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A NEW EMPIRICAL WEIGHTED MONETARY AGGREGATE FOR THE UK

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ABSTRACT

This paper utilizes an approach to long-run modelling proposed by Pesaran *et al.* (1996. Testing for the existence of a long run relationship. Mimeo, University of Cambridge) to develop an empirical weighted broad monetary aggregate for the UK. The properties of this new aggregate are contrasted with those of the corresponding simple sum and Divisia aggregates. The new weighted monetary aggregate is found to be highly stable and conforms well with standard money demand properties. The aggregate also displays sensible impulse response and persistence profiles to monetary shocks in the context of a VECM framework. Finally, the empirical weighted aggregate displays superior information content in respect of nominal income when contrasted with simple sum and Divisia aggregates using a series of St. Louis equations. Copyright © 2001 John Wiley & Sons, Ltd.

KEY WORDS: Divisia; leading indicators; monetary aggregation; monetary policy

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1. INTRODUCTION

By using an approach to long-run modelling proposed by Pesaran *et al.* (1996), this paper develops an empirically determined weighted monetary aggregate for the UK and contrasts its empirical performance with both a conventional simple sum aggregate and an aggregate based upon the increasingly popular Divisia index number methodology.

The formal targeting of monetary aggregates was introduced in many countries, including the UK and USA, during the early to mid-1970s. Monetary targets became particularly important at this time as the discipline of the Bretton Woods regime had been removed and some guiding principle for monetary policy was required. In the UK, monetary targeting using £M3 was first introduced in 1976 and £M3 was chosen as the best indicator of monetary conditions based upon the strong correlation between £M3 and nominal GDP (with a lag) established from the early 1970s. £M3 was reaffirmed as the targeted monetary aggregate in 1979–1980 and became the centrepiece of the Thatcher government's counter inflation policy under the Medium Term Financial Strategy (MTFS).

Monetary targeting failed in the UK and other countries because the chosen target aggregates did not remain stably related to key economic variables, such as nominal income and its decomposition in terms of prices (inflation) and real income (output). It is now well established, however, that the substantial financial innovations of the 1980s introduced instability into estimated demand functions for broad money, and it was largely for this reason that monetary targeting was abandoned. Indeed, the consensus at the end of the 1980s was that it was not possible to re-establish the apparent previous stability of conventional broad money demand functions, even if these were extended to include building society deposits in addition to notes and coins and bank deposits (thus defining the aggregate M4: see Hall *et al.*, 1989).

Recent empirical work (Belongia and Chrystal, 1991; Drake and Chrystal, 1994, 1997), however, suggests that the instability of broad money demand evident during the 1980s may be attributable in large part to

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the use of conventional official simple sum aggregates, which assume that the component assets are perfect substitutes. This assumption is likely to be particularly inappropriate at times of significant financial innovation involving changing interest yields on the various components of broad money. Barnett (1980, 1982) and Barnett *et al.* (1992) demonstrate that the Divisia methodology is theoretically superior to simple sum aggregation as it weights component assets according to their varying degrees of 'moneyness', and can endogenize financial innovations involving changing relative yields on component assets. Belongia and Chrystal (1991) and Drake and Chrystal (1994, 1997) find that the use of the Divisia index number methodology in the construction of monetary aggregates restores the stability of broad money demand functions over sample periods from the mid-1970s into the 1990s.

Notwithstanding the abandonment of formal broad money targeting in the UK in 1986, monetary policy still requires some guiding principles. Given that inflation is a lagging indicator, monetary policy needs, at the very least, some leading indicator of 'overheating'. Following the unsuccessful phase of at first informal, and then formal exchange rate targeting between 1987 and 1992, the UK government, in common with many other countries, subsequently adopted a policy of directly targeting the inflation rate. This policy culminated in the Bank of England being given operational independence to conduct monetary policy so as to hit an inflation target of $2.5\% \pm 1\%$. In doing so, the Bank's Monetary Policy Committee analyses a wide range of economic data, including both simple sum and the Bank's own Divisia monetary aggregates, in order to provide a forward looking assessment of inflationary pressures and prospects. Hence, the potential leading indicator properties of monetary aggregates still constitute an extremely important aspect of the conduct of monetary policy.

While it is generally conceded that Divisia monetary aggregates are theoretically superior to simple sum aggregates and, as noted above, the Bank of England and some other central banks do now monitor Divisia aggregates alongside their conventional simple sum aggregates, some reservations have been expressed by both academics and central banks/policy-makers concerning the practical applicability of Divisia aggregation. The Bank of England (Fisher *et al.*, 1993), for example, point to the practical difficulties associated with the choice of an appropriate benchmark rate of interest in the Divisia methodology (see Section 3) and to the problem posed by the introduction of new assets. Theoretically, Divisia indices are not defined when any of the monetary asset quantities are equal to zero in some time period(s). The alternative Fisher Ideal index is, however, valid in these circumstances.[1] Spencer (1986, 1994) and Ford *et al.* (1992) highlight the adjustments to the basic Divisia methodology associated with aspects of financial innovation and portfolio disequilibrium, while Barnett *et al.* (1997) and Drake *et al.* (1998, 1999) examine the necessary adjustments to the Divisia methodology required by the incorporation of risky assets. It should be noted that while all these issues can be successfully addressed, both theoretically and potentially in practice, they inevitably present problems for policy-makers in the context of the implementation of monetary policy.

Based upon the theoretical problems associated with the traditional simple sum aggregation methodology and the potential practical difficulties associated with the use of the Divisia methodology, one purpose in this paper, therefore, is to develop an alternative, empirically determined, weighted monetary aggregate. Since our purpose is to contribute to the UK policy debate in the context of monetary policy and inflation targeting, we develop this empirically weighted aggregate based upon a money-nominal income relationship and using the Bank of England's own data set for aggregate M4 (their preferred broad money aggregate).[2]

The examination of the link between monetary assets and nominal income has a long tradition, dating back to the early studies of Timberlake and Fortson (1967) and Laumas (1968), through to the more detailed statistical approaches of Tinsley *et al.* (1980) and Mills (1983a,b). Presumably because of the problems associated with monetary targeting, this approach then fell out of favour. However, there has been a resurgence of interest following the recent paper by Feldstein and Stock (1996). They address the particular issue of constructing monetary aggregates and measuring money growth in the presence of evolving financial markets and new monetary assets, and tackle this by constructing empirical monetary aggregates with the objective of providing reliable leading indicators of nominal income. In doing so they adopt two alternative methodologies: a switching regression analysis and a time varying parameter approach based on the Kalman Filter. For the purpose of this paper, however, we utilize an alternative methodology based upon a new approach to testing for the existence of a linear long-run relationship when the orders of integration in, or the form of cointegration between, the underlying regressors are not known with certainty. This approach is associated with Pesaran *et al.* (1996), henceforth PSS, and we use this technique to examine the long-run relationship between the component assets of M4 and nominal income and hence to derive an appropriate empirically weighted monetary aggregate.

The remainder of this paper is structured as follows. Section 2 outlines the PSS technique and the derivation of the empirically weighted monetary aggregate. Section 3 then contrasts the money demand properties of this new aggregate with those of the corresponding simple sum and Divisia aggregates within a VECM framework incorporating the alternative monetary aggregates, an appropriate interest rate or rental price variable, a general price index, and real income. Impulse response functions and persistence profiles are employed to investigate the effects of monetary shocks on the other variables. Section 4 attempts to verify the superior nominal income leading indicator properties of this new aggregate in the context of various forms of standard St. Louis nominal spending equations. Finally, Section 5 summarizes and concludes.

2. CONSTRUCTING A WEIGHTED MONETARY AGGREGATE FOR THE UK

The data set used to construct an empirically determined weighted monetary aggregate contains seasonally unadjusted quarterly observations from 1977:3 to 1997:2 on the logarithms of GDP at factor cost, denoted y_t, and four monetary components; x_{1t}, notes and coin plus non-interest bearing sight deposits; x_{2t}, interest bearing sight deposits; x_{3t}, time deposits; and x_{4t}, building society deposits.

These four monetary components are plotted over time in Figure 1 and reveal the distinctive features of a rapid growth in interest bearing sight deposits, a slow growth in notes and coin, and stable and consistent growth in time and building society deposits.

The approach taken to construct the weighted aggregate is that proposed by PSS. We thus begin by considering the following vector autoregressive model of order p (VAR(p)) in the vector of variables $\mathbf{z}_t = (y_t, \mathbf{x}_t')'$, where $\mathbf{x}_t = (x_{1t}, \ldots, x_{4t})'$ is the vector of monetary components

$$\mathbf{z}_t = \mathbf{b} + \mathbf{c}t + \sum_{i=1}^{p} \Phi_i \mathbf{z}_{t-i} + \varepsilon_t, \quad t = 1, 2, \ldots \tag{1}$$

where $\mathbf{b}$ and $\mathbf{c}$ are vectors of intercepts and trend coefficients and Φ_i, $i = 1, 2, \ldots, p$, are matrices of coefficients (in practice the intercept vector $\mathbf{b}$ is replaced by $\mathbf{b}^*\mathbf{S}$, where $\mathbf{S}$ is a matrix of seasonal dummies and $\mathbf{b}^*$ is the corresponding matrix of coefficients). We assume that the roots of

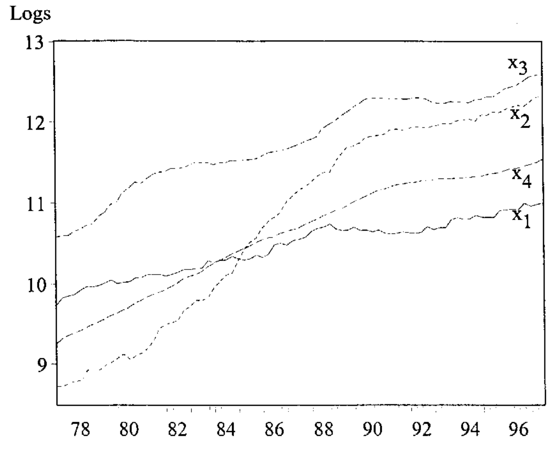

Figure 1. Monetary components.

$$\left| \mathbf{I}_5 - \sum_{i=1}^{p} \Phi_i z^i \right| = 0$$

are outside the unit circle $|z| = 1$ or satisfy $z = 1$, so that the elements of $\mathbf{z}_t$ are permitted to be either $I(0)$, $I(1)$ or cointegrated. The *unrestricted vector error correction form* of (1) is given by

$$\Delta\mathbf{z}_t = \mathbf{b} + \mathbf{c}t + \Pi\mathbf{z}_{t-1} + \sum_{i=1}^{p-1} \Gamma_i \Delta\mathbf{z}_{t-i} + \varepsilon_t, \quad t = 1, 2, \ldots \tag{2}$$

where

$$\Pi = -\left(\mathbf{I}_5 - \sum_{i=1}^{p} \Phi_i\right)$$

and

$$\Gamma_i = -\sum_{j=i+1}^{p} \Phi_j, \quad i = 1, \ldots, p-1$$

are matrices containing the long-run multipliers and the short-run dynamic coefficients, respectively.

Given the partition $\mathbf{z}_t = (y_t, \mathbf{x}_t')'$, we define the conformable partitions $\varepsilon_t = (\varepsilon_{1t}, \varepsilon_{2t}')'$ and

$$\mathbf{b} = \begin{bmatrix} b_1 \\ \mathbf{b}_2 \end{bmatrix}, \qquad \mathbf{c} = \begin{bmatrix} c_1 \\ \mathbf{c}_2 \end{bmatrix}, \qquad \Pi = \begin{bmatrix} \pi_{11} & \pi_{12} \\ \pi_{21} & \Pi_{22} \end{bmatrix}, \qquad \Gamma_i = \begin{bmatrix} \gamma_{11,i} & \gamma_{12,i} \\ \gamma_{21,i} & \Gamma_{22,i} \end{bmatrix}$$

and make the standard assumption that $\varepsilon_t = (\varepsilon_{1t}, \varepsilon_{2t}')'$ follows a multivariate i.i.d. process having mean zero, non-singular variance matrix

$$\Sigma_{\varepsilon\varepsilon} = \begin{bmatrix} \sigma_{11} & \sigma_{12} \\ \sigma_{21} & \Sigma_{22} \end{bmatrix}$$

and finite fourth moments. We also assume that $\pi_{21} = 0$, which ensures that there exists at most one (non-degenerate) long-run relationship between y_t and $\mathbf{x}_t$, irrespective of the level of integration of the $\mathbf{x}_t$ process.

With this assumption and the partitioning given above, (2) can be written in terms of the *dependent* variable y_t and the *forcing* variables $\mathbf{x}_t$ as

$$\Delta y_t = b_1 + c_1 t + \pi_{11} y_{t-1} + \pi_{12}\mathbf{x}_{t-1} + \sum_{i=1}^{p-1} \gamma_{11,i}\Delta y_{t-i} + \sum_{i=1}^{p-1} \gamma_{12,i}\Delta\mathbf{x}_{t-i} + \varepsilon_{1t} \tag{3}$$

$$\Delta\mathbf{x}_t = \mathbf{b}_2 + \mathbf{c}_2 t + \Pi_{22}\mathbf{x}_{t-1} + \sum_{i=1}^{p-1} \gamma_{21,i}\Delta y_{t-i} + \sum_{i=1}^{p-1} \Gamma_{22,i}\Delta\mathbf{x}_{t-i} + \varepsilon_{2t} \tag{4}$$

The contemporaneous correlation between ε_{1t} and ε_{2t} can be characterized by the regression

$$\varepsilon_{1t} = \omega'\varepsilon_{2t} + \xi_t \tag{5}$$

where $\omega = \Sigma_{22}^{-1}\sigma_{21}$, $\{\xi_t\}$ is an i.i.d. $(0, \sigma_\xi^2)$ process with $\sigma_\xi^2 = \sigma_{11} - \sigma_{12}\Sigma_{22}^{-1}\sigma_{21}$, and the $\{\xi_t\}$ and $\{\varepsilon_{2t}\}$ processes are uncorrelated by construction. Substituting (4) and (5) into (3) yields

$$\Delta y_t = a_0 + a_1 t + \phi y_{t-1} + \delta'\mathbf{x}_{t-1} + \sum_{i=1}^{p-1} \psi_i \Delta y_{t-i} + \sum_{i=0}^{p-1} \varphi_{12,i}\Delta\mathbf{x}_{t-i} + \xi_t \tag{6}$$

where $a_0 \equiv b_1 - \omega'\mathbf{b}_2$, $a_1 \equiv c_1 - \omega'\mathbf{c}_2$, $\phi \equiv \pi_{11}$, $\delta' \equiv \pi_{12}' - \Pi_{22}'\omega$, $\psi_i \equiv \gamma_{11,i} - \omega'\gamma_{21,i}$, $\varphi_0 \equiv \omega'$, $\varphi_i \equiv \gamma_{12,i} - \omega'\Gamma_{22,i}$. It follows from (6) that, if $\phi \neq 0$ and $\delta \neq 0$, there exists a long-run relationship between the levels of y_t and $\mathbf{x}_t$, given by

$$y_t = \theta_0 + \theta_1 t + \theta'\mathbf{x}_t + \upsilon_t \tag{7}$$

where $\theta_0 \equiv -a_0/\phi$, $\theta_1 \equiv -a_1/\phi$, $\theta \equiv -\delta/\phi$ is the vector of long-run response parameters and $\{v_t\}$ is a mean zero stationary process. If $\phi < 0$ then this long-run relationship is *stable* and (6) can be written in the error correction model (ECM) form

$$\Delta y_t = a_0 + a_1 t + \phi(y_{t-1} - \theta' \mathbf{x}_{t-1}) + \sum_{i=1}^{p-1} \psi_i \Delta y_{t-i} + \sum_{i=0}^{p-1} \varphi_{12,i} \Delta \mathbf{x}_{t-i} + \xi_t \tag{8}$$

If $\phi = 0$ in (8) then no long-run relationship exists between y_t and $\mathbf{x}_t$. However, a test for $\phi = 0$ runs into the difficulty that the long-run parameter vector θ is no longer identified under this null, being present only under the alternative hypothesis. Consequently, PSS test for the absence of a long-run relationship, and avoid the lack of identifiability of θ, by examining the joint null hypothesis $\phi = 0$ *and* $\delta = 0$ in the unrestricted ECM (6). Note that it is then possible for the long-run relationship to be *degenerate*, in that $\phi \neq 0$ but $\delta = 0$, in which case the long-run relationship involves only y_t and possibly a linear trend.

PSS consider the conventional Wald statistic of the null $\phi = 0$, $\delta = 0$ and show that its asymptotic distribution involves the non-standard unit root distribution and depends on both the dimension (here $k = 4$) and cointegration rank ($0 \leq r \leq k$) of the forcing variables $\mathbf{x}_t$. This cointegration rank is the rank of the matrix Π_{22} appearing in (4). PSS obtain this asymptotic distribution in two polar cases: (i) when Π_{22} is of full rank ($r = 4$ here), in which case $\mathbf{x}_t$ is an $I(0)$ vector process, and (ii) when the $\mathbf{x}_t$ process is not mutually cointegrated ($r = 0$ and $\Pi_{22} = 0$) and hence is an $I(1)$ process. They point out that the critical values obtained from stochastically simulating these two distributions must provide lower and upper critical value bounds for all possible classifications of the forcing variables into $I(0)$, $I(1)$ and cointegrated processes. A *bounds procedure* to test for the existence of a long-run relationship within the unrestricted ECM (6) is thus as follows. If the Wald (or related F-) statistic falls below the lower critical value bound, then the null $\phi = 0$, $\delta = 0$ is not rejected, irrespective of the order of integration or cointegration rank of the variables. Similarly, if the statistics are greater than their upper critical value bounds, the null is rejected and we conclude that there is a long-run relationship between y_t and $\mathbf{x}_t$. If the statistics fall within the bounds, inference is inconclusive and detailed information about the integration-cointegration properties of the variables is then necessary in order to proceed further. It is the fact that we may be able to make firm inferences without this information, and thus avoid the severe pre-testing problems usually involved in this type of analysis, that makes this procedure attractive in applied situations. PSS provide critical values for alternative values of k under two situations: Case 1: $a_0 \neq 0$, $a_1 = 0$ (with an intercept but no trend in (6)), and Case 2: $a_0 \neq 0$, $a_1 \neq 0$ (with both an intercept and a trend in (6)). With $k = 4$ as here, the 1% significance level bounds for the Wald statistic are, for Case 1, 19.10 and 25.61, and for Case 2, 23.09 and 28.93, while the 5% bounds are, for Case 1, 14.25 and 20.25, and for Case 2, 17.69 and 23.34.

PSS show that this testing procedure is consistent and that the approach is applicable in quite general situations. For example, Equation (6) can be regarded as an autoregressive distributed lag model in y_t and $\mathbf{x}_t$ having all lag orders equal to p. Differential lag lengths can be used without affecting the asymptotic distribution of the test statistic.

In implementing this approach, our first task is to check that the assumptions required for attention to focus solely on Equation (6) are satisfied. One underlying assumption, implicit in the discussion above, is that the maximal order of integration of the $\{\mathbf{z}_t\}$ process is unity. Unit root tests of the individual series making up $\{\Delta \mathbf{z}_t\}$ show that a unit root is rejected at the 5% level in each case. A second assumption, explicitly discussed above, is that $\pi_{21} = 0$ in (the partitioned form of) the unrestricted vector error correction (2). Estimation of this equation with p set equal to 5 (and seasonal dummies rather than a constant) produced t-statistics on the coefficients of y_{t-1} in the equations for Δx_{it}, $i = 1, \ldots, 4$, of 1.31, -0.27, -1.06 and 1.08, thus producing no evidence against the null hypothesis $\pi_{21} = 0$. A setting of $p = 5$ was thought to be an appropriate trade-off between the need to account for stochastic seasonality and the degrees of freedom available given the length and dimension of $\mathbf{z}_t$.

Having ascertained that the conditions required for (6) to be considered in isolation are satisfied, the following parsimonious specification of this equation was eventually arrived at

$$\Delta y_t = -\underset{(0.082)}{0.488}y_{t-1} + \underset{(0.041)}{0.242}(x_{1,t-1} + x_{4,t-1}) - \underset{(0.012)}{0.053}(x_{2,t-1} - x_{3,t-1}) + \underset{(0.073)}{0.163}\Delta y_{t-4} - \underset{(0.045)}{0.161}\Delta x_{1,t-1}$$
$$- \underset{(0.055)}{0.148}\Delta x_{1,t-2} - \underset{(0.041)}{0.115}\Delta x_{2,t} + \underset{(0.035)}{0.071}\Delta x_{2,t-3} - \underset{(0.168)}{0.371}\Delta x_{4,t-2} - \underset{(0.185)}{0.635}\Delta x_{4,t-4} + \text{seasonal dummies}$$

$R^2 = 0.935 \quad \hat{\sigma}_\xi = 0.0109$

AUTO(4) = 5.43[0.25] NORM = 8.71[0.01] ARCH(1) = 0.65[0.42]

HET = 2.08[0.02] RESET(1) = 0.11[0.74]

Sample: 1978:4–1997:2

Because there is evidence of heteroskedasticity, heteroskedasticity-consistent standard errors (S.E.s) of coefficients are shown in parentheses. All individual coefficients are highly significant, while the restriction linking $x_{1,t-1}$ and $x_{4,t-1}$ is easily satisfied. Most standard diagnostic checks (probability values are shown in brackets) indicate no evidence of misspecification. The significant normality statistic is a consequence of a large outlier at 1979:2, the third observation of the effective sample-if the sample is shortened to omit this observation virtually no difference is made to the estimates. Because the stability of this regression is important to what follows, we show cusum squared and recursive residual plots in Figure 2, which reveal no real evidence of parameter instability (the slight evidence of instability shown by the cusum of squares plot is probably a consequence of heteroskedasticity), and neither did a variety of other tests.

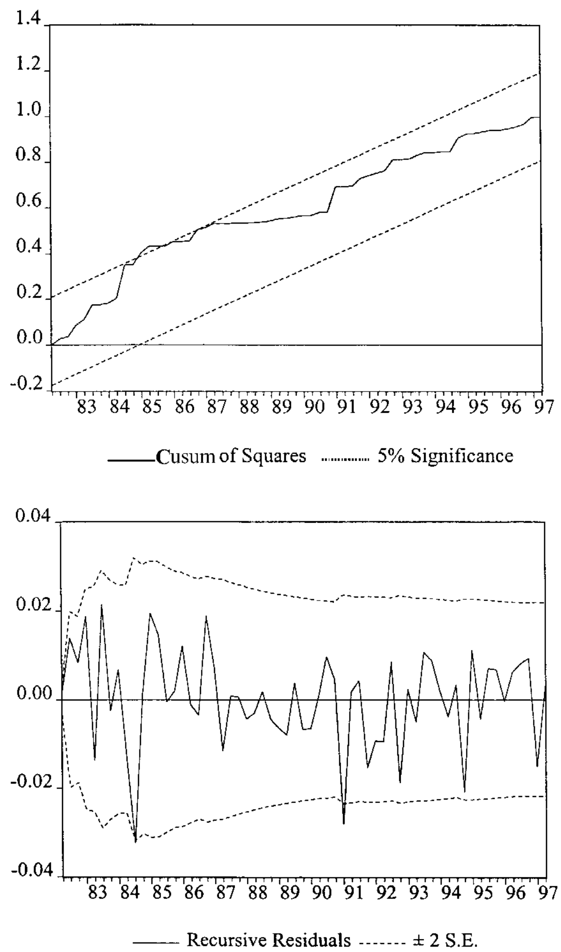

Figure 2. Cusum of squares and recursive residual plots.

The Wald statistic for testing whether there exists a long-run relationship between y_t and $\mathbf{x}_t$ produces a value of 36.00. This is well beyond the 1% significance level upper bound in both Cases 1 and 2 (note that the trend was found to be insignificant and hence has been omitted from the chosen specification). We must therefore conclude that such a long-run relationship certainly exists.

The long-run relationship implied by the model can now be used to construct an empirically determined weighted monetary aggregate. Given our estimates, the long-run relationship (7) is

$$y_t = \hat{\theta}\mathbf{x}_t = 0.5x_{1,t} - 0.1x_{2,t} + 0.1x_{3,t} + 0.5x_{4,t}$$

so that a weighted (logarithmic) monetary aggregate may be defined as $W_t = \hat{\theta}\mathbf{x}_t$. In levels, the aggregate is $(X_{1,t}X_{4,t})^{0.5}(X_{3,t}/X_{2,t})^{0.1}$, where $X_{i,t} = \exp(x_{i,t})$, $i = 1, \ldots, 4$. The implication of this is that the growth of the aggregate is driven primarily by the growth of non-interest bearing deposits and building society deposits, with the *relative* growth rate of time deposits to interest bearing deposits having only a minor impact. A possible explanation for this latter effect is that the newer time deposits are strong substitutes for the more traditional bank sight deposits. However, the presence of a negative weight on x_2 may be regarded with some concern. One possibility is that it is a consequence of omitting an interest rate variable from the model. However, if the current logarithm of the treasury bill rate and five lags are added to the parsimonious specification, none have absolute t-ratios greater than 0.62 and a Wald test for their inclusion has a probability value of just 0.98. Moreover, the implied θ weights are almost unchanged.

Of course, an important aspect of such an empirically determined aggregate is whether the weights are stable over time. To investigate this, the model was estimated recursively and used to construct a set of 'recursive weights'. These are plotted in Figure 3 from 1985, and the stability of the weights is remarkable, there being virtually no change in the weights for the last decade of the sample.

The θ weights are very different from those that would be implied by using alternative weighting methodologies, e.g. simple sum and Divisia aggregation. In particular, the Divisia methodology would accord relatively high interest bearing assets such as building society deposits, x_4, the lowest weight in the aggregation process.

3. MONEY DEMAND ANALYSIS

Although our aggregation methodology is empirical and essentially atheoretical, it is nevertheless important to examine the theoretical properties of the resultant aggregate. Thus, having derived our empirically determined weighted monetary aggregate, W_t, the purpose of this section is to analyse the

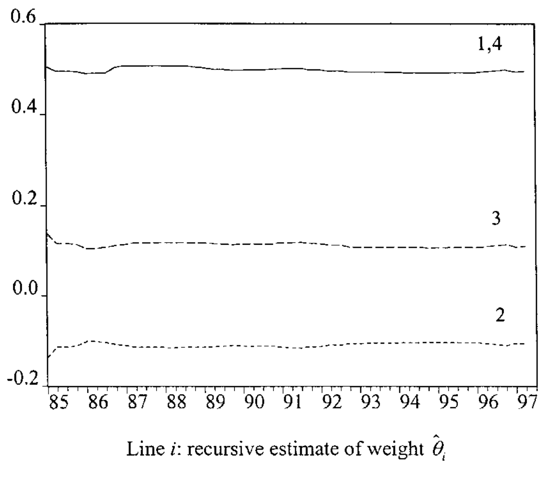

Figure 3. Recursive estimates of weights.

money demand properties of this new aggregate and to contrast these with those associated with the traditional simple sum aggregate and with a Divisia weighted monetary aggregate. The former simple sum methodology, which is still widely used by central banks around the world, accords an equal weight of unity to each component asset and therefore implicitly assumes perfect substitutability across component assets. In contrast, the Divisia methodology, pioneered in this context by Barnett (1980, 1982) and utilized initially in the UK by Mills (1983c), attempts to weight the asset components according to their degree of 'moneyness'. In order to do this, Barnett views monetary assets as providing a joint product of monetary services (transactions, liquidity, etc.) and investment services. With respect to the latter, this will entail the payment of an own rate of interest on the monetary asset in question.

Hence, in order to 'strip out' this investment services element and to construct a weighted monetary aggregate which measures only monetary services, the Divisia methodology makes use of the rental price or user cost concept for monetary assets developed by Barnett (1978)

$$\pi_{it} = (R_t - r_{it})/(1 + R_t) \tag{9}$$

Here π_{it} is the rental price of the ith asset in period t, r_{it} is the own rate of return on the asset in time t, and R_t is the benchmark rate of return. The benchmark rate of return is typically taken as the return on some asset which provides no monetary services and is held purely as a store of value.

Using these rental prices, the Divisia monetary services quantity index is a chain weighted index derived as the weighted average growth rate of the component assets. The weights are the individual asset shares in total monetary expenditure, where the latter is defined as the sum of the nominal asset quantities multiplied by their individual rental prices. The index is normalized in some base period to provide a monetary quantity aggregate in levels rather than in growth rate terms and a dual rental price index can be derived from the monetary quantity index and from the data on total monetary expenditure.

To construct the Divisia aggregate, we utilize the data on the four asset components employed previously to construct the empirical aggregate: notes and coins plus non-interest bearing sight deposits; interest bearing sight deposits; bank time deposits; and building society deposits. Data on the own rates of return for the interest bearing assets were provided by the Bank of England.[3] In order to maximize the policy relevance of our analysis, we also elect to utilize the benchmark rate of return series constructed by the Bank of England and used in the context of Equation (9) in the construction of their own Divisia aggregates (see Fisher *et al.*, 1993).

Hence, the logic of the Divisia methodology is that assets which are held primarily for their investment services (and which have a high own rate of return) will have a low rental price and will therefore be accorded a low weight in the Divisia monetary services index. Conversely, notes and coins, for example, provide no investment services and will have the highest user cost and hence the greatest weight in the aggregate. Divisia monetary aggregation has been adopted by a number of central banks in recent years. The Bank of England, for example, now publishes regular Divisia aggregates alongside the traditional simple sum aggregates (see Fisher *et al.*, 1993). We denote the logarithm of the Divisia aggregate as D_t.

In contrast to the Divisia methodology, which produces a theoretically determined weighted monetary aggregate, we have derived a weighted monetary aggregate empirically. There are strong parallels between the two methodologies, however, as both are attempting to isolate monetary services such as transactions services. The Divisia methodology attempts to do this theoretically using the rental price concept, whereas we have attempted to do this empirically by isolating a weighted monetary aggregate derived from a long-run money–nominal income relationship. With respect to the money demand literature, this approach can be likened to attempting to identify a monetary transactions aggregate, and can also be justified from a monetary policy perspective on the grounds that policy-makers are ultimately concerned with the relationship between monetary aggregates and nominal income.

Figure 4 shows the annual growth rates of D_t and W_t, along with that of the 'simple-sum' aggregate, defined as

$$S_t = \log\left(\sum_{i=1}^{4} \exp(x_{it})\right)$$

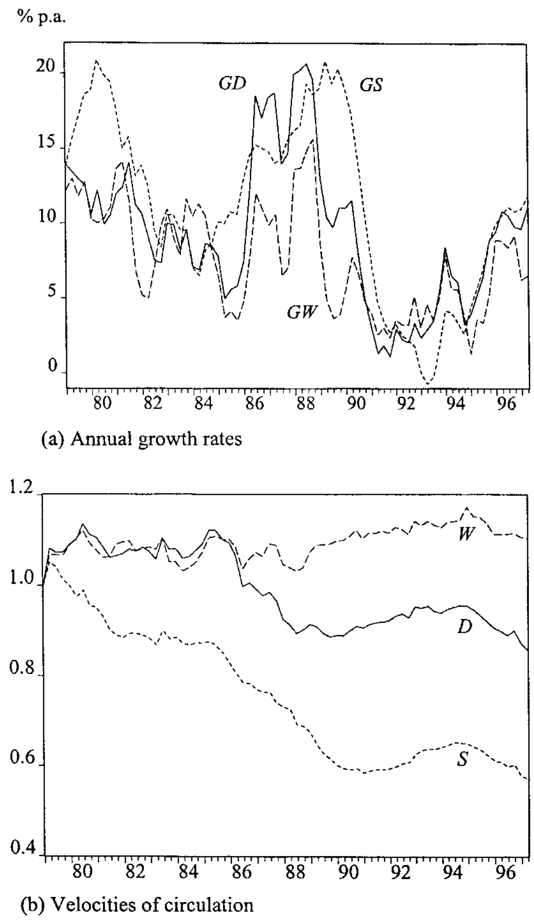

Figure 4. Weighted and simple sum money annual growth rates and velocities of circulation.

These growth rates are defined as $GD_t = 100(D_t - D_{t-4})$, $GW_t = 100(W_t - W_{t-4})$ and $GS_t = 100(S_t - S_{t-4})$. Also shown are the velocities of circulation of the three aggregates.

Noticeable and important differences between the aggregates are observed in these plots. The weighted monetary aggregate W_t typically has a lower rate of growth than the other two aggregates, most noticeably during the 1980–1982 and 1986–1989 periods, when monetary growth as conventionally defined was very high. Consequently, W_t's velocity of circulation is very stable, unlike that of the simple sum aggregate S_t and, somewhat less so, the Divisia aggregate D_t.

In terms of leading indicator properties, a casual inspection of the growth rates shows that both *GD* and *GW* declined significantly from the late 1970s to the mid-1980s, as did inflationary pressures in the UK economy. In contrast, *GS* increased significantly in the early 1980s and provided misleading signals in the context of the MTFS monetary target ranges. Similarly, in the context of the 'boom and bust' scenario of the late 1980s and early 1990s, *GW* and *GD* increased sharply from early 1986, but declined rapidly from late 1988, thus providing good leading indicators of inflationary pressure. *GS*, however, exhibited a strong increase from 1984 right up to the onset of recession in 1990.

A formal econometric comparison of the three monetary aggregates is, of course, required. To ensure that our comparison is as consistent as possible, we utilize the rental price concept in respect of the opportunity cost variable for both W_t and S_t as well as in the construction of D_t. We believe this is preferable to using an arbitrary interest rate or interest differential in the money demand specification. With W_t as defined in the previous section, the appropriate rental price variable R_{W_t} is derived as total monetary expenditure divided by $\exp(W_t)$. As the simple sum methodology assumes that the component assets are all perfect substitutes, the appropriate rental price dual to S_t is the Leontief price index, which is the minimum user cost derived from (9) using the maximum own rate of return. This is denoted as R_{S_t}. Finally, the dual rental price index for the Divisia aggregate D_t is denoted as R_{D_t}.

As our aim is to contrast the three monetary aggregates in the context of a standard money demand framework, the remaining variables used in the subsequent analysis are the logarithms of real GDP, ry_t; the retail price index, p_{RPI_t}, and the GDP deflator, p_{y_t}, with the latter two variables being alternative measures of the aggregate price level.

As all the above variables are found to be integrated of order one, the modelling framework that is appropriate is the vector error correction form of Equation (2), where the vector $\mathbf{z}_t$ is now defined as $\mathbf{z}_t = (M_t, R_t, ry_t, p_t)'$, where M_t is one of the three alternative monetary aggregates, R_t is the accompanying rental price, and p_t is either of the aggregrate price indices (we actually focus on the results obtained using p_{RPI_t}, as this index produced better results than when p_{y_t} was used). The order of the underlying VAR was selected using a combination of information criteria, likelihood ratio tests, and diagnostic checking: orders of either 5 or 6 were found to be appropriate for the various vectors.

Unlike the analysis of the previous section, however, it was found that none of the $\mathbf{z}_t$ vectors could be partitioned in such a way that only one long-run relationship existed. We thus embarked on a standard cointegration analysis: i.e., we investigated whether the long-run multiplier matrix could be partitioned as $\Pi = \alpha\beta'$, where α and β are $4 \times r$ matrices of full column rank r, $0 \leq r \leq 4$, known as the factor loading (or adjustment) matrix and cointegrating matrix, respectively, the columns of β being the cointegrating vectors. This then enables the cointegrating relations, or error corrections, $\beta'\mathbf{z}_t$ to be defined. The value of r was selected using the approach of Johansen (1995). As is well known, the manner in which the constant and trend are modelled is important in this approach, and as the trend coefficient vector $\mathbf{c}$ was found to be significantly different from zero for all $\mathbf{z}_t$s, t was restricted so that it appears only within the cointegration relation, i.e. we set $\mathbf{c} = \alpha\beta'\gamma$ so that (2) can be written

$$\Delta\mathbf{z}_t = \mathbf{b} + \alpha\beta'(\mathbf{z}_{t-1} - \gamma t) + \sum_{i=1}^{p-1} \Gamma_i \Delta\mathbf{z}_{t-i} + \varepsilon_t, \quad t = 1, 2, \ldots$$

or as

$$\Delta\mathbf{z}_t = \mathbf{b} + \alpha\mathbf{e}_{t-1} + \sum_{i=1}^{p-1} \Gamma_i \Delta\mathbf{z}_{t-i} + \varepsilon_t, \quad t = 1, 2, \ldots \tag{10}$$

Table I. Maximum likelihood tests of cointegrating rank

Null	Alternative	LR statistic	95% critical value
(a) Money aggregate D			
$r = 0$	$r \geq 1$	86.07	63.00
$r \leq 1$	$r \geq 2$	54.13	42.34
$r \leq 2$	$r \geq 3$	28.08	25.77
$r \leq 3$	$r = 4$	8.53	12.39
(b) Money aggregate W			
$r = 0$	$r \geq 1$	95.42	63.00
$r \leq 1$	$r \geq 2$	47.35	42.34
$r \leq 2$	$r \geq 3$	24.63	25.77
$r \leq 3$	$r = 4$	7.23	12.39
(c) Money aggregate S			
$r = 0$	$r \geq 1$	93.84	63.00
$r \leq 1$	$r \geq 2$	58.62	42.34
$r \leq 2$	$r \geq 3$	27.33	25.77
$r \leq 3$	$r = 4$	10.15	12.39

where $\mathbf{e}_{t-1} = \beta'(\mathbf{z}_{t-1} - \gamma t)$ is the vector of r error corrections. Table I presents likelihood ratio (LR) test statistics for determining r for the three alternative $\mathbf{z}_t$ vectors under consideration, which show that either $r = 2$ or 3 cointegrating vectors are found in each system (the test statistic for $r = 3$ in the W system is almost significant at the 95% level, but this value was not used for reasons outlined below: the maximal eigenvalue test also found r to be either 2 or 3).

To be able to interpret these cointegrating vectors, they need to be uniquely identified. With r cointegrating vectors, r independent restrictions are required to just-identify each vector: any further restrictions are then over-identifying and hence the validity of their imposition is able to be tested. Obviously, economic theory considerations should play a major role in determining these restrictions, so that interest lies in whether a long-run demand for money function and an aggregate supply function linking the price level and output can be uniquely identified. Table II presents the unique cointegrating

Table II. Uniquely identified cointegrating vectors

	$\hat{\beta}_1$	$\hat{\beta}_2$	$\hat{\beta}_3$
(a) Money aggregate D			
M	1	0	0
R	−0.244 (0.086)	0	1
ry	−3	1	2.152 (1.331)
p	−1	−0.521 (0.067)	1
t	0.00622 (0.00076)	0	−0.0224 (0.0082)
LR test of 1 over-identifying restriction: $\chi^2(1) = 0.001$ [0.973]			
(b) Money aggregate W			
M	1	0	
R	0.123 (0.082)	0	
ry	−1.586 (0.242)	1	
p	−1	−0.490 (0.066)	
t	0.00349 (0.00117)	0	
LR test of 2 over-identifying restrictions: $\chi^2(2) = 1.78$ [0.41]			
(c) Money aggregate S			
M	1	0	0
R	−0.0476 (0.0268)	0	0
ry	−3	1	1
p	−1	−0.513 (0.089)	1.483 (0.188)
t	0.00554 (0.00121)	0	0
LR test of 2 over-identifying restrictions: $\chi^2(2) = 1.20$ [0.55]			

S.E.s of unrestricted coefficients shown in ( ); probability value shown in [ ].

vectors thus identified. As the trend also appears in these vectors, its omission is a further over-identifying restriction. As is seen, the sets of over-identifying restrictions are acceptable in each system. For each system the first cointegrating vector can be interpreted as a long-run money demand function. For both the Divisia and simple sum aggregates (panels (a) and (c)), price homogeneity and an income elasticity of 3 are found, these being consistent with earlier results of Drake and Chrystal (1994) in respect of Divisia aggregates.[4] In both cases the rental price semi-elasticity is positive. For the empirically weighted aggregate (panel (b)), price homogeneity is again found but the income elasticity is halved and the rental price semi-elasticity is now negative. The second cointegrating vector is stable across all systems, being approximately $e_{2t} = y_t - 0.5p_t$. Since three cointegrating vectors can only just be rejected at the 95% level for the W system, we also investigated this extended set. Unfortunately, it was not possible to identify three independent vectors having statistically significant free parameters, so that our original choice of $r = 2$ was adhered to in further analysis.

Figure 5 shows the *persistence profiles* of the cointegrating vectors in each system (see Pesaran and Shin, 1996; Pesaran and Pesaran, 1997). These show the time profile of the response of each cointegrating vector to 'system-wide' shocks, i.e. shocks drawn from the *multivariate* distribution of ε_t, rather than from the distribution of a particular component of it. These persistence profiles are defined using the solution to (10), which is, from Pesaran and Shin (1997)

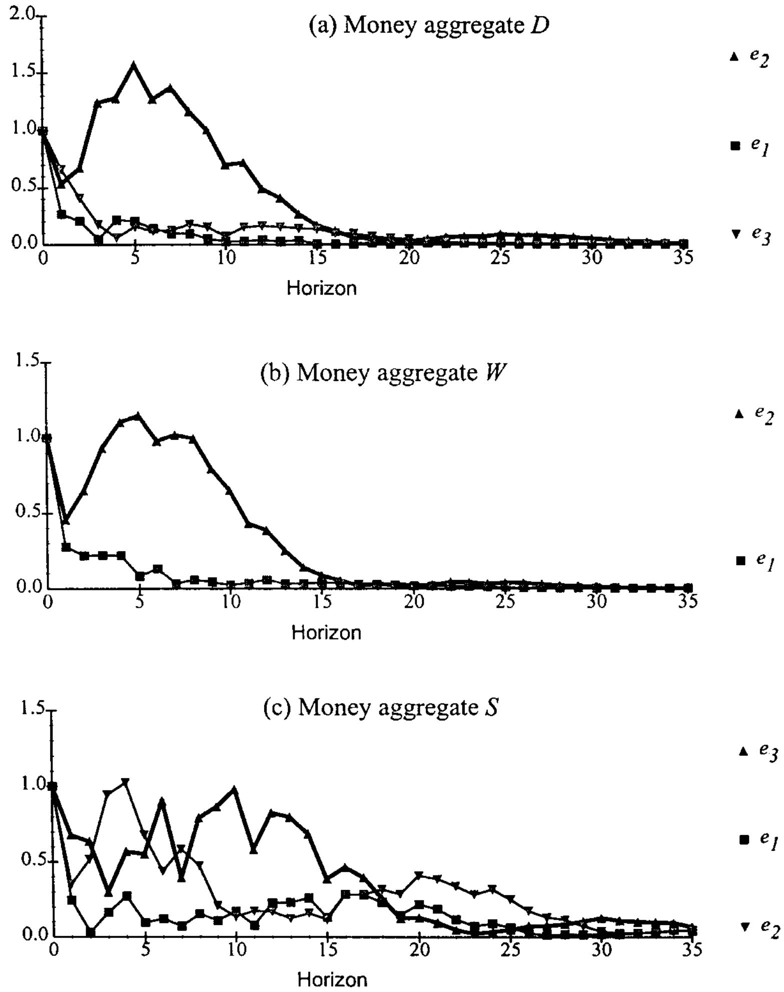

Figure 5. Persistence profiles of system-wide shocks to cointegrating vectors.

$$\mathbf{z}_t = \mathbf{z}_0 + \mathbf{b}_0 t + \mathbf{C}(1)\sum_{i=1}^{t} \varepsilon_i + \sum_{i=0}^{\infty} \mathbf{C}_i^* \varepsilon_{t-i}$$

where

$$\mathbf{b}_0 = \mathbf{C}(1)\mathbf{b} + \gamma, \quad \mathbf{C}(1) = \sum_{i=0}^{\infty} \mathbf{C}_i$$

and

$$\mathbf{C}_0^* = \mathbf{I}_4 - \mathbf{C}(1), \quad \mathbf{C}_i^* = \mathbf{C}_{i-1}^* + \mathbf{C}_i$$

The $\mathbf{C}_i$ needed to define these matrices are themselves defined in terms of the parameters of (10) as

$$\mathbf{C}_i = \sum_{j=1}^{p} \mathbf{C}_{i-j}\Phi_j, \quad \mathbf{C}_0 = \mathbf{I}_4$$

where

$$\Phi_1 = \mathbf{I}_4 - \alpha\beta' + \Gamma_1$$

$$\Phi_i = \Gamma_i - \Gamma_{i-1}, \quad i = 2, 3, \ldots, p-1$$

$$\Phi_p = -\Gamma_{p-1}$$

The persistence profile of the effect of a system wide shock on the jth cointegrating relationship is given by

$$\frac{\beta_j'(\mathbf{C}(1) + \mathbf{C}_N^*)\Sigma_{\varepsilon\varepsilon}(\mathbf{C}(1) + \mathbf{C}_N^*)'\beta_j}{\beta_j'\Sigma_{\varepsilon\varepsilon}\beta_j}$$

for $N = 0, 1, 2, \ldots$. The value of this profile will take the value of unity when $N = 0$, but should tend to zero as $N \to \infty$ if β_j is indeed a cointegrating vector. The persistence profile, when viewed as a function of N, provides information on the speed with which the effect of a system wide shock on the cointegrating relation e_{jt} disappears, and thus shows how many periods elapse before equilibrium is regained.

The 'demand for money' cointegrating vector, e_1, has almost identical profiles for both D and W, showing that a shock quickly dissipates, with equilibrium being regained after about ten quarters. For S, the return takes rather longer, approximately 6 years. The aggregate supply vector, e_2, is also identical for D and W. Here the initial shock tends to be amplified before slowly returning to equilibrium after 4 years. For S the return to equilibrium again takes about 8 years and the profile is much more cyclical. The third vector found for the D and S systems is difficult to interpret and is small and short lived for the former, but larger and much longer lasting for S.

From the general model (10), parsimonious systems of equations were then constructed. To economize on the amount of information provided, Table III just reports the estimated coefficients and S.E.s of the error corrections found to be significant in each individual equation of the three systems. The long-run demand for money function, e_1, is found to be significant in *every* equation of *every* system, while the aggregate supply vector, e_2, is significant in the money and price equations of every system, but only appears in the rental price equation for the W system and in the output equation for the S system. The third error correction appears in every equation of the D and S systems except the output equation of the former.

Figure 6 presents selected generalized impulse responses of the type proposed by Pesaran and Shin (1998). The generalized impulse response of a unit shock to the ith equation on the jth variable at horizon N is given by

$$\sigma_{ii}^{-1/2}\iota_j'(\mathbf{C}(1) + \mathbf{C}_N^*)\Sigma_{\varepsilon\varepsilon}\iota_i$$

where σ_{ii} is the ith diagonal element of $\Sigma_{\varepsilon\varepsilon}$ and ι_k is a selection vector containing unity as the kth element and zeros elsewhere. These impulse responses are invariant to the ordering of the variables in $\mathbf{z}_t$, and thus avoid the well-known problems associated with orthogonalized impulse responses.

Table III. Error correction adjustment coefficients

	$e_{1,t-1}$	$e_{2,t-1}$	$e_{3,t-1}$
(a) Money aggregate *D*			
ΔM	−0.1211	0.0749	−0.0285
	(0.0374)	(0.0333)	(0.0135)
ΔR	−0.4391	—	−0.3724
	(0.1005)		(0.0848)
Δry	0.1801	—	—
	(0.0263)		
Δp	0.0607	0.1139	0.0199
	(0.0093)	(0.0151)	(0.0061)
(b) Money aggregate *W*			
ΔM	0.0714	0.0981	
	(0.0261)	(0.0346)	
ΔR	−0.5467	0.4700	
	(0.2354)	(0.1676)	
Δry	0.1983	—	
	(0.0411)		
Δp	0.0797	0.1038	
	(0.0107)	(0.0141)	
(c) Money aggregate *S*			
ΔM	−0.3123	−0.1522	−0.2129
	(0.0685)	(0.0489)	(0.0830)
ΔR	0.4927	—	1.4282
	(0.1570)		(0.4550)
Δry	0.3593	0.2776	0.3722
	(0.0688)	(0.0549)	(0.0868)
Δp	0.0651	0.0381	0.1633
	(0.0146)	(0.0184)	(0.0329)

Attention is focused on the responses of the variables to shocks to the three monetary aggregates. A shock to the Divisia aggregate *D* leads to an initial positive response by the rental price, but which ultimately becomes negative. Output responds positively and reaches its total response after ten quarters. The price level, on the other hand, only begins to react after six quarters, but then eventually reaches a total response that is twice that of the output response. Shocks to the empirically weighted aggregate *W* have similar effects, although the rental price response is negative in both the short as well as the long run. Shocks to the simple sum aggregate *S* are rather different, however. The rental price response is now positive and there is virtually no output or price response in either the short or the long run.

Although the price response clearly dominates the output response in respect of shocks to both the *D* and *W* monetary aggregates, this evidence of a positive long-run output response is clearly at odds with the generally accepted macroeconomic notion of long-run money neutrality. It is important to note, however, that the latter is by no means universally accepted. Tobin (1965), for example, maintained that higher monetary growth should induce faster capital accumulation and productivity growth as economic agents switch out of monetary assets and into capital assets. Furthermore, in their review of the theoretical literature regarding the introduction of money into economic models, Orphanides and Solow (1990) point out that existing economic theory is somewhat inconclusive on the relationship between money growth, inflation and real output growth. They observe that 'for those that can bring themselves to accept the single consumer, infinite-horizon model as a reasonable approximation to economic life, superneutrality is a defensible presumption. All others have to be ready for a different outcome' (p. 225).

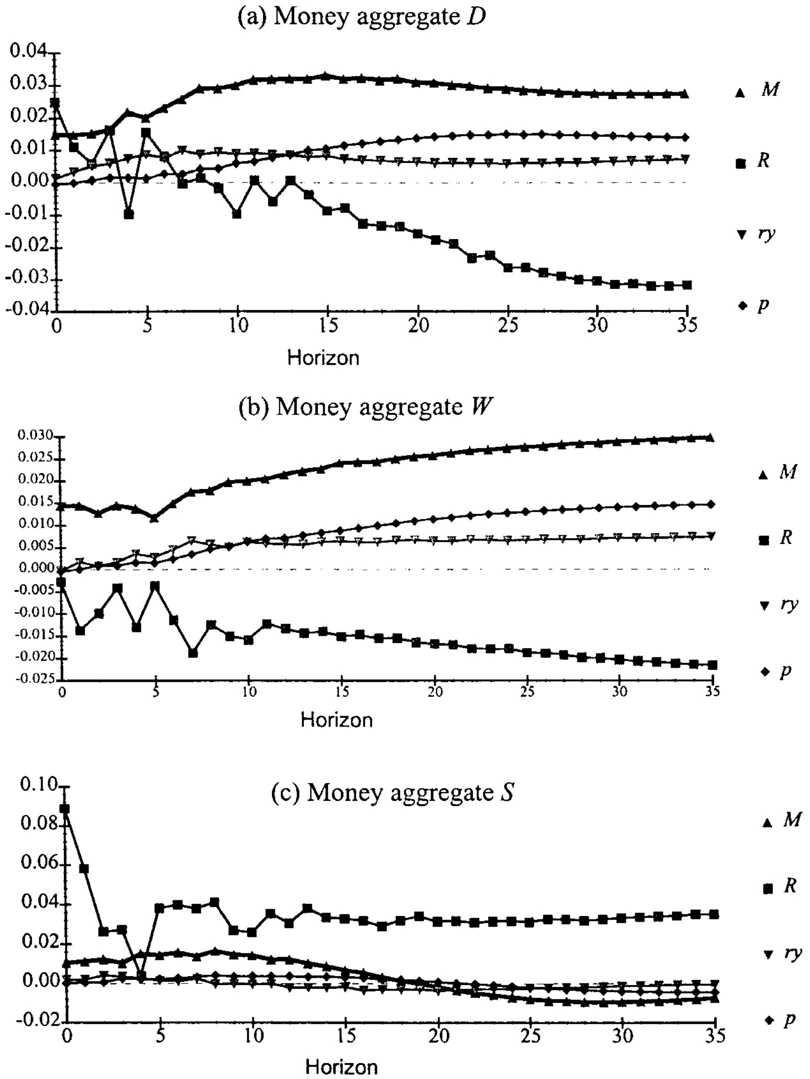

Figure 6. Generalized impulse responses to a monetary shock.

With respect to empirical work, most of the previous empirical analyses tend to support the proposition of long-run money neutrality, both across countries and across time (see, for example, Geweke, 1986; McCandless and Weber, 1994). It should be pointed out, however, that the majority of these studies utilize simple sum monetary aggregates and hence produce results consistent with our neutrality result in respect of simple sum money. Studies which have used weighted rather than simple sum monetary aggregates, however, have produced evidence against the neutrality proposition. Using Granger causality tests, Drake *et al.* (1998), for example, find that a range of alternative Divisia monetary aggregates Granger cause real output/expenditure in the long run.

Although there are acknowledged complications associated with cross-country studies, the evidence produced by these does not wholly support the money neutrality proposition. McCandless and Weber (1994), for example, do find evidence of a positive money–output correlation for the OECD sub-sample of their dataset. Further evidence that observed money–price (inflation) and money–output relationships might be sample specific is provided in a recent Bank of England study (Haldane *et al.*, 1998). Using time-averaged data from 80 countries, they find evidence of a clear one-to-one relationship between money growth and inflation and no long-run relationship between money and output growth. However, for a sub-sample of countries with average inflation below 15%, they find that the money growth-inflation relationship is less than one-for-one and that the money growth–output growth relationship is significantly positive. Clearly, this is a result which is entirely consistent with our own time series results.

Finally, it is worth noting that the evidence of a positive long-run output response to shocks to D and W is entirely consistent with our finding of a positively sloped aggregate supply relationship in the form of the second cointegrating vector in Table II. In the context of our four-equation VECM system, we can think of the monetary shocks as producing shifts in the aggregate demand schedule which would correspondingly trace out points on the aggregate supply curve in (p, ry) space. Hence, the fact that our monetary shocks are producing positive long-run output responses and positive long-run price responses of roughly twice the magnitude would lead us to expect a positively sloped aggregate supply relationship with a coefficient of approximately 0.5. This is exactly what we obtain. Conversely, for our results to be consistent with long-run money neutrality, we would expect to find a zero long-run output response and a quantity theory relationship linking money and prices one for one as the second cointegrating vector in the system (in addition to the money demand vector).

4. NON-NESTED TESTS USING A NOMINAL SPENDING EQUATION

If alternative monetary aggregates are to be useful as monetary indicators, we should be able to demonstrate that they provide superior information on final policy objectives than the currently utilized aggregates such as Divisia and simple sum. A traditional linear test of a variable's usefulness as a monetary indicator is its performance in an aggregate spending equation, typically referred to as a St. Louis equation. Clearly, it can be argued that the new empirically weighted aggregate, W_t, should perform well in a nominal income equation, given the methodology involved in its derivation. Nevertheless, this type of analysis provides a useful test of whether the PSS technique can produce a weighted monetary aggregate with genuine leading indicator properties. Furthermore, it also provides an opportunity to rank the new aggregate alongside comparable Divisia and simple sum aggregates which also potentially lay claim to having good leading indicators properties.

In this section, therefore, we compare the performance of D, W and S in the context of modified St. Louis equations. The basic St. Louis equation consists of the change in (the log of) nominal income as the dependent variable and lagged changes in (the logs of) the relevant monetary aggregate as the independent variables: the modification that we use is to also include lagged changes in nominal income and in (the logs of) nominal government expenditure, along with seasonal dummies. The modelling approach taken is that used in Mills (1983a). The best fitting model using just lags of nominal income and government expenditure was first obtained, and then lagged changes in each of the monetary aggregates were introduced, producing three competing models which can be compared on the basis of R^2, residual S.E. or BIC, for example. Using six lags of each variable, this produced a BIC ranking of 192.8 for W, 187.2 for S, and 186.7 for D. For R^2 and the residual S.E., the ranking of S and D was reversed, but both remained inferior to W.

These results suggest that, while the Divisia methodology is undoubtedly theoretically superior to the simple sum methodology, it does not always display superior information content. This result echoes those of other studies such as Drake *et al.* (1998). This type of result, in conjunction with the practical problems associated with the Divisia index approach outlined previously, may help to explain why many central banks, including the Bank of England and the US Federal Reserve, continue to utilize simple sum aggregates in monetary policy analysis despite their obvious theoretical shortcomings. The results presented in this paper, however, suggest that the new empirically weighted monetary aggregate may provide a useful alternative to both simple sum and Divisia aggregates for policy-makers.

5. CONCLUSION

This paper contributes to the monetary policy debate by devising a new weighted monetary aggregate for the UK based upon leading indicator properties for nominal income and an innovative approach to long-run modelling. From the perspective of policy-makers, the new weighted aggregate can be computed relatively easily and with minimal data requirements. Furthermore, the weights appear to be highly stable over time. The evidence of superior information content for this aggregate over the alternative Divisia and simple sum aggregates suggests that the PSS technique does provide a useful methodology for devising weighted monetary aggregates with good leading indicator properties.

Using standard cointegration techniques, we established that all three monetary aggregates produce sensible long-run money demand relationships which conform with economic priors such as price homogeneity. However, it is interesting to note that W produces a money demand function that is much closer to that of traditional empirical money demand analysis than either D or S. Whereas the weighted aggregate displayed the expected negative sign on the opportunity cost (rental price) variable, the Divisia aggregate produced a positive coefficient, although this result is not uncommon in Divisia studies (see Drake and Chrystal, 1994). Furthermore, whereas W produced a long-run income elasticity of around 1.6, both D and S produced much larger elasticities of around 3.0. Again, this evidence of relatively high income elasticties for money demand is quite common in Divisia studies (again see Drake and Chrystal, 1994).

The persistence profiles of system wide shocks to the cointegrating vectors confirmed the robust long-run equilibrating tendencies of the systems, although the speeds of adjustment were found to much quicker for W and D (around ten quarters) than for S (around 8 years). The impulse response functions of variables to monetary shocks confirmed that both W and D are potentially useful leading indicators for nominal income. Furthermore, in terms of the decomposition of nominal income changes into real income and prices, the results correspond closely to the traditional monetarist chronology. Specifically, in response to monetary shocks to both W and D, output responds positively and reaches its maximum response after around ten quarters. Prices, however, begin to increase only after around six quarters and take around 20 (for D) to 30 (for W) quarters to reach their maximum response. This latter result suggests that W may prove to be a better longer leading indicator for overheating and inflationary pressures than D. In contrast, with respect to the simple sum aggregate S, there is no evidence of a significant output or price impulse response to the monetary shock. This result, combined with the well known theoretical shortcomings of the simple sum methodology and the previous evidence of simple sum money demand instability, suggests that weighted monetary aggregates, such as W and D, should be the preferred aggregates for the conduct of monetary policy.

The impulse response functions also provide an important contribution to the much debated long-run money neutrality proposition. Although our results do show that the long-run price response for W and D is roughly twice that of the long-run output response, the latter's positive response can be taken as empirical evidence against the money neutrality proposition and the one-for-one money–price relationship evident in the Classical quantity theory.

Although not stressed in this paper, an important feature of the new aggregate W is that it has the potential to handle financial innovations, such as the introduction of new assets, as the weights can be endogenously revised based upon the evolving long-run relationship between nominal income and the component monetary assets. As stressed by Feldstein and Stock (1996), the issues of how to construct monetary aggregates and how to measure money growth in the presence of financial market innovations are crucial to the successful conduct of monetary policy. This was vividly illustrated in the USA by the 'missing M1 episode' in the early 1970s and by the 'missing M2 episode' in the early 1990s. With continual innovation in financial markets, this problem is likely to impact on UK monetary aggregates in the future and to pose potential problems for both simple sum and Divisia aggregates. Hence, this issue is likely to prove a fruitful area for future research and is one in which our new weighted monetary aggregate could potentially make a useful contribution, both in the UK and the USA and in other countries.

NOTES

1. Theoretically, Divisia indices are not defined when any of the monetary asset quantities are equal to zero in some time period(s). The alternative Fisher Ideal index is, however, valid in these circumstances.
2. Barnett *et al.* (1992) point out that the establishment of a weakly separable sub-utility function of monetary assets is a necessary precondition for the construction of admissable monetary aggregates. Testing for weak separability, however, is beyond the scope of this paper.
3. As interest rate data was provided on corporate sector and personal sector component assets separately, we construct weighted average interest rates across both sectors using the asset shares as weights. These weighted average interest rates are then assumed to be applicable to aggregate component asset holdings.
4. Income elasticities of 3 were imposed as just-identifying restrictions in both cases. Although they cannot be tested, it should be noted that the estimated 'elasticities', when only a normalization restriction was imposed on the cointegrating vector, were 2.81 for *D* and 3.22 for *S*.

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