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# HOW DOES DOMESTIC AND FOREIGN MONEY GROWTH AFFECT THE U.S. ECONOMY?

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ABSTRACT

This paper tests the hypothesis advanced particularly by McKinnon that the U.S. economy is strongly affected by the world supply of money and the U.S. effective exchange rate while the domestic money supply is of minor importance. This currency substitution hypothesis is tested by using monthly data for the floating exchange rate period 1973 - 1988. The empirical results give clear support to McKinnon's hypothesis.

CONTENTS

# 1 INTRODUCTION

Until recently it has been considered self-evident that nominal GNP and prices in the United States are mainly determined by domestic variables such as the money supply, the fiscal deficit etc. The proposition that foreign variables, including exchange rates, might play a significant role in this respect has not been taken very seriously, and even though it had been admitted that these variables have some minor effects, it has been assumed that floating exchange rates secure monetary autonomy. The developments in the 1980s surely make this closed-economy view cum the idea of monetary autonomy less obvious. Thus, it is not surprising that there have recently appeared several empirical analyses which try to assess the importance of foreign impulses. One way of explaining the growing importance of these impulses is to refer to the possibility of currency substitution which means that domestic residents hold foreign as well as domestic money balances. Specifically, multinationals and investment corporations maintain monetary balances denominated in more than one currency. By doing so, they can arbitrage interest rate differentials and fluctuations in currency rates and thus reduce the financing costs associated with worldwide operations. In the case of high-inflation or hyperinflation countries/periods, foreign currency holding is probably also motivated by the need of having some asset which performs the traditional roles of money. Some authors, for instance Brittain (1981), have claimed that currency substitution can explain the apparent instability of velocity in a number of industrial countries. McKinnon (1982, p. 324) goes even further in arguing that "in general, growth in the world money supply is a better predictor of American price inflation than is U.S. money growth".

Still, there is no wide agreement on the empirical relevance of currency substitution. Thus, for instance, Joines (1985) questions the practical importance of Brittain's (1981) findings. Cuddington (1983), in turn, claims that it is of limited importance in macro modelling but might have some implications for the estimation of money demand functions. The importance of currency substitution for the estimation of money demand functions, and more generally for monetary policy, has been stressed by several other authors, starting from Frenkel (1977) and more recently by, e.g., Fasano-Filho and Marquez (1987a,b). Even though empirical evidence in this respect is somewhat mixed, it may be fair to conclude that at least some evidence of currency substitution has been found in all empirical studies.

Clearly, McKinnon (1982), in proposing that currency substitution makes an appropriately defined world money supply rather than national money supplies relevant for studying both global and national inflation, puts forward the most far-reaching hypothesis. The interesting fact is that he also provides some empirical evidence, both in the form of stylized facts and estimation results, which all suggest that the above-mentioned hypothesis is not completely unwarranted.

As far as the econometric evidence is concerned, McKinnon (1984) exploits some simple nominal income (change rate) and inflation regression equations. To be a little bit more precise, he specifies the following set of competing equations:

(1) $$\Delta Y_t = a_0 + a_1 \Delta M_t^{us} + a_2 \Delta M_{t-1}^{us} + a_3 \Delta M_{t-2}^{us} + u_t$$

(2) $$\Delta Y_t = a_0 + a_1 \Delta M_t^{us} + a_2 \Delta M_{t-1}^{us} + a_3 \Delta EX_t + a_4 \Delta EX_{t-1} + u_t$$

(3) $$\Delta P_t = a_0 + a_1 \Delta M_t^{w} + a_2 \Delta M_{t-1}^{w} + a_3 \Delta M_{t-2}^{w} + a_4 \Delta EX_t + u_t,$$

where Y denotes the U.S. nominal GNP, P the wholesale price index, $M^{us}$ the U.S. money supply (M1), $M^w$ the world (i.e. 11 OECD countries) money supply, EX the effective U.S. exchange rate and u the error term. The equations were fitted into annual data covering the years 1958 - 1983, and the results clearly supported McKinnon's hypothesis. Thus, for example, the "closed-economy model" (1) was clearly outperformed by the composite equation (2) for the period 1972 - 1983, but not for the earlier period 1958 - 1971.

In essence, the present study contains another test for McKinnon's hypothesis. The basic differences between our and McKinnon's test procedures are the following: instead of annual data we use monthly data, concentrate solely on the floating exchange rate period 1973M5 - 1988M8, base the testing on unrestricted dynamic form (and thus, in the first place, impose no a priori lag structure for the equation to be estimated), and, finally, we do not compare separate equations but make use of the encompassing pricinple in testing non-nested models, as proposed by Mizon and Richard (1986), and specify a composite equation containing both domestic and foreign variables.

More specifically, we use the following unrestricted dynamic form as the point of departure:[1]

$$(4) \qquad \Delta X_t = a_0 + \sum_{i=1}^{k} a_i \Delta X_{t-i} + \sum_{i=1}^{k} b_i \Delta M^{us}_{t-i} + \sum_{i=1}^{k} c_i \Delta M^{w}_{t-i} + \sum_{i=1}^{k} d_i \Delta EX_{t-i} + u_t,$$

where $X = \{Y, P\}$, i.e. the model is estimated both in terms of $\Delta Y_t$ and $\Delta P_t$, the symbols being roughly the same as in the context of equations (1) - (3) (for more details, cf. section 2 below). Some preliminary analyses suggested that the proper lag length, k, is four and this value was used throughout the empirical analysis.[2]

The problem with the unrestricted form (4) is that most of the parameters cannot be estimated very precisely. Thus, the hypothesis that these parameters equal zero cannot generally be rejected at the standard levels of significance (with the F-test). Thus, we also derived a more parsimonious relationship, which turned out to be:[3]

$$(5) \qquad \Delta X_t = a_0 + a_1 \Delta X_{t-2} + a_2 \Delta M^{us}_{t-4} + a_3 \Delta M^{w}_{t-4} + a_4 \Delta EX_{t-3} + u_t.$$

This equation (5) provides a second set of results in the subsequent empirical analysis.

## 2 THE DATA

The following variables are used in this study: Y = nominal output = I•P, where I = industrial production (total manufacturing), P = the producer price index, $M^{us}$ = the United States money supply, $M^{w}$ = the OECD money supply (both M1 and M2 concepts are used here) and EX = the effective exchange rate. The latter variable is measured by using a weighted average effective exchange rate, EF (where the weights are relative GDP weights in the same way as in $M^{w}$) or, alternatively, by using the IMF (MERM) effective exchange rate index. These indices are graphed in Figure 1, and one can see that there is no qualitative difference between them although the exact numbers do, of course, differ. With both indices, a positive value of $\Delta EX$ reflects dollar appreciation - and vice versa.[4]

Most of the basic data come from the OECD Main Economic Indicators. The data for MERM quite obviously come from the IMF International Financial Statistics. Even though the data come from original tapes, they have been carefully checked and corrected using national data sources.

## 3 ESTIMATION RESULTS

Let us now turn to the estimation results which are presented in Tables 1 (for equation (4)) and 2 (for equation (5)). In the context of equation (5), estimation has been carried out by using, in addition to the OLS, Huber's robust M-estimator (for details of this estimator see Huber (1981)). This is motivated by the fact that the data contain some observations which can nearly be classified as outliers. This conclusion is also reinforced by the relatively high values of the Jarque-Bera normality test statistic.

Basically, the results can be easily summarized. Thus, the U.S. domestic money growth variable is clearly outperformed by the world money growth variable. Indeed, given the composite model (4) we cannot reject the hypothesis that the coefficients of $M^{us}$ are identically equal to zero. Moreover, the respective coefficients are either very small or even negative. The world money supply variable behaves in a completely different way: both the coefficient estimates and the respective standard deviations suggest that nominal U.S. output is indeed positively affected by the global money supply.

As far as the exchange rate variable(s) is concerned, one can conclude that, given both $M^{us}$ and $M^{w}$, it has an independent, although in some cases a marginal, effect on Y and P. Obviously, EX has a direct effect via imports and exports; however, it may also be that it signals changes in asset portfolios and can thus be interpreted as some sort of leading indicator (see e.g. McKinnon (1984) for further discussions of this point).

These results are reinforced by some time series analyses which focus on the pairwise relationships between Y, $M^{us}$, $M^{w}$ and EX. First, we estimate the cross-correlation function between Y, on the one hand, and the above-mentioned three explanatory variables on the other. Given these estimated cross-correlation functions we can compute the following $\chi^2$ statistics to test the hypothesis that the cross-correlation coefficients for lagged values of Y, or for the lagged values of one of the three X's ($M1^{us}$, $M1^{w}$, EX (= MERM)), or for both Y and X are identically equal to zero (all variables are expressed in first log differences).

| | $M1^{us}$ | $M1^{w}$ | MERM | 5 per cent critical value |
|---|---|---|---|---|
| $\chi^2_{10}$(lagged Y) | 39.06 | 16.93 | 4.09 | 18.31 |
| $\chi^2_{10}$(lagged X) | 9.05 | 11.38 | 6.28 | 18.31 |
| $\chi^2_{21}$(both Y & X) | 47.13 | 28.39 | 18.08 | 32.67 |

The values also rather clearly suggest that, first of all, causality runs from Y to $M^{us}$ but not vice versa. The relationship is rather weak between Y and $M^{w}$ so that nothing precise can be said of the direction of causation. Finally, the relationship Y and EX turns out to be contemporaneous (the contemporaneous cross-correlation coefficient is the only one which exceeds the assymptotic 5 per cent level). The causality structure between these four variables is also scrutinized by using a VAR(4) model similar to equation (4). The corresponding F-statistics are reported in Table 3. In the same way as the cross-corrrelation coefficients, these statistics indicate that $M^{w}$ causes Y, Y, in turn, causes $M1^{us}$, $M^{us}$ causes $M^{w}$, and, finally, a bidirectional causation exists between $M2^{us}$ and EX. The fact that world money is affected by lagged U.S. money is readily understandable because the former also includes the latter. The important thing, however, is that world money is clearly not Granger-caused by U.S. nominal output or the effective exchange rate. The same is not true, however, with U.S. money.

If one scrutinizes the bivariate relationships between these four variables in the frequency domain it turns out that the Y; $M^{us}$ relationship is characterized by only relatively long cycles, while with Y and $M^W$ one can discern a clear peak at high frequencies correponding to a two and half month cycle the other peaks corresponding to 3.5- and 1-year cycles (see the coherencies in Figure 2). As far as the cyclical pattern of Y; EX is concerned, very little can be said because the cohereciens do not follow any clear pattern.[5]

The above-mentioned results are surely somewhat surprising in the light of "conventional wisdom". One may suspect that the results reflect some exceptional episodes or periods during the floating exchange rate regime. To examine this possibility we estimate a varying-parameter (Kalman-filter) model in terms of equation (5). The obtained smoothed coefficient estimates for $M1^{us}$, $M1^{w}$ and MERM are displayed in Figure 3 (using the M2 concept or the weighted average effective exchange rate index (EF) did not produce any qualitative difference).[6]

Clearly, the coefficients are, after the initial volatile periods, strikingly stable and consistent with the results reported above, particularly in terms of the impotence of the U.S. money growth variable. Moreover, one cannot really say that, for instance, the "nonborrowed reserves targeting period" 1979 - 1982 would clearly differ in terms of the relative effectiveness of U.S. monetary policy.[7]

## 4 CONCLUDING REMARKS

This paper has demonstrated that currency substitution, in a way proposed by McKinnon, is of crucial importance in explaning U.S. nominal output growth. Thus, one can claim that floating exchange rates have not secured national monetary autonomy for all countries, not even for the United States. It is therefore hard to see that the United States should pursue a policy of controlling the year-to-year changes in the purely national monetary aggregates, and thus sterilizing the domestic monetary impact of foreign interventions in the exchange markets. Instead, international monetary policy co-ordination is required so that the world's money supply could be stabilized. Even though one could not agree of the proper monetary policy co-ordination rule, the important thing is to recognize that purely national monetary and credit aggregates cannot really be used as liquidity or leading indicators.

Table 1. Test statistics for lagged explanatory variables

| Y; P | $M^{us}$ | $M^{w}$ | EX | SEE | DW | J-B | definition of variables |
|---|---|---|---|---|---|---|---|
| .499 (9.20) | .116 (2.10) | .934 (2.68) | -.028 (1.79) | 1.22 | 2.01 | 8.05 | Y: M1, EF |
| .504 (9.37) | .120 (2.08) | .918 (2.70) | -.036 (2.10) | 1.22 | 2.01 | 7.53 | Y: M1, MERM |
| .426 (9.14) | .601 (0.94) | .625 (2.51) | -.069 (1.71) | 1.22 | 2.00 | 8.30 | Y: M2, EF |
| .434 (9.48) | .651 (0.90) | .609 (2.60) | -.093 (2.15) | 1.21 | 2.00 | 7.47 | Y: M2, MERM |
| .481 (8.60) | -.091 (1.10) | .344 (1.83) | .024 (1.83) | 0.91 | 1.02 | 118.47 | P: M1, EF |
| .475 (8.59) | -.083 (1.08) | .347 (1.84) | .041 (1.99) | 0.91 | 2.02 | 117.30 | P: M1, MERM |
| .408 (7.70) | -.033 (0.52) | .648 (1.47) | .013 (1.88) | 0.91 | 2.03 | 119.09 | P: M2, EF |
| .405 (7.83) | -.051 (0.52) | .642 (1.50 | .023 (2.03) | 0.91 | 2.03 | 118.20 | P: M2, MERM |

The dependent variable is either Y or P. The reported numbers are sums of the coefficient estimates, numbers in parentheses are F-test statistics for the hypothesis that these coefficients are identically equal to zero. SEE = the standard error of estimate, DW = the Durbin - Watson autocorrelation test statistic and J-B = the Jarque - Bera test statistic for the normality of residuals. The F-distribution has the following critical values with 4,167 degrees of freedom: 1.97 (10 %), 2.42 (5 %) and 3.43 (1 %). The 5 per cent critical value for the chi-square distribution with 2 degrees of freedom is 5.99.

Table 2. OLS and Robust estimation results of equation (5)

| const. | Y | $M^{us}$ | $M^{w}$ | EX | SEE | DW | J-B | estimator and definition of of variables |
|---|---|---|---|---|---|---|---|---|
| .001<br>(0.34) | .367<br>(5.45) | -.256<br>(1.45) | .768<br>(2.86) | -.086<br>(2.00) | 1.23 | 1.77 | 24.27 | OLS, M1, EF |
| .001<br>(0.27) | .370<br>(5.50) | -.253<br>(1.44) | .717<br>(2.87) | -.099<br>(2.10) | 1.23 | 1.77 | 23.03 | OLS, M1, MERM |
| -.003<br>(0.96) | .361<br>(5.37) | .191<br>(0.50) | .698<br>(2.46) | -.093<br>(2.14) | 1.23 | 1.78 | 18.76 | OLS, M2, EF |
| -.003<br>(1.08) | .364<br>(5.42) | .219<br>(0.57) | .702<br>(2.48) | -.111<br>(2.31) | 1.23 | 1.79 | 17.54 | OLS, M2, MERM |
| .002<br>(0.86) | .390<br>(6.47) | -.331<br>(2.09) | .695<br>(3.09) | -.080<br>(2.06) | 1.24 | 1.78 | 28.16 | ROB, M1, EF |
| .001<br>(0.79) | .392<br>(6.51) | -.328<br>(2.08) | .694<br>(3.10) | -.089<br>(2.10) | 1.24 | 1.78 | 26.86 | ROB, M1, MERM |
| -.001<br>(0.34) | .389<br>(6.30) | -.022<br>(0.06) | .676<br>(2.59) | -.084<br>(2.10) | 1.24 | 1.78 | 23.11 | ROB, M2, EF |
| -.001<br>(0.48) | .390<br>(6.31) | .007<br>(0.20) | .682<br>(2.61) | -.097<br>(2.19) | 1.23 | 1.79 | 21.71 | ROB, M2, MERM |

The dependent variable is Y. Numbers in parentheses are (asymptotic) t-ratios. Diagnostic statistics (not displayed) did not express any heteroskedasticity, and thus computing White's heteroskedasticity-adjusted t-ratios did not make any difference. ROB denotes Huber's M-estimator. The tuning constant 1.345 is used for all four reported equations.

Table 3. Granger-causality test statistics

| predicted variable | Y | $M^{us}$ | $M^{w}$ | EX | definition of variables |
|---|---|---|---|---|---|
| Y | 9.20 | 2.10 | 2.68 | 1.79 | M1, EF |
| Y | 9.37 | 2.08 | 2.70 | 2.10 | M1, MERM |
| Y | 9.14 | 0.94 | 2.51 | 1.71 | M2, EF |
| Y | 9.48 | 0.90 | 2.60 | 2.15 | M2, MERM |
| $M^{us}$ | 2.55 | 3.46 | 0.36 | 0.91 | M1, EF |
| $M^{us}$ | 2.48 | 3.47 | 0.36 | 1.01 | M1, MERM |
| $M^{us}$ | 1.10 | 5.66 | 0.71 | 3.01 | M2. EF |
| $M^{us}$ | 1.19 | 5.68 | 0.74 | 3.27 | M2, MERM |
| $M^{w}$ | 0.66 | 4.44 | 3.68 | 0.93 | M1, EF |
| $M^{w}$ | 0.62 | 4.41 | 3.81 | 0.96 | M1, MERM |
| $M^{w}$ | 1.40 | 3.97 | 6.67 | 0.39 | M2, ED |
| $M^{w}$ | 1.38 | 3.96 | 6.75 | 0.52 | M2, MERM |
| EX | 0.43 | 0.62 | 0.50 | 4.24 | M1, EF |
| EX | 0.63 | 0.72 | 0.43 | 3.42 | M1, MERM |
| EX | 0.18 | 2.49 | 0.35 | 4.13 | M2, EF |
| EX | 0.26 | 2.66 | 0.25 | 3.13 | M2, MERM |

Reported numbers are F-test statistics for the four lagged terms of each variable. The corresponding critical values are: $F_{.05} = 2.42$ and $F_{.01} = 3.43$.

FOOTNOTES

1) Here we abtract from all problems of identifying the underlying structural model and defining the exact meaning of causation. See e.g. Jacobs et al. (1979) for more exact treatment of these problems. See also Pikkarainen and Virén (1989) for further discussion of interpreting the results from a model similar to (4).

2) We also introduced the domestic (ex post) real interest rate as an additional explanatory variable in (4) but dropped it because the explanatory power was extremely small.

3) That was done by dropping the "insignificant" lagged terms for all four explanatory variables. The resulting specification turned out to be invariant with respect to all variable definitions of M and EX.

4) We also used quarterly data so that Y was measured by nominal GNP. A copy of the tables containing the corresponding estimation results is available upon request from the author.

5) In testing the hypothesis of zero coherency the critical values (at the 5 and 10 per cent level of significance with the F distribution) would here be, given the Tukey-Hanning window of length 40, $\hat{\rho}$ = .682 and .620, respectively (see Koopmans (1974) for details of the test procedure).

6) The standard error of estimate in this case is 1.22. Estimation was carried out using the IAS-SYSTEM 3.6 software package. As far as the standard deviations of these coefficients are concerned, some flavour can be obtained from the following values for 1980M1 and 1988M8, respectively: b:$M^{us}$ .436 and .175; b:$M^{w}$ .378 and .248; b:EX .095 and .047. Clearly, the coefficient of $M^{us}$ does typically not deviate from zero by the magnitude of two times the respective standard deviation.

7) The results for the United States were so impressive that we decided to apply the same framework also for the European Community. Thus, we estimated equation (4) using monthly EC data for the EMS period 1979M3-1988M10. Specifically, the ECU/USD rate was used for EX.

The results were to some extent analogous to those obtained for the United States. However, the relationships were generally weaker and the direction of causation could not be easily interpreted although the coefficients typically behaved according to McKinnon's hypothesis. This can be seen from the following Granger causality test statistics which are analogous to those reported in Table 3 (now $F_{.05}$ = 2.47).

| predicted variable | Y | $M^{ec}$ | $M^{w}$ | EX | definition of M |
|---|---|---|---|---|---|
| Y | 4.37 | 1.71 | 1.43 | 0.50 | M1 |
| Mec | 0.67 | 2.62 | 2.29 | 1.96 | M1 |
| Mw | 0.73 | 3.06 | 2.21 | 2.47 | M1 |
| EX | 0.33 | 1.24 | 1.53 | 2.61 | M1 |
| Y | 4.60 | 0.72 | 0.67 | 0.45 | M2 |
| Mec | 0.26 | 2.91 | 2.69 | 0.36 | M2 |
| Mw | 0.83 | 1.34 | 4.97 | 1.56 | M2 |
| EX | 0.22 | 0.41 | 1.22 | 2.18 | M2 |

Neither nominal output nor the exchange rate do appear to be caused by other variables. Instead, $M^{ec}$ and $M^{w}$ are interrelated, particularly in the case of the M1 concept of money. All in all, we can conclude that the EC data do not give similar support to McKinnon's hypothesis as do the U.S. data. Even so, these data also show that the control of purely national monetary aggregates as a basic policy tool can be questioned.

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Figure 1 U.S. Effective Exhange Rate Indices

Weighted average, 1980M1 = 100
MERM, 1980M1 = 100

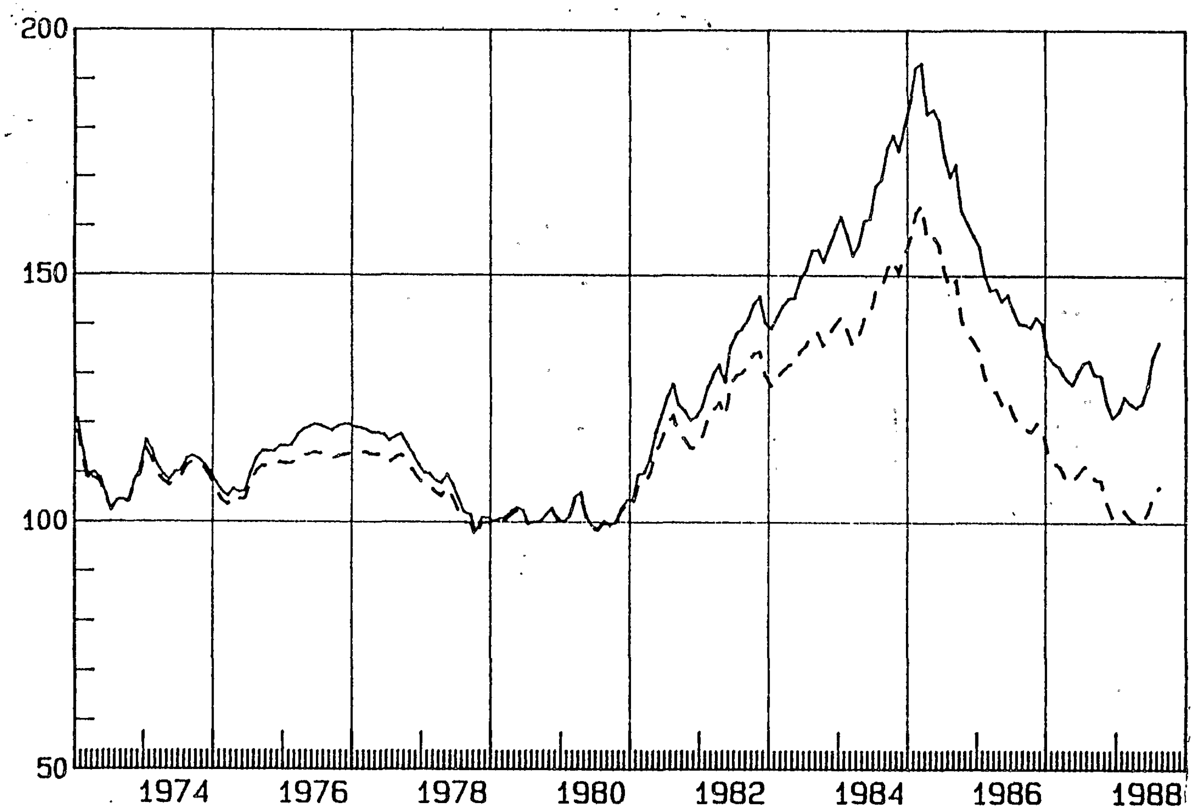
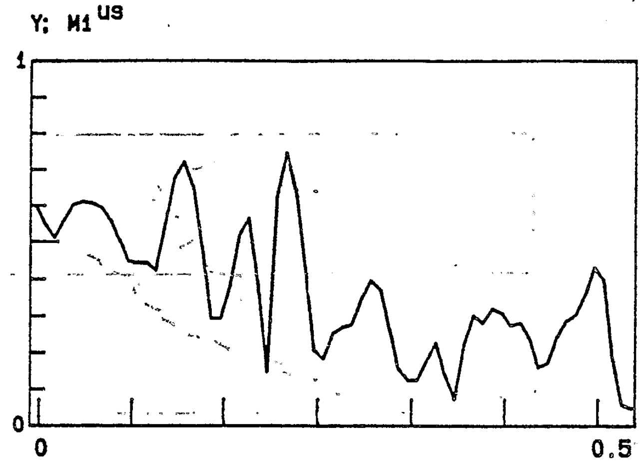
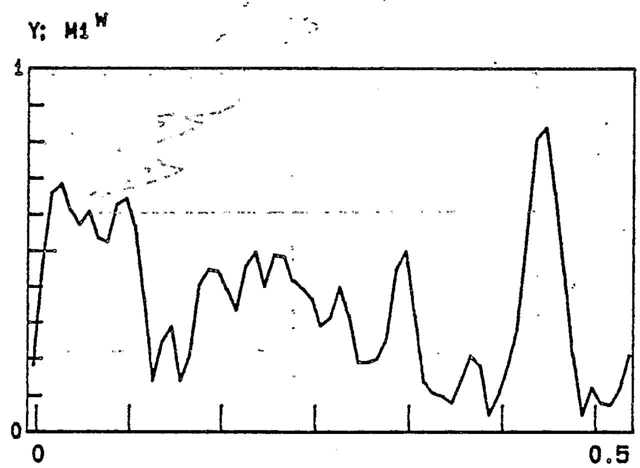


Figure 2

Coherencies between Y, $M^{US}$, $M^{W}$ AND EX



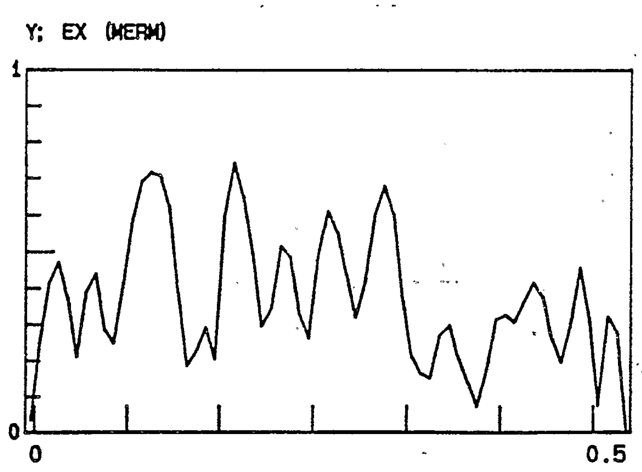
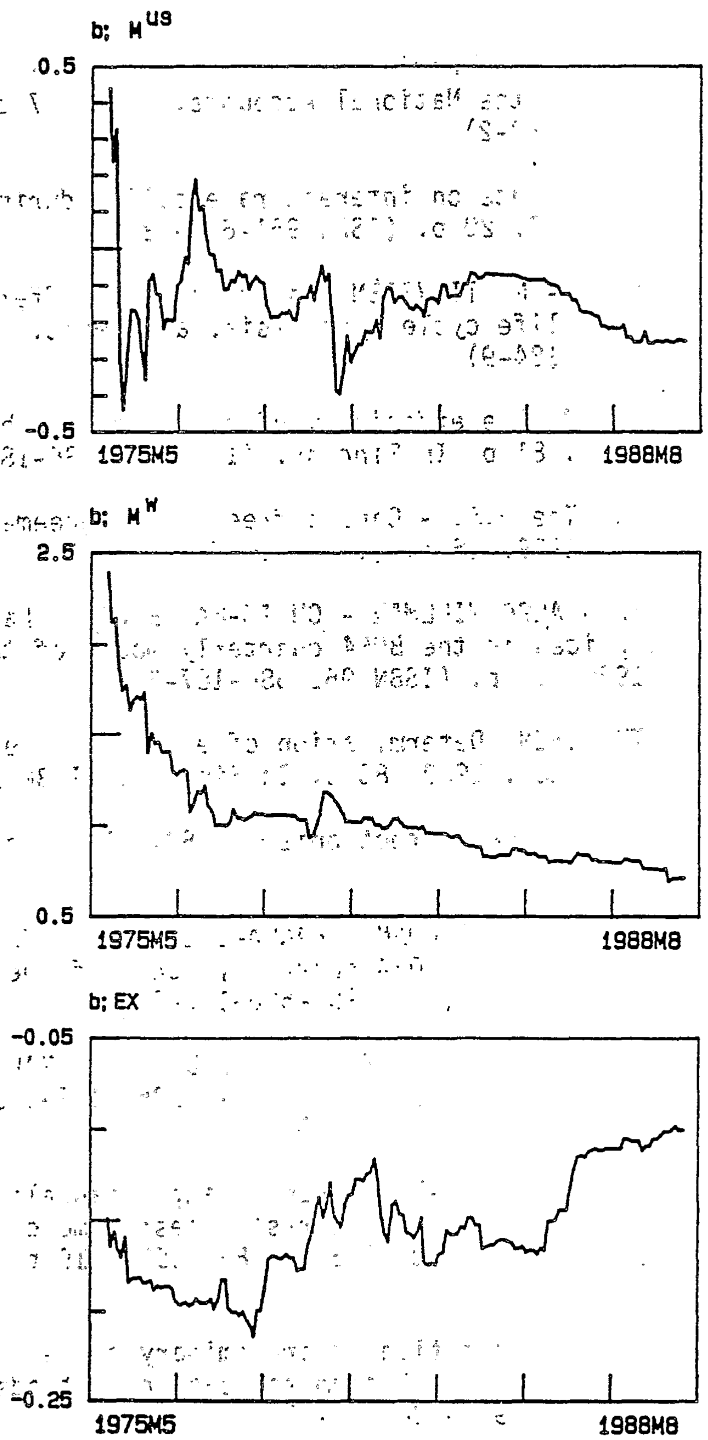



Figure 3

Time-Varying Coefficients of $M^{US}$, $M^{W}$ and EX



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