Macroeconomic Fundamentals and Exchange Rates: A Non-Parametric Cointegration Analysis*

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December 2, 2004

Abstract

This paper examines in a non-parametric setup whether a long-run relationship exists between monetary fundamentals and the dollar spot exchange rates for 19 countries. Although the Johansen's parametric approach failed to retrieve a long-relationship for any of the countries considered, the Bierens (1997a) non-parametric approach suggests that there is one cointegrating relationship for the majority of the countries considered. In addition, the [1,-1] cointegrating vector between the fundamentals and the log-level of the dollar exchange rate could not be rejected in the non-parametric formulation.

Keywords: Cointegration; Unit Root; Non-parametric Test.

JEL Classification: C14; E31; F31.

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*I would like to thank Mark P. Taylor and Lucio Sarno for helpful comments and suggestions.

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1 Introduction

Several authors have recently examined whether nominal exchange rate returns can be forecasted by monetary fundamentals. Mark (1995), Chinn and Meese (1995), Chen and Mark (1996), among others, provided empirical evidence that the long-horizon regression approach can be used in order to predict the spot exchange rate returns at longer horizons. However, according to these studies the monetary fundamentals cannot outperform in forecasting a naive random walk model of the spot exchange rate for shorter horizons, a finding first reported by Meese and Rogoff (1988).

The empirical success of the long horizon regression employed by Mark (1995) has been heavily criticized by subsequent research on the grounds of the methodology used and the underlined assumptions made. Specifically, the long-horizon regression includes, as a regressor, the deviation of the log-level of the spot exchange rate from its fundamental value, which is assumed to be stationary in Mark's setup. However, Berkowitz and Giorgianni (2001)-hereafter referred to as BG- argue that if this deviation is not stationary the long-horizon regression will be a spurious regression, while the critical values of the tests involved in the forecasting evaluation generated under the assumption of cointegration will be incorrect. Other researchers, like Kilian (1999), emphasized that the empirical results will depend on the assumed bootstrap data generating process (DGP) applied in the non-parametric bootstraps performed by Mark (1995). Subsequent empirical results reported by Faust, Rogers and Wright (2003) point out that the forecasting performance of the monetary model crucially depends on data revisions and on the use of real-time forecasts of future fundamentals compared to actual future fundamentals.

Mark and Sul (2001) formally tested for a panel cointegration between exchange rates and their fundamental values in a panel of 19 countries using a dynamic ordinary least squares approach. Their findings indicated that there is cointegration. However, their approach does not consider the possibility that the DGP of the deviation of the log-level of spot exchange rate from its fundamental value might be non-linear and/or the short-run dynamics towards the long-run relationship might be non-linear, as well, due to transaction costs' considerations. Other studies performed by Taylor and Peel (2000) and Kilian and Taylor (2003) do not distinguish between non-linearity in the DGP compared to non-linearity in the short-run adjustment. The only attempt that we are aware of is by Sekioua (2003), who examined whether a non-linear threshold autoregressive unit root exists in the deviation of the spot exchange from its fundamental value by means of tests proposed by Caner and Hansen (2001).

In addition, it is very likely that the long-run equilibrium level implied by the monetary model will vary over time depending on the exchange rate regime in operation. Sarno, Valente and Wohar (2004) provide supportive empirical evidence of the time-varying long-run equilibrium level for six industrialized countries after fitting a Markov-switching vector error correction model (VECM) with time varying parameters for a data sample starting in late 1800s.

The current study attempts to examine the cointegration between the spot exchange rate and its fundamental value at a non-parametric framework in order to account for the possibility of a non-linear DGP. The framework that we employ does not assume a specific DGP, while it enables us to test explicitly for the validity of the [1,-1] cointegrating vector between the spot and its fundamental exchange rate value. Thus, our approach is general enough without imposing a specific DGP and without ruling out the possibility of a non-linear DGP. To that end, we considered the non-parametric unit root test proposed by Breitung (2002) and the non-parametric cointegration analysis proposed by Bierens (1997a). The former has been employed by Maki (2003) in order to examine whether non-parametric cointegration exists between the nominal interest rate and the expected inflation in Japan. Coakley and Fuertes (2001) implemented Bierens' approach in order to test for a non-parametric cointegration between the spot exchange rate and its fundamental value based on the purchasing power parity (PPP).

Using quarterly data on industrial production indices, monetary aggregates and the spot exchange rate for 19 countries, we could not reject the null hypothesis of a unit root for the series considered across countries on the basis of Breitung's non-parametric test. The latter finding implies that it might be the case that a non-linearity characterizes the underlined DGP. Furthermore, we provide empirical evidence in favour of a single mean reversion process of the deviation of the spot exchange rate from its fundamental value on the basis of the Bierens' (1997a) non-parametric cointegration analysis for the majority of the countries considered. Most importantly, the [1,-1] cointegrating vector between the fundamental vale and the spot exchange rate could not be rejected for none of the countries examined, even when the non-parametric approach indicated that there are two instead of one cointegrating vectors.

In contrast, the Johansen approach provided limited support of the mean reversion hypothesis for the countries examined. We attributed the obvious discrepancy of the two tests on the possible existence of non-linear DGPs and/or non-linear short-run dynamics due to transaction costs. We also considered, as an additional explanation, the size distortions that the Johansen's approach experiences in the event of an erroneous choice of optimal lag length.

The remainder of the paper proceeds as follows. In section two, we briefly discuss the monetary model of the exchange rate determination, along with its empirical formulation. In section three, we demonstrate the non-parametric unit root and cointegration tests used in this study. Section four provides the dataset description and presents the empirical results obtained after the implementation of the parametric and non-parametric approaches considered. Finally, section five concludes the paper.

2 The Monetary Model for the Exchange Rate Determination

The starting point of the monetary model is the definition of the exchange rate as the relative price of two monies, while modelling the money and supply side of these two monies is at the heart of the monetary monetary model. The monetary equilibria at home and abroad, respectively, are provided by

$$m_t = p_t + ky_t - \gamma i_t \tag{1}$$

$$m_t^* = p_t^* + ky_t^* - \gamma^* i_t^* \tag{2}$$

where m_t, p_t, y_t and i_t stand for the log-levels of the money supply, the price level, the income and the nominal interest rate, respectively at time t; k, γ are positive constants, while asterisk denotes foreign variables and parameter values. In addition to the monetary equilibria at home and abroad, it is assumed that PPP holds continuously in time implying that the following relationship holds

$$s_t = p_t - p_t^* \tag{3}$$

where s_t is the log-level of the spot exchange rate defined as the domestic price of a unit of foreign currency. The domestic money supply will determine the domestic price level and subsequently the spot exchange rate. After subtracting (2) from (1), solving for $p_t - p_t^*$, substituting in (3) and assuming $k = k^*$ and $\gamma = \gamma^*$, we obtain

$$s_t = (m_t - m_t^*) - k(y_t - y_t^*) + \gamma(i_t - i_t^*) \tag{4}$$

Equation in (4) is the monetary model equation for the determination of the exchange rate according to which an increase at the domestic supply will induce a depreciation of the domestic currency, while an increase at the level of the domestic impact will appreciate the domestic currency. The latter is true since the price level need to fall in order to maintain equilibrium at the domestic money market, leading to an appreciation of the domestic currency through (3).

The monetary model in (4) was considered to be an appealing device in forecasting spot exchange rates. However, the early attempts by Meese and Rogoff (1983a, 1983b) towards that direction were not fruitful pointing that the naive non-change forecast of the spot exchange rate generated by a random walk could not be outperformed by (4). Further research by Mark (1995) indicated that the monetary model can be used to predict the spot exchange rate in the long-run. Specifically, Mark (1995) considered the in and out-of-sample forecasting performance of the monetary model by estimating the following formulation of (4) assuming $i_t - i_t^* = 0$

$$s_{t+\ell} - s_t = c_{0\ell} + c_{1\ell}(f_t - s_t) + \upsilon_{t+\ell,t} \tag{5}$$

where $f_t = [(m_t - m_t^*) - (y_t - y_t^*)]$ is called the fundamental value of the exchange rate under long-run money neutrality (k=1). Equation (5) is an error correction equation where the term $(f_t - s_t)$ can be considered the error term. Therefore, (5) implies that although the exchange rate and its fundamental value individually are I(1), their linear combination is stationary I(0) with a cointegrating vector [1,-1]. BG pointed out that the assumption of cointegration between the spot exchange rate and its fundamental value, when in fact such a relationship does not exist, will produce incorrect critical values for the test statistics considered. The latter is true because in this event (5) will be a spurious regression. Hence, it is necessary to test whether there is a long-run relationship between s_t and f_t and subsequently estimate (5). In order to perform this task at the present study, we use parametric and non-parametric testing procedures. The latter approach does not assume a specific DGP but it is general enough to account for the possibility of a non-linear DGP that will cause non-linear short-run dynamics towards the long-run relationship.

3 Non-Parametric Cointegration Formulation

The non-parametric approach that we implemented in this study consists of three steps. First, we test whether a non-parametric unit root is present at the log-level of the spot exchange rate and its fundamental value by means of Breitung (2002) test. Second, we test whether a non-parametric long-run relationship exists among s_t and f_t in the vector $x_t = [f_t, s_t]'$. For this purpose, we employ Bierens' (1997a) non-parametric approach. Finally, we test whether the [1,-1] cointegrating vector can be accepted for f_t and s_t.

3.1 Breitung's (2002) Non-Parametric Unit Root Test

In addition to the Phillips-Perron (1988) unit root test, we employed a non-parametric unit root test proposed by Breitung (2002). The latter is free of problems concerning the specification of the short-run dynamics and the estimation of nuisance parameters. This non-parametric unit root test decomposes the univariate process $\{x_t\}_1^n$ as $x_t = \delta' d_t + u_t$ where $\delta' d_t$ is the deterministic part with $\delta' = [\delta_1, \delta_2]$ and $d_t = [1, t]'$, while u_t is the stochastic part. If the deterministic part was not present, x_t would be consistent to the stochastic part u_t. Under the null x_t is I(1), if as $n \to \infty$, $n^{-1/2} x_{[\alpha n]} \Rightarrow \sigma W(\alpha)$, for $\sigma > 0$ the constant long-run variance, $W(\alpha)$ is a Brownian motion and [.] is the integer part. The expression for u_t is not specified giving rise to a general DGP.

In order to avoid the specification of short-run dynamics towards stationarity and the computation of an estimate for σ, Breitung proposed the following test statistic which is a variance ratio test

$$\hat{\rho}_n = \frac{n^{-4} \sum_{t=1}^{n} \hat{U}_t^2}{n^{-2} \sum_{t=1}^{n} \hat{\varepsilon}_t^2} \tag{6}$$

where $\hat{\varepsilon}_t$ is the Ordinary Least Squares (OLS) residuals such as $\hat{\varepsilon} = x_t - \hat{\delta}' d_t$ and $\hat{U}_t$ is the partial sum such that $\hat{U}_t = \hat{\varepsilon}_1 + \ldots + \hat{\varepsilon}_t$. In the event that x_t is I(0), the test statistic $\hat{\rho}_n$ converges to zero. Breitung provided simulation evidence that non-parametric unit root test outperform the standard parametric tests.

3.2 The Bierens' (1997a) Test

The non-parametric cointegration test proposed by Bierens (1997a) is based on the solutions to a generalized eigenvalue problem like Johansen's approach. However, the corresponding eigenvalue problem at Bierens' setup is formulated on the basis of two random matrices that are constructed without reference to the DGP. Specifically, Bierens assumes an observable q-variate process x_t for t=1,...,n generated from:

$$x_t = \pi_0 + \pi_1 t + y_t \tag{7}$$

where $\pi_0(q \times 1)$ and $\pi_1(q \times 1)$ are optional and trend terms, while y_t is a zero-mean unobservable process such that Δy_t is stationary. No further specification for the DGP with respect to x is required. The test is based on the following pair of random matrices that are the weighted sums of the series considered in levels and first differences, respectively

$$\begin{aligned} \hat{A}_m &= \frac{8\pi^2}{n}\sum_{k=1}^{m} k^2 (\frac{1}{n}\sum_{t=1}^{n} cos(2k\pi(t-0.5)/n)x_t) \qquad (8) \\ &\times (\frac{1}{n}\sum_{t=1}^{n} cos(2k\pi(t-0.5)/n)x_t)' \\ \hat{B}_m &= 2n\sum_{k=1}^{m}(\frac{1}{n}\sum_{t=1}^{n} cos(2k\pi(t-0.5)/n)\Delta x_t) \qquad (9) \\ &\times (\frac{1}{n}\sum_{t=1}^{n} cos(2k\pi(t-0.5)/n)\Delta x_t)' \end{aligned}$$

The ordered generalized eigenvalues $\hat{\lambda}_{1,m} \geq \ldots \geq \hat{\lambda}_{q,m}$ of the characteristic equation $det[P_n - \lambda Q_n] = 0$ are having similar properties to the corresponding eigenvalues in the Johansen framework justifying their inclusion for testing the cointegration rank r. In order for this to happen, P_n should be defined as $P_n = \hat{A}_m$, while Q_n should be chosen as $Q_n = \hat{B}_m + n^{-2}\hat{A}_m^{-1}$. Bierens (1997a) suggests that the λ_{min} test $\hat{\lambda}_{q-r_0,m}$ to test the null hypothesis $r = r_0$ against the alternative $r = r_0 + 1$. Bierens tabulates the critical values for the distribution under the null after Monte Carlo simulations. Parameter m is also tabulated for different significance levels and for various values of r_0 and q in such a way that the lower end of the power of the test is maximized.

In order to consistently estimate the rank r, the $g_m(r_0)$ is proposed, as well. This test takes the form:

$$\hat{g}_m(r_0) = \begin{cases} (\prod_{k=1}^{q} \hat{\lambda}_{k,m})^{-1} & \text{if } r_0 = 0 \\ (\prod_{k=1}^{q-r_0} \hat{\lambda}_{k,m})^{-1}(n^{2r_0}\prod_{k=q-r_0+1}^{q} \hat{\lambda}_{k,m}) & \text{if } r_0 = 1,\ldots,q-1 \\ n^{2q}\prod_{k=1}^{q} \hat{\lambda}_{k,m} & \text{if } r_0 = q \end{cases} \tag{10}$$

When $r_0 = q$ the optimal value of m is given by $m = q$, while when $r_0 < q$ the optimal value of m is given by tabulated values for different significance levels and different combinations of r,q. The rank r is consistently estimated and it is given by $\hat{r} = argmin_{r_0 \leq q} \quad g_m(r_0)$.

Having determined the dimension of the cointegration space, linear restrictions on the cointegrating vectors could be tested. For this purpose non-parametric tests could be used that employ the ordered eigenvalue solutions to the following problem:

$$det[H'\hat{A}_m H - \lambda H'(\hat{A}_m + n^{-2}\hat{A}_m^{-1})^{-1}H] = 0 \tag{11}$$

where H spans a subspace of cointegrating vectors. The critical values for the corresponding trace and lambda-max tests used to test the restrictions are given in Bierens (1997a) with m=2q.

4 Data Description and Empirical Results

Our dataset includes quarterly time series observations from 1975:Q1 to 1997:Q4 for 19 countries, namely Austria, Australia, Belgium, Canada, Denmark, Finland, France, Germany, United Kingdom (UK), Greece, Italy, Japan, Korea, the Netherlands, Norway, Spain, Sweden, Switzerland and the United States (US). Data availability determined the ending and starting points of our sample. Data on nominal spot exchange rate were end of period observations retrieved from the IFS CD-ROM (line code AE). The industrial production index (IFS line code 66) proxy the national income for the countries considered, while money was proxy by the money stock (line code 34) plus quasi money (line code 35) for all the countries, except Norway and Sweden due to availability constraints. M2 was used for Norway and M3 for Sweden from OECD's Main Economic Indicators. The resulting series for the monetary aggregates were not seasonally adjusted and we did so by implementing a one-sided moving average of the current observation plus 3 lags of the corresponding series.

In order to test for the presence of a unit root on the nominal exchange rate and its fundamental value s_t and f_t, respectively, we considered the Phillips-Perron test (PP) and the Breitung (2002) test ($\hat{\rho}_n$). With regard to the PP test, we used the truncation parameter $[cn]^k$ in the heteroskedasticity consistent covariance matrix of Newey-West with c=5 and k=0.2, following Bierens (1997b). We account for the potential size distortions that the PP test might experience in finite samples, by computing the p-values following 1000 simulations of a Gaussian AR(1) process for the first difference of every series considered. Testing results from the unit root analysis are reported at Table 1.

[Table 1]

The PP test indicated that s_t and f_t were both I(1) processes for all the countries considered on the basis of the p-values for the PP test statistic[1]. However, f_t for Norway was found to be I(0) for both simulated and non-simulated p-values. The implementation of the Breitung test[2] produced test statistics that are not statistically significant at the 5 percent significance level, implying that all the series were I(1) processes consistent to the empirical outcome of the PP test. Since the series considered are individually I(1) processes there might be the case that a linear combination of them is an I(0) process.

In order to test whether a long-run stationary relationship exists between f_t and s_t, we have to perform a cointegration analysis at a parametric and a non-parametric context. The parametric test that we consider is the Johansen cointegration test that estimates the following VECM

$$\Delta x_t = \mu + \Pi x_{t-1} + \sum_{j=1}^{L} \Gamma_j \Delta x_{t-j} + \Phi d_t + \epsilon_t \qquad \text{for t=1,...,n} \tag{12}$$

where Π is the error correction term; Γ and Φd_t is the coefficient vector and the deterministic vector, respectively; ϵ is the error vector. In order to chose the optimal lag length L we performed successive estimation of the corresponding to (12) vector autoregression (VAR) for different lag lengths and chose the L that rendered the minimum value of the Akaike information criterion[3]. Estimation results based on the Johansen test are reported in Table 2.

[Table 2]

Empirical results in Table 2 show that the cointegrating relationship between the log-level of the spot exchange rate and its fundamental value is rejected for 14 out of the 18 countries considered, implying that the use of (5) is going to be a spurious regression. Specifically, since the term $f_t - s_t$ is not I(0) the left-hand side of (5) will be more persistent, making both sides of the equation very persistent. In that case, (5) will become spurious making any inference based on that equation to be erroneous. Due to the spurious regression property of the aforementioned equation, the critical values of the test

[1] The series considered were found to be stationary at their first differences. These results are not reported here for the sake of brevity, but they are available on request.

[2] We computed Breitung's test using the Bierens (2004) *EasyReg International* package.

[3] Different information criteria did not alter the L chosen on the basis of the Akaike criterion.

statistics that will be used in order to evaluate the forecasting performance of (5) will be incorrect, as underlined by BG.

Furthermore, Johansen's cointegrating analysis indicated that there is one cointegrating relationship between f_t and s_t, only for Canada and Germany. However, the [1,-1] cointegrating vector for these countries could not be accepted as the involved likelihood ratio test was equal to 12.53 and 3.90, respectively, with corresponding p-values equal to 0 and 4 percent, respectively. In addition, Table 2 reports that there are two cointegrating relationships between f_t and s_t for Austria and the Netherlands. After imposing that the rank r is equal to 1 for Austria and the Netherlands, we tested whether the [1,-1] cointegrating vector could be accepted or not, by means of a likelihood ratio test. The latter pointed out that the former cointegrating vector could not be rejected, implying that for these countries equation (5) could be used in order to evaluate the forecasting performance of the monetary model with respect to the log-level of the spot exchange rate without the risk of estimating a spurious regression.

Bierens' test results[4], reported in Table 3, provide empirical evidence in favour of a non-parametric cointegrating relationship between f_t and s_t for 13 out of the 18 countries considered, while the [1,-1] cointegrating vector is not rejected for these 13 countries. Although the λ_{min} test indicated that there are two cointegrating relationships, the $g_m(r_0)$ test consistently estimated a number of cointegrating relationships equal to 1 for these 13 countries. The $g_m(r_0)$ test is the test that Bierens (1997a) suggests as a double-check on the findings from the λ_{min} test. Furthermore, the [1,-1] cointegrating vector was not rejected for none of the countries for which one cointegrating relationship was found on the basis of the $g_m(r_0)$ test.

[Table 3]

Moreover, for Austria, UK, Japan, Norway and Switzerland the λ_{min} and the $g_m(r_0)$ tests produced results consistent to two cointegrating relationships. In order to test whether the [1,-1] cointegrating vector could be rejected or not for these countries, we imposed a rank equal to 1 and we tested the corresponding linear restriction by computing the related trace-test. The latter was not able to reject the [1,-1] linear restriction for the variables at the x vector, implying that the [1,-1] cointegrating vector is a vector that can span the cointegrating space.

[4]We performed Bierens non-parametric cointegration analysis using the Bierens (2004) *EasyReg International* package.

The discrepancy in the inference from the Johansen and the Bierens' tests can be attributed to the fact that it is very likely that the DGP of the variables involved is non-linear resulting into non-linear short-run dynamics towards the long-run. Several studies in the literature demonstrate that it is very likely the short-run adjustment to be exhibit a non-linear process. Taylor and Peel (2000) and Kilian and Taylor (2003) demonstrate that the more the exchange rate is misaligned compared to its fundamental value the more market participants and policymakers will push the exchange rate back to its fundamental value. This is justified by the presence of transaction costs with reference to the arbitrage in international goods market.

In addition, it might be the case that the DGP itself is a non-linear process. Towards that direction, Sekioua (2003) applied non-linear threshold autoregressive unit root tests suggested by Caner and Hansen (2001) in order to examine whether $f_t - s_t$ is mean reverting. Sekioua's empirical finding indicated that linearity and nonstationarity are rejected implying a non-linear mean reversion that depends on the size of the deviation $f_t - s_t$.

Furthermore, we need to be aware that the λ_{min} test might exhibit a low power for small samples compared to Johansen's λ_{max} test on the basis of simulation evidence reported in Bierens (1997a). In that event Bierens suggests that a full parametric approach might work better than a non-parametric approach. The same simulation evidence underlined the fact that once the Johansen approach is corrected for size distortions experienced for small L, it will have a greater power. However, without the correction of these size distortions generated by an incorrect choice of L for (12), the Bierens test will outperform the Johansen test in terms of its statistical power.

5 Conclusion

We attempted in this study to examine whether a long-run relationship exists between the fundamental value of the spot exchange rate and the log-level of the latter at a parametric and a non-parametric context. The two cointegration tests that we considered were the Johansen test and the Bierens (1997a) non-parametric test. Using quarterly data on industrial production indices, monetary aggregates and the spot exchange rate, we provided empirical evidence in favour of a single mean reversion of the deviation $f_t - s_t$ on the basis of non-parametric cointegration analysis for the majority of the countries considered. Most importantly, the [1,-1] cointegrating vector between f_t and s_t could not be rejected for none of the countries examined even when the non-parametric approach indicated that there are two instead of one cointegrating vectors.

In contrast, the Johansen approach provided limited support of the mean reversion hypothesis for the countries examined. We attributed the obvious discrepancy of the two tests on the possible existence of non-linear DGPs and/or non-linear short-run dynamics due to transaction costs. We also considered as an additional explanation the size distortions that the Johansen's approach experiences in the event of an erroneous choice of optimal lag length.

The analysis performed in the current study can be extended in order to exploit simultaneously the cross section and time series information[5]. To that end, we could replicate the non-parametric analysis for the panel of the 19 countries considered by modifying the Bierens (1997a) test in order to correspond to a panel cointegration test. This could be done by augmenting (7) by a common time effect θ_t, constructing the corresponding random matrices $\hat{A}_m$ and $\hat{\beta}_m$ for the individual countries and averaging them across countries. This kind of panel data estimator is the group-mean estimator. An alternative estimator could be the pooled estimator where the random matrices are jointly estimated for the countries considered. Of course, in both cases we will have to modify accordingly the trace-test that will be used in order to test the possible linear restrictions on the cointegrating vector.

At a parametric level, Mark and Sul (2001) focused on a dynamic OLS estimator and they examined the small sample properties of weighted and non-weighted panel estimators and of the mean group estimator. They reported findings in favour of a cointegrating rela-

[5] Work in progress of the author.

tionship between f_t and s_t, while they re-established the ability of monetary fundamentals to forecast future exchange rate returns using forecasts generated by panel regressions.

Finally, the asymptotic properties of the two tests considered, namely the Johansen and the Bierens tests, should be examined in order to identify the proportion of times that the two tests individually render the correct number of cointegrating relationships in the presence of alternative non-linear short-run dynamics. For this purpose a parametric bootstrap exercise should be performed where the bootstrap DGP will be determined by a VECM with non-linear short-run dynamics determined by either exponential, logistic smooth transition functions or functions that incorporate a threshold. We leave these possibilities for a future research agenda.

References

[1] Berkowitz, J. and L. Giorgianni, (2001), "Long-Horizon Exchange Rate Predictability?", *The Review of Economics and Statistics*, 83, pp. 81-91.

[2] Bierens, H., (2004), "EasyReg International", Department of Economics, Pennsylvania State University, University Park, PA.

[3] Bierens, H., (1997a), "Nonparametric Cointegration Analysis", *Journal of Econometrics*, 77, pp. 379-404.

[4] Bierens, H., (1997b), "Testing the Unit Root With Drift Hypothesis Against Nonlinear Trend Stationarity With an Application to the US Price Level and Interest Rate", *Journal of Econometrics*, 81, pp. 29-64.

[5] Breitung, J., (2002), "Nonparametric Tests for Unit Roots and Cointegration", *Journal of Econometrics*, 108, pp. 343-363.

[6] Caner, M. and B. Hansen, (2001), "Threshold Autoregression with a Unit Root", *Econometrica*, 69., pp. 1555-1596.

[7] Chen, J. and N. Mark, (1996), "Alternative Long-Horizon Exchange Rate Predictors", *International Journal of Finance and Economics*, 1, pp. 229-250.

[8] Chinn, M.D. and R.A., Meese, (1995), "Banking on Currency Forecasts: How Predictable is Change in Money?", *Journal of International Economics*, 38, pp. 53-69.

[9] Coakley, J. and A. Fuertes, (2001), "Nonparametric Cointegration Analysis of Real Exchange Rates", *Applied Financial Economics*, 11, pp. 1-8.

[10] Faust, J., J. Rogers and J.H. Wright, (2003), "Exchange Rate Forecasting: the Errors we've Really Made", *Journal of International Economics*, 60, pp. 35-59.

[11] Kilian, L., (1999), "Exchange Rates and Monetary Fundamentals: What do we Learn from Long-Horizon Regressions?", *Journal of Applied Econometrics*, 15, pp. 491-510.

[12] Kilian, L. and M.P. Taylor, (2003), "Why is it So Difficult to Beat the Random Walk Forecast of Exchange Rates?", *Journal of International Economics*, 60, pp. 85-107.

[13] Maki, D., (2003), "Nonparametric Cointegration Analysis of the Nominal Interest Rate and Expected Inflation Rate", *Economics Letters*, 81, pp. 349-354.

[14] Mark, N.C., (1995), "Exchange Rates and Fundamentals: Evidence on Long-Horizon Predictability", *The American Economic Review*, 85, pp. 201-218.

[15] Mark, N. and D. Sul, (2001), "Nominal Exchange Rates and Monetary Fundamentals: Evidence from a Small Post-Bretton Woods Panel", emphJournal of International Economics, 53, pp. 29-52.

[16] Meese, R. and K. Rogoff, (1988), "Was It Real? The Exchange-Rate Interest Differential Relation Over the Modern Floating-Rate Period", *Journal of Finance*, 43, pp. 933-948.

[17] Meese, R. and K. Rogoff, (1983a), "Empirical Exchange Rate Models of the Seventies: Do they Fit Out of Sample?", *Journal of International Economics*, 14, pp. 345-373.

[18] Meese, R. and K. Rogoff, (1983b), "The Out-of-Sample Failure of Empirical Exchange Rate Models: Sampling Error or Misspecification?", in Jacob Frankel. ed., *Exchange Rates and International Macroeconomics*. Chicago: NBER and University of Chicago Press.

[19] Neely, C.J. and L. Sarno, (2002), "How Well do Monetary Fundamentals Forecast Exchange Rates?", *Federal Reserve Bank of Saint Louis Review*, 88, pp. 51-74.

[20] Phillips, P.C.B. and P. Perron, (1988), "Testing for a Unit Root in Time Series Regression, *Biometrika*, 75, pp. 335-346.

[21] Sekioua, S.H., (2003), "The Nominal Exchange Rate and Monetary Fundamentals: Evidence from Nonlinear Unit Root Tests", *Economics Bulletin*, 6, pp. 1-13.

[22] Sarno, L., G. Valente, and M.E. Wohar, (2004), "Monetary Fundamentals and Exchange Rate Dynamics under Different Nominal Regimes", *Economic Inquiry*, 42, pp. 179-193.

[23] Taylor, M.P. and D. Peel, (2000), "Nonlinear Adjustment, Long-Run Equilibrium and Exchange Rate Fundamentals", *Journal of International Money and Finance*, 19, pp. 33-53.

Table 1: Unit Root Test Results

		PP^a	BR^b
Australia	s	0.6900 (0.6390)[c]	0.0750
	f	0.9900 (0.9740)	0.0921
Austria	s	0.3700 (0.2970)	0.0564
	f	0.4100 (0.2420)	0.0254
Belgium	s	0.2600 (0.1190)	0.0172
	f	0.9700 (0.9040)	0.0546
Canada	s	0.5300 (0.5290)	0.0460
	f	0.7900 (0.7530)	0.0562
Denmark	s	0.3100 (0.1730)	0.0154
	f	0.5800 (0.7610)	0.0862
Finland	s	0.3000 (0.1240)	0.0157
	f	0.7600 (0.6330)	0.0174
France	s	0.3100 (0.1960)	0.0166
	f	0.7900 (0.8020)	0.0516
Germany	s	0.3500 (0.2710)	0.0533
	f	0.6900 (0.6120)	0.0214
Greece	s	0.9100 (0.8490)	0.0957
	f	0.9400 (0.8770)	0.0999
Italy	s	0.4600 (0.4630)	0.0536
	f	0.7800 (0.7470)	0.0929
Japan	s	0.6800 (0.6150)	0.0877
	f	0.6100 (0.5980)	0.0257
Korea	s	0.9300 (0.8280)	0.0561
	f	0.9800 (0.9650)	0.0708
Netherlands	s	0.3500 (0.2350)	0.0469
	f	0.3400 (0.3940)	0.0137
Norway	s	0.3600 (0.3050)	0.0326
	f	0.0100* (0.0210)*	0.0208
Spain	s	0.4600 (0.4240)	0.0458
	f	0.8700 (0.8890)	0.0910
Sweden	s	0.5300 (0.3880)	0.0481
	f	0.6400 (0.6980)	0.0399
Switzerland	s	0.3400 (0.2840)	0.0624
	f	0.5100 (0.6280)	0.0157
UK	s	0.1600 (0.0770)	0.0231
	f	0.9400 (0.8100)	0.0935

Notes: [a]Phillips-Perron ρ-test with critical value -14 at the 5 percent significance level.
[b]Breitung's $\hat{\rho}_n$ test with critical value 0.01 at the 5 percent significance level.
[c]Simulated p-values in parenthesis.
*Significant at the 5 percent significance level.

Table 2: Johansen Cointegration Analysis

	L^a	λ_{trace} $r \leq 0$ $r \leq 1$	λ_{max} $r \leq 0$ $r \leq 1$	$H_0 : \beta' = (1, -1)^b$
Australia	2	7.23	6.83	-
		0.40	0.40	
Austria	2	11.32	5.79	0.17 [0.67]
		5.54*	5.54*	
Belgium	4	7.00	6.49	-
		0.50	0.50	
Canada	4	21.65*	20.67*	12.53 [0.00]*
		0.98	0.98	
Denmark	4	7.34	6.28	-
		1.06	1.06	
Finland	4	8.15	7.76	-
		0.39	0.39	
France	2	8.75	7.87	-
		0.88	0.88	
Germany	1	15.62*	14.52*	3.90 [0.04]*
		1.10	1.10	
Greece	1	4.44	2.63	-
		1.81	1.81	
Italy	2	8.43	4.92	-
		3.51	3.51	
Japan	3	13.33	11.64	-
		1.68	1.68	
Korea	4	10.85	7.25	-
		3.59	3.59	
Netherlands	1	10.60	6.37	1.99 [0.15]
		4.23*	4.23*	
Norway	1	14.91	12.23	-
		2.68	2.68	
Spain	4	8.72	6.80	-
		1.92	1.92	
Sweden	4	10.14	9.56	-
		0.57	0.57	
Switzerland	1	6.87	6.18	-
		0.69	0.69	
UK	3	5.02	4.78	-
		0.24	0.24	

Notes: [a] Lag order sellected by the Akaike information criterion.
[b]The likelihood ratio test value for the linear restriction tested with p-values in brackets.
*Significant at the 5 percent significance level.

Table 3: Nonparametric Cointegration Analysis

	λ^a_{min} r=0/r=1 r=1/r=2	$g_m(r_0)$ $r_0 = 0, 1, 2$	$H_0 : \beta' = (1, -1)^b$
Australia	0.30×10^{-3}*	35.03×10^{4}	1.06
	2.60×10^{-2}*	$\mathbf{31.06 \times 10^{0c}}$	
		17.90×10^{1}	
Austria	0.06×10^{-3}*	19.72×10^{5}	1.11
	0.90×10^{-2}*	49.37×10^{0}	
		$\mathbf{31.80 \times 10^{0}}$	
Belgium	1.84×10^{-2}	45.09×10^{1}	1.94
	26.97×10^{-2}	$\mathbf{24.14 \times 10^{1}}$	
		13.91×10^{4}	
Canada	0.11×10^{-2}*	71.75×10^{12}	1.04
	2.10×10^{-2}*	$\mathbf{24.80 \times 10^{-8}}$	
		87.43×10^{-8}	
Denmark	0.85×10^{-3}*	25.75×10^{4}	1.13
	2.21×10^{-2}*	$\mathbf{62.44 \times 10^{0}}$	
		24.35×10^{1}	
Finland	0.30×10^{-3}*	66.14×10^{3}	1.13
	4.48×10^{-2}*	$\mathbf{59.49 \times 10^{0}}$	
		94.85×10^{1}	
France	0.30×10^{-3}*	33.19×10^{4}	1.19
	3.93×10^{-2}*	$\mathbf{15.39 \times 10^{0}}$	
		18.90×10^{1}	
Germany	0.79×10^{-2}*	11.30×10^{3}	1.16
	2.37×10^{-2}*	$\mathbf{12.43 \times 10^{2}}$	
		55.49×10^{2}	
Greece	0.10×10^{-3}*	10.63×10^{5}	1.05
	6.18×10^{-2}	$\mathbf{19.47 \times 10^{-1}}$	
		58.98×10^{0}	
Italy	0.18×10^{-2}*	69.35×10^{6}	1.03
	1.22×10^{-2}*	$\mathbf{76.03 \times 10^{-2}}$	
		90.46×10^{-2}	
Japan	0.39×10^{-2}*	12.97×10^{6}	1.35
	0.59×10^{-2}*	17.23×10^{0}	
		$\mathbf{48.34 \times 10^{-1}}$	
Korea	0.50×10^{-3}*	29.02×10^{5}	1.13
	1.94×10^{-2}*	$\mathbf{71.87 \times 10^{-1}}$	
		21.61×10^{0}	

[Table 3 Continued...]

	λ^a_{min} r=0/r=1 r=1/r=2	$g_m(r_0)$ $r_0 = 0, 1, 2$	$H_0 : \beta' = (1, -1)^b$
Netherlands	0.02×10^{-3}*	23.37×10^{11}	1.13
	1.34×10^{-2}*	$\mathbf{18.74 \times 10^{-6c}}$	
		26.84×10^{-6}	
Norway	0.05×10^{-3}*	89.46×10^{6}	1.02
	0.81×10^{-2}*	13.42×10^{-1}	
		$\mathbf{70.12 \times 10^{-2}}$	
Spain	0.45×10^{-3}*	33.42×10^{4}	1.24
	9.83×10^{-2}	$\mathbf{24.48 \times 10^{-1}}$	
		18.77×10^{1}	
Sweden	0.03×10^{-3}*	41.45×10^{8}	1.07
	2.14×10^{-2}*	$\mathbf{41.48 \times 10^{-4}}$	
		15.13×10^{-3}	
Switzerland	0.39×10^{-2}*	12.16×10^{4}	1.11
	0.80×10^{-2}*	10.01×10^{2}	
		$\mathbf{51.57 \times 10^{1}}$	
UK	0.92×10^{-2}*	47.88×10^{4}	1.19
	0.94×10^{-2}*	18.38×10^{1}	
		$\mathbf{13.10 \times 10^{1}}$	

Notes:[a]Lambda-min statistic with critical values 0.017 and 0.05 for (q=2,r=0) and (q=2,r=1), respectively, at the five percent. See Bierens (1997a), Table 2.
[b]Value of the trace test for linear restrictions with critical value for m=2q equal to 4.70 when r=1 and q=2 at the 5 percent significance level. See Bierens (1997a), Table 3.
[c] The consistent estimate of the number of cointegrating vectors is given by
$\hat{r} = argmin_{r_0 \leq 2} \quad g_m(r_0)$, m=2. Values in bold denote the minimum value of the test statistic.
*Significant at the 5 percent significance level.