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THE COMMON DEVELOPMENT OF INSTITUTIONAL CHANGE AS MEASURED BY INCOME VELOCITY: A CENTURY OF EVIDENCE FROM INDUSTRIALIZED COUNTRIES

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ABSTRACT

Previous evidence, most recently by Bordo and Jonung (1990) and Siklos (1988b, 1991), has shown on a country-by-country basis that proxies for institutional change significantly improve our understanding of the long-run behaviour of velocity and, consequently, of the demand for money.

If institutional change is a common development across industrialized countries it should have a common influence on velocity whereas the same need not be true for the other principal determinants of velocity such as income and interest rates. In statistical terms, this implies that the institutional change process should be cointegrated across countries but the conventional velocity determinants need not be.

The purpose of this study is to extend the existing evidence to study common features in velocity, income, and interest rates, across countries. The countries considered are Canada, the U.S., the U.K., Norway, and Sweden. We are relying on a sample of annual observations from 1870. The recently developed and refined techniques of testing for conintegration are used to study the common features in the series of interest.

Briefly, the evidence suggests support for the view that there exists a unique long-run relationship in velocity but not in income and interest rates and that the common feature in velocity is more apparent after rather than before World War II. However, before World War II, common features in velocity are more apparent for the U.S. and Canada, and separately, for Norway and Sweden. Finally, we find that only a model which includes institutional change proxies possesses a single common stochastic trend in the pooled time series, as well as long-run elasticities consistent with theoretical predictions. We argue that the evidence can only be understood in the context of common historical developments in the respective countries' financial systems.

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1. Introduction

The study of the long-run behaviour of velocity has intrigued many researchers who have sought to link it to the evolution of financial systems over time. Indeed, Friedman and Schwartz (1982; hereafter FS), in their seminal study, view financial sophistication as an important determinant of the long-run behaviour of velocity in the US and the UK. The aim of this study is to explore the connection between long-run velocity movements across several countries, as well as the connection between countries of its principal institutional and economic determinants.

Bordo and Jonung (1981, 1987; hereafter BJ), propose an institutional change explanation for the long-run behaviour of velocity, while Siklos (1988b, 1992) suggests that 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. These studies relied on a long sample of annual data from five industrialized countries. Many economists now agree that institutional change represents an important feature in understanding the long-run behaviour of velocity or the demand for money (e.g., Poole 1988, Darby et. al. 1987, Laidler 1982, 1985, 1990).[1]

2. Motivation

Recently developed econometric methods are now better able to address the nature of long-run relationships between time series. In particular, testing for cointegration means that the statistical evaluation of long-run relationships between time series is no longer subject to the problem of spurious regressions (Granger and Newbold 1974) which may have plagued earlier studies. Cointegration means that a set of time series act as "attractors" to each other (Granger 1990), that is, the series form an equilibrium relationship in the statistical sense. This does not preclude the possibility that any such relationship can be disturbed, even if only temporarily. On this basis Gregory

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

and Nason (1991), and Gregory and Hansen (1992), develop tests of stability for cointegrating relationships stemming from the work of Hansen (1992). Gregory and Nason find that for the same long annual US sample (1901-85) which Lucas (1988) used to demonstrate the empirical stability of long-run money demand, it is comparatively difficult to conclude that there are structural breaks in a US demand for money function of the traditional variety. When different sets of tests are used. Gregory and Hansen (1992) do find a break in a conventional U.S. money demand function during the early 1940s.

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, Hafer and Jansen 1991, and Hoffman and Rasche 1991, represent only a partial list of recent contributions in this area. Existing empirical evidence suggests that a broad monetary aggregate, usually M2, income and nominal interest rates are cointegrated over a variety of samples, at least for US data.

By contrast, empirical evidence is decidedly mixed for models which use an M1 definition of money. Hoffman and Rasche (1991) find that previous studies of M1 behaviour were misspecified and that an equilibrium relationship within a conventional money demand model can be found. Baba, Hendry, and Starr (1992) conclude otherwise since their long-run model is complex incorporating yield spreads and risk features of interest rate behaviour. Mehra (1992), and Miller (1991), also find that M1 is an unreliable variable for understanding short-run money demand behaviour. Hafer and Jansen (1991) prefer M2 over M1 for US annual data since 1915 in the sense of a finding of cointegration among the variables in a conventional money demand relation. Because the definition M2 incorporates over time the influence of financial innovations (Hester 1981), perhaps this explains a potential source of money demand instability in M1 based models (see also Baba, Hendry, and Starr 1992).

In a related development Ramey (1991) has formulated a real business cycle model in which the financial sector plays an important role and she presents empirical evidence to support the view that "technological innovations", which could be interpreted as a proxy for institutional change, are significant in US data. Recent theoretical work has also attempted to model the role of technological changes in the financial sector. Ireland (1992) 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 behaviour of velocity in particular.

This study examines the long-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 (Hendry 1986, Perron 1989). In particular, the effects of technological or institutional changes in the financial sector occur slowly. It is for this reason that we chose as long a sample as possible.

Given recent findings (BJ 1990, Siklos 1991) 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. BJ (1987, ch. 4, p. 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 behaviour of velocity. 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 of any such relationships as well as the stability of any unique cointegrating relationships which are detected. Finally, we attempt to estimate a joint velocity function for all the countries considered by pooling data for all the countries in our sample. In so doing, we improve on the earlier evidence on long-run common movements in velocity, as well as its conventional and institutional determinants, presented by BJ (1987, ch. 4) and FS (1982, ch. 7). Thus, BJ did not consider the problems which stem from the time series properties of the variables under study, especially the well-known spurious regression problem (Granger and Newbold 1974). FS, who rely on the phase-averaging technique to retrieve the secular component of a time series may also have biased their results (Campos, Ericsson, and Hendry (1990), Hendry and Ericsson (1991), but see FS (1991)).

In performing time series tests our objectives are three-fold. First, to explore whether the common features of financial systems across countries are as significant as Friedman and Schwartz (1982, ch. 7) found was true of the US and the UK, based on sophisticated measures of correlation. However, we expand the selection of countries to include Canada and Sweden.[2] 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, whether countries at similar stages of development in effect import payment technologies from other countries. A third objective of the paper 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. In a sense, then, we attempt to combine the narrative and statistical approaches to the issues of interest in this paper (as did Romer and Romer 1989).

Briefly, the results suggest support for the view that there exists a single cross-country long-run relationship in velocity but not in income and interest rates. Moreover, common features in velocity across countries are more apparent after World War II than before the war. Finally, we find that only a model which includes institutional change proxies possesses a common stochastic trend in the pooled time series case model, as well as 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.

[2] Norway is included in sub-sample estimation but could not be included in full sample estimates because of gaps in the data.

After a brief review of the institutionalist hypothesis of the long-run behaviour of velocity (section 3), and a description of econometric issues (section 4), empirical evidence confirming the above conclusions are presented (section 5). The paper concludes with a summary in section 6.

3. The Institutionalist Approach: A Review

Since much has been written about the institutionalist explanation of the long-run behaviour of velocity, advanced by Bordo and Jonung (1981, hereafter BJ, see Bordo and Jonung 1987; Siklos 1988a, 1992; Ireland 1991; Laidler 1982, 1985, 1990; Hallman, Small and Porter 1991) our review will be a brief one.

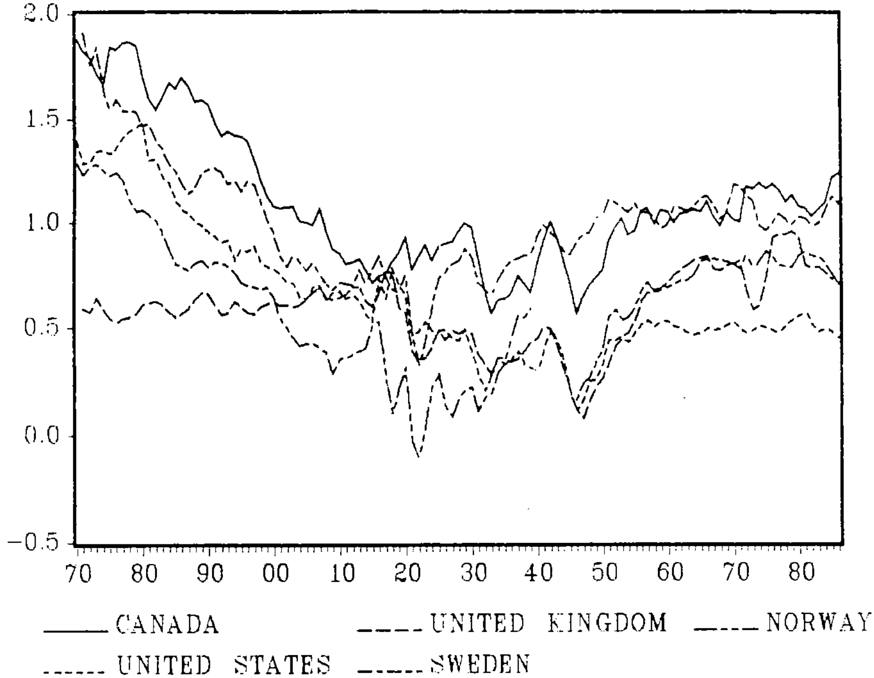
Velocity is traditionally viewed as an analogue of the demand for real balances. Consequently, it is treated as a function of income (or permanent income) and an interest rate. The latter variable serves as a proxy for the opportunity cost of holding money.[3]

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 deposits and notes. A second stage is characterized by growing financial sophistication during which the number of substitutes for notes and bank deposits grows. The combination of these two factors produces a U-shaped long-run pattern for the countries shown in Figure 1 (Bordo and Jonung 1987, ch. 2), with a falling trend that ends during the interwar period and with a rising trend starting for most countries in our sample in the mid-1940s.

[3] Specifications which examine the determinants of real balances have been preferred in part because of the finding that velocity behaves like a random walk (see Raj and Siklos 1988 for a brief survey). Nevertheless, given the difficulties surrounding tests of the random walk hypothesis (see Campbell and Perron 1991 for a recent survey), and the presence of statistical breaks in the random walk behaviour of velocity (Siklos 1988a, and Perron 1989, 1991), the empirical evidence suggests on balance that the random walk hypothesis is not a substitute for a complete model of velocity's behaviour. More on these issues below.

Figure 1

The Long-Run Behaviour of Velocity in Five Industrialized Countries*



* Annual data from BJ (1990). For Norway, no data are available for the war years (1940-45).

Bordo and Jonung attribute the downward trend in velocity before World War II to the process of monetization, and the upward trend since[4] to two developments: increasing financial sophistication and improved economic stability.[5]

BJ (1981, 1987, 1990) develop four separate proxies to capture institutional change. These are: the ratio of total non-bank financial assets to total financial assets, the currency-money ratio, the share of the labour force in non-agricultural pursuits, and a measure of economic stability. BJ (1987), and Siklos (1992), describe in detail how these variables relate to financial development, while Siklos (1992) discusses some of the drawbacks with existing proxies. Despite the problems with existing measures of institutional change no one has yet been able to suggest alternative proxies. There is, however, more of a consensus about the importance of changes in the financial sector in influencing the long-run behaviour of velocity (see Siklos 1992, and references within).

The striking similarity in the behaviour of velocity across industrialized countries suggests a common financial development in different countries despite differences in fiscal and monetary policies, in their inflationary experiences and industrial development.[6] 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 and Backus, Kehoe, and Kydland 1991, Engle and Kozicki 1991), none of the

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[4] There have been some interruptions since, such as during the 1980s when velocity in M1 levels began to level off and even decline in the US. See Darby et. al. (1987). 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.

[5] Whether the postwar period produced more stable variation in economic aggregates such as GNP or unemployment in the U.S. in particular has been the subject of a debate which remains unsettled. For a sampling of views on this subject, see, for example, Romer (1986, 1988, 1989), and Balke and Gordon (1989).

[6] By pooling time series, Bordo and Jonung (1987, ch. 4) assume that the common behaviour of velocity is explained by common ecor c determinants. They did not, however, test the validity of such an assumption.

studies, to our knowledge, have applied tests of cointegration either to test whether financial change is common across countries nor the sources of common movements if they exist.[7]

3. Econometric Specification

3.1 Background

The fundamental hypothesis of this paper may be stated 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

$$v_{it} = \beta_0' + \beta_1' \Phi_{it} + \beta_2' \Omega_{it} + \varepsilon_t \tag{2.1}$$

$\Phi_t = [y^p_t, R_t]$, where y^p is real per capita permanent income, and R is a proxy for the opportunity cost of holding money. $\Omega = [(NBFA/FA)_t, (C/M)_t, lnal_t]$, where NBFA/FA is the ratio of total non-bank financial assets to total financial assets, (C/M) is the currency-money ratio, *lnal* is the share of the labour 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, McCallum and Goodfriend 1987, and Judd and Scadding 1982, for surveys). The reasoning behind the vector Ω can be stated briefly as follows. Since NBFA/FA proxies financial development it would be expected to reduce the demand for money by increasing the number of close substitutes thereby raising velocity. BJ hypothesize that the C/M series mirrors the spread of commercial banking. Thus, as banking spreads, velocity falls at first because of growing reliance on money to settle transactions. Later, velocity rises as the financial system becomes more sophisticated, that is, as the number of substitutes for money grows. Therefore, changes in velocity are positively related to changes in C/M.

[7] We would have liked to expand the data set to consider other countries, as in Backus and Kehoe (1992), who kindly made available their data to us. We are unable to do so for at least three reasons. First, the power of the tests which are applied below fall 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 the only ones which, for historical and economic reasons, are the most likely candidates for common institutional and economic development.

The steady rise in the proportion of the labour force in non-agricultural pursuits reflects growing urbanization and, as a result, also reflects the "spread of the monetary economy" (BJ 1987, p. 34). Other things being equal BJ predict that, as this series rises over time, velocity falls.[8]

Previous research by one of the authors has concentrated on estimates of (2.1) for individual countries as in BJ (1981, 1987, 1990), Raj and Siklos (1988), and Siklos (1992). Because these studies have suggested that the Ω vector, in particular, significantly explains velocity in each of the countries considered this would suggest that if velocity is common to all countries this may be due to common features in Ω or Φ, or both. Thus, the object of the empirical analysis is to examine whether the following linear combinations are stationary, that is, whether they are cointegrated.

$$v_{it} + \delta_0' v_{jt} = \varepsilon_v \tag{2.2}$$

$$\Phi_{it} + \delta_1' \Phi_{jt} = \varepsilon_\Phi \tag{2.3}$$

$$\Omega_{it} + \delta_2' \Omega_{jt} = \varepsilon_\Omega, \quad i \neq j \tag{2.4}$$

where v is a vector to indicate that a cointegrating relationship between 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_r are stationary. To illustrate, suppose we have a sample consisting of data from two countries. The finding that a 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. Below, we consider whether one or more linear combinations of

[8] Omitted from (2.1) is BJ's measure of economic stability, a six-year moving standard deviation of the annual percent change in real per capita income. Using moving standard deviation proxies of volatility are problematical, as pointed out in Huizinga and Mishkin (1986).

velocity, its traditional and institutional determinants are stationary for the five countries in our sample.

Suppose that Φ_i and Φ_j are not cointegrated but one is unable to reject a finding of cointegration between Ω_i and Ω_j. Assuming that Φ and Ω are independent of each other, the explanation for the common movement in v_i and v_j could be due to common movements in elements of Ω_i and Ω_j. Since the latter vector proxies financial development this implies that velocity is common because financial development is common in one or more of the countries sampled, despite the lack of cointegration between income and interest rates.[9] These results also suggest that a single demand for money function common to several countries may exist, as FS (1982) suggested was true for the US and the UK. Consequently, a next logical step would be to estimate (2.2) to (2.4) jointly in a pooled sample which we also consider below. The approach outlined in (2.2) to (2.4) 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. The cointegration approach is meant precisely to address these issues. Prior to the discussion of the empirical approach we briefly discuss the econometric methodology employed in this paper.

3.2 Testing for Unit Roots and Cointegration

3.2.1 Unit Roots

There exist now a large literature about whether economic time series are stationary around a deterministic trend, or the sum of a permanent component best described as a random walk perhaps with a drift, and a transitory component which is white noise. This paper begins with the view that each of the series in equation (2.1) 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 Raj and Siklos

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 (Siklos 1988b).

(1989), and Siklos (1988a, 1988b, 1992). Testing for unit roots is a first-step in determining whether two or more series are cointegrated. The reason is that statistical theory requires that the time series under investigation be covariance stationary. Quite often, however, time series exhibit a trend but they can be rendered stationary by first-differencing, which suggests that the series has a unit root. However, if this is the case, then it is possible that two or more series possess a common unit root. Hence, a linear combination of non-stationary time series can be stationary.[10] This suggests that a (long-run) equilibrium between two or more series exists as described by a stationary line or combination called the cointegrating vector.

Despite existing unit root test results for the data used in this study some questions have been raised about whether unit root findings may be biased in the presence of a structural break in the data (Reichlin 1989, Hendry and Neale 1990). Accordingly, Table 1 presents further evidence on the unit root hypothesis using the recently developed tests by Perron (1990), see also Perron and Vogelsang (1992). For each series the year in which a statistical break is most likely is shown as well as whether the series possesses a unit root. Rejection of the unit root hypothesis means that the series is trend stationary around a "broken" trend.

However, unlike Perron's (1989) earlier test of the unit root hypothesis, in which time series were represented as stationary around a **pre-selected** broken deterministic trend, the modified test permits the data select the optimal break point.[11] Only one possible break is allowed for each series and the break-point is selected at the point at which the "t-statistic" for the null of a unit root is

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[10] 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.

[11] Zivot and Andrews (1992) develop a similar test. Apart from the consideration of whether there is a break-point in the data there is the issue of the power (i.e., the ability) of existing unit root tests to distinguish between difference stationary and trend stationary processes. See, for example, Christiano and Eichenbaum (1989).

Table 1

Test for Unit Root and Broken Trend[1]

Country	Sample	Series	k	Break[3]	Unit Root
Canada	1870-1986[2]	V	0	1922	-4.11
		y^p	1	1929	-5.11
		R	0	1964	-3.30
		C/M	0	1932	-2.13
		Inal	0	1910	-0.41
		NBFA/FA	0	1896	-3.83
US	1870-1986[2]	V	0	1945	-3.40
		y^p	1	1918	-5.15
		R	0	1966	-3.75
		C/M	0	1933	-4.91
		Inal	0	1940	-4.58
		NBFA/FA	1	1944	-3.71
UK	1870-1985[2]	V	0	1945	-5.67*
		y^p	1	1916	-4.11
		R	0	1967	-5.57*
		C/M	1	1938	-5.87*
		Inal	1	1877	-1.63
		NBFA/FA	0	1970	-4.77
Sweden	1870-1986[2]	V	1	1917	-3.11
		y^p	0	1898	-37.61*
		R	3	1960	-0.33
		C/M	1	1929	-4.01
		Inal	0	1898	-52.10*
		NBFA/FA	0	1972	-2.82

Notes:

1. Using Perron's (1990) test equation: $y_t = \mu + \theta\, DU_t + \beta t + \gamma DT_t + \delta\, D(TB)_t + \alpha y_{t-1} + \sum_{i=0}^{k} c_i\, \Delta y_{t-i} + e_t$

where
- y_t = time series at time t
- μ = constant
- t = trend
- TB = trend break dummy = t when t > TB, 0 otherwise
- DU = intercept dummy = 1 when t > TB, 0 otherwise
- Δ = difference operator
- k = value of autoregressive correction factor. Lag length chosen on the basis of Schwarz's criterion.

In the table above Break = TB, Unit Root = "t"-statistic on α.

2. For the UK the sample ends in 1985; for NBFA/FA (Sweden), the sample begins in 1880; for y^p (Canada), the sample begins in 1900; for C/M (Sweden), the sample begins in 1871.

3. All break-points are statistically significant.

* signifies rejection of the unit root hypothesis at the 5% level of significance. Critical value is -5.41 (k=0), =5.29 (k=2), -5.22 (k=5). See Perron (1990, Table II).

smallest for all possible-break points indicating the greatest likelihood of a break in the time series.[12] For all the US and Canadian series the null of a unit root still cannot be rejected. The unit root is rejected for UK velocity, interest rate and the currency-money ratio. For Sweden, permanent income and the labour force variables appear stationary around a broken trend. However, as the breaks occur very early in the sample (in 1898 in both cases) is unlikely that this will be a major factor in the tests below since data restrictions for some of the other countries would exclude observations at the break-point.[13] Accordingly, we also conducted the tests for the 1900-86 sub-sample for Sweden's permanent income and the *lnal* variable and found that we could no longer reject the null of a unit root test (statistics with the break-point in parenthesis were, respectively, -4.38 k=1 (1918), and -4.41 k=1 (1940)). It is interesting to note, however, that the test usually selects interwar or war years for a break in velocity as Bordo and Jonung (1987) hypothesized but, in every case, the chosen year differs substantially from that selected arbitrarily by BJ to explore periods of rising and falling velocity. It is also worth noting that a break in interest rates always occurs well after a break in velocity. By contrast, the year when one or more of the institutional time series breaks usually precedes when a break in velocity occurs.

As instructive as unit root tests can be in describing the univariate properties of a time series, and as a prelude to testing for cointegration, such tests are not a substitute for testing of whether linear combinations of several series are jointly stationary, a task to which we now turn.

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[12] Since the t-statistic is negative this means a large negative value is necessary to reject the null of a unit root.

[13] The unit root results in Table 1 are not particularly sensitive to the lag selection procedure in the autoregressive component of the test operation (although the timing of the break point was). The lag was chosen on the basis of the well-known Schwarz criterion (SC) because of the size of the sample. Alternatively, we would have chosen the lag length using Akaike's information criterion (AIC) which asymptotically overestimates the true lag order. Perron and Vogelsang (1992) recommend selecting the lag according to the longest lag for which the t-statistic is statistically significant. The chosen lag tends to fall somewhere in between the SC and AIC methods. Test results are, however, sensitive to whether the break-point is selected independently of the data since Perron (1989) found the break in US velocity to be 1929 although he reaches the same conclusion when the test is modified as explained above (Perron 1990) using a slightly different data set. Zivot and Andrews' (1992) finding of a break in US velocity in 1949 is closest to our findings in Table 1.

3.2.2 Cointegration

Granger (1983) introduced the notion of cointegration to describe 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 emerged which has refined and improved the original single-equation testing procedure presented in Engle and Granger (1987; EG). A rapidly growing empirical literature also exists which has applied these tests to a variety of economic problems.

The approach used here to study common features in time series is based on the work of Johansen (1990), and Johansen and Juselius (1990a,b). Among the virtues of Johansen's approach is that, unlike the EG procedure, it permits the investigator to determine the number of cointegrating relationships which may exist between the series of interest.[14] In addition, Johansen's procedure enables one to perform a variety of tests of various restrictions imposed on a model.[15] Also, since Johansen's procedure uses the vector autoregressive (VAR) approach all the variables are treated as endogenous. This is a distinct advantage in the present context since there is some question about the exogeneity of some of the series assumed to independently influence velocity (Hamilton 1989). Finally, whereas the EG procedure is estimated via Ordinary Least Squares (OLS), the Johansen procedure resorts to maximum likelihood estimation.

The Johansen procedure consists in estimating a VAR for a vector of time series X_t.

$$\Pi(L)X_t = \varepsilon_t \tag{3.1}$$

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[14] An alternative procedure, developed by Stock and Watson (1988), can also be used to determine the number of cointegrating vectors in a multivariate setting.

[15] Gonzalo (1989) shows that Johansen's procedure is at least as powerful as others under various specification errors.

where $\Pi(L)$ is a polynomial distributed lag of order k chosen so that there is no serial correlation in the residuals. Under the hypothesis that the time series in X_t have a unit root, it is convenient to transform (3.1) as follows:[16]

$$\Delta X_t = \Gamma_i \Delta X_{t-1} + \dots + \Gamma_{k-1} \Delta X_{t-k-1} + \Pi X_{t-k} + \mu + \varepsilon_t \tag{3.2}$$

where $\Gamma_{i\,=\,-1+\Pi} = 1 + \dots + \Pi_i$, $i = 1, \dots, k-1$, μ is a constant term, and $\Pi = -(I - \Pi_1 - \dots - \Pi_k)$. Equation (3.2) is simply the so-called error correction representation of (3.1) with error correction term X_{t-k} in the VAR format. Error correction imposes the long-run equilibrium relationship between the elements of X while permitting short-run deviations from this equilibrium. The matrix Π contains information about the long-run properties of the model and the finding of cointegration is determined by examining the rank of the Π matrix. When the rank of Π is zero (3.2) reduces to a VAR in first differences. The relevant economic model is then one between unrelated differenced time series. If the rank Π of r is one, then there exists a unique cointegrating relationship between the series in the VAR. When the rank is greater than zero but is less than full rank (k), then there are k cointegrating relations among the elements of X_t, and k-r common stochastic trends. Thus, there may exist a linear combination of the time series which is stationary even if the individual series themselves are not. If the series are cointegrated then one can decompose Π under the null hypothesis so that $\Pi = \alpha\beta'$, where β is the matrix of long-run parameters and α are the error correction parameters. Johansen's test proceeds by concentrating the likelihood function with respect to Π. A test for cointegration is

[16] First, subtract X_{t-1} from both sides of (3.1). Next, add $(-\Pi_1 - 1)X_{t-2} + (\Pi_1 - 1)X_{t-2}$ to the right hand side of (3.1). Next, $-(\Pi_2 + \Pi_1 - 1)X_{t-3} + (\Pi_2 + \Pi_1 - 1)X_{t-3}$ is added, and so on until (3.2) emerges.

suggested by the likelihood-ratio test which permits testing for the number of cointegrating relationships as well as other interesting restrictions.[17]

An important issue in the present context is the selection of lag length in the VAR. Again, while current research is continuing to debate the merits of different lag selection procedures lag lengths in this study were selected on the basis of the well-known Akaike Information Criterion (AIC) primarily because it tends to select relatively long lags thereby reducing the chances of certain types of specification errors.[18]

Two test statistics can be used to evaluate the number of cointegrating relationships. The trace test examines the rank of Π and the hypothesis that rank $(\Pi) \leq r$ is tested, where r represents the number of cointegrating vectors. Alternatively, the maximal eigenvalue test can be employed in which the null of r cointegrating vectors is tested against the alternative of r+1 vectors. For long samples such as the one considered here the two tests generally yield the same conclusions. Results based on the trace test only are reported below while other test results are available from the last author.

3.2.3 The Stability of Cointegrating Relationships

Even if one or more cointegrating relationship is found there is the possibility that structural breaks appear in any such relationship. Thus, for example, 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. Bordo and Jonung (1987, ch. 4) examined separately periods of falling and rising velocity and found relatively 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

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[17] Johansen and Juselius (1990) have produced critical values for cases where (3.1) contains no constant, contains a constant vector, and contains a constant vector restricted to lie in the cointegration space. These are their Tables A1, A2, and A3, respectively. Ostewald-Lenum (1992) has produced improved estimates of Johansen's critical values which are used in the empirical work which follows.

[18] Kasa (1992) similarly reports that a longer VAR specification yields better results based on diagnostic tests of the error terms of the VAR.

chosen break point is appropriate.[19] FS (ch. 7), by contrast, adjust their estimates of the relationship between secular movements in velocity between the U.S. and the U.K. by using dummy variables for wars and the Depression. Similarly, FS document the fact that while velocity movements in the U.S. and the U.K. "reflect a unified financial system" (FS, p. 337), some differences exist during the pre-1914 period. This is apparent from Figure 1 which suggests that velocity was falling in all of the countries considered, except the U.K. which exhibited only a slight fall overall in the period 1870-1914.

Several responses are available to address these issues. One is to test for structural breaks at particular known dates. For example, the dating of wars is widely agreed upon and the same is true perhaps of the Great Depression, and oil price shocks. What we do not know, however, is when the effects of a particular event will influence the aggregates under study. Moreover, 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. In this paper we therefore implement recently developed tests for stability in cointegrated relationships where the timing of such a break is unknown.[20]

We could also test for cointegration for selected sub-samples, such as the 1870-1914 period, or test for cointegration conditional on the presence of shocks arising from the two world wars, the Great Depression, and the two oil price shocks, when these are assumed to be exogenous. Tests for cointegration, however, require not just a sufficiently large sample but one over a sufficiently long period as well. Although over 100 annual observations are available caution must be exercised in

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[19] Indeed, the coefficients in their velocity model which incorporates institutional change factors show signs of a structural break in only one of the five countries considered (Canada; Bordo and Jonung 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.

[20] Perron (1989, 1991), however, finds little difference in results of univariate tests of stability relative to the case where the breaks are known or unknown. See, however, Zivot and Andrews (1992) on this point.

interpreting cointegration tests in smaller samples. Alternatively, however, we can estimate any cointegrating relationship 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.

4. Data and Empirical Results

4.1 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. In any event, our results allow comparability with earlier studies of velocity and long-run studies of US money demand using annual data (e.g., Lucas 1988, Hafer and Jansen 1988). Readers are referred to BJ (1987) for additional details about the data. Five countries are considered in the empirical results reported below. They are: the US, UK, Canada, Sweden, and Norway.

4.2 Empirical Results

This discussion of empirical results can be subdivided into three parts. First, we ask whether the series contained in (2.2) to (2.4) are cointegrated. Next, we examine whether the institutionalist hypothesis can better explain the long-run behaviour of a pooled velocity function than the conventional model as well as analyzing some of the properties of the cointegration test results. Finally, we address the question of whether any of the cointegrated relationships which are found are stable in a statistical sense.

4.2.1 Testing for Cointegration

Table 2 presents tests of cointegration for the whole sample based on the Johansen methodology outlined previously. Panel A of the table tests for cointegration using data for the four countries where the series are available for the full sample. Panel B of the table adds Norway to the list of countries considered but omits the years 1921-22 and 1940-45 in a few cases. Broadly speaking, the results are the same in both cases. We find that one cannot reject for velocity the null that a unique cointegrating relationship exists between velocity for Canada, the US, the UK, and Sweden. That is, the null hypothesis that the number of cointegrating vectors r is 1 given that r = 0 cannot be rejected. To the extent that velocity reflects income, interest rate and institutional changes, the results reflect the statistical confirmation that these countries can be viewed a single economic entity. These results would be the analogue of the Backus and Kehoe (1992) findings of striking similarities in international business cycles.

The remaining cointegration test results in Table 2 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 log levels of y^P for the four countries in our data set. Thus, if velocity is common to all the countries considered, the common trend could be partly explained by common permanent income movements. The results differ, however, when the truncated sample is considered (Panel B). There we find that at least four cointegrating vectors exist for permanent income. Thus, if we essentially exclude the turbulent war years there is evidence of a common stochastic trend in income but not of a unique equilibrium relationship for all the countries considered.

Table 2

Cointegration Test Statistics

(A) Canada - US - UK - Sweden

Series	Number of Cointegrating Vectors				Lag Length	Sample[1]
	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 - US - UK - Sweden - Norway

Series	Number of Cointegrating Vectors					Lag Length	Sample2
	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 $r \leq j$, $i \neq j$, 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 DGP.

1. Before lags are taken into account.

2. Same as above except data for 1921-22, and 1940-45, for velocity and real per capita permanent income were excluded because data were unavailable for Norway. A shift dummy was used to splice Norwegian velocity data. No shift dummies were used for the other series.

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 cointegrating vectors between interest rates. 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 behaviour of velocity, as Raj and Siklos (1989) suggested. Bordo and Jonung (1981, 1987) had earlier suggested that interest rates might be a relatively more important variable in explaining the long-run behaviour of velocity than permanent income.

The cointegration test results for the institutional proxies suggest that a single cointegrating vector is found for the currency-money ratio as well as the financial sophistication proxy (NBFA/FA). Thus, there appear to be long-run common features in institutional change. For the labour 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 (and the maximal eigenvalue test, not shown). When Norway is included the results are the same, except that instead of one cointegrating vector for the financial sophistication series we cannot reject the null of two vectors.

4.2.2 The Transmission of Institutional Change

The existence of cointegration between several of the variables considered suggest that their dynamic relationship can be modelled via a vector autoregression augmented by error correction terms. These so-called vector error correction models (VECM) are useful as a further test of the cointegration hypothesis,[21] as a device to determine the size of the error or deviation in an equilibrium relationship, and to determine which variable in a system Granger-causes other variables.[22] These are useful questions to explore in the present context since we are interested in whether departures in the series considered here are corrected slowly over time. The results of the

[21] On the meaning of error correction in different contexts, see Alogoskoufis and Smith (1991).

[22] There may be problems, however, with the use of VECMs as a tool to determine the existence of a Granger-causal relationship between time series. See Toda and Philipps (1991).

estimation of VECMs for velocity are provided in Table 3. The error correction terms, z, are statistically significant and of the correct sign in all of the regressions except in equation 2, where for Canada's velocity the error correction term is statistically insignificant. Because the error correction term is not statistically significant in the US equation, this suggests that US velocity Granger-causes velocity in the other countries. The reason is that a significant coefficient on the error correction term means that deviations from US velocity influence velocity in other countries. By the same token none of the other velocity series, with the possible exception of the UK, Granger-causes US velocity. 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 UK and Swedish cases. Further insights may be gained from a recursive estimation of the equations in Table 3. Figure 2 plots the recursive estimates, along with the standard error bands, of the error correction terms for equations 1 and 3 in Table 3. These were the equations which had the statistically significant error correction terms. It is interesting to note that the error correction coefficient is very stable in the post World War II period as well as being consistent with faster adjustment to equilibrium. This is interpreted as a reflection of the relatively greater impact of US variables in the postwar period, that is, an indication of greater international financial integration since 1946. These results provide broad support for FS's and BJ's earlier evidence of the existence of a unified financial system among the industrialized countries as well as the significant influence of U.S. velocity on velocity in the other countries.

4.2.3 A Global Velocity Function?

Because the evidence in the preceding section suggests that the five countries in our sample can be treated as one entity we perform cointegration tests to determine whether, in a pooled sample, the long-run behaviour of velocity is explained by conventional or institutional variables, or both.[23]

[23] Siklos (1992) considers this question on a country-by-country basis in the cointegration framework.

Table 3

Error Correction Models[1]

1. $$\Delta v_{UK} = -\underset{(.041)}{.075} + \underset{(.0003)}{.0001T} + \underset{(.108)^*}{.627\Delta v_{US}(-1)} - \underset{(.123)}{.151\Delta v_{US}(-2)} - \underset{(.197)}{.237\Delta v_C(-1)}$$
$$- \underset{(.195)}{.149\Delta v_C(-2)} + \underset{(.115)^*}{.254\Delta v_{UK}(-1)} + \underset{(.093)^*}{.058\Delta v_{UK}(-2)} + \underset{(.128)}{.018\Delta v_S(-1)}$$
$$+ \underset{(.131)}{.133\Delta v_S(-2)} - \underset{(.031)^*}{.069z(-1)}$$

$$R^2 = .453, \quad F(10,101) = 8.380^*, \quad SC(1) = 1.053,$$

2. $$\Delta v_C = \underset{(.024)}{.016} + \underset{(.002)}{.0003T} + \underset{(.065)}{.082\Delta v_{US}(-1)} - \underset{(.074)^*}{.181\Delta v_{US}(-2)} + \underset{(.118)}{.129\Delta v_C(-1)}$$
$$- \underset{(.117)}{.0005\Delta v_C(-2)} + \underset{(.069)}{.083\Delta v_{UK}(-1)} + \underset{(.056)}{.059\Delta v_{UK}(-2)} - \underset{(.077)}{.094\Delta v_S(-1)}$$
$$+ \underset{(.079)}{.114\Delta v_S(-2)} + \underset{(.019)}{.034z(-1)}$$

$$R^2 = .217, \quad F(10,102) = 2.833^*, \quad SC(1) = .798,$$

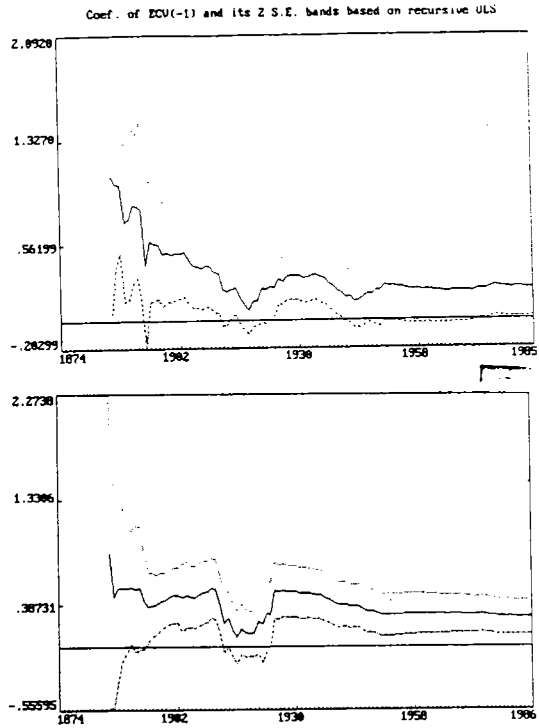
3. $$\Delta v_S = -\underset{(.029)^*}{.115} + \underset{(.0002)^*}{.0005T} + \underset{(.077)}{.117\Delta v_{US}(-1)} + \underset{(.088)}{.025\Delta v_{US}(-2)} - \underset{(.139)}{.105\Delta v_C(-1)}$$
$$- \underset{(.139)^*}{.374\Delta v_C(-2)} + \underset{(.082)}{.067\Delta v_{UK}(-1)} + \underset{(.066)}{.101\Delta v_{UK}(-2)} + \underset{(.091)^*}{.291\Delta v_S(-1)}$$
$$- \underset{(.093)^*}{.221\Delta v_S(-2)} - \underset{(.022)^*}{.077z(-1)}$$

$$R^2 = .304, \quad F(10,102) = 4.451^*, \quad SC(1) = .144,$$

1 Standard errors in parenthesis. * specifies rejection of the null at the 10% level of significance. T is a time trend, 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.

Figure 2

Recursive Estimates of Error Correction Terms in the Velocity Equation*



* Based on the VECMs reported in Table 4 for the UK (top panel) and Sweden (bottom panel).

Panel A in Table 4 tests whether permanent income and an interest rate jointly explain the long-run behaviour of velocity. Panel B of the same table adds the institutional determinants in testing whether these can also explain long-run velocity. The Table also provides estimates of the long-run elasticities 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[25] is found not to be significantly different from one at the 10% level of significance ($\chi^2(1) = 3.24$ (.07)); degrees of freedom and significance level, respectively, in parenthesis), the interest elasticity is of the wrong sign and a test of the null of a zero interest elasticity is rejected at any conventional statistical level ($\chi^2(1) = 25.59$ (.00)). When the institutional determinants of velocity are included along with the traditional determinants the results in panel B suggest that common stochastic trends exist between all the variables (one common stochastic trend when Norway is excluded) since the null of 4 cointegrating vectors cannot be rejected at the 10% level of significance. However, examination of the vectors reveals one for which the sign of all the coefficients conforms with the theoretical predictions of both the conventional and institutionalist hypotheses of velocity. Note also that the income elasticity is now considerably less than one while the interest elasticity is of the correct sign and exceeds 0.6.[26] Compared with the long-run income and interest elasticities for the U.S. derived by Hafer and Jansen (1991) for annual data from 1915 to 1988, the implied income elasticity is considerably lower (.165 versus .89) while the interest elasticity is also considerably higher (.613 versus .36) than for U.S. data alone as reported in Hafer and Jansen.

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[24] This result holds even if Norway is excluded as in Table 2.

[25] To give economic meaning to estimates of a cointegrating vector coefficients must be normalized. Following previous convention (Siklos 1992) estimates were normalized on velocity.

[26] Both are statistically significant. For income the test statistic is $\chi^2(1) = 13.38$ (.02); for the interest rate the test statistic is $\chi^2(1) = 58.59$ (.00).

Table 4

Results of Pooled Cointegration Tests: Full Sample[@]

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

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

Cointegrating Vector:[2] $[v, y^P, R] = [1, .746, -107.816]$

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

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

Cointegrating Vector: $[v, y^P, R, C/M, \mathit{lnal}, NBFA/FA] = [1, .165, .613, .856, -4.658, 2.676]$

Notes: @ See also notes to Table 2 for additional details.

1. Where Ω is set to zero.

2. Normalized on velocity.

4.2.4 The Stability of the Cointegrating Relationships

As a first step in establishing how robust the result of the previous section to sample selection, we evaluated the cointegration test statistics for the 1870-1914 sample as well as for the entire available sample conditional upon assuming exogenous shocks stemming from the two world wars, the Great Depression, and the two oil price shocks. Neither of these considerations affected the finding of a unique cointegrating vector for velocity. When tests on permanent income are conditional on the aforementioned exogenous shocks we are unable to reject the null of zero cointegrating vectors. Thus, in the absence of these shocks there is apparently no convergence in permanent income in the countries considered.[27] It may be that one consequence of these shocks is that the exchange rate regime, in particular a fixed or quasi-fixed exchange rate regime, is conducive to stimulating convergence in permanent income levels. Financial history does suggest the adoption of pegged exchange rates after the two wars (i.e., return to the Gold Standard and Bretton Woods). The only other major departure from the results in Table 2 is the finding of a unique cointegrating vector for the *lnal* variable which represents a proxy for financial development. This result reaffirms the view that, conditional on the postulated exogenous shocks, the process of urbanization is a common development in the countries under study.

A criticism of the above approach to testing stability is, of course, that the selection of the sample may be ad hoc. Another alternative has been to implement Chow type tests to determine the stability of coefficients in regression analysis. As with sub-sample estimation, however, the test requires that the timing of the break in a relationship be known a priori. This problem could lead to incorrect inferences being made as in, for example, the recently developed unit root tests applied to each of the series considered in this paper which found a break in the series in years other than the

[27] We were unable to test for cointegration between permanent income and the *lnal* variable for the 1870-1914 period due to insufficient data.

ones arbitrarily selected by Bordo and Jonung (1987). Consequently, it would seem preferable to test the stability of any cointegrating relationship which is not subject to any ad hoc selection of samples.

Gregory and Hansen (1992) 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

$$y_{1t} = \mu + \theta y_{2t} + e_t \tag{4.1}$$

One alternative hypothesis[28] is written

$$y_{1t} = \mu_1 + \mu_2\, \varphi_{t\tau} + \theta_1\, y_{2t} + \theta_2\, y_{2t}\, \varphi_{t\tau} + e_t \tag{4.2}$$

Equation (4.2) is augmented with a change in intercept (μ_2) and a change in the slope. φ is a dummy variable defined as

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

where n is the number of observations, and where φ is created for each possible break-point τ.[29] The sequence of residuals, $\hat{e}_{t\tau}$, can then be analyzed in the same manner as the test for cointegration proposed by Engle and Granger (1987), that is, by generating an augmented Dickey-Fuller (ADF) statistic for each τ. We analyze the statistical stability of the cointegrating relationships summarized in (2.2) to (2.4). BJ (1987, Table A.2) report that in their preferred velocity specification for individual countries, which include proxies for institutional change, only Canada fails to pass the no-constancy hypothesis when the sample is split at the point where velocity begins to rise. The results of implementing the Gregory-Hansen test are shown in Figure 3 which plots the sequence of ADF statistics for the cointegrating regressions (2.2) and (2.3). There is no apparent instability in any of the cointegrating relationships considered. Thus, the equilibrium relationship describing velocity,

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[28] Gregory and Hansen (1992) consider several other alternative cointegrating regressions which are, with one exception, nested in the specification considered below.

[29] Following previous convention in this literature, 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. See Hansen (1992).

permanent income, and the interest rate across countries does not appear to be subject to a regime shift.[30] Although there is a peak in the ADF statistic for velocity in 1927 it is not statistically significant even at the 10% level of significance.[31]

5. Conclusions

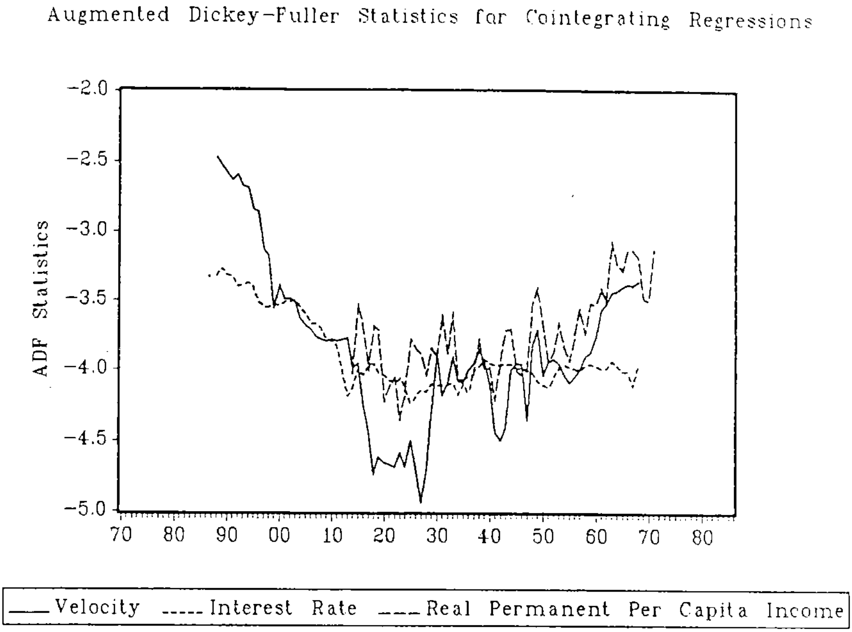
This paper has utilized the econometric techniques of cointegration and error correction to investigate whether institutional factors represent a common element in explaining the long-run behaviour of velocity in five industrialized countries. Relying on recent work which suggests that institutional factors are important determinants of velocity's behaviour 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 behaviour of velocity.

[30] The interest rate test must be interpreted with some caution because, according to Table 2, there are two cointegrating vectors for R. The Gregory-Hansen approach, by relying on the Engle-Granger (1987) methodology, implicitly assumes the existence of a unique cointegrating relationship.

[31] It is worth noting that when Chow tests are generated from the VARs estimated recursively parameter stability is rejected for velocity in 1945, and in 1976 for the interest rate. No parameter instability was found for real per capita permanent income. Hansen (1991) discusses some of the problems with Chow tests based on recursive estimation.

Figure 3

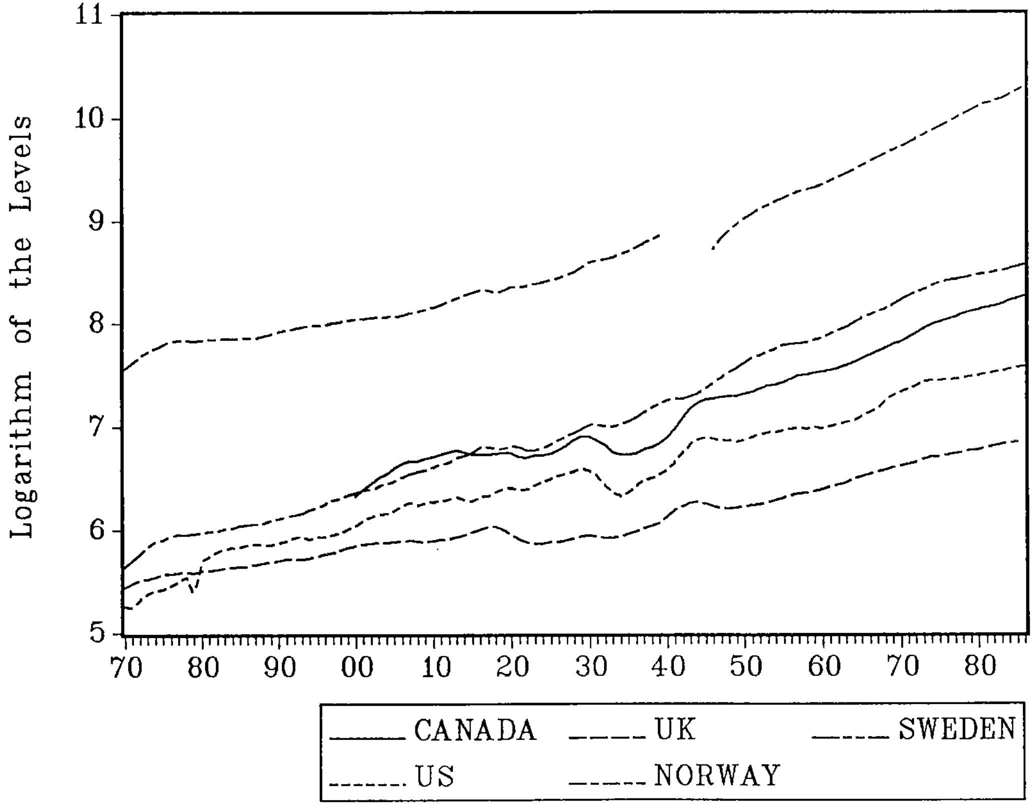
<u>**Test for Stability in Cointegrated Relationships***</u>



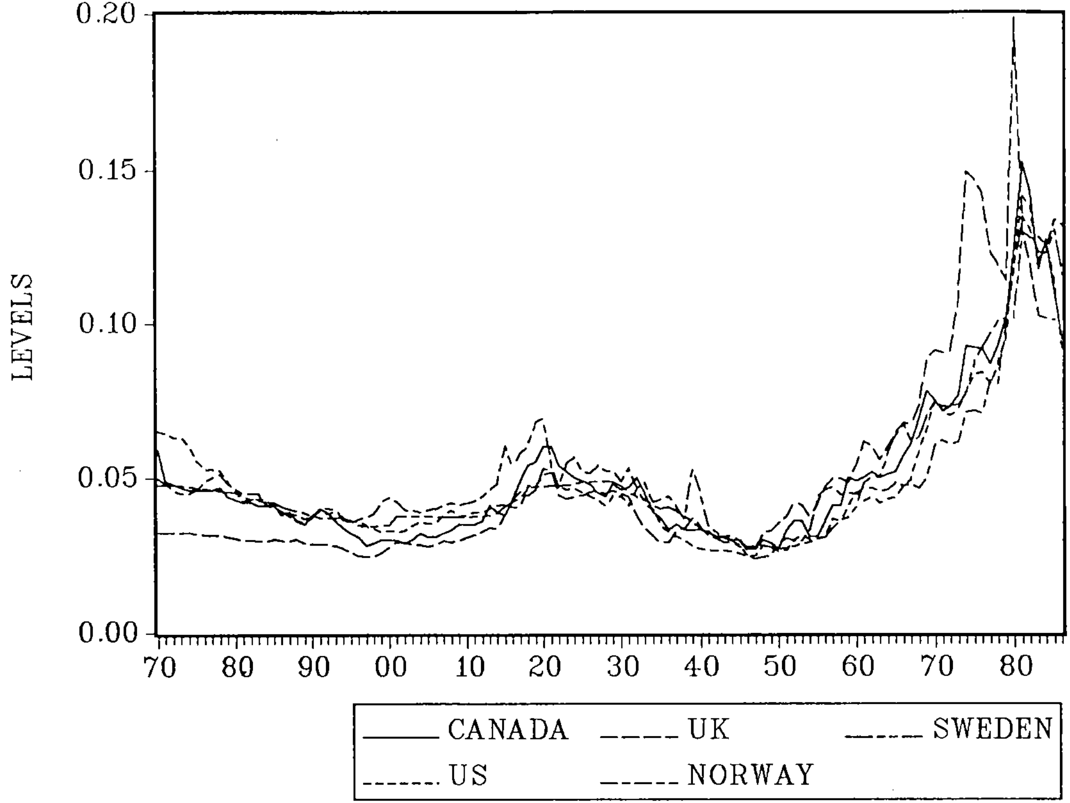
* Based on Gregory and Hansen (1992). The null is the standard cointegrating regression as defined in the text. The alternative is labelled the regime shift model. The test is applied to the largest ADF statistic (in absolute values). The critical value is -5.75 at the 10% level (see Table 1C in Gregory and Hansen, 1992).

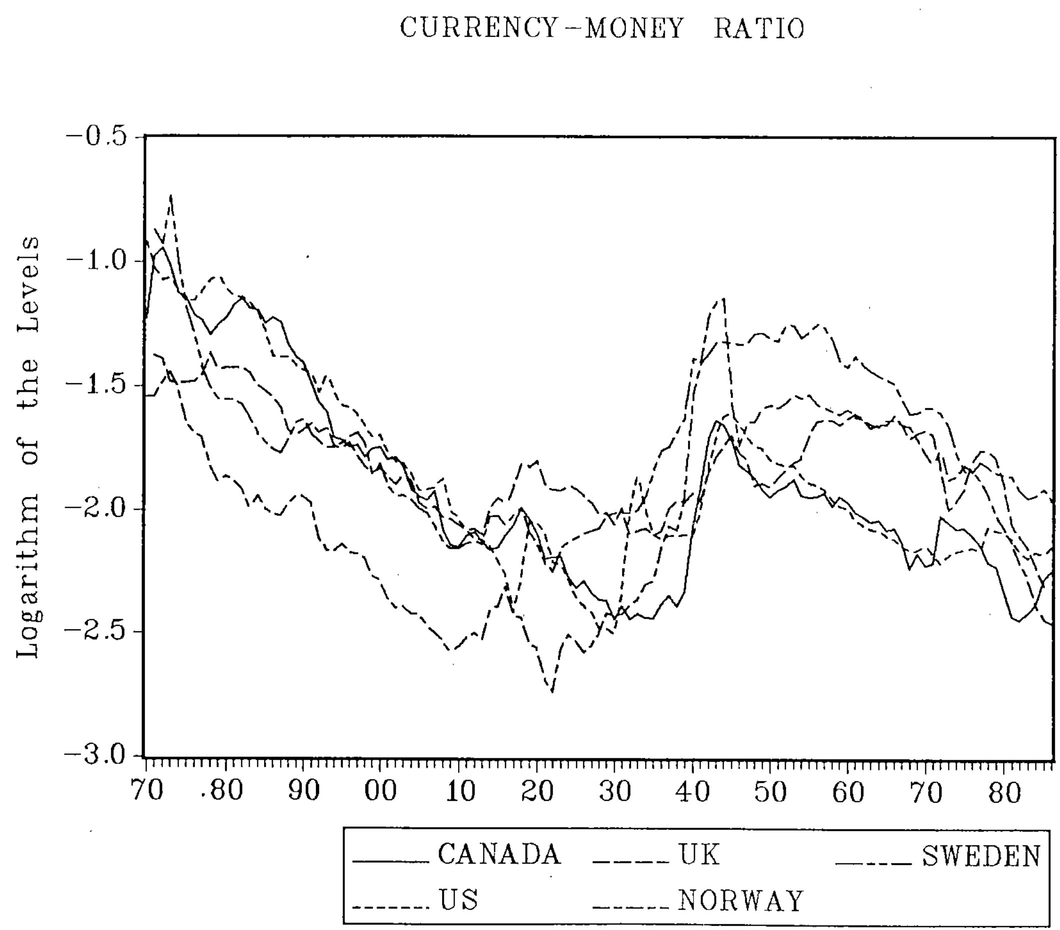
Appendix

REAL PER CAPITA PERMANENT INCOME



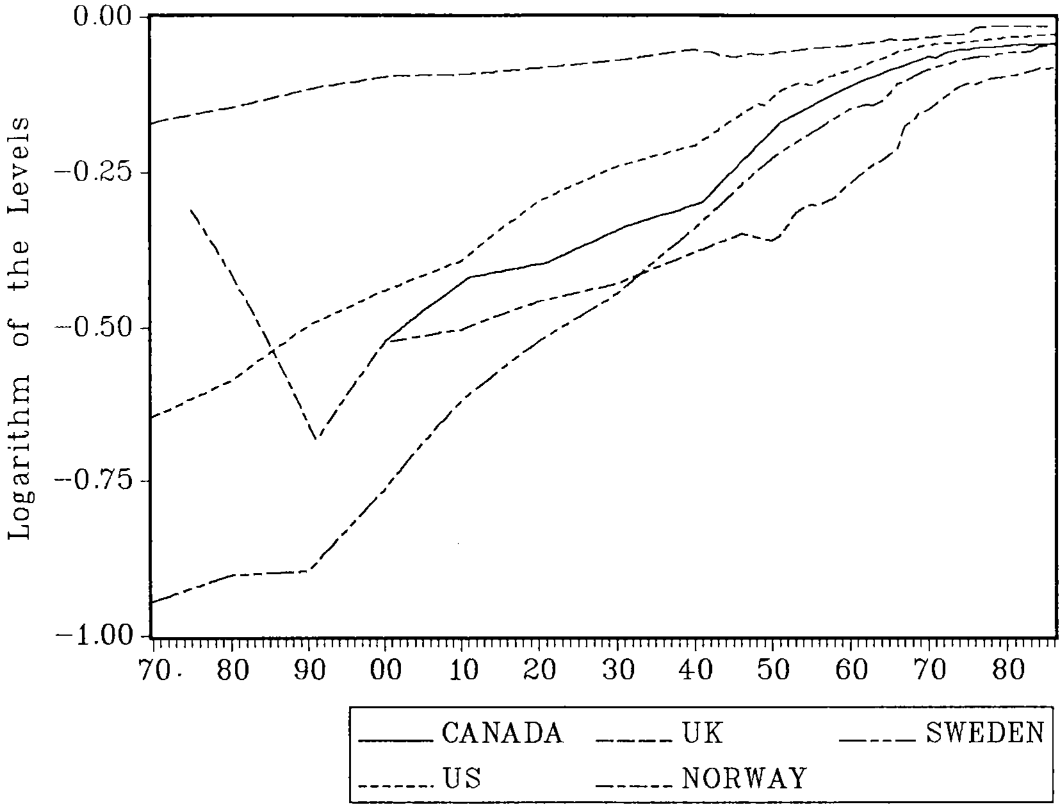
LONG-TERM INTEREST RATE





CURRENCY-MONEY RATIO
Logarithm of the Levels
-0.5
-1.0
-1.5
-2.0
-2.5
-3.0
70
.80
90
00
10
20
30
40
50
60
70
80
CANADA
UK
SWEDEN
US
NORWAY

NON-AGRICULTURAL LABOUR FORCE as a Percent of the LABOUR FORCE



NON-FINANCIAL ASSETS to TOTAL FINANCIAL ASSETS

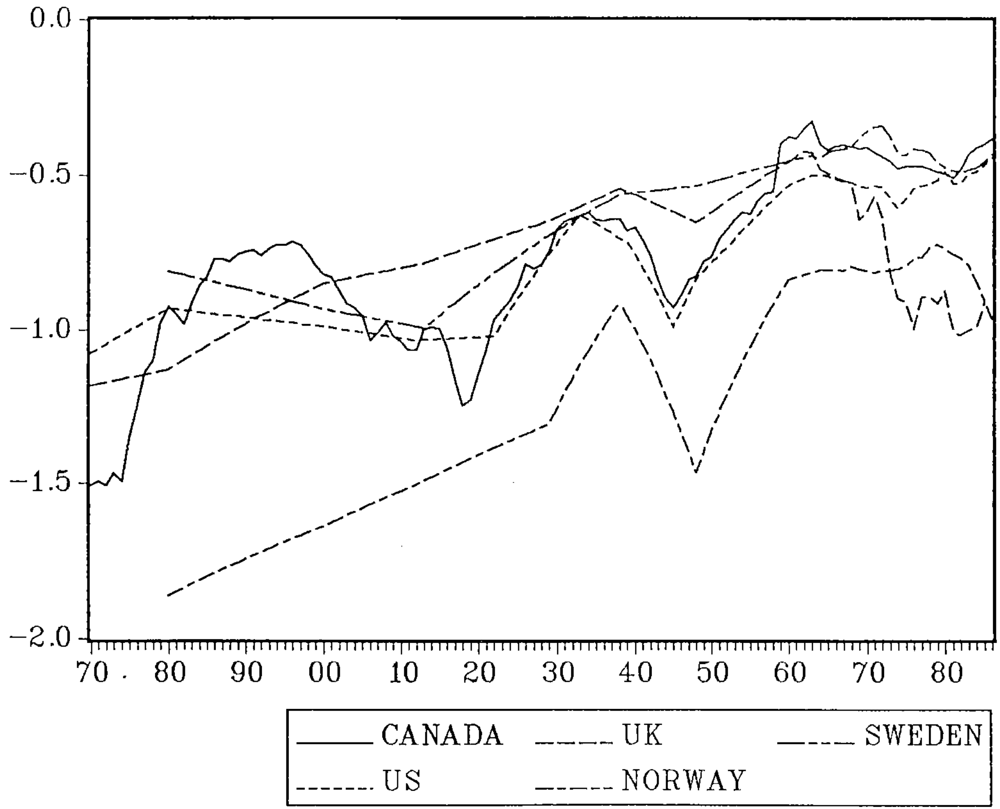


Table A1

Further Cointegration Test Statistics[1]

(A) Canada - U.S. - U.K. - Sweden - Norway: 1870-1914

Series	Number of Cointegrating Vectors					Lag Length	Sample
	0	1	2	3	4		
v	74.90*	39.23	15.19	2.98	.16	2	1870-1914
y^P			@				
R	73.59*	40.18	21.23	10.15	3.77	2	1870-1914
$\dot{C}$/M	61.81*	38.04	17.46	7.94	0.10	2	1871-1914
lnal			@				
NBFA/FA	73.97*	43.96*	23.32	9.12	1.89	2	1880-1914

Notes: @ insufficient data to test.

1. See Table 2 for additional details about the test procedure and a description of the null and alternative hypotheses.

(B) Canada - U.S. - U.K. - Sweden: 1870-1986[2]

Series	Number of Cointegrating Vectors				Lag Length	Sample
	0	1	2	3		
v	47.79*	19.59	7.59	2.92	2	Full
y^P	36.84	16.71	6.31	.75	2	Full
R	62.62*	29.15*	13.40*	.34	2	Full
C/M	52.62*	14.74	4.73	.73	2	Full
lnal	56.84*	24.00	13.07	2.62	2	Full
NBFA/FA	46.75*	26.92*	11.15	1.65	2	Full

2. Conditional on separate dummies for the two world wars (1914-19, 1939-41), the Depression (1929-32), and the two oil price shocks (1973-74, 1978-79).

Table A2

Further Pooled Cointegration Test Results[@]

(A) Conventional Velocity Model (excluding Norway)

Model[1]	Number of Cointegrating Vectors			Lag Length
(2.1)	0	1	2	
Statistic	44.92*	18.21	4.87	8

Cointegrating Vector: [v, y^P, R] = [1, .201, -62.553]

(B) Institutionalist Model (excluding Norway)

Model	Numbers of Cointegrating Vectors						Lag Length
(2.1)	0	1	2	3	4	5	
Statistic	153.20*	98.38*	56.75*	36.22*	16.31*	1.48	7

Cointegrating Vector: [v, y^P, R, C/M, *lnal*, NBFA/FA] = [1, 1.39, 3.65, -19.49, -14.62, -4.21]

@ See notes to Tables 2 and 4.

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