THE LONG AND SHORT RUN DETERMINANTS OF THE VELOCITY OF BROAD MONEY: SOME INTERNATIONAL EVIDENCE

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ABSTRACT

This paper investigates the long and short run determinants of the velocity of broad money in five industrial countries. The velocities are first decomposed into their long run trend and short run cyclical components, using the multivariate stochastic detrending approach advanced by Vahid and Engle (1993) and Hecq, Palm, and Urbain (2000). The paper then offers evidence on the determinants of the trend and cyclical components for each of the sample countries. Our evidence indicates that while fundamentals explain velocity in the long run, short run velocity is impacted by shocks to the fundamentals and also by global contagion factors. (JEL: G15, C32)

INTRODUCTION

The behavior of the M2 velocity of money continues to be the subject of contentious theoretical and empirical research in the literature. Theory suggests that the existence of velocity as a stable function of a few key macroeconomic variables, such as the levels of the interest rate and wealth, is essential to the use of monetary aggregates as important components of monetary policy (Friedman, 1956, 1988; Laidler, 1997). This fact largely accounts for the continuing research regarding the empirical issue of whether velocity is indeed stable, or at least predictable, albeit with mixed results. Gould and Nelson (1974), Nelson and Plosser (1982), Haraf (1986), Friedman and Kuttner (1992) and Serletis (1995) have presented evidence which they interpreted as inconsistent with the stability of velocity. However, later developments in the concept of cointegration clarified that the findings of a random walk (by Gould and Nelson, 1974, and Nelson and Plosser, 1982) do not rule out the possibility that the demand for money and thus velocity may bear a stable relationship to other variables. Latane (1962), Meltzer (1963), Wilbratte (1975), Lucas (1980), Heitzel and Mehra (1989), Siklos (1993), Choudhry (1996), Bordo, Jonung, and Siklos (1997), Mehra (1997), Anderson and Rasche (2001), and Mendizabal (2006) have found evidence in support of the proposition that velocity is stable, especially for broadly defined money. A possible weakness of the existing work on the behavior of the velocity, however, is the absence of a distinction between its trend and cyclical components, as each may be influenced by completely different sets of determinants. As a result, there is risk that statistical tests of the cyclical volatility of velocity in response to short run shocks may be misinterpreted as instability in the trend velocity in the long run.

The present study attempts to advance the evidence regarding the behavior of velocity by identifying its trend and cyclical components, using a recently developed multivariate decomposition approach. This approach relies on extensions of the work of Beveridge and Nelson (1981) to the multivariate case by Stock and Watson (1988) and Gonzalo and Granger (1995), who showed that each member of a group of interrelated variables can be represented as a linear combination of several random walks and cycles. It was further demonstrated that, in the event some of the random walks are common to some of the variables, i.e., if the variables are cointegrated in the sense of Engle and Granger (1987) and Johansen (1988), the multivariate composition can be more parsimoniously expressed in terms of the observable levels of the variables.

Engle and Kozicki (1993) and Vahid and Engle (1993) further advanced the application of this technique by showing that the variables in a multivariate system can have both common stochastic cycles and trends, with the total number of common cycles and trends equal to or exceeding the number of variables in the system. Vahid and Engle also demonstrated that, by taking the presence of common cycles and trends into account, one can more efficiently estimate the permanent and transitory components. A restriction of the Vahid-Engle decomposition, however, is the condition that the total number of common trends and cycles precisely equal the number of variables in the system. Thus, in order to address the more general case in which the number of common trends and cycles exceeds the number of variables, we rely on the less restrictive decomposition technique of Proietti (1997) and Hecq, Palm, and Urbain (2000).

Using these methods, the present study investigates the common time series properties of the velocity of money in five major industrial countries. As the first step in our investigation, we establish that for each of the sample countries, velocity is both difference stationary and autoregressive. Next, we test for the presence and number of common stochastic trends and cycles in the time series. Since this number is found to exceed the total of five variables, violating the conditions of the Vahid and Engle model, we employ the method of Proietti (1997) and Hecq, Palm, and Urbain (2000) to extract the permanent and transitory components of the velocity series. We complete our tests by demonstrating that the long run trends in velocity are explained in terms of such fundamentals as the interest rate, stock prices as a proxy for wealth, and oil prices. In particular, our findings shed light on the role of stock prices on the demand for money and hence on velocity. Specifically, we show that for four of the five countries in our sample, the substitution effect of the stock market, which decreases the demand for money and thus increases velocity as wealth holders shift funds from money to stocks, dominates the wealth effect, which causes the asset holders to increase their money holdings in response to stock price appreciation (Friedman, 1988). Our results also provide evidence that the short run fluctuations in velocity reflects increased short term volatilities in such variables as stock and oil prices as well as contagion effects among countries.

The organization of this paper is as follows: The following section provides a description of the decomposition approach employed. The next section presents the data used and the empirical results, and the final section concludes.

THE MODEL

Friedman (1956) provides a forceful case for the importance of velocity to monetary policy, basing his assertion on both theoretical underpinnings and statistical evidence. Essential to Friedman's argument is the stability of the velocity of money in the standard equation of exchange, which can be written as follows:

$$MV = PY \tag{1}$$

where M = the money supply broadly defined, V = the income velocity of broad money, P = the general price level, and Y = real national output. Clearly, money and nominal income can have a stable relationship, provided that velocity can be characterized as a stable function of a limited number of macroeconomic variables. To show that velocity is indeed a stable function, Friedman first reinterprets the quantity theory of money as essentially a theory of the demand for money, thus breaking from his classical predecessors. More specifically, Friedman posits that the demand for real money can be expressed as a function of identifiable and measurable economic variables representing the opportunity cost of holding money, the return on financial assets, income, and wealth, as follows:

$$M/P = f(Y,W,R_M, R_B, R_E,u) \tag{2}$$

where Y = real income, W = nonhuman wealth, R_M = the return on monetary assets, R_B = the return on bonds, R_E = the return on equities, and u signifies the unavoidable presence of an undetermined set of variables affecting the demand for money. Thus, all of the arguments in equation 2 are candidate variables which may significantly affect money demand in empirical testing. Meltzer's (1963) pathbreaking work demonstrated that the demand for money is stable over a long sample period and for various subperiods characterized by widely differing economic conditions, and that it can be explained in terms of a limited set of variables, the nominal interest rate and wealth. Based upon Meltzer's findings, equation 2 can be simplified to become:

$$M/P = Y\ f(R_M,W,u) \tag{3}$$

Finally, by combining equations 1 and 3, we can derive the velocity function as follows:

$$V = PY/M = 1/\ f(R_M,W,u) = V(R_M,W,u) \tag{4}$$

Thus, it is clear that velocity itself is a function of the same observable variables as the demand for money.

As the foregoing makes clear, the stability of the demand for money and hence of velocity is at the heart of the monetarist argument, expressed in the extensive empirical research devoted to this issue. To the extent that velocity may be characterized by trend and cyclical components with different stability characteristics, failure to distinguish between them in the estimation process may account for the conflicting findings of previous research. The decomposition approach presented in the next section is designed to shed light on this important issue.

DECOMPOSITION APPROACH

Our empirical model follows the assumption that the data are I(1) and can be described by a VAR in levels of order *p* as follows:

$$y_t = \prod{}_1 y_{t-1} + \prod{}_2 y_{t-2} + \cdots + \prod{}_p y_{t-p} + \mu_t \,, \qquad (5)$$

where y_t is a vector of *n* difference-stationary variables and μ_t is a vector of Gaussian white noise disturbances. The above VAR can be restated in the first differences of the variables as follows:

$$\Delta y_t = \prod y_{t-1} + \Gamma_1 \Delta y_{t-1} + \Gamma_2 \Delta y_{t-2} + \cdots + \Gamma_{p-1} \Delta y_{t-p+1} + \mu_t \qquad (6)$$

Elements of y_t are cointegrated if they have linear combinations that are stationary. Johansen (1988) has shown that if there are *r* cointegrating vectors, the rank of Π equals $r < n$, such that Π can be factored as the product of two *n* x *r* matrices $(\Pi = -\beta\alpha')$, where α represents the matrix of *r* cointegrating vectors and β represents the matrix of adjustment coefficients. Under the cointegration assumption, equation 6 can be expressed in the following error correction model (VECM) form:

$$\Delta y_t = -\beta Z_{t-1} + \Gamma_1 \Delta y_{t-1} + \Gamma_2 \Delta y_{t-2} + \cdots + \Gamma_{p-1} \Delta y_{t-p+1} + \mu_t \qquad (7)$$

where $Z_t \equiv \alpha' y_t$ represents the error correction terms. It is also seen from equation 7 that the presence of cointegration provides a more parsimonious dynamic representation of the model by reducing the number of parameters to be estimated, thus leading to more efficient estimation results.

The above VECM can be represented even more parsimoniously if, following Engle and Kozicki (1993), we consider the possibility that the first differences of the variables also share serial correlation common features. The serial correlation common features arise if there are linear combinations of the first differences of the variables that are white noise, i.e., that are not characterized by autoregression. Vahid and Engle (1993) show that, for difference stationary variables, if there is a linear combination that eliminates serial correlation in their first differences, this same combination will also eliminate common cycles in their levels. It is thus possible to have an *n* x *s* matrix γ of rank *s*, such that $\gamma' \Delta y_t$ is a vector of white noise processes. The *s* linear combinations within γ are cofeature vectors. Furthermore, as proven by Vahid and Engle (1993), the cofeature vectors are always linearly independent of the cointegrating vectors, so that the total number of cointegrating and cofeature vectors cannot exceed the dimension of the system $(r + s \leq n)$. Since the number of common stochastic trends = *n* – *r,* and the number of common stochastic cycles = *n* – *s*, the above requirement can also be written as $(n - r) + (n - s) \geq n$, that is, the total number of the common stochastic trends and common stochastic cycles cannot be smaller than the dimension of the system.

Estimation of the above VECM in the presence of common trends and cycles requires the implementation of several tests. The first is a unit-root test to determine whether all variables of the system are I(1). In addition, it is necessary to determine *p*, the order of the VAR in levels, which can then be used to determine *r*, the rank of the cointegration matrix, as well as to estimate the cointegrating vectors using the Johansen (1988) method. Given the number of cointegrating vectors, *r*, Vahid and Engle (1993) show how to determine the number of cofeature vectors by searching for linear combinations of Δy_t whose correlation with the elements of the relevant past information, i.e., the lagged values of the first differences of the variables as well as the error correction terms in the VECM, are zero. This can be achieved by computing the canonical correlations between the first differences of the variables and the right-hand side of the VECM. The correlations that are not statistically different from zero represent linear combinations of Δy_t that are independent of the relevant past, and thus represent the cofeature vectors. In addition, the following test statistic can be used to test for the null hypothesis that the dimension of the cofeature space is at least *s*:

$$C(s) = -(T - p - 1)\sum_{i=1}^{s} \log(1 - \lambda_i), \quad (8)$$

where λ_1 ,…, λ_s are the *s* smallest squared canonical correlations between Δy_t and the right-hand side of the VECM, *T* is the number of observations, and *p* is the order of the VAR in levels. Under the null, this statistic has a χ^2 distribution with $s(np + r) - s(n - s)$ degrees of freedom.

Once the numbers of cointegrating and cofeature vectors are determined, it is possible to attempt a permanent-transitory decomposition of the variables in the system. Since our results to be presented later in the paper indicate that $r + s < n$, we need to decompose the data by following the approach outlined in Proietti (1997) and Hecq, Palm, and Urbain (2000). (To conserve space, we omit the mathematical derivations and refer the reader to a more detailed presentation of this approach in Shirvani and Wilbratte, 2007.)

EMPIRICAL RESULTS

The present study is based on OECD quarterly data for 5 major industrial nations, France, Germany, Japan, UK, and US, spanning the period 1978:1 – 1998:4, a sample period roughly coincident with the liberalization of the world financial markets. In addition, the endpoint of 1998 largely eliminates any possible shocks arising from the changeover to the euro as the EMU currency. As noted, the money supply definition employed is the M2 measure, defined as currency, demand deposits, savings, and time deposits, and velocity is expressed in logarithms. Before the decomposition methodology outlined in the preceding section can be implemented for our data, we perform several tests, based on a VAR order of 2, as suggested by the

Akaike (1974) information criterion. The first test is the augmented Dickey-Fuller (1979) test to determine whether the underlying data are difference stationary against the null that they are trend stationary. As Table 1 indicates, at the 5 percent level of significance, all variables are I(1).

TABLE 1
UNIT ROOT TEST RESULTS
(Dickey and Fuller, 1979)

Country	Levels	First Differences
France	-2.91	-5.48*
Germany	-2.24	-7.34*
Japan	-1.71	-5.78*
UK	-3.31	-4.25*
US	-2.52	-4.13*

* Indicates significant at the 5 percent level.

Next, we conduct the Johansen (1988) cointegration test to determine the number of cointegrating vectors. As shown in Table 2, at the 5 percent level of significance, we find only one cointegrating vector, indicating the presence of 4 common stochastic trends among the velocities of the 5 sample countries.

TABLE 2
COINTEGRATION TEST RESULTS
(Johansen, 1988)

Null Hypothesis	Eigenvalue Test	Trace Test
$r = 0$	38.12*	84.09*
$r \leq 1$	19.45	45.97
$r \leq 2$	16.90	26.52
$r \leq 3$	8.37	9.61
$r \leq 4$	1.24	1.24

* Indicates significant at the 5 percent level.

Next, we test whether our data are characterized by common stochastic cycles. A necessary underlying condition is that the data be autoregressive in first differences, as the absence of autoregression allows no scope for common serial correlation. One such autoregression test is the Lagrange multiplier (LM) test, in which the first difference of each variable is regressed on the first differences of all the other variables, including the error correction term, and the resulting TR^2 (T = number of observations, and R^2 = coefficient of multiple determination) has a χ^2 distribution with $n + r - 1$ degrees of freedom. Our test results, in which all the corresponding test statistics are significant at the 5 percent level, strongly reject the null of no autoregression. Having thus shown that our data are autoregressive in first differences, we now test for the presence of serial correlation common features.

Vahid and Engle (1993) propose two different methods of testing for the presence of common feature vectors. The first method is based on the residuals of the regression of the first differences of one of the variables on the first differences of the remaining variables in the system. To deal with the simultaneity problem, we use the two-stage least squares estimation method, using the right-hand-side variables of the error correction model 3 as instruments. The residuals from this regression are then tested to determine whether they are white noise, that is, independent from the instruments, by conducting a standard Lagrange multiplier (LM) test. The LM statistic will have a χ^2 distribution with $n(p-1)+r$ degrees of freedom. We perform a series of regressions, with the first differences of each country's velocities alternately serving as the dependent variable and the other four serving as independent variables. Following this method, we obtain χ^2 values ranging from 0.89 to 1.85, none of which are significant at the 0.05 level, indicating the absence of serial correlation among the estimated residuals. Thus, our two-stage least squares results provide evidence of the presence of common cycles.

The two-stage least squares method outlined above suffers from two limitations. First, there is the normalization problem, meaning the results of the tests for common cycles may be sensitive to the choice of the dependent variable. Second, the method has the limitation that, while it can indicate the presence of cofeature vectors, it does not specify the number of such vectors. To deal with these shortcomings, we use the alternative method proposed by Vahid and Engle (1993), and described in equation 8. The results of this test for the variables in our model are presented in Table 3. As shown in the table, we can reject the null of four or more cofeature vectors, indicating the presence of at most three cofeature vectors, at the 5 percent significance level.

TABLE 3
Canonical Correlation Test Results
(Vahid and Engle, 1993)

Null hypothesis	C(s, j)	Degrees of freedom
$s \geq 1$	0.52	2
$s \geq 2$	2.66	6
$s \geq 3$	13.54	12
$s \geq 4$	58.13*	20
$s \geq 5$	121.20*	30

* Indicates significant at the 5 percent level.

To confirm the presence of three cofeature vectors, we also conduct a likelihood ratio (LR) test of the overidentifying restrictions of the constrained VECM given by 7 above versus the unconstrained VECM given by 6. This test exploits the fact that the difference between the log determinants of the likelihood functions of the two VECM versions has a χ^2 distribution with $s(np+r)-s(n-s)$ degrees of freedom. Under the hypothesis of three cofeature vectors, the related LR test statistic (34.69) is significant at the 5 percent level, indicating that the overidentifying restrictions based on the assumption of 3 cofeature vectors can be rejected. As a result, we also conduct another LR test under the null of only two cofeature vectors. This time, the related LR test statistic (2.85) is not significant at the 5 percent level, indicating that the overidentifying restrictions based on 2 cofeature vectors cannot be rejected. Thus, the rest of this paper proceeds on the maintained hypothesis that there are only two cofeature vectors. The presence of one cointegrating and two cofeature vectors, of course, also implies the presence of four common trends and three common cycles.

Having established the presence of four common trends and three common cycles, we proceed to estimate the permanent and transitory components of the velocities of the sample countries. Since the total number of the common trends and cycles equals 7, which is greater than the 5 variables in the model, we rely on the decomposition method proposed by Proietti (1997) and Hecq, Palm, and Urbain (2000). Using this method, we obtain the permanent and transitory components for each of the velocities of the countries in our sample. The permanent, or trend, components are plotted against the actual velocity levels in Figures 1 through 5, with the vertical differences between each pair of curves representing the corresponding transitory, or cyclical, component. Clearly, in each case, actual velocity follows its corresponding trend closely. However, for each country, there remains a visible cyclical component as well, shown as the vertical distance between the lines of the diagrams. The diagrams reveal clearly that while the trends in velocity follow closely the actual values, as one would expect from our statistical test results, the presence of cyclical deviations from the trends are clearly observable.

With the velocities of the five sample countries decomposed into their trends and cycles, it is possible to explore the possible determinants of each of these components. The following results are presented as evidence of the ability of the trend-cycle decomposition to glean new information regarding velocity from economic time series data. The trend components are hypothesized to respond to such long run “fundamentals” as market interest rates, wealth as measured by the ratio of stock values to national income, technological advances in the payments system, and oil prices as a highly visible indicator of the potential future inflation rate.

Since we employ the broad definition of money (which includes interest-bearing deposits) in this study, interest rate movements represents not variations in the opportunity cost of holding money but the return on the assets within this aggregate and thus is expected to have a positive effect on the demand for M2 and, hence, a negative effect on velocity. Our use of the ratio of stock prices to national income is partially adapted from Friedman (1988), who argues that stock prices may affect the demand for money and thus velocity either positively or negatively, depending upon whether the substitution or wealth effect of stocks dominates. In addition, because we express stock prices relative to income, i.e., to the size of the economy, the ratio can also reflect the degree of financial sophistication in a given economy. We would expect increases in this ratio to reflect enhanced financial sophistication, given that equities are largely outside the provenance of less sophisticated investors, and also to reflect a concomitant improvement in the technology of the payments system, so that increases in this variable should affect velocity positively. In short, an increase in the relative size of the stock market in the economy could potentially have three distinct effects on velocity: a wealth effect (lowering velocity), a substitution effect (raising velocity), and a financial development effect (again raising velocity). Finally, we would expect rising oil prices to lower velocity by signaling an increase in the rate of inflation and, given the slow adjustment of nominal interest rates to higher inflation, to increase the opportunity cost of holding money.

FIGURE 1

PERMANENT COMPONENT OF M2 VELOCITY:

France

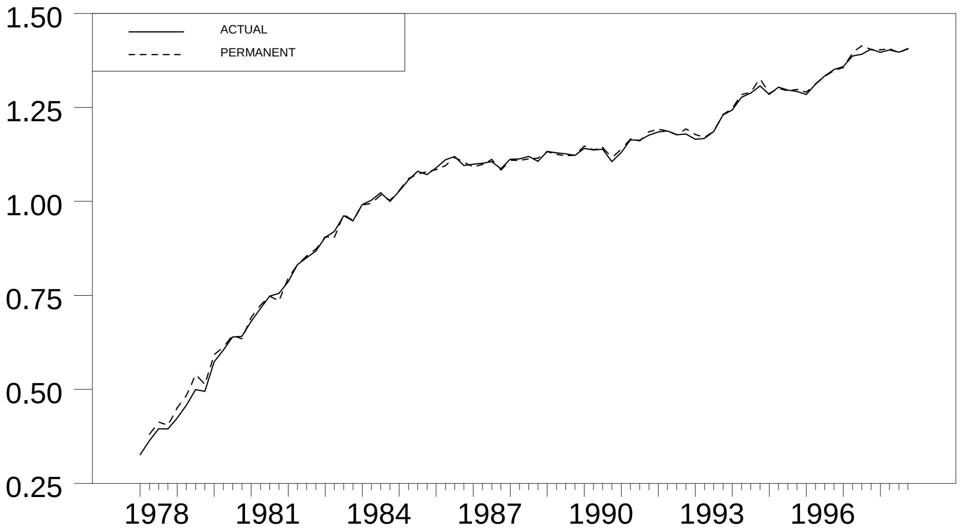


Figure 2
Permanent Component OF M2 Velocity:

Germany

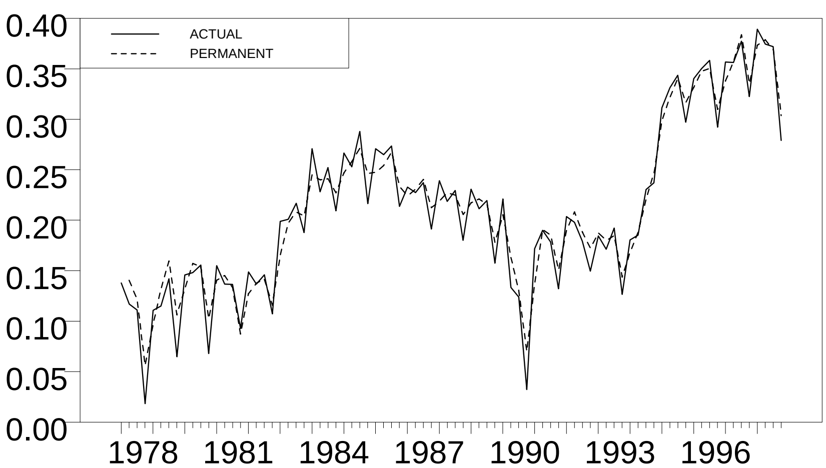


Figure 3
Permanent Component of M2 Velocity:
Japan

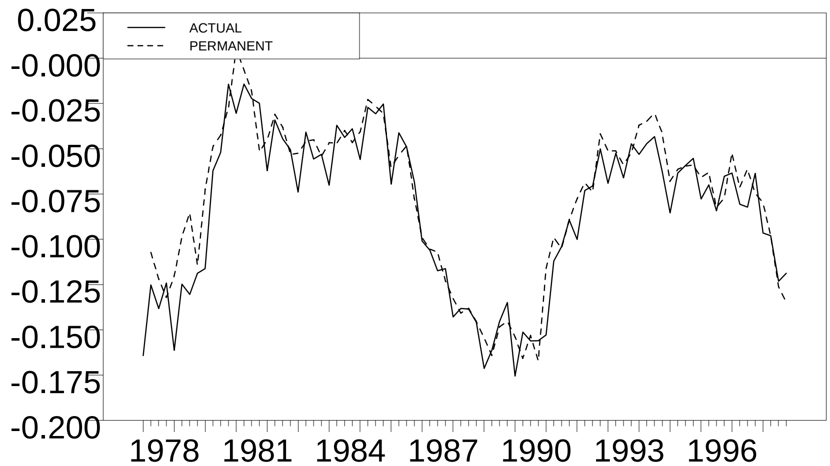


Figure 4
Permanent Component of M2 Velocity
UK

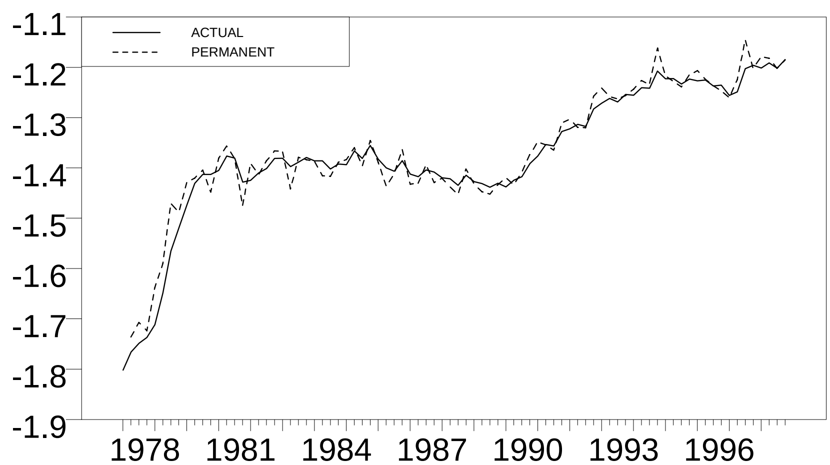
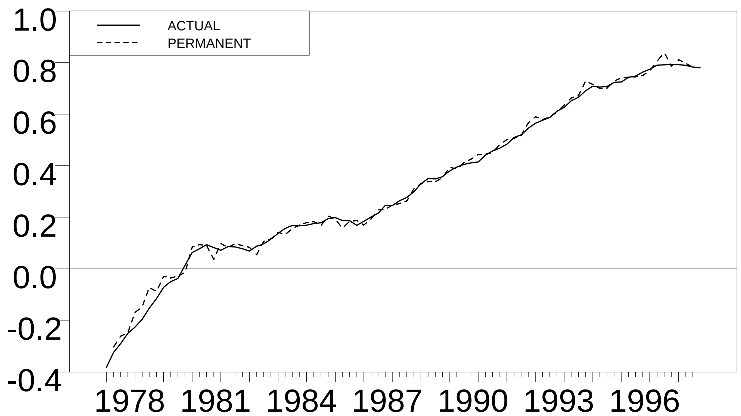


Figure 5
Permanent Component of M2 Velocity

US



The appropriate statistical tests to identify the long run drivers of velocity depend, of course, on the time series properties of the trends and their potential determinants. Since most of these variables are found to be I(1), we cannot employ standard regression analysis, as the relevant confidence tests would be spurious (Granger and Newbold, 1974; Phillips, 1986). Under these conditions, the appropriate estimation technique is the cointegration approach. Stock and Watson (1993) offer a relatively simple dynamic OLS method as a direct way of estimating the cointegrating vectors, as well as testing the significance of the cointegrating coefficients, using the standard distributions.

The dynamic OLS procedure involves regressing the level of the dependent variable on the levels of the explanatory variables (to obtain the cointegrating vector) as well as on past and future changes in the explanatory variables (to adjust for the nonstationarity of the variables in levels). As shown in Table 4, the interest rate coefficient is negative and significant at the 5 percent level in all cases. The sign of the coefficients is as expected, as rising interest rates increase the demand for broadly defined money, and thus lower its velocity, due to the inclusion of interest-bearing deposits in this definition of money. On the other hand, changes in stock prices have both a positive wealth effect and a negative substitution effect on the demand for broad money, with the net effect being positive or negative, depending on which of these two effects dominates. For example, should the positive wealth effect of rising stock prices dominate, then stock prices will positively affect the demand for money and thus negatively impact velocity. The findings reported in Table 4 indeed indicate that this is the case only for Japan, while for the other four countries, the substitution effect dominates. Thus, in most cases, our evidence fails to support Friedman's (1988) expectation that the wealth effect may dominate. Clearly, the booming stock markets in the four non-Asian countries, which lasted through the end of the sample period of this study, attracted funds from bank deposits. In Japan's case, however, the relatively short lived stock market boom (1984-1989) was followed by a prolonged economic and stock market decline from 1990 through the end of the sample period and real estate prices declined as well. This huge destruction of wealth apparently decreased the demand for assets generally in Japan, including broad money. On the other hand, the oil price coefficient is positive, as expected, for every country, and significant for four of the five. Apparently, this highly visible indicator of inflation indeed increases velocity, given the level of interest rates.

TABLE 4
LONG RUN DETERMINANTS OF VELOCITY
BASED ON DYNAMIC OLS
(Stock and Watson, 1988)

Country	Interest Rate	Stock Prices	Oil Prices
France	-3.96*	0.36*	0.14
Germany	-2.61*	0.12*	0.17*
Japan	-1.24*	-0.04*	0.15*
U.K.	-1.52*	0.21*	0.28*
U.S.	-2.50*	0.92*	0.43*

* indicates significant at the 5 percent level.

To explain the short run cyclical fluctuations of the velocities around their long run trends, we perform a number of Granger (1969) causality tests, considering two distinct sets of short-term influences. First, we construct a series of variability coefficients (defined as the standard deviation divided by the mean for a moving window of four quarters) to capture the measured volatility in such psychologically important variables as interest rates, stock markets, and oil prices. In addition, we test for possible contagion effects among the countries in the sample by including in our causality tests the short term cycles of other countries.

TABLE 5
GRANGER CAUSALITY TEST RESULTS
FOR SHORT-RUN DETERMINANTS OF VELOCITY
(Granger, 1969)

F Statistics of Joint Significance				
	Volatility Effects			Contagion Effects
Country	Interest Rates	Stock Prices	Oil Prices	
France	3.32*	16.20*	2.55	48.41*
Germany	4.57*	0.79	6.27*	4.34*
Japan	15.25*	27.24*	13.23*	24.75*
U.K.	20.86*	4.64*	12.89*	14.09*
U.S.	1.13	6.50*	4.86*	5.95*

* indicates significant at the 5 percent level.

The causality test results reported in Table 5 are F tests of joint significance. They reveal that, while velocity in the short run is affected by changes in the fundamentals, as expected from the long-run results, velocity is also affected by the short-run contagion effects emanating from temporary shocks occurring in the other countries of the sample. For example, spillover effects of liquidity crises such as the 2007 problems in the subprime and commercial paper markets will be felt outside the countries in which the shocks originate.

CONCLUSION

The stability of the velocity of broad money plays a crucial role in the monetarist literature. Should velocity be subject to representation as a stable function of a few key fundamental variables, it is possible to derive a predictable relationship between changes in the broad money and nominal income. This paper addresses the issue of the determinants of velocity by distinguishing between long and short run factors. Failure to make this distinction may give rise to apparent instability in the long run behavior of velocity, due to instability of its short run behavior, arising from psychological and international contagion factors. Using recently developed decomposition techniques in a multivariate context, the present paper first breaks down the behavior of velocity for each country in the sample into its long run trend and short run cyclical components. Next, the paper estimates the determinants of each of the two components, using appropriate econometric techniques. Our results indicate that the trend components of velocity are indeed explainable in terms of a limited number of key macroeconomic fundamentals--interest rates, wealth, and oil prices. In addition to these factors, we find that the cyclical movements are influenced by such determinants as the contagion effects of short term economic and financial shocks affecting velocities in countries other than those in which they originate.

These findings thus lend confirmation to empirical findings that velocity is stable and explainable in terms of a number of economically rational determinants in both the short run and the long run. However, the decomposition also reveals that the effects of economic shocks to velocities of money are global in scope through contagion effects. They suggest that policy makers in each country should be vigilant with regard to events in other major economies, as short-term shocks such as the 2007 mortgage lending problems which have affected financial institutions in the US and the UK may also have adverse international contagion effects. While it is accepted that domestic monetary policy cannot be conducted in a vacuum, this paper clarifies that a key component of the international transmission mechanism is the velocity of broad money. In addition, given the finding that these contagion effects are short-term in nature, central banks would do well to react quickly to global events through closely concerted policy coordination.

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