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AGGREGATION, INFORMATION AND THE INSTABILITY OF ASSET DEMANDS

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Since the mid-1970s, monetary economists have been puzzled by the instability of the *M1* money demand function. Here, we broaden the investigation of this instability and its linkage to financial innovations, along two lines of inquiry. First, we consider whether the instability extends to the broader aggregates. Then, we investigate whether the means of aggregation is the source of this instability using the new aggregate based on the Divisia index. Through a battery of tests for functional stability of econometric relationships, we confirm instability across all asset demand equations. The finding is robust with respect to the breadth of the aggregate and the means of aggregation. We also examine the usefulness of the two competing aggregates, simple sum and Divisia, in the time series nexus in terms of their information content. We find the Divisia aggregate cannot improve the explanatory power of a normal income regression nor is it a superior target variable in a small macro model.

1. Introduction

Aggregation is a means of collecting many pieces of information in a single construct. It is attempted because it is believed that the appropriate index conveys significant information. In the case of financial assets, the Federal Reserve pools numerous asset supplies, through simple summation, into four aggregates, *M1*, *M2*, *M3*, and *L*, each more comprehensive than the other. One or all of these aggregates are viewed as central to our understanding of the economy and its movements. The division between groups is functional. *M1* contains the most liquid assets; these are essentially funds that are available for current purchases. *M2* contains small time deposits that are not instantaneously available. *M3* and *L* then contain assets of progressively longer duration.[1]

The financial innovations since the early 1970s have succeeded in blurring these functional distinctions. A variety of interest bearing but completely liquid assets has emerged. The most notable is the Negotiable Order of Withdrawal (NOW) account that has been part of *M1* since 1982. In addition, technological innovations have accelerated the mobility of funds

[1] The largest element of *M1* is checkable deposits. The primary component of *M2* is small time deposits. The bulk of *M3* is in large time deposits and money market mutual funds. *L* adds Treasury securities and commercial paper.

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between different functional categories. Elements of *M3* can become *M1* with just a telephone call or a few strokes of the computer keyboard.

Some have argued that these changes have resulted in substantial shifts in the behavior of the monetary aggregates and their role in the economy. As many have noted, the data seem to suggest such a change in the behavior of at least the lower level of monetary aggregates. Since the mid-1970s, monetary economists have been puzzled by the instability of the *M1* money demand function. We investigate here the link between this instability and these financial innovations, using two lines of inquiry. First, we consider whether the instability extends to the broader aggregates. Then we investigate whether the means of aggregation is the source of this instability. With regard to the latter, we consider a new aggregate based on the Divisia index proposed by the Federal Reserve and academic economists.[2]

Through a battery of tests for functional stability of econometric relationships, we confirm instability across all asset demand equations in the 1970s. The finding is robust with respect to the breadth of the aggregate and the means of aggregation. Something fundamental shifted in agents' decision rules during this period. Our analysis confirms this shift, but leaves the source as a question for further investigation.

We also examine the usefulness of the two competing aggregates, simple sum and Divisia, in the time series nexus in terms of their information content. We try to formally identify the informational distinctions between the two aggregates. We find the Divisia aggregate cannot improve the explanatory power of a nominal income regression nor is it a superior target variable in a small macro model. We remain open-minded, but far from convinced, about the utility of the new aggregates.

2. The 'missing money'

Nearly fifteen years have passed since the money demand equation was temporarily declared a closed line of research. Goldfeld's (1973) consensus specification incorporated the theoretical transactions-based framework of Baumol (1952) and Tobin (1956) and provided compelling empirical answers to questions of income and interest elasticities.

As is well known, the literature was re-opened by Goldfeld himself a mere three years later when his equation began to systematically overpredict the amount of real money balances actually held by the public. A vast literature suggested, without much success, innumerable places where the 'missing money' might be found. A good survey of these post-1976 developments may

[2] William Barnett has been the most forceful advocate of the new indices. He has written numerous articles on the topic, one of which may be found in the references. The Federal Reserve has been computing and publishing Divisia indices for several years. We cover these in greater detail in section 6.

be found in Judd and Scadding (1982). This literature has not come close to reforming the consensus of 1973. Our own work[3] has suggested some reasons as to why this has been the case. The function underwent structural change, in a statistical sense, much before it began to track poorly.

We wish to extend that work here by applying a number of the same tests to a broader range of assets. Looking at the higher level aggregates strengthens our earlier findings and supports strongly the notion of a change in regime. We begin with a review of the appropriate methodology.

3. Tests for stability

A great deal of work has been done in econometrics in the area of parameter stability. A general approach to the problem was proposed by Brown, Durbin and Evans (1975), hereafter BDE. A more recent survey of the literature may also be found in Dufour (1982).

When examining the stability of structural relationships, the analysis cannot center upon the characteristics of the OLS residuals because, in general, they are not independent and distributed $N(0,\sigma^2)$. BDE proposed a highly powerful method for transforming the residuals into a form in which departures from structural constancy could be assessed.

The BDE approach consists of estimating the regression coefficients recursively adding one observation at a time. Beginning with the basic linear regression model,

$$y_t = X_t\beta_t + \mu_t, \qquad t = 1, \ldots, T, \quad \mu_t \sim N(0, \sigma_t^2), \tag{1}$$

where the subscript t indicates that we will allow the parameter vector $\boldsymbol{\beta}$ to vary over time. If we let K equal the dimension of the $\boldsymbol{\beta}$ vector, we take the least squares estimate of β based on the first K observations, and then re-estimate successively $T-K$ times.

The null hypothesis H_0 is joint for both the stability of the regression parameters and the variance,

$$H_0\colon \beta_1 = \beta_2 = \cdots = \beta_T = \beta, \qquad \sigma_1^2 = \sigma_2^2 = \cdots = \sigma_T^2 = \sigma^2. \tag{2}$$

We obtain recursive residuals in the following manner. Let b_r be the OLS estimate of β based on the first r observations,

$$b_r = (X_r' X_r)^{-1} X_r' y_r. \tag{3}$$

[3] Mizrach and Santomero (1986). We isolate four regimes of monetary behavior in the 1959–1984 period.

We find the one-period ahead forecast error,

$$v_r = y_r - x_r b_{r-1}. \tag{4}$$

Under the null hypothesis, v_r should have mean zero and variance $\sigma^2 d_r$.

$$d_r = [1 + x_r'(X_{r-1}'X_{r-1})^{-1}x_r]^{\frac{1}{2}}. \tag{5}$$

Then, dividing v_r by d_r, we obtain a set of standardized prediction errors,

$$w_r = (y_r - x_r b_{r-1})/[1 + x_r'(X_{r-1}'X_{r-1})^{-1}x_r]^{\frac{1}{2}}, \tag{6}$$

having constant variance σ^2.

These recursive residuals, w_r, have nice properties which enable us to test quite easily deviations from normality. BDE demonstrate that under the null

$$\mathrm{E}(w_r w_s) = 0 \quad \text{for } r = s.$$

The first CUSUM test involves plotting over time

$$W_r = (1/\hat{\sigma}) \sum_{K+1}^{r} w_j, \qquad r = K+1, \ldots, T, \tag{8}$$

where $\hat{\sigma}$ is the unbiased estimate of σ.

$$\hat{\sigma} = (S_T / T - K)^{\frac{1}{2}}. \tag{9}$$

To a good approximation under the null hypothesis,

$$\mathrm{E}[W_r] = 0; \quad \mathrm{var}(W_r) = r - K; \quad \mathrm{cov}(W_r, W_s) = \min(r,s) - K. \tag{10}$$

That allows us to construct probability bounds,

$$y = d + c(t - K), \qquad K \leqq t \leqq T, \tag{11}$$

such that the probability that the CUSUM plot will cross the upper or lower bound is less than α where d and c are defined by

$$d = a(T-K)^{\frac{1}{2}}, \qquad c = 2a/(T-K)^{\frac{1}{2}}. \tag{12}$$

Approximate values corresponding to confidence levels of 99%, 95%, and 90% are 1.143, 0.948, 0.850. The higher the confidence level, the wider the bounds.

A second test, the CUSUM of squares test, is aimed at testing 'haphazard rather than systematic' types of movement that the first test is designed to capture. In the CUSUM test, the bounds will be violated after a consistent series of over-or-under predictions of small magnitude or a relatively few errors of large magnitude. For these errors to be detected though, they must be systematic, i.e., of the same sign. The CUSUM of squares negates the sign of the residuals and will detect changes solely on the basis of magnitude. It will thus illuminate significant but random changes in the regression coefficients or covariance matrix.

One plots against time,

$$S_r = \sum_{j=K+1}^{r} w_j^2 \Big/ \sum_{j=K+1}^{T} w_j^2 = S_r / S_T, \tag{13}$$

where S_r is the sum of the squared residuals of the regression based on the first r observations. The recursive residuals are related to the residual sum of squares in a convenient fashion.[4]

$$S_r = S_{r-1} + w_r^2. \tag{14}$$

A 95% confidence limit approximation for the plot of this statistic is

$$y_t = t/T + 1.36/(T-K)^{\frac{1}{2}}. \tag{15}$$

In many applications, we have the additional complication of having a non-scalar covariance matrix,

$$\mu_t \sim \mathrm{N}(0, \sigma^2 \Sigma), \tag{16}$$

where Σ is a $T \times T$ positive definite matrix. We know there exists a matrix P, which, if known, would enable us to make the transformation $P'\Sigma P$ back to the basic model. In general, we estimate p by either iterative least squares (Cochrane–Orcutt) methods or by maximum likelihood [e.g., Beach–McKinnon (1978)].

There is no assurance that the residuals from the true p and $\hat{p}$ will have the same asymptotic distribution as Durbin (1970) illustrates. Dufour (1982) has suggested considering a grid of values of p, possibly inside some neighborhood of $\hat{p}$, as a test of robustness.

4. Outline of specification and tests

The CUSUM methodology is ideal for testing the structural stability of the

[4]See Brown, Durbin and Evans (1975, Lemma 2).

money demand function without imposing any breakpoints a priori. We utilize the standard Goldfeld functional form which may be written quite generally as

$$m_t = \alpha_0 + \alpha_1 (M_{t-1}/p_t) + \alpha_2 Y + \alpha_3 OC_t. \tag{17}$$

m_t is the aggregate money index, *M1* to *L*, divided by the current period implicit price deflator for personal consumption, p_t.[5] We postulate a nominal adjustment hypothesis, dividing the lagged dependent variable by the current period deflator. Y is real personal consumption expenditures. OC_t represents our vector of opportunity cost variables. For the simple sum aggregates, OC_t is the six-month commercial paper rate. All the data are logged.

Estimation was done correcting for an AR(1) disturbance using the Beach–McKinnon maximum likelihood technique.[6] Tests for stability rely on the residuals from this equation.

5. Simple sum aggregate results

The first test of stability was conducted on the full range of aggregates *M1* to *L*. Recursive sample estimation beginning with 1970:1–1970:8 up to 1984:8 was undertaken. There were some problems with the fit of the interest rate variable, especially for *M3*, but with rare exceptions, coefficients were of the correct signs. The point estimates on the lagged dependent variable averaged 0.879, 0.931, 0.984, 0.991 for *M1* through *M4*, respectively. The serial correlation coefficients for each aggregate averaged 0.236, 0.545, 0.604, and 0.413.

CUSUM of squares tests rejected functional stability for all four aggregates. *M1* is fairly placid, crossing the lower boundary in June 1973, returning for good in July 1980. *M3* is the most problematic. It lies above the confidence limits from the beginning of the sample until November 1972 but remains steady from therein. See figs. 1 and 3.

M2 and *L* show much later breaks. *L* flirts with the lower boundary beginning in August 1974 and is always below the lower limit from February 1977 until its return in October 1979, as we see from fig. 4. *M2* first crosses the lower boundary in August 1975. It returns in dramatic fashion in December 1982 as can be seen in fig. 2.

Pinpointing exact points of discontinuity requires us to use Chow tests.

[5] In a quarterly equation, GNP and its deflator are generally used, but they become available only on a quarterly basis. Evidence from our other work suggests this substitution is of little empirical consequence.

[6] Betancourt and Kelejian (1981) have shown that the Cochrane–Orcutt algorithm can converge on non-global minima. For a truly robust estimate of the covariance matrix, a grid search procedure is called for. This is especially true given the problems of distinguishing the contribution of the lagged dependent variable and an AR(1) disturbance.

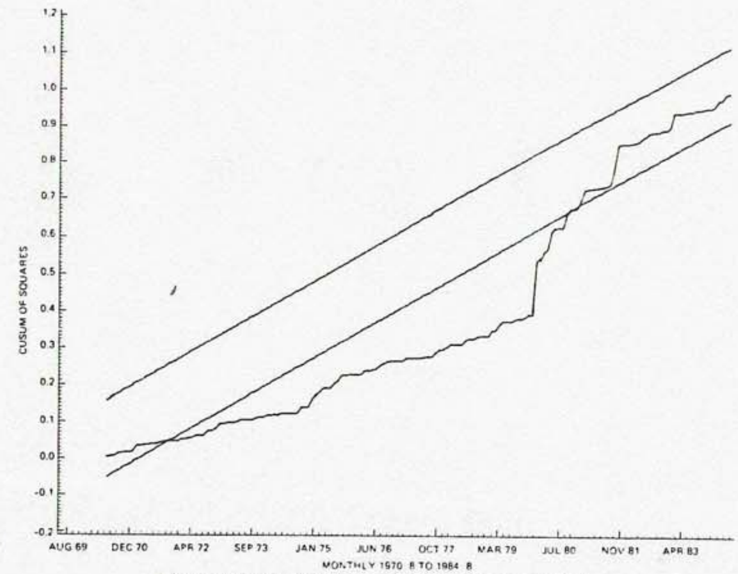

Fig. 1. CUSUM of squares – simple sum aggregate *M1*.

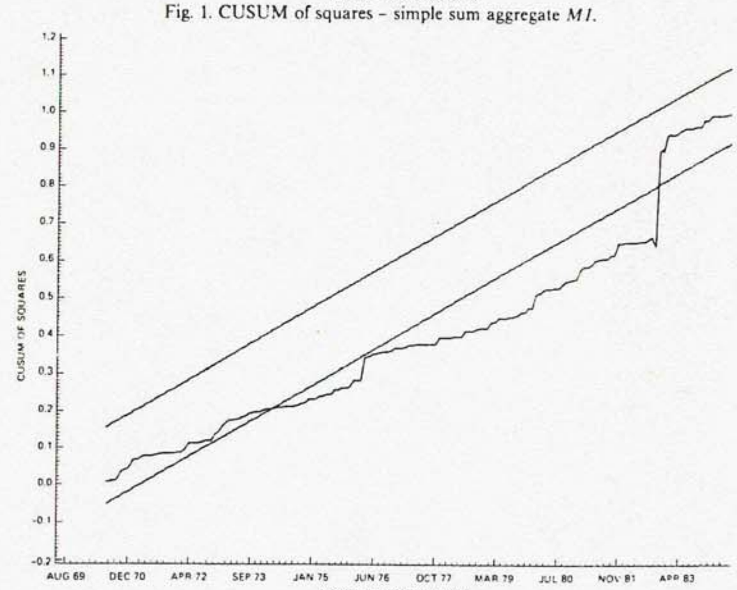

Fig. 2. CUSUM of squares – simple sum aggregate *M2*.

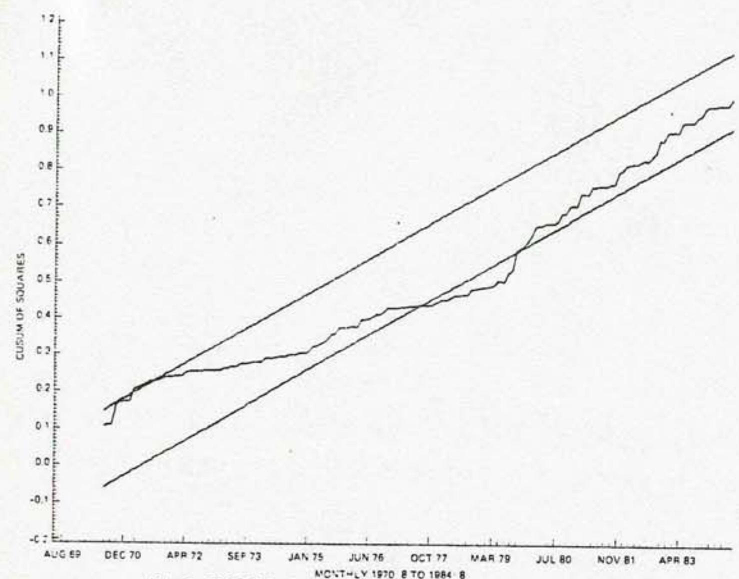

Fig. 3. CUSUM of squares – simple sum aggregate *M3*.

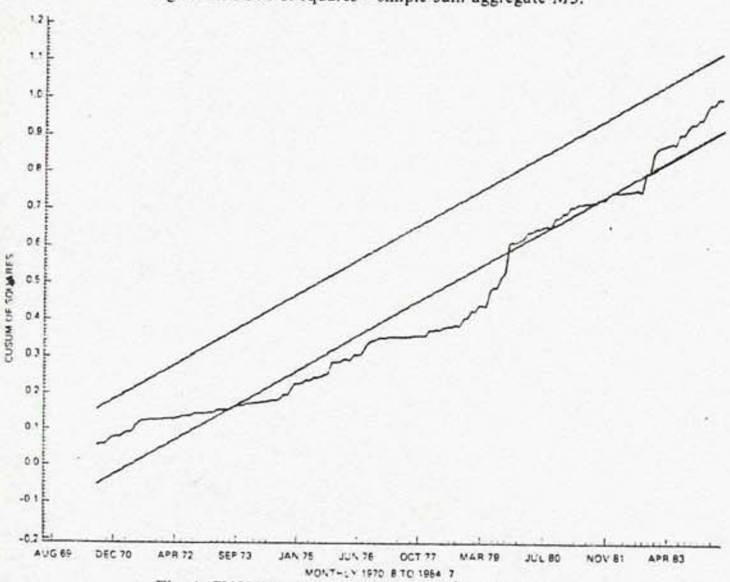

Fig. 4. CUSUM of squares – simple sum aggregate *L*.

Moving sample Chow tests, where the breakpoint is moved through the sample, confirmed these findings and revealed a number of early structural breaks.

The Chow statistic for *M2* has a global peak in 1982:12. *M3* and *M4* have global peaks in 1970:8, local peaks in 1974:7 and 1974:6, and then 1980:9 and 1980:5. *M1* has a life of its own, with a peak in 1974:12, a global peak in 1980:3, and a local blip in 1982:1.

The data, therefore, suggests clear instability in all the monetary aggregates tested. As noted above, some have suggested that this is due to the simple sum approach used to aggregate various monetary components. To see if this is the case, we now turn to the Divisia methodology to confirm if it can solve this substantial instability problem. We begin with a review of the theory.

6. The theory of Divisia aggregation[7]

It is clear both theoretically and empirically that the components of the monetary aggregates are not perfect substitutes. Yet the simple sum indices implicitly assume that money and Treasury bills are perfect substitutes, entering the index with an equal weight, and that shifts from one to the other are irrelevant. *L* would be completely unaffected by what is an important change in asset composition.

A theoretical literature on index numbers has developed to deal with this type of problem in a broad sense, proposing indices that formalize the way in which a composite vector of commodities may be properly aggregated.[8] Intuitively, we must find an index that, though it consists of many components, can be treated in utility terms like a single good. The existence of such an aggregate implies certain strong properties on the agent's utility function.[9] As Samuelson and Swamy (1974) note '. . . an economic quantity index . . . must itself be a cardinal indicator of ordinal utility'.[10]

Diewert (1976) has proposed what he calls the 'superlative' class of quantity index numbers. A quantity index is said to be superlative if it is exact for an aggregator function which can provide a second order approximation to any arbitrary, twice continuously differentiable, linearly homogenous aggregator function.

William Barnett (1980), then of the Federal Reserve, has advanced the theory and application of index number theory to the monetary aggregates. He has proposed use of the Tornqvist (1936)–Theil (1967) discrete time approximation of the Divisia quantity index, which is a member of Diewert's superlative class. We may express the discrete index in log form as

[7]This section draws heavily on Barnett (1980).

[8]Much of this literature is surveyed at a very high technical level in Diewert (1981).

[9]Diewert (1976) formally shows that this implies additive separability and homotheticity of the utility function.

[10]Samuelson and Swamy (1974, p. 598). This is also quoted in Barnett (1980).

$$\log Q_t^* - \log Q_{t-1}^* = \sum_k s_{jt}^- (\log m_{jt} - \log m_{jt-1}), \tag{18}$$

where $s_{jt} = \pi_{jt} m_{jt} / \sum_k \pi_{kt} m_{kt}$ is the expenditure share on the jth monetary asset, m. $s_{jt}^- = 0.5(s_{jt} - s_{jt-1})$.

What is essential about the index is that the components are multiplied by their user costs π_{jt},

$$\pi_{jt} = [p_t^*(R_t - r_{it})(1 - T_t)]/[1 + R_t(1 - T_t)], \tag{19}$$

where p_t^* is the true cost-of-living index, r_{it} is the own current holding period yield on component i, R_t is a benchmark maximum expected holding period yield, and T_t is the marginal tax rate. In computing the aggregates, all the factors in (19) cancel out of the share weights, s_{jt}^-, except for $(R_t - r_{it})$, which is a measure of the opportunity cost of holding the lower yielding asset. The growth rate of the index is a weighted average of the growth rate of the components.

To the extent that liquidity is inversely related to yield, the index measures the transactability of monetary assets, with currency providing the greatest 'moneyness' of all assets. In the example cited above of a shift from currency to Treasury bills, the index would show a decline in the 'monetary services' provided by the existing pool of assets.

In May 1985, the Federal Reserve made a comprehensive revision to their computation of the Divisia indices.[11] Because of the problem of incorporating new assets, the Fed has used the Fisher Ideal index,

$$Q_t = Q_{t-1} \left[\frac{\sum(\pi_{it} m_{it}) \sum(\pi_{it-1} m_{it})}{\sum(\pi_{it} m_{it-1}) \sum(\pi_{it-1} m_{it-1})} \right]^{\frac{1}{2}} \tag{20}$$

The Fed computes these indices *MSI1*, *MSI2*, *MSI3*, *MSI4*, using the asset stocks included in simple sum *M1–L*. There are 27 assets and corresponding own rates. We use these data in the stability and information tests that follow.

7. Divisia aggregate results

A case for the Divisia aggregates would, we argue, have to explain or at least reduce the structural instability we found in section 4. If in fact asset substitution was the source of the underlying functional instability reported above, then Divisia aggregation would not show parameter variation over the same interval. This is the issue investigated below.

[11] Farr and Johnson (1985) have called the revised data the monetary services indices. Construction of the indices is discussed in detail here.

In the Divisia equations, OC_t is obtained by dividing the total expenditure on the assets by the quantity index leaving, as Barnett (1980) shows, a functional price index depending solely on prices. The other variables are unchanged.

Recursive sample estimation was again undertaken. Coefficient estimates were of the right signs throughout, although the coefficient on the lagged dependent variable (LDV) exceeded one at some point in the sample for all the aggregates except for *MSI1*. Point estimates on the lagged dependent variable averaged 0.817 for *MSI1*, 0.937 for *MSI2*, 0.972 for *MSI3*, and 0.942 for *MSI4*. The corresponding covariance estimates showed strong first-order serial correlation, averaging 0.521, 0.808, 0.811, and 0.801, for *MSI1–MSI4*. The small gains in the speed of adjustment that are obtained using the Divisia constructs seem to come at the expense of much stronger serial correlation.[12]

The CUSUM tests decisively reject functional stability in all four cases. The CUSUM of squares for *MSI1* crosses a lower 95% confidence boundary plot in October 1972 and returns in January 1981; for *MSI2* the critical points are February 1972 (lower) and May 1981 (return); and finally, for *MSI3* and *MSI4* September 1972 (lower) and January 1972 (lower) with both returning in August 1981. See figures 5 to 8.

As a test for the robustness of the covariance estimate, we repeated the CUSUM procedure for a grid of autoregressive parameter values ranging from -0.9 to $+0.9$ by steps of 0.1. For the more reasonable positive estimates from $+0.1$ to $+0.9$, all four Divisia aggregates show breaks ranging from September 1971 to May 1972, all within a few months of the maximum likelihood estimates. The negative estimates ranged from December 1971 to June 1975, save for three outliers, all in December 1978 (*MSI1*, $p = -0.7$, *MSI2*, $p = -0.2$, and *MSI3*, $p = -0.3$). Even in the extremely unlikely case that we drastically mis-estimated the covariance matrix, the data seem to be sending us a very strong message of structural instability.

On the Chow tests, *MSI2–MSI4* show local peaks in the ratio of the sum of squared residuals in 1970:8, 1978:12, and 1981:4 (global). The CUSUM responds only slowly, as errors accumulate, but strong residuals do occur near the points that the CUSUM crosses the confidence intervals. *MSI1* is unstable for all splits in the sample until 1981:1, after which the F-statistic remains well below the critical level for all but one month (1982:7). Its global peak is in 1978:12.

These results are especially damaging for the high level aggregates. This is true even at *MSI4* where we have virtually the entire set of traded assets. This suggests some more fundamental instability in total asset demand.

[12] As the discussion in section 5 noted, in the stock adjustment model, it is quite difficult to discern between the contribution of the lagged dependent variable and the serial correlation parameter. The sum of these two is higher for the Divisia aggregates.

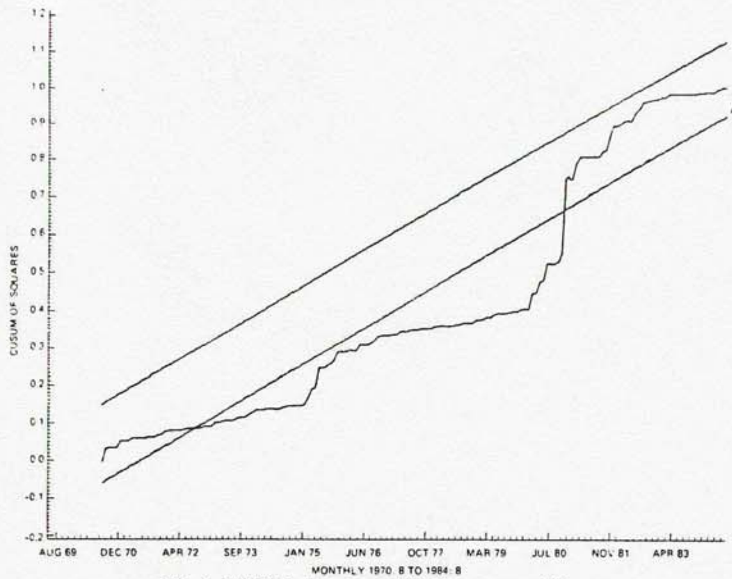

Fig. 5. CUSUM of squares – Divisia aggregate *M1*.

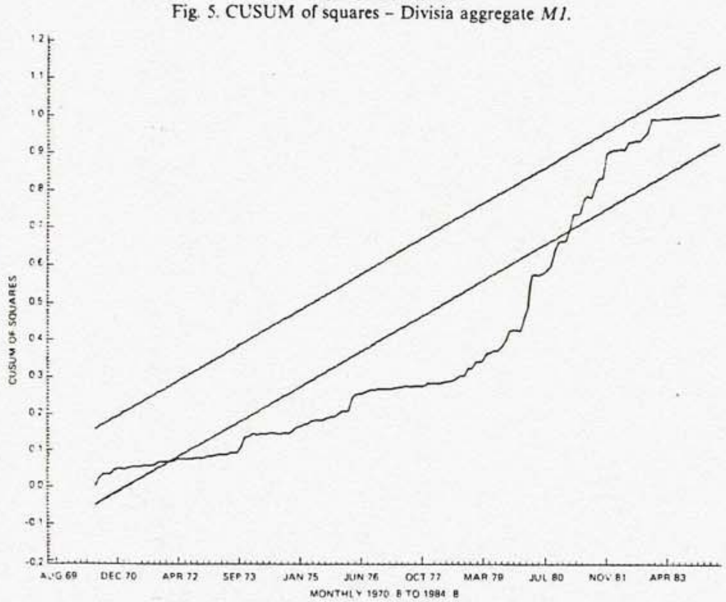

Fig. 6. CUSUM of squares – Divisia aggregate *M2*.

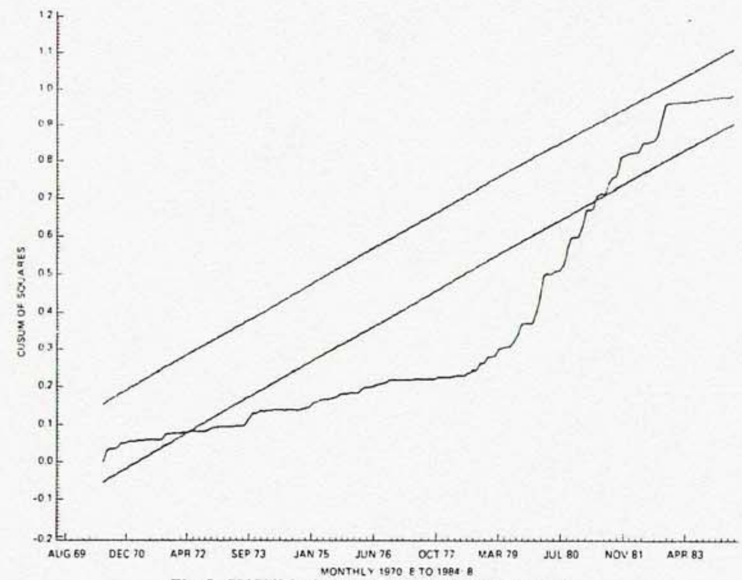

Fig. 7. CUSUM of squares – Divisia aggregate *M3*.

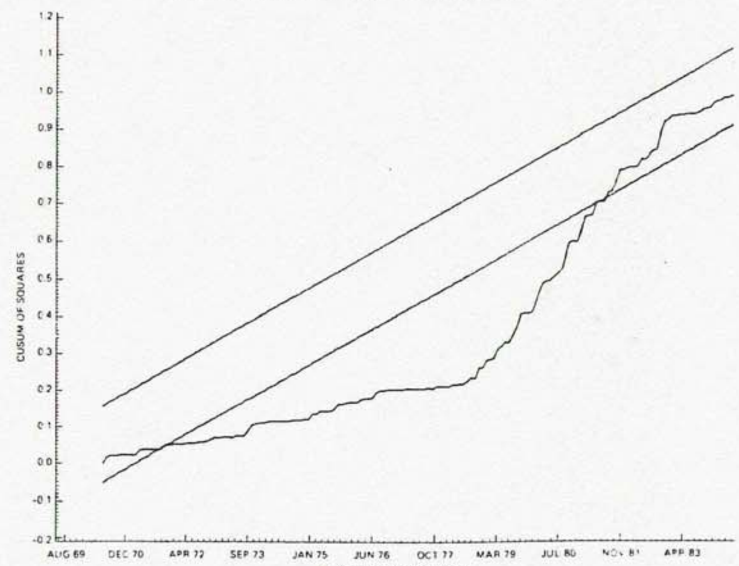

Fig. 8. CUSUM of squares – Divisia aggregate *M4*.

8. A comparison of Divisia and simple sum results

While both sets of aggregates can be regarded as being unstable, there seems to be no direct time correspondence between the CUSUM of squares plots. An analysis of the residuals, though, shows that the recursive forecast errors seem to be responding to similar phenomena. A six-month moving average of the absolute value of the residuals (the date shown is the startdate of the six-month period) revealed several periods of strong error concentrations common to the aggregates.

M1 and *MSI1* have strong residual activity in early 1975 and mid to late 1980. These periods generally coincide with two widely recognized periods of monetary instability – the missing money in the mid-1970s and the Volcker reserve policy in late 1979. *M2* and *MSI2* share a period of strong residuals in early 1976 and 1980, but the big shift in *M2* in December 1982 is not present in *MSI2*. *M3* and *M4* and their Divisia counterparts also exhibit large errors in 1980 and the big jump in December 1982. In late 1982, it should be noted, the strict reserve targeting begun by Volcker was relaxed considerably.

9. Which money matters?

9.1. A test of efficiency

Central to the usefulness of the Divisia monetary aggregates is its relevant and relative information in explaining the behavior of aggregate income. From this perspective, it is appropriate to ask which money matters?

Our first approach was to test whether simple sum or Divisia aggregated money contains sample information not contained in the other. Specifically, we test the orthogonality of the error term from an autoregression of nominal GNP on money and its own lagged values to other sample information sets.[13]

$$Y_t = \sum_{i=1}^{8} M_{t-1} + \sum_{i=1}^{8} Y_{t-1} + C_0 + u_t, \tag{21}$$

where Y_t is the log of nominal GNP, M_t is one of the four simple sum aggregates. C_0 is a constant term, and u_t is a random error term. A test of complete rationality would be whether

$$\mathrm{E}[u_t \mid \Omega_{it-1}] = 0 \quad \text{for } i = 1, \ldots, n, \tag{22}$$

for a potentially infinite set of possible information sets Ω_i available at time

[13] All regressions in this section utilize quarterly averages of the monthly data over the interval 1970:2–1984:2. Eight lags, excluding contemporaneous lags, were used in all tests.

$t-1$. We confine ourselves here to testing whether u_t is orthogonal to the Divisia aggregates. We regress $\hat{u}_t$ on a constant term and lagged values of the Divisia indices.

$$\hat{u}_t = C_0 + \sum_{i=1}^{8} MSI_{t-1} + e_t \tag{23}$$

F-tests confirm that the errors from (23) are efficient with respect to their Divisia counterparts. The sample information in the Divisia aggregates is thus fully incorporated into the simple sum indices.

9.2. Reduced form estimates of GNP

Ultimately, of course, we're interested in explaining *real* income movements. We next compare the fit of the various aggregates in the context of a simple IS/LM macroeconomic model,

$$\text{IS } y_t = \beta_1(p_t - {}_{t-1}p_t^*) + \beta_2 y_{t-1} + u_t, \tag{24}$$

$$\text{LM } M_t - p_t = \beta_3 y_t - \beta_4 r_t + e_t, \tag{25}$$

where y_t is log of real GNP, p_t is the implicit price deflator for GNP, r_t is logged user costs where m_t is a Divisia aggregate, and the log of the commercial paper rate where simple sum indices are used, ${}_{t-1}p_t^*$ is the one-period ahead forecast error of an eight-quarter autoregression on the implicit price deflator as a proxy for price expectations. u_t and e_t are white noise error terms.

We take the equilibrium solution of the model and estimate as if it were a reduced form for GNP,

$$y_t = \beta_5(m_t - p_t) + \beta_6(p_t - {}_{t-1}p_t^*) + \beta_7 y_{t-1} + \beta_8 r_t + w_t, \tag{26}$$

where $\beta_5 = 1/2\beta_3$, $\beta_6 = \beta_1/2$, $\beta_7 = \beta_2/2$, $\beta_8 = \beta_4/2\beta_3$, and $w_t = u_t/2 + e_t/2\beta_3$. The standard errors of this macro model uniformly favor the hypothesis that it is simple sum money rather than 'monetary services' that explains real income movements better. Simple sum *M2* performs the best, but any of the simple sum aggregates beats any of the Divisia aggregates. Results with quarterly data from 1970:2 to 1984:2 may be found in table 1.

10. Summary

Our tests have confirmed an instability in money demand functions stemming beyond the technique of aggregation. On the source of this

Table 1

Aggregate	Eq. (26) Standard error of	
	Divisia	Simple sum
M1	0.01140	0.01043
M2	0.01133	0.00935
M3	0.01139	0.01044
M4.L	0.01133	0.01049

instability we offer no clear answer. What we can report is that the factors responsible for this instability seem to affect all the aggregates in qualitatively the same manner. Our evidence suggests that the overall demand for the group of financial assets that enter the indices has gone through a major upheaval in the last fifteen years. Divisia aggregation does not call into question this finding, nor does it offer any insight into these shifts other than to say that asset substitution centering around money characteristics is not the source. Strong negative residuals from 1975–1977 using either the *MSI* or simple sum aggregates indicate that the missing money is not simply a fallacy of aggregation. One period-ahead forecast residuals using a Divisia monetary aggregate deteriorate considerably with changes in the institutional environment, even for the high level aggregates.

Turning to the information content of the series, tests of efficiency confirm that no sample information is gained by substituting in Divisia aggregates into time series equations for nominal GNP. Tracking real income movements, also, favors the simple sum aggregates. Therefore, the time series does not support the theoretically superior alternative. Empirical evidence does not, as yet, demonstrate Divisia aggregation to be a preferable practical alternative.

References

Barnett, W., 1980, Economic monetary aggregates – An application of index number and aggregation theory, Journal of Econometrics 14, 11–48.

Barnett, W., E. Offenbacher and P. Spindt, 1984, The new Divisia monetary aggregates, Journal of Political Economy 92, 1049–1085.

Baumol, W., 1952, The transactions demand for cash: An inventory theoretic approach, Quarterly Journal of Economics 66, Nov., 545–556.

Beach, C.M. and J.G. MacKinnon, 1978, A maximum likelihood procedure for regression with autocorrelated errors, Econometrica 45, 51–58.

Betancourt, R. and H. Kelejian, 1981, Lagged endogenous variable and the Cochrane–Orcutt procedure, Econometrica 49, July, 1073–1078.

Brown, R., J. Durbin and J. Evans, 1975, Techniques for testing the constancy of regression relationships over time, Journal of the Royal Statistical Society, Series B, 37, 149–192.

Diewert, W.E., 1976, Exact and superlative index numbers, Journal of Econometrics 4, 115–145.

Diewert, W.E., 1981, The economic theory of index numbers: A survey, in: A. Deaton, ed., The theory and measurement of consumer behavior (Cambridge University Press, Cambridge).

Dufour, J., 1982, Recursive stability analysis of linear regression relationships, Journal of Econometrics 19, 31–76.

Durbin, J., 1970, Testing for serial correlation in least squares regression when some of the regressors are lagged dependent variables, Econometrica 38, 410–421.

Farr, H. and D. Johnson, 1985, Revisions in the monetary services (Divisia) indexes of monetary aggregates, Special studies paper no. 189 (Federal Reserve Board, Washington, DC) May.

Goldfeld, S., 1973, The demand for money revisited, Brookings Papers on Economic Activity 3, 577–646.

Goldfeld, S., 1976, The case of the missing money, Brookings Papers on Economic Activity 3, 683–739.

Judd, J. and J. Scadding, 1982, The search for a stable money demand function: A survey of the post-1973 literature, Journal of Economic Literature 20, 993–1023.

Mizrach, B. and A. Santomero, 1986, The stability of money demand and forecasting through changes in regime, Review of Economics and Statistics 68, May, 324–328.

Samuelson, P.A. and S. Swamy, 1974, Invariant economic index numbers and canonical duality: Survey and synthesis, American Economic Review 64, 566–598.

Theil, H., 1967, Economics and information theory (North-Holland, Amsterdam).

Tobin, J., 1956, The interest elasticity of the transactions demand for cash, Review of Economics and Statistics 38, Aug., 241–247.

Tornqvist, L., 1936, The Bank of Finland's consumption price index, Bank of Finland Monthly Bulletin 10, 1–8.