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INSTITUTIONAL HYPOTHESIS OF THE LONG-RUN INCOME VELOCITY OF MONEY AND PARAMETER STABILITY OF THE EQUILIBRIUM RELATIONSHIP

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SUMMARY

It has recently been argued that when the conventional specification of M2 income velocity is extended to include proxies for two types of institutional change, as emphasized by Bordo and Jonung (1987,1990), corresponding to the processes of monetization and increasing financial sophistication of financial developments, this extended model is stable in the sense that one can reject the null hypothesis of no cointegration against the alternative of a single cointegrating vector. There may be implications that such an equilibrium relation is a structural income velocity of money function. The evidence based on century-long data from 1880 to 1986 presented in this paper about parameter instability of the cointegrating vector of velocity with its determinants for Canada, Norway, Sweden, and the United Kingdom casts doubt on this interpretation. The evidence is based on using formal stability tests. Moreover, it has an 'eyeball' support from the sequential estimates of various parameters of the cointegrating relationship including income and interest semi-elasticities.

1. INTRODUCTION

The proposition expounded by Milton Friedman (1959), that the long-run demand for money (or its rearrangement as a velocity) is a stable function of a small set of variables, has been subjected to considerable scrutiny in recent years, often using new techniques of econometric investigation and new multi-country century-long data. A partial list of recent contributions includes Lucas (1988), Friedman and Schwartz (1991), Hendry and Ericsson (1991), Bordo and Jonung (1987, 1990), Barnett *et al.* (1992) and Stock and Watson (1993). These contributions have generated further discussion and comments, which are briefly reviewed later in this paper. The renewed intellectual interest in this issue appears to be motivated both by the desire to account for the perceived causes of apparent instability or predictive failure of the conventional specification of the income velocity of money, and to seek a sound basis for monetary policy. This issue is quite important from the standpoint of policy since, increasingly, the burden of macroeconomic policy appears to have shifted to monetary policy tools, given that existing budget imbalances faced by policy makers in several countries have robbed them of available fiscal instruments (cf. Friedman, 1988; Laidler, 1990,: p. 143).

A popular modelling strategy in recent years has been to treat variables in the demand for money function as jointly endogenous and specify them as a reduced-form vector autoregression error-correction model. The question of interest in this framework is to find out if a unique (or a multiple cointegrating) equilibrium relationship exists (cf. Granger, 1981; Engle and Granger, 1987) among them (cf. Hafer and Jansen, 1991; Hoffman and

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Rasche, 1991; Siklos, 1993, among others). The answer to this question can be obtained by, for example, testing for a rank condition in a matrix of parameters of the error correction term in the reduced-form VAR model, which is taken as the data-generation process. Additionally, alternative structural hypotheses in the form of linear restrictions on either the cointegrating vector(s) or the weights (or basis) vector are often performed to determine if these behavioural restrictions are supported by the cointegrating space for a given information set corresponding to a given behavioural hypothesis. The objective then is to identify a structural relationship corresponding to the cointegrating vector(s). One popular approach for using this strategy is to apply the maximum likelihood-based testing procedure developed by Johansen (1988, 1991) and Johansen and Juselius (1990). The use of this procedure permits determination of the number of cointegrating vectors in the information set as well as tests for linear restrictions on the cointegrating parameter space, including the test for weak exogeneity of the variables to yield a parsimonious behavioural model consistent with the data.

In this paper we investigate the stability property of the unique cointegration or equilibrium relation for log of income velocity of M2 money with log of income, interest rate, log of currency–money ratio (*cm*), and the log of ratio of financial assets to total assets (*pa*) in Canada, Norway, Sweden, and the United Kingdom using the framework in Hansen (1992). Such a relation was recently reported by Siklos (1993) for these countries using annual data from 1880 to 1986. Hansen's framework tests the null of cointegration in a set of variables against the alternatives of no cointegration (H_1), no cointegration with an episodic structural break at an unknown point in time (H_1') and no cointegration with a gradual shift in parameter according to random walk (H_1''), based on estimates of the cointegrating vector obtained by the fully modified (FM) asumptotically efficient method of Phillips and Hansen (1990). The first configuration yields the Lagrange Multiplier test of the null of cointegration against the alternative of no cointegration, and the remaining two configurations yield the tests of parameter stability. This investigation is motivated by the theoretical arguments in Hamilton (1989a) and Hallman *et al.* (1991) that the use of *cm* and *fa* proxy variables for institutional change make it difficult to interpret the velocity relationship as a structural relationship, since the use of these proxies amount to circularity in definition. The analysis is also motivated by the empirical result in Hallman *et al.* (1991) showing that income velocity of money for M2 is cointegrated with the fraction of labour force not employed in agriculture as a proxy for institutional change for the USA; their results are also shown to hold for other countries by Raj (1994). These results indicate that there is no unique way to obtain a cointegrating relationship between velocity and institutional variables. Thus, our prior belief is that the equilibrium relation for income velocity in this paper is unlikely to have a structural interpretation. One of the conclusions of this paper is that while the null of cointegration against H_1 is generally not rejected, we are able to reject the null against both H_1'; and H_1'' lending support to our prior belief.

The rest of the paper is organized as follows. Section 2 contains a brief review of Bordo and Jonung's Institutional Hypothesis. Section 3 briefly describes the alternative tests of parameter stability from Hansen (1992) for the I(1) regressors model. Besides performing these formal tests, we have also undertaken another type of analysis whereby sequential estimation of the cointegration relationship is carried out using an asymptotically efficient method starting from a subset of observations and progressively enlarging the sample one observation at a time. Section 4 presents our empirical results. The final section contains some conclusions.

2. INSTITUTIONAL HYPOTHESIS OF THE LONG-RUN VELOCITY BEHAVIOUR

Background and General Discussion of Issues

Bordo and Jonung's (1987) historical study of the long-run behaviour of velocity is unique in that it looks into the nature and causes of a U-shaped secular trend in velocity on a century-long scale for a number of countries. This study stresses the importance of interpreting the effects of financial innovations in the past 15–20 years as possible symptoms of the instability of the velocity function in a longer-run context as opposed to recent episodes of institutional change that have received considerable attention. Their thesis is that the observed secular pattern of velocity can be interpreted more effectively in terms of the evolutionary technical progress taking place in the financial sector of the economy over the long run rather than a few episodic changes on which others have focused. Furthermore, the financial evolutionary changes they outlined are hypothesized to influence velocity in two competing ways, each dominant at two different stages of the secular process of industrialization of a particular country.

In the first phase, increasing monetization of the economy takes place such that velocity is expected to decline, since demand for money (or transaction balances) grows faster than income. Bordo and Jonung go on to define the process of monetization as the proliferation of commercial banking and the spread of a money economy. The currency-to-money ratio proxy is used by Bordo and Jonung to measure the degree of monetization.

The second phase of secular industrialization, roughly represented by the post-war period, is associated with increasing financial sophistication such that the velocity is expected to rise as methods are developed to economize on cash balances and an active search for money substitutes takes place. The ratio of total non-bank financial assets to total financial assets proxy is used by Bordo and Jonung to measure the degree of financial development during the second phase. This is, in essence, Bordo and Jonung's Institutional Hypothesis about the U-shaped secular trend of velocity.

Theoretical Underpinnings

Theoretical support for Bordo and Jonung's hypothesis can be obtained from traditional theories of long-run velocity (cf. Laidler. 1985; Ireland, 1991, Section 2]) as well as from the use of a representative agent subject to a cash-in-advance constraint model developed by Lucas and Stokey (1987) and Lucas (1987) after a slight modification to account for an exogenous change in payment technology as shown by Hamilton (1989a, pp. 337–342). Moreover, antecedents of the Institutional Hypothesis can be traced back to the pioneering work of well-known institutional economists such as Wicksell (1936, Chapter 6, Section c), and Fisher (1963, Chapter v, Section 3) among others. It can also be seen as a generalization of the Friedman and Schwartz (1982) monetary trends study in which the importance of the role of financial sophistication is anticipated but modelled as an intervention variable (i.e. as episodic change) causing level shifts in the demand for money function with income and interest rate as two key determinants. However, it is important to point out that Bordo and Jonung's hypothesis differs from others in that they see institutional changes affecting velocity in an on-going manner.

Critical Review of Existing Empirical Evidence

Bordo and Jonung (1987) provide empirical support employing three quite different methodologies such as regressions, a detailed case study of a particular country, and a cross-section analysis. Accordingly, a variety of independent bases of empirical evidence are available. Nevertheless, this evidence raises a number of methodological and econometric issues relating to the measurement, identification, integrativeness and endogeneity of the variables involved. These issues have been raised by a number of researchers, such as Raj and Siklos (1988), Muscatelli (1988), and Hamilton (1989a).

While Bordo and Jonung (1990) updated their earlier work based on time-series data by adding more recent data, and employed econometric techniques less prone to earlier criticisms than those used in their study, some of the fundamental issues pertaining to endogeneity and leaving out long-run information from their specification in first-difference form continue to remain. Accordingly their conclusions about the support for the hypothesis from time-series data remain puzzling and difficult to interpret for reasons outlined in Hamilton (1989a) and Huizinga (1990).

In view of these shortcomings some researchers have used either a new data set or newer econometric techniques to empirically test for the institutional hypothesis. For example, Ireland (1991) has used US regional data since 1929 and obtained results consistent with Bordo and Jonung's hypothesis. Following Bordo and Jonung's framework, Hallman, *et al.* (1991, Section iv) model M2 velocity for the USA as a cointegration relation with a fraction of the labour force not employed in agriculture and a level-shift dummy. They exclude other variables earlier used by Bordo and Jonung's study from the velocity function on the grounds that their inclusion would impair interpreting the fitted regression as the long-run steady-state value of velocity (Hallman *et al.*, 1991, footnote 26). Theoretical underpinnings and motivation for using this information set is given in their paper.

3. PARAMETER INSTABILITY TESTS WHEN VARIABLES ARE INTEGRATED OF ORDER ONE

It is a standard practice in applied econometrics to perform alternative model adequacy tests such as the test of parameter constancy on regression models. These tests are important since parameter non-constancy is indicative of model misspecification. In a recent paper Hansen (1992) extended some relevant tests for parameter stability for regression models with stationary regressors to regression models with integrated regressors of order 1, using the fully modified (FM) estimation method of Phillips and Hansen (1990). Specifically, he derived the large-sample distributions of Lagrange Multiplier (LM) tests for parameter stability against several alternatives of interest in the context of cointegrated regressions. This test was first proposed by Gardner (1969).

The use of Hansen's parameter stability testing framework which will be briefly described shortly is instructive in the sense that it permits testing for parameter stability as well as for cointegration. This is possible through the use of Hansen's test statistics, L_c, which tests only for martingale variation in the constant of cointegration regression. Thus, this statistic tests the null of cointegration against the alternative of no cointegration. The use of this type of test instead of the usual test of the null of no cointegration may be appealing since its use allows the analyst to control for the probability of rejecting a valid economic model.

A common form of model misspecification encountered frequently in the conventional money demand function with the real income and interest rate as its arguments is concerned with parameter instability. It has recently been claimed that when this conventional specification is extended to include proxies for two types of institutional change, as emphasized by Bordo and Jonung, the extended model is stable in the sense that one can reject the null hypothesis of no cointegration against the alternative of a single cointegrating relationship between velocity and its determinants for a number of countries such as Canada, Norway, Sweden, and the

United Kingdom using Johansen's MLE framework. In our empirical analysis, using the same century-long data as in the previous study for these countries, we will argue that while one cannot generally reject the null hypothesis that velocity and its determinants are cointegrated against the alternative of no cointegration, the cointegrating vector in all countries suffers from a gross form of parameter instability. This argument is developed by employing the formal tests for parameter instability recently developed by Hansen (1992) in regressions with I(1) processes using the efficient FM estimator of the cointegrating vector. In addition to using the evidence from formal tests of parameter stability, we trace out how FM estimates of key parameters of the cointegrating relation change as the sample period is varied to provide eyeball support from the formal test statistics.

The use of the FM estimator is motivated by the availability of asymptotic theory of parameter instability for this method, as well as the large sample critical values for the asymptotic distributions of these tests. The semi-parametric FM single-equation estimator of the unique cointegrating vector is asymptotically equivalent to full-information methods such as the maximum likelibood parametric estimator proposed by Johansen (1988, 1991) and Johansen and Juselius (1990). The FM method corrects for the finite-sample bias in the OLS estimator of the static cointegration regressions in order to obtain median-unbiased and asymptotically normal estimates of the parameters of long-run parameters of the cointegrating vector. Specifically this method modifies the least-squares estimates of the static equilibrium regression in two ways. First, it accounts for both the serial correlation in the residuals as well as mitigating the effects of the second-order bias of the least-squares estimator of the static regression. Second, it corrects for the long-run simultaneous-equation bias to permit the use of conventional asymptotic procedure for inference in the estimated regression. In summary, the use of the FM method not only achieves reduction of second-order bias of the least-squares estimates but also simplifies the statistical inference by correcting standard errors of coefficients in the cointegrating regression.

It may be pointed out that although in principle the parameter stability tests, described below using the Phillips–Hansen FM estimator, can easily be obtained for any other asymptotically efficient method of estimates of the cointegrating vector such as the MLE estimator of Johansen or 'the leads and lags' estimator of Stock and Watson (1993). However, this line of enquiry is not pursued due mainly to the unavailability of the asymptotic theory for such tests.

Hansen proposed a total of two tests of parameter stability. The null hypothesis for each of these two tests is the same; that is, there exists a unique cointegration among variables with stable parameters. The first test of parameter instability specifies the alternative of a structural change in parameters, possibly resulting from an unspecified episodic event or events. This test is in the spirit of Quandt (1960), who proposed specifying the alternative hypothesis for parameter instability as a single structural break of unknown timing. Specifically, the test labelled as the Sup$F$ test, is computed as

$$\mathrm{Sup}F = \operatorname*{Sup}_{t/n \in J} F_{nt},$$

where F_{nt} is the F-test statistic corresponding to the classical Chow (1960) test or the spilt sample test for a fixed t. A word of caution about the use of this test statistic is provided by Hansen. If it is found that this test rejects parameter stability, then Hansen believes that one should not interpret, based on this evidence alone, that there are two cointegrating regimes. The reasons why such a conclusion may be inappropriate is that the rejection of the null hypothesis cannot be taken to imply acceptance of the alternative hypothesis. The only statistically justifiable conclusion one can draw from the rejection of the null hypothesis is that the null of parameters constancy of the long-run parameters of the equilibrium regression is rejected by the data at the chosen significance level. The second test statistic is the Mean$F$, which is suitable for testing the null of cointegration with constant parameters against the alternative of a gradual shift in the parameters. This test statistic is computed simply as the average value of the F_{nt} statistics defined before, and is computed for $t/n \in J$.

The finite-sample performance of the size and power of Hansen's tests of parameter instability in regressions with I(1) processes, using the FM estimation method is available from the Monte Carlo study of Gregory and Nason (1993). They provide evidence suggesting that the size and power of these tests are reasonable, although their evidence corresponds to systems of size smaller than those being used in this paper. To partially readdress this potential lack of knowledge of finite sample properties for the large system used in this paper, we report evidence from a variety of sensitivity analyses outlined below. First, stability testing is performed using the FM estimates with a variety of kernels. In addition we have obtained estimates of the cointegrating vector corresponding to a variety of asymptotically efficient methods such as the single-equation 'leads and lags' estimation methods; the estimates of the cointegrating vector from these methods will be compared and contrasted with those estimates from Johansen's method. The analysis is performed corresponding to two model specifications: the first model postulates that a deterministic cointegration restriction holds for the cointegrating vector. In other words, it is postulated that the cointegrating vector eliminates both the stochastic and deterministic trends from the velocity and its determinants. The second model postulates that the variables in the velocity function are stochastically cointegrated (for further details consult a recent survey paper by Ogaki, 1993).[1]

4. EMPIRICAL RESULTS

The FM estimates[2] of the cointegrating vector, along with various test statistics, are reported below. The cointegrating relation corresponds to the log of M2 velocity (v); the log of real permanent per capita income (y^P); the interest rate (r), which is measured in per-cent as (say) 0·05 for an interest rate of 5%; the log of currency–money ratio (cm) and log of the ratio of non-bank financial to total financial assets (pa). The data consist of annual observations on five variables listed above for the period from 1870 to 1986. Specifically, the data for Canada are for the period from 1900 to 1986; Norway from 1880 to 1986, excluding observations for the Second World War period from 1939 to 1945 during which the country was under German occupation; Sweden from 1880 to 1986; the United Kingdom from 1876 to 1985. The data are drawn from studies by Bordo and Jonung (1987, Appendix IA; 1990, Appendix I).[3] A brief description of the data is provided in the Appendix to this paper.

The results reported below are based on two reasonable model specifications for the income velocity of M2 money. Model I for the velocity has an intercept term in it. This model specifies

[1] A commonly specified data generation process (DGP) for the vector of variables for the velocity function allows for both a linear deterministic trend and a noise component; the noise component is specified to contain elements with unit roots as well as stationary components. Following Johansen's (1988) notation, this DGP can be written as a vector error-correction VAR system with an intercept, with a certain restriction (for example, see Campbell and Perron, 1991). If this restriction is not supported by the data, then the cointegrating vectors among the variables in the velocity function that annihilate the nonstationarity in the noise component will not annihilate the nonstationarity in the trend component. Accordingly, when this condition does not hold, Model II is more appropriate to use as Model I is misspecified. The use of linear trend in the error-correction VAR model has implications for cointegration testing and causality as shown by Perron and Campbell (1993). However, the use of Model II has a potential drawback in that the presence of trends in estimated autoregressions creates a downward bias of the long-run parameters as shown by Andrews (1991a) when the sample size is not large.

[2] These calculations were performed using the computer code supplied by Bruce Hansen.

[3] The data were supplied by Pierre L. Siklos.

that the velocity and its determinants are deterministically cointegrated, using the terminology of Ogaki and Park (1990). Model II for the velocity includes an intercept and a time trend in the level regression. As pointed out earlier, this model specifies that velocity and its determinants are stochastically cointegrated. The FM method uses a kernel estimator with the plug-in bandwidth recommended by Andrews and Monahan (1992) for the estimates of the parameters of the cointegrated vector. The t-statistics are also computed using similar modifications (cf. Phillips and Hansen, 1990). It is generally contended that the choice of a kernel for the FM method is less important than the choice of a bandwidth parameter. However, in a recent study by Haug (1994), who undertook a Monte Carlo study on the small sample performance of the alternative tests for cointegration, it was argued that the choice of kernel is quite important. In view of the conflicting arguments about the importance of choice of kernel, we reported alternative parameter stability test statistics for the Quadratic Spectral (QS) and Bartlett (B) kernels or weighting functions. Also we have reported alternative parameter stability test statistics along with the associated p-values of these test statistics in Table I. The estimate of the plug-in bandwidth parameter, labelled as M, corresponding to the choice of alternative kernels is also reported. The values of test statistics are computed over the trimmed region $(0{\cdot}15n,\ 0{\cdot}85n)$ following the recommendation of Andrews (1991b) and Andrews and Monahan (1992); the symbol n in the trimming region represents the sample size. A low p-value, below 0·10 (say), for a particular test statistic is interpreted as the instability of the parameters of the cointegrating vector; that is, the null hypothesis of cointegration is rejected against a specified alternative hypothesis.

Two important conclusions emerge from these results. First, test results based on alternative L_c statistics and the associated p-values show that one cannot generally reject the null hypothesis that velocity is cointegrated. This is especially true if the Bartlett kernel is used to compute the PM estimates and Model II is viewed as the appropriate characterization of the data-generation process. However, in general, the null of cointegration cannot be rejected for Canada, Norway, or Sweden irrespective of the choice model or the kernel used. Also, non-rejection of the null of cointegration for the United Kingdom is sensitive to choice of both the kernel and the model used for the PM estimation. The last result can be interpreted to imply that the cointegration result for the United Kingdom is fragile. Another important result is that while the velocity and its determinants are cointegrated, the parameters of the cointegrated vector are strongly unstable for all four countries irrespective of the model used or the kernel used in computing the estimates based on the efficient FM method. This conclusion is based on the results from the use of both Sup$F$ and Mean$F$ tests and their associated p-values. Later we provide informal evidence based on plots of sequential estimates of the cointegrating vector to provide eyeball support for this conclusion.

The sequence of F-statistics as defined by Hansen (1992) for structural change for the four countries corresponding to the FM estimates with the QS kernel for Model I are plotted in Figures 1 to 4 for Canada, Norway, Sweden, and the United Kingdom. The horizontal axis on all the graphs have the calendar date, except for Norway, for which the calendar date is replaced by observations from 1 to 100. Similar plots (not shown here) are obtained for the choice of the other kernel and the model II. These figures also display the 5% critical values for its largest value (Sup$F$), its average value (Mean$F$), and for a fixed known break point. The sequential F-statistic largely depicts the U-shaped pattern similar to those of the velocity function for each of the four countries. Moreover, plots of the sequence of F-statistics suggest the nature of the instability are on-going rather than short-lived, since the Mean$F$ critical value lies below the F-statistic plot in almost all cases. Similarly the *F-statistic in each case exceeds the Sup$F$* critical value for each of the four countries in all years instead of during a few economic or political episodic shocks. This

Table I. Lagrange multiplier test for the null of cointegration and parameter stability test statistics

Model I: $v = \mu + \Theta_y y_t^p + \Theta_r r_t + \Theta_{cm} cm_t + \Theta_{fa} fa_t + e_t$

Model II: $v = \mu + \Theta t + \Theta_y y_t^p + \Theta_r r_t + \Theta_{cm} cm_t + \Theta_{fa} fa_t + e_t$

Country	Canada: 1900–86			Norway: 1880–1986, excluding 1939–45			Sweden: 1880–1986			United Kingdom: 1876–1985		
Model (Kernel)[a]	Alternative test statistic			Alternative test statistic			Alternative test statistic			Alternative test statistic		
	L_c	Sup$F$	Mean$F$	L_c	Sup$F$	Mean$F$	L_c	Sup$F$	Mean$F$	L_c	Sup$F$	Mean$F$
Model I (QS)	0·73 (0·11)[b] $M=2·55$	$1·45 \times 10^2$ (0·01) $M=2·55$	$2·86 \times 10^1$ (0·01) $M=2·55$	0·54 (0·20) $M=1·68$	$9·60 \times 10^1$ (0·01) $M=1·68$	$4·13 \times 10^1$ (0·01) $M=1·68$	0·56 (0·20) $M=2·12$	$1·40 \times 10^3$ (0·01) $M=2·12$	$2·50 \times 10^2$ (0·01) $M=2·12$	1·46 (0·01) $M=3·73$	$1·04 \times 10^2$ (0·01) $M=3·73$	$4·09 \times 10^1$ (0·01) $M=3·73$
Model I (B)	0·71 (0·12) $M=2·55$	$1·37 \times 10^2$ (0·01) $M=2·55$	$2·74 \times 10^1$ (0·01) $M=2·55$	0·54 (0·20) $M=1·56$	$1·10 \times 10^2$ (0·01) $M=1·56$	$4·78 \times 10^1$ (0·01) $M=1·56$	0·49 (0·20) $M=2·10$	$1·25 \times 10^3$ (0·01) $M=2·10$	$2·24 \times 10^2$ (0·01) $M=2·10$	0·89 (0 05) $M=4·17$	$6·54 \times 10^1$ (0·01) $M=4·17$	$2·46 \times 10^1$ (0·01) $M=4·17$
Model II (QS)	0·80 (0·13) $M=2·55$	$9·14 \times 10^1$ (0·01) $M=2·55$	$2·99 \times 10^1$ (0·01) $M=2·55$	0·58 (0·20) $M=2·03$	$1·32 \times 10^4$ (0·01) $M=2·03$	$1·76 \times 10^3$ (0·01) $M=2·03$	0·15 (0·20) $M=2·57$	$2·26 \times 10^2$ (0·01) $M=2·57$	$2·02 \times 10^1$ (0·01) $M=2·57$	1·17 (0 03) $M=3·97$	$1·01 \times 10^2$ (0·01) $M=3·97$	$3·28 \times 10^1$ (0·01) $M=3·97$
Model II (B)	0·76 (0·15) $M=2·50$	$8·37 \times 10^1$ (0·01) $M=2·50$	$2·78 \times 10^1$ (0·01) $M=2·50$	0·6 (0·20) $M=1·99$	$1·55 \times 10^4$ (0·01) $M=1·99$	$2·07 \times 10^3$ (0·01) $M=1·99$	0·16 (0·20) $M=2·64$	$6·26 \times 10^2$ (0·01) $M=2·64$	$4·98 \times 10^1$ (0·01) $M=2·64$	0·77 (0·15) $M=4·45$	$8·10 \times 10^1$ (0·01) $M=4·45$	$2·44 \times 10^1$ (0·01) $M=4·45$

[a] QS = Quadratic Spectral kernel; B = Bartlett kernel.

[b] A p-value of '0·10' means greater than or equal to 0·10. A p-value below 0·10 is interpreted as the null is rejected. These p-values are calculated as outlined in Hansen (1992).

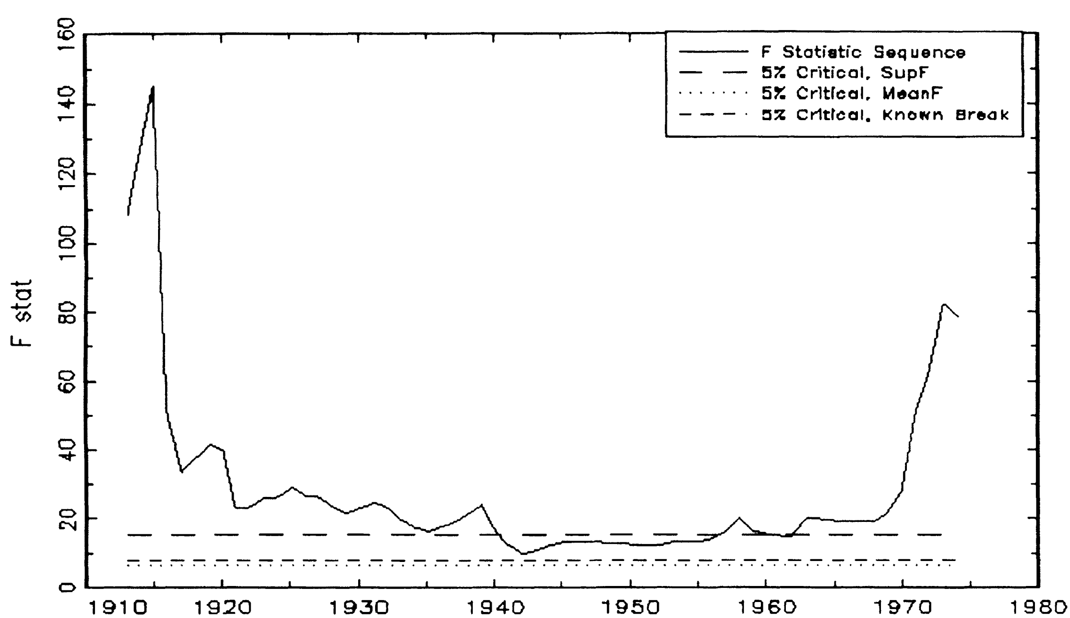

Figure 1. The F-statistic sequence for Canada with 5% critical values of Sup$F$ (---), Mean$F$ (...), and known break (___) values

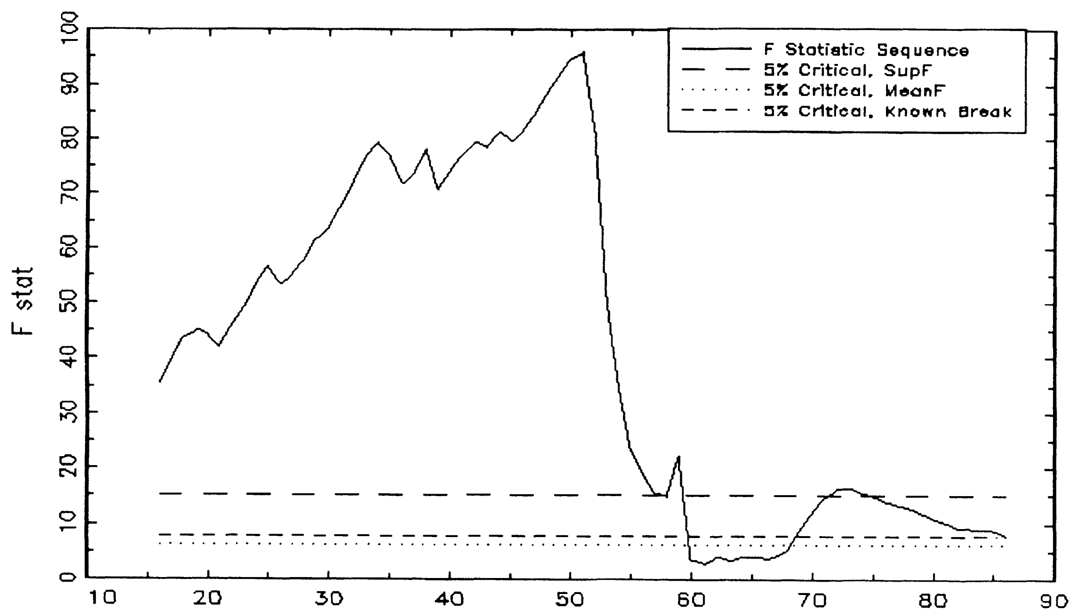

Figure 2. The F-statistic sequence for Norway with 5% critical values of Sup$F$ (---), Mean$F$ (...), and known break (___) values

evidence suggests that the Bordo–Jonung specification for the velocity cannot be repaired for any of the four countries simply by adding a dummy variable or two. One possible way to obtain a stable relationship is to impose some suitable theoretical structure on the demand for money function by imposing restrictions on the income elasticity and interest rate semi-elasticity.

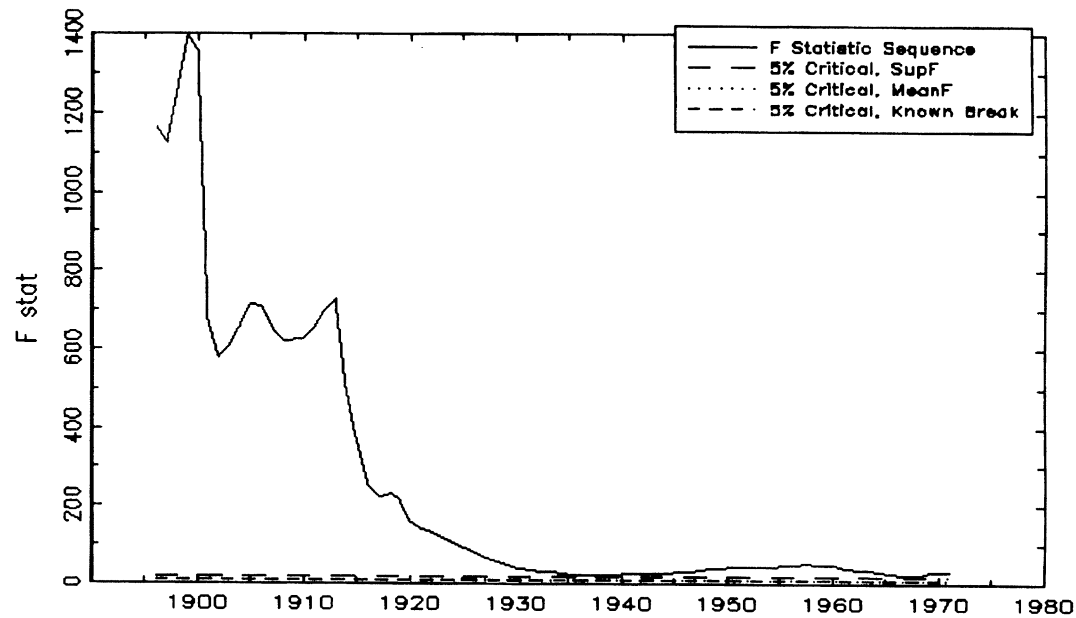

Figure 3. The FM estimate sequence of the income elasticity for Canada plus and minus two standard errors of estimates

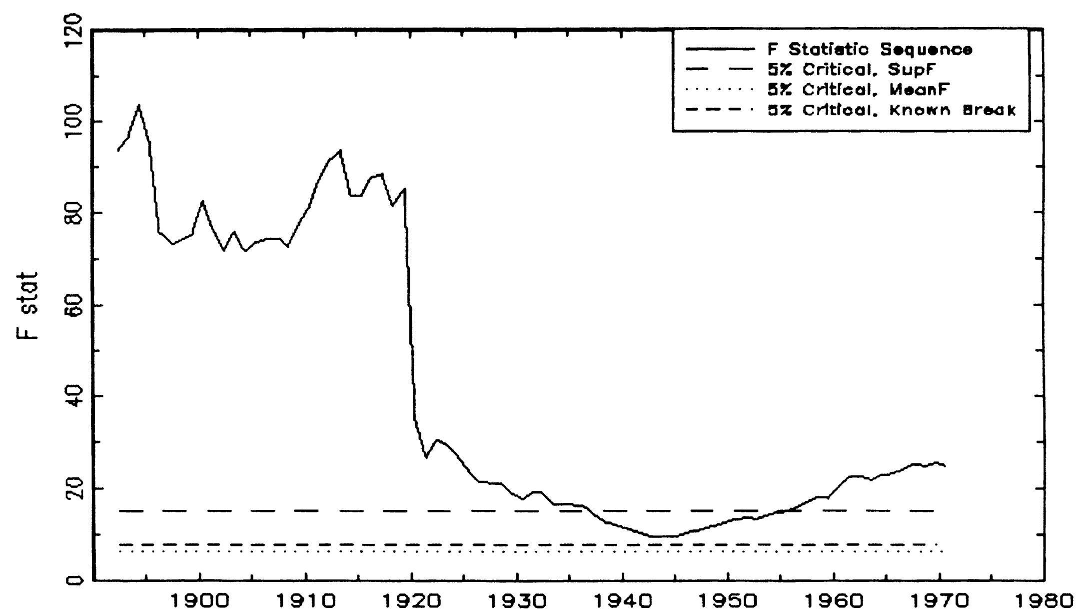

Figure 4. The F-statistic sequence for the United Kingdom with 5% critical values of Sup$F$ (---), Mean$F$ (...), and known break (___) values

The FM estimates of the cointegrating vector (income elasticity and interest rate semi-elasticity and other parameters of the long-run relation) for the two models and different kernels are given in Table II. The following conclusions can be drawn from these estimates. First, the point estimates of elasticities of alternative determinants of velocity and the interest semi-elasticity are fairly

Table II. Estimates of long-run parameters of the cointegrating relation

Estimation method[a]	Model I				Model II			
	θ_{vp}	θ_r	θ_{cm}	θ_{fa}	θ_{vp}	θ_r	θ_{cm}	θ_{fa}
	Canada: 1900–1986							
DOLS	−0·31	6·69	0·50	0·53	0·47	4·24	0·29	0·78
	(0·18)[b]	(2·93)	(0·30)	(0·26)	(0·19)	(1·87)	(0·17)	(0·15)
DGLS	0·03	2·14	0·08	0·23	0·39	3·41	0·20	0·61
	(0·21)	(2·68)	(0·23)	(0·25)	(0·27)	(2·90)	(0·26)	(0·37)
FM(QS)	−0·55	10·01	0·68	0·93	0·31	5·38	0·37	0·93
	(0·17)	(2·45)	(0·23)	(0·25)	(0·20)	(1·61)	(0·15)	(0·17)
FM(B)	−0·53	9·82	0·67	0·90	0·32	5·31	0·37	0·92
	(0·16)	(2·30)	(0·22)	(0·24)	(0·20)	(1·60)	(0·15)	(0·16)
JOH[2]	−1·36	22·58	1·98	1·93	0·16	8·36	0·70	1·15
	Norway: 1880–1986 (excluding Second World War years)							
DOLS	−0·29	7·27	0·82	0·03	1·61	−3·64	0·04	0·41
	(0·50)	(6·17)	(0·27)	(0·64)	(0·18)	(1·64)	(0·09)	(0·15)
DGLS	−0·07	3·75	0·57	−0·18	1·54	−1·74	0·10	0·41
	(0·12)	(1·48)	(0·07)	(0·15)	(0·19)	(1·97)	(0·10)	(0·16)
FM(QS)	−0·08	4·54	0·72	−0·18	1·61	−3·33	0·05	0·42
	(0·33)	(4·80)	(0·26)	(0·48)	(0·16)	(1·50)	(0·08)	(0·11)
FM(B)	−0·07	4·80	0·73	−0·23	1·59	−3·08	0·06	0·42
	(0·32)	(4·71)	(0·26)	(0·47)	(0·15)	(1·48)	(0·08)	(0·11)
JOH[4]	−0·06	−0·11	0·66	−0·24	1·18	−2·84	0·22	0·23
	Sweden: 1880–1986							
DOLS	−0·36	9·39	0·76	0·29	0·94	5·51	0·48	0·78
	(0·08)	(2·02)	(0·16)	(0·35)	(0·31)	(1·58)	(0·12)	(0·27)
DGLS	−0·12	6·87	0·77	−0·24	0·91	6·74	0·57	0·71
	(0·18)	(3·01)	(0·20)	(0·48)	(0·27)	(1·29)	(0·11)	(0·26)
FM(QS)	−0·34	9·45	0·86	0·12	1·05	5·82	0·51	0·63
	(0·08)	(1·44)	(0·13)	(0·33)	(0·30)	(1·54)	(0·14)	(0·33)
FM(B)	−0·34	9·52	0·85	0·13	1·01	6·08	0·53	0·60
	(0·08)	(1·44)	(0·13)	(0·33)	(0·29)	(1·51)	(0·14)	(0·32)
JOH[2]	−0·56	19·59	1·46	−0·49	2·53	4·54	0·42	0·75
	United Kingdom: 1876–1985							
DOLS	−0·27	7·69	0·22	0·14	1·06	7·51	0·10	0·67
	(0·21)	(2·03)	(0·17)	(0·27)	(0·16)	(0·61)	(0·05)	(0·10)
DGLS	−0·06	4·85	0·24	0·04	1·12	7·33	0·09	0·69
	(0·11)	(1·06)	(0·11)	(0·15)	(0·19)	(0·71)	(0·06)	(0·12)
FM(QS)	−0·47	9·76	0·18	0·45	1·18	7·06	0·06	0·82
	(0·13)	(1·35)	(0·12)	(0·17)	(0·22)	(0·80)	(0·08)	(0·13)
FM(B)	−0·43	9·51	0·19	0·38	1·18	7·04	0·07	0·79
	(0·14)	(1·49)	(0·13)	(0·19)	(0·23)	(0·83)	(0·08)	(0·14)
JOH[10]	−0·23	6·19	0·27	0·33	1·46	7·62	0·18	0·97

[a] The DOLS and DGLS estimates were obtained for two leads and lags for the variables on the right-hand side of the velocity model. QS represents the Quadratic Spectral; B represents the Bartlett kernel.
[b] The figures in parentheses are standard errors.

robust to the choice of the kernel used for the FM method within a given model. However, the point estimate of income elasticity is positive in Model II while it is negative in Model I. In other words, long-lun parameter estimates of the velocity function are not robust to the choice of model used for estimation. One interpretation of this lack of robustness of estimates across models is that both models are misspecified, so one should not expect similar point estimates across either models or methods of estimation.

In Table II we have also provided estimates of cointegrating vector using the lead–lag estimators developed by Stock and Watson (1993), among others. These estimates along with those from the FM method of estimation of the cointegrating relationships correspond to the triangular representation of the stochastic process used by Phillips (1991), among others. Specifically, the reported results correspond to the dynamic ordinary least squares (DOLS) and the dynamic generalized least squares (DGLS) methods of estimating of the long-run relationships. The point estimates of income elasticity and interest rate semi-elasticity from the cointegration regression from the DGLS and FM methods are quite similar.

In Table II, we have also given estimates of the cointegrating vector using Johansen's (JOH) maximum likelihood estimation (MLE) method, which is one of the popular methods of estimation and statistical inference for the long-run relationships. These estimates correspond to the lag length for the vector autoregressive (VAR) model specified in the bracket after the symbol JOH in this table; this lag length was chosen by using the multivariate Akaike Information Criterion. This method uses the error-correction representation of the vector process instead of the triangular representation of Phillips (1991). The Monte Carlo evidence obtained by Gonzalo (1994) favours the use of MLE method in view of the fact this method is not only efficient in large samples but it also performs extremely well in finite samples. A more informed view of the performance of the MLE emerges when the exact sample properties of the MLE estimator analysed by Phillips (1994) is taken into consideration. One of Phillips's results is that the MLE estimates can be unreliable in finite samples. An empirical illustration of the Phillips's result for the MLE method of estimation is provided by Stock and Watson (1993).[4] These new set of findings about the performance of the MLE method in finite samples cast doubt about the generality of the recommendation for its use by Gonzalo (1994). A minimum requirement for the MLE to perform adequately in finite samples is that the underlying statistical VAR model for estimation of the long-run relationships is a good description of the underlying data-generation process; this would require that the influence of episodic large shock(s) is adequately modelled before performing statistical inference about the long-run relationships.

In summary, there is no *a priori* reason to expect that the estimates from the MLE method for the velocity models analysed in this paper will be similar to those obtained from the other single-equation methods since some of the requirements to obtaining an adequate description of the data-generation process are not met; specifically the data do contain a number of episodic changes whose influence is not modelled.

The divergence of estimates from the MLE method and the single equation methods in finite samples may occur for a yet another reason. It is known that the estimates of the cointegrating

[4] It can be argued that results based on Stock and Watson's (1993) Monte Carlo (MC) experiments, which are obtained for the models involving a smaller number of parameters as compared to the models used in this paper, do not exactly apply. In view of this it was suggested that we performed a pilot MC study using a model and estimates in this paper to investigate finite sample performance of the MLE estimator relative to the single equation FM estimator. One of the conclusions of this limited investigation was that the FM estimator in some instances performed better than the MLE estimator in simulation experiments. This result is consistent with the MC results obtained by Stock and Watson (1993) but differs from the MC result obtained by Gonzalo (1994). In order to settle this issue satisfactorily, one would need a more extensive set of MC experiments, which we hope to undertake at some future date.

vector obtained from using the MLE method impose matrix normalization differently from the single-equation methods; whereas the triangular single-equation methods impose normalization prior to estimation the MLE method applies normalization after estimation is completed.

The MLE estimates for each of the two models given in Table II are qualitatively similar to those estimates of parameters obtained from several single-equation methods. However, these estimates are sometimes quantitatively different from those from single-equation methods, especially in model I for Canada and Sweden. While in principle the quantitative difference can be attributed to one or more of the possible reasons given above, it is conjectured that the quantitative difference between the MLE and the single-equation methods is indicative of the general inadequacy of the economic model rather than inherent finite sample superiority of a particular method over other methods used for estimation. Finally, the JOH estimates do not exactly replicate Siklos's results; this difference is due to a slight difference in the lag lengths for the VAR models used and/or difference in the starting dates for the sample period used for estimation.

The DGLS estimates use corrections for AR(1) and AR(2) errors in all models and countries with the exception of Norway and United Kingdom, corresponding to model I. For these two cases the corrections required were MA(1) and MA(2), and MA(1) and AR(2). The inference based on the DGLS method for income elasticity, with the exception of the United Kingdom for Model II, reveals that the null of unit income elasticity cannot be rejected. The estimate of the interest semi-elasticity is generally positive in both models. However, as pointed out earlier this point estimate is not robust either across models or methods. Moreover, some may argue that their inferences based on these estimates are meaningless in view of the instability of parameters.

The plots of sequential parameter estimates correspond to the FM method with QS kernel for Model I are presented in Figures 5 to 8. These figures give visual support for the earlier conclusions about parameter instability of the cointegrating vector for Canada, Norway, Sweden, and the United Kingdom in respect to the income elasticity, interest semi-elasticity, and elasticities for proxy for monetization and financial sophistication variables. The first value of the FM estimate and estimate of the parameter ±2 standard errors of the estimate in these figures correspond to the first 50 observations; thereafter we sequentially added one observation at a time, until all the observations were included to obtain subsequent values given in their figures. These displays clearly show parameter stability except for Canada for some coefficient estimates, suggesting that the institutional hypothesis operationalized by including the *cm* and *fa* variables into the conventional specification for the velocity has failed to yield a constant parameter structural relationship. Indeed, it appears proxy variables such as *cm* and *fa* may have contributed to the instability of the velocity function. Theoretical support for the above empirical result are provided by Hamilton (1989a) and Hallman *et al.* (1991). These results suggest that structural changes resulting from mixtures of normal distribution are likely to be an important ingredient for characterizing the velocity relation based on ideas outlined by Perron (1989), Hamilton (1989b), Chen and Tiao (1990), Gregory and Hansen (1992), and Raj (1993), among others.

Interpreting Instability in the Velocity Function

The results of the instability of the Bordo–Jonung specification can be interpreted in one of two ways. One interpretation is that the parameter specification considered in this paper cannot be given a structural interpretation since it is essentially a semi-reduced form relationship as argued by Hamilton (1989a) and Hallman *et al.* (1991) on theoretical grounds.

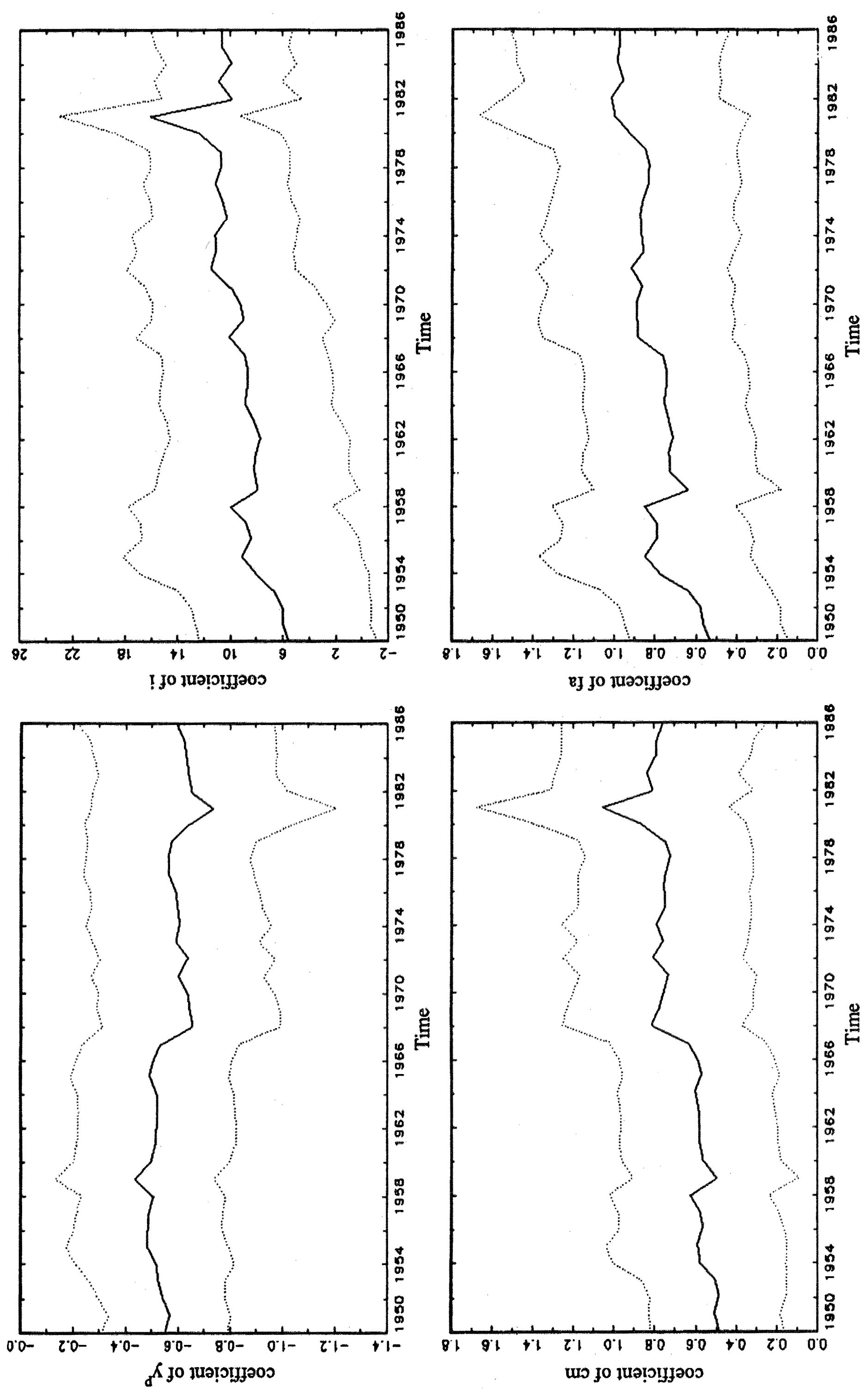

Figure 5. The sequence of FM estimates of parameters of long-run velocity function for Canada plus and minus two standard errors of estimates

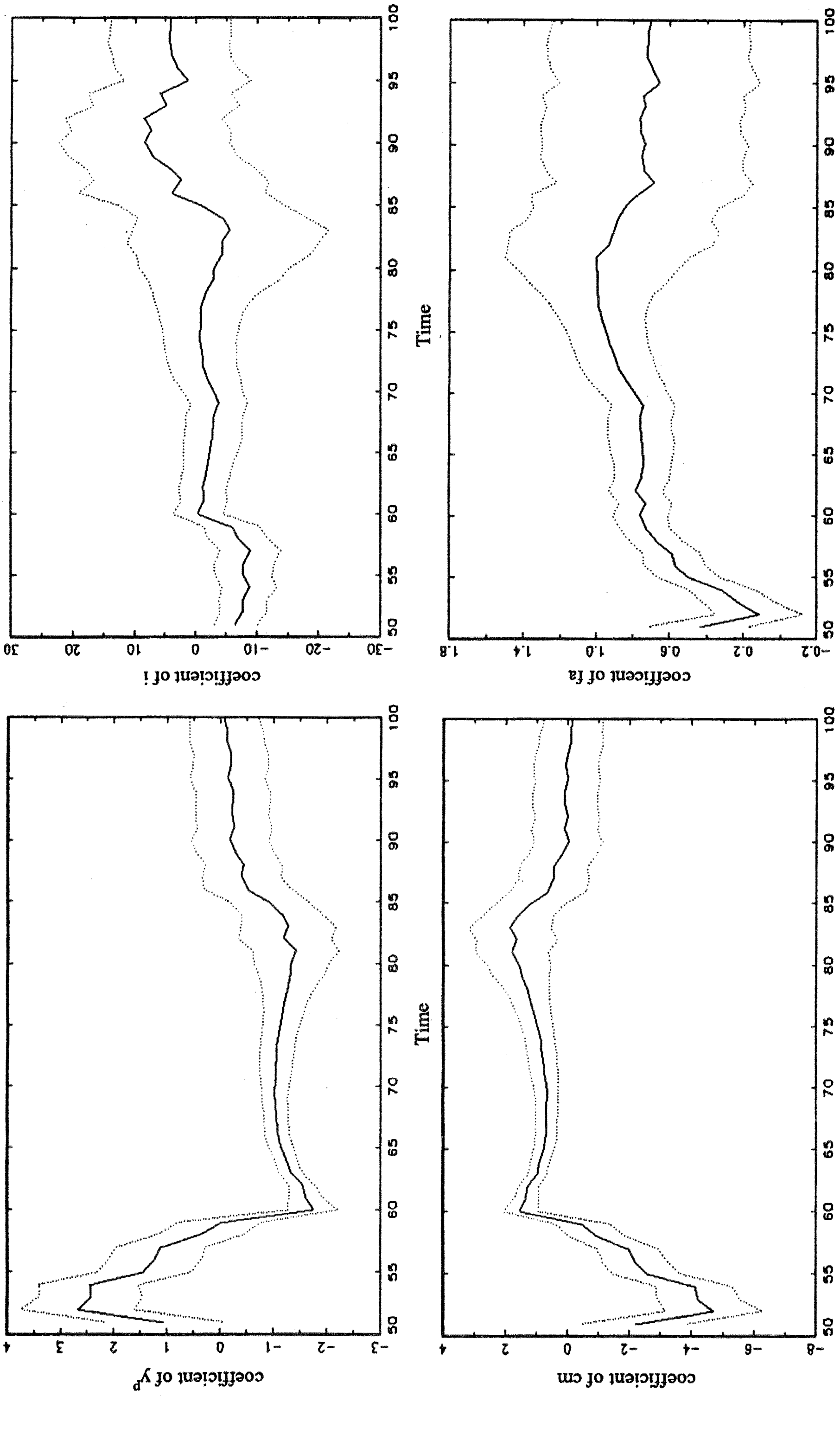

Figure 6. The sequence of FM estimates of parameters of long-run velocity function for Norway plus and minus two standard errors of estimates

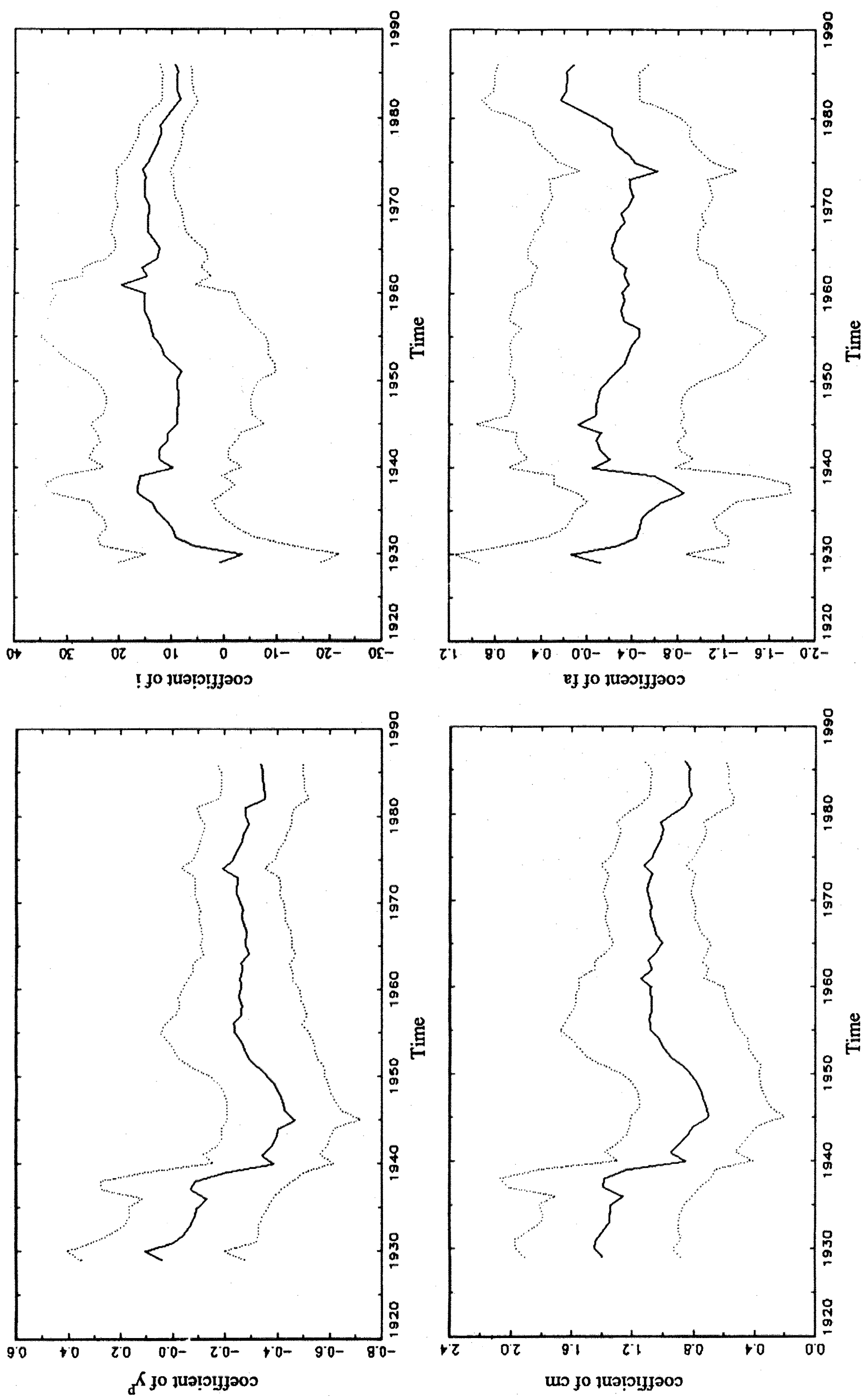

Figure 7. The sequence of FM estimates of parameters of long-run velocity function for Sweden plus and minus two standard errors of estimates

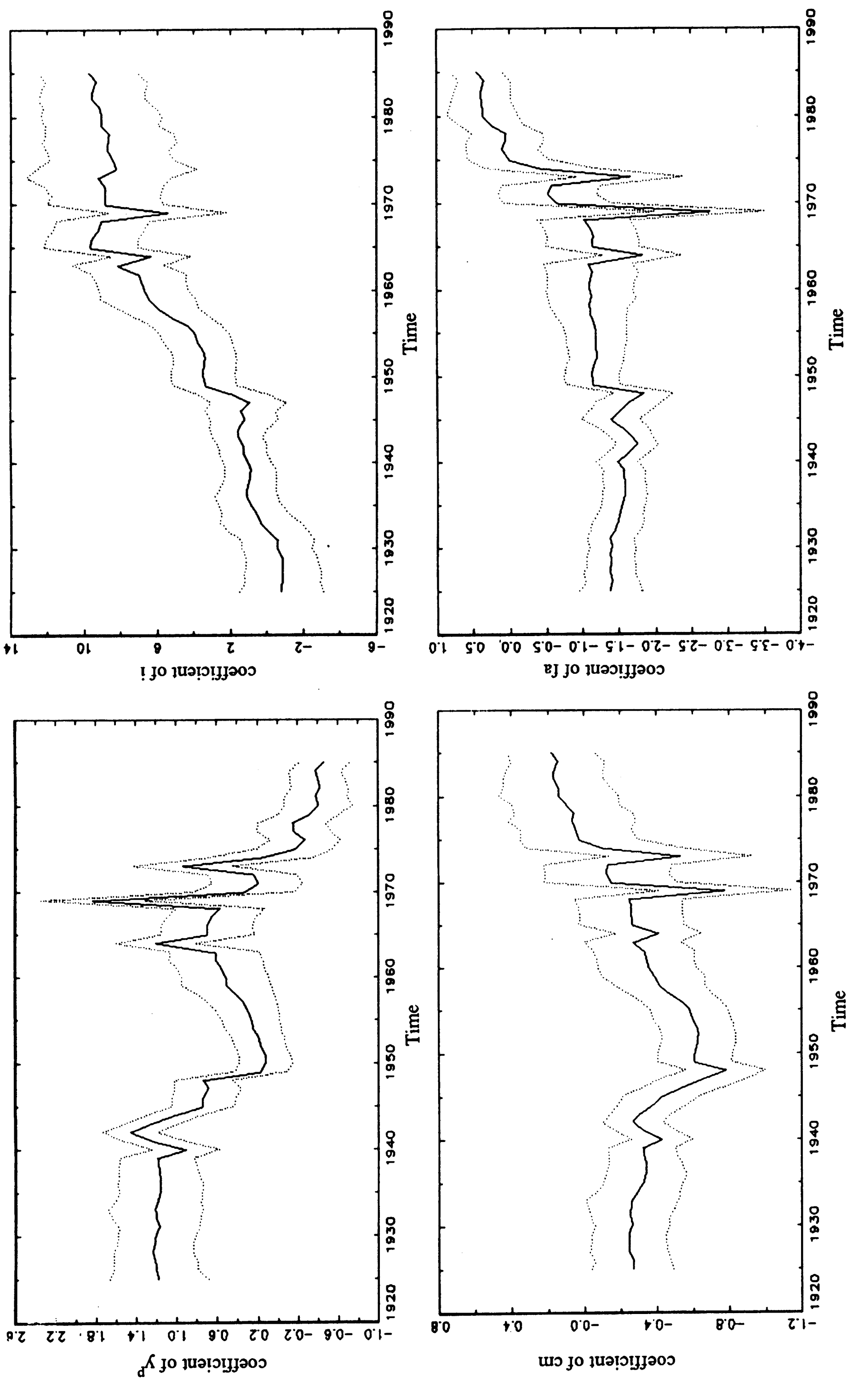

Figure 8. The sequence of FM estimates of parameters of long-run velocity function for the United Kingdom plus and minus two standard errors of estimates

There is another possible interpretation. If institutional change is a plausible explanation of at least some of the on-going changes in the long-run velocity function in a number of countries over the last decade, as stressed by both Wicksell and Fisher long ago, and more recently by Bordo and Jonung, then it might be impossible to obtain a stable relationship. Moreover, it suggests that policy rules that attempt to impose some sort of fixed targets on the monetary aggregates to ensure monetary stability are unlikely to be successful for at least two reasons. One reason is that these rules will likely provoke reaction on the part of those affected by them. The second is that these fixed rules tend to limit the discretionary abilities of monetary authorities when it is most needed to allow continual monitoring of the private sector's response to their action or behaviour in controlling the growth of the money supply. A similar argument is made by Laidler (1982, Chapter 5) and Goodhart (1989) based on Lucas's critique against using fixed rules. Finally, if the transmission mechanism works through changes in income and interest rates as a consequence of shocks to money supply and if the supply of the money is affected by institutional and technical factors, then it is quite possible to envisage a situation where the demand for money will appear unstable and unpredictable due to unpredictable changes in institutions and/or technological changes in the payment system as it works through the financial system's dynamic mechanisms. It should therefore hardly be surprising to find that a linear velocity function with fixed parameters is an inadequate description of the velocity function. More research is needed to discriminate between these two different interpretations.

5. CONCLUSIONS

The idea of a stable long-run demand for money has received considerable attention and scrutiny in recent years after becoming fragmented and sidelined in the mid-1970s and 1980s. It continues to have empirical and policy relevance in contemporary settings, as persuasively argued by Laidler (1990) in his introductory essay to the special volume for the *Journal of Policy Modelling*. He further argued that modelling short-run dynamics of money holdings when they are away from long-run equilibrium remains generally ill-understood.

Bordo and Jonung (1987, 1990) proposed a theory of velocity's long-run behaviour that accounts for the instability in the conventional money-demand specification. According to their hypothesis, the monetization of the economy causes velocity to fall over time during the early phases of financial development. In the later phases of financial development, however, increasing financial sophistication leads to the invention of a variety of money substitutes, causing velocity to rise over time. It was recently shown that only when the conventional specification is extended to include the currency–money ratio and the ratio of total non-financial assets to total assets variables as proxies for institutional change is it possible to reject the null of no cointegration against the alternative hypothesis of a single cointegrating vector. This result appears to hold for the most part when the null of cointegration is tested against the alternative of no cointegration using the LM test, but recent research has shown that there is no unique way to obtain an equilibrium relationship between income velocity of M2 money and its determinants. For example, Hallman *et al.* (1991) found that income velocity for M2 money for the USA is cointegrated with the fraction of labour force not employed in agriculture variable, which is another proxy for institutional change suggested by Bordo and Jonung (1987). This result has been shown to hold for other countries by Raj (1994). Similarly, Raj (1993) has shown that income velocity of M2 money forms a cointegrating relationship with income and interest rate (the traditional determinants of the velocity function for the four countries analysed in this paper using the same data) once the role of a large shock is taken into account.

This evidence suggests that the long-run restrictions implied by cointegration are necessary but not sufficient to identify a structural relationship between income velocity of M2 money and its determinants without proper regard economic theory and measurement issues.

In summary, we have provided empirical evidence in support of earlier theoretical arguments (cf. Hamilton, 1989a; Hallman *et al.*, 1991, among others) suggesting that, while income velocity of M2 money might be influenced by on-going institutional change variables, there is a need to exercise some care in how their influence is modelled. Adding proxy variables into the velocity function entails the perils of converting a structural equilibrium relationship into a semi-reduced-form relationship.

APPENDIX

The income velocity of M2 money is measured as the ratio of income (as defined below) to the average M2 money stock held in the non-public institutions (which is also defined below). The real per capita permanent income (or wealth) is measured as the weighted average of real per capita income using Friedman's weights (see Bordo and Jonung, 1987, Appendix IB). The interest rate is measured as the short-term rate for all countries, except Canada, Norway, and Sweden for which the interest rate is measured as the long-term bond yield. The variable *cm* is measured as the ratio of currency over money. Finally, the variable *fa* is measured as the ratio of total non-bank financial assets to total financial assets.

The money value of income is measured as the Gross National Product at market prices for Canada and the United States, the Gross Domestic Product at market prices for Norway and Sweden, and the Net National Product for the United Kingdom. The money stock M2 is defined as the sum of currency + demand deposits + savings deposits, except for the United Kingdom for the period after 1967. For the United Kingdom the money stock data after 1967 is for M3, since data for M2 are not available.

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