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INSTITUTIONAL CHANGE AND THE VELOCITY OF MONEY: A CENTURY OF EVIDENCE

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We study common features in the income velocity of money, income, and interest rates for Canada, the U.S., the U.K., Sweden and Norway using annual data from 1870. The recently developed and refined techniques of testing for cointegration are employed.

The evidence suggests there is a unique long-run relationship in velocity but not *in income and interest rates. Moreover, we find that only a model which includes institutional change proxies is properly specified. We argue that the evidence is best interpreted in the context of common historical developments in the respective countries' financial systems.* (*JEL* E41, N2, O11)

I. INTRODUCTION

The study of the long-run behavior of velocity has intrigued many researchers who have sought to link it to the evolution of financial systems over time. Indeed, the approach taken by Bordo and Jonung [1987] (BJ) explains the long-run portion of velocity in five countries by institutional factors including monetization and financial development, while the seminal study by Friedman and Schwartz [1982] (FS) argues that financial sophistication is an important determinant of the long-run behavior of velocity in the U.S. and the U.K. The aim of this study is to further explore the connection between long-run velocity movements across several countries, as well as the relationship between countries of its principal institutional and economic determinants.

This line of research is important for a number of reasons. It allows us to demonstrate that the demand for real balances cannot be adequately expressed by a few aggregates alone and that institutional variables need be included. If technological changes in the financial system are found to influence the demand for money, this has implications for the question of whether the demand for real balances is likely to be stable over time. This also impinges on the links which are thought to exist between monetary aggregates and economic activity. Finally, as Boughton [1992] argues, international comparisons of the demand for money reveal not only that institutional factors represent an important determinant of velocity, but that these appear to differ in the short-run across countries. Given the increasingly global nature of financial markets, it is of interest to explore whether the common development of long-term institutional changes among selected countries represents only a postwar phenomenon.

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ABBREVIATIONS
BJ: Bordo and Jonung
FS: Friedman and Schwartz
EG: Engle and Granger

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BJ [1981; 1987] suggest that institutional changes explain much of the long-run behavior of velocity. Siklos [1993], following upon BJ, confirms that in order to generate a long-run statistical model of velocity, a conventional model of velocity (as a function of real income and the nominal interest rate) needs to be augmented with institutional change proxies. Many economists now contend that institutional change represents an important element in explaining the long-run behavior of velocity or the demand for money (e.g., Laidler [1993]).[1]

This study examines the long- and short-run relationship of velocity for a sample of five industrialized countries using annual data beginning in 1870. Since the "long run" in economics need not be the same for all problems, an important issue is the selection of the sampling frequency of the data (e.g., Perron [1991]). In particular, the effects of technological or institutional changes in the financial sector occur slowly, necessitating as long a sample as possible.

Given recent findings by BJ [1990] and Siklos [1993], which empirically demonstrate that institutional change is common to each country, it would seem natural to ask whether there are common features in financial changes across countries. To investigate such a possibility, we perform a variety of tests to determine if velocity, and each of its individual determinants are, separately, cointegrated across countries. We also examine the short-run dynamics and the stability of any unique cointegrating relationship which is detected using some recently developed statistical tests. Finally, we attempt to estimate a joint velocity function by pooling data for all the countries in our sample.[2] In so doing, we improve on the earlier studies of long-run common movements in velocity and its conventional and institutional determinants, presented by BJ [1987, ch. 4] and FS [1982, ch. 7].

In performing time-series tests our objectives are three-fold. First, we wish to explore whether the common features of financial systems across countries are as significant as FS [1982, ch. 7] found them to be for the U.S. and the U.K., based on more sophisticated measures of correlation. We also expand the selection of countries to include Canada and Sweden.[3] Second, an analysis of the common features of institutional change across countries could shed some light on the speed with which technological changes are transmitted across countries. That is, do countries at similar stages of development in effect import payments technologies from other countries? A third objective is to ascertain whether certain historical features, which would presumably have had an impact on financial development, can be detected in the data. It is here that structural stability tests serve a useful purpose. Briefly, the results lend support to the view that there exists a single cross-country long-run relationship in velocity, but not in income and interest rates. We also find that only a model which includes institutional change proxies is both properly specified in the pooled time-series case and produces long-run elasticities consistent with theoretical predictions. In general, we argue that the statistical evidence can only be understood in the context of common historical developments in the respective countries' financial systems. Another important conclusion is that studies which suggest that international linkages between financial developments are essentially a post-World War II phenomenon are misleading. The common development of institutional changes in the financial system is not a recent occurrence, for we find signs of such behavior dating back to at least the beginning of this century.

After a brief review of the institutionalist hypothesis of the long-run behavior of velocity (section II), and a description of econometric issues (section III), empirical evidence confirming the above conclusions are presented (section IV). The paper concludes with a summary in section V.

1. An exception is Rasche [1987] who does not find institutional change to be important. However, his testing procedure is univariate in nature, not multivariate as is the case in the present paper.

2. BJ [1987, 48] pool their data to show that the influence of institutional change variables on velocity is similar in all the countries examined, suggesting that common forces underlying the institutionalist proxies explain the common behavior of velocity. But their study confounds short-run and long-run influences since they could not rely on recent advances in time-series analysis.

3. Norway is included in subsample estimation but could not be included in full sample estimates because of gaps in the data.

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II. THE INSTITUTIONALIST APPROACH: A REVIEW

Since much has been written about the institutionalist explanation of the long-run behavior of velocity advanced by BJ [1981] (see also BJ [1987; 1990], Siklos [1993], Ireland [1991], Laidler [1993], and Hallman, Porter and Small [1991]), we present only a very brief review here.

Velocity is traditionally viewed as an analogue of the demand for real money balances. Consequently, it is treated as a function of income (or permanent income) and an interest rate. The latter variable serves as a proxy for the opportunity cost of holding money.[4] BJ suggest that, in addition to its traditional determinants, velocity is a function of institutional changes in the financial system. These institutional developments proceed in roughly two phases. First, most economies experience a monetization phase. During this period money is used more intensively to settle transactions. At the same time, the speed with which the banking system spreads throughout the economy produces rapid growth in the use of currency and deposits. A second stage is characterized by growing financial sophistication during which the number of substitutes for bank notes and deposits grows. The combination of these two factors produces a U-shaped long-run pattern in velocity for the countries considered in this paper. As shown in Figure 1, velocity demonstrates a falling trend that ends during the interwar period and a rising trend starting for most countries in our sample in the mid-1940s. The downward trend in velocity before World War II is attributed to the process of monetization, and the upward trend since to two developments:[5] increasing financial sophistication and improved economic stability.[6]

The striking similarity in the behavior of velocity across industrialized countries suggests the possibility of common financial developments in different countries despite differences in fiscal and monetary policies, in their inflationary experiences, and industrial development. Alternatively, the shared economic features might be due to similar experiences in income or interest rate patterns. For example, existing economic and historical evidence suggests that while there are several common features in macroeconomic aggregates such as GNP and consumption across countries (e.g., Backus and Kehoe [1992]), none of the studies, to our knowledge, has applied tests of cointegration either to determine whether financial change is common across countries or to attempt to isolate the sources of common movements if they exist.[7]

Many authors have applied tests of cointegration to determine whether the traditional determinants of the demand for money, namely income and interest rates, are cointegrated with some measure of the money stock. Miller [1991] and Hafer and Jansen [1991] represent only a partial list of recent contributions in this area. Existing empirical evidence for U.S. data suggests that a broad monetary aggregate (usually M2), income and nominal interest rates are cointegrated over a variety of samples, at least for U.S. data.

By contrast, empirical evidence is decidedly mixed for models which use an M1 definition of money. Baba, Hendry, and Starr

4. Specifications which examine the determinants of real balances have been preferred in part because of the finding that velocity behaves like a random walk. Nevertheless, given the difficulties surrounding tests of the random walk hypothesis (see Campbell and Perron [1991]), and the presence of statistical breaks in the random walk behavior of velocity (e.g., Perron [1989]), the empirical evidence suggests, on balance, that the random walk hypothesis is not a substitute for a complete model of velocity's behavior.

5. There have been some interruptions in the rising trend, such as during the 1980s in the U.S. when velocity in M1 levels began to level off and even decline. Since these unexpected changes have, in hindsight, been ascribed to the slow pace of regulatory reform in the face of financial innovations, there is still greater reason to consider the possibility of a relationship between velocity and institutional change.

6. Whether the postwar period produced more stable variation in economic aggregates such as GNP or unemployment, particularly in the U.S., has been the subject of a debate which remains unsettled.

7. We would have liked to expand the data set to consider other countries, as in Backus and Kehoe [1992], who kindly made their data available to us. We are unable to do so for three reasons. First, the power of the tests which are applied below falls with the number of variables in the model. Second, we are unable to produce estimates of institutional change for countries other than the five considered in this paper. Third, the countries examined are, for historical and economic reasons, relatively good candidates for common institutional and economic development. Thus, for example, the gold standard as well as trading relationships linked the U.S., the U.K., and Canadian economies, and the same is true of the links between Sweden and Norway, on the one hand, and the U.K., Sweden, and Norway on the other, by virtue of exchange rate regimes and geographical proximity.

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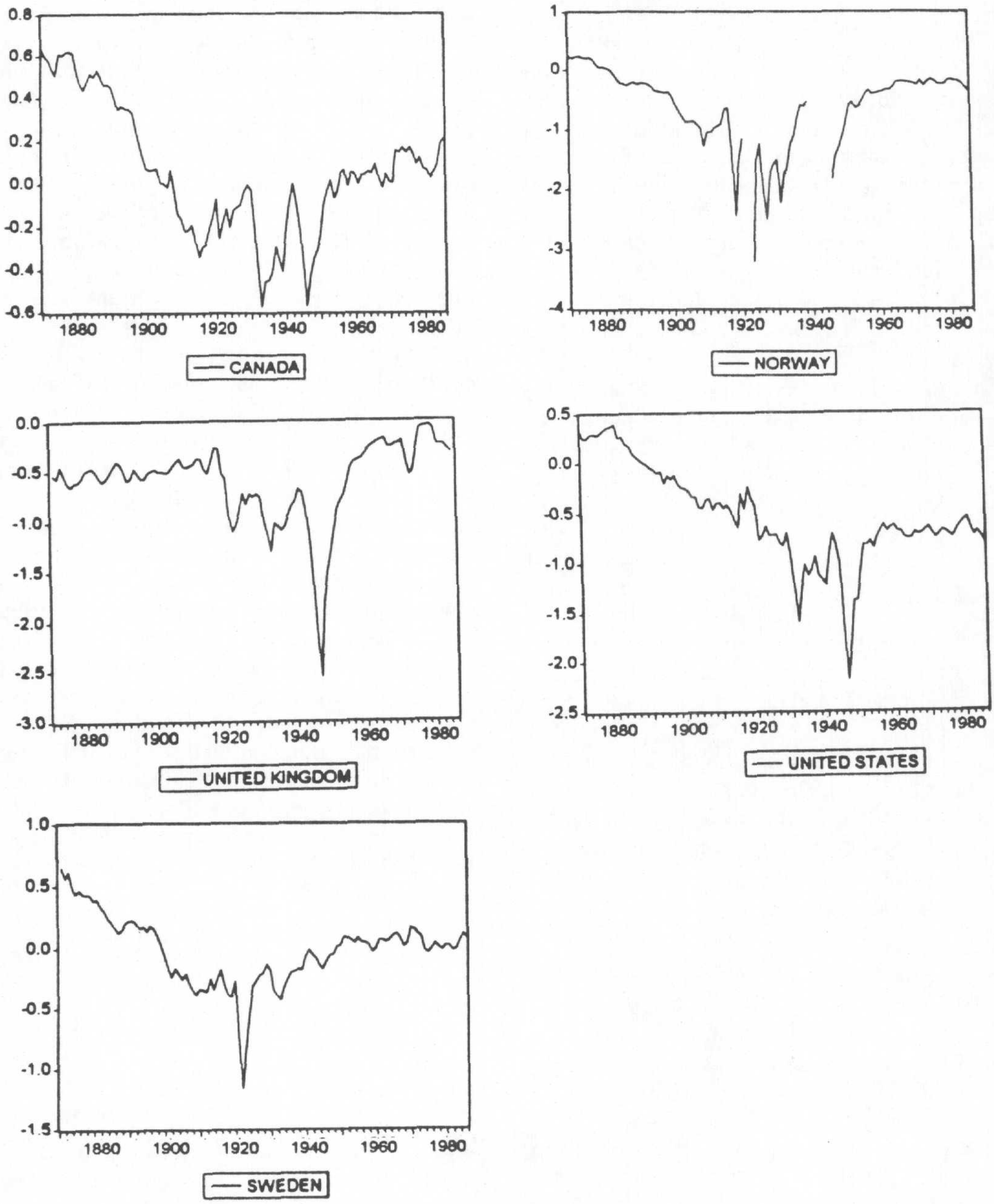
FIGURE 1
Income Velocity (Log) Levels in Five Countries[a]



[a]Generally measured as the logarithm of GNP (or a proxy) divided by M2 (or its equivalent). See BJ [1987] and Siklos [1993] for greater details and data sources. Gaps for Norway reflect missing observations.

[1992] augment their long-run M1 model to incorporate yield spreads and risk features of interest rate behavior. Miller [1991] finds that M1 is an unreliable variable for understanding short-run money demand behavior. Hafer and Jansen [1991] prefer M2 over M1 for U.S. annual data since 1915, in the sense of finding cointegration among the variables in a conventional money demand relation. Because the M2 definition incorporates the influence

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over time of financial innovations, this may explain a potential source of money demand instability in M1-based models (see also Baba, Hendry, and Starr [1992]).

Recent theoretical work has also attempted to model the role of technological changes in the financial sector. Ireland [1994] incorporates two of the features which are central in the empirical work to follow, namely monetization and growing financial sophistication, in a general equilibrium model which is capable of reproducing empirical facts about the long-run behavior of velocity.

III. ECONOMETRIC SPECIFICATION

Specifications Tested

Cointegration describes the relationship between two or more time series which appear to share a common trend as a statistical description of the long run in economics. A large literature has refined and improved the original single-equation testing procedure presented in Engle and Granger [1987] (EG). Tests for cointegration have also become popular as one way, albeit not the only one, of conducting inference on time series which are non-stationary and drift over time.

The fundamental approach of this paper may be described as follows. Let velocity, v_t, be determined by its traditional determinants, denoted by the vector Φ, and its institutional proxies, denoted by the vector Ω. The institutionalist hypothesis may then be written

$$(1) \qquad v_{it} = \beta_0' + \beta_1' \Phi_{it} + \beta_2' \Omega_{it} + \varepsilon_t.$$

The vector $\Phi_t = [y^p_t, R_t]$ where y^p is real per capita permanent income, and R is a proxy for the opportunity cost of holding money. The vector of institutional proxies Ω=[$(NBFA/FA)_t$, $(C/M)_t$, $lnal_t$], where *NBFA/FA* is the ratio of total non-bank financial assets to total financial assets, (*C/M*) is the currency-money ratio, and *lnal* is the share of the labor force in non-agricultural pursuits. The index i identifies a particular country and t is time.

The theoretical rationale for the vector Φ is well known (see Goldfeld and Sichel [1990] for a survey). The motivation behind the vector Ω can be stated briefly as follows (see also BJ [1987]). Since *NBFA/FA* proxies financial development, its increase would be expected to reduce the demand for money by increasing the number of close substitutes and thereby raising velocity. The *C/M* series mirrors the spread of commercial banking since, as the banking habit spreads, the use of deposits increases relative to the use of currency (i.e., *C/M* falls). This increases the desired holdings of real balances and reduces velocity. Thus, changes in velocity are positively related to changes in *C/M*. The steady rise in the proportion of the labor force in non-agricultural pursuits reflects the "spread of the monetary economy" (BJ [1987, 34]) and growing urbanization. Other things being equal, BJ predict that, as this series rises over time, velocity falls.[8]

Previous research (BJ [1981; 1987; 1990] and Siklos [1993]) has concentrated on estimates of equation (1) for individual countries. The present application pools data for up to five countries and estimates a "North Atlantic Velocity Function" to determine whether a common velocity function can be identified. If so, institutional change may be a common feature for at least the countries considered in our sample. In addition, because previous studies have suggested that the Ω vector, in particular, significantly explains velocity in each of the countries considered, this would suggest that if a velocity function is common to all countries, it may be due to common features in Ω or Φ, or both. Thus, one objective of the present empirical analysis is to examine whether the following linear combinations are also stationary, that is, whether they are cointegrated.

$$(2) \qquad v_{it} + \delta_0' v_{jt} = \varepsilon_v,$$

$$(3) \qquad \Phi_{it} + \delta_1' \Phi_{jt} = \varepsilon_\Phi,$$

$$(4) \qquad \Omega_{it} + \delta_2' \Omega_{jt} = \varepsilon_\Omega\,, i \neq j,$$

where v is a vector to indicate that a cointegrating relationship for velocity exists between a time series for countries i and j, where j can represent values for one or several countries and the residuals ε are stationary. For example, suppose we have a sample con-

8. Omitted from equation (1) is BJs measure of economic stability, a six-year moving standard deviation of the annual percent change in real per capita income. Using a moving standard deviation measure of volatility is problematical in econometric estimation.

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sisting of data from two countries. The finding that a relationship with cointegrating vector [1 –1] is stationary would imply that velocity in country i is cointegrated with velocity in country j, thereby establishing a long-run statistical relationship between velocity in the two countries. The same arguments can be extended to the case where three or more countries are considered and to the relationship between the variables in equations (2) to (4).

Next, suppose that Φ_i and Φ_j are not cointegrated, while one is able to reject the absence of cointegration between Ω_i and Ω_j, and that $\beta_1' = 0$ in equation (1). In that case the explanation for the common movement in v_i and v_j would be explained by common movements in elements of Ω_i and Ω_j. Since the latter vector proxies financial development, this implies that financial development is common in one or more of the countries sampled.[9] Alternatively, of course, it may be that the cointegrating relationship in velocity is explained by common trends in income and interest rates (e.g., as in when $\beta_2' = 0$ in equation (1)). Clearly, some combination of the two extremes is also possible, although we intend to empirically demonstrate that, at least in the long run, one can easily reject the hypothesis that $\beta_2' = 0$ in terms of equation (1), and that this is due to the properties of equation (4). It is the omission of the vector Ω in much of the previous work in this area that we wish to draw attention to. The approach outlined in equations (2) to (4) also begs the question whether any long-run relationship is stable and whether we can identify the transmission of institutional factors from one country to another. Gregory and Hansen [1996] develop tests of the stability of cointegrating relationships. They find a break in a conventional U.S. money demand function during the early 1940s, but the evidence against stability is not strong.

Testing for Unit Roots and Cointegration

There exists a large literature examining whether economic time series are stationary around a deterministic trend, or are the sum of a permanent component best described as a random walk (perhaps with a drift) and a transitory component. We assume, based on existing empirical evidence, that each of the series in equation (1) possesses a unit root. This contention is based on results of several available unit root tests. Test results using the present data set have been presented in Siklos [1993].[10] Testing for unit roots is often viewed as a first step in determining whether two or more series are cointegrated.[11]

The approach used here to study common features in time series is based on the work of Johansen [1991], and Hansen and Johansen [1992]. Since several expositions of the technique are now available in the literature (e.g., Johansen [1991], Johansen and Juselius [1990], Hafer and Jansen [1991]), we refer readers to these sources for the details about the so-called Johansen procedure for tests of cointegration. Johansen's procedure also enables the investigator to perform a variety of tests of various restrictions imposed on a model.

Nevertheless, there are a number of issues which need to be considered when applying Johansen's procedure. First, is the selection of the lag length in the VAR. Lag lengths in this study were selected on the basis of Akaike's Information Criterion (AIC), primarily because it tends to select relatively long lags thereby reducing the chances of certain types of specification errors.

Two test statistics can be used to evaluate the number of cointegrating relationships, namely the trace test and the maximal eigenvalue test. For long samples, such as the one considered here, the two tests generally yield the same conclusions. Results based only on the trace test are reported below while other test results are available upon request.[12]

9. One of the criticisms of BJ is that the elements of those vectors are not independent enough of each other in principle. Empirically, however, the problem does not appear to be a very serious one.

10. Questions have been raised about whether unit root findings may be biased in the presence of a structural break in the data. An appendix (available upon request) presents results based on one recently developed test (see Siklos [1993]) which generally confirms the existence of a unit root in the series considered here despite the possibility of a break in the series.

11. But, as in Johansen and Juselius [1990], while such tests are useful guides to the possibility of finding a cointegrating relationship, they are not sufficient tests for cointegration.

12. Critical values depend on whether the VAR contains a constant, a constant vector, and a constant vector restricted to lie in the cointegration space. These are Tables A1, A2, and A3, respectively, in Johansen and Juselius [1990]. Osterwald-Lenum [1992] has produced improved estimates of Johansen's critical values. These are used in the empirical work which follows.

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The Stability of Cointegrating Relationships

While wars or the Great Depression may not have influenced the long-run common pattern in velocity across the countries considered, it is nevertheless possible that these events may have interrupted the relationship which exists between the time series. BJ [1987, ch. 4] separately examined periods of falling and rising velocity and found few differences across countries in the latter period which largely coincides with the post-World War II era. They did not, however, rely on a statistical test to determine whether their chosen break-point is appropriate.[13] FS [1982, ch. 7] adjust their estimates of the relationship between secular movements in velocity between the U.S. and the U.K. by including dummy variables for wars and the Depression. Similarly, they document the fact that while velocity movements in the U.S. and the U.K. "reflect a unified financial system" (FS [1982, 337]), some differences exist during the pre-1914 period. This is apparent from Figure 1 since it suggests that velocity levels were falling in all of the countries considered, except the U.K. which exhibited only a slight overall decline in the period 1870–1914.

Several responses are available to address these issues. One is to test for structural breaks at particular known dates. However, unless we catalog all of the events which can impinge on the financial relationship between countries, we cannot be certain that the most significant structural break has been accounted for. For this reason it is preferable to rely on tests for which the date of the structural break is unknown. Therefore, we implement recently developed tests for stability in cointegrated relationships where the timing of such a break is unknown.[14] Alternatively, the stability of any cointegrating relationship can also be explored by estimating the relevant relationships recursively. In each recursion a new observation is added and the model is reestimated. Previously used observations are not discarded. This approach enables us to examine the evolution of any postulated relationship over time and is thus not subject to the criticism of ad hoc sample selection.[15] We can also explore the stability of long-run relationships by generating rolling estimates where a fixed proportion of the sample is analyzed in a sequential fashion. This procedure has the advantage of giving each additional observation in the rolling regressions equal weight as opposed to the declining weights inherent in the recursive approach. Either approach seems preferable to estimation over selected subsamples though such testing was also conducted (results not shown).[16]

13. Indeed, the coefficients in their velocity model incorporating institutional change factors show signs of a structural break in only one of the five countries considered, namely Canada (BJ [1987, Table A.2]). This could be a sign either that the importance of institutional factors permeates the entire sample or that the break points were inappropriately chosen.

14. An alternative to tests for breaks in certain years or over time is to select some suitably long subsample. The only sample sufficiently long to conduct cointegration tests is the Gold Standard period 1870–1914. We also tested for cointegration conditional on the presence of shocks arising from the two world wars, the Great Depression, and the two oil price shocks, where these are assumed to be exogenous. Our conclusions are unaffected.

15. In the context of cointegrated relationships, however, there are additional considerations to keep in mind. One can constrain the short-run component of a model and estimate recursively long-run parameters or one can, as in most studies, fix the long-run and estimate the short-run recursively. Hansen and Johansen [1992] argue that the first two are most useful for the analysis of structural breaks in long-run relationships as they do not rely on the identification of the individual cointegrating vectors. In this paper, we follow the former approach although it is to be noted that inference can be different under the two approaches. Again, see Hansen and Johansen [1992] for the details.

16. Gregory and Hansen [1996] propose new tests of stability in the context of cointegrated relationships. Their test posits that the null is the standard cointegration equation. Thus, for two series y_{1t} and y_{2t}, the standard cointegrating regression is written

$$y_{1t} = \mu + \theta y_{2t} + e_t.$$

One alternative hypothesis is written

$$y_{1t} = \mu_1 + \mu_2 \varphi_{t\tau} + \theta_1 y_{2t} + \theta_2 y_{2t} \varphi_{t\tau} + e_t.$$

This last equation is augmented with a change in intercept (μ_2) and a change in the slope. We define a dummy variable φ as

$$\varphi = \begin{cases} 0, \text{ if } t \leq [n, \tau] \\ 1, \text{ if } t > [n, \tau] \end{cases}$$

where n is the number of observations and where φ is created for each possible break point τ. The usual notation is $[n\ \tau]$ where τ is defined in the interval $[.15n, .85n]$. Some trimming of the sample is required because the test statistic is not, strictly speaking, defined over all of n. The sequence of residuals can then be analyzed in the same manner as the test for cointegration proposed by EG [1987], that is, by generating an augmented Dickey-Fuller (ADF) statistic for each τ.

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TABLE I
Cointegration Test Statistics

(A) Canada – U.S. – U.K. – Sweden

Series	Number of Cointegrating Vectors				Lag Length	Sample[a]
	0	1	2	3		
v	46.74*	18.09	8.32	2.64	2	1870–1985
y^P	63.62*	23.77	6.28	.11	5	1900–1985
R	80.00*	41.78*	12.76	.99	3	1870–1985
C/M	46.61*	20.50	8.80	2.89	5	1871–1985
lnal	67.71*	39.49*	19.05*	7.87*	5	1900–1985
NBFA/FA	48.53*	25.27	12.51	2.17	5	1880–1985

(B) Canada – U.S. – U.K. – Sweden – Norway

Series	Number of Cointegrating Vectors					Lag Length	Sample[b]
	0	1	2	3	4		
v	88.03*	36.57	18.06	5.53	1.03	3	1870–1985
y^P	108.59*	70.11*	34.01*	2.83	2.83	5	1875–1985
R	99.67*	64.13*	29.66*	2.35	2.35	5	1870–1985
C/M	76.01*	47.12*	22.53	7.03	3.26	5	1871–1985
lnal	102.83*	65.15*	41/44*	22.91*	10.10*	5	1900–1985
NBFA/FA	100.88*	53.42*	27.51	10.71	2.29	5	1880–1985

Notes:

*signifies rejection of the null that $r = i$ vs $r \leq j$ $u \neq j$, where r is the number of cointegrating vectors, at the 10% level of significance (trace test). Critical values are from Osterwald-Lenum [1992] who recalculated the values in Johansen and Juselius [1990]. The tests assume that the series are trended variables with a trend in the GDP.

[a]Before lags are taken into account.

[b]Same as above except data for 1921–1922, and 1940–1945, for velocity and real per capita permanent income, were excluded because data were unavailable for Norway. An intercept dummy was used to splice Norwegian velocity data, which are non-existent for the years 1940–1945 inclusive, because of a substantial shift in velocity levels between 1938 and 1946. Shift dummies were not found to be necessary for the other series.

IV. DATA AND EMPIRICAL RESULTS

Data

The annual data used in this study are updated from BJ [1987]. The sample begins in 1870 and ends in 1986. Given the difficulty of updating some of the institutional change proxies (particularly the *NBFA/FA* series) the data could not be readily extended beyond 1986. Five countries are considered in the empirical results reported below. They are the U.S., U.K., Canada, Sweden, and Norway.

Testing for Cointegration. Table I presents tests of cointegration based on the Johansen methodology for equations (2) to (4) separately for the whole sample. Panel (A) of the table tests for cointegration using data for the four countries in which the series are available for the full sample. Panel (B) of Table I adds Norway to the list of countries considered but omits the years 1921–1922 and 1940–1945 because data for Norway were non-existent. For velocity, the results are the same in both cases. We find that one cannot reject the null that a unique cointegrating relationship exists between velocity for Canada, the U.S., the U.K., and Sweden. To the extent that velocity reflects income, interest rate and institutional changes, the results reflect the statistical confirmation that velocity levels in these countries are attracted to each other in a statistical sense. These results would also be the analogue of the Backus and Kehoe [1992] findings of striking similarities in international business cycles.[17]

17. An alternative way of stating the above result is to say that with four countries a single cointegrating vector implies three common stochastic trends. Thus, several factors may be driving velocity levels in the countries considered.

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The remaining cointegration test results in Table I seek to determine whether, separately, other determinants of velocity are cointegrated. Our findings may be summarized as follows. One cannot reject the null of a single cointegrating vector between y^p for the four countries in our data set. The results differ, however, when the truncated sample is considered (panel B). There we find that at least three cointegrating vectors exist for permanent income. Thus, if we control for the war years, there is evidence of possibly one common stochastic trend in income but not of a unique equilibrium relationship for all the countries considered. This means that any one country's income is representative of income-level movements of all the countries considered here. In general, however, since the number of cointegrating vectors for all series except velocity and *lnal*, rises when Norway is added, this would suggest that Norway belongs in the long-run analysis considered.

There is also no evidence of a single cointegrating vector in interest rates for either case considered. Instead, one cannot reject the null of two or three cointegrating vectors (i.e., two common trends) between interest rates depending on whether Norway and the war years are excluded. Given the findings of different numbers of cointegrating vectors in Table I, it may be of interest to test whether data from some countries can be excluded altogether from the long-run relationships considered so far. This is accomplished by testing the significance of the long-run coefficients in the cointegrating relationships, as well as testing the hypothesis of weak exogeneity in the error correction representation of the estimated models (see below). Results (not shown) indicate that one cannot exclude any country's velocity in Table I (panel A or B) from the cointegrating vector. Thus, (log) velocity levels in all five countries are related to each other in the long-run.[18] This result was confirmed not only by testing whether an individual country's velocity can be excluded but also via tests to determine whether separate cointegrating relationships could be found over the whole sample for Canada and the U.S., the U.S. and the U.K., or the U.K. and Sweden.[19]

Turning to real per capita income, we found that a unique cointegrating relationship among income levels could be identified by excluding Canada while, for the long-term interest rate, both Canada and Sweden can be omitted from a long-run relationship. Further tests indicate that separate long-run relationships between the U.S., the U.K. and Sweden appear to exist for the interest rate. Therefore, the findings for permanent income and the interest rate imply that permanent income may be a relatively more important determinant of the long-run behavior of velocity, as Siklos [1993] suggested. BJ [1981; 1987] suggest that interest rates might be a relatively more important variable in explaining the long-run behavior of velocity than permanent income. The cointegration test results for the institutional proxies, at least in panel A of Table I, suggest that a single cointegrating vector is found for (*C/M*) as well as the financial sophistication proxy (*NBFA/FA*). Thus, there appear to be long-run common features in institutional change. For the labor force variable, there does not appear to be a single common stochastic trend as the null of four cointegrating vectors is rejected by the trace test. When Norway is included, one cannot reject the null of two vectors. Exclusion tests performed on the currency-money ratio proxy for institutional change reveal that one cannot exclude any of the countries, thus giving rise to a single cointegrating vector with the four countries considered. For the *NBFA/FA* proxy for financial development one can omit Canada so that a unique cointegrating relation with the U.S., U.K., and Sweden adequately describes the long run for this series.

The Transmission of Institutional Change

The results in the previous section suggest that the dynamic relationship among many of

18. As a referee correctly points out, the mere fact that we cannot exclude any of the countries considered in the long-run relation does not imply that, had we included other countries not considered here, these would also be excluded. Indeed, there are good reasons to believe, based on previous work (e.g., BJ [1987]), that the appropriate vector of countries one might want to consider should be larger. Footnote 7 describes the data limitations we faced. Our objective here is simply to confirm that the group of countries considered here can be viewed as a single entity in a narrow sense.

19. There is a chance, however, that a separate long-run relationship exists between U.S. and U.K. velocity, but no such separate cointegrating relationship could be found between either the U.S. and Canada or the U.K. and Sweden.

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TABLE II
Error Correction Models

ΔV_{UK}	=	.067	$+.627\Delta V_{US}(-1)$			
		(.03)	(.108)*			
		$+.25\Delta V_{UK}(-1)$				$-.064z(-1)$
		(.114)*				(.031)*
Summary Statistics: $R^2 = .453$, $F(9,102) = 9.38$*, $SC(1) = .98$						
ΔV_C	=	.032		$-.182\ \Delta V_{US}(-2)$		
		(.021)		(.074)*		
						$.034z(-1)$
						(.019)*
Summary Statistics: $R^2 = .206$, $F(9,103) = 2.907$*, $SC(1) = .793$						
ΔV_S	=	$-.084$				$-.351\ \Delta V_C(-1)$
		(.025)*				(.140)*
				$+.303\Delta V_S(-1)$	$-.207\Delta V_S(-2)$	$-.076z(-1)$
				(.092)*	(.094)*	(.023)*
Summary Statistics: $R^2 = .279$, $F(9,103) = 4.428$, $SC(1) = .0003$						
ΔV_{US}	=	$-.0001$		$-.38\Delta V_{US}(-2)$		
		(.04)		(.15)*		
						$.01z(-1)$
						(.04)
Summary Statistics: $R^2 = .13$, $F(9,103) = 1.68$, $SC(1) = 6.56$*						

Note: To conserve space only the statistically significant coefficients, the constant and the error correction terms are shown. Standard errors in parenthesis. * signifies rejection of the null at the 10% level of significance. Δ is the difference operator, R^2 is the coefficient of multiple determination, F is the test for the joint statistical significance of the regressors (degrees of freedom in parenthesis), and SC is the test of first-order serial correlation in the residuals. The sample is the same as in Table IA before differencing and lags.

the series considered can be modeled via a vector autoregression augmented by error correction terms. Vector error correction models (VECM) are useful as a further test of the cointegration hypothesis, as a device to determine the size of the error or deviation in an equilibrium relationship, and to determine which variable in a system is weakly exogenous relative to other variables in the model. To illustrate, the results of the estimation of *VECM*s for velocity are provided in Table II. The error correction terms, z, are statistically significant and of the correct sign in all of the regressions except in the U.S. equation, where the error correction term is statistically insignificant. Hence, these results suggest that U.S. velocity is weakly exogenous of velocity in the other countries. Together with the results considered in Table I, there is confirmation of a unique cointegrating relation in velocity among the four countries considered. The size of the error correction terms is small, suggesting that adjustment to equilibrium is slow, in the order of approximately 7% to 8% per year in the U.K. and Swedish cases for example. Such an outcome is not surprising since institutional change is believed to take place slowly and has long-lasting effects on the demand for money, as discussed in BJ [1987, ch. 3]. Further insights may be gained from Figure 2 which plots estimates based on a rolling regression with a fixed sample size of 40 years, along with the standard error bands, of the error correction terms for the U.K. and Sweden equations in Table II.[20] The top panel of Figure 2 reveals that the size of the error

20. It is unclear, a priori, how wide the window or fixed portion of the sample should be. The size of the error correction term would suggest a minimum of at least 20 years. Windows of 20, 40 and 60 years were considered.

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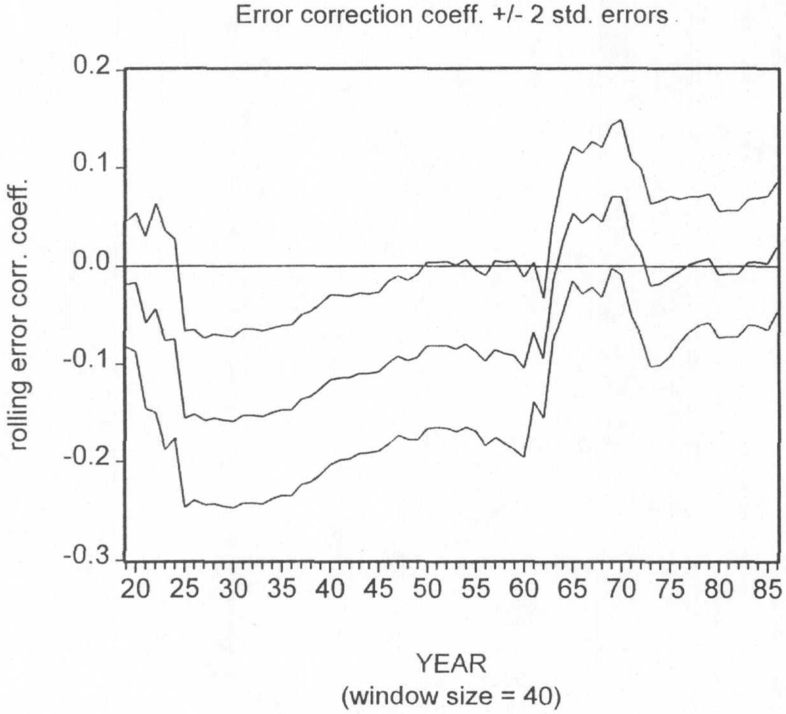
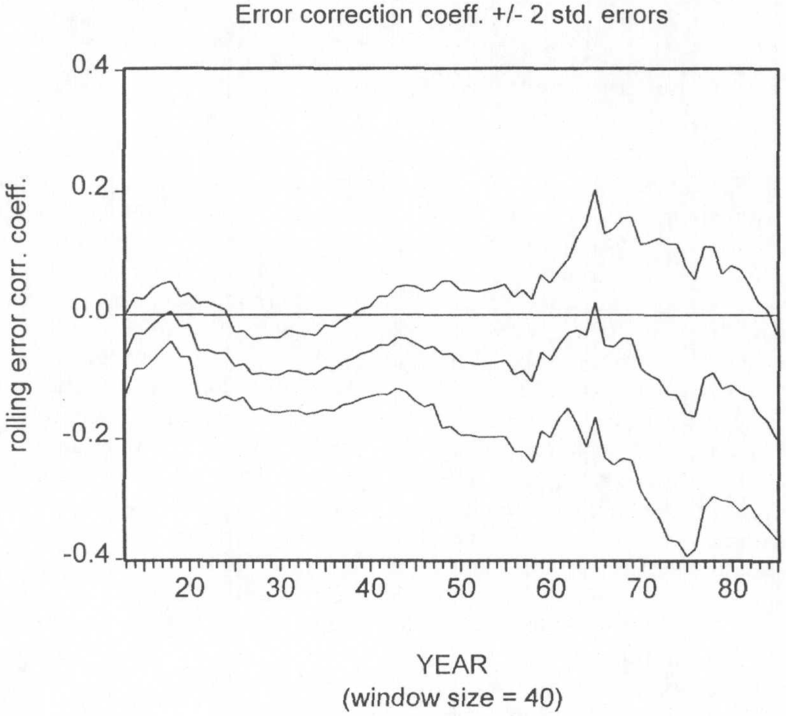
FIGURE 2
Rolling Estimates of Error Correction Coefficients in the Velocity Equation*





*Based on the VECMs reported in Table II for the U.K. (top panel) and Sweden (bottom panel).

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correction coefficients becomes larger for the U.K. following 1967 (the year of the sterling devaluation), indicative of relatively faster adjustment toward equilibrium in the post-Bretton Woods period.[21] This is interpreted as a reflection of the relatively greater impact of U.S. variables in the postwar period, that is, an indication of growing international financial integration since 1946. For Sweden, the impact of U.S. velocity is strongest during the 1925–1945 period but becomes more stable thereafter.[22] These results provide support for FSs and BJs earlier evidence of the existence of a unified financial system among the industrialized countries, as well as the dominant influence of U.S. velocity on velocity in the other countries. Figure 2 also suggests the possibility, because the size of the coefficients are rather different between the pre- and postwar samples, that a structural stability problem might exist.

A North Atlantic Velocity Function?

We performed additional cointegration tests to determine whether the long-run behavior of velocity is explained by conventional or institutional variables, or both, in a panel data set. This is accomplished by stacking the data on the conventional and institutional variables for all five countries.[23] Panel A in Table III tests whether permanent income and an interest rate jointly explain the long-run behavior of velocity (i.e., model 1 with $\beta_2' = 0$). Panel B of Table III adds the institutional determinants to test whether these can also explain long-run velocity. Table III also provides estimates of the long-run elasticities (i.e., the cointegrating vector) of each of the determinants with respect to velocity.

Panel A suggests that we are unable to reject the null of a single cointegrating vector between velocity, permanent income and an interest rate.[24] However, whereas the income elasticity is found not to be significantly different from one at the 10% level of significance, the interest elasticity is of the wrong sign and a test of the null of a zero interest elasticity is rejected.[25] On this basis, the conventional velocity model appears to be misspecified. When the institutional determinants of velocity are included along with the traditional determinants, the results in panel B lead to the conclusion that one cannot reject the null of four cointegrating vectors at the 10% level. Restrictions need to be imposed, therefore, in order to identify a cointegrating vector. It seems reasonable, based on the results of panel A, as well as the results for the individual series considered above, to impose unitary elasticity and a zero elasticity on the *lnal* variable.[26] The resulting vector is one for which the signs of all the coefficients conform with the theoretical predictions of both the conventional and institutionalist hypotheses of velocity, and thus this appears to be a well-specified model of velocity. Since the mean interest rate over all five countries is .054, this implies a North Atlantic interest elasticity of 0.22 which is well within the range found by FS [1982, Table 6.11] for the U.S. and the U.K., and Hafer and Jansen [1991] for the U.S.

The Stability of the Cointegrating Relationships

To establish the robustness of the results of the previous section to sample selection, we evaluated the cointegration test statistics first for the 1870–1913 sample and then by increasing the sample five years forward until the full sample was reached. In the case of the conventional model of velocity (equation (1) with $\beta_2' = 0$) the null cannot be rejected for any post-World War II sample for Canada; for the U.S. and Sweden there is a tendency for

21. Based on recursive estimates, the error correction coefficient is rather stable in the post-World War II period for both the U.K. and Sweden as well as being consistent with faster adjustment to equilibrium.

22. Essentially, comparable results were obtained from the recursions.

23. Clearly, by stacking the variables in such a manner we are assuming that the same specification works well for all five countries. The evidence in BJ [1987; 1990] and Siklos [1993] suggests that this is appropriate and that the coefficients in the model for individual countries are roughly similar. Given the potential loss of degrees of freedom we did not pursue additional refinements.

24. This result holds even if Norway is excluded as in Table I, panel A.

25. To give economic meaning to estimates of a cointegrating vector, coefficients must be normalized. Following previous convention (Siklos [1993]) estimates were normalized on velocity.

26. Results are similar if *lnal* is not constrained to zero, but the evidence in Table I is quite decisive about the lack of statistical significance of this variable. The various restrictions were imposed on all possible cointegrating vectors.

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TABLE III
Results of Pooled Cointegration Tests: Full Sample[a]

(A) Conventional Velocity Model (Canada – U.S. – U.K. – Sweden – Norway)

Model[b] (2.1)	Number of Cointegrating Vectors 0	1	2	Lag Length
Statistic	45.67*	11.29	2.52	8

Cointegrating Vector:[c] $[v, y^P, R]$ = [1, .746, –1.078]

Tests: $y^P = 1(\chi^2(1) = .078)$; $y^P = 1$ and $R = 0$ ($\chi^2(2) = 31.62$*); $R = 0$ ($\chi^2(1) = 25.59$*)

(B) Institutionalist Model of Velocity (Canada – U.S. – U.K. – Sweden – Norway)

Model (2.1)	Numbers of Cointegrating Vectors 0	1	2	3	4	5	Lag Length
Statistic	162.96*	91.34*	57.41*	32.70*	8.99	.02	7

Cointegrating Vector[2]: $[v, y^P, R, C/M, lnal, NBFA/FA]$ = [1,1,4.428,1.682,0,3.159]

Notes:

[a]See also notes to Table I for additional details about the manner in which Norway was added to the model. Dummy variables for the Great Depression and the two oil price shocks were also included as exogenous variable. Conclusions were not, however, affected by the inclusion of such dummies.

[b]Where Ω is set to zero.

[c]Normalized on velocity.

the same null not to be rejected for samples ending after 1958 (and beginning in 1873); for the U.K. there is considerable variation, with the inability to reject the null concentrated in pre-World War II samples as well as the full sample (1872–1985). By contrast, when the institutionalist model is considered, it is generally not possible to find cases where the null of no cointegration cannot be rejected, the only exception being the U.S. for the full sample (1873–1986). Similar results were obtained when the cointegration tests were generated via rolling regressions.

We also considered a number of other tests of the stability of the cointegrating relationships given in Tables I and III. Test statistics which complement the results of Table I, panel A, were also generated.[27] The null of a single cointegrating vector cannot be rejected beginning with the 1870–1983 sample only, so that the cointegrating rank is not, strictly speaking, constant. This points to a possible source of instability in equation (2). However, it should also be noted that the cointegrating rank is stable if a 15% significance level is chosen throughout all subsamples. The evidence against the stability of equation (2) is, therefore, not overwhelming. We also examined whether our findings for equation (1) reported in Table III would be affected by how the countries in the sample were grouped together. Thus, we compared estimates of equation (1), with and without the institutional variables, and found the results in Table III to be generally robust only when the institutionalist variables are added to the model. In other words, for the conventional velocity model, one cannot reject the null of no cointegration for any subsample (as in Table III, panel A). Hence, the pooled cointegration test results are specific to the full sample only, but the same was not found to be true for the institutionalist model where the null of zero

27. The interest rate test must be interpreted with some caution because, according to panel A, Table I, there are two cointegrating vectors for R. A drawback of the Gregory and Hansen [1996] approach is that, by relying on the EG [1987] methodology, it implicitly assumes the existence of a unique cointegrating relationship. We did find a peak in the ADF statistic for velocity in 1927, but it is not statistically significant even at the 10% level of significance.

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cointegrating vectors is easily rejected. The results are the same, at the 12.5% level of significance, as those shown in panel B of Table III when all countries are considered in a panel. Again, while it is possible that the "North Atlantic" velocity function is unstable, the statistical evidence against stability is not particularly strong.

Finally, the results of implementing the Gregory-Hansen [1996] test reveal that there is no apparent instability in the cointegrating relationships considered. Thus, the equilibrium relationship describing velocity, permanent income, and the interest rate across countries (i.e., as in equations (2) and (3)) does not appear to be subject to a regime shift.

V. CONCLUSIONS

This paper utilizes cointegration and error correction modeling to investigate the role of institutional factors in explaining the long-run behavior of the income velocity of money in five industrialized countries. Relying on recent work which suggests that institutional factors are important determinants of velocity's behavior in individual western industrialized countries, we asked in this paper whether these factors can explain the common U-shaped pattern of velocity for over a century of data for these countries. Notwithstanding the difficulties in measuring and assessing financial development and innovations (Boughton [1992]), the evidence presented in this paper suggests that institutional change is a good candidate to explain the striking similarities in the long-run behavior of velocity.[28] The importance of institutional factors is also reinforced by the finding that it is comparatively hard to detect instability in the long-run velocity model augmented with financial change proxies.[29]

The implications of our findings are important for at least three reasons. First, studies of the long-run behavior of velocity are inadequate if they exclude the impact of technological changes affecting the financial sector. Second, our empirical results clearly demonstrate that financial change is transmitted across countries. Third, the common features detected in institutional factors in velocity are not simply a post-1945 phenomenon; instead they emerged early this century if not earlier.

28. BJ [1987] and Silos [1993] have attempted to construct other determinants of institutional change, including proxies for expected inflation, with little effect on the empirical results for the institutionalists hypothesis.

29. Nor is this finding necessarily due to the length of the sample. Siklos [1993] shows that instabilities are more evident for narrower rather than the broader monetary aggregates.

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