CHARLES LIEBERMAN*

The Long-Run and Short-Run Demand for Money, Revisited

1. INTRODUCTION

THE MONEY MARKET plays a critical role in virtually all theories of income determination and is, as a result, one of the keystones of econometric modeling. Because of its central role, and to obtain unbiased and useful forecasts, considerable effort has been expended to specify and test the competing theories of money demand. Thus, much of the literature of the 1960s and early 1970s focused on whether the evidence favored the transactions or asset (or utility) theories of money demand. (The asset (utility) model appears in [16, 35]; the inventory-theoretic (transactions) approach appears in [1, 37]. Extensive reviews of the literature are available in [29, 3].) Papers by Chow [9], M. Friedman [15], Laidler [30, 31], and Meltzer [35] stand out as classics in theoretical and empirical work on money demand and are buttressed by an enormous supporting cast of papers and books that also suggest that money is demanded because it provides utility and is held for asset motives in contrast to the transactions motives stressed by the Keynesian model. The neoclassical resurgence that emphasizes monetary aggregates as opposed to money-market conditions has been fostered, in part, by these findings.

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More recently, the role of wealth in the money-demand equation has also become central to studies concerned with whether fiscal policy produces portfolio effects that crowd out or crowd in investment. Recent papers [2, 14] suggest that portfolio crowding out may be substantial if wealth is an important determinant of money demand. If, on the other hand, wealth does not belong in the money-demand equation, the portfolio effects generated by bond-financed fiscal policy may actually crowd in investment.

Whereas the earlier empirical literature strongly favored wealth over income in the money-demand equation, more recent studies [20, 33] suggest the opposite conclusion. Thus, a reexamination of the literature would appear to be appropriate to reconcile the inconsistent empirical findings. This paper reports the findings of a reexamination of one of these major papers [9] to cast light on whether money demand is transactions or utility based.\footnote{I wish to thank Gregory C. Chow for graciously providing me with a complete copy of his data set.} Although this review employs Chow's data and reexamines his contribution to the literature, the findings reported are of general concern for empirical work on the demand for money. This review indicates that the standard empirical findings contain two errors.

First, as is well known, but often overlooked in empirical research, money-demand equations tend to be subject to serial correlation. It is recognized that equations that are subject to serial correlation may provide misleading results, but too little effort has been expended in the money-demand area in an attempt to avoid the possible problems that may result.\footnote{Coefficient estimates of equations subject to serial correlation are theoretically unbiased but inefficient. The presence of serial correlation was noted as a problem in the middle 1960s by Courchene and Shapiro [10] but few researchers attempted to correct or adjust their equations for serial correlation. Granger and Newbold [21] argue further that the presence of serial correlation indicates that the estimated equation is misspecified and, as a result, the coefficient estimates are unreliable.} It seems desirable to review the basic money-demand formulation to reestimate these equations and adjust for serial correlation. Quite a few of Chow's results are changed markedly after his equations are adjusted for serial correlation.

A second problem, a lack of stability in estimated coefficients, is suggested by several studies [6, 28, 29, 31, 20]. Laidler [28, 29, 31] and Cagan and Schwartz [6], among others, have noted that the estimated income or interest-rate elasticities are sensitive to the period over which the equations are estimated. Even greater discrepancies from the earlier findings are reported by Goldfeld [20] and Lieberman [33], who find that the transactions model of money demand outperforms the asset model over the post–World War II period, contrary to the findings of the earlier writers. And the accumulating body of evidence including [19, 32] regarding the period 1974–76 and 1978–79 suggests that wide-ranging institutional and technological changes in the money markets make extrapolations of even the best-fitting money-demand equations hazardous.

Evidence presented here suggests that one possible cause of the volatility of coefficient estimates is the many institutional changes that occurred in 1933 and that may have resulted in structural changes in the money-demand equation. If this is true, the money-stock series is not meaningfully consistent for periods before and after 1934. Long-run studies that combine data over this breakpoint may therefore...
produce misleading results. One source of structural change, the prohibition by law of the payment of interest on demand deposits, would tend to cause studies of $M_1$ demand using pre-1934 data to provide results more consistent with post-1934 $M_2$ or $M_3$ relationships than later $M_1$ equations. Moreover, equations that begin with the 1934 data and adjust for serial correlation will be shown to be consistent with the recent postwar studies [20, 33]. Thus, when properly estimated, Chow’s formulation and data support the Keynesian transactions theory of money demand rather than the asset theory, as had been previously believed. Moreover, these findings favor the efficacy of fiscal policy and the crowding in of investment from debt-financed expenditures.

2. THE LONG-RUN DEMAND FOR MONEY, REVISITED

This review of the standard money-demand formulations employs the same annual data and time period (1897–1958) employed by Chow [9] in his well-known demand-for-money study. Chow’s data were employed since his findings are consistent with much of the work performed in this area and because of Chow’s attempts at reproducing or refining the work of Meltzer [35], Friedman [5], and Brunner and Meltzer [5]. Each of Chow’s reported regressions were replicated to verify that the data and the findings proved, in fact, common to both studies.

The regression package employed, a version of TSP, reported all of the usual regression results including the Durbin-Watson statistics. Virtually all of the equations estimated by Chow were found to be subject to serial correlation. Some Durbin-Watson statistics were as low as 0.44, indicating a serious problem with serial correlation. This problem is nevertheless not unique to Chow’s paper. Many other money-demand studies also report no Durbin-Watson statistics even though it is well known that serial correlation is commonly found in money-demand equations. Other money-demand studies that do report Durbin-Watson statistics generally reveal serially correlated errors but the researchers rarely attempt to cope with the problem, perhaps because of the nature of the consequences of the autocorrelated error terms.

Econometric theory indicates that serial correlation in well-specified equations leads to coefficient estimates that are inefficient but unbiased. Thus, the expected value of the estimated coefficient is the true coefficient. Even so, coefficient estimates could deviate substantially from their true values, whereas the standard errors, which are underestimated, will suggest much better fits than is, in fact, the case (see [26, chap. 10]).

Large samples, such as those employed in [9, 30, 31, 35, 15], and the unbiased properties of the coefficient estimates in equations with serially correlated errors suggest the possibility that the findings of these papers might be accepted with some confidence. Unfortunately, the degree of inefficiency is also related to the magnitude of the autocorrelation of the errors and the variables (see [26, 21]).

Granger and Newbold [21] take the much stronger position that any equation subject to serial correlation is misspecified and, as a result, that the coefficient estimates are unreliable.
Many of the variables employed in these analyses are known to be highly autocorrelated and, as already suggested, the degree of autocorrelation of the errors is substantial. In fact, many of the coefficient estimates are affected radically when the equations are reestimated adjusting for serial correlation, despite the large number of observations employed in the analysis.

Table 1 reports the changes in Chow’s regression estimates brought about by adjusting for serial correlation using the Cochrane-Orcutt technique. Since the results of adjusting the estimated equations with the Hildreth-Liu scanning technique were virtually identical to those produced by the Cochrane-Orcutt method, these findings are not reported, and since the equations, when adjusted for first-order serial correlation, yielded residuals that were serially uncorrelated, second- and higher-order adjustments are unnecessary.

<table>
<thead>
<tr>
<th>Equations</th>
<th>k</th>
<th>$Y_p$</th>
<th>$Y$</th>
<th>bond</th>
<th>$R^2$</th>
<th>S.E.</th>
<th>D.W.</th>
<th>p</th>
</tr>
</thead>
<tbody>
<tr>
<td>A</td>
<td>0.66</td>
<td>1.097</td>
<td>-0.729</td>
<td>0.992</td>
<td>0.063</td>
<td>0.59</td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td>(2.91)</td>
<td>(62.74)</td>
<td>(-15.92)</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>B</td>
<td>3.32</td>
<td>0.718</td>
<td>-0.388</td>
<td>0.507</td>
<td>0.039</td>
<td>1.64</td>
<td>0.94</td>
<td></td>
</tr>
<tr>
<td></td>
<td>(2.04)</td>
<td>(5.17)</td>
<td>(-4.33)</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>C</td>
<td>-0.15</td>
<td>0.983</td>
<td>-0.545</td>
<td>0.983</td>
<td>0.090</td>
<td>0.71</td>
<td></td>
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</tr>
<tr>
<td></td>
<td>(0.49)</td>
<td>(43.61)</td>
<td>(-8.08)</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>D</td>
<td>8.40</td>
<td>0.273</td>
<td>-0.321</td>
<td>0.409</td>
<td>0.040</td>
<td>1.67</td>
<td>0.96</td>
<td></td>
</tr>
<tr>
<td></td>
<td>(8.36)</td>
<td>(3.56)</td>
<td>(-3.37)</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>E</td>
<td>-0.65</td>
<td>1.051</td>
<td>0.042</td>
<td>-0.720</td>
<td>0.992</td>
<td>0.064</td>
<td>0.59</td>
<td></td>
</tr>
<tr>
<td></td>
<td>(2.78)</td>
<td>(7.23)</td>
<td>(0.32)</td>
<td>(-13.48)</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>F</td>
<td>3.48</td>
<td>0.583</td>
<td>0.111</td>
<td>-0.357</td>
<td>0.525</td>
<td>0.039</td>
<td>1.64</td>
<td>0.94</td>
</tr>
<tr>
<td></td>
<td>(2.14)</td>
<td>(2.79)</td>
<td>(1.00)</td>
<td>(-3.75)</td>
<td></td>
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</tr>
</tbody>
</table>

Although Chow estimated linear and log-linear equations, the equations reported here are exclusively log-linear since the log-linear formulation has emerged as the preferred specification. Chow’s use of nominal variables instead of real magnitudes is inappropriate, however, unless the elasticities of money demand with respect to the scale variables are unity. If the elasticities are unity, the misspecification of employing nominal instead of real variables produces no bias. Since this condition is not met, as will be demonstrated shortly, only the properly specified real-magnitude equations are reported here.

The first two equations of Table 1 employ permanent income as the scale variable. Equation 1A is unadjusted for serial correlation and confirms the unitary permanent income elasticity reported by Chow as well as a large interest elasticity.

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4Permanent income that is estimated as a long distributed lag of past and present income will clearly be highly autocorrelated.

5Since one central focus of this paper is to eliminate the serial correlation problem that plagues money-demand estimation to obtain unbiased efficient coefficient estimates, the time-series techniques in [4, 36] were employed to estimate the autocorrelations in the data to determine the order of the serial-correlation problem. These results also indicated that the serial correlation was first order.

6Linear equations were also estimated; these also proved subject to serial correlation and the coefficient estimates were also sensitive to the adjustment for autocorrelation.

7Chow did, in fact, test for this necessary condition, but his findings were misleading due to the presence of serial correlation.
This equation is, however, subject to serial correlation as indicated by a Durbin-Watson statistic of 0.59.

When equation 1A is reestimated adjusting for serial correlation, the coefficient estimates change substantially. Equation 1B suggests a permanent-income elasticity significantly below unity and an interest-rate elasticity that is halved in value. The fit, as measured by the \( R^2 \), is reduced from 0.992 to 0.507 and the standard error is also reduced.\(^8\) Thus, these results reveal that previous studies such as Chow’s overstated the magnitude of both the permanent income and the interest rate effects.

Equations 1C and 1D report the regression results of equations that employ current income in place of permanent income. Equation 1C is unadjusted for serial correlation, and once again the problem of serial correlation in the unadjusted equations is evidenced by a low Durbin-Watson statistic of 0.71. Estimated coefficients are also affected substantially when the equation is reestimated adjusting for serial correlation. The adjusted equation reveals a much decreased income elasticity that is significantly below unity. Once again, the interest elasticity is roughly halved, from \(-0.545\) to \(-0.321\). As with permanent income, the \( R^2 \) statistics are reduced by adjusting for serial correlation and the standard errors are also reduced.

When current income and permanent income are included in the same equation, as in equation 1E, the permanent income elasticity is significant and near unity and income is insignificant, as first reported by Chow. Nevertheless, the equation is also subject to serial correlation, as indicated by a Durbin-Watson of 0.59. When the equation is adjusted for serial correlation, as in equation 1F, income remains insignificant and permanent income retains its significance, although once again the magnitude of the effect is approximately halved.

It should be noted that not all of Chow’s reported regressions exhibit such significant changes when they are adjusted for serial correlation. Equations that employ wealth as a scale variable are affected insignificantly either by replacing nominal values with real values or by adjustments for serial correlation. Table 2 reports four such equations that are not markedly affected by adjustments for serial correlation although, as will be shown, they are biased due to the poor quality of the data.

The first two equations of Table 2 employ wealth as the only scale variable. The coefficients of equation 2A are very similar to Chow’s equation 1.5, which was also estimated linear in logs with nominal instead of real magnitudes. Adjusting for serial correlation changes these results only minimally, as noted by equation 2B. When income is added to these equations and the equations are adjusted for serial correlation, the coefficient estimates remain largely unaffected, as noted by equations 2C and 2D.

\(^8\)Chetty [8] also made an effort to correct Chow’s work for serial correlation. His findings proved only modestly different from those originally reported by Chow, perhaps in part because Chetty employed a Bayesian autocorrelation parameter instead of the Cochrane-Orcutt or Hildreth-Liu procedures and estimated only nominal equations instead of the properly specified real equations. \( R^2 \) statistics and standard errors are not strictly comparable across equations whenever the regressions are adjusted for serial correlation, since the dependent variables are a function of the estimate of \( \rho \), the serial correlation coefficient. Equations with different \( \rho \) estimates have different dependent variables and different variances.
TABLE 2
DEMAND-FOR-MONEY REGRESSIONS, 1897–1958

<table>
<thead>
<tr>
<th>Equations</th>
<th>$k$</th>
<th>$W$</th>
<th>$Y$</th>
<th>bond</th>
<th>$R^2$</th>
<th>S. E.</th>
<th>D-W</th>
<th>$p$</th>
</tr>
</thead>
<tbody>
<tr>
<td>A</td>
<td>-2.47</td>
<td>1.057</td>
<td>-0.703</td>
<td>0.986</td>
<td>0.081</td>
<td>0.43</td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td>(-7.57)</td>
<td>(48.89)</td>
<td>(-11.97)</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>B</td>
<td>1.65</td>
<td>0.988</td>
<td>-0.617</td>
<td>0.963</td>
<td>0.038</td>
<td>2.00</td>
<td>0.69</td>
<td></td>
</tr>
<tr>
<td></td>
<td>(-3.55)</td>
<td>(30.63)</td>
<td>(-8.37)</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>C</td>
<td>-1.62</td>
<td>0.627</td>
<td>0.408</td>
<td>-0.628</td>
<td>0.990</td>
<td>0.071</td>
<td>0.54</td>
<td></td>
</tr>
<tr>
<td></td>
<td>(-4.52)</td>
<td>(5.65)</td>
<td>(3.95)</td>
<td>(-11.39)</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>D</td>
<td>-1.36</td>
<td>0.732</td>
<td>0.255</td>
<td>-0.525</td>
<td>0.972</td>
<td>0.034</td>
<td>2.00</td>
<td>0.68</td>
</tr>
<tr>
<td></td>
<td>(-3.23)</td>
<td>(9.32)</td>
<td>(7.41)</td>
<td></td>
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</tr>
</tbody>
</table>

At this point, it may be useful to reconsider the objectives of these equations. Equations including both current income and wealth are useful in attempting to discriminate between transactions and asset theories of demand. Such an equation will fail in its mission if institutional changes have occurred that have altered the uses of money (however it may be defined). In fact, major changes did occur and may have resulted in the specialization of money into the various roles for which these tests were designed.

Before 1913 and the Federal Reserve Act, legal reserve requirements did not in all cases differentiate time from demand deposits. Moreover, banks often made no clear distinction between these two types of deposits. If prior to 1913 banks did not bother differentiating time from demand deposits, it is unclear why individuals should be viewed as making any clearly meaningful distinction between demand and time deposits.

Unfortunately, the data between 1913 and 1934 are also somewhat suspect. Prior to the passage of the Banking Act of 1933, banks were permitted to pay interest on demand deposits. Thus, the motives for holding demand deposits may have differed from the motives manifested in today’s world where such deposits earn no interest. Even if demand deposits are held today purely for transactions purposes and all asset demand is manifested in the form of other deposits or assets, prior to 1933 demand deposits may have also satisfied some asset demand. After all, demand deposits earned a positive pecuniary return. And whatever the motivations for holding money before 1933, it seems plausible that transactions motives became more important during 1933 as a result of the changes in the law.10

9Some of the recognized weaknesses of the early data include incompleteness due to the unavailability of deposit data for some banks in addition to the difficulty of separating deposit data into demand and time deposits. The latter problem is of primary concern for this analysis. See [17, p. 211–3, 329–31] for extensive discussions on the quality of the money-stock data. Friedman, for one, has long preferred to use $M_2$ instead of $M_1$ as a measure of the money stock, in part because of the data-quality problem.

10It might seem appropriate to continue estimating these equations spanning the changes of 1933 if only we added an own rate of return on money, as several studies have done. Presumably, the pecuniary own rate of return would decline in 1933 and the nonpecuniary return, typically in the form of foregone service charges, would increase. This procedure is inappropriate, however, since it fails to distinguish between average and marginal rates of return. Once services charges are forgiven, the marginal return to adding a dollar to a demand deposit is zero. And it is the marginal rate that matters. Using a somewhat different approach, recent work by Klein [27] assumes that all excess returns caused by deposit-rate ceilings are paid implicitly to depositors proportionately with respect to deposits. Even if all excess returns are competed away—a proposition open to dispute—the excess returns may be distributed in part to depositors, lenders, to the government (as taxes), or squandered as management perquisites and as higher costs. It is therefore unclear how much of the excess return, if any, represents a marginal return on deposits.
Restated somewhat, the Banking Act of 1933 may have resulted in a degree of specialization in the roles of $M_1$ and time deposits that did not exist previously. $M_1$ quite possibly became more specialized as a medium of exchange, whereas time deposits become more specialized as a near-money store of value. As a result of the institutional changes of 1913 and 1933, the meaningfulness of a single narrow definition of money is unclear. Modern empirical researchers may be trying to find evidence for theories that make distinctions that in fact were made only some time after their first observation.

To test this proposition, a subset of Chow’s data set from 1934 through 1958 (excluding the war years of 1941–45) was used to reestimate all of the equations reported in Tables 1 and 2. Table 3 reports subperiod results for the equations that test the performance of income and asset effects simultaneously. Some additional equations, many of which also show substantially different estimated coefficients when they are adjusted for serial correlation, are reported in the appendix.

### Table 3
Demand-for-Money Regressions, 1934–1958

<table>
<thead>
<tr>
<th>Equations</th>
<th>$k$</th>
<th>$\gamma$</th>
<th>$\gamma_p$</th>
<th>W</th>
<th>bond</th>
<th>$R_2$</th>
<th>S.E.</th>
<th>D-W</th>
<th>$\rho$</th>
</tr>
</thead>
<tbody>
<tr>
<td>A</td>
<td>-1.96</td>
<td>0.200</td>
<td>0.861</td>
<td>-0.959</td>
<td>0.973</td>
<td>0.070</td>
<td>1.03</td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td>(-1.09)</td>
<td>(0.51)</td>
<td>(1.79)</td>
<td>(-5.33)</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>B</td>
<td>1.25</td>
<td>0.597</td>
<td>0.221</td>
<td>-0.270</td>
<td>0.869</td>
<td>0.028</td>
<td>1.63</td>
<td>0.74</td>
<td></td>
</tr>
<tr>
<td></td>
<td>(0.98)</td>
<td>(3.65)</td>
<td>(1.20)</td>
<td>(-2.18)</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>C</td>
<td>-1.31</td>
<td>-2.87</td>
<td>1.444</td>
<td>-0.601</td>
<td>0.988</td>
<td>0.047</td>
<td>1.21</td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td>(-2.23)</td>
<td>(-1.26)</td>
<td>(5.24)</td>
<td>(-7.14)</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>D</td>
<td>2.11</td>
<td>0.680</td>
<td>0.097</td>
<td>-0.226</td>
<td>0.827</td>
<td>0.029</td>
<td>1.81</td>
<td>0.76</td>
<td></td>
</tr>
<tr>
<td></td>
<td>(1.44)</td>
<td>(3.02)</td>
<td>(0.32)</td>
<td>(-1.73)</td>
<td></td>
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<td></td>
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<td></td>
</tr>
</tbody>
</table>

Equation 3A, which is unadjusted for serial correlation, seemingly supports the findings of Chow [9], Laidler [30, 31], and Meltzer [35] that asset demand dominates transactions demand over long historical samples. Even so, the Durbin-Watson statistic reveals that the equation is not unambiguously free of serially correlated error terms.\(^{11}\)

Equation 3A was also reestimated adjusting for serial correlation and the results are reported as equation 3B. Once again, the coefficient estimates of the equations adjusted for serial correlation are markedly different from their unadjusted variants but they are also different from the equations estimated over the longer period, 1897–1958. Using a Chow test and a comparison of equations 2D and 3B, the null hypothesis that no structural shift occurred is easily rejected with an $F$ statistic of 6.77 as compared to the critical value of 3.16.\(^{12}\)

Although the Durbin-Watson statistics of equation 3A fall into the indeterminate region, the estimated serial-correlation coefficient of equation 3B of 0.74 is highly significant with a $t$ value of 4.61, whereas the Durbin-Watson statistic falls within

\(^{11}\)The small number of remaining observations drastically increases the size of the indeterminate region of the Durbin-Watson test by substantially reducing the lower bound and reducing the upper bound modestly.

\(^{12}\)Strictly speaking, the Chow test is inappropriate here, since the equations are adjusted for serial correlation and each equation has a different estimated serial-correlation coefficient. The serial-correlation coefficients at 0.68, 0.74 and 0.63 for the period 1897–1933 are, however, reasonably close.
the serially independent region of the test. Moreover, the estimated coefficients of these equations now support the transactions theory of money demand and reverse the conclusions drawn by Chow [9], Meltzer [35], Laidler [28], and many others. The wealth coefficient is small and statistically insignificant. The income coefficient, however, is statistically significant and demonstrates economies of scale in money demand. Although the inventory-theoretic approach to money demand has long suggested economies of scale, empirical evidence supporting this view has been sparse, especially with pre-World War II data.

The new interest-rate coefficient is also noteworthy in that it is much smaller (in absolute value) than is characteristic of equations unadjusted for serial correlation. Moreover, the interest rate coefficient is marginally significant or marginally insignificant in many different equations suggesting that interest-rate effects on the demand for money may be somewhat smaller than previously believed. Although recent papers [20, 33, 6] have provided evidence of small interest-rate elasticities of money demand for the postwar era, studies of the pre-World War II period have invariably suggested larger effects. Those larger effects are now seen to be, in large part, a result of the serial-correlation problem.

When permanent income was used as a measure of asset demand in place of wealth, the changes resulting from adjusting for serial correlation also proved quite marked. Using the 1934–58 subperiod, equations unadjusted for serial correlation show significant permanent-income coefficients and negative but insignificant current-income coefficients, as indicated by equation 3C. After the necessary serial-correlation adjustments were made, the coefficient of permanent income was reduced to insufficence, whereas the current-income coefficient of about 0.6 was highly significant, as presented by equation 3D.

3. THE "SHORT-RUN" DEMAND FOR MONEY, REVISITED

Chow’s short-run equations (those including the lagged dependent variable) were also reestimated adjusting for serial correlation over the 1897–1958 and 1934–58 periods. The findings tend to confirm the long-run estimates presented earlier (which excluded the lagged dependent variable). Four sample equations, all of which are estimated in real terms, are presented below. Equations (1) and (2) employ the Cochrane-Orcutt technique to adjust for serial correlation. Equation (3) employs an instrumental-variable technique to eliminate biases introduced by including a lagged dependent variable in an equation subject to serial correlation.14

13The serial-correlation coefficients are not significantly different from unity, however. The suitability of a first-order serial-correlation adjustment was also corroborated by examining the autocorrelations of the differences using a first difference and the 0.74 estimate produced by the least-squares regression. In both cases, all autocorrelations were statistically insignificant.

14See [26, 22] for extensive discussions on the econometric difficulties introduced by serial correlation and lagged dependent variables. For example, the coefficient of the lagged dependent variable will be biased upwards and the speed of adjustment will be biased downwards. Instrumental-variable techniques are suggested in [26, 11] to obtain consistent and more efficient estimated coefficients.
The serial-correlation adjustments were obtained by a Hildreth-Liu search and the equations presented are those with the lowest estimated standard error. Equation (4) employs Hatanaka’s [25] residual adjusted Aitken estimator to obtain consistent and efficient coefficient estimates.

\[
\ln M/P = -0.44 + 0.192 \ln GNP/P + 0.352 \ln W/P \\
(\bar{0.50}) (3.02) (3.87) \\
- 0.392 \ln bond + 0.440 \ln (M/P)_{t-1} \\
(-6.69) (5.35)
\]

\[R^2 = 0.995 \quad \text{S.E.} = 0.32 \quad \rho = 0.30 . \tag{1}\]

Equation (1), which is estimated over the period 1897–1958, displays a number of noteworthy changes from Chow’s published equation 3.3. The equation suggests that both asset and transactions motives are important in explaining \( M_1 \) demand, at least over the entire period for which the equation is estimated. In addition, equation (1) suggests a somewhat faster speed of adjustment of 56 percent per year, as compared to Chow’s 48 percent, but as noted by Griliches [22] and Fair [11], the estimated speed of adjustment is likely to be understated in equations containing lagged dependent variables and subject to serial correlation.

These findings are even more strongly supported when the equations are reestimated over the 1934–58 period, as reported below.

\[
\ln M/P = 0.91 + 0.474 \ln GNP/P - 0.050 \ln W/P \\
(0.84) (1.97) (-0.16) \\
- 0.352 \ln bond + 0.501 \ln (M/P)_{t-1} \\
(-2.45) (3.46)
\]

\[R^2 = 0.988 \quad \text{S.E.} = 0.033 \quad \rho = 0.18 . \tag{2}\]

Equation (2) suggests, first of all, that wealth is not significant in explaining money demand over the 1934–58 period. Thus, the equation supports a transactions theory of money demand. The speed of adjustment, which is estimated at about 50 percent, is still rather low, as suggested by Griliches [22] and Fair [11].

Since equations that include lagged dependent variables are subject to estimation problems due to the correlation between the errors and the lagged dependent variables even when adjusted for serial correlation, an instrumental-variable technique was employed to obtain more reliable coefficient estimates (see [26, Chap. 10, esp. pp. 316–20]). The exogenous variables and their lagged values as

15 The estimation procedure employed in this analysis combines the instrumental-variable techniques of Leviatan [34] and Wallis [38] and a Hildreth-Liu scan. The instrumental-variable technique suggested by Fair [11] was also employed with no change in the results from those reported here.
well as population and a time trend were employed as instruments for the lagged dependent variable.\textsuperscript{16} A search routine was employed to estimate the serial-correlation coefficient and to provide more efficient as well as consistent coefficient estimates.\textsuperscript{17} The equation with the lowest standard error is reported as equation (3).

\[
\ln M/P = 1.09 + 0.615 \ln GNP/P + 0.282 \ln W/P
\]
\[
- 0.276 \ln bond - 0.080 \ln (M/P)_{t-1}
\]
\[
(0.70) \quad (3.18) \quad (1.10)
\]
\[
(\text{S.E.} = 0.031, \quad \rho = 0.72)
\]

Equation (3) confirms the previous findings that wealth is not an important explanatory variable in an $M_1$ equation and the lagged dependent variable is now also statistically insignificant. An insignificant lagged dependent variable suggests that the speed of adjustment is insignificantly different from 100 percent per year. Thus the instrumental-variable technique suggests a faster speed of adjustment than the previous results, which, as noted in [11, 22], understate the true speed of adjustment. Moreover, these results are not affected materially by deleting wealth or lagged money from the equation. For example, the income elasticity of 0.601 of equation (3) is insignificantly different from the 0.597 estimate of equation 3B.

Since the estimation procedure is so crucial when equations that include lagged dependent variables are suspected of autocorrelated errors, Hatanaka’s [25] efficient residual adjusted Aitken estimator was also employed, and these results are reported as equation (4).

\[
\ln M/P = 0.17 + 0.527 \ln Y/P + 0.186 \ln W/P - 0.339 \ln bond
\]
\[
(0.14) \quad (2.11) \quad (0.62) \quad (\text{S.E.} = 0.035, \quad \hat{\rho} = 0.45)
\]
\[
+ 0.219 \ln (M/P)_{t-1} - 0.11 \mu_{t-1}
\]
\[
(0.78) \quad (\text{S.E.} = 0.035, \quad \hat{\rho} = 0.45)
\]

The instruments used in the first stage include the contemporaneous and lagged values of the exogenous variables as well as a time trend and population. Other instruments were also employed with no meaningful changes in the findings. These results confirm the findings of the instrumental-variable-scanning estimator. The wealth coefficient remains insignificant and the income coefficient suggests economies of scale and retains its significance. The insignificant lagged dependent-variable coefficient indicates that the speed of adjustment is not meaningfully different from 100 percent per year. And the results are also quite similar to the

\textsuperscript{16}The fit of the lagged dependent variable in the first stage was quite good with an $R^2$ of 0.98.
\textsuperscript{17}The coefficient estimates were very robust with respect to different values of rho, the serial correlation coefficient, during the search.
long-run estimates of equation 3B. Thus, all of the 1934–58 results suggest that wealth is not an important explanatory variable in an M1 equation, at least over that period. Moreover, the results are consistent with our intuitive and casual observations that the money market clears well within one year.

The findings presented here may be compared to those obtained by Feige [12], whose work distinguished between adjustments between desired and actual money stocks and between current and expected income as determinants of the desired level of money balances. As in the results reported here, Feige [12] found that the adjustment between desired and actual money stocks was complete within a year. He then inferred that the significance of the lagged dependent variable in previous work was derived from the effects of expected income. Moreover, the elasticity of expected income with respect to current income was 0.37, very close to the 0.40 estimate in [6].

The results reported here, however, indicate that when expected income (or nonhuman wealth) is explicitly included in the analysis, income remains significant and permanent income and wealth are insignificant. Moreover, the measure of permanent income employed here is Friedman’s own measure and would, as a result, be extremely close to Feige’s estimates of permanent income. The differing conclusions are probably due to differences in sample periods. Feige’s estimates cover 1915–63 and therefore span the structural break uncovered for 1933. The analysis reported here avoids this problem and the misspecification produced by the structural shift.

A comparison of Goldfeld’s [20] more recent findings with the reestimates of Chow’s findings (adjusted for serial correlation) is also of interest. Goldfeld [20] employed the Cochrane-Orcutt routine to adjust for serial correlation and showed wealth to be statistically insignificant in a money-demand equation; his long-run income elasticity was estimated at 0.68 [20, p. 583]. The results reported here are insignificantly different from Goldfeld’s postwar estimates. All of the long-run real-income elasticities estimated over the 1934–58 period fall within the 0.6 to 0.9 range, and wealth was also insignificant in all of the equations estimated. Although the interest elasticities reported here are larger than in [20] (in absolute value), that may be largely a result of the use of a short-term commercial-paper rate in [20] as opposed to the bond rate in this analysis. Thus, the traditional formulations, when adjusted for serial correlation and estimated using data after 1933, provide coefficient estimates consistent with the postwar Goldfeld analysis [20].

The results reported here, which are based upon the more traditional

---

18There is some overlap in the time period between Goldfeld’s data, which began with the 1952 second-quarter observation, and with data employed in this paper, which ended in 1958. Unfortunately, avoiding this overlap would have left the analysis with too few degrees of freedom. Even so, the consistency of the finding is comforting in that Goldfeld’s analysis includes many observations not included in this paper and vice versa.

19Cagan and Schwartz [6] have recently suggested that the interest sensitivity of money demand has declined as a result of technological change and the growth of money substitutes. Unfortunately, their tests, which compare estimated interest-rate coefficients from the postwar period to equations estimated over the 1921–31 period, are biased in favor of their thesis. As shown above, M1 equations estimated before 1934 are not comparable to later M1 equations. Moreover, the interest sensitivity of money over the 1934–58 period is not inconsistent with Goldfeld’s [20] postwar 1952–73 estimates, although they are larger than the postwar findings obtained by Lieberman [3].
income-transactions model, may also be compared with the postwar debits-
transactions model recently reported by Lieberman [33]. The income model, like
the debits model, indicates complete adjustment within one year, significant
transactions coefficients, and insignificant wealth or permanent-income coeff-
cients. Moreover, the income and debits approaches also yield comparable
coefficient estimates. Since the elasticity of debits with respect to income is about
two, the estimated elasticity of money with respect to debits, which ranged from
about 0.3 to 0.4, is consistent with the income elasticities reported here, which
ranged from about 0.6 to 0.9 [33, p. 310].

4. CONCLUSIONS AND IMPLICATIONS

Several conclusions may be drawn from this analysis. The first conclusion
suggests that money-demand equations that are subject to serial correlation may
well provide misleading empirical results unless they are properly adjusted to
remove this difficulty. Although coefficients estimated in the presence of serial
correlation are theoretically unbiased (if the equation is specified correctly), ad-
justing for serial correlation may lead to marked changes in the estimated co-
efficients and may reveal very different underlying relationships. The empirical
findings reported in this paper suggest that money-demand equations in particular
may provide misleading coefficient estimates unless they are adjusted to remove the
existing serial correlation.

The second primary conclusion to be drawn from this analysis is that long-run
studies that span the pre- and post-1933 periods may be misleading because of the
major changes in the structure of the financial system and in the motives for holding
money that appear to have occurred. Before 1934, the motives for holding $M_1$ may
have been different, at least in part, because interest payments on demand deposits
were permitted. Thus, to a larger extent than today, $M_1$ may have been held as a
store of value. Thus, researchers who have attempted long-term estimates of money
demand have overlooked an important change in the structure of the money-demand
equation, which has biased their findings.

Although long-run $M_1$ studies appear to be unreliable, it is still feasible to
examine the long-term behavior of broader monetary aggregates. Thus, whenever it
is necessary to study the behavior of the aggregates before 1934, efforts should be
made to build up to the more reliable $M_2$ or $M_1$ measures, which span the asset and
transactions motives, instead of employing a structurally dissimilar component like
$M_1$. Even if studies of the broader aggregates are contemplated, structural changes
in the underlying components require that such empirical research be undertaken
only with considerable care.\footnote{Previous long-term $M_2$ studies are considerably more reliable than the long-run $M_1$ studies, provided
that, whenever necessary, the appropriate adjustments for serial correlation have been made. See [30, 31, 15], and others for long-run $M_2$ studies. None was adjusted for serial correlation, however.}

The third primary conclusion suggested by this study is that the transactions
model of money demand appears to be more appropriate than the asset or utility
demand model, at least over the post-1933 period, in sharp contrast to the findings of many previous studies. Wealth and permanent income are statistically insignificant in all cases where current income enters as a measure of the volume of transactions. Moreover, the speed of adjustment of the money market appears to be quite rapid, as commonly believed. These findings are also consistent with the postwar findings of Goldfeld [20] and Lieberman [33], whose results also favor the transactions model. Thus, the evidence implies that the narrow money stock is held first and foremost for transactions purposes and that neither net worth nor permanent income aid significantly in explaining demand for these balances.

Finally, the empirical findings reported an absence of wealth effects on money demand, which tends to favor proponents of the efficacy of fiscal policy and those who suggest that the portfolio effects of debt-financed fiscal policy may crowd in capital investment. As demonstrated by B. Friedman [14], bond-financed fiscal policy supports the crowding in of investment if there are no wealth effects on money demand.

APPENDIX: TABLE 4
DEMAND-FOR-MONEY REGRESSIONS, 1934–1958

<table>
<thead>
<tr>
<th>Equations</th>
<th>k</th>
<th>y</th>
<th>w</th>
<th>Yp</th>
<th>Bond</th>
<th>R²</th>
<th>S.E.</th>
<th>D.W</th>
<th>ρ</th>
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<td>(-1.96)</td>
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