Exchange Rate Predictability and Monetary Fundamentals in a Small Multi-Country Panel*

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Abstract

In this paper a panel of vector error correction models based on a common long-run relationship is utilized to test whether the Euro exchange rates of Canada, Japan and the United States have a long-run link with monetary fundamentals. We use both exchange relationships relative to the full EMU area (with synthetic aggregates for the pre-EMU period) and relative to Germany solely. Compared to existing cointegration frameworks our approach provides more evidence that the aforementioned exchange rates are consistent with a rational expectations-based monetary exchange rate model based on a common long-run relationship, albeit with a long-run impact of relative income that is higher than predicted by the theory. As a next step we analyze the out-of-sample fit of this common long-run exchange rate model relative to naive random walk-based forecasts. These forecasting evaluations indicate that the monetary fundamentals-based common long-run model is superior to both random walk-based forecasts and standard cointegrated VAR model-based forecasts, especially at horizons of 2 to 4 years.

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JEL classification: C12, C23, F31.

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1 Introduction

If monetary exchange rate models are valid representations of long-run exchange rate behaviour then the exchange rate will only deviate from its monetary fundamentals in the short-run. Within an efficient foreign exchange market long-run currency pricing will consequently reflect the (true) market expectation regarding the future behaviour of monetary fundamentals. The purpose of this study is to test the empirical validity of monetary model-based foreign exchange pricing, and in testing this we follow a two-tier route. First, we want to establish whether there is a long-run link between exchange rates and monetary fundamentals according to the monetary exchange rate model through cointegration tests. Also, we want to exploit the forward-looking feature of the efficient market-based monetary exchange rate model to analyze the out-of-sample effects on exchange rate movements of this model.

In our empirical analysis we focus on the Euro exchange rates of Canada, Japan and the United States [U.S.] for the 1975-2000 period. The establishment of the European Monetary Union [EMU] in 1999 in our view provides the ultimate test of the forward looking feature of the efficient market-based monetary model, as both the moment as well as the institutional set-up were publicly announced well in advance. The expected behaviour of the EMU-based monetary fundamentals should therefore already have been discounted in the respective spot Euro exchange rates before 1 January 1999. For our proxy of both the Euro exchange rates as well as the EMU-based monetary fundamentals we at first make use of constructed synthetic EMU data from the 7 major EMU member states. As a second alternative we follow Arnold and de Vries (2000) who suggests that the historical behaviour of EMU money demand can best be approximated by extrapolating the behaviour of the country which mimics the European Central Bank's [ECB] line of conduct best instead of constructing artificial synthetic EMU data. Hence, we also use solely German data as a proxy of pre-1999 EMU behaviour and extrapolate the behaviour of Germany-based monetary fundamentals beyond 1 January 1999.

A number of recent studies, in particular Mark (1995) and Chinn and Meese (1995), tried to exploit the assumed long-run link between exchange rates and monetary fundamentals, and they claim that current monetary fundamentals-based disequilibria can predict the exchange rate four years ahead, both in-sample and out-of-sample for the U.S. dollar exchange rates of Canada, Germany, Japan, Switzerland and the U.K. over the 1973-1991 period. This long horizon predictability, however, breaks down in Groen (1999) for out-of-sample tests on these exchange rates within a 1973-1994 sample. Groen relates his result to the absence of cointegration between exchange rates and monetary fundamentals, *i.e.* a stationary linear combination of the exchange rate and its monetary fundamentals is absent. Hence, the issue of cointegration is of major importance in establishing a predictable link between exchange rates and their monetary fundamentals.

Based on the cointegrated vector autoregressive [VAR] framework of Johansen (1991), however, Sarantis (1994) and Groen (2000) do not find evidence for cointegration based on the monetary model for a large number of OECD-countries' exchange rates relative to the U.S. dollar, the Deutsche Mark [DM] and the pound-sterling. On the other hand, Groen (2000) finds that the use of cross-section regressions for a large number of countries, or tests for cointegration within a fixed individual effect multi-country panel data model, result in more empirical evidence in favor of the monetary exchange rate model. Like Groen (2000), Mark and Sul (2001) use a panel of bilateral exchange rates for 17 OECD countries over the period 1973-1997 to analyze the empirical features of the monetary model. The results in Mark and Sul indicate that there is cointegration based on the monetary model and that monetary fundamentals significantly predict future exchange rate returns using panel regression estimates with fixed time effects. Hence, the analysis of the monetary exchange rate model across multiple countries simultaneously seems to provide more positive results than the single country monetary model.

The panel cointegration frameworks described in Groen (2000) and Mark and Sul (2001) are essentially pooled versions of the two-step Engle and Granger (1987) procedure and therefore only allow for a limited amount of cross-country heterogeneity. From the power analysis in Groen (2002) it becomes clear that the power gain of this panel Engle-Granger framework over a pure time series approach may be limited for panels with a small number of countries. The power analysis in Groen (2002) shows, on the other hand, that in panels with a limited amount of countries the Groen and Kleibergen (2003) approach has a significant precision gain over pure time series-based techniques. The panel cointegration framework developed by Groen and Kleibergen (2003) is based on a panel of the individual vector error correction [VEC] models across the constituting countries. Another attractive feature of the aforementioned framework is that one can test whether or not the long-run relationships are identical across the countries given heterogeneous short-run dynamics.

From the results in Berkowitz and Giorgianni (2001) and Groen (1999) it becomes clear that one has to have cointegration based on the monetary exchange rate model in order to find a significant out-of-sample fit of this structural exchange rate model. We therefore first test whether there is cointegration based on the monetary exchange rate model for the entire sample both through the purely time series-based Johansen (1991) cointegrated VEC framework as well as through the panel cointegration framework from Groen and Kleibergen (2003). Next, following the seminal work of Meese and Rogoff (1983) we analyze the out-of-sample performance of our cointegrated and panel cointegrated exchange rate models relative to that of the random walk model.

In the remainder of this paper, Section 2 summarizes the cointegration restrictions implied by the monetary exchange rate model and shows what its implications are in an efficient market context. The issue of how to approximate the behaviour of EMU aggregates over the pre-EMU era is discussed in Section 3. Section 4 discusses the two (panel) cointegration procedures used, followed by an application of the two procedures on our data. The methodology and the results of the out-of-sample forecasting evaluation of the monetary exchange rate model can be found in Section 5. Finally, Section 6 contains conclusions.

2 The Monetary Exchange Rate Model and Efficient Markets

The link between exchange rates and macroeconomic fundamentals such as money quantities and real income can be made explicit in the context of the monetary exchange rate model, see *e.g.* Mussa (1976). Within this framework it is assumed that both purchasing power parity [PPP] and uncovered interest rate parity [UIP] hold. It is also assumed that in logarithms there is a stable linear relationship between the logarithm of the demand for real balances on the one hand and the log of real income and the nominal interest rate on the other hand. In this stable log-money demand relationship income elasticity is denoted as $\phi > 0$ and interest semi-elasticity equals $\omega > 0$, and these elasticities are assumed to be common across countries.

Suppose now that we have a variable ζ_t, which is the composite of temporary PPP, UIP, and respective money demand disturbances, and that this variable complies for notational convenience with a stationary AR(1) model:

$$\zeta_t = \rho\zeta_{t-1} + \xi_t, \tag{1}$$

with $0 < \rho < 1$ and $\xi_t \sim$ i.i.d.$(0,\ \sigma_\xi^2)$. Combining ζ_t with PPP, UIP and stable log-money demand functions at home and abroad yields a rudimentary monetary exchange rate model, and the log-exchange rate solution of the this model is a present value relationship of the form:[1]

$$s_t = c + \frac{1}{1+\omega}\sum_{i=0}^{\infty}\left(\frac{\omega}{1+\omega}\right)^i \mathrm{E}_t(f_{t+i}) + \left(\frac{\zeta_t}{1-\rho}\right), \tag{2}$$

where:

$$f_t = (m_t - m_t^*) - \phi(y_t - y_t^*).$$

In accordance with Campbell and Shiller (1987) and Campbell *et al.* (1997, Chapter 7)

[1]We define s_t as the log of current domestic price of one unit of foreign currency, m_t (m_t^*) is the log of the current domestic (foreign) money supply and y_t (y_t^*) is the log of current domestic (foreign) real income.

deviations from (2) can be written as a function of expected future growth rates of the monetary fundamentals, *i.e.*

$$s_t - f_t = c + \sum_{j=1}^{\infty} \left(\frac{\omega}{1+\omega} \right)^j \mathrm{E}_t(\Delta f_{t+j}) + \left(\frac{\zeta_t}{1-\rho} \right). \tag{3}$$

The asset market approach implies, through (2), that when the monetary fundamentals are I(1) the exchange rate also has a unit root. Through (3) the monetary exchange rate model has the implication that within efficient foreign exchange markets $s_t - f_t$ is stationary as the right hand side of (3) only contains I(0) variables and thus s_t, $(m_t - m_t^*)$ and $(y_t - y_t^*)$ are cointegrated with the cointegrating vector: $\boldsymbol{\beta} = (1 \;\; -1 \;\; \phi)'$. Equation (3) also makes it clear that current deviations between s_t and f_t are induced by discounted expectations of both future growth rates of f_t and the rate at which current disturbances in the PPP, UIP and relative money demand relationships die out in the future.

3 Approximating Pre-1999 EMU Behaviour

The focus on Euro exchange rates and EMU-based monetary fundamentals raises the issue of what data one should use for the pre-EMU era. A number of papers that study the behaviour of European Monetary System [EMS]-wide money demand functions, most notably Kremers and Lane (1990), have utilized weighted averages of the individual data of the member countries. In general these studies find favorable evidence for a very stable EMS-wide money demand function. From the description of the efficient market-based monetary exchange rate model in Section 2 it becomes clear that currency pricing is based on the expected behaviour of money demand at home and abroad. As such the usage of artificial European-wide monetary aggregates could prove to be favorable for tests of the monetary model. Indeed, when la Cour and MacDonald (2000) utilize a monetary model-based vector error correction model for the ECU/U.S. dollar rate based on synthetic European Union-wide monetary aggregates they find ample evidence *pro* the monetary model, both in-sample as well as out-of-sample. Given the aforementioned results we shall start off the analysis in Sections 4.2 and 5 of the Euro exchange rates of Canada, Japan and the U.S. plus the corresponding monetary fundamentals based on synthetic EMU aggregates for the pre-1999 period. In constructing these synthetic EMU aggregates we follow the approach of Beyer *et al.* (2001) based on variable-weight aggregators of individual growth rates and which are in level terms spliced together with the EMU-wide data for the period from 1999 onwards.[2]

However, as formalized by Arnold and de Vries (2000), the use of these weighted averages of the individual series for the pre-1999 period can overstate the stability in the underlying relative money demand function on which the monetary exchange rate model is based. Crucial in this line of reasoning is the difference in the co-movements amongst the national monetary aggregates of the EMU countries before and after the establishment of EMU. The introduction of a common monetary policy amongst the EMU countries induced a strong co-movement in the corresponding monetary aggregates as changes in monetary policy and money demand affect all these aggregates simultaneously from 1999 onwards. In the pre-EMU era, however, national monetary policies amongst the future EMU member states diverged substantially, especially during the 1970s, as witnessed

[2]Appendix A contains a more elaborate description of the construction of our synthetic pre-1999 EMU aggregates.

by the substantial movements in the bilateral exchange rates relative to Germany over this period for many current EMU member states. The resulting lack of co-movements across the national monetary aggregates could therefore induce, through the law of large numbers, a degree of stability in the weighted averages of the national monetary aggregates for the pre-1999 period which is by definition not present from 1999 onwards.[3] Hence, in the EMU period area-wide money aggregates exhibit a different behaviour than during the pre-EMU period. As an alternative Arnold and de Vries (2000) propose to extrapolate the behaviour of the country whose central bank design is closest to the design of the European Central Bank. As there is a consensus in the literature that the ECB design is inspired by that of the Bundesbank, we alternatively use the German Deutsche Mark [DM] as the numeraire currency and use the corresponding monetary fundamentals of Germany *vis-à-vis* Canada, Japan and the U.S.

4 Cointegration Tests of the Monetary Model

Several methods are available to test whether or not the log exchange rate s_t is cointegrated with the log monetary fundamentals $(m_t - m_t^*)$ and $(y_t - y_t^*)$. Traditionally, these methods are based on time series data and the most frequently used method is the VAR-based framework of Johansen (1991). However, Otero and Smith (2000) show in Monte Carlo studies that the power of this particular approach to reject the null of no cointegration

[3] Consider for example modeling Dutch-German money demand since March 1983 when the Dutch guilder/DM exchange rate has remained stable. As a consequence the Arnold and de Vries (2000) critique would not hold in this particular case. Indeed, based on an extensive empirical analysis they show that the transition to EMU did not cause a significant increase in the co-movement between Dutch and German monetary aggregates.

in the face of a persistent alternative depends on the span of the data sample. As the post-Bretton Woods sample covers a quite short time span one would expect that methods such as those of Johansen (1991) have difficulty in verifying the cointegration restriction as summarized in (3). Alternatively, one could apply panel-based techniques in which inference is based on an artificially extended number of observations. In Section 4.1 we discuss such a panel-based approach and contrast it with the time series-based framework of Johansen (1991). Both cointegration frameworks are subsequently applied in Section 4.2 on our sample of exchange rates and monetary fundamentals of Canada, Japan and the U.S. relative to both the full EMU-area and Germany in order to test the cointegration restriction of the monetary exchange rate model within an efficient Euro foreign exchange market.

4.1 Cointegration Tests for Time Series and Panels

To test for cointegration between s_t, $(m_t - m_t^*)$ and $(y_t - y_t^*)$ we can use the VEC framework of Johansen (1991), *i.e.*

$$\Delta X_t = \sum_{s=1}^{3} \chi_s \bar{D}_s + \boldsymbol{\alpha} \left(\boldsymbol{\beta}' \quad - \boldsymbol{\beta_0}' \right) Z_{t-1} + \sum_{j=1}^{p-1} \Gamma_j \Delta X_{t-j} + \varepsilon_{it}. \tag{4}$$

In (4) the 3×1 vector X_t equals

$$X_t = \left(s_t \quad (m_t - m_t^*) \quad (y_t - y_t^*) \right)',$$

$\Delta X_t = X_t - X_{t-1}$, $Z_{t-1} = (X_{t-1}' \quad 1)'$, $\bar{D}_s$ is a zero-mean seasonal dummy and ε_{it} is a 3×1 vector of white noise disturbances. The $1 \times r$ vector $\boldsymbol{\beta_0}$ is a vector of intercept terms, $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ are $3 \times r$ matrices of adjustment parameters and cointegrating vectors, respectively, and r is the cointegrating rank value of VEC model (4). Note that this specification of the deterministic part of (4) implies that the intercepts appear only in the long-run relationships, which is in compliance with (3).

The Johansen (1991) likelihood ratio statistic for the null of r cointegrating vectors versus the alternative of a stationary VAR, which has a non-standard asymptotic distribution, can be used to determine the proper value of the cointegrating rank r in (4). After the proper cointegrating rank is determined likelihood ratio tests are used to test restrictions on the r cointegrating vectors. As these tests are conducted conditional on the cointegrating rank they have standard limiting distributions. Validity of the monetary model within VEC (4) implies a reduced rank value $r = 1$ and a cointegrating vector, normalized on s_t, equal to $(\bar{\boldsymbol{\beta}}' \quad -\bar{\boldsymbol{\beta}}_{\mathbf{0}}') = (1 \quad -1 \quad \phi \quad -c)$, which complies with $-\phi < 0$ in (2).

VEC model (4) can therefore be used to *separately* test the cointegration restriction in (3) for each bilateral exchange rate in our sample. However, both the log exchange rates s_{it} and the corresponding spreads $s_{it} - f_{it}$ exhibit in reality co-movements across countries due to contemporaneous correlation. This contemporaneous correlation is caused by the fact that we analyze bilateral exchange rates relative to the same *numeraire* and movements related to the base country therefore induces contemporaneous correlation. Also, common factors such as herd behaviour, contagion and so on, which are not part of our model, can cause the aforementioned co-movements. Groen (2000) shows that within a cross-section of 14 OECD countries long-run DM exchange rate changes are related to long-run changes in the corresponding monetary fundamentals based on common parameter values. Hence, based on these arguments one has to analyze all the bilateral exchange rates in our sample simultaneously as applying the VEC (4) for each exchange rate separately will then be based on inefficiently estimated parameters and thus decrease the power of this approach.

Alternatively like Groen and Kleibergen (2003) one can stack VEC models like (4), constructed for each of our 3 exchange rates, into one system:

$$\begin{array}{rcl} \Delta X_{1t} = & \sum_{s=1}^{3} \chi_{1,s}\bar{D}_s + \boldsymbol{\alpha}_1 \left(\boldsymbol{\beta}_1' \quad -\boldsymbol{\beta}_{0,1}'\right) Z_{1,t-1} + \sum_{j=1}^{p_1-1} \Gamma_{j,1}\Delta X_{1,t-j} + \varepsilon_{1t} \\ \vdots & \vdots \qquad\qquad \vdots \qquad\qquad \vdots \qquad\qquad \vdots \\ \Delta X_{Nt} = & \sum_{s=1}^{3} \chi_{N,s}\bar{D}_s + \boldsymbol{\alpha}_N \left(\boldsymbol{\beta}_N' \quad -\boldsymbol{\beta}_{0,N}'\right) Z_{N,t-1} + \sum_{j=1}^{p_N-1} \Gamma_{j,N}\Delta X_{N,t-j} + \varepsilon_{Nt} \end{array} \qquad (5)$$

and estimate this system under the assumption of a *common* cointegrating rank $r_1 = \cdots = r_N = r$. Groen and Kleibergen (2003) show that maximum likelihood estimation of the parameters in a panel VEC model such as (5) can be achieved by applying iterative Generalized Method of Moment [GMM] estimators. Again, as in the time series approach of Johansen (1991), the corresponding likelihood ratio tests on the common cointegrating rank value r have a non-standard limiting distribution. We consider and test different possible specifications of the cointegrating vectors in our panel VEC model which can be defined as:

Definition 4.1 *In the N-country panel VEC model which corresponds with (5) we test the following hypotheses:*

- $\boldsymbol{B}(r)|\boldsymbol{A}$ *tests the null hypothesis that for each cross-sectional unit i the cointegrating rank value equals r versus the alternative hypothesis that for each cross-sectional unit i we have a full rank value while not assuming $\boldsymbol{\beta}_1 = \cdots = \boldsymbol{\beta}_N$ in (5).*

- $\boldsymbol{C}(r)|\boldsymbol{A}$ *tests the null hypothesis that for each cross-sectional unit i the cointegrating* *rank value equals r versus the alternative hypothesis that for each cross-sectional unit i we have a full rank value assuming $\boldsymbol{\beta}_1 = \cdots = \boldsymbol{\beta}_N = \boldsymbol{\beta}$ in (5).*

When the monetary exchange rate model is appropriate within panel VEC (5) we thus expect to find the following test results:

- (5) has a common cointegrating rank value of $r = 1$ combined with $\boldsymbol{\beta}_1 = \cdots = \boldsymbol{\beta}_N = \boldsymbol{\beta}$,
- the common cointegrating vector must comply with the parameter restrictions of the monetary model, and these restrictions are tested through a likelihood ratio statistic which is conditional on the common cointegrating rank r and thus has a standard (χ^2) asymptotic distribution.

4.2 Testing the Cointegration and Parameter Restrictions

Quarterly Euro exchange rates and the accompanying monetary fundamentals of Canada, Japan and the U.S. are analysed for the period 1975-2000. These monetary fundamentals are based on M3 monetary aggregates and real GDP, for the respective bilateral relationships in the pre-1999 period we use synthetic aggregates based on weighted averages of the data for Austria, Finland, France, Germany, Italy, the Netherlands and Spain. We repeat our analysis on quarterly DM exchange rates and Germany-based monetary fundamentals of Canada, Japan and the U.S. for the same time span.[4]

Any VAR-based inference is dependent on the choice of the lag order, and this equally applies to the panel VEC-based cointegration tests as these are based on stacking individ-

[4]For a more detailed description of the data and its sources, see Appendix A.

ual VAR systems (see Section 4.1). We first construct unrestricted VEC models as in (4) with $r = 3$ for Canada, Japan and the U.S.[5] The lag order p for each of these countries is determined through the minimum of the Akaike Information Criterion [AIC] computed for each of $p = 0, \ldots, 8$.

We than analyze our three exchange rates simultaneously (either relative to the full EMU area or Germany solely) within a stacked system or panel of the individual VEC models such as in (5). In particular, we first test the common cointegration rank value across Canada, Japan and the U.S. within (5) through the common cointegration rank likelihood ratio statistics of Groen and Kleibergen (2003) based on the lag orders of the individual systems in Table 3.[6] The panel VEC-based statistics are reported in Table 1 and the first row of this table contains both the computed value of the likelihood ratio statistic for the null of no cointegration across the three countries and the corresponding critical values. The result in this first row indicates that the null of no cointegration across the countries has to be rejected irrespective of the choice of the *numeraire*. In the second row of Table 1 we report the test results for the null of a common cointegration rank $r = 1$ based on country specific cointegrating vectors. From the values in this row one can observe that neither *vis-à-vis* the full EMU area nor Germany can this null be rejected.

An interesting testable hypothesis is whether the single cointegrating vectors of Canada,

[5]Unreported unit root tests indicate that s_t, $(m_t - m_t^*)$ and $(y_t - y_t^*)$ are I(1) for our bilateral relationships, which is in compliance with *e.g.* MacDonald and Taylor (1993) and de Vries (1994).

[6]The usage of the individual lag orders is based on the 'bottom-up' modelling strategy for restricted VAR models from Lütkepohl (1993, pp.182-183) such that the appropriate lag orders for the panel VEC model are set equal to the optimal AIC-based lag order for each country-specific VEC sub-model individually.

Japan and the U.S. are, apart from the intercept term, identical across these countries. The fourth row of Table 1 contains the outcome of a likelihood ratio test of the joint hypothesis of a common cointegrating rank $r = 1$ combined with a common cointegrating vector (apart from the intercept). This test result reveals that we have across Canada, Japan and the U.S. cointegration based on a single cointegrating vector per country with parameters which are homogeneous across both the full EMU-based exchange rates as well as the Germany-based exchange rates.

The discussion of the monetary model in Section 2, however, showed that cointegration based on a single equilibrium relationship is not enough for the monetary model to be empirically valid. It should also be the case that within the cointegrating vector $\boldsymbol{\beta} = (\beta_s\ \beta_{(m-m^*)}\ \beta_{(y-y^*)}\ \ -\beta_0)'$ the restriction $\beta_s + \beta_{(m-m^*)} = 0$ should be valid, which implies that the normalized cointegrating vector corresponds with

$$\begin{pmatrix} \bar{\boldsymbol{\beta}} \\ -\bar{\boldsymbol{\beta}}_{\mathbf{0}} \end{pmatrix} = (1\ \ -1\ \ \phi\ \ \ -c)', \tag{6}$$

as suggested in Section 2. The first and third columns of Table 2 contains the normalized cointegrating vector for (5) estimated under $r = 1$ and a common cointegrating vector. The signs of both the relative money elasticities as well as the relative income elasticities are in compliance with the theory. We compute the likelihood ratio statistic for $\mathrm{H}_0 :\ \beta_s + \beta_{(m-m^*)} = 0$ versus $\mathrm{H}_1 :\ \beta_s + \beta_{(m-m^*)} \neq 0$ for the common cointegrating vector of the Canada/Euro, the Canada/Germany and the U.S./Germany relationships, and this test statistic has a $\chi^2(1)$ distribution as the restriction only applies to one parameter. Values of the likelihood ratio test statistic can be found in the fourth row of Table 2. This test result verifies that the estimated relative money elasticity is not significantly different from

1 both in case of full EMU-area data and in case of German data, and the corresponding normalized common cointegrating vectors can be found in the second and fourth column of Table 2. From these two columns it also becomes apparent that the estimated relative real income elasticities are quite similar across the two *numeraire* choices as both parameters are close to 3. Equality of both long-run relationships hinges on the null whether the relative income elasticity is identical across the two relationships and the row of Table 2 contains a likelihood ratio test of this restriction. Irrespective of the *numeraire* for the three exchange rates relative real income elasticity is not significantly different from 3. Therefore, the choice of the base currency has no influence on the common long-run monetary model-based exchange rate relationship.

One of the factors which causes deviations between the log spot exchange rate and the log monetary fundamentals to occur within our forward-looking monetary model is the persistence of disturbances in the underlying PPP, UIP and relative money demand relationships, see (3). However, a high rate of persistence of the aforementioned disturbances, *e.g.* when ρ in (3) is close to 1, in combination with the short post-Bretton Woods time span would make it very difficult for pure time series-based tests on individual exchange rates to detect strong empirical evidence for our monetary model. The precision gain from the panel VEC framework can potentially circumvent this problem, and in order to show this we repeat the exercise for the bilateral rates separately.

We conduct, sequentially, the Johansen (1991) likelihood ratio 'trace' test for the null hypothesis of r cointegrating vectors in (4) for each bilateral rate, and these are reported in Table 3. The results in Table 3 show that in case of the full EMU-based relationships we only can reject the null of no cointegration, accept the null of a single cointegration relationship between the log exchange rate and the log monetary fundamentals, and accept the parameter restrictions of the monetary model for the Canada/Euro rate. The cointegration test results seem to be more favourable for the Germany-based relationships as the null of no cointegration is rejected and the null of 1 cointegrating vector is accepted for Canada and the U.S. However, the corresponding likelihood ratio test on whether we have a unity relative money elasticity in the cointegrating vector indicate that we cannot accept this restriction of the monetary model, see the fourth column of Table 3.

Thus, in contrast to the individual time series models our panel VEC models for both the full EMU-based exchange rates and the DM-exchange rates of Canada, Japan and the U.S. provides empirical evidence for the validity of both the cointegration and parameter restrictions of the monetary exchange rate model. The resulting estimated equilibrium relationships in the second and fourth columns of Table 2 do, however, exhibit a high relative real income elasticity value of 3. This value of income elasticity is particularly high in comparison with estimates in the money demand literature, suggesting that relative income is correlated with the exchange rate for other reasons than the effect on money demand at home and abroad. One possible explanation could be that one of the building blocs of the monetary model, PPP, does not hold. Instead of complying with PPP, the real exchange rate could be driven by the Balassa (1964)-Samuelson (1964) effect. In the Balassa-Samuelson framework fast growing economies exhibit a higher productivity growth in the traded goods sector than in the non-traded goods sector and thus the relative price of non-traded/traded goods for such an economy would rise quickly. The corresponding bilateral real exchange rate relative to slower growing economies would therefore show a sustained appreciation. If productivity is correlated with real income the high value of relative income elasticity could be due to the impact of this Balassa-Samuleson effect on the exchange rate. Indeed, in case of the Euro/dollar exchange rate Chinn and Alquist (2000) find more favourable evidence for a sticky-price monetary model if they include proxies for the relative non-tradable/tradable price. Alternatively, the high relative income elasticity value of 3 could be due to overreaction in the response of FX market participants to real income news. This would be in compliance with the findings of Andersen *et al.* (2001), in which there is a relatively large high frequency exchange rate response to real news and a relatively low response to nominal news.

5 Out-of-Sample Evaluation of Exchange Rate Movements

Since the seminal paper of Meese and Rogoff (1983) it has become common practice in international finance to subject structural exchange rate models to an out-of-sample forecasting competition *vis-à-vis* non-structural models such as the random walk model. This section deals with the out-of-sample performance of our monetary model-based panel VEC model for the EMU-based and Germany-based exchange rates of Canada, Japan and the U.S. relative to both non-structural models and purely time series-based approaches. The methodology through which we measure this out-of-sample performance is described in Section 5.1. Section 5.2 presents the results of our forecasting evaluations.

5.1 Methodology

Meese and Rogoff (1983) compared post-sample predictions of several monetary exchange rate models for the log level of the exchange rate with those of a random walk or 'no change' model at forecasting horizons up to 1 year. Mark (1995) and Chinn and Meese (1995) conducted a similar exercise in which they compared the out-of-sample exchange rate change predictions of current error-correction terms with the predicted change, *i.e.* a zero change, of the random walk model at horizons up to 4 years. As this has become standard in international finance, we shall also follow this approach and compare the out-of-sample exchange rate forecasts (in levels) of either (4) or (5) for Canada, Japan and the U.S. with those of a random walk. Our evaluation criterion for the log exchange rate level is the root of the mean of squared forecast errors [RMSE]

$$\text{RMSE} = \sqrt{\frac{1}{T - t_0 - h} \sum_{t=t_o}^{T-h} e_{s,t+h}^2}, \tag{7}$$

where t_0 is the first observation in the forecast period, h is the forecasting horizon and $e_{s,t+h}$ is the forecast error of the model-generated prediction of the log exchange rate level relative to the *observed* log exchange rate level.

The forecasts are generated in a recursive manner. Suppose that our first h-period ahead forecast has to be generated at observation t_0 ($t_0 < T$). Consequently, we first estimate on a sample which runs up to t_0 either (4) under $r = 1$ for each of our bilateral rates separately or (5) based on one common equilibrium relationship jointly for all our bilateral rates. For (5) we impose in the recursive re-estimation a common cointegrating vector (1, –1, 3) and re-estimate the constrained intercepts and short-term dynamics.

Similarly in case of (4) we impose in the sub-sample estimates a unity relative money elasticity and a negative relative real income elasticity in the equilibrium relationships and re-estimate the remaining parameters. Based on these estimates we generate our forecasts for the log exchange rate levels at all forecasting horizons h. For $h > 1$ the exchange rate forecasts in the individual cointegrated VAR system (4) and the panel VEC model (5) are generated in a dynamic manner, *i.e.* forecasts for log exchange rates and monetary fundamentals in the previous quarter are used to generate the exchange rate forecast for the current quarter. These two steps are repeated for the observations $t_0+1, t_0+2, \ldots, T-h$.[7] In order to be able to evaluate the behaviour of our monetary model-based forecasts, we construct the ratio of RMSE (7) based on our recursively generated predictions from either (4) or (5) relative to that of the random walk model. For our monetary model-based cointegrated VEC models to be valid these ratios should be smaller than 1.

If our monetary model-based (panel) VEC models have empirical validity they should outperform the random walk-based forecasts. However, both the monetary model-based (panel) VEC models as well as the random walk model impose an identical order of integration for the log exchange rate, *i.e.* I(1). Using a moving average representation for the h-period ahead forecast errors based on level forecasts of I(1) variables, in which the current h-period ahead forecast error is a function of the forecast errors up to $t + h - 1$, Christoffersen and Diebold (1998) show that for $h \to \infty$ the forecast error variance tends to infinity. Hence, a ratio of RMSE measures, which is a measure of the forecast error

[7]Note that the recursive re-estimation of either (4) or (5), under the appropriate restrictions, is based on a new AIC-based selection of the appropriate lag order for the individual VAR systems for each recursive step.

variance, for two forecast models that impose I(1) on the level forecasts could potentially be very close to 1. In order to circumvent this problem and to be able to measure the added value of the restriction of monetary model-based cointegration for exchange rate forecasts we therefore simulate through a parametric bootstrap procedure p-values for each of our estimated RMSE ratios for H_0: RMSE ratio = 1 versus H_1: RMSE ratio < 1.

In this parametric bootstrap procedure we generate for each bilateral relationship artificial I(1) series of the log exchange rate and log monetary fundamentals, based on a random walk for the log exchange rate and autoregressive models for the log fundamentals, which are not cointegrated with each other in order to comply with the null hypothesis. We then apply both (4) and (5) through the aforementioned recursive forecasting procedure on these artificial series to generate artificial equivalents of our RMSE ratios *vis-à-vis* the random walk model. The artificial RMSE ratios from 5,000 parametric bootstrap simulations are then combined with the empirical estimates of the RMSE to compute the p-values. Appendix B contains a more detailed description of the parametric bootstrap estimates of the p-values for H_0: RMSE ratio = 1 versus H_1: RMSE ratio < 1.

5.2 Forecasting Evaluation

We evaluate the forecasting performance of our monetary model-based (panel) VEC models, *i.e.* (4) and (5) with $r = 1$ (in case of (5) based on a common cointegating vector) with monetary model-based parameter restrictions, for the forecast period 1989.I-2000.IV. This forecast period includes important phenomena like the unification of Germany and the launch of EMU, and as such it provides us with an appropriate testing sample to analyze the out-of-sample properties of the efficient market-based monetary exchange rate model. Like recent papers on exchange rate predictability such as Mark (1995) and Chinn and Meese (1995), we consider as forecasting horizons (in quarters) $h = 1$, 4, 8, 12 and 16.

Mark (1995) and Chinn and Meese (1995) concluded that at forecasting horizons of three to four years monetary fundamentals predict exchange rate movements better than the random walk model. Berkowitz and Giorgianni (2001) and Groen (1999), amongst others, showed that these results were potentially spurious as these forecasts were not necessarily based on cointegrated models. But, the relatively short span of the post-Bretton Woods sample and the persistence of short-run dynamics makes it hard to detect monetary model-based cointegration. From the analysis in Section 4.2 it is apparent, however, that in this context the panel VEC framework provides enough precision gain over the individual time series VEC approach to detect monetary fundamentals-cointegration, even for a relatively small cross-section dimension. We therefore can expect to have positive results for exchange rate predictions based on the panel VEC model.

The results for the EMU-based exchange rate level predictions *vis-à-vis* the random walk can be found in the first two columns of Table 4. These two columns report RMSE ratios for the log exchange rate level of an individual time series monetary model-based VEC model (first column) or the monetary model-based panel VEC model (second column) relative to random walk forecasts. The pure time series VEC-based forecasting models are never able to significantly outperform naive random walk-based exchange rate forecasts. Forecasts based on the monetary panel VEC model, however, significantly outperforms the random walk model at horizons of 2 to 4 years in case of the Canada/Euro rate, horizons of 3 to 4 years for the Japan/Euro rate and the 4 year horizon solely for the U.S./Euro rate.

Qualitatively, we can draw the same conclusions from the results of a similar exercise for the Germany-based exchange rates as reported in the last two columns of Table 4. That is, the time series-based cointegrated VAR model under monetary model restrictions applied separately on each individual rate is never able to significantly outperform the random walk in terms of out-of-sample exchange rate forecasts. The monetary panel VEC model on the other hand has a significantly better exchange rate forecast performance than random walk forecasts at medium-to-long term forecasting horizons: 1 to 4 years in case of the Canada/Germany relationship, 3 to 4 years in case of the Japan/Germany relationship and 2 to 4 years in case of the U.S./Germany relationship. The slightly better out-of-sample performance of the monetary panel VEC model *vis-à-vis* Germany than relative to the whole EMU-zone could possibly be due to the recognition of the similarity between the institutional frameworks of the pre-1999 Bundesbank and the current ECB by currency market participants.

6 Conclusions

We investigate in this paper both the in-sample as well as the out-of-sample fit of the efficient markets-based monetary exchange rate model in order to assess the long-run efficiency of (pseudo-)Euro exchange rate movements. We follow both the standard pure time series-based cointegrated VAR approach and the panel VEC approach of Groen and Kleibergen (2003) to analyze the cointegration and long-run parameter restrictions of the monetary exchange rate model. In doing this we have to deal with the issue what data to use for the EMU area in the pre-1999 period, *i.e.* the period before EMU was established. As a result we both look at the bilateral relationships of Canada, Japan and the U.S. relative to the full EMU-area (with pre-1999 synthetic data spliced together with EMU data from 1999 onwards) and relative to Germany. Our empirical investigation indicates that modeling the three aforementioned exchange rates simultaneously within a panel structure is necessary in order to be able to find some consistency of exchange rate movements with the monetary exchange rate model irrespective of the choice of *numeraire*, albeit that we find a larger long-run impact of relative income than implied by theory.

The out-of-sample forecasting evaluation confirms that a multi-country panel data structure is necessary in order to find any empirical evidence *pro* long-run efficiency based on monetary fundamentals. Both for the Japanese, Canadian and U.S. bilateral exchange rate relationships relative to the whole EMU-zone as well as Germany the monetary fundamentals-based panel VEC approach yields superior exchange rate forecasts *vis-à-vis* random walk forecasts at horizons of two, three and four years. The usage of pure time series VEC models under monetary model restrictions, however, never helps to significantly outperform random walk-based exchange rate predictions. These results confirm that for monetary fundamentals-based exchange rate forecasting models to be successful, they should make use of techniques that have more power to properly identify the corresponding cointegration restrictions.

In summary we can conclude that the major Euro exchange rates have exhibited a proper low frequency comovement with corresponding monetary fundamentals. This is especially the case when we assume that the behaviour of Germany-based monetary fundamentals has been appropriate for both the pre-1999 period as well for the period since 1999 when EMU was established.

Appendices

A The Data

The main source for our dataset are the *International Financial Statistics* [IFS] from the IMF, and we use quarterly data which start in the first quarter of 1975 and end in the last quarter of 2000. We have chosen 1975.I as a starting date as only from this date onwards we were able to find reliable data for a significant amount of EMU member states. Our synthetic EMU series for the pre-EMU era, *i.e.* the period before 1999, are constructed based on the data for Austria, Finland, France, Germany, Italy, the Netherlands and Spain. For the construction of our synthetic EMU data series we make use of the method proposed by Beyer *et al.* (2001), which is based on constructing a weighted sum of the *growth rates* of the respective individual series.

In general terms we apply the Beyer *et al.* (2001) approach as follows. Suppose we have for our 7 EMU countries a series X, than the relative growth rate of the synthetic aggregate X_t equals

$$\Delta x_t = w_{1,t-1}\Delta x_{1,t} + \cdots + w_{7,t-1}\Delta x_{7,t}, \tag{A.1}$$

with $w_{1,t} + \cdots + w_{7,t} = 1$ and $\Delta x_{i,t} = \ln(x_{i,t}) - \ln(x_{i,t-1})$. When the variable X equals money the weights $w_{1,t}, \ldots, w_{7,t}$ are constructed based on the relative sizes of the national money stocks denoted in a common currency, *i.e.*

$$w_{i,t} = \frac{M_{i,t}/S_{i,t}}{\sum_{i=1}^{7}(M_{i,t}/S_{i,t})}, \quad i = 1, \ldots, 7, \tag{A.2}$$

where $S_{i,t}$ is the exchange rate of currency i in terms a common base currency and in our case we have chosen the U.S. dollar as the common base currency.[8] For all other variables we use the relative sizes of national real GDP expressed in a common currency,

$$w_{i,t} = \frac{Y_{i,t}/S_{i,t}}{\sum_{i=1}^{7}\left(Y_{i,t}/S_{i,t}\right)}, \quad i = 1, \ldots, 7. \tag{A.3}$$

We obtain the unlogged levels of the X series through

$$X_t = \frac{\bar{X}_T}{\exp(\sum_{j=1}^{T}\Delta x_j)} \exp\left(\sum_{j=1}^{t}\Delta x_j\right), \quad t = 1, \ldots, T, \tag{A.4}$$

where $\bar{X}_T$ is the first official EMU observation for variable X and therefore $T = 1999.I$. In compliance with the monetary exchange rate model X_t can either be the exchange rate relative to Canada, Japan and the U.S., money or real GDP. The synthetic pre-1999 Euro exchange rates of Canada, Japan and the U.S. are constructed through cross-rates for the 7 EMU countries based on the observed U.S. dollar rates (based on line 'ae' in IFS). As our measure of money we use the IMF definition of M3, *i.e.* 'money' (line 34 in IFS) plus 'quasi-money' (line 35 in IFS), and this measure of money is *not* seasonally adjusted. Note that for German M3 and real GDP we correct the series for the unification by fitting an autoregressive model to the log first differences with an unification dummy for 1991.I (and seasonal dummies in case of M3) included. As a next step we rescale the level of the pre-unification series with the estimated coefficient of the aforementioned unification dummy. Finally, real GDP is proxied by gross domestic product [GDP] in volume terms (line 99B in IFS). Hence, through (A.1) and (A.4) we splice the synthetic series for the exchange rates, real GDP (both based on (A.3)) and M3 (based on (A.2) over the pre-1999

[8] Beyer *et al.* (2001) show that the choice of base currency is not important as only the implied cross rates matter.

span, based on our 7 EMU member states data, together with the official EMU data for the period from 1999 onwards into one 1975.I-2000.IV series.

As motivated in the text, we can alternatively use German behaviour as a proxy for EMU behaviour in the pre-EMU era. For the DM exchange rates of Canada, Japan and the U.S. this implies that we use the observed U.S. dollar exchange rates, and calculate the DM exchange rates through cross-rates relative to the U.S.. For the period from 1999 onwards we use the Euro exchange rates for these three countries rescaled in DM terms through the Euro conversion rate of the DM (*i.e.* 1 Euro = 1.95583 DM).[9]

For Canada, Japan and the U.S. we compute the relative M3 money supplies *vis-à-vis* either the (synthetic-)EMU-area or Germany solely for the *entire* sample. The relative real income series are similarly constructed relative to either the (synthetic-)EMU-area or Germany solely.

[9] See the website of the European Central Bank, http://www.ecb.int.

B Computing P-values for the RMSE Ratios

The p-values used for testing H_0: RMSE ratio $= 1$ versus H_1: RMSE ratio < 1 for the estimated RMSE ratios in Table 4 are computed through parametric bootstrap procedures. We base the computation of these p-values on a system of *non-cointegrated* I(1) VAR models for the bilateral exchange rates and corresponding monetary fundamentals, *i.e.*

$$\begin{aligned}
\Delta s_{it} &= \epsilon_{it}^{s} \\
\Delta(m-m^*)_{it} &= \delta_{i1} + \sum_{j=1}^{p_i} \gamma_{i1j}\Delta(m-m^*)_{i,t-j} + \epsilon_{it}^{(m-m^*)} \\
\Delta(y-y^*)_{it} &= \delta_{i2} + \sum_{j=1}^{p_i} \gamma_{i2j}\Delta(y-y^*)_{i,t-j} + \epsilon_{it}^{(y-y^*)}, \quad i=1,\ldots,3 \quad t=1,\ldots,T,
\end{aligned} \tag{B.1}$$

where the innovations are assumed to comply with a multivariate normal distribution

$$(\epsilon_{1t}^{s}\ \epsilon_{1t}^{(m-m^*)}\ \epsilon_{1t}^{(y-y^*)}\ \cdots\ \epsilon_{3t}^{s}\ \epsilon_{3t}^{(m-m^*)}\ \epsilon_{3t}^{(y-y^*)})' \sim \mathrm{N}(\mathbf{0},\ \Omega),$$

and Ω is an unrestricted 9×9 covariance matrix. Both relative to the whole EMU-zone as well as solely Germany, we estimate (B.1) and the corresponding Ω across the $i = 1,\ldots,3$ bilateral relationships with feasible Generalized Least Squares [FGLS] on the quarterly 1975.I-2000.IV sample for Canada, Japan and the U.S. with the individual lag orders from Table 3.

In (B.1) the log exchange rate is, as in the standard approach, assumed to be generated by a random walk, whereas the log monetary fundamentals are I(1) but have no long-run relationship with each other or the log exchange rate. We therefore can use the estimated restricted VAR models (B.1) to generate artificial data that comply with H_0: RMSE ratio $= 1$. In order to compute the p-values we go through the following sequence of steps for each set of bilateral exchange rates of Canada, Japan and the US relative to the EMU-zone and Germany respectively:

1. We generate 154 pseudo-innovations $\epsilon_{1t}^{s}\ \epsilon_{1t}^{(m-m^{*})}\ \epsilon_{1t}^{(y-y^{*})}\ \cdots\ \epsilon_{3t}^{s}\ \epsilon_{3t}^{(m-m^{*})}\ \epsilon_{3t}^{(y-y^{*})}$ for (B.1) from the estimated multivariate normal distribution $\mathrm{N}(\mathbf{0},\ \hat{\Omega})$, where $\hat{\Omega}$ allows for contemporaneous correlation.

2. Using the pseudo-innovations from step 1 we generate 154 observations of non-cointegrated pseudo I(1) series s_{1t}, s_{2t}, s_{3t}, $(m-m^{*})_{1t}$, $(m-m^{*})_{2t}$, $(m-m^{*})_{3t}$, $(y-y^{*})_{1t}$, $(y-y^{*})_{2t}$ and $(y-y^{*})_{3t}$ from (B.1). The first 50 observations are deleted to correct for any initial-value bias. The remaining 104 observations are in compliance with our quarterly 1975.I-2000.IV sample.

3. We apply the recursive forecasting procedure from Section 5.1, based on either the pure time series-based monetary cointegrated VAR model or the monetary panel VEC model, on the 9 pseudo I(1) series from step 2 to generate RMSE ratios relative to random walk forecasts for each of our forecasting horizons $h = 1$, 4, 8, 12 and 16.[10]

These 3 steps are replicated 5,000 times. Using the resulting 5,000 artificial RMSE ratios for each horizon h we can compute for each estimated RMSE ratio the p-value for H_0: RMSE ratio $= 1$ versus H_1: RMSE ratio < 1.

[10] The recursive forecasting procedure from Section 5.1 involves a recursive AIC-based selection procedure for the lag order in either (4) or (5). As this is repeated in each iteration of our parametric bootstrap algorithm, we allow up to a certain for model uncertainty in the computation of the p-values.

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Table 1: Cointegration rank tests in the panel VEC models, 1975:I-2000:IV[a]

	syn. EMU data	German data	90%	95%	99%
LR[B(0)|A]	96.11**	95.29**	85.97	90.38	99.03
LR[B(1)|A]	29.35	42.92	46.03	49.27	55.76
LR[B(2)|A]	5.95	11.27	18.35	20.63	25.36
LR[C(1)|A]	39.92	48.61	50.70	54.15	61.03
LR[C(2)|A]	11.72	17.04	23.40	25.99	31.33

[a] Rows denoted with "LR[B(r)|A]" and "LR[C(r)|A]" contain likelihood ratio tests on the null of a common cointegrating rank value of r assuming heterogeneous and homogeneous cointegrating vectors respectively, see also Definition 4.1. The columns "90%" ("95%") ["99%"] are the 90% (95%) [99%] quantiles of the appropriate limiting distribution computed with the procedures from Groen (2002, Appendix) and $^{*(**)[***]}$ indicates a rejection of the null hypothesis at these quantiles.

Table 2: Normalized common cointegrating vectors in the panel VEC model, 1975:I-2000:IV[a]

	synthetic EMU data		*German data*	
	$\bar{\boldsymbol{\beta}}'_{\mathrm{C}}$	MON.EL.	$\bar{\boldsymbol{\beta}}'_{\mathrm{C}}$	MON.EL.
s_{it}	1	1	1	1
$(m_{it} - m^*_t)$	−0.50	−1	−0.83	−1
$(y_{it} - y^*_t)$	1.16	2.73	3.02	3.16
$\mathrm{LR}[(\beta_s + \beta_{(m-m^*)}) = 0]$	1.91	—	0.29	—
	(0.17)		(0.59)	
$\mathrm{LR}[(\beta_s - 3\beta_{(y-y^*)}) = 0]$	—	0.12	—	0.18
		(0.73)		(0.67)

[a] Column $\bar{\boldsymbol{\beta}}'_{\mathrm{C}}$ provide the normalized estimated common cointegrating vectors, and column "MON.EL" is the normalized cointegrating vectors with proportionality between the log exchange rate and the log relative money supply imposed. Row $\mathrm{LR}[(\beta_s + \beta_{(m-m^*)}) = 0]$ contains a likelihood ratio test of the restriction of a unity relative money elasticity on the respective non-normalized common cointegrating vector, whereas row $\mathrm{LR}[(\beta_s - 3\beta_{(y-y^*)}) = 0]$ reports a likelihood ratio test of the restriction that relative income elasticity equals -3 on the respective non-normalized common cointegrating vector under a unity relative money elasticity. The corresponding $\chi^2(1)$ p-values are reported in parentheses.

Table 3: Cointegration tests for the individual exchange rates, 1975:I-2000:IV[a]

	Lags	LR(0|3)	LR(1|3)	LR(2|3)	MON.EL.	INC.EL.	β
Canada/Euro	3	32.08*	9.05	2.35	0.14 (0.70)	7.24 (0.00)	(1 -1 2.06)'
Japan/Euro	3	26.74	11.75	2.06	–	–	–
U.S./Euro	1	27.77	9.86	2.59	–	–	–
Canada/Germany	3	41.12***	11.98	3.37	14.98 (0.00)	–	–
Japan/Germany	3	27.98	10.03	3.73	–	–	–
U.S./Germany	4	39.27**	6.96	2.19	15.75 (0.00)	–	–
90%		31.88	17.79	7.50			
95%		34.80	19.99	9.13			
99%		40.84	24.74	12.73			

[a] The column denoted with 'Lags' contains the order of first differences in (4) determined with AIC. LR($r|3$) denotes the values of the Johansen (1991) likelihood ratio test statistic for H_0: rank(Π) = r versus H_1: rank(Π) = 3 in (4). The symbol * (**) [***] indicates rejection of H_0 at the 10% (5%) [1%] significance level. The row '90%' ('95%') ['99%'] contains the asymptotic 90% (95%) [99%] quantile for LR($r|3$) under the null, see Johansen (1996, Table 15.2). The column denoted with "MON.EL." contains, if $r = 1$ is accepted, the likelihood ratio test of the restriction of a unity relative money elasticity on the non-normalized cointegrating vector (*i.e.* $\hat{\beta}_s + \hat{\beta}_{(m-m^*)} = 0$, see text). The column "INC.EL" reports for the Canada/Euro rate a likelihood ratio test for the restriction on the unnormalized cointegrating vector under unity relative money elasticity that relative income elasticity equals 0 (*i.e.* $\hat{\beta}_{(y-y^*)} = 0$). The corresponding $\chi^2(1)$ p-values are reported in parentheses. Column β finally reports for the Canada/Euro rate the normalized cointegrating vector.

Table 4: Forecasting evaluation of models based on monetary fundamentals, 1989:I-2000:IV[a]

	h	*vis-à-vis EMU* Ind.VEC	*vis-à-vis EMU* Pan.VEC	*vis-à-vis Germany* Ind.VEC	*vis-à-vis Germany* Pan.VEC
Canada	1	1.153	1.019	1.197	0.997
		(0.981)	(0.538)	(0.994)	(0.193)
	4	1.074	0.929	1.206	0.967
		(0.869)	(0.120)	(0.996)	(0.092)
	8	1.202	0.884	1.641	0.878
		(0.811)	(0.081)	(0.987)	(0.044)
	12	1.128	0.883	1.732	0.812
		(0.551)	(0.047)	(0.960)	(0.038)
	16	1.092	0.801	1.844	0.781
		(0.416)	(0.047)	(0.928)	(0.042)
Japan	1	1.045	1.057	0.969	1.024
		(0.394)	(0.414)	(0.091)	(0.185)
	4	0.989	1.016	0.972	1.009
		(0.168)	(0.351)	(0.113)	(0.281)
	8	1.013	0.947	1.071	1.013
		(0.235)	(0.122)	(0.435)	(0.243)
	12	1.128	0.676	1.117	0.845
		(0.521)	(0.010)	(0.501)	(0.067)
	16	1.128	0.714	1.161	0.868
		(0.471)	(0.023)	(0.527)	(0.099)
U.S.	1	1.095	1.023	1.215	1.019
		(0.898)	(0.600)	(0.998)	(0.487)
	4	1.020	0.956	1.125	0.999
		(0.471)	(0.430)	(0.965)	(0.346)
	8	1.061	0.956	1.199	0.883
		(0.439)	(0.178)	(0.785)	(0.047)
	12	1.052	0.915	1.279	0.802
		(0.369)	(0.138)	(0.773)	(0.032)
	16	0.882	0.639	1.411	0.663
		(0.117)	(0.010)	(0.793)	(0.013)

[a] The table reports RMSE ratios of monetary fundamentals-based versus random walk-based exchange rate predictions; simulated p-values for H_0 : ratio = 1 versus H_1: ratio < 1 are in parentheses (see Appendix B); "h" indicates the forecasting horizons (in quarters); "Ind.VEC" report the outcomes for the individual country VEC models and "Pan.VEC" those for the panel VEC model.