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Abstract

We find strong evidence of a stable "money demand" relationship for MZM and M2M through the 1990s. Though the M2 relation breaks down somewhere around 1990, evidence has been accumulating that the disturbance is well characterized as a permanent upward shift in M2 velocity, which began around 1990 and was largely over by 1994. Taken together, our results support the hypothesis that households permanently reallocated a portion of their wealth from time deposits to mutual funds. Although this reallocation may have been induced by depository restructuring, we argue that the substitution could be explained by appropriately measured opportunity cost.

Key Words: Money Demand; VECM;

JEL Classification: E41

* The views stated herein are those of the authors and are not necessarily those of the Federal Reserve Banks of Cleveland or St. Louis or of the Board of Governors of the Federal Reserve System.

1. Background

It is incumbent upon advocates of a role for money in the conduct of monetary policy to demonstrate that certain measures of money are reliably linked to objectives of policy, which may include intermediate targets or ultimate goals. For example, based on data from 1960-1992 Feldstein and Stock (1994) investigate the relationship between M2 and nominal GDP. They conclude that the relationship is sufficiently strong and stable to warrant a further investigation of the link between M2 and nominal GDP. Experience through the mid 1990s, however, revealed a sharp unexplained increase in M2 velocity, undermining the evidence of Feldstein and Stock. At face value, the 1990s experience suggests that the issue of a stable relationship between money and GDP remains unresolved.

Empirical investigations by Friedman and Kuttner (1992), Miyao (1996), and Estrella and Mishkin (1997) examine evidence of significant relationships between money (variously defined) and nominal income or between money and real income or prices separately. Friedman and Kuttner find that inclusion of data from the 1980s weakens the postwar time-series evidence for such relationships, and focusing on data after 1970 destroys this evidence altogether. They look at both joint significance tests and evidence on cointegrating relationships. Estrella and Mishkin examine joint significance and parameter stability using Chow tests. They also conclude that the link between money and economic activity breaks down in the 1980s. Miyao focuses on a cointegrating M2 demand relation, using an array of testing methods and allowing for both constants and trends. Although test results are mixed for subsamples before 1990, virtually no results support the existence of a cointegrating relation through the mid 1990s. In sum, the promising empirical conclusions of Feldstein and Stock (1994) that established predictive content for M2 in a vector error correction setting do not seem to find support in data that extend though the mid 1990s.

These recent studies seriously question how useful broad monetary aggregates are in explaining the behavior of nominal economic activity. A key issue here is whether there remains any empirical evidence for a stable relationship linking broad aggregates, nominal economic activity, and the opportunity cost of maintaining money balances. If there is no evidence that such a relationship exists, there is no empirical basis for assigning any role in monetary policy to a broad aggregate. If a stable relationship can actually be established from the data, the questions remain: Are the deviations from the long-run equilibrium sufficiently small and free from persistence to facilitate policy-making objectives or to allow pursuit of monetary rules? Do deviations from the long-run equilibrium actually portend changes in the pace of the nominal economy? Is the relation resilient to innovations that characterize the macro economy?

We reexamine this issue with a battery of cointegration tests. All cointegration tests and estimates are based upon models that span three broad aggregate definitions—M2, MZM, and M2 minus small time deposits (M2M)—and three distinct scale measures that are available at monthly frequency—the index of coincident indicators used by Estrella and Mishkin (1997), the index of industrial production, and personal income.[1] Money balances are converted to real measures using two distinct price deflators, the CPI and the "chain-weighted" price deflator for personal income. The robustness of conclusions reached in the empirical analysis is explored using alternative samples and alternative lag-length parameterizations. The tests reveal considerable evidence that stability is maintained by a relation linking each of the broad monetary aggregates, the scale of the economy, and the relevant measure of the opportunity cost of maintaining aggregate money balances. However, it is clear that the relation can appear unstable if there is no accounting for episodes when distinct changes occurred in the incentives

[1] MZM, or money at zero maturity, includes all financial instruments that are, or can be easily converted into, transactions balances without penalty or risk of capital loss. MZM includes M1, savings deposits, including money market deposit accounts, and both institutional and retail money market mutual funds. Unlike M2, MZM does not include small time deposits (e.g., retail certificates of deposit). This concept was first proposed by Motley (1988). M2M is simply M2 less these small time deposit accounts.

for maintaining the small time deposit portion of M2. We demonstrate, by comparisons with M2M and MZM, that observed instabilities of the M2 relationship discussed by Carlson and Parrot (1991), Partlan and Wenniger (1992), Darin and Hetzel (1994), and Lown et al. (1997) may be isolated in the small time deposit component.

Can the existence of a broad aggregate relation, formed from any of the monetary aggregate measures, assist monetary policy makers? We explore the issue by examining the deviations from the cointegrating equation formed from estimates of the long-run relation. Next, we examine whether an equilibrium nominal income level may be extracted from the relation and measure how nominal income responds to deviations from this equilibrium. After we control for widely documented episodes of financial innovation, this exercise is analogous to measuring predictive content of information sets in VECMs as in Feldstein and Stock (1992). Finally, we apply the technical framework popularized by King, Plosser, Stock, and Watson (1991) to examine how long it takes to reestablish the long-run relationship after perturbations to fundamental permanent or transitory shocks that prevail in a simple vector autoregressive representation of the system.

We find considerable support for a stable long-run relation with minimal short-run perturbation in an M2 specification of the money demand relation whereas modest support is found for a similarly defined MZM specification. The results suggest, in contrast to Estrella and Mishkin(1997), that the aggregate money measures do indeed maintain predictive content for the nominal representation of the macroeconomy once we account for the effects of financial innovation in the early 1990s. Also, we find that the equilibrium cointegrating relationships we identify are reestablished within a reasonably short period of time following particular innovations to the VECM representation of the system. These results are quite robust to the varied specifications we examine, suggesting that continued attention to broad aggregate relations offers the possibilities discussed in the literature on a policy role for money.

2. The breakdown in the M2 relation.

The essence of the criticism leveled at M2 in recent years may be summarized by the distinct departure from the presumed stable cointegrating relation for M2 that was observed in the 1990s. This departure, illustrated in Figure 1, reveals the sharp contrast between the stable relation that prevailed in pre-1990 samples and more recent performance. The top panel in Figure 1 depicts the monthly time series of the natural log of M2 velocity (VELM2CN), using the coincident indicator series (as in Estrella and Mishkin) to proxy scale over the period 1964:1 to 1998:12. The lower panel portrays the opportunity cost of maintaining M2 balances against the effective yield on U.S. three-month Treasury bills (RM2E).[2] The figure reveals that before 1990, M2 velocity fluctuated only about 15% with the observed fluctuation clearly positively correlated with increases in M2 opportunity cost.[3] Perhaps the most telling portion of this period is the 1990-94 interval, where in the first two years the opportunity cost fell sharply (from 0.024 to 0.005) with no discernible change in velocity. In 1993 and 1994 opportunity cost first leveled off and then surged sharply (from 0.01 to 0.03) with again, no discernible response from M2 velocity, which grew at a fairly constant rate from 1991 through 1994.

The significance of the 1990's break from the historical pattern is revealed in Figure 2. This Figure contains actual velocity and the levels of velocity that are predicted from bivariate (velocity and opportunity cost) cointegrating equations through 1989 for a six-lag specification. The pre 1990 historical relation is simply inconsistent with the sharp increase in velocity observed in the early 1990's. Figures 1 and 2 clearly reveal that the positive correlation between

[2] M2 opportunity cost is defined here as the difference between the yield on the three-month Treasury bill and the share-weighted rates of return on M2 components. Similarly, the opportunity costs of measures of money considered below are defined as the difference between the three-month Treasury bill yield and the share-weighted rates of returns on the measures components.

[3] Comparable velocity-opportunity cost relations can be obtained using the index of industrial production to proxy the scale of the economy as examined below in some of the cointegration tests. M2 velocity measured using personal income to proxy scale displays a distinct linear trend though 1989, but deviations from the trend correspond to variation in opportunity cost. In the 1990s the trend stability of M2 personal income velocity clearly disappeared in a fashion analogous to the deterioration in mean stability depicted

opportunity cost and M2 velocity from an earlier period is not maintained in the 1990s, except for perhaps in 1995 through 1998.[4]

3. Testing for the stability of a broad aggregate relation

3.1. An investigation of M2

Formal evidence of the well-documented breakdown in stability of the broad M2 relation is available from tests of the cointegration proposition. Table 1 offers cointegration tests of the M2 relation:

$$\ln(M / P)_t - \beta_1 \ln(y) + \beta_2 oppcost_t = \varepsilon_t ,$$

with M measured as M2. The CPI is the price index measure in the upper panel and the chain-weighted personal income deflator is the price index in the lower panel. Proxies for *y* include the index of coincident indicators *(cn)*, the index of industrial production *(ip)* and real personal income *(pi/p).* The opportunity cost is measured as ln(1+*rtbe*/100)-log(1+*m2ownr*/100), where *rtbe* is the effective yield on three-month Treasury bills and *m2ownr* is the share-weighted yield on the M2 components.[5]

Estimates in Table 1 span the period 1964:1-1989:12.[6] Cointegration tests are obtained using Johansen's (1988) trace statistic, estimated from VECM models with 6 and 9 lags using the traditional specification that allows for linear trends in the data but no trend in the cointegrating relationship. That is, we ask whether a linear equilibrium relationship exists between real M2,

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for the coincident indicator velocity in Figure 2. The distinction between trend stability of the long-run relation and mean stability is important for the cointegration specifications we employ in the analysis below.

[4] As discussed in the preceding footnote, similar conclusions can be reached using the index of industrial production to proxy scale. If we use personal income, the distinct linear trend that prevails in velocity prior to 1990 is not maintained from 1990-1994, so the deviations from M2 personal income velocity about a constant linear trend throughout the 1964-1998 period cannot be traced to opportunity cost fluctuations in the 1990s.

[5] This approach to measuring opportunity cost is tantamount to a semi-log specification.

[6] We follow the convention of Moore, Porter and Small (1990) and begin the estimation sample in 1964:1. Conclusions from our empirical analysis are not significantly altered by inclusion of the early 1960s in the sample.

scale of the economy, and opportunity cost that eliminates all stochastic and deterministic trends from the data. Recognizing that the personal income velocity exhibits a distinct trend over the sample, we also estimated the models scaled by personal income with a specification that allows for quadratic trends in the data and a linear trend in the cointegrating relation. These specifications are denoted by *(t)* and appear in the last two rows of each panel of Table 1.

With a few exceptions, the results in Table 1 are consistent with a single cointegrating relation of the form depicted above. Instead of focusing on a particular lag-length specification or scale measure, we prefer to examine the preponderance of evidence at the two lag-length specifications, three distinct scale measures, and two price deflator indices. In general, we find that little significant residual persistence remains in the residuals of the six-lag specifications, but for comparison with Estrella and Mishkin, we present the results obtained with nine lags as well. The long-run income elasticity of real money balances is in the vicinity of 1.0, and the estimated interest semi-elasticity ranges from -2.7 to -4.5 in the CPI models and is -2.5 to -3 in the chain-weighted price index specifications. The cointegration tests in Tables 1 and 2 formalize the breakdown in the cointegrating relation that is apparent in Figure 2. Table 2 contains comparable estimates and test statistics obtained from exactly the same specification with the sample extended through 1998:12. Clearly, there is less evidence of cointegration when the tests incorporate the additional 108 data points taken from the 1990s. None of the 12 trace statistics is significant at the 5% level. Estimates of the scale elasticities are uniformly smaller and considerably less precise, while estimates of interest semi-elasticities are considerably larger (often taking on unrealistically high levels) and are typically very imprecise.

3.2. An investigation of MZM and M2M

Our contention is that well documented changes in the incentives to maintain small time deposits throughout the 1990s are responsible for the observed instability in the M2 money demand relation. We establish this position by examining the stability of relations based on alternative monetary aggregate measures: MZM and M2M. MZM refers to the money at zero maturity measure discussed by Poole (1991) and Motley (1988), and examined recently by Carlson and Keen (1996). If the origin of the instability is indeed small time deposits, these aggregate measures should be immune from the financial innovations that induced instabilities in the M2 relation in the 1990s.

Figure 3 depicts the relationship between the natural log of MZM velocity (VELMZMCN) in the upper panel, again using the coincident indicator series to proxy scale, over the period 1976:1 to 1998:12.[7] The lower panel portrays the opportunity cost of maintaining MZM balances against the effective yield on U.S. Treasury bills (RMZME) over the same period. Taken together, the panels reveal that the MZM velocity movements in the 1990s, unlike those of the M2 counterpart, continue to mirror movements in MZM opportunity cost. The M2M aggregate may be used to produce a pattern of velocity and opportunity cost movements that is indistinguishable from Figure 3. This suggests that velocity measures constructed from monetary aggregates that exclude small time deposits do not display the pronounced change in trend that appeared in M2 velocity in the 1990s.

Tests of cointegration based on the MZM and M2M measures appear in Tables 3 and 4, where we focus on the last 23 years of the sample (1976:1-1998:12).[8] These tests reveal very strong evidence of cointegration for both MZM and M2M at nearly all lag-length specifications

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[7] The 1976:1-1998:12 period was chosen to match an interval suggested by Carlson and Keen, who demonstrate that money demand relations based upon MZM and M2M exhibit instabilities prior to 1976.

[8] We also investigate the stability of these relationships from 1964 to present. The corresponding tests fail to detect evidence of deterministic cointegration. Our hypothesis is that the absence of a cointegrating relationship before 1975 reflects the measurement error in the computed opportunity cost for small time deposits. In principle, an appropriate measure should be net of both proportional and fixed transactions costs. We hypothesize that fixed transactions costs on small time deposits were so large that their *net* yield was essentially zero. Thus, households had no incentive to diversify into such assets before the mid 1960s. However, we find evidence that one can control for such effects by including a variable that measured the maximum fed funds rate to date. The inclusion of the ratchet proxy effectively captures the linear trend in MZM and M2M income velocity during the 1964-1975 period. Hence, there is significant evidence of cointegrating MZM and M2M relationships throughout the sample, once we allow for the possibility that transactions costs shift with changes in the ratchet variable. These results appear in Table 8 and are discussed below.

and income measures and across both price index measures. The income elasticities are in the vicinity of unity and are statistically significant, while a sizable and statistically significant opportunity cost elasticity is observed throughout.

The results in Tables 3 and 4 may be contrasted with similar tests based upon M2 over the 1976:1-1998:12 sample periods which are presented in Table 5. The latter are very similar to the results depicted in Table 2 because they incorporate the instabilities of the 1990s. There is no significant evidence of a cointegrating M2 money demand relationship in this sample, and income elasticities are low and often insignificant, while opportunity cost elasticities are, as in Table 2, generally large and imprecise. The cointegration tests clearly reveal strong evidence of a stable cointegrating relation comprised of monetary aggregates MZM and M2M–suggesting that the problem with M2 clearly relates to portfolio changes in recent years that impact M2 but not MZM and M2M.

3.3. Explaining instabilities in M2 in the 1990s

Numerous studies, including Duca (1992, 1995), Collins and Edwards (1994), Darin and Hetzel (1994), Orphanides et al. (1994), Carlson and Keen (1996), and Mehra (1997), have found an association between the M2 demand shortfall and the meteoric growth in stock and bond mutual funds, which are not included in M2. The consensus of this literature is that mutual funds are relatively close substitutes for time deposits, and what occurred in the early 1990s was a massive reallocation in household portfolios from time deposits to bond funds and to a lesser extent, stock funds. The question remains, however: Why did such a reallocation occur at the time it did?

The research cited above leads us to conclude that the timing largely reflected the confluence of two factors. First, innovations reduced transactions costs of mutual funds and increased their accessibility to households. These instruments allowed individuals to buy into a diversified portfolio of long-term bonds, while maintaining some liquidity with check-writing privileges. These features made them increasingly attractive as substitutes for time deposits. Second, beginning the late 1980s, many depositories, especially thrift institutions, found themselves in poor financial condition. A large number of such institutions failed to meet minimum capital standards and hence were constrained from acquiring assets and thus from competing aggressively for funds (see Carlson and Parrott [1991]).

The restructuring of depositories acted as a catalyst in the development of mutual funds, especially bond mutual funds. Bond funds are subject to capital losses in the short run, but in the long run they yield relatively higher rates than deposit instruments. When short-term interest rates began falling in 1989, the mutual fund industry intensified marketing strategies that informed households about bond funds, which were yielding significantly higher returns. Capital-constrained thrifts were effectively limited in their pricing responses. As a result, many households, apparently for the first time, diversified their portfolios out of M2 deposits into bond mutual funds. It appears now that for many of these households, bond funds have become a permanent and significant part of their portfolios, thus supplanting bank CDs.

In principle, one would expect that substitution from time deposits to bond funds should be associated with a corresponding change in the relative returns. Feinman and Porter (1992), Mehra (1997), and Lown et al. (1997) examine evidence on this issue. Lown et al. find some evidence that capital-constrained thrifts offered lower rates for time deposits than did unconstrained thrifts.[9] They conclude, however, that the differentials are too small to account for the extent of the M2 slowdown. Mehra and Feinman and Porter develop more comprehensive

[9] In principle, a properly measured opportunity cost should also allow for relative risk premiums. When the deposit insurance fund was solvent, there was no reason to introduce a risk premium in the measurement. However, if the deposit insurance fund becomes insolvent, if thrifts are being closed, and if long-term time deposits are being "called" and the public is aware of the problem (such as after the passage of FIRREA in 1989), then the issue of risk or call premiums, in principle, should be considered in constructing a properly measured opportunity cost. Of course, in the available measures, nothing like this is considered, and it is not clear how to make such measurements. Without some mismeasurement of the marginal opportunity cost, all that would be observed is a shift in the allocation of small time deposits among institutions, not a problem in the slowdown in the growth of the total. Moreover, the threat of insolvency could account for only a transitory effect on M2 demand/velocity.

measures of opportunity cost that include capital market yields as alternative rates. These approaches, however, also fall short as a complete explanation.

Our hypothesis is that a truly comprehensive measure of opportunity cost should also net out transactions costs. This may not be a significant problem when transactions costs are small and time invariant. Since many households were unaware of mutual funds as deposit substitutes, the effective transactions cost was infinite for these instruments; hence, the opportunity cost of holding M2 versus mutual funds was zero for such households. However, as households learned about the advantages of bond funds over time deposits, the *effective* opportunity cost of M2 rose in the aggregate. *Measured* opportunity cost thus understated the true rise of M2 opportunity cost during the period when many households learned about bond funds. Moreover, the *permanent* reduction in transactions cost can explain the *permanent* downward shift in M2 demand (upward shift in velocity). Unfortunately, data are not available to measure the discovery of bond funds by most households and hence the decline in their transactions cost. The relative stability of M2 velocity since the beginning of 1994, however, suggests that the transactions cost for obtaining and redeeming bond funds has stabilized around some small, negligible level.

The question remains: Can this innovation be modeled as an interval shift in the stability of the relationship, or does it permanently destroy the historical relationship between M2 velocity and opportunity cost? Figure 1 provides some indication that since late 1994 the relationship between M2, inflation, output, and interest rates has been reestablished. This is consistent with a one-time shift in the level of M2 relative to income. Such an outcome would be the case if the driving forces behind the velocity shift—the restructuring of the credit markets and financial innovation—had worked themselves out, leaving only a one-time effect on M2 demand. In light of this, we examine the evidence for a stable M2 monetary aggregate relationship using a smooth-shift variable, which we define as 0.0 before 1990, 1.0 after 1994, and increasing linearly in between, to proxy the innovations that gave rise to the instabilities observed in the early 1990s.

Though it is virtually impossible to pin down the precise date of these financial innovations, this shift accords with findings of Whitesell (1997) and Orphanides and Porter (1998). Using annual data, Whitesell employs a procedure that allows him to identify both the timing and the magnitude of the velocity shift. Whitesell estimates that a sharp upward shift in long-run M2 velocity essentially begins in 1990 and is largely completed by 1994. Orphanides and Porter use a regression-tree approach to estimate structural changes in the M2 velocity opportunity cost relationship. Their results lead them to conclude that the equilibrium of M2 velocity experienced an upward shift that occurred over a short period in the 1990s.

Using quarterly data, Mehra (1997) obtains similar results from a single-equation error correction specification of M2 that includes a slope shift in the opportunity cost measure in 1990:QI. Essentially, he augmented a standard M2 demand regression to include a bond rate spread variable, defined as the difference between the nominal yield on ten-year Treasury bonds and the own rate of return on M2. He compares simulations of this regression over the 1990:QI-1996:QIV period, first using actual values of the bond rate spread over the prediction interval and then repeating the simulations with the spread set to zero. The difference leads him to conclude that the weakness in M2 in the early 1990s is due to households' substitution out of M2 and into stock and bond mutual funds.

Setting the bond rate to zero in the one simulation is tantamount to assuming that transactions costs preclude portfolio diversification. That is, in the earlier period the transactions costs are such that the bonds are not a viable alternative to M2 components. The innovations in bond funds are presumed to have effectively eliminated much of the transactions cost of holding capital instruments.

Any substitution into stock and bond funds was not immediate. The penalty of early withdrawal of time deposits meant that deposit holders had incentive to hold them to their full term, and portfolio diversification could only occur as long-term small CDs matured.[10] This suggests that it may not be possible to identify a single measure of the effects of financial intermediation and that a proxy such as a broken time trend may be a viable surrogate.[11]

3.4. Cointegration tests based upon the experience of the 1990s

The literature described in the previous section suggests that a proxy for the financial innovations of the 1990s can be formed from a broken linear trend that is 0.0 prior to 1990 and 1.0 after 1994 and increases in equal increments in the 1990-1994 interval. Tables 6 and 7 depict cointegration tests obtained with this variable (t9094) added as a deterministic drift to the cointegrating relation using the full 1964:1-1998:12 sample and the 1976:1-1998:12 sample.[12] The results reveal considerable evidence of cointegration throughout the two samples, while estimates of income and opportunity cost elasticities are quite similar across the 1964-1998 and 1976-1998 samples.[13] In sharp contrast to the results in Table 2, the full sample results, with the t9094 adjustment, yield long-run elasticities that are remarkably similar to those obtained in the 1964-1989 sample depicted in Table 1. Income elasticities are in the vicinity of 1.0 (for the personal income models this is especially true when we allow for a trend in the cointegrating

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[10] Perhaps equally important, it took time for households to learn about the increased accessibility of bond funds.

[11] Unfortunately, it is not possible to discriminate between the hypothesis that there was a one-time reduction in transactions costs and a gradual adjustment to this reduction as CDs matured, versus a gradual reduction in transactions costs.

[12] The variable t9094 is similar to the interaction term Mehra (1997) includes in his regression D. His variable equals the spread between the ten-year Treasury and the own rate on M2 from 1989:QI to 1996:QIV and equals zero otherwise. Thus, his variable mimics a broken linear trend in 1980:QI. The data are too limited in duration to discriminate between these approaches.

[13] Because it is not evident when the breakdown began or ended, we ran these tests for alternative specifications of the trend break variable. The results were robust for a wide range of trend specifications. Indeed, virtually the same results are obtained from specifications that restrict the trend to 1991-1993. From Figure 1, it is this three-year period where movements in velocity and opportunity costs diverge the most from their long-term patterns.

relationship). Estimates of the interest rate semi-elasticity are quite precise and uniformly smaller than those obtained from the MZM and M2M aggregate specifications.

Comparison of the results in Table 2 with Table 6 and, separately, Table 5 with Table 7 demonstrates how important t9094 is for stability. The results suggest that once adjustments are made for widely documented financial innovations, the previously observed instabilities in the M2 relationship are overcome.

While our results suggest there is support for a stable cointegrating relation among the variables of a money demand function, several questions remain. First, which of the money supply measures, M2 or MZM, yields money demand relations that are most conducive to providing an understanding of the link between monetary policy and macro aggregates? Second, is there evidence of predictive content in specifications that embody these money demand relations? Third, how resilient are these fundamental relations to stochastic shocks that characterize the macro economy? We address these questions below.[14]

4. Aggregate Stability and Policy Usefulness

4.1. Analysis of cointegrating equations

In this section we examine whether evidence of a cointegrating relation can reveal useful information about the role of monetary aggregates in explaining short-run movements in nominal income. In addressing this issue, we first compare the cointegrating relations for the MZM and M2 models.[15] The cointegrating equations (CEs)—defined as

[14] In an earlier version of this paper, Carlson et al., (1998), we examined these questions after allowing for the possibility that more than one cointegrating vector characterizes this trivariate system. A comparison of the cointegration trace tests in Table 6 with standard critical values provides mixed evidence on the issue of one or two cointegrating relations. Extensive tests, both likelihood ratio and Horvath-Watson tests for prespecified vectors (velocity and opportunity costs), fail to reveal definitive evidence that the system is characterized by either one or two vectors. The test statistics compiled in these experiments are available on request. Our earlier work suggests that the basic conclusions (predictive content and resilience to shocks) are essentially unaltered in the one- and two-vector specifications. To facilitate comparison with prior literature we analyze the dynamics of the system using only the one-vector specification.

[15] Again, relations for M2M-based measures (not depicted) are observationally indistinguishable from the MZM diagrams. Also, the statistical tests applied to the M2M model yield results that are very similar to

$\ln(M / P)_t - \beta_1 \ln(y) + \beta_2 opp\cos t_t$ and formed from the vector estimates obtained over the 1964:1 to 1998:12 samples—appear in Figures 4 through 6. They are normalized to mean zero to facilitate comparison with the M2 relation (ZM2PI6, ZM2IP6, and ZM2CN6) adjusted for effects of the t9094 financial innovation proxy, and the MZM relation adjusted for the ratchet variable. These CEs are all based upon the six-lag specifications obtained with the chain-weighted price index for deflating the monetary aggregates and personal income. The estimated coefficients used to form these CEs appear in the lower panel of Table 6 for the M2 CEs. The estimates for the MZM CEs were obtained with the ratchet variable. Results for the MZM 1964-1998 sample, estimated with the ratchet variable discussed in footnote 8, appear in Table 8. The evidence for cointegration in the MZM models using the ratchet variable is comparable to that observed for MZM in the 1976-1998 sample. In each of the lag specifications and for all aggregates, there is evidence of cointegration: the income elasticities are in the proximity of one, and the interest semi-elasticity is relatively high.

Figures 4 through 6 reveal that, regardless of the measure used to proxy scale, the cointegrating equation for M2 yields substantially smaller errors than does its MZM counterpart. Specifically, the standard errors of the M2 CE errors are about 2% over the entire sample, with the CE based upon industrial production exhibiting slightly higher errors and the CE based upon personal income slightly lower errors. The standard errors of the corresponding MZM-based CE errors are approximately 2.5 times higher. This differential remains even in the last ten years of the sample, except in the case of the models scaled with industrial production, where the standard errors of the cointegrating equations are about the same in the M2 and MZM models. These results suggest that departures from the estimated long-run equilibria may be the smallest in models based upon the M2 aggregate, with sharp departures from equilibrium observed in the

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those forthcoming from the MZM specification. For clarity, we choose to present only the results from the MZM analysis.

MZM specification during periods where interest rates are high and volatile. However, there is a general tendency for the errors in the equations to move together, with both displaying a positive deviation from zero at the end of the sample.

4.2. Equilibrium nominal income

It remains to be determined whether the stable cointegrating relations depicted in Figures 4 through 6 yield predictive content for the short-run behavior of nominal income. We derive an expression for equilibrium nominal income from our basic cointegrating relation:

$$-\ln(M / P)_t + \beta_1 \ln(y) - \beta_2 opp\cos t_t = -\varepsilon_t .$$

Define $Nomy_t \equiv \ln(P)_t + \ln(y)_t$, and when the income elasticity $\beta_1 = 1$, nominal income may be isolated in the cointegrating equation. In this case, let $Nomy_t^* \equiv \ln(M)_t + \beta_2 opp\cos t_t$ so that $Nomy_t - Nomy_t^* = -\varepsilon^*_t$ denotes deviations from the cointegrating equation.

We limit our "predictive content" analysis of the equilibrium nominal income concept to models based upon MZM and M2 estimated with the chain-weighted deflator, where generally more evidence of unitary income elasticity is observed across the specifications. The results in Tables 4, 6, and 8 suggest that the income elasticities for the coincident index and for industrial production are generally not significantly different from 1.0, as implied by the nominal income specification above, when the chain-weighted price index is used to proxy price movements. The personal income elasticity is not significantly different from 1.0 in specifications that allow for a linear trend in the cointegrating relation.

We re-estimated the long-run relations based upon the PC deflator specifications with the unitary income elasticity in place for each of the monetary aggregate specifications. Two separate specifications are employed for each aggregate. Estimates obtained from a bivariate specification of M2 velocity and opportunity cost appear in the top panel of Table 9. In this case, the unitary income elasticity restriction is imposed on the long- and short-run representation of the system. These results are consistent with a cointegrating vector with interest semi-elasticity about the same as that observed in the unconstrained models. In the lower panel of Table 9 we present the interest elasticities estimated with a trivariate M2 model with the unitary income elasticity imposed only as a long-run restriction. Test statistics and estimated interest elasticities for this velocity-constrained specification also appear in Table 9. Likelihood ratio test results suggest that there is considerable support for a velocity-constrained long-run M2 relation, and the ADF tests suggest that the constrained relation is indeed stationary.

Table 10 parallels the Table 9 results for the MZM chain weighted price specification using the “post-ratchet” 1976-1998 sample. Generally the estimated interest elasticities are not affected by the imposition of the unitary income elasticity. However, the likelihood ratio statistics reveal some evidence against the unitary income elasticity restriction. We proceed with the analysis with the restriction in place since the ADF statistics suggest that the MZM velocity restricted relation exhibits stationarity.

We extracted a cointegrating equation (based upon the six-lag specification) from the lower panels of Tables 9 and 10 for each of the three scale specifications. Deviations from the M2 equilibrium nominal income values, now defined as *Nomy-Nomy** (ZVM2CN6, ZVM2IP6, and ZVM2PI6) appear in Figure 7, and deviations from the MZM equilibrium nominal income values (ZVMZMN6, ZVMZMIP6, and ZVMZMPI6) appear in Figure 8. These deviations are comparable to the negative of the CE representations in Figures 4 through 6. In each case, the *Nomy** estimate is adjusted for the broken trend t9094 and, in the case of personal income, for a linear trend throughout the sample. As observed in Figures 4-6, the MZM deviations are considerably larger than M2 counterparts. Also, deviations for both specifications are highly correlated and generally negative in the most recent part of the sample.

We explored the predictive content of these *Nomy-Nomy** deviations using the nominal income equation from a VECM formed from the nominal representation of each of the following:

ln(y)+ln(P)=Nomy, ln(m), and oppcost.[16] The cointegrating relations used in this investigation are taken from the "real" specification (ln(y), ln(M/P), and oppcost) with estimates displayed in Tables 9 and 10. With unitary income elasticity in place, deviations from the "real" and "nominal" long-run representations are of course identical. The question we pose in this section is simply: Do deviations from a textbook representation of money demand, estimated from the "real" representation of the system, have predictive content for the short-run behavior of nominal income?

The nominal income growth equation used as a basis for these tests consists of six or nine lags of nominal income growth, six or nine lags of nominal money growth, and six or nine lags of opportunity cost changes and is estimated with and without separately including lags of Δ *ln(P)*. The structure was augmented with the CEs discussed above.[17]

Table 11 contains standard F-tests for the significance of the lagged M2 growth terms and the error correction term in the nominal income equation based upon all three income measures for the period 1964:1-1998:12. Test statistics are formed from autoregressive specifications that exclude and include changes in the log of the PC deflator as a separate regressor. The experiments reveal that M2 maintains predictive content. The error-correction

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[16] A nominal representation of the system is obtained by specifying the long-run "money demand" relationship in its nominal form and imposing unitary income elasticity: $-\ln(M)_t + \ln(y * P) - \beta_2 opp\cos t_t = -\varepsilon_t$, so that the relevant variables in the VECM are nominal balances, nominal income, and opportunity cost. This representation is consistent with the formulations used by Estrella and Mishkin (1997) and Feldstein and Stock (1996).

[17] It is noteworthy that the nominal income representation we examine is not simply the transformation of the output equation in the VECM that we used to estimate the long-run money demand equation in the preceding sections. In that sense, the empirical question we pose in this section is whether deviations from the long-run "real" representation of the system have predictive content for a distinct short-run nominal income specification. Of course, an observationally equivalent long-run relation could be estimated from the outset as the nominal relation: $-\ln(M)_t + \ln(y * P) - \beta_2 opp\cos t_t = -\varepsilon_t$. In this case, the corresponding VECM is comprised of nominal balances, nominal income, and opportunity costs. The nominal income equation in this VECM serves as the foundation for our experiments in this section, but the long-run relations used to form the CEs are taken from the real representation of the system used throughout this paper. In separate exercises, we have re-estimated all long-run structures using this nominal representation and found that the estimates of the long-run relations are very similar to those obtained from the "real" representations. Results of these exercises are available on request.

term is statistically significant regardless of scale specification and lag length or whether inflation is included separately in the nominal income equation. The lagged M2 growth terms are also significant in portending positive growth in nominal income.

The test statistics in Table 12 summarize the case for predictive content using an analogous VECM specification for MZM. The sample and equilibrium estimates correspond to Table 10. There remains some evidence of predictive content in the MZM specifications, though the case is not as strong as it is for M2. Error correction terms are significant only when personal income is used to measure the scale of the economy. Also, there is information content for the lagged money variables when the coincident indicator and industrial production are used to measure scale.

Extensive experiments with alternative specifications and samples lead to the conclusion that is forthcoming from Tables 11 and 12. The M2 models reveal significant predictive content, while specifications based upon the MZM model yield modest evidence of predictive content for fluctuations in nominal income.

Why do M2 models and MZM models generate different conclusions? First, the M2 models that offer the most convincing evidence of predictive content are estimated over 1964:1-1998:12. Experiments that confine the analysis to post-1975 or even post-1983 samples (not depicted) offer some support, but not the strong case that is contained in Table 11. Lagged M2 growth terms are not significant in the shortened sample, similar to the conclusions reached by Estrella and Mishkin. The sum of the lagged money growth coefficients remains positive, but is statistically significant only in the coincident indicator specification. However, the error correction specification clearly makes a difference, as illustrated by Feldstein and Stock. Even in the post-1983 sample, deviations of nominal income from the values predicted by the estimated long-run equilibrium (*Nomy-Nomy**) do have predictive content in the M2 specifications. Short-run nominal income growth is clearly correlated with recent deviations of nominal income from a measure of *Nomy** that is defined after accounting for financial innovations of the 1990s as discussed above.

The evidence remains weaker for MZM even when we conduct the predictive content experiments (not depicted) with the long sample (1964-1998, including ratchet) or a shorter, post-1983 sample. The evidence in the long sample is about the same as that portrayed for the post-1976 sample in Table 12. The evidence for the post-1983 sample is slightly stronger with two of the scale measures, industrial production and personal income, yielding specifications that portend predictive content.

We speculate that there is less evidence of predictive content in the MZM-based models because the interest elasticity estimates are significantly larger–making the deviations from the money demand equilibrium larger and longer-lived. This does not necessarily imply that deviations from an MZM money demand equilibrium are a completely uninformative signal. Indeed as discussed above the MZM deviations are highly correlated with the M2 deviations. However, they may contain considerably more "noise" than their M2 counterparts. We explore the resilience of these relations to stochastic shocks in the section below.

4.3. Dynamics

Another gauge of the robustness of the money demand specification is to determine how quickly the relation regains stability following shocks or innovations that impact the economy. We applied a variation of the original King, Plosser, Stock, and Watson (1991) approach to identify the fundamental innovations in our cointegrating system and then tracked the response of the equilibrium money demand relationship to these innovations.[18]

King et al. apply the Granger Representation Theorem to express the moving average representation of a p dimensional cointegrated system with rank r as a function of p-r permanent and r distinct transitory innovations. In our specification this implies that the long-run properties of the data are dictated by two permanent innovations. We identify them as a technology shock normalized to ultimately increase real output by 1% in the long run and an innovation that leaves a permanent positive imprint on opportunity cost in the long run. The remaining shock in the system is defined as a shock to the money demand relation, which is, of course, transitory.

The response of selected variables in the system to these innovations appears in Figures 9 through 11. There is a separate graph for each of the shocks that are identified. In separate panels of each graph we have included the responses with all three scale variables: the coincident indicator, industrial production, and real personal income. Each panel provides a comparison of the responses in the model using M2 (the solid lines) with that in the model using MZM (dashed lines). The right-hand panels in each graph show the adjustment of the cointegrating vector to the various shocks. The left-hand panels show the adjustment of the variable on which the shock is normalized.

The responses of the scale variable to the technology shock in Figure 9 indicate approximately the same response patterns regardless of scale measure and regardless of which measure of money appears in the model. The technology shock is defined to leave a 1% imprint on the scale variable in the long run, and will have no long-run impact on the cointegrating vector, since this variable is stationary. The interesting patterns are in the short- and intermediate-run responses. In five of the six cases, it takes several months before a positive technology shock begins to influence the scale variable, the exception being the MZM model with real personal income. In the latter case, the initial impact of this shock on real income is a substantial fraction of the ultimate impact. In all cases, the response of the scale variable effectively achieves its long-run value in 18-24 months.

The most interesting response pattern is observed in the right-hand panels, where it can be seen that departures from the money demand equilibrium for the M2 models are

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[18] For the technical details of the procedures used to identify the innovations, see Hoffman and Rasche

inconsequential after about a six-month period. Over the first twelve months, departures in the equilibrium to shocks that leave a long-run imprint on the scale variable of 1% result are no more than 0.3%. Clearly, departures of the M2 demand equilibrium due to a technology shock are small.

The short-run departures from the money demand equilibrium for the MZM models are larger than the corresponding measures for the M2 model and are somewhat more persistent. In the models that use the coincident indicator or industrial production, these deviations have returned close to zero within a year after the shock. The only model with considerable persistence of the deviations from money demand equilibrium is that for MZM with real personal income as the scale variable.

The impulse response pattern from the opportunity cost shock is portrayed in Figure 10. This graph shows the responses to the second permanent shock in the model, which we have labeled an opportunity cost shock because it is scaled to have a long-run effect of 1% on opportunity costs.[19] The left-hand panels in Figure 10 show the dynamic responses of the opportunity cost variable in all of the models to this shock. In every case, the immediate effect is to increase the opportunity cost measure. The effect of the shock on this variable then damps out to the equilibrium in a cyclical fashion. In some cases, the impact effect is larger than the equilibrium; in some cases, it is smaller. But in all cases, there is some overshooting of the equilibrium change as the adjustment progresses over time. The right-hand panels of Figure 10 indicate the response of the money demand equilibrium to this shock. In all cases, the effects are qualitatively the same. The initial effect is a positive deviation from the long-run equilibrium, which is reversed within about six months. The negative deviations increase in absolute value

(1996) and Rasche (1997).

[19] While we observe considerable persistence in the opportunity cost data, it is difficult to explain why, absent regulatory restrictions on deposit rates or changes in reserve requirements, own rates do not eventually adjust to permanent shocks to effective rates of return on Treasury bills.

until about nine months, after which the effects die off. Most of the deviations from equilibrium have been eliminated after eighteen months.

Figure 12 shows the impulse response patterns for the single transitory shock in the model. This shock has been scaled so that it has an impact effect of one percent on real balances. Since it is a transitory shock, its long-run effect on all variables is zero. The response functions of real balances are shown in the left-hand panels of Figure 12 for all three scale variables and both measures of money. The effects here on real balances are persistent, generally taking around two to three years to die out. In two cases (the M2 model with the coincident indicator and the MZM model with personal income), the maximum response of real balances is not felt immediately. In the former case the maximum response is delayed over a year; in the latter case about six months. The responses of the money demand equilibrium are shown in the left-hand panels. For all three scale variables, the immediate money demand disequilibrium is larger in absolute value for MZM than it is in the equivalent M2 model. This reflects a larger response of the demand for MZM to changes in its opportunity cost than is the case for the demand for M2. The initial disequilibrium of M2 demand varies in sign, depending on the scale variable used in the model. In contrast, the initial deviations in the MZM models are always positive. When industrial production is used as the scale variable, the money demand disequilibrium dies out within a year for both measures of money. When the coincident indicator or personal income is used as the scale variable, the disequilibrium is more persistent, but in all cases is quite small after about two years have elapsed. When viewed across the three scale variables, there is no unique ranking of the two money shocks based on the persistence of money demand disequilibrium to this type of shock.

Our analysis of fundamental relations that underlie the cointegrated system reveals that the money demand equilibrium is quite resilient to fundamental innovations. A stable and reliable relationship linking a broad monetary aggregate to the nominal economy must be resilient to shocks that commonly occur. The analysis of this section reveals that the M2 and MZM relationships offer substantial support for such resiliency.

Conclusion

During the 1980s, M2 evolved to become the primary intermediate target of monetary policy. Since the early 1990s, however, the reliability of money measures as targets or indicators of monetary policy has been called into question. In 1993 the FOMC "downgraded" the role of M2, and no single variable has been identified to take its place. This action finds support in Estrella and Mishkin (1997), who conclude that whatever their informational content may have been in earlier time periods, the monetary aggregate—the monetary base and M2 in particular—do not seem to provide adequate and consistent information at present in the United States. The results of this paper suggest that such a conclusion seems too extreme.

Specifically, we find strong evidence of a stable "money demand" relationship for MZM and M2M through the 1990s. Though the M2 relation breaks down somewhere around 1990, evidence has been accumulating that the disturbance is well characterized as a permanent upward shift in M2 velocity, which began around 1990 and was largely over by 1994. Taken together, our results support the hypothesis that households permanently reallocated a portion of their wealth from time deposits to mutual funds. Although this reallocation may have been largely induced by depository restructuring, we argue that the substitution could be explained by appropriately measured opportunity cost. Unfortunately, data do not exist that would allow a test of this hypothesis.

In the absence of an appropriate measure of opportunity cost, we construct a linear shift variable, which we based on evidence of independent studies that showed the effects were contained within a period beginning in 1990 and ending before 1994. When we control for such a shift, we find that M2 continues to have information content consistent with the pre-1990 sample. Our results are robust to three distinct scale measures—the coincident indicator series, industrial production, and personal income—and two price measures—the CPI and the chained-weighted personal income deflator.

In separate experiments, we examine the predictive content of both M2- and MZM-based money demand relations for short-run fluctuations in nominal income. We also monitor the response of the relation to fundamental innovations identified as real and nominal impulses. The specifications employed in the analysis are quite extensive. We examine three distinct scale measures in samples that span up to 30 years, using two separate autoregressive specifications of each model. Our results suggest that there is indeed an argument for reconsideration of an aggregate money demand function as an anchor in a model linking monetary aggregates to nominal economic activity. Departures of actual nominal income from equilibrium nominal income estimates extracted from the money demand relationships clearly have predictive content over a sample period inclusive of both the new operating procedures experiment (1979:10-1982:8) and the depository restructuring of the early 1990s. The results are very consistent across the various alternative specifications for M2 but less so for MZM. Moreover, the relationships we identify are quite resilient to innovations identified in the analysis. Generally, the long-run money demand relationship adjusts to regain equilibrium within an eighteen month period following the innovations we identify in the analysis.

We recognize, however, that velocity shifts like that occurring to M2 in the 1990s are not all that uncommon across monetary aggregates. Estrella and Mishkin (1997) suggest that a way of characterizing this situation is to think of velocity shocks as noise that obscures signals from monetary aggregates. In a regime like that of 1990s, in which nominal income and money growth and inflation are all subdued, the signal-to-noise ratio is likely to be low, making monetary aggregates poor guides for monetary policy. As Estrella and Mishkin also note, there may be time periods where pronounced changes in money and inflation produce a high signal-to-noise ratio. The challenge to policymakers is to understand the consequences of these episodes and to do a better job anticipating them. It would be tragic if policy makers waited for high inflation and money growth before they recognized any signal in the aggregates.

Finally, our results do not lead us to conclude that M2 should immediately regain its intermediate target status. With limited evidence on the re-emergence of a stable long-run velocity, such reliance on M2 would be unwarranted. Targeting M2 is one thing. Ignoring it, however, is yet another. We interpret our results as suggesting that it would be unwise to ignore persistent periods of above-trend M2 growth. Ignoring M2 altogether would be especially unwise if one believes that unpredictable velocity shocks are most likely to be the product of financial innovation, which in turn leads to permanent increases in velocity. Under such a situation, policy makers may be more inclined to tolerate slow money growth, as they did in the early 1990s. In contrast, tolerating rapid M2 growth in the face of a permanent upward shift in velocity is a recipe for inflation.

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Table 1
M2 1964:1-1989:12

	Johansen Trace Tests			Real balances	Scale Measure			oppcost	time
Ar lag	**H0:r = 0**	**H0:r = 1**	**H0:r = 2**	**Ln(m/p)**	**log (cn)**	**log(ip)**	**ln(pi/p)**	**ln(1+rtbe)-ln(1+ownr)**	**t**
				CPI Deflator					
6	35.7	9.8	.5	1	-.859			4.692	
					(.03)			(.63)	
9	25.8	6.6	.4	1	-.857			4.283	
					(.03)			(.62)	
6	38.1	11.4	1.4	1		-.821		4.422	
						(.02)		(.46)	
9	26.9	8.5	2.3	1		-.817		4.374	
						(.02)		(.52)	
6	65.1	21.6	1.3	1			-.813	2.904	
							(.01)	(.20)	
9	54.4	14.8	1.2	1			-.814	2.790	
							(.01)	(.19)	
6(t)	77.9	29.1	6.8	1			-.960	2.670	.00038
							(.04)	(.16)	
9(t)	67.4	22.1	5.9	1			-.951	2.609	.00032
							(.04)	(.15)	
				PC Deflator					
6	27.4	12.4	.6	1	-1.025			4.230	
					(.04)			(1.1)	
9	22.0	9.1	.6	1	-1.043			3.023	
					(.03)			(.61)	
6	35.7	14.8	1.4	1		-1.021		2.418	
						(.02)		(.42)	
9	28.5	12.1	2.8	1		-1.016		2.604	
						(.02)		(.46)	
6	60.4	19.4	2.3	1			-.844	2.485	
							(.01)	(.18)	
9	52.6	12.4	2.6	1			-.842	2.531	
							(.01)	(.16)	
6(t)	68.5	21.9	7.1	1			-.994	2.552	.00042
							(.05)	(.15)	
9(t)	61.9	16.3	6.7	1			-.971	2.583	.00036
							(.04)	(.13)	

Note: Lag lengths pertain to the number of lags in the VECM specification of the model. The 5% critical values for the deterministic cointegration tests are 29.7, 15.4, and 3.7 and the 5% critical values for the trend in the vector cointegration tests (denoted by (t)) are 34.6, 18.2 and 3.7. Standard errors for the estimated long-run parameters appear in parentheses.

Table 2
M2 1964:1-1998:12

	Johansen Trace Tests			Real balances	Scale Measure			oppcost	Time
Ar lag	**H0:r = 0**	**H0:r = 1**	**H0:r = 2**	**Ln(m/p)**	**log (cn)**	**log(ip)**	**ln(pi/p)**	**ln(1+rtbe)-ln(1+ownr)**	**T**
				CPI Deflator					
6	21.8	6.9	.1	1	-.623			7.345	
					(.09)			(2.5)	
9	24.6	10.6	.2	1	-.681			5.341	
					(.06)			(1.6)	
6	22.7	5.6	.1	1		-.555		9.211	
						(.08)		(3.0)	
9	25.7	10.6	.1	1		-.600		9.175	
						(.08)		(3.1)	
6	29.2	8.8	.1	1			-.226	18.37	
							(.53)	(18)	
9	28.7	11.1	.5	1			-2.90	-105.5	
							(14.5)	(707)	
6(t)	36.4	16.4	6.8	1			-3.86	-10.06	.0066
							(3.2)	(16.3)	
9(t)	36.7	17.8	5.4	1			-2.72	-7.25	.0043
							(1.4)	(9.1)	
				PC Deflator					
6	20.1	7.0	.1	1	-.666			18.36	
					(.33)			(18.1)	
9	25.5	9.4	.1	1	-.771			134.4	
					(1.78)			(97.1)	
6	22.3	7.1	.1	1		-.652		15.78	
						(.18)		(9.74)	
9	27.2	9.2	.1	1		-.776		86.69	
						(.80)		(309.7)	
6	22.6	9.0	1.2	1			-.461	9.622	
							(.23)	(6.63)	
9	22.9	11.1	1.4	1			-.108	31.85	
							(1.6)	(77.3)	
6(t)	31.2	18.6	7.5	1			-2.156	2.013	.0035
							(.37)	(1.04)	
9(t)	30.2	18.4	7.7	1			-2.729	-3.729	.0049
							(1.7)	(9.10)	

Note: Lag lengths pertain to the number of lags in the VECM specification of the model. The 5% critical values for the deterministic cointegration tests are 29.7, 15.4, and 3.7 and the 5% critical values for the trend in the vector cointegration tests (denoted by (t)) are 34.6, 18.2 and 3.7. Standard errors for the estimated long-run parameters appear in parentheses.

TABLE 3
M2M 1976:1-1998:12

	Johansen Trace Tests			Real balances	Scale Measure			Oppcost
Ar lag	**H0:r = 0**	**H0:r = 1**	**H0:r = 2**	**ln(m/p)**	**log (cn)**	**log(ip)**	**ln(pi/p)**	**ln(1+rtb3)- ln(1+ownr)**
				CPI Deflator				
6	44.5	7.2	.1	1	-.858			4.757
					(.03)			(.22)
9	33.5	5.8	.1	1	-.853			4.792
					(.03)			(.22)
6	36.4	7.0	.1	1		-.717		5.245
						(.03)		(.27)
9	28.0	6.5	.1	1		-.700		5.539
						(.03)		(.31)
6	42.5	10.4	.1	1			-.860	4.055
							(.04)	(.29)
9	31.9	6.7	.1	1			-.869	3.931
							(.04)	(.29)
				PC Deflator				
6	39.1	7.4	.1	1	-.991			4.078
					(.03)			(.22)
9	31.9	6.4	.1	1	-.984			4.133
					(.03)			(.21)
6	27.1	7.2	.1	1		-.818		4.83
						(.04)		(.35)
9	24.3	7.1	.1	1		-.995		5.190
						(.04)		(.41)
6	40.6	7.6	.1	1			-.872	4.020
							(.03)	(.25)
9	34.1	5.7	.1	1			-.871	4.037
							(.03)	(.23)

Note: Lag lengths pertain to the number of lags in the VECM specification of the model. The 5% critical values for the deterministic cointegration tests are 29.7, 15.4, and 3.7 and the 5% critical values for the trend in the vector cointegration tests (denoted by (t)) are 34.6, 18.2 and 3.7. Standard errors for the estimated long-run parameters appear in parentheses.

Table 4 MZM
(1976:1-1998:12)

	Johansen Trace Tests			Real balances		Scale Measure		Oppcost
Ar lag	**H0:r = 0**	**H0:r = 1**	**H0:r = 2**	**ln(m/p)**	**log (cn)**	**log(ip)**	**ln(pi/p)**	**ln(1+rtbe)-ln(1+ownr)**
CPI Deflator								
6	38.0	7.4	.1	1	-1.076 (.03)			4.660 (.25)
9	29.3	6.4	.1	1	-1.072 (.03)			4.705 (.26)
6	33.4	7.2	.1	1		-.898 (.03)		5.274 (.29)
9	27.2	7.7	.1	1		-.881 (.04)		5.569 (.33)
6	34.8	8.9	.1	1			-1.106 (.05)	3.52 (.39)
9	25.7	5.7	.1	1			-1.116 (.06)	3.430 (.43)
PC Deflator								
6	35.7	7.3	.1	1	-1.210 (.03)			3.947 (.24)
9	30.3	7.3	.2	1	-1.200 (.03)			4.030 (.24)
6	26.9	7.8	.2	1		-1.002 (.04)		4.814 (.37)
9	25.1	9.1	.3	1		-.979 (.05)		5.164 (.42)
6	35.0	6.5	.1	1			-1.072 (.04)	3.829 (.30)
9	29.5	5.6	.2	1			-1.068 (.04)	3.872 (.28)

Note: Lag lengths pertain to the number of lags in the VECM specification of the model. The 5% critical values for the deterministic cointegration tests are 29.7, 15.4, and 3.7 and the 5% critical values for the trend in the vector cointegration tests (denoted by (t)) are 34.6, 18.2 and 3.7. Standard errors for the estimated long-run parameters appear in parentheses.

Table 5
M2 1976:1-1998:12

	Johansen Trace Tests			Real balances	Scale Measure			oppcost	time
Ar lag	**H0:r = 0**	**H0:r = 1**	**H0:r = 2**	**ln(m/p)**	**log (cn)**	**log(ip)**	**ln(pi/p)**	**ln(1+rtbe)-ln(1+owr)**	**t**
					CPI Deflator				
6	15.3	4.0	.1	1	-.436			6.352	
					(.10)			(1.51)	
9	20.1	6.7	.1	1	-.421			7.081	
					(.10)			(1.71)	
6	16.1	3.2	.1	1		-.378		7.056	
						(.08)		(1.56)	
9	23.8	6.2	.1	1		-.361		7.705	
						(.07)		(1.48)	
6	14.9	4.4	.1	1			-.122	13.12	
							(.48)	(11.0)	
9	18.6	5.4	.1	1			-.2.99	-65.61	
							(14.6)	(403.6)	
6(t)	26.1	13.2	4.2	1			-3.284	-1.460	.0052
							(1.19)	(2.6)	
9(t)	29.2	15.2	5.3	1			-3.067	-4.741	.0046
							(1.5)	(6.1)	
					PC Deflator				
6	15.8	3.9	.1	1	-.433			8.971	
					(.20)			(4.14)	
9	22.9	6.5	.1	1	-.366			11.447	
					(.25)			(6.01)	
6	17.9	4.2	.1	1		-.414		8.474	
						(.12)		(2.6)	
9	26.1	6.8	.1	1		-.341		10.975	
						(.15)		(3.9)	
6	11.3	3.8	.1	1			-.323	8.017	
							(.26)	(4.87)	
9	15.6	5.2	.1	1			14.111	359.51	
							(560)	(13635)	
6(t)	25.7	10.2	3.1	1			-2.291	3.167	.0037
							(.27)	(.42)	
9(t)	28.9	15.2	4.9	1			-2.493	3.240	.0042
							(.30)	(.46)	

Note: Lag lengths pertain to the number of lags in the VECM specification of the model. The 5% critical values for the deterministic cointegration tests are 29.7, 15.4, and 3.7 and the 5% critical values for the trend in the vector cointegration tests (denoted by (t)) are 34.6, 18.2 and 3.7. Standard errors for the estimated long-run parameters appear in parentheses.

Table 6
M2 with t90-94 1964:1-1998:12

	Johansen Trace Tests			Real balances	Scale Measure			Oppcost	time
Ar lag	**H0:r = 0**	**H0:r = 1**	**H0:r = 2**	**ln(m/p)**	**Log (cn)**	**log(ip)**	**ln(pi/p)**	**ln(1+rtbe)-ln(1+ownr)**	
					CPI Deflator				
6	35.5	13.0	1.1	1	-.856			3.178	
					(.02)			(.40)	
9	31.3	11.3	.9	1	-.847			2.782	
					(.03)			(.45)	
6	51.0	16.5	2.8	1		-.818		3.885	
						(.02)		(.30)	
9	45.0	17.1	4.6	1		-.816		3.553	
						(.02)		(.34)	
6	49.0	18.6	.8	1			-.828	1.642	
							(.01)	(.31)	
9	44.3	15.4	.5	1			-.819	1.606	
							(.01)	(.32)	
6(t)	60.6	22.7	8.7	1			-1.070	1.894	.00083
							(.06)	(.20)	
9(t)	54.7	19.9	7.0	1			-1.037	1.814	.00077
							(.06)	(.22)	
					PC Deflator				
6	32.5	12.1	1.6	1	-1.075			1.285	
					(.02)			(.48)	
9	31.9	11.1	1.6	1	-1.058			.818	
					(.03)			(.68)	
6	47.8	17.4	3.9	1		-1.006		2.545	
						(.02)		(.41)	
9	48.9	17.2	5.7	1		-1.003		1.979	
						(.02)		(.49)	
6	45.5	14.8	2.5	1			-.856	1.632	
							(.01)	(.24)	
9	40.6	13.1	2.6	1			-.846	1.689	
							(.01)	(.24)	
6(t)	52.1	15.8	5.1	1			-1.073	2.091	.00087
							(.07)	(.19)	
9(t)	47.8	15.8	5.9	1			-1.033	2.058	.00077
							(.06)	(.19)	

Note: Lag lengths pertain to the number of lags in the VECM specification of the model. The 5% critical values for the deterministic cointegration tests are 29.7, 15.4, and 3.7 and the 5% critical values for the trend in the vector cointegration tests (denoted by (t)) are 34.6, 18.2 and 3.7. Standard errors for the estimated long-run parameters appear in parentheses.

Table 7
M2 with t90-94 1976:1-1998:12

	Johansen Trace Tests			Real balances	Scale Measure			oppcost	time
Ar lag	H0:r = 0	H0:r = 1	H0:r = 2	ln(m/p)	Log (cn)	log(ip)	ln(pi/p)	ln(1+rtbe)-ln(1+ownr)	
				CPI Deflator					
6	38.1	11.1	.1	1	-.839			3.245	
					(.03)			(.20)	
9	39.7	14.1	1.8	1	-.824			3.348	
					(.02)			(.18)	
6	43.3	12.9	1.5	1		-.799		3.851	
						(.03)		(.25)	
9	45.4	18.4	2.7	1		-.814		3.643	
						(.04)		(.30)	
6	24.5	7.1	.1	1			-.864	1.424	
							(.05)	(.51)	
9	25.1	5.9	.3	1			-.851	1.325	
							(.04)	(.59)	
6(t)	42.7	13.2	5.2	1			-1.340	1.867	.0016
							(.11)	(.21)	
9(t)	43.3	13.9	3.6	1			-1.297	1.872	.0015
							(.10)	(.21)	
				PC Deflator					
6	42.3	11.9	1.5	1	-.991			2.386	
					(.02)			(.16)	
9	45.1	16.7	2.1	1	-.983			2.409	
					(.02)			(.17)	
6	39.8	14.4	1.6	1		-.964		2.756	
						(.05)		(.39)	
9	44.6	16.1	2.5	1		-1.059		1.515	
						(.09)		(.76)	
6	24.9	6.9	1.5	1			-.874	1.783	
							(.03)	(.29)	
9	26.9	8.1	1.9	1			-.861	1.763	
							(.03)	(.29)	
6(t)	37.6	9.2	.5	1			-1.407	2.414	.0018
							(.14)	(.17)	
9(t)	41.4	12.9	1.6	1			-1.347	2.391	.0017
							(.13)	(.15)	

Note: Lag lengths pertain to the number of lags in the VECM specification of the model. The 5% critical values for the deterministic cointegration tests are 29.7, 15.4, and 3.7 and the 5% critical values for the trend in the vector cointegration tests (denoted by (t)) are 34.6, 18.2 and 3.7. Standard errors for the estimated long-run parameters appear in parentheses.

Table 8
MZM with ratchet 1964:1-1998:12

	Johansen Trace Tests			Real balances	Scale Measure			oppcost	time
Ar lag	**H0:r = 0**	**H0:r = 1**	**H0:r = 2**	**ln(m/p)**	**Log (cn)**	**log(ip)**	**ln(pi/p)**	**ln(1+rtbe)-ln(1+ownr)**	
				CPI Deflator					
6	38.0	8.2	.1	1	-.979 (.06)			5.930 (.50)	
9	34.0	9.7	.1	1	-.973 (.06)			5.87 (.52)	
6	40.4	6.2	.1	1		-.846 (.04)		6.150 (.43)	
9	42.6	8.6	.1	1		-.840 (.04)		6.225 (.40)	
6	36.9	12.2	.4	1			-.958 (.07)	5.600 (.62)	
9	30.6	12.8	.1	1			-1.019 (.08)	4.904 (.64)	
				PC Deflator					
6	32.9	9.4	.1	1	-1.127 (.06)			5.230 (.53)	
9	32.9	11.5	.1	1	-1.112 (.06)			5.169 (.52)	
6	34.4	7.6	.1	1		-.966 (.05)		5.678 (.50)	
9	39.6	10.7	.3	1		-.958 (.05)		5.703 (.45)	
6	31.7	8.7	.1	1			-1.018 (.06)	4.940 (.56)	
9	27.2	9.3	.1	1			-1.058 (.07)	4.397 (.61)	

Note: Lag lengths pertain to the number of lags in the VECM specification of the model. The 5% critical values for the deterministic cointegration tests are 29.7, 15.4, and 3.7 and the 5% critical values for the trend in the vector cointegration tests (denoted by (t)) are 34.6, 18.2 and 3.7. Standard errors for the estimated long-run parameters appear in parentheses.

Table 9
M2 with t90-94 unitary income elasticity constraint personal consumption deflator 1964:1-1998:12

	Johansen TraceTests			**Real balances**		**Scale Measure**		**oppcost**
Ar lag	**H0:r = 0**	**H0:r = 1**		**ln(m/p)**	**Log (cn)**	**log(ip)**	**ln(pi/p)**	**Ln(1+tb)-ln(1+owr)**
				Bi-variate Specification				
6	16.9	1.7		1	-1.0			.878 (.80)
9	19.6	1.3		1	-1.0			.06 (1.2)
6	38.4	8.8		1		-1.0		2.443 (.42)
9	41.1	9.7		1		-1.0		1.915 (.50)
6 (t)	33.0	2.2		1			-1.0	1.780 (.23)
9 (t)	24.1	2.1		1			-1.0	1.612 (.33)

Trivariate Specification

Ar lag	**ln(m/p)**	**ln(cn)**	**ln(ip)**	**Ln(pi/p)**	**oppcost**	**trend**	**ADF***	$\chi^2_{(1)}$
6	1	-1			1.352 (.543)		-2.84	4.83
9	1	-1			0.331 (.662)		-3.11	2.30
6	1		-1		2.576 (.394)		-4.02	.82
9	1		-1		1.990 (.426)		-3.71	.02
6(t)	1			-1	1.959 (.177)	.0004 (.0002)	-4.47	1.03
9(t)	1			-1	2.009 (.163)	.0004 (.0002)	-4.12	.64

* McKinnon critical values for the ADF test are 1%: -3.44, 5%: -2.87

TABLE 10
MZM: unitary income elasticity constraint, personal consumption deflator, 1976:1-1998:12

	Johansen TraceTests			Real balances	Scale Measure			oppcost
Ar lag	**H0:r = 0**	**H0:r = 1**		**ln(m/p)**	**Log (cn)**	**log(ip)**	**ln(pi/p)**	**ln(1+tb)-ln(1+owr)**
				Bi-variate Specification				
6	10.3	.1		1	-1.0			5.063 (.49)
9	7.9	.2		1	-1.0			5.18 ( )
6	19.1	1.5		1		-1.0		4.612 (.32)
9	14.1	1.5		1		-1.0		4.626 (.37)
6(t)	30.3	3.2		1			-1.0	3.852 (.29)
9(t)	26.5	4.4		1			-1.0	3.880 (.28)

Trivariate Specification

Ar lag	ln(m/p)	ln(cn)	ln(ip)	ln(pi/p)	oppcost	ADF*	$\chi^2_{(1)}$
6	1	-1			5.418 (.493)	-2.99	17.50
9	1	-1			5.586 (.507)	-3.69	13.59
6	1		-1		4.829 (.318)	-4.42	.00
9	1		-1		5.021 (.344)	-4.13	.12
6	1			-1	4.165 (.250)	-3.51	3.68
9	1			-1	4.211 (.233)	-3.49	3.80

* McKinnon critical values for the ADF test are 1%: -3.44, 5%: -2.87

Table 11
Predictive content of M2 in nominal income growth regressions based upon chained personal income deflator specifications with t9094 1964:1-1998:12

Variable				
Scale Measure	**lag**	**(Nomy – Nomy*)$_{t-1}$**	**Δln M2$_{t-i}$**	**$\Sigma\Delta$ln M2$_{t-i}$**
CN: exclud ΔP	6	-.0190*	2.01	.2182*
		(.007)	(.06)	(.07)
	9	-.0149*	2.16*	.3060*
		(.005)	(.02)	(.08)
IP: exclud ΔP	6	-.0499*	2.12*	.5638*
		(.016)	(.05)	(.16)
	9	-.0630*	2.08*	.7528*
		(.015)	(.03)	(.19)
PI: exclude ΔP	6	-.0825*	1.36	.2016
		(.023)	(.23)	(.12)
	9	-.0750*	1.34	.2626*
		(.027)	(.21)	(.14)
CN: includ ΔP	6	-.0190*	1.87	.2109*
		(.007)	(.08)	(.07)
	9	-.0163*	2.39*	.3214*
		(.005)	(.01)	(.08)
IP: includ ΔP	6	-.0490*	1.97	.5453*
		(.016)	(.06)	(.16)
	9	-.0636*	2.17*	.7698*
		(.015)	(.02)	(.18)
PI: includ ΔP	6	-.0938*	1.71	.2332*
		(.023)	(.12)	(.11)
	9	-.0875*	1.77	.3148*
		(.027)	(.07)	(.14)

Note: Estimates for Nomy* are obtained from the trivariate specification in the lower panel of Table 9. In each case the long run equilibrium is demeaned and adjusted for the shift in velocity captured by t9094. In the personal income specification the vector is also adjusted for the possibility of a trend in the long run relationship. The * denotes significance at the 5% critical level.

Table 12
Predictive content of MZM in nominal income growth regressions based upon chained personal income deflator 1976:1-1998:12

Variable				
Scale Measure	**lag**	**$(Nomy - Nomy^*)_{t-1}$**	**$\Delta \ln M2_{t-i}$**	**$\Sigma\Delta \ln M2_{t-i}$**
CN: exclud ΔP	6	.0027	1.62	.0736*
		(.006)	(.14)	(.03)
	9	.0093	2.30*	.1143*
		(.006)	(.02)	(.04)
IP: exclud ΔP	6	-.0042	1.53	.1928*
		(.014)	(.17)	(.08)
	9	.0096	1.55	.3362*
		(.015)	(.13)	(.11)
PI: exclude ΔP	6	-.0355*	1.09	-.1236*
		(.012)	(.37)	(.06)
	9	-.0391*	.79	-.1368
		(.014)	(.62)	(.08)
CN: includ ΔP	6	-.0009	1.60	.0635*
		(.007)	(.4)	(.03)
	9	.0098	2.43*	.1159*
		(.006)	(.01)	(.04)
IP: includ ΔP	6	-.0046	1.51	.1885*
		(.32)	(.17)	(.08)
	9	.0119	1.55	.3547*
		(.015)	(.13)	(.11)
PI: includ ΔP	6	-.0396*	.80	-.0934
		(.012)	(.57)	(.06)
	9	-.0429*	.88	-.1035
		(.014)	(.54)	(.08)

Note: Estimates for Nomy* are obtained from the trivariate specification in the lower panel of Table 9. In each case the long run equilibrium is demeaned. The * denotes significance at the 5% critical level

Figure 1. Log of M2 velocity (using the coincident index for scale) and the opportunity cost of M2 balances, 1964:1-1998:12.

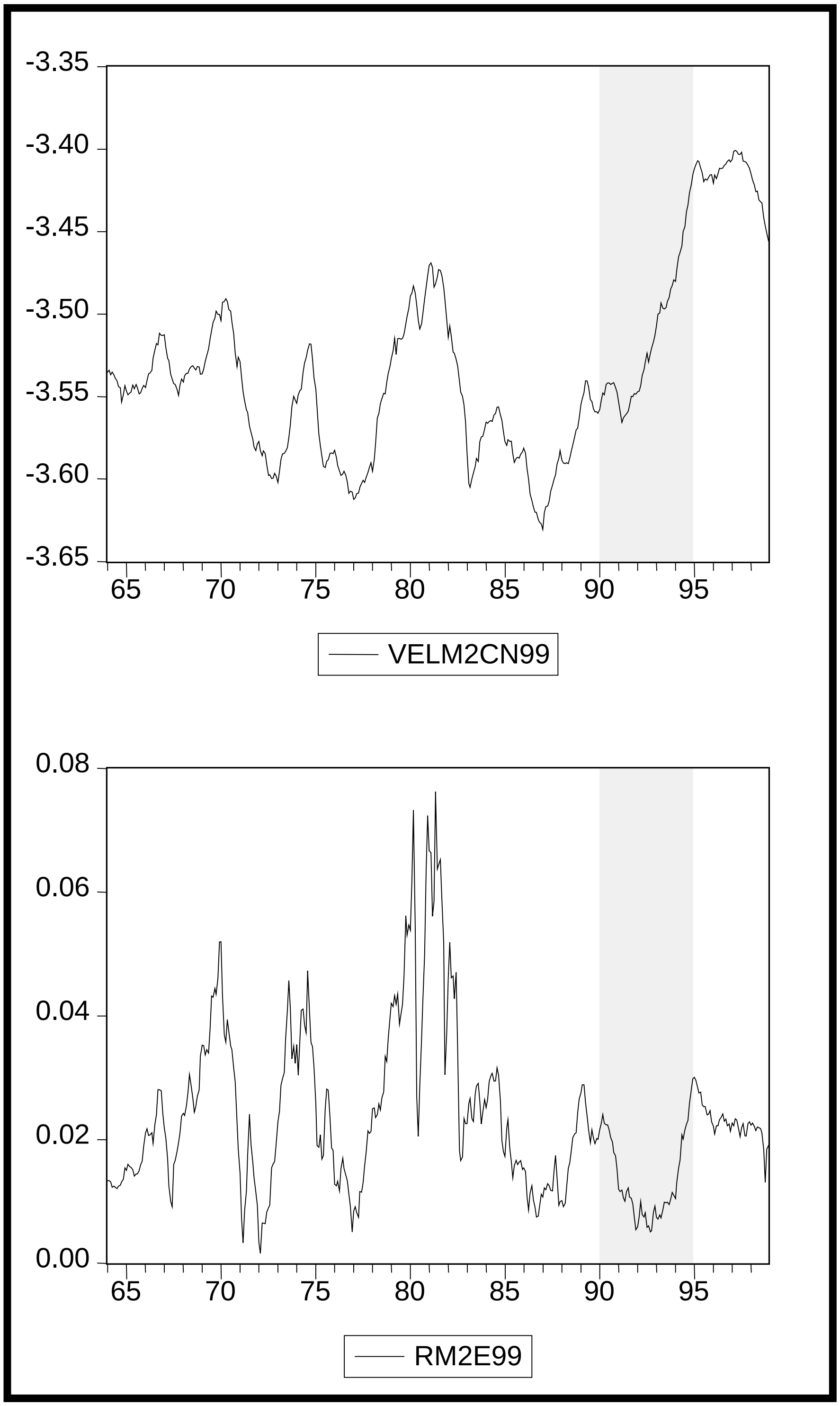
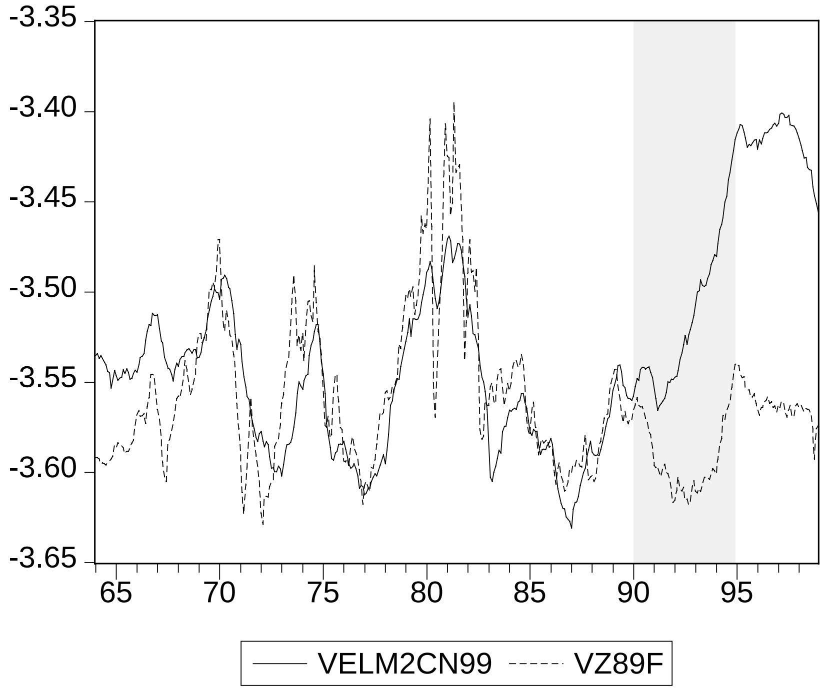
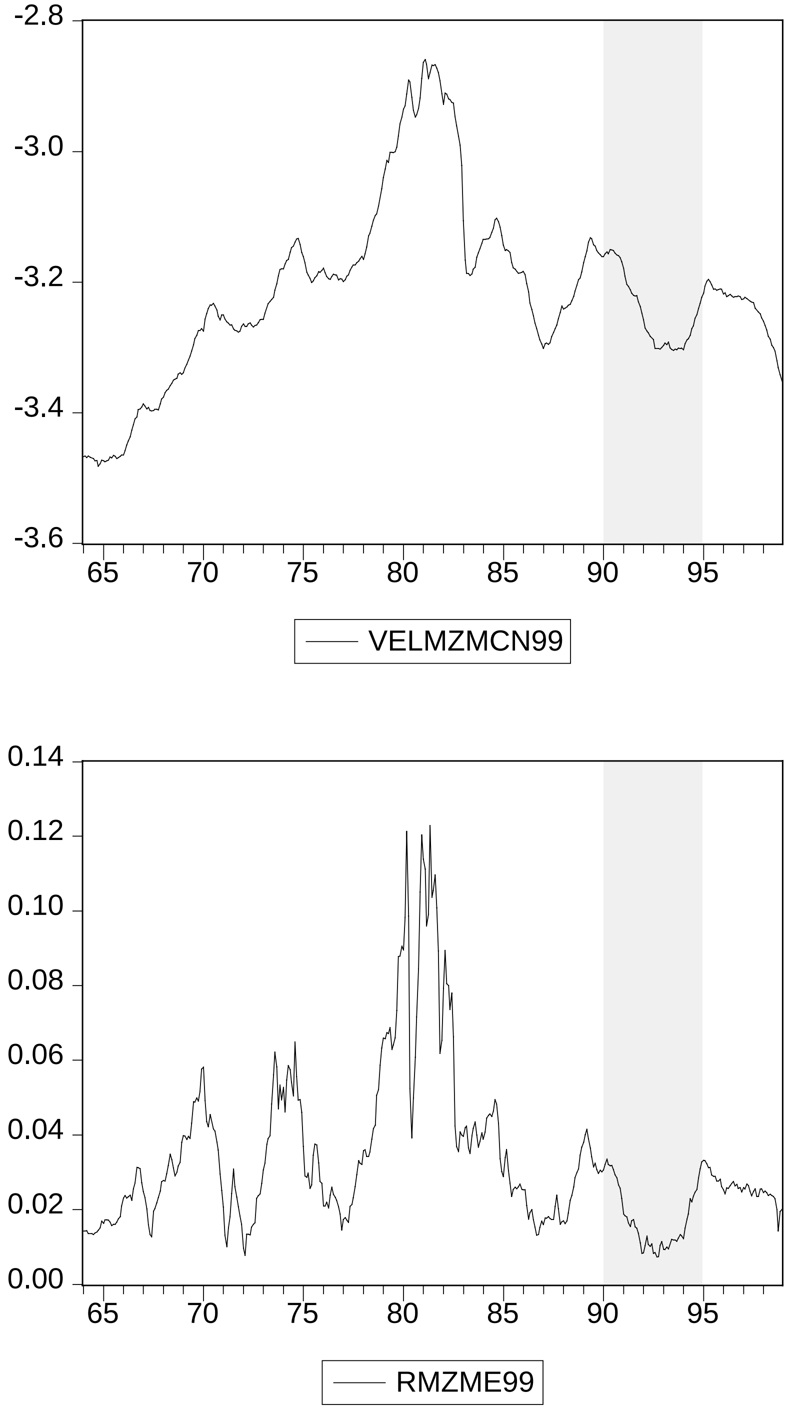


Figure 2. Log of actual and predicted M2 “coincident indicator” velocity, 1964:1-1998:12.



Note: The dashed line depicts equilibrium velocity as predicted from a bivariate cointegrating regression of velocity and opportunity cost (six lag spefication), estimated through 1989.

Figure 3. Log of MZM velocity using the coincident index for scale and the opportunity cost of M2 balances (1964:1-1998:12).



Figures 4. Cointegrating equations for the M2 and MZM models coincident index, 1964:1 1998:12.

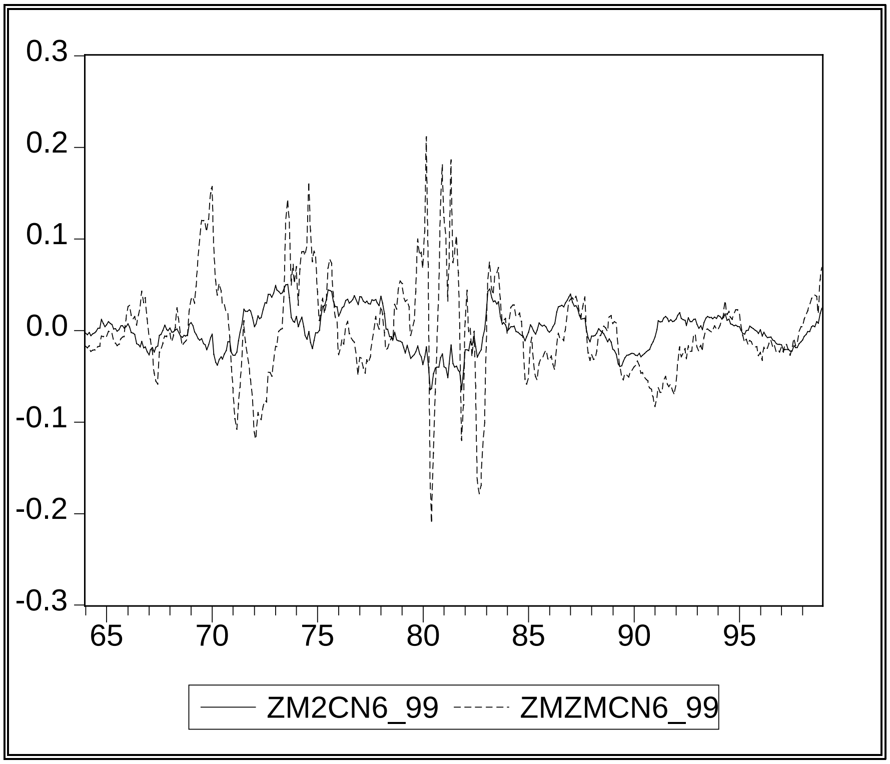


Figure 5.Cointegrating equations for the M2 and MZM models industrial production index, 1964:1-1998:12.

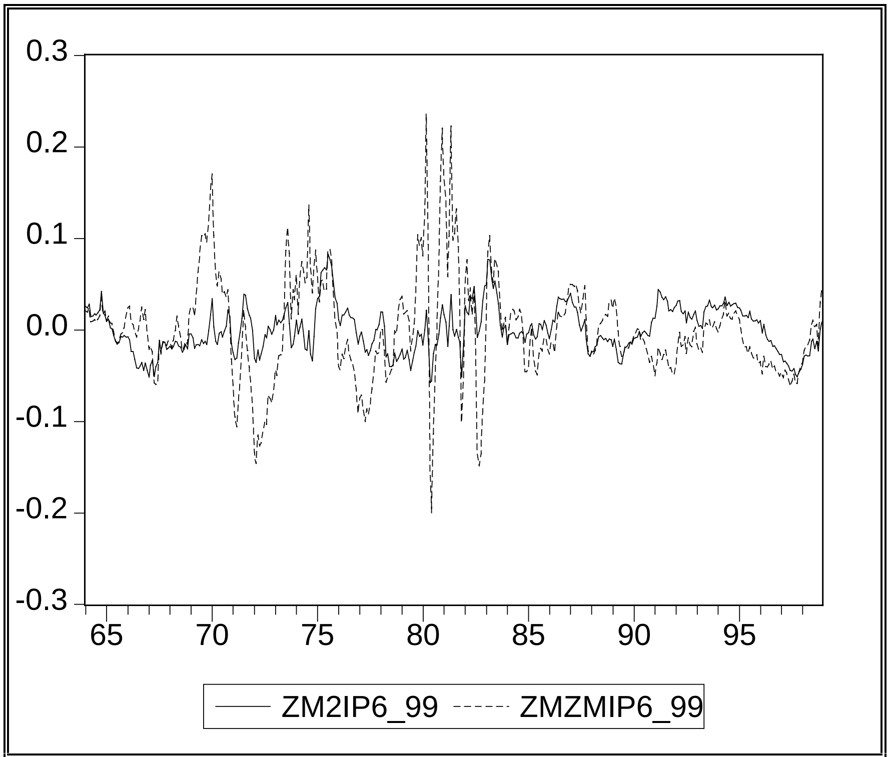


Figure 6. Cointegrating equations for the M2 and MZM models, personal income, 1964:1-1998:12.

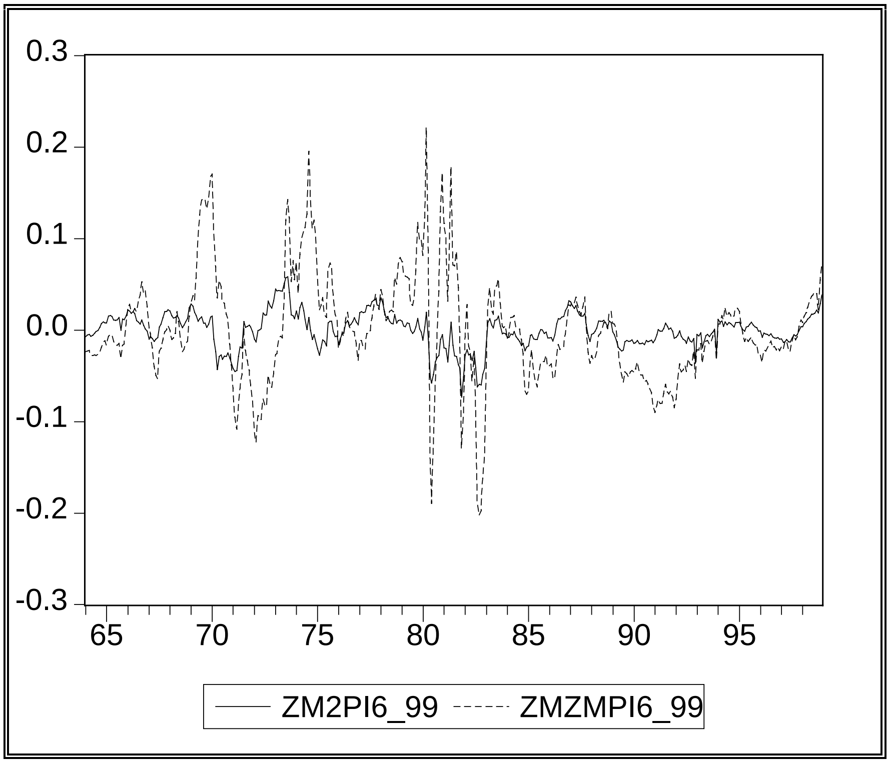


Figure 7 M2 cointegrating equation errors for all scale measures, velocity constrained equilibria (nomy-nomy*), 1964:1-1998:12.

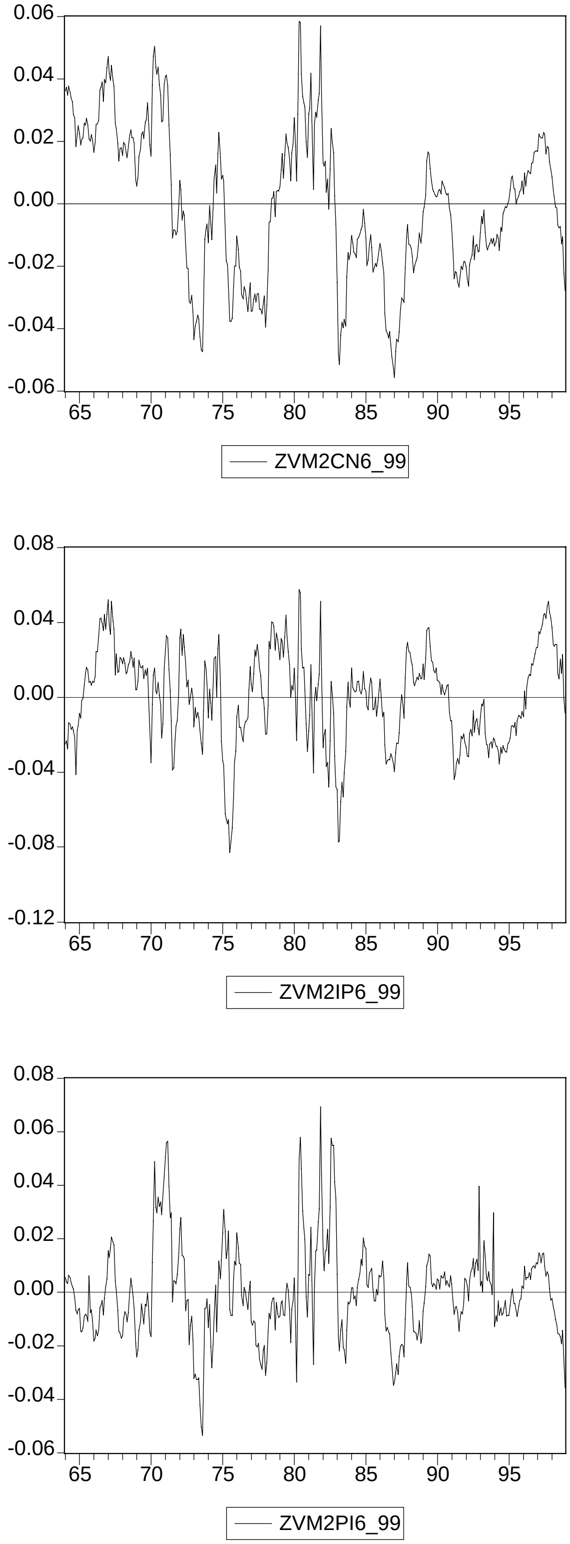


Figure 8. MZM cointegrating equation errors for all scale measures, velocity constrained equilibria (nomy-nomy*), 1964:1-1998:12

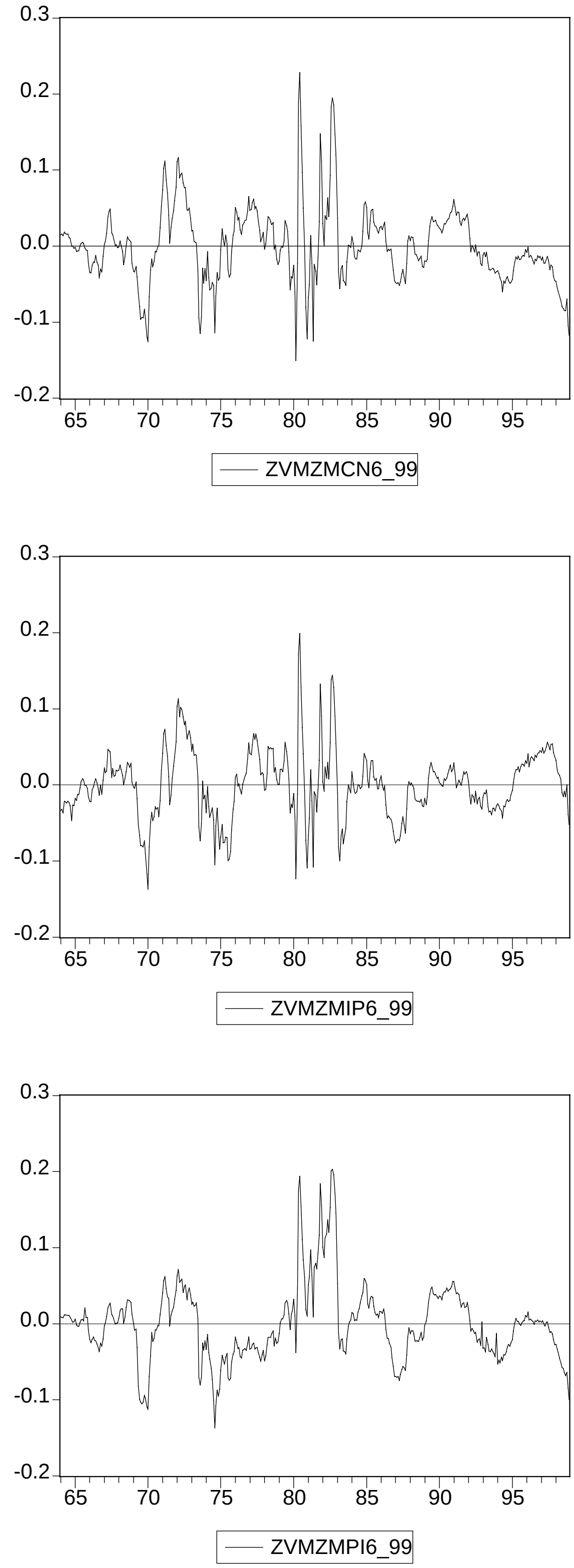


Figure 9. Impulse response functions

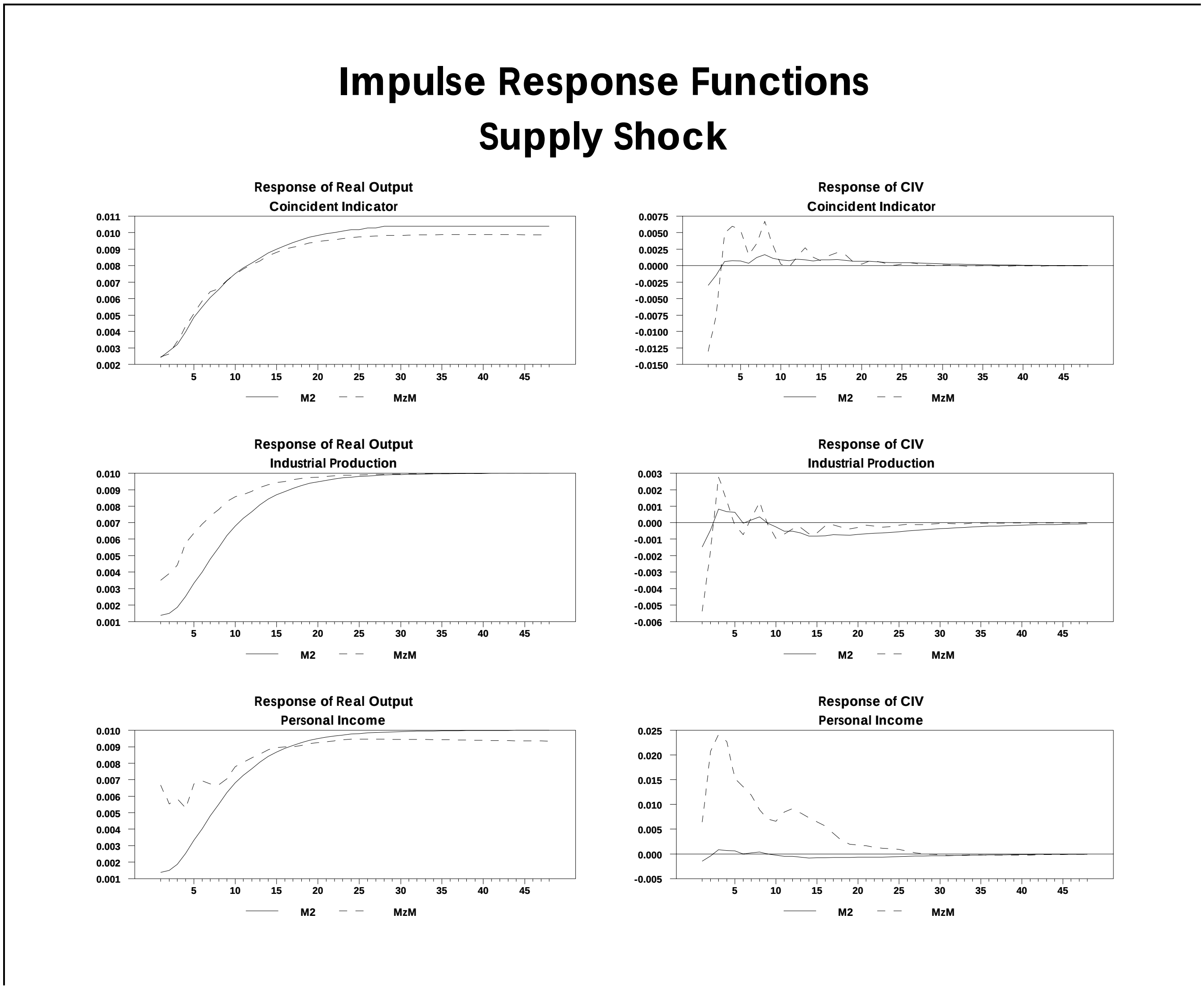


Figure 10. Impulse response functions.

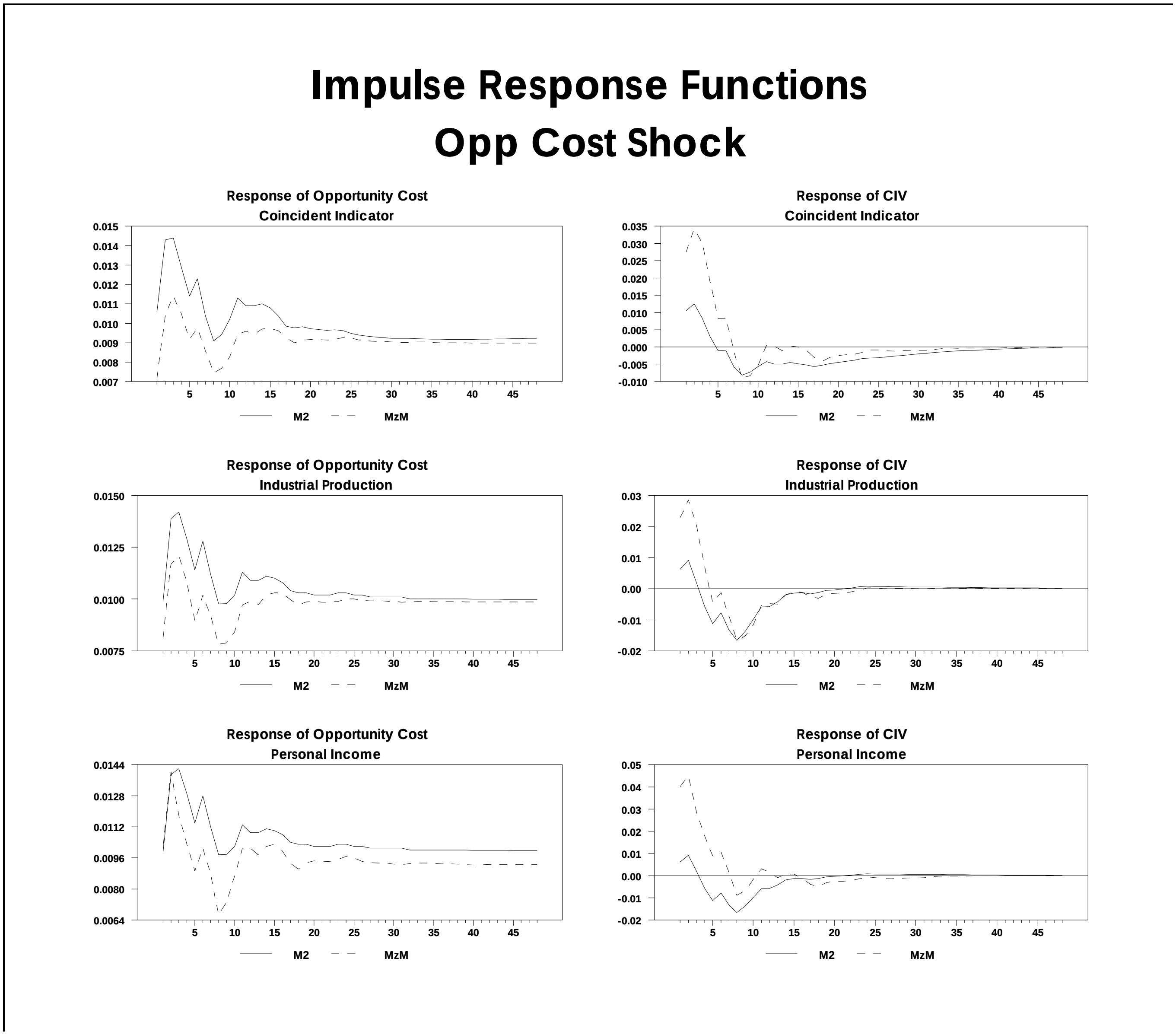


Figure 11. Impulse response functions.