged asset/debt transactions, about the probability of completing transactions at set or reservation prices, and the level of interest rates. In addition, there is generally the possibility of discretionary transactions or alterations in the size or timing of planned transactions as a means of adjusting cash balances. Such discretionary transactions generally involve additional fixed costs which are well known to lead to "big-S/little-s" inventory rules so that money balances will be adjusted back to an optimal level only if cumulative shocks push them outside potentially wide control bands around an optimal periodic path.

The key element of the buffer-stock approach as we view it is not merely that inventory behavior of this sort can result in deviations of aggregate money holdings from the average amount associated with existing output, prices, and interest rates. Rather it is that such deviations are predictable as an inherent part of the adjustment to a money shock.

A money shock is implemented by the Federal Reserve buying assets at, say, a higher price (lower yield) than incorporated in the expectations of the economic agents. As a result, total nominal balances are higher in aggregate than would be held on average given the levels of interest rates, income, and prices. On this view, a central bank open market operation will result in a much smaller interest rate movement in the short run than would be required to permanently induce agents to hold the current money balances at current income and prices.

Some critics assert that this means that the buffer-stock models thus do little more than introduce a temporary Keynesian liquidity trap with the apparent interest elasticity of the demand for money rising from near minus infinity toward zero as time progresses. On this view, buffer-stock models "merely" pertain to shifts in the IS-LM curves during the adjustment period and not to the ultimate effect. Those of us who have toiled in these vineyards would argue (a) most of the interesting macroeconomics involves the adjustment period and not the long run in which the only (nontrivial) effect is on the price level and, (b) that money-balance adjustment may involve more than asset transactions -- in particular, expenditures for goods and services may be accelerated or retarded as the most cost-effective way to adjust cash balances.[2] This latter portfolio adjustment version of the real balance effect implies distinctly non-Keynesian short-run shifts in the IS curve.

Laidler (1984) contrasts buffer-stock models with, on the one hand, the Keynesian approach which assumes that agents are always on their long-run money-demand functions and that interest rates adjust with output and prices initially sticky and, on the other hand, the real-business-cycles or "neo-Austrian" approach in which people are always on their long-run money-demand functions and prices adjust. Sometimes effective marketing of ideas or automobiles may require a bit of exaggeration, but we do believe that there is a real kernel of truth in Laidler's characterization. Central bank policy is generally formulated as if the quantity of money responds fairly quickly to given interest-rate changes in pretty much the same way whether these interest rate changes stem from open-market policy actions or other LM shifts or from, say, changes in tax policy or inflation expectations. That is, changes in monetary policy are seen to work through -- rather than disturbing -- a stable long-run demand-for-money function. We shall return to this theme in Section IV.

Stability of the Long-Run Demand for Money

Possibly another difference of nuance with Laidler is whether or not economists can write down a stable relation determining the long-run average quantity of money as a function of prices, income, interest rates, and perhaps a few other appropriate variables.[3] We suspect that there are permanent -- or, at least, very long lasting -- shifts in the relationship due to the uneven pace of technological and regulatory progress. We report some surprising, though far from definitive, evidence to the contrary in Section II below, but believe that even if our agnosticism is warranted, the implications are primarily of an econometric and not a policy nature.

If the long-run demand-for-money function has a non-stationary or barely stationary disturbance in the level form, we shall argue that nothing significant is lost since differencing appears to induce stationarity. Both the Carr-Darby and Gandolfi-Lothian functions were introduced primarily as price equations. The former made minimal adjustment to introduce shock-absorber ideas into the then-popular Chow (1966) model of partial adjustment to a stable long-run money demand function. The latter model was much more general since it allows for error processes which may or may not be consistent with partial adjustment to a stable or randomly-walking long-run money demand function.

The Carr-Darby and Darby-Stockman Functional Forms

Following Chow (1966), Carr and Darby (1981) posited a long-run money demand function of a standard form:[4]

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$$(m-p)_t = a + b\ y_{pt} + c\ i_t \quad , \tag{1}$$

where m is the logarithm of the nominal quantity of money, p is the logarithm of the price level, y_p is the logarithm of real permanent income, i is the nominal interest rate, the subscript t signifies the time period and the superscript d signifies demand. A fraction h of any difference between $(m-p)^d_t$ and $(m-p)_{t-1}$ is eliminated (reflected in the prices used to deflate m_t or in the RHS variables) each period except where it supports current transitory transactions (y_{Tt} measured as $\log(y_t/y_{pt})$). Furthermore, a fraction f of any unexpected change in the nominal money supply ($um_t = m_t - m^*_t$) will be held as money besides what would otherwise be associated with the other variables. That is, the Carr-Darby estimating form is:

$$(m-p)_t = ha + hb\ y_{pt} + hc\ i_t + (1-h)\ (m-p)_{t-1} + d\ y_{Tt} + f\ um_t + e_t \quad , \tag{2}$$

where e is the disturbance term.

Darby and Stockman (1983, hereafter DS) found that the Carr-Darby functional form could be usefully generalized by entering um as a distributed lag running from the current quarter. DS proposed this change to allow for a more complicated adjustment pattern than implicit in Chow's original formulation. Another advantage of the distributed lag form is that it does not so severely constrain the money-shock term to a single quarter expectation formation: The expectations reflected in the underlying income-expenditure plans may have been formulated not last quarter but several quarters earlier. Indeed, differently dated expectations may be relevant to different agents. Suppose, for example, that nominal money growth was some constant plus an independent drawing of a normally distributed random variable x_t. Then nominal money expected two quarters ago for current time t is simply $m_t - x_t - x_{t-1}$, and um_t is identically equal to x_t. Thus, if um_t and um_{t-1} entered with coefficients of 1 in the DS form in this case, then current prices, income, and interest rates would depend on money expected two quarters previous and not on current nominal money.

In our empirical work in Section II, we used the DS version of the Carr-

Darby specification to allow for both longer expectational formation periods and for more complicated adjustment processes. In addition, because we believe that quarterly movements in nominal money are primarily exogenous, we followed DS in writing the equation with the ratio of prices to money as the LHS variable:

$$(p-m)_t = ha + hb\ y_{pt} + hc\ i_t + (1-h)(p-m)_{t-1} + d\ y_{Tt} + \sum_{j=0}^{n} f_j\ um_{t-j} + e_t\ , \qquad (3)$$

where the expected signs are reversed from the money-demand form. That is, hb, d, f < 0; hc > 0; and $-1 < (1-h) < 0$.

The Gandolfi-Lothian Functional Form

Gandolfi and Lothian (1983) were concerned that variants of the Chow money demand functions like the Carr-Darby and DS equations would suffer from an upward bias in the coefficient of $(m-p)_{t-1}$ if the disturbances in equation (2) were positively autocorrelated. This autocorrelation seems likely on the assumption that there are permanent shocks to the long-run money demand function such as in equation (1). Accordingly, they specified an estimation equation in which partial adjustment was implicit in a second-order autoregressive process for the residuals.

The basic approach started with equation (1) augmented by a transitory income term. Repeated substitution in a Lucas supply function yielded:

$$(p-m)_t = a + b\ y_{pt} + c\ i_t + \sum_{j=0}^{\infty} [(h^j\ f)/(1+f)\ um_{t-j}] + u_t\ , \qquad (4)$$

where $u_t = \rho_1 u_{t-1} + \rho_2\ u_{t-2} + e_t$. In the absence of an infinite sample, the summation in equation (4) could be truncated at i = n where money shocks reached a negligible effect on current y. Elaborating on ideas of Griliches (1961), Gandolfi and Lothian argued that their specification allowed for potential partial adjustment to a long-run money demand function with a non-stationary disturbance as well as transitory disturbances. Of course, for many if not all purposes it makes little difference whether the underlying process is conceived of as one of partial adjustment of error correction, but we return to that issue in Section III below.

II. Money and Prices and the Demand for Money: Estimation Results

Our initial focus empirically was on the GL and DS price equations. Since these equations performed well in the original eight-country sample for the period 1957 to 1976, we made only minor changes in their functional forms and reestimated them using updated U.S. data. Our replication of the DS equation, however, yielded unsatisfactory results, parameter instability across the two sample periods, with the coefficient on the lagged dependent variable, being particularly poorly behaved. Consistent with the findings described immediately below for the more general GL equation and in the cointegration tests reported in the next section of the paper, we believe that this failure is largely a reflection of the near-random walk behavior of real money balances.[5] Since the GL specification did not suffer from such problems, we concentrate on those results.

The one major difference between this and the earlier analyses is in the use of the newer monetary definitions. In place of old M2, the sum of currency held by the public and commercial bank demand and time deposits, we use the current and much more broadly defined M2 series. Consistent with this use of new M2, we also take account of the interest return on M2 in computing the opportunity cost variables. As an alternative definition of money, we use M1A, M1 as currently defined less other checkable deposits.[6]

The first equation in Table 1 reproduces the original GL estimates from International Transmission; the second is our reestimation with the new data for the original sample period.[7] With the exception of the constant term, which reflects the much higher level of current than old M2, the replication yields results that are highly similar to the original results. The real permanent income elasticity is approximately unity, as before. The coefficients on the money-shock terms follow the same general pattern, with the lag perhaps shortening somewhat and their magnitude diminishing. The error structure of the new equation is approximately the same as the old. Again, a second-order correction is necessary to allow for the observed autocorrelation of the residuals. Moreover, the estimated pattern is very much the same as in the GL study. The first-order coefficient again is well above unity; the sum of the first and second-order coefficients is approximately unity. The inference drawn in the earlier study of two types of shocks affecting the levels relationship -- one-time permanent shocks and transitory shocks with rather long-lived effects -- appears to continue to hold.[8] The implication of this finding that a differenced form of the equation is to be preferred, therefore, also continues to hold.[9] The one disconcerting aspect of these results is the continued failure of the interest-rate term to enter the regression significantly. We return to this issue below.

The next set of estimates is based on data extended through the fourth quarter of 1988. Rather remarkably, given the developments that have occurred in the 1980s, these data yield essentially identical results to those for the earlier data that we have just described. The standard error of estimate in this equation, moreover, is even a touch lower (.0050 versus .0053) than in the earlier one. The only noticeable difference between the two sets of estimates is the shift in the intercept term (5.3461 here versus 5.4490 above).[10] Stability, in the sense of convergence of inflation and monetary growth appears to be maintained.

The remaining GL equations present evidence on two issues -- the effect of price shocks and the form of the opportunity cost variable. In the original GL study, real oil prices, when added contemporaneously and with lags to equation (4), appeared to play some role, but it was difficult determining how much. This remains true when the equation is replicated. Several t values in each regression are in the neighborhood of two, but the rest are below unity. Taken as a group they are significant at the .01 level in the shorter sample, but surprisingly, not significant at the .05 level in the full sample.[11] Also of interest is the fact that inclusion of the oil-price terms does not affect the estimates of the equation's error structure. Failure to include oil does not, therefore, explain the pattern of errors -- or, more specifically, the existence of gradual adjustment to shocks other than monetary shocks as we have measured them.

In the final GL regressions for M2, which are reported in Table 2, we experiment with the opportunity-cost variable. We include lagged as well as contemporaneous values and use four alternative definitions, the rate on 4-to-6 month prime commercial paper, the AAA bond yield and the spread between each and the rate paid on M2. In most instances, the coefficients of contemporaneous and lagged value of the four variables have t values greater than two. Our ignoring lags, therefore, appears to be the reason that we failed to find an effect of opportunity-cost variables in the original specification of the equation. The magnitude of such effects, however, remains very much an open question, given the well-known simultaneity problem that plagues any investigation of the short-run money-interest-rate relationship.[12] The coefficients on the money-shock terms generally decline in absolute value when the lagged interest-rate variables are added and fewer prove statistically significant at conventional levels.

Table 3 reports the results of GL regressions using M1A in place of M2. The problem here is that the income coefficients for both the shorter period and the full period are positive. Given the form in which the regression is run, this implies a negative income elasticity of real money balances, which makes no sense at all. One possible explanation for this result is specification bias caused by some omitted variable. An alternative is that the differencing inherent in the correction for autocorrelation has removed too much information from the data, with the result that the the permanent income variable is in fact something closer to transitory income. In this instance, the positive coefficient would simply be indicative of a short-run Phillips Curve.[13]

Table 4 contains the results of our reestimation of the DS version of the Carr-Darby specification. As pointed out at the beginning of this discussion, these results were unsatisfactory in one very important respect. Implicit in this specification is the constraint of a return to level equilibrium. As already demonstrated, this constraint appears inconsistent with the data. The results reported in Table 4, estimated coefficients on the lagged dependent variable of nearly unity in two instances and greater than unity in one, coupled with an insignificant permanent income coefficient in one of these instances provide further evidence that this constraint is invalid.[14]

III. Cointegration and the Errors in the Money-Price Relationship

To analyze the error structure of the money-price relationship further, we conducted a series of tests for cointegration between the price level and the two alternative definitions of money used above. Before turning to the results of this exercise, let us first say something about the logic underlying the tests.

Two variables, which themselves have to be differenced to be stationary, are said to be cointegrated if there is some linear combination of the two that is stationary.[15] Cointegrated variables, therefore, have the property that even though both may be subject to upward or downward shifts over time, there is a linear combination of the two that is not. In the case of the general level of prices and the level of the stock of money, cointegration implies that given a long enough period of time there will be a convergence of the one to the other. Correspondingly, in the case of their first differences, cointegration translates into eventual convergence of the rate of inflation and the rate of monetary expansion.

The estimation results suggested that the log levels of the two series were not in fact cointegrated, but that their first differences were. The second-order error model that we estimated for M2 suggested the existence of one-time shocks to the levels that had permanent effects. Given such a shock, there would be no tendency for the price level to converge to the level of the nominal stock of money. The average rate of inflation would, however, converge to the average rate of monetary growth. Our object here is to reevaluate this evidence using the tests of cointegration.

To conduct these tests, we followed the two-step procedure outlined in Engle and Granger (1987), first estimating the "equilibrium" or "cointegrating" regression, which in the simplest case took the form:

$$p_t = \alpha_1 + \beta_1 m_t + u_{1t} \quad , \tag{5}$$

where p now represents either the log price level, or its first difference, and m either the nominal stock of money, or its first difference. We also experimented with two alternative specifications, one to allow for the effects of real permanent income, the other to allow for the effects of both real permanent income and our opportunity cost variables:

$$p_t - m_t = \alpha_2 + \beta_2 y_{pt} + u_{2t} \; , \tag{6}$$

and

$$p_t - m_t - y_{pt} = \alpha_3 + \beta_3 r_t + u_{3t} \; . \tag{7}$$

We then used variants of the Dickey-Fuller (1979) test for unit roots, to examine the stationarity of the u_t. The equations underlying these tests were of the two general forms:

$$u_t - u_{t-1} = \lambda_1 u_{t-1} + v_{1t} \; , \tag{8}$$

and

$$u_t - u_{t-1} = \lambda_2 u_{t-1} + \sum_{k=1}^{K} \gamma_k [u_{t-k} - u_{t-k-1}] + v_{2t} \tag{9}$$

Of interest in each instance is the coefficient on the level of the lagged error term. A negative and statistically significant value of the coefficient is a rejection of the hypothesis of a unit root and, hence, of non-stationarity. This, in turn, implies cointegration -- either in the levels or the first differences, as the case may be. The first variant of this test, based on equation (8) is referred to as the DF (Dickey-Fuller) test; the second, based on equation (9), is referred to as the ADF (augmented Dickey-Fuller) test. Additionally, we made use of the Durbin Watson statistic from the cointegrating regression, which provides a further test of cointegration.[16]

Table 5 contains the results of these tests. For the levels, we used all three specifications of the cointegrating regression; for the first differences we used (5) and (6) alone. Consistent with the estimation results reported above, these tests, in the main, provide evidence of cointegration of rates of change. For the levels of the series, this is not generally the case. The exceptions, however, stand out. For M2, we do find some evidence of cointegration after we have allowed for the influence of variables affecting money demand. Taken in conjunction with the estimation results, these findings raise an intriguing set of questions, questions for which, unfortunately, we have only partial answers.

On one level, there is the statistical problem pointed out by Darby (1983) in his study of real exchange rates of distinguishing between an exceedingly slow adjustment process in the levels of the variables and the absence of adjustment. In the instances in which we fail to reject the hypothesis of a unit root (for the levels), the coefficient on the level of the lagged error term in the test regression is generally less than .05, implying a first-order autoregressive coefficient of .95 or greater per quarter. It could be, therefore, that we simply have too few degrees of freedom to differentiate between that and a coefficient of unity.[17]

Alternatively, it could be that failure to reject is a reflection of our ignoring the effects of changes in the arguments of the (conventional) long-run demand for money function. In the one instance in which the test results point fairly strongly to rejection, we have allowed for the effects of both interest rates and real income on the real quantity of money demanded.

We suspect, however, that this is at best only part of the explanation. For one thing, it conflicts with the estimation results presented in the previous section in which we have also made allowance for income and opportunity-cost variables. In addition, it raises the question of what shock-type variables could be responsible for the deviations from equilibrium. Money and price shocks are obvious candidates, but our proxies still left unaccounted-for errors.

A further plausible candidate is financial innovation of the type that has taken place over the past two decades, from the introduction of CDs in the early 1960s to commercial bank MMDAs in the early 1980s. Much of this innovation was the result of the monetary acceleration that began in the late 1960s and continued through the 1970s combined with regulatory restrictions on commercial-bank behavior, principally restrictions on the payment of interest on deposits. In its initial phases, the monetary acceleration led to increases in the costs of holding monetary assets as nominal interest rates on market-traded financial instruments rose and deposit-rate ceilings became binding. This was the period of the "missing money," of "disintermediation" and, not just coincidentally, of greatly increased financial innovation. Then, in the 1980s, as nominal interest rates declined across the board and the regulatory environment changed, rates paid on deposits rose. A situation of rising opportunity costs of holding deposits had given way to one of falling opportunity costs.

In principle, this process could explain the existence of long-lived and serially correlated departures of the price-level from its long-run equilibrium path, departures of the type that would be consistent with the very slow adjustment coefficients implicit in many of the cointegrating regressions. In principle also, however, such innovation could also produce permanent effects, new deposits of differing "moneyness" than the old engendering one-time shifts in the demand for money.[18]

IV. Implications for Monetary Policy

The buffer-stock approach to monetary analysis emphasizes the dynamic relationship among the money stock, interest rates, income, and prices. Despite a general disrepute which has fallen upon monetary relations during the 1980s, we have seen that the GL model has done rather well in the decade since it was formulated: not only are coefficients stable when the data for the 1980s are added, but the standard error slightly declines over the longer period. The DS model does not display the same stability as the sample is extended and appears to be saved primarily by the lagged dependent variable. We believe that this reflects the presence of long-lasting if not permanent disturbances to the long-run demand for money function.

These results complement those reported by Darby, Mascaro, and Marlow (1989) in which St. Louis-style relations relating real GNP growth and inflation to past monetary growth have performed much better than reputed. Unlike that study -- in which M1A seemed at least as good an indicator as M2 -- and Rasche (1988) who found no structural break for M1A, this paper presents evidence that supports a focus on M2 as an indicator of monetary policy. Since the M2 equations performed well in those other studies as well, perhaps the overall conclusion is that M2 remains a useful indicator of the impact of monetary policy upon the U.S. economy.

A second conclusion which is supported by our results is that the adjustment process is a long drawn-out affair of the sort that the buffer-stock literature has long emphasized. That is, the relationship among money, income, and interest rates now depends upon money shocks which occurred two or more years earlier. Slow money growth in one year will be largely offset by induced rapid velocity growth, but velocity will grow more slowly the next several years as a predictable part of the adjustment process.

Base drift (basing monetary growth targets on the fourth quarter of the preceding year rather than the midpoint of the previous target) combined with gradual adjustment of growth targets raises concerns in the buffer-stock framework. Indeed, it is well known that base drift combined with leaning against interest rate movements can lead to cumulative deflationary or inflationary spirals. Base drift might be justified if the effects of monetary policy were felt within the year, but that is clearly not the case on these estimates. Perhaps the ultimate refuge for base drift is an appeal to superior information -- unavailable to the models estimated here -- as to what is a permanent shock to the long-run demand for money. By accommodating those shocks, it is argued, the long-run stability of the price level is enhanced. Even this argument may fail, if the provocative cointegration results reported in Table 5 for the levels of M2 hold up in future research. In that case, any base drift ultimately shows up in the price level and increases the prediction error for the price level relevant for long-term contracts.

V. Summary and Conclusions

We have examined the stability of two buffer-stock models of price adjustment in response to monetary shocks over the 1980s. Like others, we find that the DS model which relies on a lagged dependent variable is not stable over the last decade despite an apparent ability to fit the data ex post. In contrast, the GL model which uses an explicit error process instead of the lagged dependent variable is stable both as to coefficients and explanatory power. This success is in sharp contrast to conventional wisdom that money-demand functions have totally collapsed in the 1980s.

We believe that the success of the GL model relates to its ability to handle both permanent, or very persistent, shocks and transitory shocks. Although the model is not inconsistent with nonstationarity in the disturbances to the long-run demand for money, a separate investigation of the cointegration properties of M2 and money demand variables suggests the possibility that these disturbances display very persistent shocks which are eventually eliminated. The Federal Reserve System's current study of the long-run indicator properties of $P^* = M2/y$ will doubtless contribute to our understanding of these issues.

Table 5

Tests for Cointegration between the Logarithms of Money and the Price Level in the United States, 1959:I-1988:IV

Cointegrating Regression	DW	DF	ADF
ln(M1A)			
p vs. m	0.03	-1.213	-2.165
p-m vs. y_p	0.02	-1.229	-2.091
p+y-m vs. $i_{ct\ -3}$	0.12	-1.235	-1.000
ln(M2)			
p vs. m	0.02	-1.883	-2.469
p-m vs. y_p	0.05	-1.396	-2.963
p+y-m vs. (i_c-i_m)	0.35	-3.996	-3.083
Δln(M1A)			
Δp vs.Δm	0.47	-4.134	-2.416
Δ(p-m) vs. Δy_p	1.05	-6.258	-3.468
Δln(M2)			
Δp vs. Δm	0.50	-4.236	-2.541
Δ(p-m) vs. Δy_p	0.78	-5.160	-3.931

Source: See text.

Notes: DW is the Durbin Watson statistic for the cointegrating regression, DF is the Dickey-Fuller statistic and ADF the augmented Dickey-Fuller test statistic. For the log levels we used four lagged values of the differences in the errors in the ADF test regression; for the log differences, we used two. The cointegrating regressions are equations (5), (6) and (7) in the text. Critical values for the tests of the hypothesis of no cointegration are:

	.01	.05	.10
DW	0.51	0.39	0.32
DF	-4.07	-3.37	-3.03
ADF	-3.73	-3.17	-2.91

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NOTES

1. The Carr-Darby model was derived from Darby (1972), tested successfully in Laidler (1980) and elsewhere, extended in Darby and Stockman (1983), criticized by MacKinnon and Milbourne (1984), and defended in Carr, Darby, and Thornton (1985). Laidler (1985) places these basic approaches in the broader stream of empirical analysis of buffer-stock models.

2. For evidence on long-run relationships, see Lothian's (1985) study using cross-country data for 20 industrial countries in the post-WWII period and Friedman and Schwartz's (1982) study of the United States and the United Kingdom for the period 1873-1975.

3. See Bordo and Jonung (1987, 1990 forthcoming) for cross-country evidence on the general role of institutional factors in the demand for money and Lothian (1976) on the more specific question of financial innovation and the stability of money-demand functions in an environments of high and variable inflation and regulatory restrictions on bank behavior.

4. Mc Callum and Goodfriend (1987) refer to equations like equation (1) as a "portfolio-balance relationship". They derive such a relationship from a "proper" money demand function, that they derive in turn from a model of household behavior in which agents hold real money balances to economize on shopping time. Like Goodfriend earlier (1985), however, they question the slow estimated speeds of adjustment obtained in empirical models based on (1) and fitted to short-run data, pointing to measurement error as a likely source of (downward) bias.

5. Darby and Stockman estimated these equations in the context of an eight-country simultaneous model. At the time, computer limitations precluded their using more general (ARIMA) processes to estimate the error structures of the individual equations in their model.

6. Following Darby, Mascaro and Marlow (1989), we adjusted the M1A series for the introduction of interest-bearing chevcking accounts in 1982:I. A description of this adjustment and the sources and methods used in deriving the other data are available in a separate appendix from the authors.

7. For M2 we replicated the expected money equation used by Gandolfi and Lothian (1983); for M1A we used a simple autoregressive model. In future work, we plan to experiment with expected money equations based on the reaction functions developed in Darby and Lothian (1989).

8. Shifts of this sort have been uncovered in other studies of the demand for money. Gandolfi and Lothian (1976), for example, found that individual yearly cross-state money demand functions were subject to intercept shifts over the period 1929-1968. Friedman and Schwartz (1982) found it necessary to use dummy variables to

allow for the influence of war-related uncertainty on the demand for money in the United States in the WWII period and its immediate aftermath. Also in this regard see the series of studies by Gould and Nelson (1974) and Gould, Nelson, Miller and Upton (1978).

9. As is well-known, correction for auto correlation is equivalent to a series of quasi-differencing operations applied to all of the variables in the equation, which in the case of a second-order correction, we can write for any variable x as

$$x_t - \rho_1 x_{t-1} - \rho_2 x_{t-2.}$$

As in GL, we obtain estimates of ρ_1 and ρ_2 such that $\rho_1 \approx 1 - \rho_2$. In the case of an exact relationship, the second order correction would reduce to:

$$x_t - (1-\rho_2) - \rho_2 x_{t-2} ,$$

or equivalently, to a first-order correction of the first differences of x,

$$(x_t - x_{t-1}) - \rho_2 (x_{t-1} - x_{t-2.}).$$

Rasche (1988) and Darby, Mascaro and Marlow (1989) are recent studies that have found first-differenced formulations to work well empirically. Also see the references cited in Gandolfi and Lothian (1983).

10. We also used analysis of variance to examine the temporal stability of this relationship. We can reject the hypothesis that the separate regressions for the subperiods ending in 1976:IV and beginning in 1977:I are not homogeneous at the .01 level.

11. Consistent with the findings reported in The International Transmission of Inflation, these effects also remain relatively small in magnitude.

12. Cagan (1965), in his analysis of the cyclical role of money, contains one of the best discussions that we know of of the problems inherent in attempting to separate effects of this sort. Also see Sims (1980).

13. One potential source of omitted variable bias is failure to adequately account for the effects of financial innovation. The introduction of interest-yielding checkable deposits in 1982 is an obvious candidate, since there is a sizeable and abrupt shift in the level of the M1A series between fourth-quarter 1981 and first-quarter 1982. We have made some allowance for this shift, however, as stated in note 5 above. In addition, have we experimented with dummy variables as an alternative adjustment procedure. This yielded virtually identical results to those reported in the table.

The alternative explanation, that differencing is responsible for these results, while plausible, does not explain the much better performance of the M2 regressions.

14. Boughton and Tavlas (1990, forthcoming) report more satisfactory results using the CD specification both for the United States and a number of foreign countries. We are inclined to attribute the difference to their much shorter sample period (1973:I-1985:IV), in which case the problem of temporal instability that we have pointed to remains.

15. See the discussion of cointegration in Engle and Granger (1987).

16. See Engle and Granger (1987) for significance levels for both the Durbin Watson and the t statistics.

17. The recent exchange-rate literature provides an interesting parallel: Whether one rejects or fails to reject the hypothesis of stationarity of the real exchange rate (or of the purchasing-power-parity relationship) appears to be highly dependent on the length of the sample period. Autoregressive models of the errors estimated for long and short periods are often very similar. The ability to distinguish between very slow adjustment processes for the levels and no adjustment at all, as Frankel (1986) and Huizinga (1986) have conjectured, appears to be a question of underlying degrees of freedom. Results reported in Enders (1989) and Lothian (1989), which show stationarity (in some instances trend-stationarity) of real exchange rates in long-term time series, are consistent with this hypothesis. For a general discussion of the difficulties involved in testing for unit roots see Cochrane (1988).

18. There may well be effects other than those arising purely from changes in the opportunity cost of holding real balances. Some of these innovations involve changes in the characteristics of deposits. In principle, this could alter elasticities and otherwise affect the demand for money function. See Lothian (1976) for a discussion of these issues.