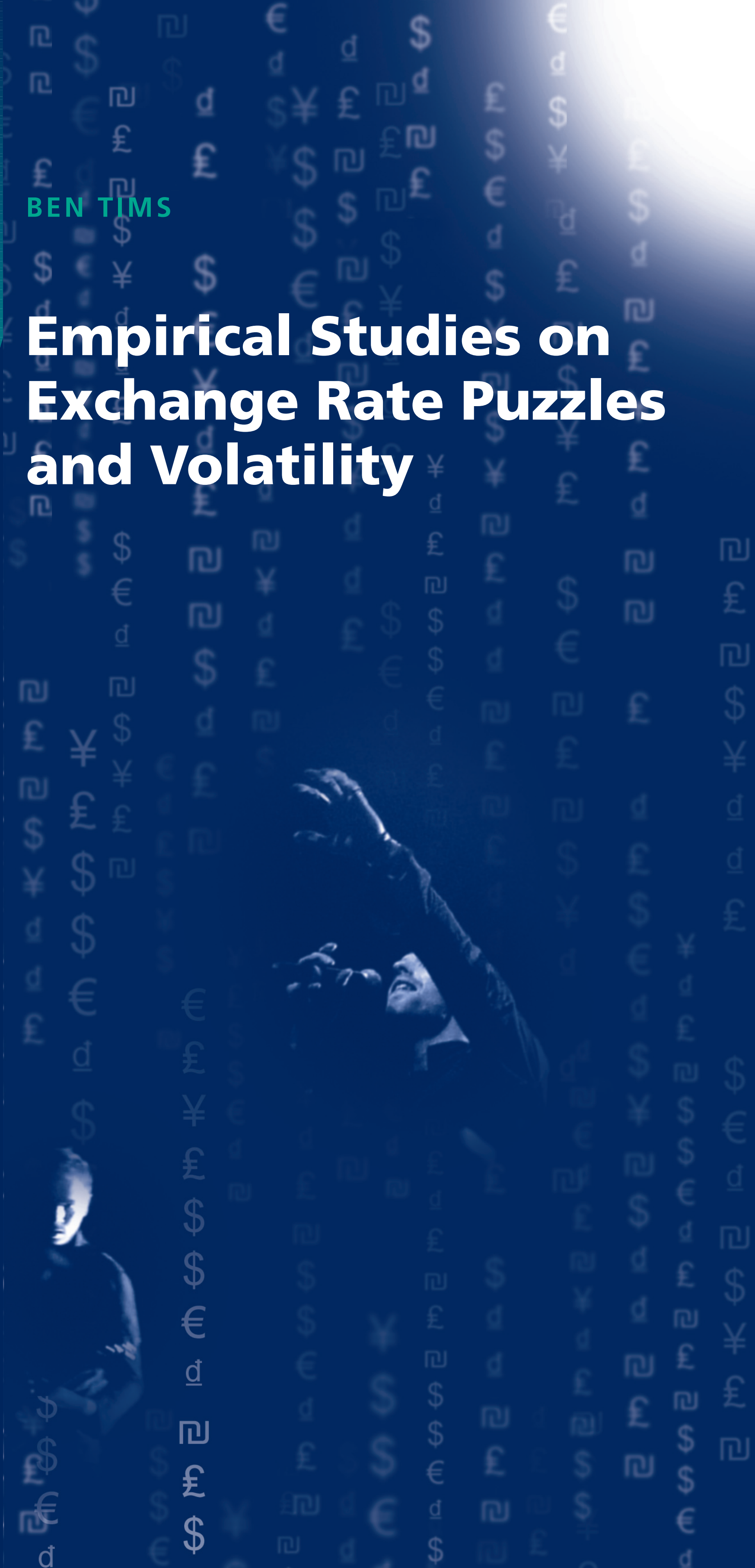
BEN TIMS

Empirical Studies on Exchange Rate Puzzles and Volatility



Empirical Studies on Exchange Rate Puzzles and Volatility

Empirical Studies on Exchange Rate Puzzles and Volatility

Empirische studies naar wisselkoerspuzzels en -volatiliteit

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"Nobody said it was easy"
The scientist, Coldplay

Voorwoord

Mijn proefschrift is eindelijk klaar en ik kan niet wachten totdat ik het eerste exemplaar ritueel kan gaan verbranden. Dat het zo'n tijd geduurd heeft is vooral aan mezelf te wijten en ben dankbaar dat "men" (in het bijzonder ERIM en de vakgroep) zo geduldig is geweest. Daarom wil ik alle familie, vrienden, collega's die direct danwel indirect betrokken zijn geweest bij dit proces bedanken voor de steun en het vertrouwen in de afgelopen jaren. Zonder jullie was het me zeker niet gelukt om dit proefschrift af te krijgen.

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Ben Tims
Rotterdam, september 2006

Contents

Chapter 1

Introduction

Money makes the world go 'round but to make money go 'round the world exchange rates are of paramount importance. It is hard to imagine a world without exchange rates and it will take a long time (if possible at all) before a global currency will arise that makes the current exchange rate system obsolete.

The theory and empiricism behind foreign exchange (FX) rates have proved to be very interesting and relevant in international financial economics. For example, the introduction of the euro in 1999 is a very recent example of this. This major event gives rise to important theoretical and practical issues (e.g. what is the effect of the introduction of the euro on the integration process in the euro area?). Another example is the current debate on the potential collapse of the dollar due to the reemergence of the twin deficits which also occurred during the middle of the eighties (see e.g. Obstfeld and Rogoff (2005)). Other issues like the determination and behavior of FX rates, market efficiency and currency crises are just a few topics of the vast exchange rate literature.

The focus in this thesis will be on the Purchasing Power Parity (PPP) puzzle, the Uncovered Interest rate Parity (UIP) puzzle, and on FX volatility. PPP and UIP are two of the most well-known exchange rate puzzles and serve as foundations for many economic models so it is important to investigate the validity of these parities. The volatility of exchange rates plays a crucial role in measuring risks in international trade and portfolio flows.

1.1 Purchasing Power Parity (PPP)

PPP states that (consumer and producer) goods should have the same price across countries when expressed in the same currency. Most studies investigate the validity of this parity for a basket of goods at an aggregate level using consumer and/or producer price indices instead of for individual goods (although there are exceptions of the latter such as the well-known Big Mac index published by the Economist). In real terms (as opposed to nominal terms) this means that the real exchange rate should equal unity at any given point in time. This version of PPP is called *absolute* PPP where the price ratio equals the exchange rate. *Relative* PPP states that the change in the exchange rate should equal the difference in inflation rates (change in prices). In this case the real exchange rate should be constant over time (not necessarily equal to one).

In theory, any deviation from PPP should result in an immediate change in prices and/or the exchange rate in such a way that the parity still holds. However, due to barriers like transaction costs, non-tradability etc., PPP does not hold perfectly in practice. The theoretical consensus is that these deviations are supposed to be temporary so that prices should converge to each other. This means that the real exchange rate should be mean-reverting/stationary and does not contain a unit root; after the real exchange rate is affected by an exogenous shock (whether caused by a change in prices or a change in the nominal exchange rate), the series should return to its mean level. The degree to which mean reversion takes place is often measured by the half-life which denotes the time needed for the series to return to a level that equals half the size of the original shock.

There is a vast history of PPP literature. A clear evolution in the methodology to assess the validity of this hypothesis can be discerned. First the parity was investigated in a univariate framework. However, due to the low power of the unit root test the null hypothesis of a unit root does not need to be rejected even if the real exchange rate is in fact stationary. A solution that has been offered is to increase the sample size (see e.g. Edison (1987) and Lothian and Taylor (1996)). According to Froot and Rogoff (1995) stationary (monthly) data covering a period of 72 years are needed to accept PPP with a half-life of three years. However, most exchange rate data covering such lengthy period are sampled under a fixed and floating exchange rate regime.

Under a fixed regime the exchange rate is fixed by (one of) the authorities of the underlying countries while under a floating regime the exchange rate is determined by supply and demand. The choice for a fixed or floating regime can be such that this regime is the best in keeping a stable macroeconomic performance as measured by for example price level, output, and consumption (see e.g. Caramazza and Aziz (1998)). Most of the developed countries follow a floating regime. An example of a fixed exchange rate regime was the Bretton Woods fixed exchange rate system which lasted from 1946 to 1973. A more recent example is the former regime in China that pegged the currency to the U.S. dollar until July 2005. Needless to say that the behavior of the exchange rate is different under a fixed or floating regime thereby contaminating inferences made on the PPP hypothesis. So increasing the time dimension in a way that different exchange rate regimes are included in the sample is not an option.

Another solution which has been offered to solve the lack of power of the unit root test in the univariate framework is to take a multivariate/panel approach by increasing the cross-sectional dimension instead of the time dimension. Subsequently the PPP hypothesis is tested by measuring the mean reversion of the panel as a whole. Using this multivariate panel model more support for stationary real exchange rates was found (see among others Abuaf and Jorion (1990)).

One of the most influential critiques on this approach was made by O'Connell (1998). He states that the neglect of the possible dependence between the real exchange rates in the panel can lead to spurious rejection of the unit root. Accounting for cross-sectional dependence O'Connell is unable to reject the unit root hypothesis. A second problem with this approach was that it does not allow for heterogeneous serial correlation of the real exchange rates in the panel. Papell and Theodoridis (2001) and Wu and Wu (2001) reckon with both the serial and cross-correlation properties and find evidence supporting the PPP hypothesis.

A third problem concerns the restriction of the same degree of mean reversion for all real FX rates in the panel. One would expect that the speed of mean reversion of the real exchange rate between two countries would depend on their relative proximity, their mutual trade regulations etc. such that the homogeneous confinement of the degree of mean reversion across real exchange rates is too restrictive. The effect this homogeneity has on inferences on the PPP hypothesis and on the power properties of the unit root test will be further discussed in this thesis.

1.2 Uncovered Interest rate Parity (UIP)

UIP states that the expected change in the spot exchange rate is equal to the forward premium (difference between the forward exchange rate and the spot exchange rate), see e.g. Hodrick (1987). In other words, the forward premium should be an unbiased predictor of the future exchange rate return. The forward premium is equal to the difference of the interest rates of the two countries underlying the exchange rate.

Therefore the UIP hypothesis can also be stated as the equivalence of the expected exchange rate return with the interest rate differential.

The common approach to assess the validity of the UIP hypothesis is to regress the realized exchange rate return on the forward premium or equivalently the interest differential and test whether the slope coefficient equals unity which is the case if UIP holds (see e.g. Fama (1984)). However, the literature finds consistently that the forward exchange rate is biased (see Hodrick (1987) and Engel (1996) for an overview): the slope is often significantly different from unity and often negative.

Several attempts are made to explain this UIP puzzle (or forward premium puzzle); see Sarno and Taylor (2003) and Sarno (2005) for a treatment of these offered solutions. For example, Bansal (1997) and Wu and Zhang (1996) investigate the asymmetric relationship between the future exchange rate return and the interest rate differential in such a way that positive and negative interest rate differentials are allowed to have a different impact on the exchange rate return. Bansal (1997) also investigates the relationship between the future exchange rate return and nonlinear transformations of the forward premium.

Other studies examine a nonlinear relationship between the exchange rate return and the forward premium to provide an explanation for the UIP hypothesis. Examples are Baillie and Kiliç (2006) and Sarno *et al.* (2005) who define a model where a distinction is made between different regimes where UIP does and does not hold.

This thesis will focus on explaining the future spot exchange rate return by (asymmetric and nonlinear versions of) the forward premium using linear and nonlinear models.

1.3 Exchange rate volatility

The third part of this thesis concerns the volatility of exchange rates which is an important topic in international economics. What model to use to measure volatility needs to be resolved.

Three of the most popular frameworks that deal with measuring volatility are the Realized Volatility (RV) models (see e.g. Andersen *et al.* (2001b) and Andersen *et al.* (2001a)), the Generalized AutoRegressive Conditional Heteroscedasticity (GARCH) models (see Engle (1982) and Bollerslev (1986)), and the Stochastic Volatility models (see Taylor (1986)). RV models are nonparametric models and deliver efficient volatility estimates making use of high-frequency (intradaily) data. The GARCH and SV models are both parametric models. The major distinction between these two models is that for the GARCH models expected volatility is a function of past, realized, observable information like returns while for the SV models expected volatility is a function of latent variables (see e.g. Andersen *et al.* (2006)). Estimation of the GARCH model is relatively easy while this is more cumbersome for the SV model.

For both GARCH and SV multivariate extensions exist. Not surprisingly, mainly due to the increase in parameters this hampers estimation even more. As a solution to this the use of another kind of data is pursued in this thesis. Instead of return data that is used to proxy for volatility, the maximum and minimum values of the exchange rate over a given period are utilized to construct another volatility approximation. The range is defined as the logarithm of the high price (maximum value for exchange rate) minus the logarithm of the low price. In other words, the range can be seen as the maximum "return" an investor could have made if he/she would have invested at the low price and divested at the high price.

The range is not only an accurate proxy for volatility (see Andersen and Bollerslev (1998) and Alizadeh *et al.* (2002)) but also has beneficial distributional properties. These properties will be exploited in this thesis in the estimation of a multivariate SV model to extract country-specific news information.

1.4 Outline

This thesis consists of three parts. Part I comprises of chapters 2 and 3 which present two studies on Purchasing Power Parity. Part II (chapters 4 and 5) consists of two studies on Uncovered Interest rate Parity while a new model for exchange rate volatility is presented in Chapter 6 in Part III. The summary and concluding remarks are given in the final chapter.

In Chapter 2 of this thesis the validity of the PPP hypothesis within the euro area and for the euro versus other major currencies is investigated. This is done by using a multivariate panel approach with and without the restriction of a homogeneous mean reversion across exchange rates in the panel. In the homogeneous setting the PPP hypothesis is investigated for the panel as a whole while in the heterogeneous case the validity of PPP is assessed for each individual currency pair in the panel.

Furthermore, the influence of the European economic integration process on the stationarity of the real exchange rates for the euro area countries is determined. In particular, the effect of the adoption of the Maastricht Treaty in 1992 and the introduction of the euro in 1999 on inferences for PPP are examined.

The multivariate panel model used in this chapter is based on the model of Flôres *et al.* (1995) that allows for heterogeneity of mean reversion across countries and for cross-sectional dependence. This model is extended to allow for serial correlation as well. Monte Carlo simulations are conducted to derive the critical values for the multivariate unit root tests. Simulations are also performed to analyze the power of the univariate unit root test versus the multivariate heterogeneous unit root test.

The results in this chapter give rise to a further analysis of the finite-sample properties of various multivariate unit root tests which are discussed in Chapter 3. A distinction is made between estimation and testing of the multivariate panel model homogeneously and heterogeneously which results in three different multivariate methodologies. For the first methodology estimation and testing are both done homogeneously. This is a common approach in the literature to evaluate the PPP hypothesis. The second methodology estimates the degree of mean reversion individually for each currency pair but the unit root test is still performed for the panel as a whole. Although heterogeneous mean reversion across real exchange rates is allowed in the second model, for both models it holds that no information is provided how many and which real exchange rates in the panel are stationary should the unit root hypothesis be rejected. Finally the multivariate model where both estimation and testing are done heterogeneously is considered. Again Monte Carlo simulations are performed to assess the (finite-sample) properties of these three multivariate methodologies.

Chapter 4 investigates the performance of linear models in describing and predicting exchange rate returns. Several insights of the current literature are combined into a single framework. The model of Clarida and Taylor (1997) forms the basis for the model in this chapter. Clarida and Taylor (1997) show that their model, which uses information from the interest rate term structure (spot and forward exchange rates with different maturities), superiorly outperforms the standard benchmark Random Walk model in predicting the exchange rate which is very remarkable. Their results contrast sharply with many other studies that are not able to beat the Random Walk model. In this thesis their results are verified.

Furthermore the multivariate model of Clarida and Taylor (1997) is extended to encompass the ideas of Bansal (1997) and Wu and Zhang (1996) who find an asymmetric relationship between the forward premium and the exchange rate return. More specifically, the sign and higher moments of the forward premium seem to be important in explaining the exchange rate. It will be investigated if these variables contain information that can be exploited in predicting the exchange rate.

In Chapter 5 the focus will be on the possible nonlinear relationship between the exchange rate return and the interest rate differential. Baillie and Kiliç (2006) employ a dynamic logistic smooth transition regression model to reveal nonlinearity. In this chapter this model is used as the reference model. UIP is assumed and deviations from this parity are measured by the nonlinear part of the model. The hypothesis that UIP holds is investigated by constructing bounds on the nonlinear part and test whether the deviations are significant from zero.

The model is extended to a multivariate model by collecting the UIP equations into a panel. In this way possible cross-sectional dependence between the exchange rate returns can be accounted for. Furthermore, the fact that the U.S. interest rate is a common factor across the equations can be exploited. The possible influence of these factors on the validity of the UIP hypothesis is examined.

In Chapter 6 the multivariate SV model to measure exchange rate volatility is introduced. According to Alizadeh *et al.* (2002), the range-based volatility measure, which is based on high and low prices, is an efficient estimate for volatility. Using the advantageous property that the log range is approximately normally distributed, the multivariate SV model as discussed in this chapter can be estimated efficiently. Furthermore, by decomposing the exchange rates into currency-specific news factors as is done by Mahieu and Schotman (1994) the exchange rate volatility consists of the sum of the two independent variances of the currencies underlying the exchange rate which leads to a parsimonious model. It is investigated whether the currency specific news factors are able to pick up country specific events.

The summary and concluding remarks are reported in Chapter 7.

Part I

Purchasing Power Parity

Chapter 2

Purchasing Power Parity and the Euro Area[1]

2.1 Introduction

Economic integration within Europe has progressed rapidly over the past decades. The introduction of the euro in January 1999 constituted the culmination of the monetary integration process that effectively started with the establishment of the European Monetary System in 1979. These developments may have important implications for the behavior of real exchange rates, not only within Europe, but also between the euro area and other countries.

Although the depreciation of the euro against the dollar in the period 1999-2000 and the subsequent appreciation has attracted a lot of attention, both in the popular financial press and in academic research (see Portes (2001) for extensive discussion), remarkably few empirical studies examine the behavior of real exchange rates for the euro area. In particular, only a very limited number of academic papers study the hypothesis of purchasing power parity (PPP) for the euro.

There are at least three reasons why research on PPP within the euro area is interesting and relevant. First, PPP is one of the central theoretical concepts in international economics. The transition of the euro area countries toward a single currency forms a unique opportunity to test the hypothesis of PPP. Second, the convergence of price levels (and thus the behavior of real exchange rates) within the European Monetary Union (EMU) is an important issue for public policy makers. This is highlighted by a recent speech by European Central Bank President Wim

[1] This chapter is based on Koedijk *et al.* (2004).

Duisenberg on concerns about divergent price developments among euro area countries.[2] Third, studying PPP in the euro area is also interesting from the perspective of asset pricing and portfolio management. While nominal exchange rate risk disappeared within the euro area in 1999, differences in inflation may entail nontrivial real exchange rate risk. The finding that real exchange rates within the euro area still exhibit considerable variation would have important implications for financial markets and asset managers.

In addition to the question whether PPP holds within the euro area, we are also interested in analyzing the behavior of the euro versus other major currencies, including the dollar, the pound, and the yen. As the economies in the euro area gradually converge, the euro area can increasingly be regarded as a single economic entity. Before the introduction of the euro, several studies (e.g. Pollard (1998) and Portes and Rey (1998)) suggested that the euro would rival the dollar as a major international currency. Recent evidence presented by the Bank for International Settlements (2003) suggests that the role of the euro in international financial markets is growing. This calls for an extensive investigation of PPP between the euro area as a separate economic entity versus other major industrialized economies.

We are aware of only two studies that directly examine PPP within the euro area or for the euro versus other currencies.[3] Lopez and Papell (2004) study the convergence to PPP in the euro area over the period 1973-2001. Their study involves quarterly data on exchange rates and CPI indices for the euro area and a number of related countries. Lopez and Papell find that the evidence in favor of PPP is considerable and clearly stronger within the euro area than between the euro area and other (European) countries. In particular, they present evidence that convergence to PPP was set in motion around the adoption of the Maastricht Treaty in 1992. Chinn (2002) uses data on the "synthetic" euro-dollar exchange rate for the period 1985-2001 to test the hypothesis of PPP. The results indicate that PPP is rejected when consumer price indices are used, although Chinn suggests that the euro-dollar rate exhibits more evidence of stationarity when producer price indices are employed. The paper documents a stable long-run relationship between the real euro-dollar rate, productivity differentials, and the real price of oil.

[2] See W.F. Duisenberg, "Are different price developments in the euro area a cause for concern?," speech delivered at the 2000 meetings of the Financial Services Industry Association in Dublin (available at `www.ecb.int`). We also refer to European Central Bank (1999).

[3] A number of related papers do not explicitly test for PPP. Rogers (2001) and Lutz (2002) investigate price level convergence within the euro area using European city price data and a data set consisting of a number of final goods prices, respectively. Clostermann and Schnatz (2000) analyze the determinants of the euro-dollar exchange rate without separately considering PPP.

This chapter examines PPP both within the euro area and between the euro area and a number of other major economies. Our contribution is threefold. First, we take into account that the extent to which PPP holds may exhibit important differences across countries. Most recent research on PPP imposes a common speed of mean reversion in unit root tests for a panel of real exchange rates. We strongly argue that PPP may well hold for some currency pairs and not for others. We employ a methodology that exploits the cross-sectional dependence between real exchange rates in order to enhance the power of the tests, but allows for different speeds of mean reversion for each individual currency in our sample. Second, our chapter analyzes PPP of the euro area as a distinct economic area versus a panel of other major economies. We study the behavior of the real exchange rate of the euro relative to the British pound, the Canadian dollar, the Danish krone, the Japanese yen, the Norwegian kroner, the Swedish krona, the Swiss franc, and the U.S. dollar over more than two decades. The use of the synthetic euro exchange rate before 1999 may shed valuable light on the validity of the PPP hypothesis for the euro versus other currencies in the future. Third, we use a very recent set of monthly data that enables us to assess the effect of the introduction of the euro in 1999.

Our results show that the unit root hypothesis can be rejected for the euro area over the period 1973-2003 when the speed of mean reversion is considered to be the same for all currencies. Relaxing this restriction, however, reveals that PPP is a reasonable hypothesis for some currency pairs, but not for others. This suggests that accounting for inter-country differences is of great importance in empirical studies of PPP. The assumption that a panel of real exchange rates exhibit a common speed of mean reversion is generally too restrictive. Our analysis of PPP between the euro area and other major economies reveals that the unit root hypothesis for the panel of real exchange rates against the euro can be rejected. However, with heterogeneous mean reversion we present evidence in favor of PPP between the euro area and Switzerland only over the period 1979-2003.

We investigate the influence of the European economic integration processes on the stationarity of real exchange rates. In particular, we examine whether the adoption of the Maastricht Treaty in 1992 and the introduction of the euro in 1999 have fueled a convergence toward PPP. We confirm the finding of Lopez and Papell (2004) that especially the former event had an important impact on the stationarity of real exchange rates in the euro area. Strong evidence in favor of PPP for the full panel of euro area currencies is detected after 1992. The convergence process toward PPP is rather diverse for individual currency pairs, however.

The remainder of the chapter is structured as follows. In Section 2.2 we describe the methodology. Section 2.3 provides the data description. We examine the behavior of real exchange rates within the euro area in Section 2.4, while Section 2.5 discusses our findings on PPP of the euro versus other major currencies. An assessment of the power of the univariate and the multivariate unit root tests is provided in Section 2.6. Section 2.7 concludes.

2.2 Methodology

For each country (currency) i ($i = 1, ..., N$) we define the log real exchange rate at time t ($t = 1, ..., T$) as follows:

$$R_{i,t} = e_{i,t} - e_{0,t} + p_{0,t} - p_{i,t} \tag{2.1}$$

where $R_{i,t}$ is the logarithm of the real exchange rate, $e_{i,t}$ is the logarithm of the nominal exchange rate expressed in units of currency i per dollar, $e_{0,t}$ is the logarithm of the nominal exchange rate expressed in units of the numeraire currency 0 per dollar, $p_{0,t}$ is the logarithm of the consumer price index of the country used as numeraire, and $p_{i,t}$ is the logarithm of the consumer price index in country i.

If PPP holds perfectly, the real exchange rate is constant. In practice, testing for PPP boils down to investigating whether the log real exchange rate shows mean-reverting behavior. This is usually done by means of a unit root test. If the null-hypothesis of a unit root is rejected, the real exchange rate is mean-reverting and therefore real exchange rates tend to revert to their PPP level in the long run. If the series contain a unit root, however, there is no mean-reversion and PPP does not hold.

There has been a clear evolution in the methodologies employed in PPP studies. Early papers predominantly use univariate unit root tests. However, the lack of power of the Dickey-Fuller unit root test can deter rejection of the unit root in favor of PPP even though the log real exchange rate under consideration is, in fact, stationary. Increasing the length of the sample has been offered as a solution (see e.g. Edison (1987) and Lothian and Taylor (1996). Froot and Rogoff (1995) show that a very long time series is needed to overcome the power problem.[4] This implies that data from both fixed and floating rate periods have to be used, which blurs the interpretation of the results.

As an alternative way of increasing the power of the unit root tests, many studies turn to panel data models, see e.g. Abuaf and Jorion (1990), Jorion and Sweeney (1996), and Frankel and Rose (1996). Imposing a common mean-reversion coefficient

[4] They show that 72 years of stationary (monthly) data is needed to accept PPP with a mean reversion coefficient of 0.981 (equivalent to a half-life of three years).

for all real exchange rates results in relatively strong evidence in favor of PPP. A principal problem with the panel approach is formulated by O'Connell (1998), who demonstrates that spurious rejections of a unit root can occur when cross-sectional dependence is unaccounted for. Imposing severe restrictions on the variance-covariance matrix in a panel of real exchange rates leads to serious biases in the size and the power of the test. Accounting for cross-sectional dependence, O'Connell is unable to reject the unit root hypothesis in a panel of 64 countries over the period 1973-1995. Abandoning the restrictions on the variance-covariance matrix in panel studies has the implication that the results become invariant to the choice of the numeraire currency.

Papell and Theodoridis (2001) and Wu and Wu (2001) point out that the results of O'Connell are only valid when the serial correlation properties of all real exchange rates are the same. They essentially propose a multivariate version of the Augmented Dickey-Fuller (ADF) test. As both lag length and the serial correlation coefficients are heterogeneous across real exchange rates, the choice of the numeraire currency can make a difference. Both papers present evidence against the unit root null in a panel of currencies over the recent float.

In the present chapter, we contend that another culprit in recent panel studies of PPP is generally overlooked. The usual assumption of a common mean reversion coefficient across all real exchange rates is also excessively restrictive. Intuitively, we would expect that the speed of mean reversion depends on, for example, the relative proximity of countries, their mutual trade regulations, and the openness of their economies. Hence, even within an economically integrated region such as the euro area, the extent to which violations of PPP occur is likely to be dependent on the countries examined. Econometrically, Flôres *et al.* (1995) contend that unit root tests are better behaved when different speeds of mean reversion are allowed.

A number of previous studies incorporate heterogeneous mean reversion in the panel methodology. Koedijk *et al.* (1998) focus on the symmetry and proportionality conditions in the PPP relation, however, and do not perform a unit root test. Im *et al.* (2003) and Wu and Wu (2001) allow the slope coefficients in the panel unit root tests to differ across exchange rates, but they propose a test statistic for the validity of PPP for the *full* panel of currencies. This complicates the interpretation of rejecting the null hypothesis, as the null hypothesis will be violated if one or more of the real exchange rates is stationary. The tests do not provide any guidance as to which particular real exchange rates are stationary. Taylor and Sarno (1998) suggest a test statistic that only rejects the null if all real exchange rates are stationary, but this test does not facilitate the evaluation of PPP on an individual country basis. We reckon that while multivariate tests of PPP may be necessary for power considerations, the issue whether individual real exchange rates are stationary is still interesting and germane.

In this chapter we extend the heterogeneous Seemingly Unrelated Regression (SUR) methodology employed by Flôres *et al.* (1999) to test the PPP hypothesis. This model is not only able to cope with the cross-sectional dependence, but also with the different speeds of mean reversion across real exchange rates. Flôres et al. develop unit root tests that can be applied to each individual currency in the panel. The model can be expressed as follows:

$$R_{i,t} = \alpha_i + \beta_i R_{i,t-1} + u_{i,t}, \tag{2.2}$$

where $u_{i,t}$ is a stationary error term, and α_i and β_i are the intercept and the mean reversion parameters, respectively. Note that in this model contemporaneous correlations between the error terms $u_{i,t}$ are allowed. However, Flôres et al. assume that the serial correlation properties of each real exchange rate are the same. As Papell and Theodoridis (2001) and Wu and Wu (2001) show, however, allowing for heterogeneous serial correlation is important. Therefore, we extend the model of Flôres et al. as follows

$$R_{i,t} = \alpha_i + \beta_i R_{i,t-1} + \sum_{k=1}^{l_i} \gamma_{i,k} \Delta R_{i,t-k} + u_{i,t}, \tag{2.3}$$

where l_i denotes the number of lags needed for currency i.[5,6]

The SUR model is estimated in the following way. First, for each currency i, we apply OLS to Equation (2.3). The covariance matrix of the error terms is used as the weighting matrix in a Feasible Generalized Least Squares (FGLS) procedure.[7] Second, the estimated parameters are utilized to construct new residuals, which in turn result in a new estimate for the covariance matrix and so on. This process is repeated until convergence takes place.

[5]The value of l_i is determined by the recursive t-statistic procedure of Campbell and Perron (1991) applied to each individual log real exchange rate. This means that for currency i we choose the value of l_i by first setting l_i to some maximal value, l_{max}, then estimate Equation (2.3) by OLS and subsequently test whether the last included lag is statistically significant. If so, then l_i is set to this value, else the model is estimated by setting l_i to l_{max} - 1. The procedure is repeated until a significant value of l_i is found. When no lag is significant then l_i is set to 0. Following Wu and Wu (2001), we set l_{max} to 24 and use a 10% significance level.

[6]A similar methodology is developed by Breuer *et al.* (2002). However, our approach provides for a more balanced view of the unit root hypothesis, because we assess various alternative formulations of the null-hypothesis in line with Flôres et al. (1999). Moreover, the empirical analysis of Breuer et al. is restricted to quarterly data up to 1998.

[7]Note that in a standard SUR model the degrees of freedom needed for the calculation of the covariance matrix of the error terms equals $T-2$. We have to make a correction for the inclusion of lagged changes in the real exchange rate in the model. To that effect, we set the degrees of freedom to $T-2-\mathit{entier}[(li+lj+1)/2]$, where $\mathit{entier}[x]$ rounds x down to the nearest integer value.

We use Monte Carlo simulations to derive critical values for the test statistic τ for the multivariate tests, analogous to Dickey and Fuller (1979). For the heterogeneous model, we follow Flôres et al. (1999) and distinguish between three different null-hypotheses for each individual currency. Under H_0^1, $\alpha_i = 0$ and $\beta_i = 1$ for all currencies i. Under H_0^2, we compute the critical values for each currency i by setting $\alpha_i = 0$ and $\beta_i = 1$, and $\beta_j = \hat{\beta}_j$, $j \neq i$. The third null-hypothesis, H_0^3, is a more conservative approach than H_0^2 which involves a two-step procedure. In the first step, we make use of the critical values for H_0^1 and define I_1 as the set of currencies for which the latter null-hypothesis is rejected. In the second step we compute critical values for $H_0^3 : \alpha_i = 0$ for all currencies i, $\beta_j = \hat{\beta}_j$, $j \in I_1$, and $\beta_j = 1$, $j \notin I_1$.

The Monte Carlo simulations for the computation of the critical values for each individual currency i involve five steps. First, given the estimation of the γ_k parameters of the model, we compute the residuals under the null-hypothesis and compute the covariance matrix. Second, we generate N error terms $u_{i,t}$ (T times) from a multivariate normal distribution with mean zero and the covariance matrix (note that this matrix accounts for cross-dependence across the exchange rates). Third, given the estimated parameters γ_k, we compute simulated exchange rate series using Equation (2.3) with the simulated error terms $u_{i,t}$ from step 2, and the three null-hypotheses H_0^1, H_0^2, and H_0^3. Fourth, we estimate the parameters in the SUR regression (2.3) with the simulated exchange rate series and compute the test statistic $\tau = (\beta - 1)/s(\beta)$. Finally, we replicate the first four steps 1000 times and derive critical values for the test statistic from its sample distribution. Alternatively, empirical p-values can be calculated as the fraction of times the observed test statistic using the actual empirical data series is exceeded in the replications.[8] For the model with homogeneous mean reversion parameters, we only simulate critical values for H_0^1.

2.3 Data

The empirical analysis presented in this chapter consists of two parts. First, we study PPP within the euro area. For this purpose we collect a dataset of consumer price index (CPI) and nominal exchange rate data for Austria, Belgium, Finland, France, Germany, Greece, Italy, the Netherlands, Portugal, and Spain for the period 1973:02-2003:03.[9] CPI data and period-ending exchange rates against the U.S. dollar are

[8] Note that for H_0^2 N simulations need to be performed to obtain the p-values for each currency.

[9] The first 25 observations are used to compute the lagged exchange rate changes needed for the ADF tests.

obtained from International Financial Statistics.[10] In the second part of the chapter we study the real exchange rate behavior of the euro versus the British pound, the Canadian dollar, the Danish krone, the Japanese yen, the Norwegian kroner, the Swedish krona, the Swiss franc, and the U.S. dollar. Nominal exchange rates against the dollar and CPI data for Canada, Denmark, Japan, Norway, Sweden, Switzerland, the United Kingdom, and the United States for the period 1978:12-2003:03 are taken from IFS. Because the euro/dollar rate is only available from January 1999, we use the "synthetic" euro from the ECB.[11] In order to construct the CPI data for the euro area we use the geometric weighted average method as described on page 11 in Maeso-Fernandez *et al.* (2001). Ireland is discarded from the analysis because the CPI is only available as of 1997. Luxembourg is excluded because of its currency union with Belgium.

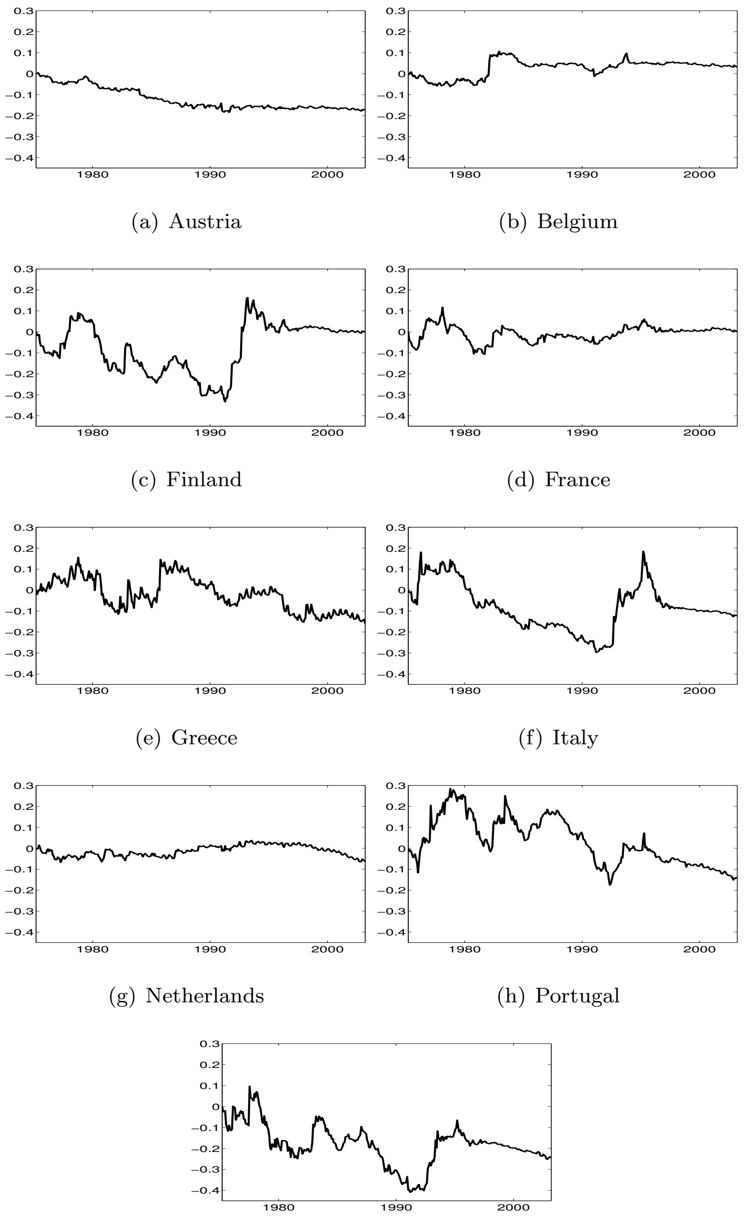
2.4 PPP within the euro area

This section discusses our empirical analysis of PPP within the euro area. Figure 2.1 depicts the (log) real exchange rates against the DMark over the period March 1975-March 2003.[12] Two striking observations can be made from inspecting the graphs. First, there are large differences in the time-series behavior of the real exchange rates against the DMark. The graphs for e.g. Greece and Portugal appear inconsistent with short-run mean-reverting behavior, as the real exchange rates for these countries cross their mean values relatively infrequently. On the other hand, the graph for France is suggestive of a stationary real exchange rate. Second, the degree to which real exchange rates fluctuate is substantially smaller in the past decade than before. Especially since the introduction of the euro in 1999 real exchange rates have been relatively stable. Notably for Belgium, Finland, France, Italy, and Spain, the graphs suggest that real exchange rates against the DMark have exhibited strong mean-reverting behavior in the most recent years.

[10] The CPI data for the Netherlands seems to contain an error. In January 1981, there is an unexplainable upward shift in the index from 70.4 to 87.1, which is reversed in February 1987 by a drop from 102.2 to 83.4. We have tried to correct for this data error by subtracting the average shift (amounting to 17.75) from all the data points between these dates. Our results do not materially depend on the correction method.

[11] This synthetic euro series is the "ECB reference exchange rate, US Dollar/Euro, 2:15 pm (C.E.T.), against ECU up to December 1998," which is available at the website of the European Central Bank (`www.ecb.int`).

[12] Note that, without loss of generality, we can normalize the first observation for each real exchange rate to be equal to zero, because we use price index data and not actual prices.



(a) Austria

(b) Belgium

(c) Finland

(d) France

(e) Greece

(f) Italy

(g) Netherlands

(h) Portugal

(i) Spain

Figure 2.1: Euro Area Real Exchange Rates Against the DMark
This figure presents the (log) real exchange rates of Austria, Belgium, Finland, France, Greece, Italy, the Netherlands, Portugal, and Spain versus Germany for the sample period 1975:03-2003:03.

As we study the convergence to PPP within the euro area in relation to the ongoing integration process, it is of interest to examine how exchange rates were actually set in the EMU. This necessitates a brief discussion of the history of European monetary integration. After the collapse of the Bretton Woods system, the Smithsonian Agreement of December 1971 provided for an expansion of the band within which exchange rates were allowed to move from 1 percent to 2.25 percent. Members of the European Economic Community (EEC), however, decided on a narrower band of 1.125 percent of their currencies. This regime was referred to as the "snake in the tunnel," as the European currencies moved closely together within the wider band allowed for other currencies. The snake was considered unsuccessful in limiting exchange rate fluctuations and several countries were forced to leave the system. In March 1979, the snake arrangement was replaced by the Exchange Rate Mechanism (ERM), which was part of the broader European Monetary System (EMS) designed to establish a "zone of monetary stability" in Europe. Within the ERM, each currency was kept within a band of ± 2.25 percent around central parity. The arrangement was represented by the parity grid, a system of par values among ERM currencies called the "ERM central rates." After the crisis in 1992, in which the Italian lira and the British pound were forced to leave the ERM, the monetary authorities adopted wider bands of ± 15 percent around central parity. In the mid-1990s, Austria, Finland, and Greece joined the ERM, while Italy rejoined in 1996. In the years before the adoption of the euro, the ERM currencies moved very close to their central rates. On January 1, 1999, the irrevocable euro conversion rates of the 11 EMU member states currencies were set on the basis of the bilateral ERM central rates.[13] Greece joined the EMU in January 2001.

The determination of the ERM central rates is a subject that has received little attention. The bilateral ERM rates were not set on the basis of a thorough analysis of economic fundamentals and equilibrium exchange rates, but were rather based on the exchange rates in the snake in the tunnel arrangement (and can thus be traced back to the Bretton Woods system). Without reliable data on absolute price levels in Euro area countries, it is very hard to ascertain which currencies were overvalued and which were undervalued in the ERM. Hence, establishing whether individual currencies were subsumed in the euro below, above or at the (long-run) PPP level is hardly feasible. Although the ERM was a managed exchange rate system, it is unlikely that the exchange rates could deviate substantially from what the market considers to be the fundamental rate. Indeed, the ERM was characterized by frequent and substantial realignments, to a large extent due to market pressure. This suggests

[13] Detailed information can be found in the "Joint Communiqué on the Determination of the Irrevocable Conversion Rates for the Euro," issued by the European Union on May 2, 1998.

that large discrepancies from fundamental market values among euro area currencies were not to be expected at the introduction of the euro. On the other hand, the vast PPP literature shows that even completely flexible exchange rates do not always tend to trade at their PPP level. Moreover, the EMU central banks committed themselves to ensuring that the closing rates on December 31, 1998 would be equal to the central rates, so minor discrepancies are not unlikely. Assessing whether the price levels of individual countries could be expected to adjust in the first years after the establishment of the EMU in order to reestablish absolute PPP is an extremely intricate issue.

Table 2.1: Correlations of Euro Area Real Exchange Rates.
This table shows correlations of the first differences in the real exchange rates $R_{i,t} - R_{i,t-1}$ of nine euro area countries versus the Deutsche mark over the period 1973:02-2003:03.

Country	Aus	Bel	Fin	Fra	Gre	Ita	Net	Por	Spa
Austria	1.000	0.341	0.205	0.241	-0.012	0.202	0.245	0.141	0.259
Belgium		1.000	0.194	0.365	0.111	0.219	0.429	0.164	0.272
Finland			1.000	0.314	0.260	0.507	0.059	0.315	0.424
France				1.000	0.246	0.418	0.289	0.374	0.344
Greece					1.000	0.282	0.183	0.178	0.299
Italy						1.000	0.161	0.304	0.480
Netherlands							1.000	0.155	0.227
Portugal								1.000	0.389
Spain									1.000

Table 2.1 displays pair wise correlations between changes in the real exchange rates for the sample period February 1973-March 2003 of nine countries in the euro area. Virtually all correlations are positive and substantial, suggesting that accounting for cross-sectional dependence is vital in an analysis of PPP for the euro area. Table 2.2 presents the results of the univariate ADF unit root test for all real exchange rates in the sample. Columns present the estimated parameters, their standard errors, the half-life of PPP deviations expressed in months, the maximum number of lags for each exchange rate, and the test statistic τ. The last two rows depict the 10% and 5% critical values for the test statistic, obtained from Dickey and Fuller (1979). Table 2.2 shows that the unit root can be rejected for four countries: Belgium, Fin-

land, France, and Spain. For all other real exchange rates against the DMark there is no evidence against the unit root null. As argued in earlier studies, it may well be the case that the failure to reject the unit root is due to the lack of power in univariate tests.

The results of the SUR tests with homogeneous mean reversion depicted in Table 2.3 show that multivariate analysis has important advantages. When a common mean reversion coefficient is assumed, the unit root hypothesis can be rejected at the 5% level for our panel of real exchange rates in the euro area. Critical values are obtained using Monte Carlo simulations as discussed in Section 2.2. The half-life of PPP deviations is estimated to be approximately three years, which is consistent with previous research. Note that the critique of O'Connell does not apply for this model, as we do not impose any restrictions on the covariance matrix. The results in Table 2.3 support the conclusion of Lopez and Papell (2004) that there is evidence in favor of PPP for the euro area over a recent period.

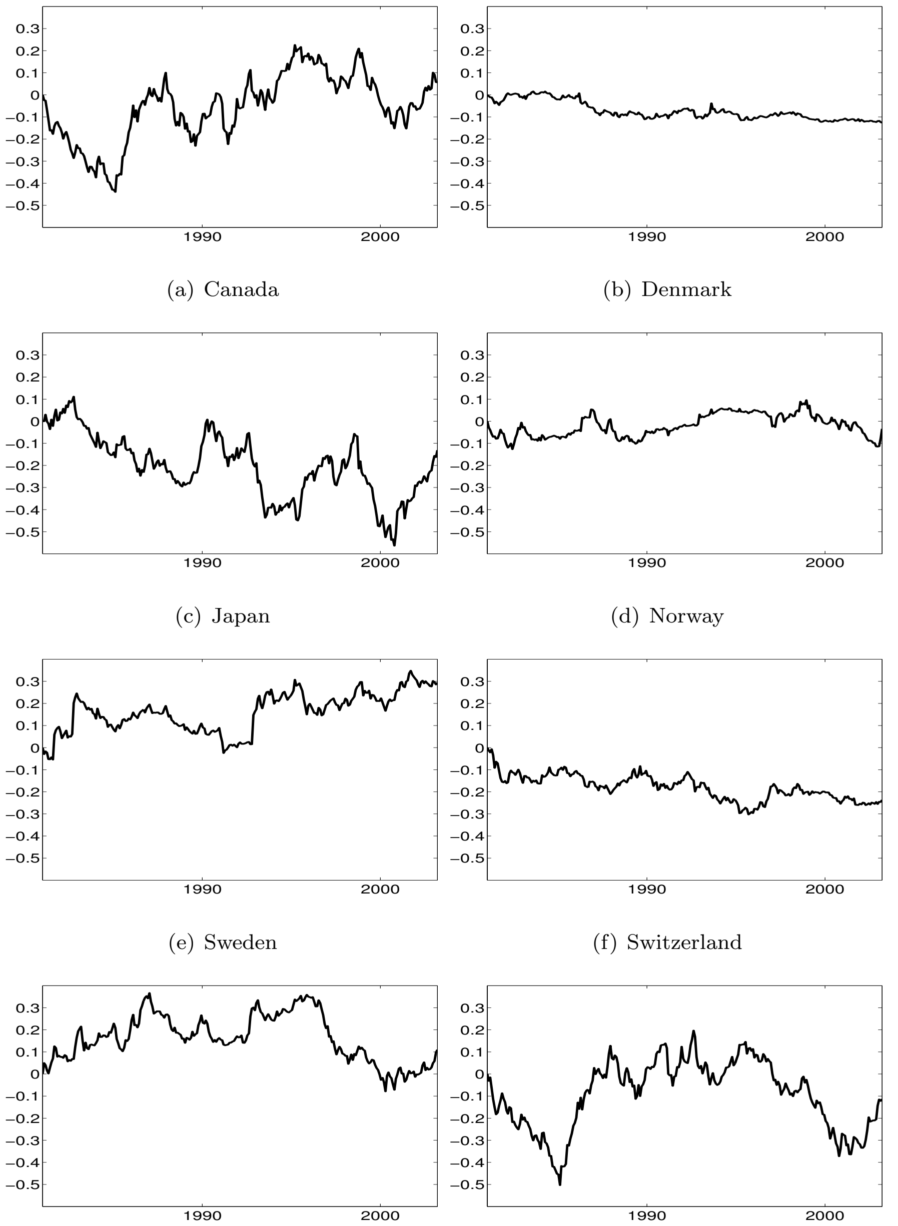
From an economic viewpoint, however, the restriction that the mean reversion parameter should be the same across countries seems unjustifiably restrictive. The univariate tests indicate that half-lifes substantially differ across real exchange rates. We are interested in the PPP hypothesis for individual currency pairs. Therefore, we present the results of a SUR model with heterogeneous mean reversion coefficients in the top panel of Table 2.4. The final column of the table depicts the test statistic τ for distinct unit root tests for each of the countries in the sample. Simulated 10% critical values under the null-hypothesis that all real exchange rates have a unit root (hypothesis H_0^1) are depicted in brackets. The results demonstrate that accounting for differences across countries is important. There is substantial variation in the estimates of β across countries. Half-lifes vary from less than a year for France to more than five years for Austria, with an extreme estimate of over a hundred years for Portugal. The hypothesis of a unit root can be rejected for some real exchange rates, but not for others. There is evidence in favor of PPP for Finland, France, and Spain (all at the 5% level). The bottom panel of Table 2.4 reflects empirical p-values for the three different null-hypotheses H_0^1, H_0^2, and H_0^3. Simulations of the p-values under H_0^2 assume stationarity of all other series, while H_0^3 assumes stationarity for some of the series. For H_0^3 and H_0^2, the unit root can also be rejected for Italy at the 10% level.

We are interested in the issue whether the continuing economic integration in the euro area has set off a convergence process toward PPP. We investigate PPP for a number of subperiods. As Monte Carlo simulations for heterogeneous SUR models with 10 currencies are very intensive in terms of computing time, we restrict ourselves to the period before the adoption of the Maastricht Treaty in 1992, the period before the introduction of the euro in 1999, and the full sample period. Table 2.5 shows the p-values under the null-hypothesis H_0^1 for three different periods: 1975:03-1991:12, 1975:03-1998:12, and 1975:03-2003:03. The result for the homogeneous SUR model confirm Lopez and Papell's (2003) inference that the Maastricht Treaty triggered convergence toward PPP. There is very little evidence in favor of PPP before 1992, while p-values indicate strong rejection of the unit root when the post-Maastricht Treaty period is included. The p-values for the heterogeneous model provide a more balanced view on the convergence process toward PPP. Simulated p-values clearly decrease over time for six of the nine currencies versus the DMark. The effect is most notable for Belgium, Italy, and Spain. However, for the guilder-mark rate there does not seem to be any convergence, while p-values substantially increase for the Greek drachma and the Portuguese escudo. This suggests that the case for convergence is not as clear-cut as previous studies imply. For the Netherlands, this effects seems to be due to the fact that Dutch inflation has been persistently higher than German inflation since 1996, probably related to the much lower level of unemployment. Greek and Portuguese inflation have also been considerably higher than in Germany for an extended period. A tentative explanation is the "inflation catch-up" phenomenon. This entails a temporarily higher inflation in low-price countries due to the convergence of price levels across Europe. The post-euro sample period is too short to resolve the issue whether these developments are transitory.

The results in Table 2.3 indicate that PPP seems to hold within the euro area when a common mean reversion coefficient is assumed. Shocks that cause a divergence from PPP are generally halved within three years. However, the evidence in Table 2.4 shows that it is vital to take cross-country differences into account. Half-lifes of PPP deviations turn out to exhibit considerable differences across real exchange rates. With different mean reversion coefficients, we find evidence in favor of PPP for several countries, but for other countries the unit root null cannot be rejected.

2.5 PPP in a panel of the euro and other major currencies

This section applies the approach employed in Section 2.4 to the euro area as a separate economic entity versus other major economies. We examine PPP for the real exchange rates of the British pound, Canadian dollar, the Danish krone, the Japanese yen, the Norwegian kroner, the Swiss franc, the Swedish krona, and the U.S. dollar with the euro as the numeraire. Figure 2.2 shows that real exchange rates against the (synthetic) euro exhibit substantial variation over time, both before and after



(a) Canada (b) Denmark

(c) Japan (d) Norway

(e) Sweden (f) Switzerland

(g) U.K. (h) U.S.

Figure 2.2: Real Exchange Rates of Several Major Currencies Against the Euro. This figure presents the (log) real exchange rates of Canada, Denmark, Japan, Norway, Sweden, Switzerland, the U.K., and the U.S. versus the euro area for the sample period 1978:12-2003:03.

the introduction of the euro. There are also notable differences between the shapes of the different graphs. For example, the graph for Switzerland seems indicative of a stationary rate, while the graphs for Japan and the U.S. seems to suggest a unit root. In Table 2.6 the pair wise correlations between changes in the real exchange rates are displayed. Correlations are generally positive and too high to neglect, especially the correlation between the rates of the Canadian dollar and the U.S. dollar against the euro. Therefore, accounting for cross-sectional dependence is imperative when investigating PPP for this panel of countries. Tables 2.7 to 2.9 present the results of respectively the univariate ADF tests, the SUR model with homogeneous mean reversion, and the SUR model with heterogeneous mean reversion.

The results of the univariate analysis in Table 2.7 indicate that only for the U.S. the unit root is rejected at the 10% level. The parameter estimates for the SUR model with homogeneous mean reversion presented in Table 2.8 imply that there is evidence in favor of PPP for the full panel of exchange rates at the 10% level. Relaxing the restriction that PPP holds equally well for each currency in the sample produces mean reversion coefficients that differ significantly across rates (see Table 2.9). The half-life of PPP deviations is approximately one year for the euro-Norwegian kroner and the euro-Swiss franc series and almost four years for the euro-Danish krone rate. Evidence of PPP is only detected between the euro area and Switzerland. In particular, there seems to be no evidence for PPP between the euro area and the U.S. The bottom panel of Table 2.9 shows that the outcomes remain basically unaffected under different null-hypotheses.

Progressing European economic integration does not only affect the behavior of real exchange rates within the euro area, but is also likely to have an impact on PPP in a panel of the euro and other major currencies. We again analyze the influence of the adoption of the Maastricht Treaty in 1992 and the introduction of the euro in 1999 on the evidence for PPP. Table 2.10 shows the development of the p-values under H_0^1 with the euro as numeraire. For the full panel of currencies, the unit root is rejected irrespective of the time period. For the individual currency pairs, no clear pattern arises. Simulated p-values do not tend to decline over time, not even for the non-EMU European countries in the sample. Neither the Maastricht Treaty nor the introduction of the euro has a reliable effect on the test results.

Taken together, the evidence for PPP in a panel of the euro and other major currencies is ambiguous. We report evidence in favor of PPP for the full panel, but the unit root hypothesis cannot be rejected for the individual real exchange rates in the sample, with the exception of the euro-Swiss franc rate. Again, this suggests that the assumption of a common mean reversion coefficient for all real exchange rates in the sample is excessively restrictive. The conclusion that PPP holds for the panel of currencies, while in actual fact all but one of the real exchange rates contain a unit root, is indefinite.

2.6 The power of multivariate versus univariate tests

The methodology of multivariate unit root tests was introduced into the PPP literature primarily because of power considerations. Univariate unit root tests are known for being relatively inept in distinguishing between the unit root null and stationary alternatives. A number of studies have shown that the statistical power of panel methodologies that impose homogeneous mean reversion across real exchange rates is much higher. This chapter contends that allowing for heterogeneous mean reversion is important from an economic perspective. A germane issue is whether alleviating the restriction of a common mean reversion influences the power of the multivariate test. This section present the results of an analysis of the power of the univariate ADF test versus the power of the heterogeneous SUR ADF test.

The power functions of the univariate and multivariate tests are computed by Monte Carlo simulations. For the univariate case we first estimate Equation (2.3) separately for all real exchange rates. Second, we adjust the residuals of the estimation in such a way that they reflect various alternative hypotheses. This adjustment is done in a similar fashion in the calculation of the critical values of the ADF test. We follow Taylor and Sarno (1998) and employ the following values of β_i: 0.990, 0.975, 0.950, 0.925, and 0.900. These rates of mean reversion correspond to half-lifes of PPP deviations of, respectively, 69, 27, 14, 9, and 7 months. Third, we perform 1000 replications by generating error terms from a normal distribution with mean 0 and the variance of the adjusted residuals, which are used to construct simulated real exchange rates. Finally, we derive the power from the fraction of times the unit root null is rejected in favor of the stationary alternative using the critical values derived under H_0^1. In order to obtain the power function of the multivariate test, we simulate real exchange rate series from a multivariate normal distribution with mean zero and the covariance matrix of the residuals, which are adjusted for the alternative hypotheses after the multivariate estimation of Equation (2.3).[14] The power function

[14] For computing time considerations, we have decided not to iterate the estimation procedure until convergence takes place but take a single step approach to calculate the power of the models. Unreported results show that this generally has no influence on the parameter estimates, standard errors, and test statistics of the OLS ADF and SUR ADF models. The results are available from the authors on request.

can be constructed by evaluating the simulated cumulative distribution function of the heterogeneous SUR ADF test statistics at the critical value obtained under H_0^1.[15]

As we are also interested in the influence of the length of the data set on the power of alternative models the Monte Carlo analysis is repeated for different values of T. The number of observations in the simulations is set to the original number of monthly time-series observations in our sample (which is 337 for the countries within the euro area and 267 for the euro and other major currencies), 400 and 500. We use a significance level of 10%. As presenting and comparing the individual power functions for all currencies in the sample is cumbersome, we aggregate the results across all individual currencies and present average power functions.[16]

Figure 2.3 exhibits the average power functions of the univariate and the heterogeneous SUR model for different values of T for the countries within the euro area with the DMark as numeraire. We observe that the power of both the univariate and the multivariate test is relatively low. The fact that the tests are able to distinguish between the null hypothesis that $\beta_i = 1$ and the alternative hypothesis that $\beta_i = 0.990$ in only 20 percent of the cases does not strike us as an insurmountable problem. If the half-life of deviations from PPP is almost 7 years, few economists would regard the series as perfectly stationary. However, even when the half-life is less than 1.5 years, the tests often fail to reject the unit root. For example, for the original number of observations ($T = 337$) there is a probability of at most 70% that the unit root hypothesis is rejected while the true β_i equals 0.950, which corresponds with a half-life of only 14 months. Even for $\beta_i = 0.900$ (7 months half-life) there is still a probability of more than 10% that the stationarity of the real exchange rate is not detected by the statistical test. The power increase for a higher number of observations is remarkably limited. Even when the data set would span a period of more than 40 years, there is at most a 70% probability that the unit root is rejected for a real exchange rate with a half-life of 27 months.

The most remarkable conclusion from Figure 2.3, however, is that the average power of the univariate unit root tests is at least as large as the average power of the heterogeneous SUR ADF test. Evidently, with the DMark as the numeraire currency, using multivariate tests does not lead to a higher power when compared to univariate tests. A plausible explanation is that estimating a heterogeneous SUR model involves a very large number of parameters. In a model with 10 real exchange rates, a maximum of 279 parameters ($9 \times 10/2$=45 elements of the covariance matrix,

[15]Note that for the univariate model the power functions of the real exchange rates in the sample are constructed separately. For the heterogeneous SUR model, the power of the unit root test for currency i depends on currency j. We set $\beta_j = 1$ for $j \neq i$.

[16]Power functions for individual real exchange rates are available from the authors.

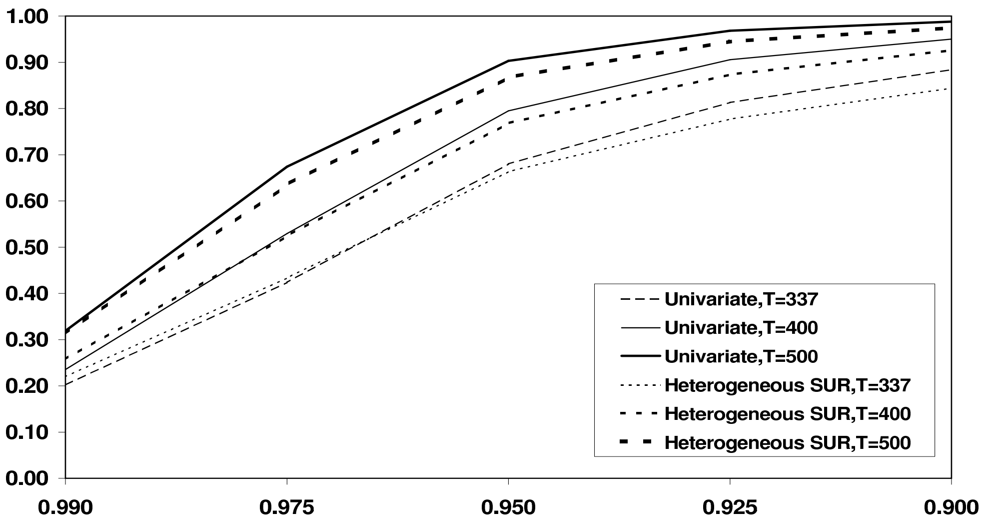


Figure 2.3: Estimated Power Functions of Unit Root Tests of Euro Area Real Exchange Rates: Univariate ADF versus SUR ADF. This figure presents the estimated power functions for both the univariate ADF and the heterogeneous SUR ADF unit root tests of the real exchange rates of nine euro area countries versus the Deutsche mark. The power functions reflect the average power of the unit root test for all nine individual currency pairs in the sample. Substantial variations in power exist, however, across the real exchange rates in the sample. The estimated power functions presented in this picture are based on Monte Carlo simulations with, respectively, 337, 400, and 500 time-series observations for each of the real exchange rates.

9×24=216 serial correlation terms and 9×2=18 coefficients) needs to be estimated. The solution to this estimation problem is not straightforward. O'Connell (1998) convincingly argues against imposing restrictions on the covariance matrix, while Papell and Theodoridis (2001) and Wu and Wu (2001) show that accounting for the (heterogeneous) serial correlation properties of real exchange rates is important. This chapter stresses the economic rationale of looking at individual currency pairs and removing the restriction of a common mean reversion parameter. While imposing complete uniformity across real rates leads to seemingly precise estimates, this is hard to reconcile with the heterogeneity in the observed behavior of individual real exchange rates.

Power functions vary considerably across real exchange rates. For example, the power of the unit root test against the alternative of $\beta_i = 0.990$ is 0.45 for Belgium and only 0.16 for Austria and Greece. Differences between univariate and multivariate power also depend on the country. While the power of the univariate test is higher for most countries, the multivariate test is more powerful for others.

Figure 2.4 compares the power functions of the univariate and multivariate tests for the euro and other major currencies with the euro as the numeraire currency. Again, the average power is quite low, although slightly higher than in Figure 2.3. With $T = 267$, the heterogeneous SUR model accepts stationarity for only about 60% of the cases when the half-life equals 14 months, while the rejection probability increases to no more than 80% when the half-life equals 7 months. The power of the univariate tests is considerably lower. When the half-life amounts to 14 months the probability of rejections is below 50%, while a half-life of 7 months is not enough to reject the unit root in 40% of the cases. Power does increase with sample size, but even in the multivariate case and over 40 years of data, a real exchange rate with a half-life of PPP deviations equal to less than 2.5 years is indicated to exhibit a unit root in almost 30 percent of the cases. Notable differences exist between individual currencies. For some real exchange rates in the sample, the power of the univariate tests is actually higher.[17]

Overall, our analysis indicates that the power of both univariate and multivariate tests of a unit root in real exchange rates is relatively low. Even real exchange rates that exhibit a half-life of one to two years are quite likely to be earmarked as non-stationary. While the power increases with the number of time-series observations available, even as many as 15 years of extra monthly observations after Bretton-Woods would not lead to high statistical power levels for mean reversion parameters smaller than 0.95. Furthermore, the power of univariate ADF tests is in some situations as least as high as the power of heterogeneous SUR ADF tests. This implies that researchers interested in the PPP hypothesis for individual real exchange rates do not necessarily benefit from SUR estimation. Evaluating the PPP for individual currency pairs is a precarious exercise. This is not likely to change as more data will become available in the near future.

An alternative for the classical approach of testing for PPP is Bayesian analysis. Classical tests for the null hypothesis of a unit root generally have moderate power

[17] In order to assess the influence of the serial correlation structure on the power of the PPP tests, we re-estimate the OLS ADF and SUR ADF models with the number of lags set to the maximum value of 24 for all countries. In general, the results indicate weaker evidence for PPP due to the decrease in degrees of freedom. In addition, we estimate both models with the restriction that all exchange rates have the same serial correlation properties ($\beta_i = \beta$ and $\gamma_{i,k} = \gamma_k$ for all i and k). The power does not increase for the euro area, but for the euro compared to other major currencies there are a few differences. However, as Papell and Theodoridis (2001) and Wu and Wu (2001) make a strong case against these restrictions, we do not pursue this any further.

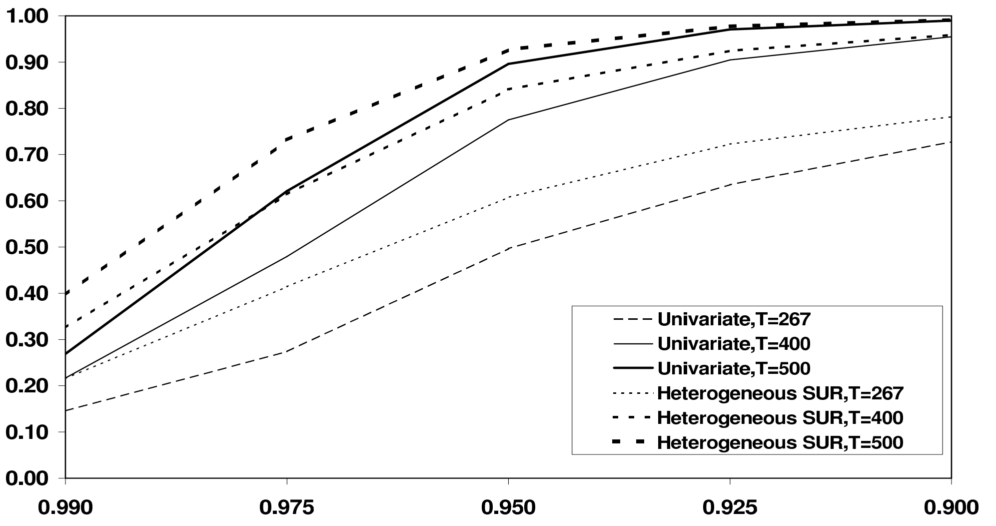


Figure 2.4: Estimated Power Functions of Unit Root Tests of the Real Exchange Rates of Several Major Currencies Against the Euro: Univariate ADF versus SUR ADF. This figure presents the estimated power functions for both the univariate ADF and the heterogeneous SUR ADF unit root tests of the real exchange rates of several major international currencies versus the euro. The power functions reflect the average power of the unit root test for all nine individual currency pairs in the sample. Substantial variations in power exist, however, across the real exchange rates in the sample. The estimated power functions presented in this picture are based on Monte Carlo simulations with, respectively, 267, 400, and 500 time-series observations for each of the real exchange rates.

against stationary alternative hypotheses. DeJong *et al.* (1992) show that tests of a trend-stationary null against the unit root alternative suffer from the same problem. An advantage of Bayesian approaches to evaluating PPP is that the unit root and stationarity hypotheses can be treated symmetrically. Bayesian methods allow for an assessment of the probability of a unit root in the data by evaluating the Bayesian posterior odds ratio.

Early empirical studies indicates that a Bayesian analysis of PPP may lead to different conclusions than classical tests. For example, Schotman and Van Dijk (1991) study the stationarity of eight real exchange rates over the period 1973-1988. They find that although classical tests are unable to reject the unit root null at the 5% level for all series, a Bayesian posterior odds analysis indicates that for six out of eight series the hypothesis of stationarity is as least as likely as the unit root hypothesis. A major problem with the Bayesian approach, however, is the specification of the prior distribution. This has been the topic of extensive debate. DeJong and Whiteman (1991b) show that only priors that assign a very low probability to the trend-stationarity support the classical results that most economic time-series contain a unit root. With a flat prior, however, stationarity is generally supported. Phillips (1991) demonstrates that flat priors, presumed uninformative by definition, in fact favor stationarity over the unit root hypothesis. Hence, the use of flat priors may seem objective, but is actually likely to bias the results in the direction of stationarity. DeJong and Whiteman (1991a) challenge this conclusion and question the priors used in Phillips' approach. They contend that there is a strong case for stationarity in many economic time-series. Koop (1992) employs a variety of alternative priors and concludes that "... the failure of classical procedures to reject the unit root hypothesis is not necessarily proof that a unit root is present with high probability" (p. 65).

Despite the problems involving the choice of the prior distribution and the complexity of computing posterior odds analytically, Bayesian approaches constitute an important alternative way of assessing unit roots in economic time-series. We consider the Bayesian analysis of the implications of European monetary integration for the behavior of real exchange rates to be an interesting area for further research.

2.7 Conclusions

In this chapter we study the effects of the ongoing economic integration in Europe on the behavior of real exchange rates. Specifically, we analyze the convergence toward purchasing power parity (PPP) within the euro area as well between the euro area and other major economies. The results are important for researchers in international economics, monetary policy makers as well as asset managers and investment practitioners.

We examine the unit root hypothesis for a panel of real exchange rates over three different periods in order to assess the impact of the Maastricht Treaty signed in 1992 and the introduction of the euro in 1999 on the stationarity of real exchange rates. In contrast to previous studies, we employ a Seemingly Unrelated Regression (SUR) methodology that allows the rate of mean version to vary across countries. We reckon that this heterogeneous SUR approach provides a more balanced and comprehensive view on PPP. Economically, we would expect that the speed of mean reversion is not the same across exchange rates and depends on, for example, the relative proximity of countries, their mutual trade regulations, and the openness of their economies. Following Papell and Theodoridis (2001) and Wu and Wu (2001), we account for heterogeneous serial correlation by performing Augmented Dickey-Fuller (ADF) tests. The O'Connell critique does not apply for our model because we do not impose any restrictions on the covariance matrix.

Our contribution is threefold. We stress the importance of incorporating different mean reversion parameters for different currency pairs and report results of unit root tests for all individual real exchange rates in the sample. In addition to an analysis of PPP within the euro area, we use "synthetic" euro data to study the validity of PPP between the euro area as a distinct economic entity and other major economies. Moreover, we use more recent as well as more frequent data than employed in recent PPP studies, which facilitates a detailed analysis of recent developments.

We find evidence in favor of PPP within the euro area with the DMark as numeraire. The half-life of PPP deviations over the period 1973-2003 is approximately three years. There has been a clear convergence process toward PPP within the euro area in the past decade. The adoption of the Maastricht Treaty has played an important role in this process. Accounting for intra-euro area differences in mean reversion across real exchange rates proofs to be vital. Half-lifes of PPP deviations vary widely across different currencies in the sample. Convergence processes toward PPP also show important differences across countries.

Our argument that focusing on individual real exchange rates in addition to the full panel of exchange rates is essential is underlined by the results of the unit root tests on a panel of major currencies including the (synthetic) euro. With the euro as the numeraire, the unit root hypothesis is rejected at the 5% significance level for the full panel of exchange rates over the period 1979-2003. With the exception of Switzerland, however, PPP does not hold between any of the individual countries and the euro area. There is no evidence that the increased economic integration in Europe has affected the evidence for PPP between the euro and other major currencies in a consistent way.

An analysis of the power of the univariate and heterogeneous SUR ADF tests suggests that caution should be applied in the interpretation of the test results. Monte Carlo simulations indicate that the power of both univariate and multivariate tests is relatively low. Moreover, although the power of the heterogeneous SUR ADF test generally exceeds the power of the univariate test, the differences are remarkably limited. This suggests that research on the PPP hypothesis for individual real exchange rates does not necessarily benefit importantly from adopting a multivariate approach. Concluding, evaluating PPP for individual currency pairs is a precarious exercise. Different currency pairs display different speeds of mean reversion and this calls for a heterogeneous unit root test. However, the power of the heterogeneous

SUR ADF tests is limited and not much higher than the power of the univariate tests, even for panels covering more than 40 years of monthly data. This inference underlines the merits of studying long-run time series of real exchange rates, as conducted by e.g. Edison (1987) and Lothian and Taylor (1996). An interesting alternative approach is a Bayesian analysis of unit roots in real exchange rates. We leave this suggestion for further research.

Table 2.2: Univariate ADF Unit Root Tests of Euro Area Real Exchange Rates Against the DMark. This table presents the results of the univariate ADF unit root tests of the real exchange rates of nine euro area countries versus the Deutsche mark over the period 1975:03-2003:03. For each exchange rate, we run the following regression: $R_{i,t} = \alpha_i + \beta_i R_{i,t-1} + \sum_{k=1}^{l_i} \gamma_{i,k} \Delta R_{i,t-k} + u_{i,t}$, where $R_{i,t}$ is the log of the real exchange rate. The value of l_i is determined by the recursive t-statistic procedure of Campbell and Perron (1991). The critical values have been obtained from Dickey and Fuller (1979). * and ** denote the significance of the test statistic $\tau_i = (\beta_i - 1)/s(\beta_i)$ at the 10% and 5% level, respectively.

Country	α_i	s.e.	β_i	s.e.	half-life	l_i	τ_i
Austria	-0.0022	0.0008	0.9878	0.0054	56	16	-2.26
Belgium	0.0009	0.0005	0.9712	0.0108	24	18	-2.66*
Finland	-0.0019	0.0012	0.9738	0.0096	26	17	-2.73*
France	-0.0006	0.0006	0.9255	0.0231	9	23	-3.23**
Greece	-0.0012	0.0012	0.9703	0.0171	23	18	-1.74
Italy	-0.0020	0.0013	0.9800	0.0103	34	23	-1.94
Netherlands	-0.0004	0.0004	0.9739	0.0150	26	24	-1.73
Portugal	0.0002	0.0013	0.9814	0.0120	37	21	-1.55
Spain	-0.0070	0.0026	0.9639	0.0131	19	17	-2.76*
10% critical value							-2.57
5% critical value							-2.88

Table 2.3: SUR ADF Unit Root Tests of Euro Area Real Exchange Rates Against the DMark: Common Mean Reversion Coefficient. This table presents the results of the SUR ADF unit root test of the real exchange rates of nine euro area countries versus the Deutsche mark over the period 1975:03-2003:03. We estimate the following system: $R_{i,t} = \alpha_i + \beta R_{i,t-1} + \sum_{k=1}^{l_i} \gamma_{i,k} \Delta R_{i,t-k} + u_{i,t}$, where $R_{i,t}$ is the log of the real exchange rate and the value of l_i is taken from the OLS ADF unit root test results. Hence, we impose the restriction that the estimate of the mean reversion coefficient is independent of the currency. The critical values have been obtained using Monte Carlo Simulations. * and ** denote the significance of the test statistic $\tau = (\beta - 1)/s(\beta)$ at the 10% and 5% level, respectively.

Country	α_i	s.e.	β	s.e.	half-life	l_i	τ
Austria	-0.0032	0.0005	0.9814	0.0033	37	16	-5.69**
Belgium	0.0006	0.0004				18	
Finland	-0.0014	0.0010				17	
France	-0.0002	0.0006				23	
Greece	-0.0010	0.0012				18	
Italy	-0.0019	0.0011				23	
Netherlands	-0.0004	0.0004				24	
Portugal	0.0002	0.0012				21	
Spain	-0.0041	0.0013				17	
10% critical value							-4.79
5% critical value							-5.21

Table 2.4: SUR ADF Unit Root Tests of Euro Area Real Exchange Rates Against the DMark: Heterogeneous Mean Reversion Coefficients. This table presents the results of the SUR ADF unit root test of the real exchange rates of nine euro area countries versus the Deutsche mark over the period 1975:03-2003:03. We estimate the following system: $R_{i,t} = \alpha_i + \beta_i R_{i,t-1} + \sum_{k=1}^{l_i} \gamma_{i,k} \Delta R_{i,t-k} + u_{i,t}$, where $R_{i,t}$ is the log of the real exchange rate and the value of l_i is taken from the OLS ADF unit root test results. The critical values [10% cv] as well as the empirical p-values have been obtained using Monte Carlo Simulations. * and ** denote the significance of the test statistic $\tau_i = (\beta_i - 1)/s(\beta_i)$ at the 10% and 5% level, respectively.

Country	α_i	s.e.	β_i	s.e.	half-life	l_i	τ_i [10% cv]
Austria	-0.0022	0.0007	0.9892	0.0050	64	16	-2.14 [-2.75]
Belgium	0.0007	0.0005	0.9764	0.0099	29	18	-2.38 [-2.80]
Finland	-0.0024	0.0012	0.9682	0.0084	21	17	-3.81** [-3.00]
France	-0.0005	0.0006	0.9337	0.0198	10	23	-3.35** [-3.06]
Greece	-0.0011	0.0012	0.9797	0.0163	34	18	-1.24 [-2.81]
Italy	-0.0024	0.0013	0.9761	0.0089	29	23	-2.68 [-3.06]
Netherlands	-0.0004	0.0004	0.9750	0.0135	27	24	-1.86 [-2.92]
Portugal	-0.0006	0.0013	0.9995	0.0106	1310	21	-0.05 [-2.98]
Spain	-0.0080	0.0023	0.9595	0.0112	17	17	-3.63** [-3.00]

p-values	Model H_0^1	Model H_0^2	Model H_0^3
	τ_i	τ_i	τ_i
Austria	0.288	0.195	0.265
Belgium	0.221	0.167	0.199
Finland	0.019**	0.006**	-
France	0.049**	0.021**	-
Greece	0.666	0.621	0.616
Italy	0.203	0.081*	0.075*
Netherlands	0.426	0.297	0.370
Portugal	0.961	0.895	0.881
Spain	0.028**	0.006**	-

Table 2.5: SUR ADF Unit Root Tests of Euro Area Real Exchange Rates: Subperiod Results. This table presents the p-values of the SUR ADF unit root tests under H_0^1 of the real exchange rates of nine euro area countries versus the Deutsche mark over the periods 1975:03-1991:12, 1975:03-1998:12 and 1975:03-2003:03. The empirical p-values have been obtained using Monte Carlo Simulations. * and ** denote the significance of the test statistic $\tau_i = (\beta_i - 1)/s(\beta_i)$ at the 10% and 5% level, respectively.

p-values	1975:03-1991:12	1975:03-1998:12	1975:03-2003:03
Model H_0^1	τ_i	τ_i	τ_i
	Homogeneous mean reversion coefficient		
	0.384	0.012**	0.010**
	Heterogeneous mean reversion coefficients		
Austria	0.575	0.484	0.288
Belgium	0.463	0.275	0.221
Finland	0.060*	0.009**	0.019**
France	0.077*	0.049**	0.049**
Greece	0.330	0.449	0.666
Italy	0.798	0.164	0.203
Netherlands	0.358	0.380	0.426
Portugal	0.260	0.855	0.961
Spain	0.186	0.057*	0.028**

Table 2.6: Correlations of the Real Exchange Rates of Several Major Economies. This table shows correlations of the first differences in the real exchange rates $R_{i,t} - R_{i,t-1}$ of several major international currencies versus the euro over the period 1978:12-2003:03.

Country	Can	Den	Jap	Nor	Swe	Swi	U.K.	U.S.
Canada	1.000	0.015	0.334	0.346	0.325	-0.189	0.280	0.893
Denmark		1.000	0.099	0.111	0.055	0.249	-0.298	-0.011
Japan			1.000	0.170	0.155	0.175	0.156	0.373
Norway				1.000	0.425	-0.074	0.167	0.351
Sweden					1.000	-0.069	0.198	0.289
Switzerland						1.000	-0.193	-0.154
U.K.							1.000	0.281
U.S.								1.000

Table 2.7: Univariate ADF Unit Root Tests of the Real Exchange Rates of Several Major Currencies Against the Euro. This table presents the results of the univariate ADF unit root tests of the real exchange rates of several major international currencies versus the euro over the period 1981:01-2003:03. For each exchange rate, we run the following regression: $R_{i,t} = \alpha_i + \beta_i R_{i,t-1} + \sum_{k=1}^{l_i} \gamma_{i,k} \Delta R_{i,t-k} + u_{i,t}$ where $R_{i,t}$ is the log of the real exchange rate. The value of l_i is determined by the recursive t-statistic procedure of Campbell and Perron (1991). The critical values have been obtained from Dickey and Fuller (1979). * and ** denote the significance of the test statistic $\tau_i = (\beta_i - 1)/s(\beta_i)$ at the 10% and 5% level, respectively.

Country	α_i	s.e.	β_i	s.e.	half-life	l_i	τ_i
Canada	-0.0015	0.0021	0.9699	0.0149	23	21	-2.02
Denmark	-0.0016	0.0011	0.9827	0.0136	40	24	-1.26
Japan	-0.0062	0.0033	0.9718	0.0137	24	16	-2.06
Norway	-0.0010	0.0010	0.9589	0.0221	17	23	-1.86
Sweden	0.0060	0.0030	0.9686	0.0179	22	20	-1.75
Switzerland	-0.0075	0.0031	0.9625	0.0166	18	13	-2.26
U.K.	0.0052	0.0028	0.9701	0.0148	23	22	-2.03
U.S.	-0.0031	0.0022	0.9662	0.0130	20	13	-2.61*
10% critical value							-2.57
5% critical value							-2.88

Table 2.8: SUR ADF Unit Root Tests of the Real Exchange Rates of Several Major Currencies Against the Euro: Common Mean Reversion Coefficient. This table presents the results of the SUR ADF unit root test of the real exchange rates of several major international currencies versus the euro over the period 1981:01-2003:03. We estimate the following system: $R_{i,t} = \alpha_i + \beta R_{i,t-1} + \sum_{k=1}^{l_i} \gamma_{i,k} \Delta R_{i,t-k} + u_{i,t}$ where $R_{i,t}$ is the log of the real exchange rate and the value of l_i is taken from the OLS ADF unit root test results. Hence, we impose the restriction that the estimate of the mean reversion coefficient is independent of the currency. The critical values have been obtained using Monte Carlo Simulations. * and ** denote the significance of the test statistic $\tau = (\beta - 1)/s(\beta)$ at the 10% and 5% level, respectively.

Country	α_i	s.e.	β	s.e.	half-life	l_i	τ
Canada	-0.0011	0.0020	0.9781	0.0046	31	21	-4.75*
Denmark	-0.0019	0.0006				24	
Japan	-0.0051	0.0021				16	
Norway	-0.0006	0.0009				23	
Sweden	0.0045	0.0015				20	
Switzerland	-0.0047	0.0012				13	
U.K.	0.0039	0.0015				22	
U.S.	-0.0020	0.0019				13	
10% critical value							-4.54
5% critical value							-4.95

Table 2.9: SUR ADF Unit Root Tests of the Real Exchange Rates of Several Major Currencies Against the Euro: Heterogeneous Mean Reversion Coefficients. This table presents the results of the SUR ADF unit root test of the real exchange rates of several major international currencies versus the euro over the period 1981:01-2003:03. We estimate the following system: $R_{i,t} = \alpha_i + \beta_i R_{i,t-1} + \sum_{k=1}^{l_i} \gamma_{i,k} \Delta R_{i,t-k} + u_{i,t}$ where $R_{i,t}$ is the log of the real exchange rate and the value of l_i is taken from the OLS ADF unit root test results. The critical values [10% cv] as well as the empirical p-values have been obtained using Monte Carlo Simulations. * and ** denote the significance of the test statistic $\tau_i = (\beta_i - 1)/s(\beta_i)$ at the 10% and 5% level, respectively.

Country	α_i	s.e.	β_i	s.e.	half-life	l_i	τ_i [10% cv]
Canada	-0.0007	0.0020	0.9853	0.0076	47	21	-1.93 [-3.40]
Denmark	-0.0014	0.0010	0.9865	0.0124	51	24	-1.09 [-2.92]
Japan	-0.0071	0.0031	0.9676	0.0125	21	16	-2.60 [-2.83]
Norway	-0.0011	0.0010	0.9519	0.0200	14	23	-2.40 [-2.99]
Sweden	0.0056	0.0028	0.9709	0.016	24	20	-1.82 [-2.87]
Switzerland	-0.0092	0.0029	0.9527	0.0159	14	13	-2.98* [-2.84]
U.K.	0.0055	0.0026	0.9684	0.0132	22	22	-2.39 [-2.86]
U.S.	-0.0015	0.0020	0.9835	0.0068	42	13	-2.41 [-3.42]

p-values	Model H_0^1	Model H_0^2	Model H_0^3
	τ_i	τ_i	τ_i
Canada	0.581	0.156	0.547
Denmark	0.716	0.612	0.676
Japan	0.163	0.104	0.135
Norway	0.238	0.157	0.208
Sweden	0.483	0.340	0.458
Switzerland	0.074*	0.051*	-
U.K.	0.236	0.190	0.225
U.S.	0.383	0.102	0.364

Table 2.10: SUR ADF Unit Root Tests of the Several Major Real Exchange Rates: Subperiod Results. This table presents the p-values of the SUR ADF unit root tests under H_0^1 of the real exchange rates of several major international currencies versus the euro over the periods 1981:01-1991:12, 1981:01-1998:12 and 1981:01-2003:03. The empirical p-values have been obtained using Monte Carlo Simulations. * and ** denote the significance of the test statistic $\tau_i = (\beta_i - 1)/s(\beta_i)$ at the 10% and 5% level, respectively.

p-values	1981:01-1991:12	1981:01-1998:12	1981:01-2003:03
Model H_0^1	τ_i	τ_i	τ_i
	Homogeneous mean reversion coefficient		
	0.004**	0.036**	0.076*
	Heterogeneous mean reversion coefficients		
Canada	0.116	0.587	0.581
Denmark	0.616	0.460	0.716
Japan	0.648	0.399	0.163
Norway	0.024**	0.197	0.238
Sweden	0.430	0.106	0.483
Switzerland	0.002**	0.062*	0.074*
U.K.	0.058*	0.017**	0.236
U.S.	0.308	0.450	0.383

Chapter 3

Estimation and Tests of Purchasing Power Parity

3.1 Introduction

Since the early 1990s, researchers interested in testing the hypothesis of purchasing power parity (PPP) have turned to multivariate testing procedures in order to increase the statistical power.[1] Initial applications of multivariate analysis to real exchange rates imposed severe restrictions on the structure of the model. Two of these restrictions have been successfully challenged in the recent literature. First, O'Connell (1998) questions the common assumption in the PPP literature that the real exchange rates are cross-sectionally independent.[2] He demonstrates that spurious rejections of the unit root null can occur when cross-sectional dependence is neglected. In response to his critique, nearly all subsequent research relaxes this restriction and thus takes cross-sectional dependence into account. Second, Papell and Theodoridis (2001) and Wu and Wu (2001) criticize the prevalent restriction that the serial correlation properties of all real exchange rates in the panel are the same. Both papers show that assuming a restrictive homogeneous serial correlation structure weakens the evidence against the unit root null. In line with these findings, recent panel studies on PPP abandon this second restriction as well.

[1] Examples of early multivariate PPP studies include Abuaf and Jorion (1990), Frankel and Rose (1996), and Jorion and Sweeney (1996). We refer to Rogoff (1996) and Taylor and Taylor (2004) for reviews of the literature.

[2] Specifically, O'Connell (1998) examines the restriction that the variance-covariance matrix of the residuals in the panel model is diagonal.

Yet, the use of a third important restriction on the structure of multivariate models of real exchange rates is still widespread in the academic literature on PPP. The vast majority of recent empirical studies assume a common mean reversion coefficient across all real exchange rates.[3] From an economic perspective, the justification for the assumption that PPP holds equally well for all country pairs is weak. The speed of mean reversion of a real exchange rate between two countries should depend on, for example, their relative proximity, their mutual trade regulations, and the openness of their economies. The econometric consequences of imposing homogeneous mean reversion for the properties of multivariate PPP tests have not been thoroughly investigated to date. Exploratory econometric research on the properties of panel data models (notably Robertson and Symons (1992) and Pesaran and Smith (1995)) suggests that the homogeneity assumption in dynamic panel models may have serious consequences. Pooling heterogeneous panel data can lead to biases in the parameter estimates, as a result of which estimation results are potentially misleading. An important question is to what extent this affects the inferences drawn from multivariate studies on the PPP hypothesis.

This chapter analyzes the finite-sample properties of various multivariate unit root tests employed for investigating PPP. In order to assess the consequences of the homogeneity restriction, we compare three different multivariate estimation and testing methodologies. The first methodology involves homogeneous estimation of the mean reversion parameters and a unit root test on the validity of PPP for the full panel of real exchange rates. This methodology, or a variation thereof, is applied by a large number of recent empirical papers on PPP (examples are provided in footnote 3). The second methodology entails estimating the model heterogeneously, but still testing the PPP hypothesis jointly for all series in the panel. This means that while any inferences about PPP still concern the panel as a whole, differences in mean reversion across countries are allowed for. This approach is taken by, among others, Taylor and Sarno (1998), Im *et al.* (2003), Wu and Wu (2001). In addition, we propose an alternative methodology in which both estimation and testing are performed heterogeneously.

We employ Monte Carlo simulation in order to examine the empirical performance of the three multivariate methodologies. Our Monte Carlo experiments are based on a sample of the real exchange rates between five of the world's largest economies (Canada, the euro area, Japan, the U.K., and the U.S.) over the period 1978:12-2003:12. We demonstrate that when the assumption of homogeneous mean

[3] See e.g. Choi (2004), Fleissig and Strauss (2000), Frankel and Rose (1996), Jorion and Sweeney (1996), Levin *et al.* (2002), Lothian (1997), Lopez and Papell (2004), O'Connell (1998), Papell (1997), Papell (2002), and Papell and Theodoridis (2001).

reversion in the panel of real exchange rates is violated, the methodology that relies upon homogeneous estimation and testing suffers from important adverse properties. First, homogeneous estimates of the mean reversion parameter can exhibit serious biases. Second, large estimation uncertainties arise as a result of the homogeneity restriction. This implies that the statistical power of the homogeneous test against the unit root null is generally limited. Third, the power function is not monotonically increasing when the mean reversion parameters generated under the alternative hypothesis move away from the unit root null. These results indicate that a homogeneous estimation methodology can lead to potentially misleading inferences about the validity of the PPP hypothesis. This chapter offers critical insights into the consequences of imposing parameter restrictions in multivariate tests for PPP by showing that these testing methodologies should not only take cross-sectional dependence and heterogeneous serial correlation into account, but also heterogeneous mean reversion.

The remainder of the chapter is structured as follows. In Section 3.2 we describe the methodology. Section 3.3 provides the data description. The results of our Monte Carlo simulations are presented in Section 3.4. Section 3.5 concludes.

3.2 Methodology

3.2.1 Multivariate tests of PPP

This section discusses the three different multivariate estimation and testing methodologies we examine in this study. For each country (currency) i ($i = 1, ..., N$) we define the log real exchange rate at time t ($t = 1, ..., T$) as follows:

$$R_{i,t} = e_{i,t} - e_{0,t} + p_{0,t} - p_{i,t} \tag{3.1}$$

where $R_{i,t}$ is the logarithm of the real exchange rate, $e_{i,t}$ is the logarithm of the nominal exchange rate expressed in units of currency i per dollar, $e_{0,t}$ is the logarithm of the nominal exchange rate expressed in units of the numeraire currency 0 per dollar, $p_{0,t}$ is the logarithm of the consumer price index of the country used as numeraire, and $p_{i,t}$ is the logarithm of the consumer price index in country i.

The three methodologies analyzed in this chapter are all based on the following multivariate regression:

$$R_{i,t} = \alpha_i + \beta_i R_{i,t-1} + \sum_{k=1}^{l_i} \gamma_{i,k} \Delta R_{i,t-k} + u_{i,t}, \tag{3.2}$$

where β_i are the mean reversion parameters, α_i are the intercepts, $\gamma_{i,k}$ are the coefficients on the lagged real exchange rate returns, l_i denotes the number of lags needed for currency i, and $u_{i,t}$ is a stationary error term: $u \sim N(0, \Sigma)$. The null-hypothesis of a unit root is expressed by $H_0 : \beta_i = 1$ for all currencies i. The three methodologies differ in the estimation of the parameters in Equation (3.2) as well as in the way the test of the unit root hypothesis is performed. For all three methodologies, we incorporate the suggestions made by O'Connell (1998) to allow for contemporaneous correlations between the error terms $u_{i,t}$ and by Papell and Theodoridis (2001) and Wu and Wu (2001) to allow the serial correlation parameters to vary across exchange rates.

In the first methodology, estimation of Equation (3.2) and testing the unit root hypothesis are carried out homogeneously. That is, the restriction $\beta_i = \beta$ is imposed for all i and the null-hypothesis and alternative hypothesis can thus be expressed as $H_0 : \beta = 1$ and $H_A : \beta < 1$, respectively. This implies that the PPP hypothesis is only evaluated for the panel as a whole and not for individual country pairs. A large number of recent empirical papers adopt (a variation of) this methodology (see footnote 3). For instance, O'Connell (1998) applies this methodology in order to study the restriction that exchange rates are cross-sectionally independent. Accounting for cross-sectional dependence, he finds no empirical evidence for PPP in a panel of 64 countries over the period 1973:I-1995:IV. As a second example of homogeneous estimation and testing, Lopez and Papell (2004) investigate convergence towards PPP within the euro zone and between the euro zone and other countries. Using data over the period 1973:I-2001:IV, they present evidence of convergence toward PPP within the euro zone, but not for the real exchange rates of the euro versus other currencies.

The second methodology we investigate involves heterogeneous estimation of the mean reversion parameters β_i in Equation (3.2), but still testing the unit root hypothesis in a homogeneous way. This means that while inferences about PPP concern the entire panel and no statements can be made about individual country pairs, the speed of mean reversion is allowed to differ across countries. The null-hypothesis can be expressed as $H_0 : \beta_i = 1 (\forall i = 1, ..., N)$, while there are several possibilities for the alternative, for example, $H_A : \exists j \in \{1, ..., N\} : \beta_j < 1$ or $H_A : \forall i \in \{1, ..., N\} : \beta_i < 1$. The interpretation of the first alternative is that at least one of the exchange rates is stationary, the second states that all exchange rates are mean-reverting. This implies that rejection of the unit root null does not necessarily provide information on how many and which real exchange rates in the panel are stationary. An example of a study using this methodology is Taylor and Sarno (1998). They estimate an equivalent multivariate model as in Equation (3.2) and perform a joint unit root test on all N equations by constructing a standard Wald test statistic, which they refer to as the Multivariate Augmented Dickey-Fuller (MADF) test. The alternative hypothesis for this test is $H_A : \exists j \in \{1, ..., N\} : \beta_j < 1$. Critical values for this test are obtained by Monte Carlo simulation. The MADF test rejects the unit root null in a panel containing the UK, France, Germany, Japan, and the U.K. (with the U.S. as numeraire country) over the period 1973:I-1996:II. Taylor and Sarno (1998) suggest an alternative testing methodology, based on the Johansen likelihood ratio (JLR) test for cointegration, which tests the unit root null hypothesis versus the alternative that $H_A : \forall i \in \{1, ..., N\} : \beta_i < 1$. The JLR test also rejects the null hypothesis of non-stationarity for their panel of exchange rates. Other papers that use a methodology in which estimation is performed heterogeneously, but the unit root null hypothesis is defined for all real exchange rates in the panel simultaneously, are Im *et al.* (2003) and Wu and Wu (2001). Im *et al.* (2003) develop a unit root test with the same null hypothesis as described above, but with alternative hypothesis $H_A : \beta_i < 1, i = 1, ..., N_1, \beta_i = 1, i = N_{1+1}, ..., N$. Hence, a rejection of the null-hypothesis in this model does not necessarily mean that all exchange rates in the panel are mean-reverting. Wu and Wu (2001) apply the tests developed in the paper of Maddala and Wu (1999) and in an earlier version of the paper of Im *et al.* (2003). Both tests are based on univariate ADF regressions and have $\forall i \in \{1, ..., N\} : \beta_i < 1$ as alternative hypothesis.[4] Wu and Wu (2001) take heterogeneous serial correlation and cross-sectional dependence into account and document substantial evidence for PPP in a panel of 20 industrial countries over the period 1973:II-1997:IV.

The third methodology entails both heterogeneous estimation and heterogeneous testing of Equation (3.2). The unit root null-hypothesis and alternative hypothesis for every real exchange rate $i (i = 1, ..., N)$ are $H_0 : \beta_i = 1$ and $H_A : \beta_i < 1$, respectively. In this framework, the PPP tenet is evaluated for each individual real exchange rate in the panel. Flôres *et al.* (1999) apply such a methodology and report evidence in favor of long-run PPP for the G10 currencies versus the U.S. dollar in the period 1973:01-1994:12. Koedijk *et al.* (2004) enhance this approach by allowing for heterogeneous serial correlation. They investigate a panel of ten countries within the euro area over the period 1973:02-2003:03 as well as for the real exchange rates of the euro versus other major currencies over the period 1978:12-2003:03. The empirical evidence shows that the mean reversion properties differ importantly across real exchange rates. Three out of the nine real exchange rates within the euro area are mean reverting and only one out of the eight real exchange rates versus the euro is stationary. Finally, Engel *et al.* (1997) estimate a panel of real exchange rates for eight cities in four countries and two continents and allow for different speeds of mean reversion for within-country, within-continent, and cross-country city pairs.

[4] Note that this alternative hypothesis is different from the one used in the published version of the Im, Pesaran, and Shin paper.

They fail to reject the unit root hypothesis based on any of these mean reversion parameters.

3.2.2 Estimation procedure

All three methodologies rely upon Seemingly Unrelated Regression (SUR) estimation of Equation (3.2). Before we estimate the model, we determine the number of lags l_i for each currency i by applying the recursive t-statistic procedure of Campbell and Perron (1991) to each individual log real exchange rate.[5] The estimation procedure is then as follows. First, for each currency i, we apply OLS to Equation (3.2). Second, the resulting covariance matrix of the error terms is used as the weighting matrix in a Feasible Generalized Least Squares (FGLS) procedure to estimate the full panel.[6,7] When applicable, we impose the homogeneity restriction in the second step.[8] For the methodology in which the model is estimated and tested homogeneously, we compute the usual ADF test statistic $\tau = (\beta - 1)/s(\beta)$ to evaluate the unit root null-hypothesis. In order to test the unit root hypothesis homogeneously when estimation is carried out heterogeneously, we employ the MADF test as described in Taylor and Sarno (1998). For the methodology with heterogeneous estimation and testing, inferences about the stationarity of the individual real exchange rates are based on the individual ADF τ-statistics $\tau_i = (\beta_i - 1)/s(\beta_i)$.

We use Monte Carlo simulations to derive critical values for the test statistics of the various multivariate tests. In the first step, given the estimated parameters of the model α_i, β_i, $\gamma_{i,k}$, and l_i we compute residuals and the corresponding covariance matrix. Second, we generate N error terms $u_{i,t}$ (T times) from a multivariate normal distribution with mean zero and this covariance matrix. Third, given the estimated

[5]This means that for currency i we choose the value of l_i by first setting l_i to some maximal value, l_{max}, then estimate Equation (3.2) by OLS and subsequently test whether the last included lag is statistically significant. If so, then l_i is set to this value, else the model is estimated by setting l_i to $l_{max} - 1$. The procedure is repeated until a significant value of l_i is found. When no lag is significant then l_i is set to 0. Following Wu and Wu (2001), we set l_{max} to 24 and use a 10% significance level.

[6]The estimation procedure can be extended by iterating between estimating parameters, constructing the corresponding residuals, and using the covariance matrix of these residuals in the FGLS step. However, following Taylor and Sarno (1998), we limit this iterative process to one iteration for computational reasons. Unreported results show that this does not have an important effect on the resulting estimates.

[7]In a standard SUR model the degrees of freedom needed for the calculation of the covariance matrix of the error terms equals T-2. We have to make a correction for the inclusion of lagged changes in the real exchange rate in the model. To that effect, we set the degrees of freedom to T-2-entier[$(l_i + l_j + 1)/2$], where entier[x] rounds x down to the nearest integer value.

[8]Note that these homogeneity restrictions only apply to the mean reversion parameters (β_i) and not to the intercepts (α_i).

parameters $\gamma_{i,k}$ and l_i, assuming $\alpha_i = 0$, and imposing the null-hypothesis $\beta_i = 1$, we compute simulated exchange rate series using Equation (3.2) on the basis of the simulated error terms $u_{i,t}$ from step 2. Fourth, given the value of l_i, we estimate the parameters in Equation (3.2) with the simulated exchange rate series and compute the test statistic. We replicate step 2 to step 4 1,000 times and derive critical values for the test statistic from its sample distribution. Alternatively, empirical p-values can be calculated as the fraction of times the observed test statistic using the actual empirical data series is exceeded in the replications.

3.2.3 Monte Carlo simulations

Our purpose is to investigate the finite-sample properties of the three methodologies, and specifically potential estimation biases and the statistical power of the unit root test. Our approach is inspired by Taylor and Sarno (1998) and is based on Monte Carlo experiments. We simulate real exchange rate series for several data generating processes (DGP's) using Equation (3.2) as described above but instead of setting $\beta_i = 1$ for all exchange rates to construct the distribution of the ADF test statistic under the unit root null-hypothesis (see step 3 of Section 3.2.2), we assign other values to the mean reversion parameters. We follow Taylor and Sarno (1998) and use the following set of values for $\beta_i : \{0.99, 0.975, 0.95, 0.925, 0.9\}$. These rates of mean reversion correspond to half-lives of PPP deviations amounting to, respectively, 69, 27, 14, 9, and 7 months (computed as $\ln(0.5)/\ln(\beta\ )$). The power functions of the tests are subsequently constructed by comparing the simulated values of the test statistic to the 5% critical value obtained using the procedure described in Section 3.2.2. In order to gain further insight into the power properties of the multivariate tests, we monitor the estimate of the mean reversion parameter and the estimate of the uncertainty of this parameter. In particular, we assess the bias of the mean reversion parameter estimate that can arise because of the homogeneity restriction (see e.g. Robertson and Symons (1992) and Pesaran and Smith (1995) as well as the small sample (see e.g. Murray and Papell (2002)).

In order to avoid confusion in our description of the properties of the three different methodologies in subsequent sections, we introduce the following notational shorthand. We refer to the first methodology as the HoHo methodology (homogeneous estimation and testing), to the second as the HeHo methodology (heterogeneous estimation and homogeneous testing), and to the final as the HeHe methodology (heterogeneous estimation and testing).

3.3 Data

We collect consumer price index (CPI) and nominal exchange rate data for Canada, the euro area, Japan, the U.K., and the U.S. for the period 1978:12-2003:12. CPI data and period-ending exchange rates against the U.S. dollar are obtained from International Financial Statistics. Because the euro - dollar rate is only available from January 1999, we use the "synthetic" euro from the ECB.[9] In order to construct the CPI data for the euro area we employ the geometric weighted average method as described on page 11 in Maeso-Fernandez *et al.* (2001). Ireland is discarded because the CPI is only available as of 1997. The first 25 observations are used to compute the lagged exchange rate changes needed for the ADF tests. This implies that 276 time-series observations are used for the estimation of the model and we also use $T = 276$ in the Monte Carlo simulations.

Figure 3.1 shows the development of the log real exchange rates against the U.S. dollar over the period January 1981 through December 2003. As is obvious from the plots, the selection of countries in our sample has not been driven by the aim to maximize the heterogeneity of the mean reversion properties of the real exchange rates in our panel. In particular, the real exchange rates of the euro and the British pound against the U.S. dollar seem to exhibit very similar time-series patterns. Based on the graphs, both series also appear to share common characteristics with the dollar - yen series, while only the behavior of the Canadian dollar versus the U.S. dollar looks markedly different. The similarities across the real exchange rates are to a considerable extent driven by the large swings in the relative value of the dollar in the 1980s as well as in the past five years. Using a different numeraire currency could potentially lead to more pronounced heterogeneity in the mean reverting behavior of the real exchange rates in the panel. Further heterogeneity would probably be introduced by adding countries that are less well integrated with the world economy, such as emerging markets in Asia, Latin-America, and Eastern Europe.[10] Hence, we bias the results against finding important adverse consequences resulting from the homogeneity restriction in multivariate models.

Table 3.1 presents summary statistics for the real exchange rates as well as results of univariate ADF unit root tests. The correlations between the real exchange rates of the euro, pound, and yen against the dollar are substantial. This underlines

[9] This synthetic euro series is the "ECB reference exchange rate, U.S. dollar/euro, 2:15 pm (C.E.T.), against ECU up to December 1998," which is available at the website of the European Central Bank (www.ecb.int).

[10] Note that heterogeneity can also arise within closely integrated economic regions. Koedijk *et al.* (2004) present evidence that real exchange rates within the euro area exhibit important differences in mean-reverting behavior.

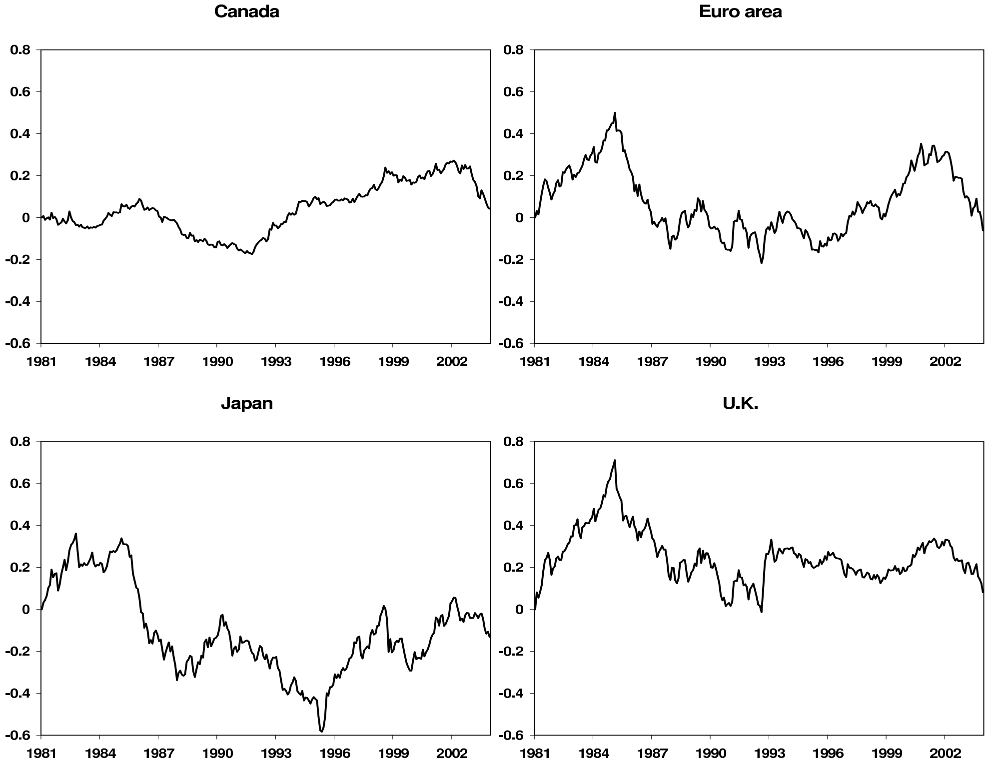


Figure 3.1: Real Exchange Rates Against the U.S. Dollar. This figure presents the (log) real exchange rates of Canada, the Euro area, Japan, and the U.K. versus the U.S. for the sample period 1981:01-2003:12.

the importance of accounting for cross-sectional dependence in panel tests of PPP. Moreover, the bottom panel of Table 3.1 shows that the serial correlation properties are not identical across the real exchange rate series. The optimal lag length in the univariate unit root tests, determined with the procedure of Campbell and Perron (1991), varies from 13 for the euro - dollar series to 22 for the yen - dollar rate. Hence, it is important to allow for heterogeneous serial correlation. Notwithstanding the similarities among three of the four real exchange rate series in Figure 3.1, the estimates of the α_i and β_i parameters in Equation (3.2) show considerable differences across the countries examined. The estimated half-life of PPP deviations ranges from about 1 year (13 months) for the U.K. to almost 3.5 years (41 months) for Canada. The unit root hypothesis can be rejected for the U.K. at the 5% level and for the euro area at the 10% level.

Table 3.1: Summary Statistics and Univariate ADF Unit Root Tests. This table presents summary statistics and the results of the univariate ADF unit root tests of the real exchange rates of Canada, the Euro area, Japan, and the U.K. versus the U.S. over the period 1981:01-2003:12 (276 time-series observations). We estimate the following equation: $R_{i,t} = \alpha_i + \beta_i R_{i,t-1} + \sum_{k=1}^{l_i} \gamma_{i,k} \Delta R_{i,t-k} + u_{i,t}$ where $R_{i,t}$ is the log of the real exchange rate and the value of l_i is determined by the recursive t-statistic procedure of Campbell and Perron (1991). The critical values have been obtained from Dickey and Fuller (1979). * and ** denote significance of the test statistic $\tau_i = (\beta_i - 1)/s(\beta_i)$ at the 10% and 5% level, respectively.

Country	Mean	St.dev.	Skewness	Kurtosis	Correlation coefficient		
Canada	0.0401	0.1171	0.1368	2.1163	Canada	Euro Area	Japan
Euro Area	0.0769	0.1568	0.4770	2.3235	0.3925		
Japan	-0.1078	0.2087	0.4063	2.5411	-0.0106	0.7553	
U.K.	0.2523	0.1235	0.8751	4.4304	0.1438	0.7368	0.5362
	α_i	s.e.	β_i	s.e.	half-life	l_i	τ_i
Canada	0.0003	0.0009	0.9833	0.0079	41	16	-2.10
Euro Area	0.0022	0.0021	0.9657	0.0127	20	13	-2.70*
Japan	-0.0021	0.0023	0.9819	0.0108	38	22	-1.69
U.K.	0.0135	0.0048	0.9464	0.0178	13	21	-3.01**
10% critical value							-2.57
5% critical value							-2.88

3.4 Empirical analysis

This section investigates the finite-sample properties of the three multivariate methodologies discussed in Section 3.2. We do not intend to offer an exhaustive treatment of the properties of unit root tests in heterogeneous panels. Rather, the goal is to illustrate how homogeneous estimation and testing of a panel can affect the inferences, in particular when examining the PPP hypothesis in a panel consisting of real exchange rates.

For all three methodologies we estimate the parameters of model 3.2 and test the PPP hypothesis by comparing the (multivariate) unit root test statistics to their corresponding simulated critical values. Table 3.2 displays the results of estimating and testing PPP with the HoHo methodology, imposing homogeneous mean reversion. The half-life is estimated at roughly 2.5 years and the unit root hypothesis is rejected

Table 3.2: SUR ADF Unit Root Tests: Common Mean Reversion Coefficient (HoHo methodology). This table presents the results of the SUR ADF unit root test of the real exchange rates of Canada, the Euro area, Japan, and the U.K. versus the U.S. over the period 1981:01-2003:12. We estimate the following system: $R_{i,t} = \alpha_i + \beta R_{i,t-1} + \sum_{k=1}^{l_i} \gamma_{i,k} \Delta R_{i,t-k} + u_{i,t}$, where $R_{i,t}$ is the log of the real exchange rate and the value of l_i is taken from the OLS ADF unit root test results. Hence, we impose the restriction that the mean reversion coefficient does not vary across countries. The critical values have been obtained using Monte Carlo Simulations. * and ** denote significance of the test statistic $\tau = (\beta - 1)/s(\beta)$ at the 10% and 5% level, respectively.

Country	α_i	s.e.	β	s.e.	half-life	l_i	τ
Canada	0.0005	0.0009	0.9765	0.0056	29	16	-4.23**
Euro Area	0.0016	0.0019				13	
Japan	-0.0027	0.0021				22	
U.K.	0.0061	0.0023				21	
10% critical value							-3.59
5% critical value							-3.95

(p-value = 0.023). This suggests that the PPP tenet is a reasonable description of the panel as a whole. Note that the τ-test and the Wald test are equivalent in case the unit root hypothesis is tested homogeneously. Table 3.3 presents the results for the HeHo and HeHe methodologies. Both methodologies require heterogeneous estimation of the α_i and β_i parameters. Again, the estimated half-lives of PPP deviations differ importantly across the real exchange rates, varying from less than 1.5 years to almost 3.5 years. In the HoHo methodology, this heterogeneity is obscured by the common mean reversion parameter estimate of 29 months. Table 3.3 shows much less clear-cut evidence in favor of PPP. The Wald (MADF) test rejects the unit root null only at the 10% level for the panel as a whole (HeHo methodology) and the ADF τ-statistic (HeHe methodology) is only significant for the euro - dollar rate (5% level).

Inferences regarding the validity of the PPP hypothesis in our panel are highly dependent on the chosen methodology. While the HoHo methodology yields strong rejection of the unit root hypothesis, the HeHe methodology indicates that only one of the four real exchange rates is stationary. Adopting the HeHo methodology leads to an intermediate result, a rejection of the unit root at the 10% level. A thorough understanding of the consequences of methodology choice requires an analysis of the finite-sample properties of the unit root test across the three methodologies. Therefore, we perform a Monte Carlo study on the empirical performance of the three methodologies. Naturally, when comparing the performance of alternative methodologies to detect unit roots in panels, the most interesting experiments concern DGP's that contain both stationary series and series with unit roots. Hence, we follow Taylor and Sarno (1998) and analyze the properties of the methodologies when, respectively, one, two, and three of the simulated series have a root less than unity. Subsequently, we study a variety of DGP's that contain only stationary exchange rates.

3.4.1 Monte Carlo evidence on panels with both stationary and non-stationary series

In this section we examine DGP's in which one or more of the real exchange rates in the panel contain a unit root under the alternative hypothesis. First, we consider DGP's with three non-stationary series. The mean reversion parameter (β_i) of the remaining real exchange rate is set equal to values from the set $\{0.99, 0.975, 0.95, 0.925, 0.9\}$. We then apply the HoHo, HeHo, and HeHe methodologies to estimate the parameters in Equation (3.2) and compute the test statistics. Comparing these statistics with the critical values yields a power function for each of the methodologies. We repeat this procedure for all four exchange rates. Figure 3.2 displays the power functions of the HoHo and HeHo tests for the stationarity of the panel as a whole as well as the HeHe tests evaluating the stationarity of the four exchange rate series individually. Chart titles indicate which of the simulated real exchange rate series is stationary. The first thing that leaps to the eye is the remarkable behavior of the power function of the unit root test of the HoHo methodology. The power of this methodology to reject the unit root null is very low and often even lower than the size of the test (5%). Furthermore, if the β of the mean reverting exchange rate decreases from 0.99 to 0.9, the power is not monotonically increasing. Especially for Canada, the power decreases when the moves further away from the unit root null. This is remarkable, as the power of the test should increase when the panel as a whole becomes more stationary.

The behavior of the four individual unit root tests in the HeHe methodology is essentially as expected. The power to reject the unit root null for the non-stationary series is very low and roughly equal to the 5% probability of a type I error. The power function of the HeHe test for the stationary series is the highest of all methodologies and clearly increases when β moves further away from unity. The power to reject the null depends on the particular real exchange rate that is stationary. For example, for a half-life of 14 months ($\beta = 0.95$) the power is 0.83 for the real exchange rate

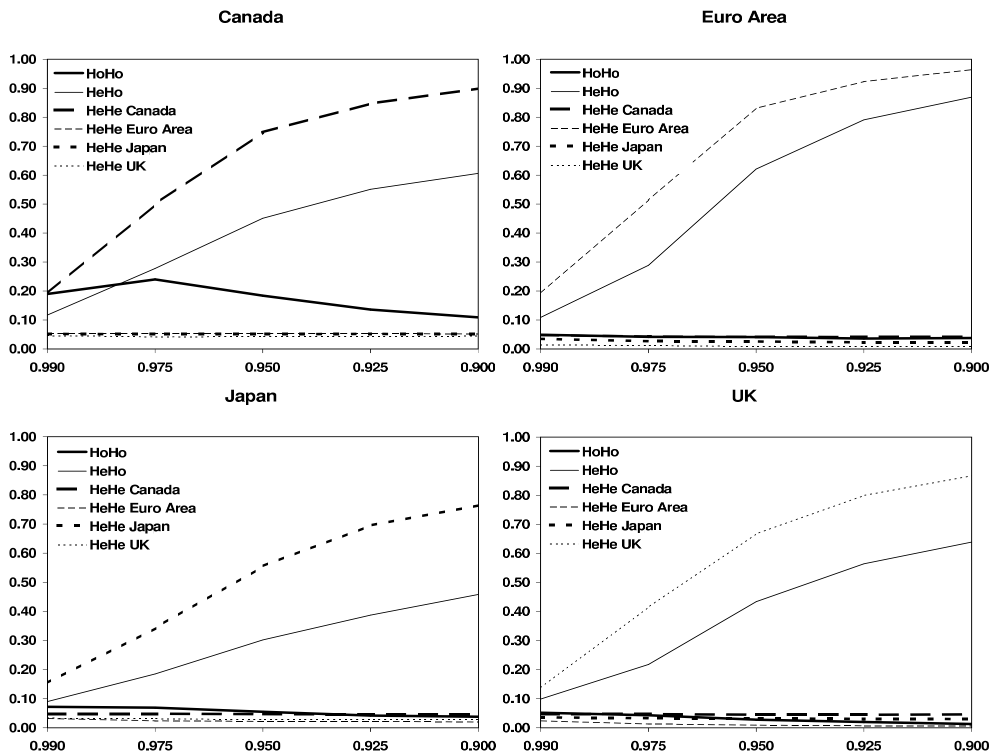


Figure 3.2: Power Functions Multivariate Unit Root Tests: Heterogeneous Mean Reversion Coefficients (3 Unit Roots). This figure presents the estimated power functions for the multivariate models in which (i) the mean reversion parameters are estimated homogeneously and the unit root test is performed for the entire sample of real exchange rates (HoHo model), (ii) the mean reversion parameters are estimated heterogeneously, but the unit root test is performed for the entire sample of real exchange rates (HeHo model), and (iii) the mean reversion parameters are estimated heterogeneously and the unit root test is performed for each individual real exchange rate separately (HeHe model). The estimated power functions presented in this figure are based on Monte Carlo simulations (see Section 3.2 for a detailed description) for the real exchange rates of Canada, the Euro area, Japan, and the U.K. versus the U.S. for the period 1981:01-2003:12. The power functions are estimated under the restriction that three out of four mean reversion parameters are equal to one, while the remaining mean reversion parameter (mentioned in the chart titles) decreases from 0.99 to 0.90.

of the euro area, while it is only 0.56 for the exchange rate of Japan. Hence, the influence of the country or currency of study on the power of the test seems to be substantial. Note that although the power for the HeHe test is high relative to the other methodologies, even at low half-lives there is a considerable probability of failing to reject the true null-hypothesis.

The power of the HeHo methodology increases when β decreases, but is considerable lower than the power of the HeHe methodology. For example, when the real exchange rate of Canada is mean reverting (see top left graph), the HeHe methodology exhibits a power of 0.75 at a half-life of 14 months, while the HeHo methodology has a power of only 0.45. Although this is a natural implication of the fact that the HeHe test is an individual unit root test while the HeHo methodology evaluates the stationarity of the panel as a whole, it also underlines the role the chosen methodology can play in assessments of the PPP hypothesis.

Figure 3.3 and 3.4 show the power of the unit root tests for the three different methodologies when the number of exchange rates in the panel that are non-stationary is equal to two and one, respectively. The mean reversion parameters of the remaining stationary series decrease simultaneously from 0.99 to 0.9. We observe that although the power of the HoHo methodology improves slightly with fewer unit roots in the sample, the power function again does not monotonically increase when β is reduced from 0.99 to 0.9. For example, when the real exchange rates of Canada and Japan are stationary, the power of the panel unit root test is twice as high when the β's of Canada and Japan are equal to 0.975 compared to the case where both β's equal 0.9. The irregular properties of the HoHo methodology seem to become more pronounced when the number of unit roots in the panel is reduced.[11]

The power functions of the HeHe methodology reveal that the power does not crucially depend on the number of non-stationary series in the panel, but it does still vary considerably across the currencies. Again, the power of this test to reject the null is very close to the size of the test for the non-stationary series. The power of the HeHo methodology increases when the number of unit roots in the panel decreases, because the panel as a whole becomes more stationary. When comparing the power properties of the HeHe and the HeHo methodologies, we note that the power is similar for DGP's in which two out of the four exchange rates are stationary. When only

[11] Note that the degree of heterogeneity in the panels analyzed in Figure 3.3 (3.4) is actually relatively limited, as the mean reversion parameters are identical for two (three) out of the four series. The drawbacks of the HoHo methodology could be aggravated by introducing more heterogeneity in the panel and hence we bias the result against finding important differences between the alternative methodologies.

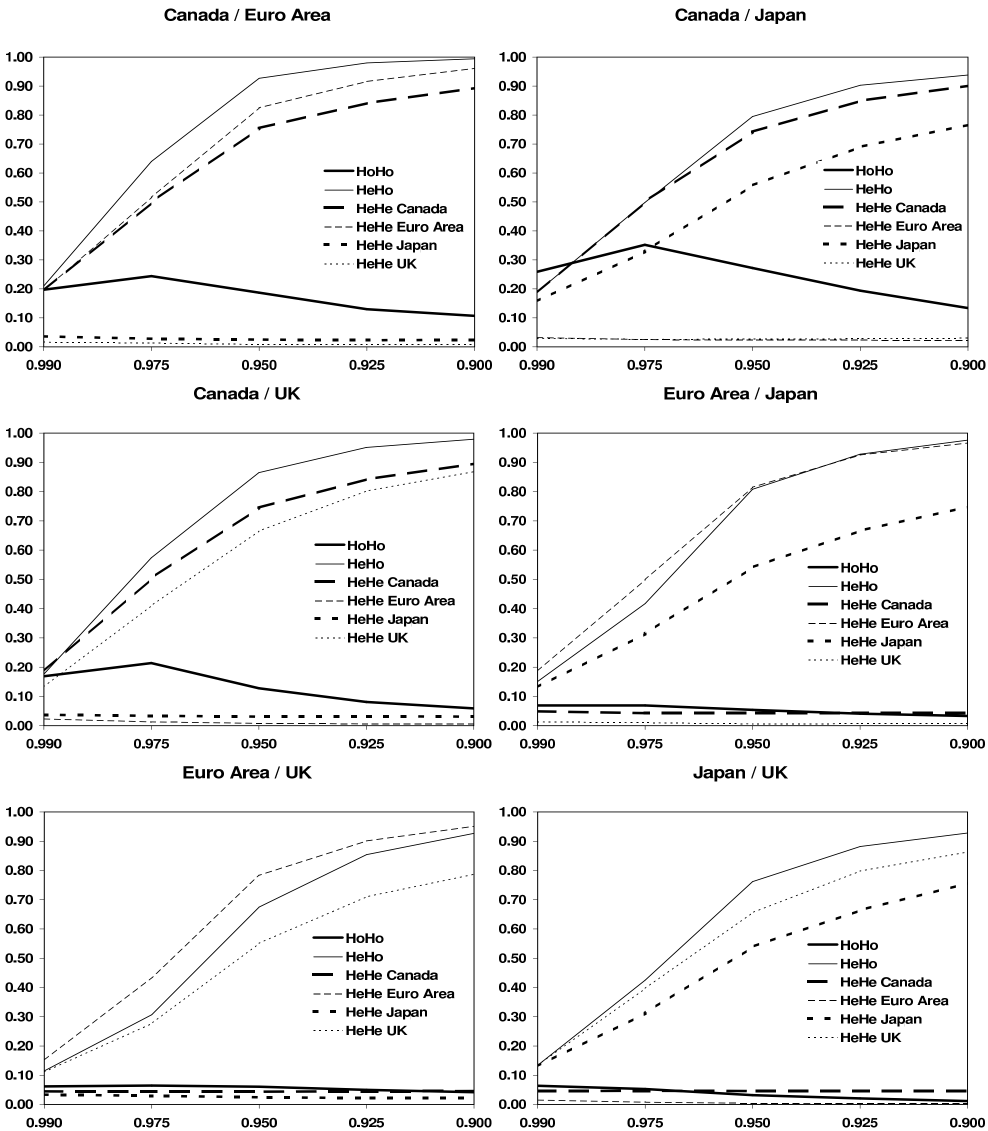


Figure 3.3: Power Functions Multivariate Unit Root Tests: Heterogeneous Mean Reversion Coefficients (2 Unit Roots). This figure presents the estimated power functions for the HoHo, HeHo, and HeHe models (described in Figure 3.2). The functions are estimated under the restriction that two out of four mean reversion parameters are equal to one, while the remaining 2 parameters (mentioned in the chart titles) decrease from 0.99 to 0.90.

one of the exchange rates in the panel contains a unit root, the power of the HeHo test is generally higher.

This chapter aims at studying the finite-sample properties of various alternative methodologies for evaluating the PPP hypothesis for a panel of real exchange rates. Our analysis of the power functions demonstrates that the HeHo methodology is generally most powerful when there are three or more stationary series in the panel. The HeHe methodology succeeds relatively well in identifying individual stationary series. The power properties of the HoHo methodology can be described as undesirable, as the power is low and often not monotonically increasing when the mean reversion parameter decreases. In order to gain more understanding of these adverse properties and thus of the consequences of imposing homogeneous mean reversion, we analyze the distribution of estimated β coefficients under the HoHo methodology.

Figures 3.5, 3.6, and 3.7 display the distribution of the (homogeneously) estimated mean reversion parameter β under the HoHo methodology when the panel contains, respectively, three, two, and one unit roots. The mean reversion parameter of the stationary exchange rates in the panel are assumed to decrease simultaneously from 0.99 to 0.5.[12,13] We added $\beta = 0.5$ (corresponding to a half-life of merely 1 month) to the set of $\{0.99, 0.975, 0.95, 0.925, 0.9\}$ in order to highlight the impact of the homogeneity restriction in extreme cases. The histograms also present the values of μ, defined as the average of the true mean reversion parameters, and b, the (homogeneously) estimated mean reversion parameter.

The results are striking. We observe that when the mean reversion parameter of the stationary series in the panel decreases, the distribution of the estimated β of the total panel first gradually moves to the left. This can be expected, as the panel as a whole becomes more stationary. However, when the mean reversion parameter reaches a value of 0.95 or less, there is a clear movement to the right, implying higher estimates of the homogeneous β. As the value of b slowly increases, while the mean of the true mean reversion parameters (μ) steadily drops, this implies an increasingly large bias in the homogeneous parameter estimate. Second, marked changes to the shape of the distribution occur. Moving from the top graphs to the bottom, the dispersion and skewness of the distribution initially increase notably,

[12] The distributions are based on the same Monte Carlo simulations as for the analysis of the statistical power of the tests, but 10,000 instead of 1,000 simulations are used to obtain a smoother representation of the distributions.

[13] Note that for a given number of (non-)stationary exchange rates in the panel, there are several possibilities to assign the unit root(s) to the series. For example, in Figure 3.5 we choose the Canadian dollar - U.S. dollar series to be stationary, but we could have chosen any of the three other series. The results for combinations other than displayed in Figures 3.5 to 3.7 are similar and available from the authors.

signifying increasing uncertainty about the parameter estimate. However, when the stationary series in the panel become even more mean reverting, the distribution again becomes more dense and less skewed. This "boomerang" effect is more severe when the panel contains fewer series with unit roots. Hence, the more the panel as a whole moves further away from the unit root null-hypothesis, the more the distribution of the estimated β starts to resemble the one under the null-hypothesis again. Remarkably, the homogenous estimation methodology seems to "ignore" mean reversion parameters that deviate considerably from the non-stationary null and the unit roots in the panel dominate the stationary series in the estimation.

The extraordinary character of this feature of the HoHo methodology is best illustrated by the bottom panel of Figure 3.7. This panel concerns a DGP in which only the euro - dollar rate contains a unit root, and all other three real exchange rates have half-lives of 1 month ($\beta = 0.5$). The mean of the estimated mean reversion parameters (denoted by b) under the HoHo methodology is 0.993, implying a bias of no less than 0.368 relative to the average of the true β's in the panel under the null-hypothesis ($\mu = 0.625$). These findings reinforce the conclusion that imposing a common mean reversion coefficient in heterogeneous panels can lead to testing methodologies with seriously adverse properties, including low and non-monotonic statistical power and severe biases in the parameter estimates.

3.4.2 Monte Carlo evidence on panels with only stationary series

In this section we investigate the power functions of the HoHo, HeHo, and HeHe methodologies for DGP's in which all real exchange rates are stationary. As a benchmark, we first investigate the situation in which all four exchange rates exhibit the same mean-reverting behavior. That is, the mean reversion parameters β_i ($\forall i$) are equal to the same value $\beta \in \{0.99, 0.975, 0.95, 0.925, 0.9\}$. The results are displayed in Figure 3.8. Not surprisingly, the HoHo methodology performs very well. Its power is the highest of all methodologies and it monotonically increases with a decrease in the mean reversion parameter. This can be explained by the fact that the homogeneous model is not misspecified for this DGP. The HeHo methodology is only slightly less powerful than the HoHo methodology, however. For β's of 0.95 and lower (corresponding with half-lives of 14 months and less), both testing methodologies are able to reject the unit root in virtually all cases. When the half life is around 2.5 years ($\beta = 0.975$), the probability that non-stationarity is rejected still amounts to 92% (HoHo) and 79% (HeHo). The power functions of the HeHe methodology behave similarly to those generated under DGP's with both stationary and non-stationary series. Homogenous estimation and testing thus leads to a power advantage when the real exchange rates indeed have the same mean-reverting behavior.

As panel data sets in which the series have exactly the same mean reversion parameters are unlikely to be encountered in practice, more interesting DGP's involve real exchange rates with different rates of mean reversion.[14] In Figure 3.9 we examine the power properties of the multivariate methodologies by setting the β_i's equal to the estimated values of the heterogeneous SUR model (see Table 3.3) and decrease the mean reversion parameter of one of the series at the time. As the DGP's in each of the four panels of Figure 3.9 are generated under the alternative hypothesis, it is not surprising that the methodologies that assess the stationarity of the panel as a whole have the highest power to reject the null. The HeHo methodology generally achieves the highest statistical power. The HoHo methodology is also relatively powerful, but exhibits the same properties as discussed above, notably for the situation in which the mean reversion of the pound - dollar rate is varied (bottom right panel). The power functions of the HeHe methodology have regular shapes: increasing for the exchange rate of which the mean reversion parameter decreases and essentially flat for the other series.

Figure 3.10 displays the power functions of the three methodologies for DGP's with a decreasing value of β_i for one of the exchange rates, while we assign the values $\{0.99, 0.975, 0.95\}$ to the remaining mean reversion parameters (these three values are assigned to the remaining real exchange rate series in such a way as to maximize the resemblance with the estimated values in Table 3.3). The plots demonstrate how sensitive the behavior of the HoHo test is to heterogeneity in the panel. The heterogeneity in the DGP's of Figure 3.10 is only moderately larger than the heterogeneity in the DGP's of Figure 3.9, but the power of the HoHo methodology deteriorates considerably and the non-monotonicity effect becomes stronger. The power functions of the HeHo and HeHe methodologies are basically unaffected.

We take a closer look at the empirical performance of the HoHo model by plotting the distribution of the estimated β coefficients for the DGP's analyzed in Figure 3.10. To preserve space, Figure 3.11 only shows the distributions for the case in which the mean reversion of the Canadian dollar - U.S. dollar rate is varied. The distributions for the other exchange rates display a similar pattern. The "boomerang" effect documented in the previous section is evidently not contingent on the presence of unit roots in the panel. Again, the distribution of the estimated common mean reversion

[14]Even if we restrict ourselves to parameters from the set $\{0.99, 0.975, 0.95, 0.925, 0.9\}$, the number of DGP's we can generate is very large. In Figures 9 and 10 we therefore present the power functions of the HoHo, HeHo, and HeHe methodologies for an interesting subset of all possible DGP's. More results can be obtained from the authors.

coefficient at first shifts to the left and becomes more dispersed. For lower half-lives of the Canadian dollar - U.S. dollar real exchange rate, however, the distribution moves back to the right.[15] Even without a unit root in the panel, the least stationary series appear to dominate stationary series with a root that deviates from unity. Consequently, the HoHo methodology can also lead to serious biases in the estimated mean reversion parameters in panels with only stationary real exchange rates.

In order to illustrate the differences between the finite-sample properties of the HoHo and HeHe methodologies, we also plot the distribution of the estimated β's under the HeHe methodology. Figure 3.12 presents histograms for the heterogeneously estimated β's for the same DGP as in the top-left panel of Figure 3.10 and as in Figure 3.11. Again, μ denotes the average of the true mean reversion parameters, and b represents the heterogeneously estimated mean reversion parameter for the specific series under examination. Note that for clarity we use different x-axes for the β distributions for Canada. As the half-life of the exchange rate of Canada versus the U.S. increases, the distribution of the estimated β for Canada shifts to the left, while the distributions for the other countries virtually stay the same. For Canada, the estimate b moves in lockstep with the decreasing true mean reversion parameter β. For the euro area, Japan, and the U.K., the true mean reversion parameters are equal to $\{0.975, 0.99, 0.95\}$ respectively, while - independent of the decrease in Canada's β - the estimated mean reversion parameter b equals $\{0.965, 0.98, 0.934\}$. This indicates that the HeHe methodology leads to a relatively accurate estimate of individual mean reversion parameters in heterogeneous panels, even in small samples.

3.4.3 Evaluation of the three methodologies

The evidence presented in sections 3.4.1 and 3.4.2 suggests that restricting the mean reversion parameter to be homogeneous across different real exchange rates can have detrimental consequences for the validity of the outcomes of multivariate unit root tests. Irregularly shaped power functions and potentially large biases in the parameter estimates arise in panels with and without non-stationary series and even when the degree of heterogeneity in the panel is relatively limited. These findings constitute compelling reasons for using a methodology that allows for heterogeneous *estimation* of the mean reversion parameters.

The issue whether researchers interested in the PPP hypothesis should also perform unit root *tests* heterogeneously depends on the purpose of the study. If the hypothesis of interest is primarily whether one or more of the real exchange rates in the panel are stationary, the researcher's main objective is probably to achieve max-

[15] Note that the accompanying decrease in dispersion is less pronounced than in Figures 3.5-3.7.

imum power. The relative power of the HeHo and HeHe methodologies is dependent on the number of unit roots present in the panel. As this number is unknown, it may be valuable to use both methodologies.

An important drawback of homogeneous tests of PPP is that rejecting the unit root hypothesis does not provide any guidance as to how many real exchange rates are stationary, let alone which.[16] Furthermore, from many perspectives the relevant question is not whether a panel of real exchange rates as a whole can be considered to be stationary, but whether PPP holds between individual countries.[17] Research that is directed at this question should use a methodology that evaluates the PPP hypothesis for individual currency pairs in the sample. An important additional advantage of such a testing methodology is that the power is independent of the number of non-stationary series in the panel.

Alternative approaches to evaluate the PPP hypothesis for individual exchange rates include studying long-run time series (see e.g. Edison (1987) and Lothian and Taylor (1996)) and Bayesian analysis of unit roots in real exchange rates (see e.g. Koop (1992)). Bayesian methods treat the unit root and stationarity hypotheses symmetrically and allow for an assessment of the probability of a unit root in the data by evaluating the Bayesian posterior odds ratio.

3.5 Conclusions

Froot and Rogoff (1995) show that a researcher interested in testing the PPP hypothesis needs 72 years of stationary (monthly) data to reject the null-hypothesis of a random walk when the real exchange rate has a half-life of three years. Since the early 1990s, many researchers have turned to multivariate unit root tests in order to overcome this power problem. However, inferences drawn from multivariate testing methodologies are adversely affected by a number of restrictions imposed on the multivariate model. Two of these restrictions have been defied by O'Connell (1998), who emphasizes the importance of allowing for cross-sectional dependence, and by Papell and Theodoridis (2001) and Wu and Wu (2001), who show that it is vital to take heterogeneous serial correlation into account. However, a third critical restriction in multivariate tests for PPP is still widely used: numerous studies impose a homogeneous rate of mean reversion across all real exchange rates in the panel.

[16] This pitfall is also discussed by Taylor and Sarno (1998). They suggest a test statistic that only rejects the null hypothesis if all real exchange rates are stationary, but this test neither reveals how many nor which series are stationary when the null is not rejected.

[17] This is essential for many practically oriented studies of PPP. As an example, asset managers may be interested in the validity of PPP in relation to hedging specific currency risks.

The economic rationale for this restriction is arguable, as the validity of PPP between two countries should depend on various economic, institutional, and possibly even geographic characteristics specific to those countries. Econometrically, imposing this restriction may affect the test outcomes in important ways that have not been comprehensively investigated in the literature.

This chapter studies the finite-sample properties of three different multivariate estimation and testing methodologies. We examine the consequences of the homogeneity restriction on tests for PPP by means of Monte Carlo simulation. The first methodology we investigate involves homogeneous estimation of the mean reversion parameters and performing a unit root test on the panel as a whole. In the second methodology, estimation is performed heterogeneously but testing is done homogeneously. In the third methodology, the mean reversion parameters are estimated heterogeneously and the null-hypothesis of non-stationarity is tested for each individual exchange rate in the sample individually.

Our findings uncover important adverse properties of the methodology with homogeneous estimation and testing in the presence of heterogeneity in the data generating process. The statistical power of this testing methodology is relatively low and, remarkably, does not increase when the mean reversion parameter is decreased. In addition, we document significant biases in the estimated mean reversion parameters. In particular, when one or more of the mean reversion parameters are decreased while the remaining parameters stay the same, the homogenous mean reversion parameters estimate increasingly underestimates the average degree of stationarity in the panel.

These properties are observed for both panels with and panels without nonstationary real exchange rates and arise even when the heterogeneity across the real exchange rates in the panel is limited. The homogeneity restriction only leads to higher power when all exchange rates are stationary and mean reversion parameters are very similar. The power functions of the other two methodologies behave in a regular way. Heterogeneous estimation and homogeneous testing generally leads to substantial power, while the third, fully heterogeneous, methodology performs well in detecting stationarity for individual series, especially when the number of unit roots in the panel is high. These findings highlight the importance of taking heterogeneous mean reversion into account in multivariate tests of PPP.

Our study is related to recent work by Imbs *et al.* (2005). They demonstrate that heterogeneity in the dynamics of sectoral price indices may induce significant biases in estimates of mean reversion parameters based on aggregated price indices. Our analysis shows that when aggregate real exchange rates are used for testing the PPP hypothesis, it is important to recognize that mean reversion properties may differ across countries.

Table 3.3: SUR ADF Unit Root Tests: Heterogeneous Mean Reversion Coefficients (HeHo and HeHe methodology). This table presents the results of the SUR ADF unit root test of the real exchange rates of Canada, the Euro area, Japan, and the U.K. versus the U.S. over the period 1981:01-2003:12 (276 time-series observations). We estimate the following system: $R_{i,t} = \alpha_i + \beta_i R_{i,t-1} + \sum_{k=1}^{l_i} \gamma_{i,k} \Delta R_{i,t-k} + u_{i,t}$, where $_{Ri,t}$ is the log of the real exchange rate and the value of l_i is taken from the OLS ADF unit root test results. The critical values [10% cv] have been obtained using Monte Carlo Simulations. * and ** denote significance of the test statistic $\tau_i = (\beta_i - 1)/s(\beta_i)$ at the 10% and 5% level, respectively.

Country	α_i	s.e.	β_i	s.e.	half-life	l_i	τ_i [10% cv]	Wald [10% cv]
Canada	0.0003	0.0009	0.9833	0.0078	41	16	-2.13	21.69*
							[-2.69]	[20.25]
Euro Area	0.0025	0.0020	0.9636	0.0098	19	13	-3.72**	
							[-3.15]	
Japan	-0.0026	0.0023	0.9770	0.0096	30	22	-2.39	
							[-2.80]	
U.K.	0.0100	0.0038	0.9607	0.0135	17	21	-2.90	
							[-3.10]	

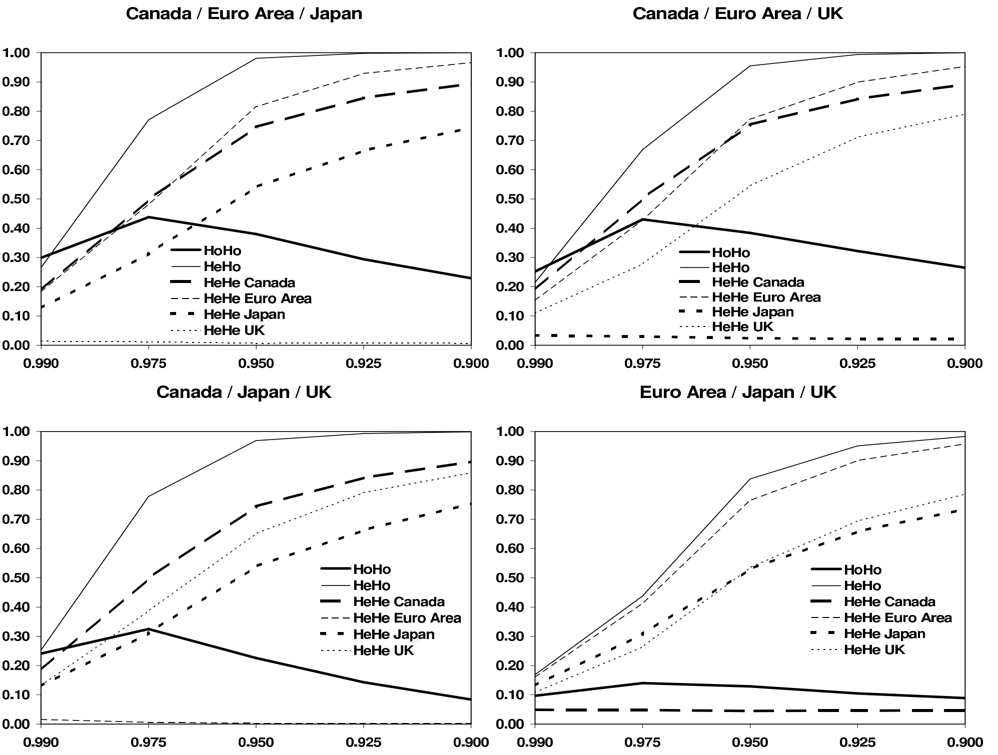


Figure 3.4: Power Functions Multivariate Unit Root Tests: Heterogeneous Mean Reversion Coefficients (1 Unit Root). This figure presents the estimated power functions for the HoHo, HeHo, and HeHe models (described in Figure 3.2). The functions are estimated under the restriction that one out of four mean reversion parameters is equal to one, while the remaining 3 parameters (mentioned in the chart titles) decrease from 0.99 to 0.90.

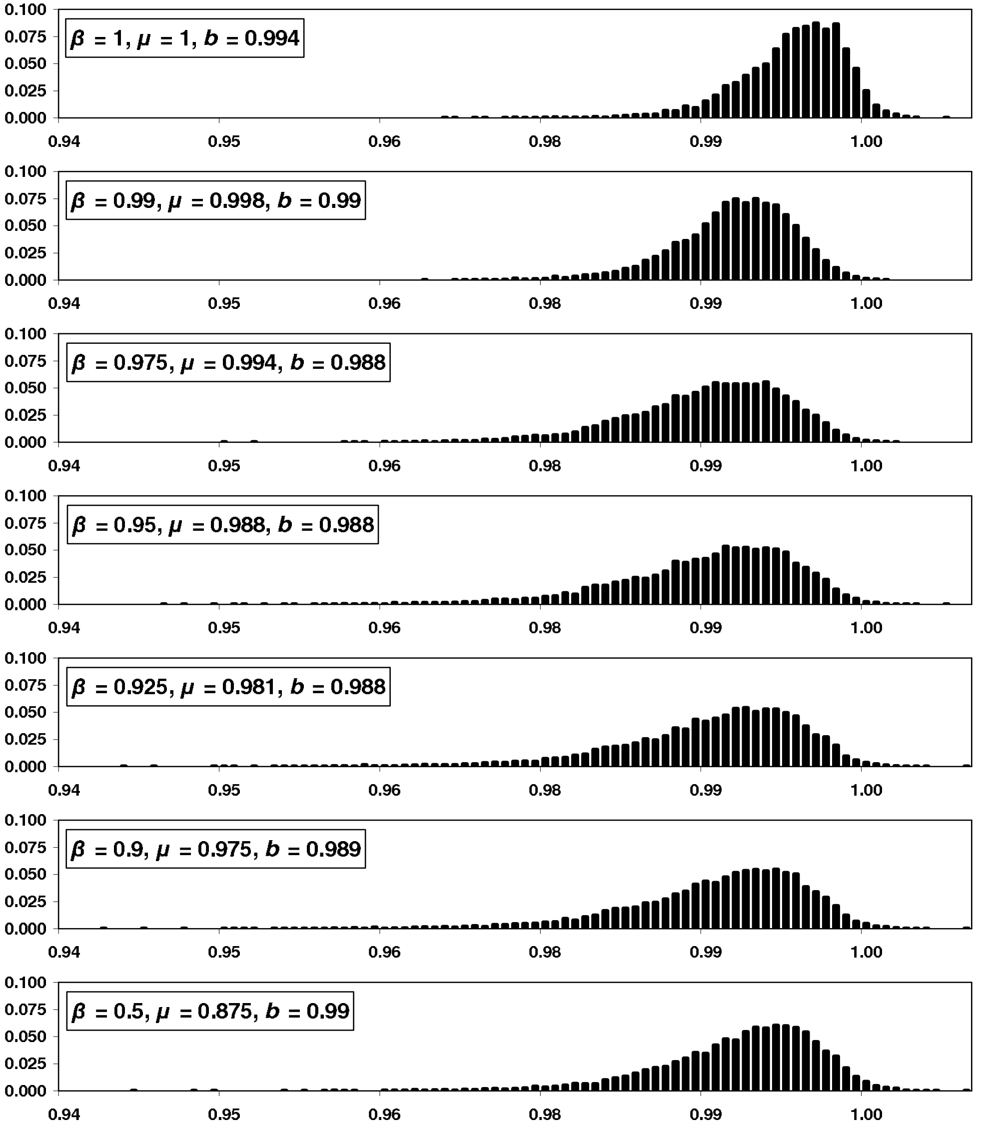


Figure 3.5: Histograms of Common Mean Reversion Coefficient Estimates (3 Unit Roots). This figure presents histograms of multivariate estimates of a common mean reversion coefficient when the true parameter values for the set of countries {Canada, Euro area, Japan, U.K.} equal $\{\beta, 1, 1, 1\}$, where β decreases from 1 to 0.5.

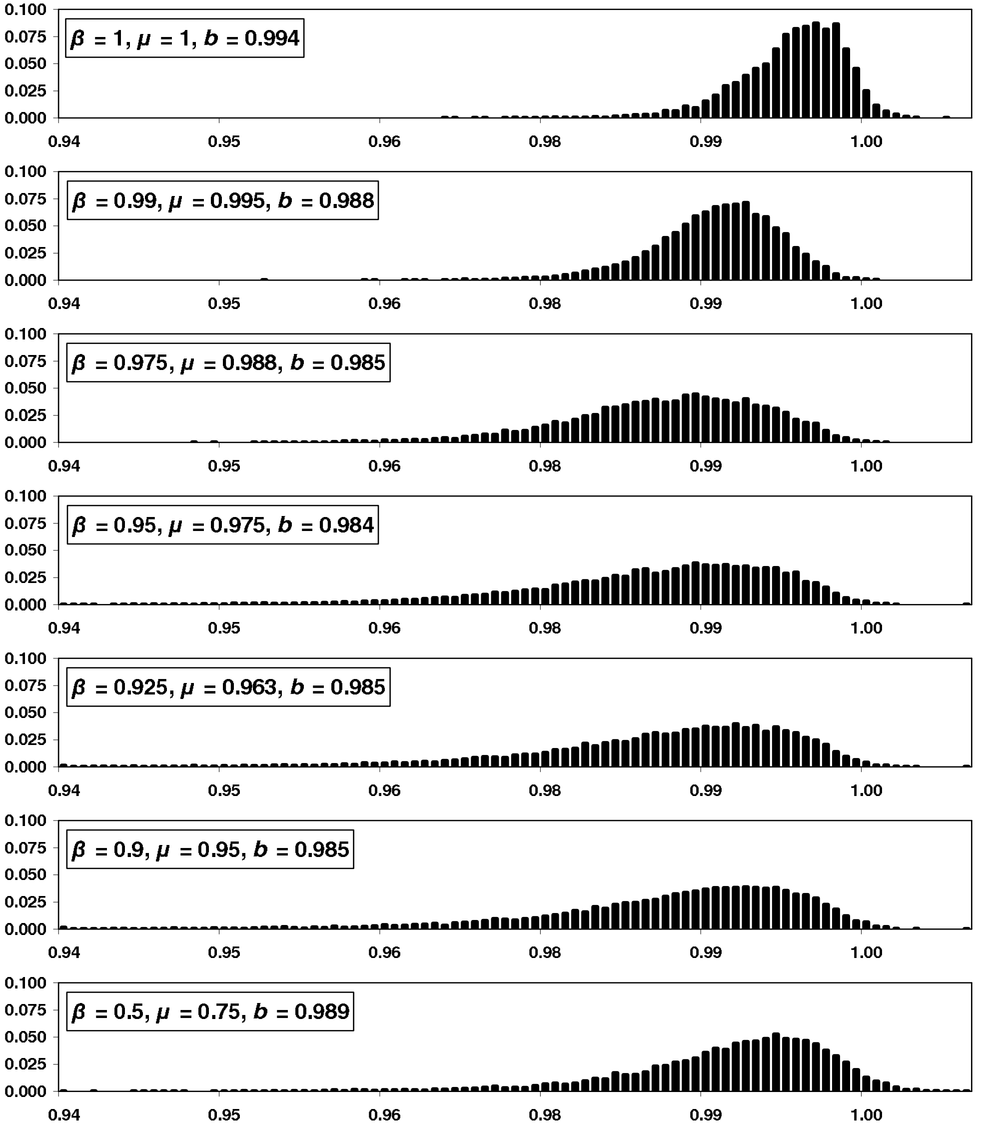


Figure 3.6: Histograms of Common Mean Reversion Coefficient Estimates (2 Unit Roots). This figure presents histograms of multivariate estimates of a common mean reversion coefficient when the true parameter values for the set of countries {Canada, Euro area, Japan, U.K.} equal {β, 1, β, 1}, where β decreases from 1 to 0.5.

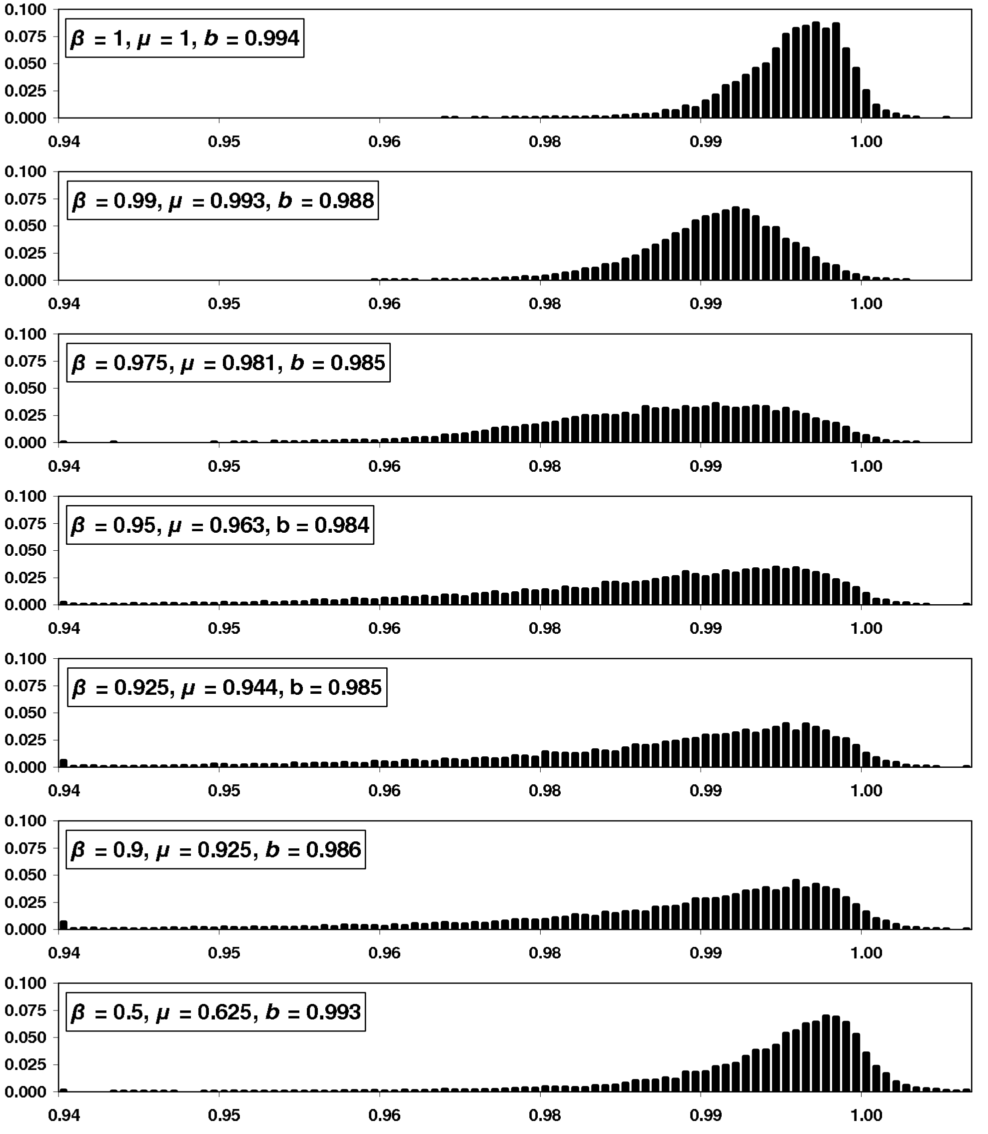


Figure 3.7: Histograms of Common Mean Reversion Coefficient Estimates (1 Unit Root). This figure presents histograms of multivariate estimates of a common mean reversion coefficient when the true parameter values for the set of countries {Canada, Euro area, Japan, U.K.} equal $\{\beta, 1, \beta, \beta\}$, where β decreases from 1 to 0.5.

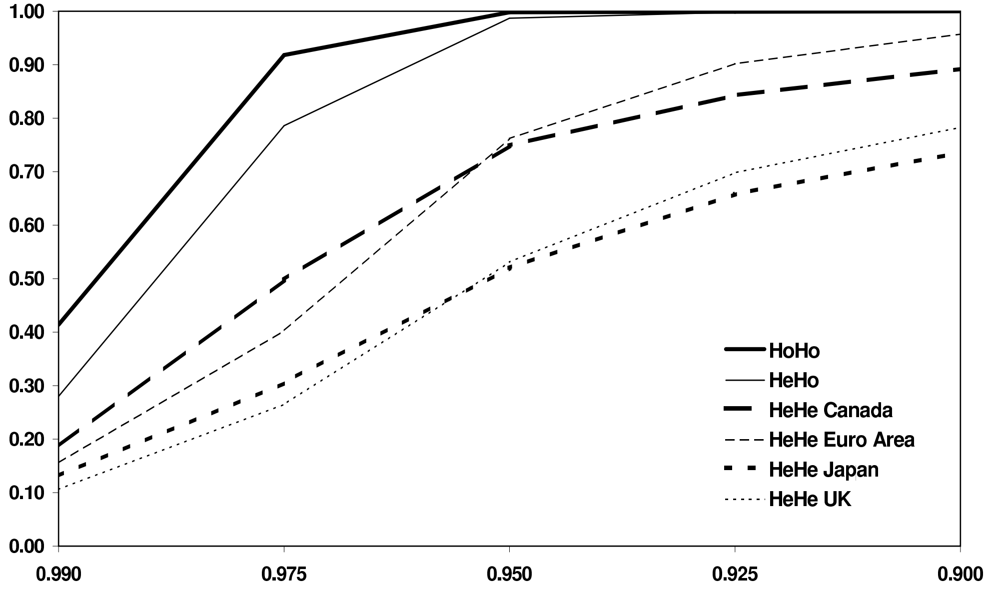


Figure 3.8: Power Functions Multivariate Unit Root Tests: Homogeneous Mean Reversion Coefficients. This figure presents the estimated power functions for the multivariate models in which (i) the mean reversion parameters are estimated homogeneously and the unit root test is performed for the entire panel of real exchange rates (HoHo model), (ii) the mean reversion parameters are estimated heterogeneously, but the unit root test is performed for the entire panel of real exchange rates (HeHo model), and (iii) the mean reversion parameters are estimated heterogeneously and the unit root test is performed for each individual real exchange rate separately (HeHe model). The estimated power functions presented in this figure are based on Monte Carlo simulations (see Section 3.2 for a detailed description) for the real exchange rates of Canada, the Euro area, Japan, and the U.K. versus the U.S. for the period 1981:01-2003:12. The power functions are estimated under the restriction that all mean reversion parameters in the sample of four real exchange rates are equal and decrease from 0.99 to 0.9.

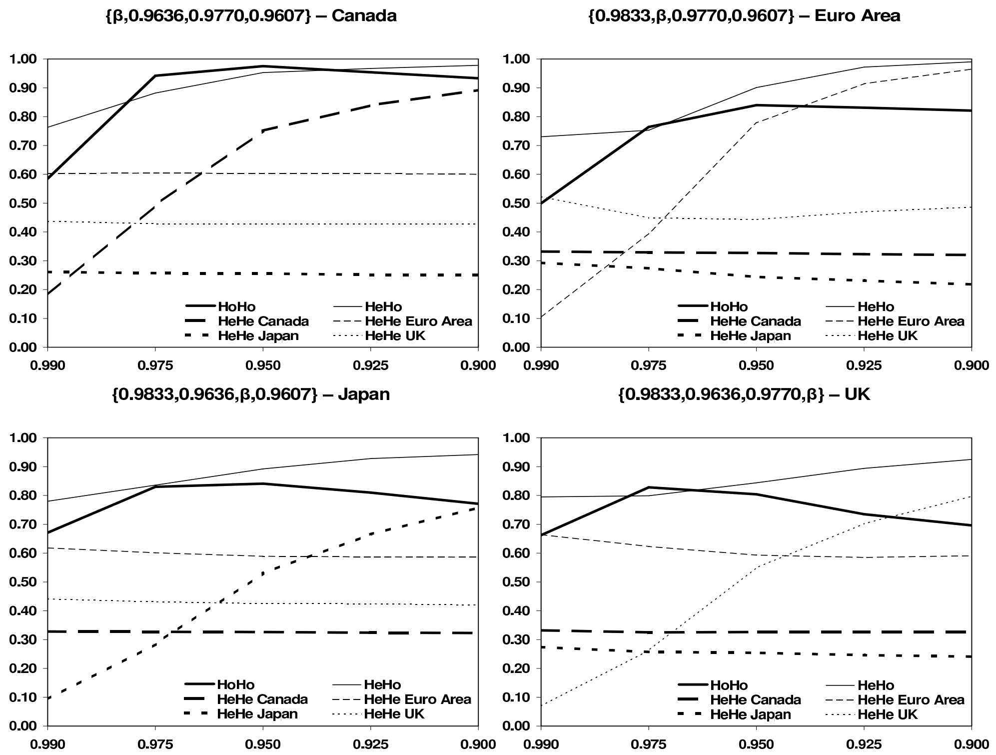


Figure 3.9: Power Functions Multivariate Unit Root Tests: Heterogeneous Mean Reversion Coefficients (No Unit Roots). This figure presents the estimated power functions for the HoHo, HeHo, and HeHe models (described in Figure 3.8). The functions are estimated under the restriction that three out of four mean reversion parameters are fixed at a level smaller than one (chosen from the set {0.990, 0.975, 0.950, 0.925, 0.900}), while the remaining parameter (mentioned in the chart titles) decreases from 0.99 to 0.90.

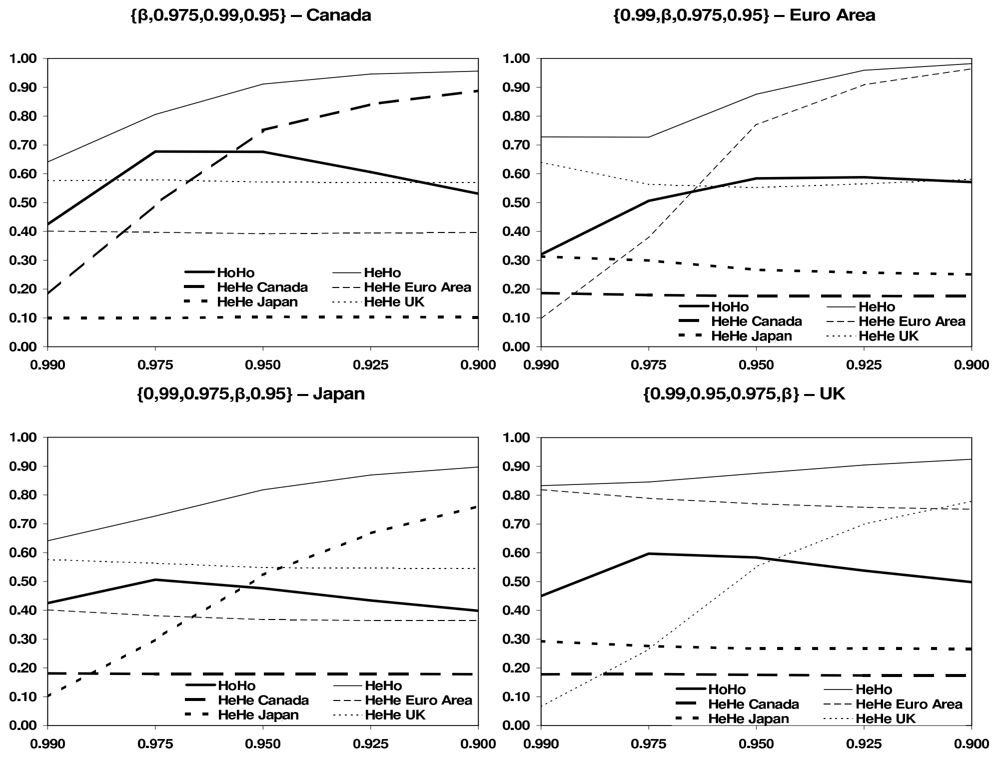


Figure 3.10: Power Functions Multivariate Unit Root Tests: Heterogeneous Mean Reversion Coefficients (No Unit Roots). This figure presents the estimated power functions for the HoHo, HeHo, and HeHe models (described in Figure 3.2). The functions are estimated under the restriction that three out of four mean reversion parameters are fixed at a level smaller than one (chosen from the set {0.990, 0.975, 0.950}), while the remaining parameter (mentioned in the chart titles) decreases from 0.99 to 0.90.

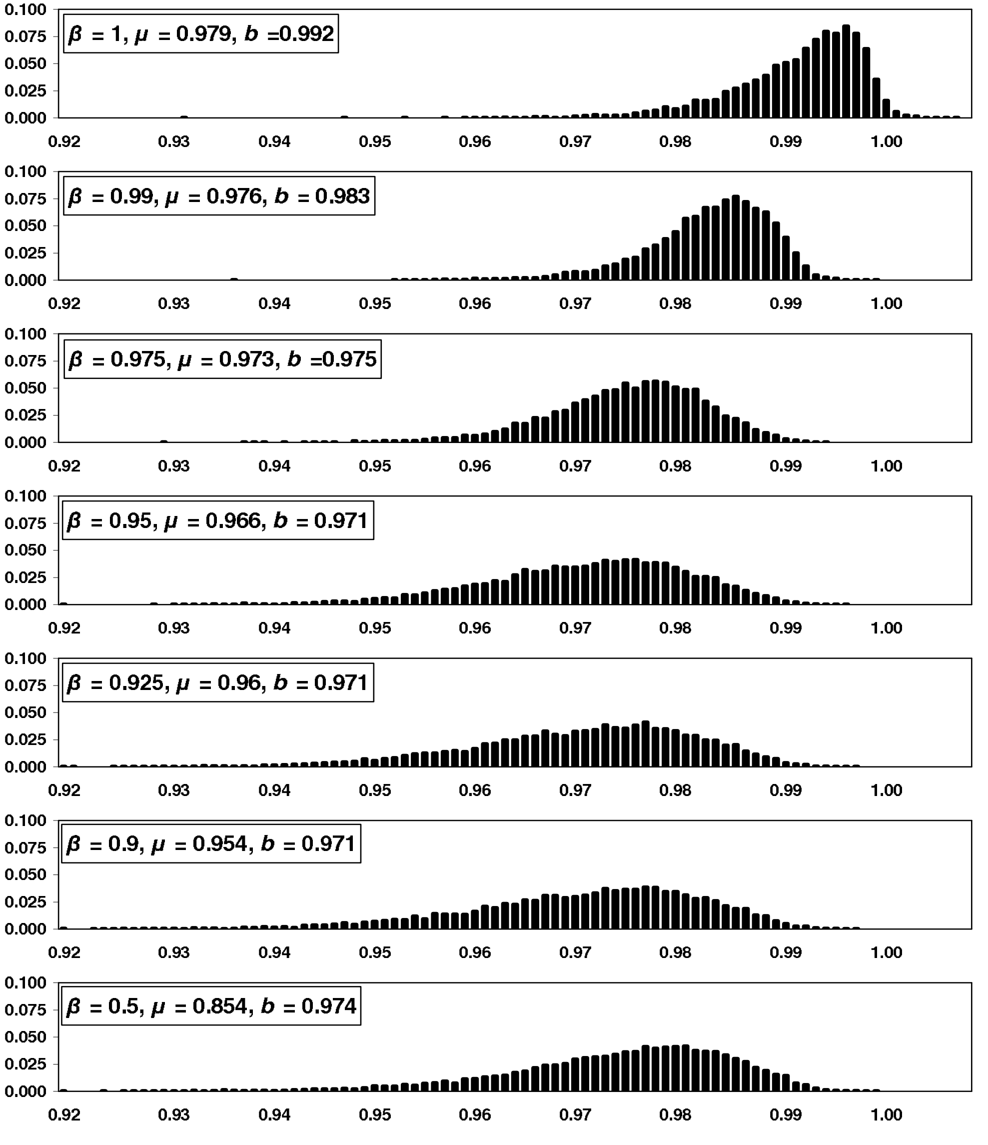


Figure 3.11: Histograms of Common Mean Reversion Coefficient Estimates (No Unit Roots). This figure presents histograms of multivariate estimates of a common mean reversion coefficient when the true parameter values for the set of countries {Canada, Euro area, Japan, U.K.} equal {β, 0.975, 0.99, 0.95}, where β decreases from 1 to 0.5.

Canada

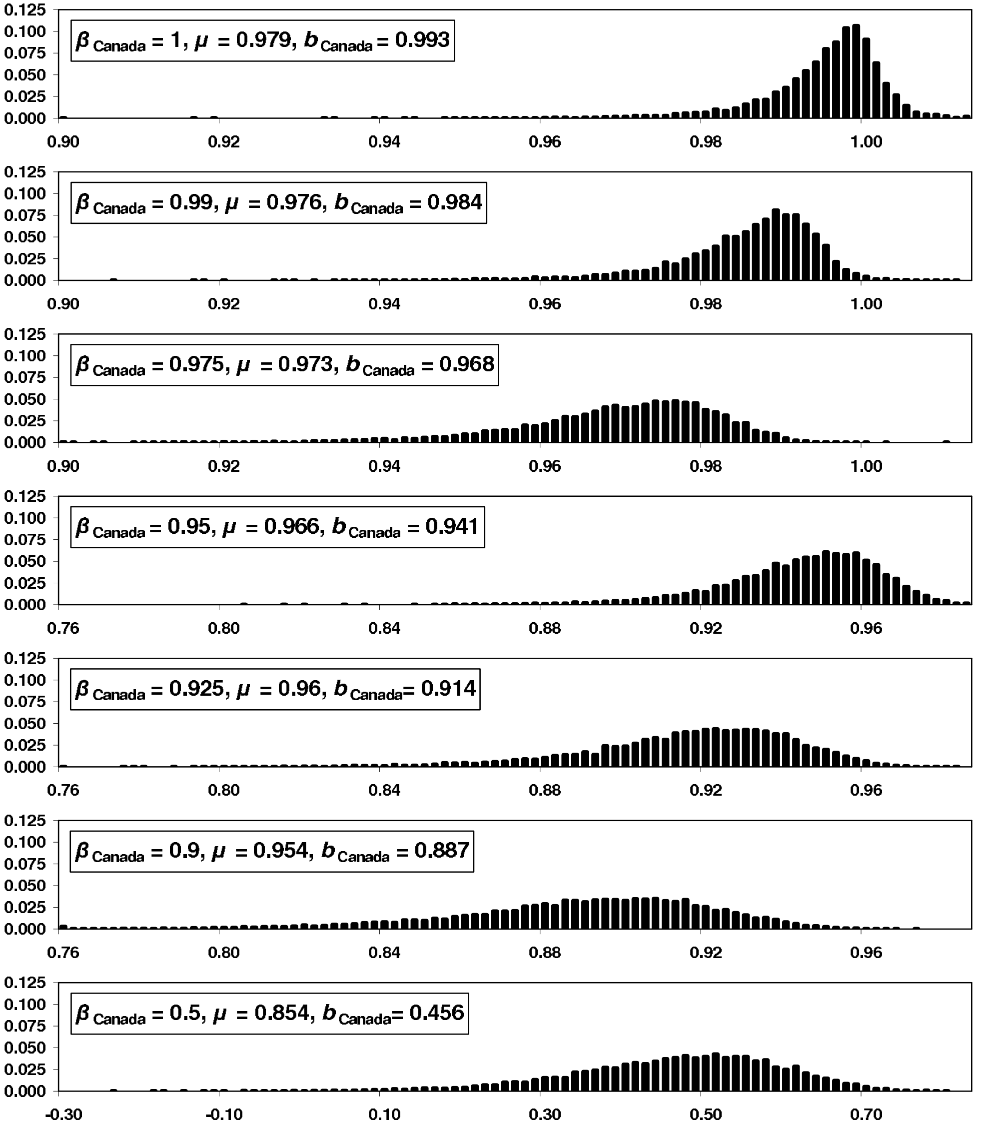


Figure 3.12: Histograms of Heterogeneous Mean Reversion Coefficient Estimates (No Unit Roots). This figure presents histograms of multivariate estimates of heterogeneous mean reversion coefficients when the true parameter values for the set of countries {Canada, Euro area, Japan, U.K.} equal {β, 0.975, 0.99, 0.95}, where β decreases from 1 to 0.5.

Euro area

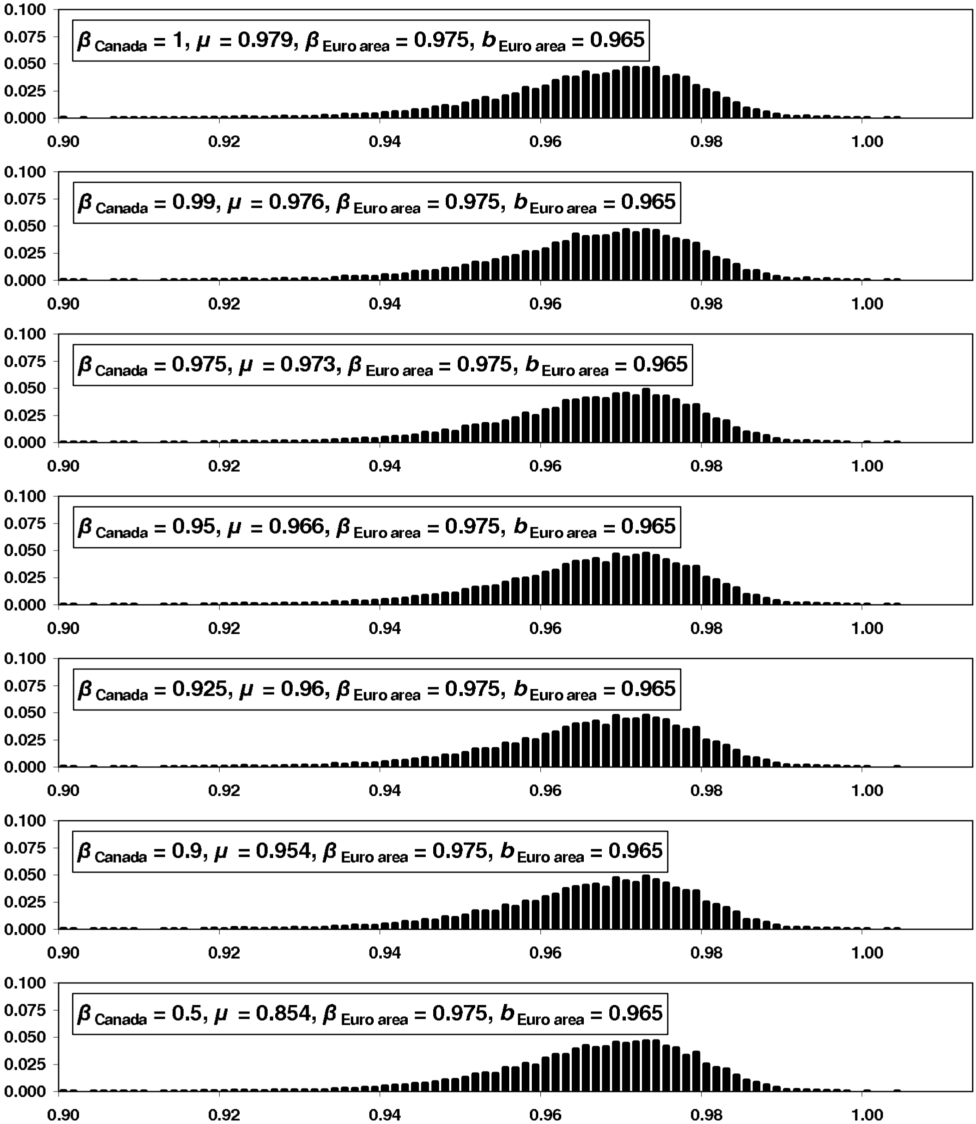


Figure 3.12: Histograms of Heterogeneous Mean Reversion Coefficient Estimates (No Unit Roots) - *Continued.*

Japan

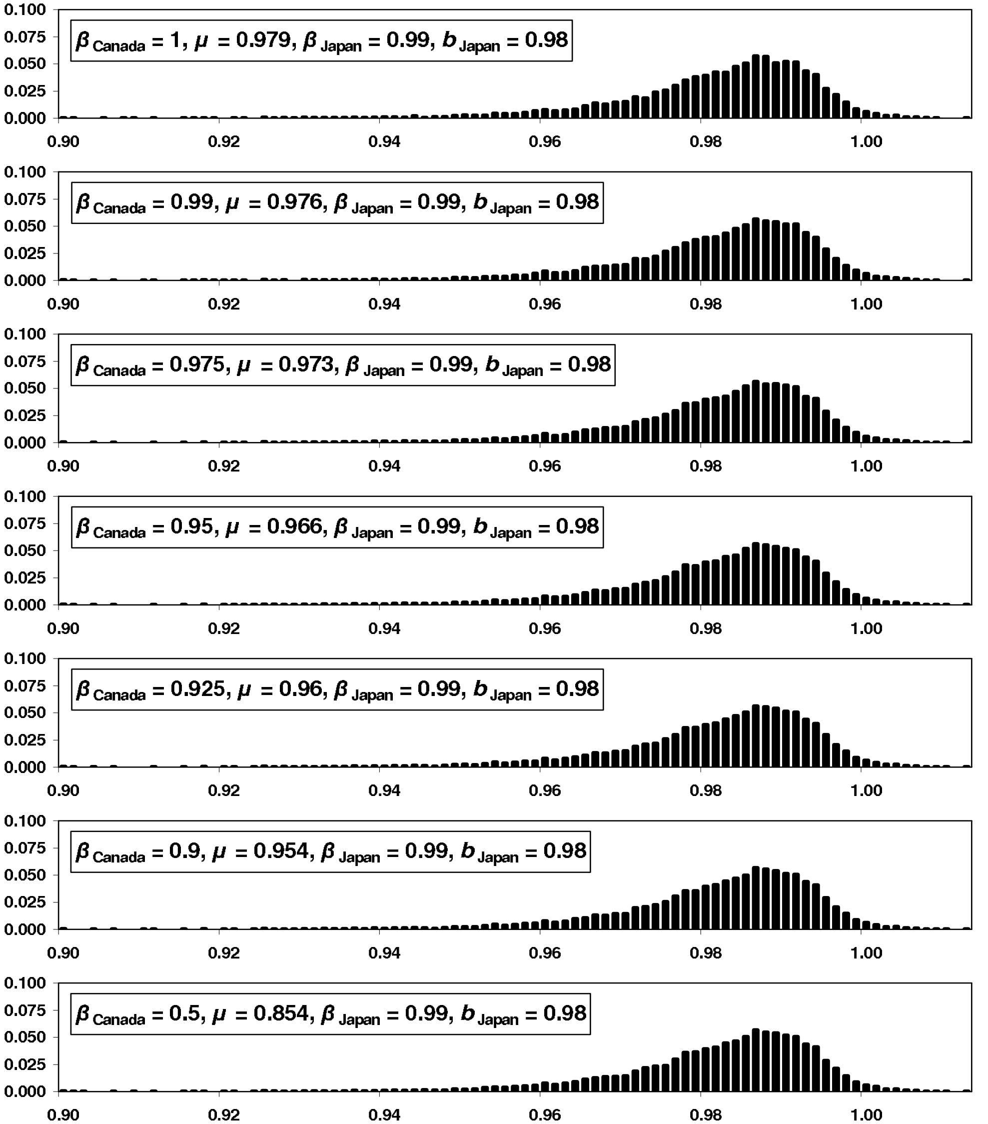


Figure 3.12: Histograms of Heterogeneous Mean Reversion Coefficient Estimates (No Unit Roots) - *Continued.*

U.K.

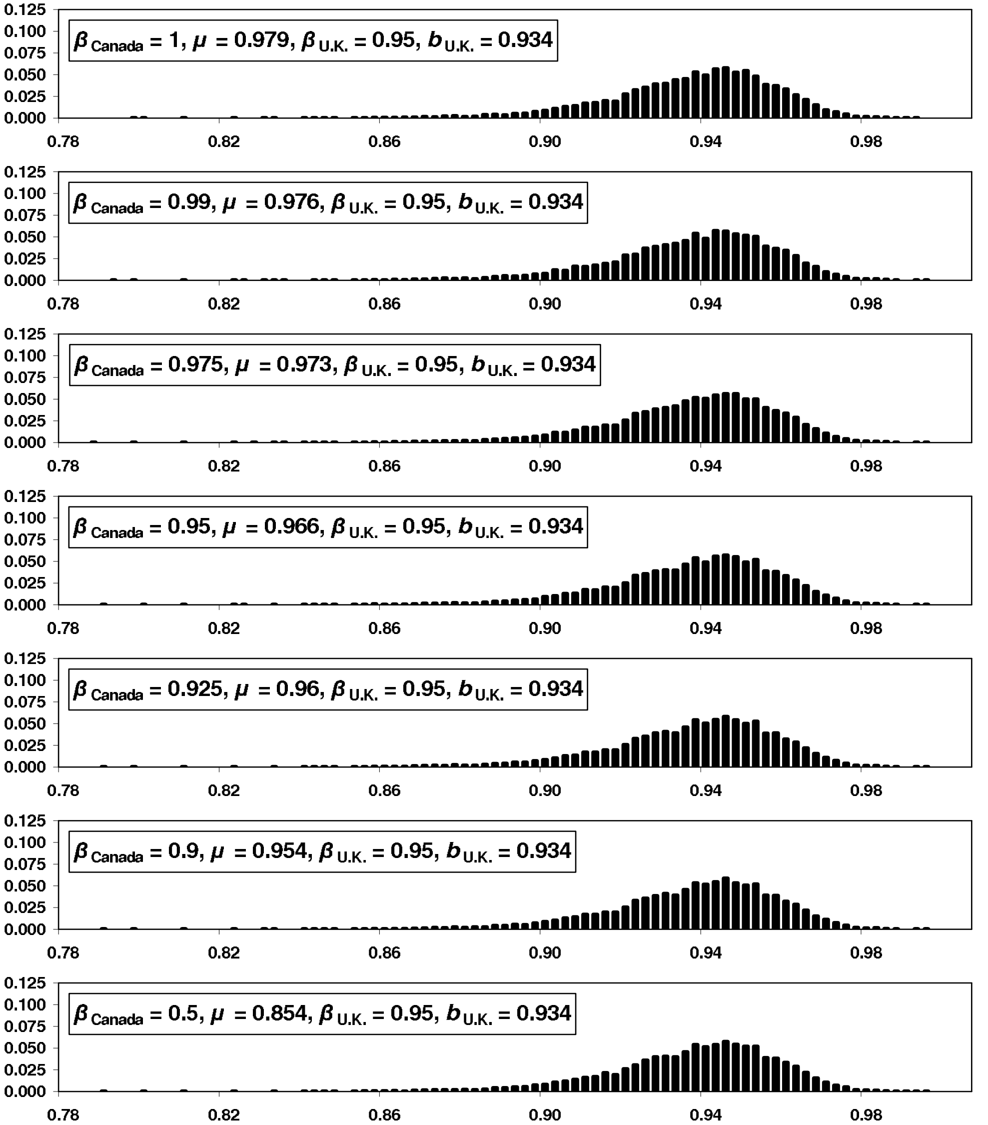


Figure 3.12: Histograms of Heterogeneous Mean Reversion Coefficient Estimates (No Unit Roots) - *Continued.*

Part II

Uncovered Interest rate Parity

Chapter 4

Exchange Rate Predictability Using Linear Vector Error Correction Models

4.1 Introduction

In the last decade more and more investors with foreign exposures have become aware that exchange rate movements can have a significant impact on local currency investment returns. This has lead to a whole new industry that has focused explicitly on managing foreign currency risks (See for example Record (2004) for an overview). An important prerequisite for this industry is that returns and risks on currency risk management strategies need to be understood. In this respect exchange rate models play an important role to measure these risks and returns. Particularly, in the international finance literature the question whether or not exchange rates are predictable still has an important position on the research agenda. Based on the seminal studies by Meese and Rogoff (1983b) and Meese and Rogoff (1983a) many authors have tried to beat the random walk benchmark and many made a futile attempt. However, according to some fundamental economic theories exchange rates should be related to other economic variables thereby implying that these latter variables should be informative in describing and/or predicting exchange rates.

Two of the most popular frameworks to the present day for describing and predicting exchange rates are the Purchasing Power Parity (PPP), which has been discussed in the two previous chapters, and the Uncovered Interest rate Parity (UIP) to which we turn in this and the next chapter. The UIP hypothesis states that the expected exchange rate return should be equal to the forward premium, which is the proportion by which a country's forward exchange rate exceeds its spot exchange rate (see for example Solnik (2004)). In other words, if UIP holds then the forward premium is an unbiased estimator for the future return on the spot exchange rate. Because the forward premium is equal to the difference between the nominal interest rates of the underlying countries due to arbitrage arguments, UIP implies that the high interest rate country is expected to have a depreciating currency.[1] The intuition behind this is that for an international investor the expected gain due to an appreciation of the home/foreign currency should be offset by a larger interest rate in the foreign/home country.

A simple test for this hypothesis in a linear regression framework is to regress the future exchange rate return on the forward premium and test if the slope coefficient (say β) equals unity (see e.g. Fama (1984)). However, there is not much empirical support for UIP. Many studies find that the estimated slope coefficient β is negative and significantly different from unity for different exchange rates and for different sample periods (see e.g. Froot and Thaler (1990) and Engel (1996)) suggesting that the currency of the high interest rate country is appreciating instead.[2] Chinn and Meredith (2004), Chinn and Meredith (2005) and Chinn (2006) have similar results for short horizons, however, using longer-maturity (5 and 10 year) bonds they do find evidence for UIP at long horizons.

The main contribution of this chapter is to assess the performance of linear models in describing and predicting exchange rate returns in a framework which looks similar but is not exactly equal to the UIP framework that is usually used. The basic model that we apply is the one used in Clarida and Taylor (1997), who try to exploit information possibly present in the interest rate term structure to predict exchange rates and show that the term structure of differential interest rates of both currencies can be helpful in improving exchange rate forecasts. For each exchange rate in their data set (all U.S. dollar denominated) they propose a linear Vector Error Correction Model (VECM) on the spot and forward rates with different maturities.

The main motivation of applying the VECM model of Clarida and Taylor (1997) is that not only they seem to find evidence consistent with long-run UIP but, even more striking, in contrast with many studies they apparently find that their model

[1] Due to the Covered Interest rate Parity (CIP), which holds predominantly (see for example Taylor (1995) and the references therein), the forward premium is equivalent to the nominal interest rate differential (see also footnote 11 in Section 4.3 for more details).

[2] In Froot and Thaler (1990) it is stated that "the average coefficient across some 75 published estimates is -0.88".

does have superior out-of-sample performance compared to the random walk (RW) model.[3] Furthermore, we will extend the VECM model to incorporate the findings of Bansal (1997) and Wu and Zhang (1996) who show that the sign of the forward premium has in-sample information on explaining violations of UIP. Bansal (1997) also finds that nonlinear transformations of the forward premium (within a linear model) are in-sample statistically significant in explaining the average return. We also want to investigate if these transformations of the forward premium are informative when predicting the exchange rate return. More specifically, we replace the lagged forward premia which are present as explanatory variables in the linear VECM of Clarida and Taylor (1997) by asymmetric and nonlinear variants thereof as proposed by Bansal (1997) and Wu and Zhang (1996) to describe and forecast the return of several U.S. dollar denominated exchange rates.

Our results in this chapter suggests that the in-sample and out-of-sample evidence reported by Clarida and Taylor (1997) are not supported. The forward premia do not add value in describing and predicting the exchange rate return. Furthermore, the forecasting performance of the model by including asymmetric and nonlinear variants of the forward premia do not improve the model either. In particular, we do not find the superior performance of the VECM model relative to the RW model as reported by Clarida and Taylor (1997).

This chapter is structured as follows. Section 4.2 deals with the linear models and the forecasting framework. The data and the results are discussed in sections 4.3 and 4.4 respectively, followed by the conclusions in Section 4.5.

4.2 Methodology

In this section we will discuss the methodology used in this chapter. This section is divided into two parts. The first part introduces and discusses the linear models of Clarida and Taylor (1997), Bansal (1997), and Wu and Zhang (1996) which we will

[3] On page 358 of Clarida and Taylor (1997) it is stated that "The dollar-mark and dollar-yen estimates are also similar in that all of the forward premiums or error correction terms enter with a negative coefficient. This implies that, when dollar interest rates are high relative to yen and mark rates at any point in the term structure...there is a tendency for the dollar to depreciate. This is qualitatively consistent with long-run uncovered interest rate parity...". The notation for the forward premia is opposite to the notation usually used in this strand of literature which explains why the finding of a negative coefficient is in line with evidence for UIP. Note however that no statement about the length of the long-run horizon is made which is in contrast with the papers by Chinn where the spot exchange rate return and the forward premium are calculated over horizons of 5 to 10 year.

use to base our in- and out-of-sample analysis on. The second part of this section explains the methodology to measure the (relative) out-of-sample performance.

4.2.1 Models

Clarida and Taylor (1997) (henceforth denoted as CT97) try to predict the exchange rate return making use of possible information present in the term structure of differential interest rates of both of the currencies underlying the exchange rate. They note that although many studies find that the slope coefficient β in the regression of the exchange rate return on the forward premium is significantly smaller than unity (which is contrary to what UIP describes), β is also often significant different from (and often smaller than) zero and therefore they claim that the forward premium should have informational content in describing/predicting the exchange rate return. Inspired by the framework of Hall *et al.* (1992) they propose a linear Vector Error Correction Model (VECM) for each currency for the system containing the spot and forward rates with four different maturities.

In general a system of time series variables can be represented by a VECM when all individual series contain a unit root and one or more independent cointegrating relations between the variables in the system exist (see Engle and Granger (1987) for this VECM representation theorem). Each cointegrating relation is a linear combination of the variables in the system such that the resulting series is stationary. This cointegrating relation can be seen as a long-run equilibrium and is incorporated in the model as an error correction mechanism that steers the system towards this long-run equilibrium when deviations from this equilibrium occur. Note that the number of *independent* cointegrating relations in a system is at most the number of series in the system minus one.[4]

The VECM of CT97 is as follows

$$\Delta y_t = \mu + \Gamma_1 \Delta y_{t-1} + \Pi^* (\boldsymbol{s}_{t-1} - \boldsymbol{f}_{t-1}) + \varepsilon_t \tag{4.1}$$

where $y_t = [s_t, f_{4,t}, f_{13,t}, f_{26,t}, f_{52,t}]'$ with s_t the log spot exchange rate and $f_{i,t}$ the log forward exchange rate with maturity i in weeks with $i \in \{4, 13, 26, 52\}$, $\boldsymbol{s}_t$ a vector of size 4 with for each element the same s_t, $\boldsymbol{f}_t = [f_{4,t}, f_{13,t}, f_{26,t}, f_{52,t}]'$, μ is a vector of size 5, Γ_1 is a 5x5 matrix and Π^* is a 5x4 matrix which contain parameters that need to be estimated."

For this model of CT97 to be valid, the spot and the forward exchange rates of this system should not only have to contain a unit root, or equivalently be integrated of order 1 (in short I(1)) but they also have to be cointegrated with each other, sharing

[4] See for example Verbeek (2000) for a discussion on the implementation of a VECM.

precisely one common stochastic trend. In this case, [1, -1] should be the cointegrating vector for the spot rate and each possible forward rate in the system because the forward premia enter their model as the long-run equilibria. For the second condition to hold, it would be sufficient to show that all possible forward premia are stationary.[5] In that case the system would have four independent cointegrating relations which is the maximum number in this system.

CT97 first show that a unit root is present in the log spot exchange rate (U.S. dollar denominated) and the log forward exchange rates for all four different maturities for each of the countries Germany, Japan and the United Kingdom. Many other studies have shown that the spot exchange rate and the forward exchange rate are both I(1), see for example Cornell (1977) and Meese and Singleton (1982) so CT97's finding is in agreement with the general consensus. Furthermore, CT97 also find evidence that all forward premia are stationary implying that there is (only) one common stochastic trend that is the driving force behind the system of the spot and forward exchange rates for each currency. This result is not straightforward because other studies find mixed results when the stationarity of the forward premia is investigated. Some find that the forward premium is I(0) like Hai *et al.* (1997), some find that it is I(1) such as Crowder (1994), and even others show that the forward premium is fractionally integrated e.g. Baillie and Bollerslev (1994). Engel (1996) notes that the different conclusions in these and other studies on the stationarity of the forward premium (often on the same set of currencies) is likely due to the applied test statistics and not so much to the used sampling period. In sum, with these results CT97 validate the use of their VECM to model the dynamics of the system consisting of the log spot exchange rate and the log forward exchange rates. With this model they find that in- and out-of-sample the term structure has informational content in explaining the exchange rate return. Therefore we will use this model for our data set, which is similar to the one of CT97, as well.

Furthermore, we will also modify the VECM in this chapter to accommodate for other information which might be relevant to improve in describing and predicting the exchange rate. Here we use findings of another stream of literature that analyzes the validity of the UIP hypothesis. We will use the studies by Bansal (1997) and Wu and Zhang (1996) who both test if an asymmetric relationship between the exchange

[5] Note that evidence against a stationary forward premium does not necessarily mean that the spot and forward rate are not cointegrated but it does mean however that the cointegrating vector is not [1, -1].

rate return and the forward premium exists.[6] To be more precise, they investigate whether a positive forward premium has a different impact on the exchange rate return than a negative forward premium (or forward discount) has. The model can be written as[7,8]

$$\Delta s_{t+1} = \alpha + \beta_1 (f_t - s_t)^+ + \beta_2 (f_t - s_t)^- + \varepsilon_t \tag{4.2}$$

Both studies report that for a negative forward premium the estimated slope coefficient, β_2, is still negative and significantly different from one, corresponding with the general finding that the currency of the high interest rate country is rather inclined to appreciate than to depreciate. However, for the case where the forward premium is positive, the slope β_1 is positive and not significantly different from unity.[9] Wu and Zhang (1996) do not have an explanation for this asymmetry while Bansal (1997) argue that asymmetry across countries of the market price of interest rate risk could be a possible explanation.

In addition to the asymmetric relationship between the forward premium and the exchange rate return, Bansal (1997) tests the inclusion of several nonlinear variations of the forward premium in an effort to explain the exchange rate return by estimating

$$\Delta s_{t+1} = \alpha + \beta_1 (f_t - s_t) + \beta_2 (f_t - s_t)^2 + \beta_3 (f_t - s_t)^3 + \varepsilon_t \tag{4.3}$$

He finds that these variables have a significant in-sample impact in explaining the change in the log spot exchange rate (dollar denominated).

[6] Formally Bansal (1997) and Wu and Zhang (1996) use the nominal interest rate differential of the underlying currencies instead of the forward premium but these two variables are equal due to the covered interest rate parity, which holds predominantly (see for example Taylor (1995) and the references therein). Therefore, in this chapter we will report their findings in terms of forward premia as opposed to using nominal interest rate differentials.

[7] In the model of Wu and Zhang (1996) not only the slope coefficient but also the intercept depend on the sign of the forward premium. We will not pursue this any further.

[8] $(f_t - s_t)^+$ and $(f_t - s_t)^-$ are defined respectively by

$$(f_t - s_t)^+ = \begin{cases} f_t - s_t, & \text{if } f_t - s_t > 0 \\ 0, & \text{if } f_t - s_t \leq 0 \end{cases}$$

and

$$(f_t - s_t)^- = \begin{cases} 0, & \text{if } f_t - s_t > 0 \\ f_t - s_t, & \text{if } f_t - s_t \leq 0. \end{cases}$$

[9] Bansal (1997) and Wu and Zhang (1996) define the exchange rate as the U.S. dollar price for one unit of foreign currency. This definition has no effect on the sign of the coefficient of the forward premium because the exchange rate return changes accordingly. However, when the distinction is made between a forward premium and forward discount then the corresponding coefficients in Bansal (1997) and Wu and Zhang (1996) have the opposite sign with respect to our coefficients for the forward premium and forward discount. To avoid confusion we have translated their results according to Equation (4.2).

In this chapter we want to investigate whether the VECM of CT97 combined with the insights of Bansal (1997) and Wu and Zhang (1996) not only does a better job in-sample but also improves the out-of-sample performance as documented by CT97. With this in mind we will combine the VECM specification as in Equation (4.1) and incorporate the asymmetric or nonlinear variants of the forward premium in the model. Note that we replace each forward premium $(s_{t-1} - f_{i,t-1})$ where $i \in \{4, 13, 26, 52\}$ with its asymmetric or nonlinear counterpart in order to limit the proliferation of parameters. In particular we will estimate the following models

$$\Delta y_t = \mu + \Gamma_1 \Delta y_{t-1} + \Pi^* x_{t-1} + \varepsilon_t \tag{4.4}$$

where Π^* is a 5x4 matrix except for Equation (4.5) (see below) where Π^* is a 5x8 matrix. We have the following cases for x_{t-1}:

$$x_{t-1} = [((\boldsymbol{s}_{t-1} - \boldsymbol{f}_{t-1})^+)', ((\boldsymbol{s}_{t-1} - \boldsymbol{f}_{t-1})^-)']', \tag{4.5}$$

$$x_{t-1} = (\boldsymbol{s}_{t-1} - \boldsymbol{f}_{t-1})^+, \tag{4.6}$$

$$x_{t-1} = (\boldsymbol{s}_{t-1} - \boldsymbol{f}_{t-1})^-, \tag{4.7}$$

$$x_{t-1} = (\boldsymbol{s}_{t-1} - \boldsymbol{f}_{t-1})^2 \tag{4.8}$$

and

$$x_{t-1} = (\boldsymbol{s}_{t-1} - \boldsymbol{f}_{t-1})^3 \tag{4.9}$$

We note that the replacement of the forward premia by asymmetric or nonlinear variants transforms the models from VECM into VARX models because in these cases x_{t-1} are lagged exogenous variables. See Lütkepohl (1993) for more details.

4.2.2 Forecasting

In this part we give a detailed description of the forecasting methodology that is used to predict the exchange rate returns and relate the out-of-sample performance of the models to each other.

To measure the out-of-sample performance of our model we relate its forecasts to the forecasts of two other models. The first benchmark model we use is the standard random walk (RW) model which simply states that exchange rate returns are not predictable.

$$\Delta s_t = \kappa + \varepsilon_t \tag{4.10}$$

where κ is a constant.

The second model is a fourth-order Vector Auto Regression (VAR(4)) model which CT97 also use in their paper as benchmark for the VECM. The model has the same dependent variable $y_t = [s_t, f_{4,t}, f_{13,t}, f_{26,t}, f_{52,t}]'$ as the VECM model. This model is described by

$$\Delta y_t = \mu + \sum_{j=1}^{4} \Gamma_j \Delta y_{t-j} + \varepsilon_t \tag{4.11}$$

We will also add $x_{i,t-1}$ with $i \in \{4, 13, 26, 52\}$ to this VAR(4) model in a similar way as is done for the VECM model so that the in-sample and out-of-sample performance of both models can be compared directly. In other words, we can directly compare the forecasting performance of the models from Equation (4.4) and (4.11) when we add one of the specifications of Equation (4.5) to (4.9).

The actual calculations needed to obtain the forecasts are done as follows. First, a window is defined using, say, n observations over which the model for the exchange rate return is estimated. Second, for the 52 weeks after the estimation period forecasts are made, i.e. $\Delta\hat{s}_{n+h}$ for $h \in \{1, 2, ..., 52\}$. This is done by rolling the model forward using in-sample based information only. To be more precise, the one-week ahead forecast is constructed in a straightforward way using the estimated model. For example, if we use the VECM of Equation (4.1) we have that

$$\Delta\hat{y}_{n+1} = E_n [\Delta y_{n+1}] = \hat{\mu} + \hat{\Gamma}_1 E_n [\Delta y_n] + \hat{\Pi}^* E_n [(\boldsymbol{s}_n - \boldsymbol{f}_n)] =$$

$$= \hat{\mu} + \hat{\Gamma}_1 \Delta y_n + \hat{\Pi}^* (\boldsymbol{s}_n - \boldsymbol{f}_n)$$

For the two-week ahead forecast we have to use the one-step ahead forecast, otherwise out-of-sample information would be used, and so on. In other words

$$\Delta\hat{y}_{n+h} = E_n [\Delta y_{n+h}] = \hat{\mu} + \hat{\Gamma}_1 E_n [\Delta y_{n+h-1}] + \hat{\Pi}^* E_n \left[(\boldsymbol{s}_{n+h-1} - \boldsymbol{f}_{n+h-1})\right] =$$

$$= \hat{\mu} + \hat{\Gamma}_1 \Delta\hat{y}_{n+h-1} + \hat{\Pi}^* (\hat{\boldsymbol{s}}_{n+h-1} - \hat{\boldsymbol{f}}_{n+h-1})$$

for $h = \{2, 3, ..., 52\}$.[10] In total we generate weekly forecasts for the first 52 weeks following the estimation period. The forecasts of the log exchange rates can be

[10] This procedure is applied for all models, however, we note that for the models where asymmetric or nonlinear variants of the forward premium enter the equation, rolling the model forward is not the most appropriate procedure to estimate $E_n [\Delta y_{n+h}]$ for $h = \{2, 3, ..., 52\}$. To show this, let $g(.)$ be an asymmetric or nonlinear function as in Equation (4.5) to (4.9) such that $x_{n+h-1} = g(\boldsymbol{s}_{n+h-1} - \boldsymbol{f}_{n+h-1})$. Then Jensen's inequality states that $E_n \left[g(\boldsymbol{s}_{n+h-1} - \boldsymbol{f}_{n+h-1})\right]$ is not necessarily equal to $g(E_n \left[(\boldsymbol{s}_{n+h-1} - \boldsymbol{f}_{n+h-1}\right])$ (note that the max-, min- and power-functions are all convex or concave functions). A more appropriate way to construct forecasts would be to simulate several paths for Δy_{n+h} for $h = \{2, 3, ..., 52\}$ from the model under consideration where distributional assumptions for the errors have to be made and estimate $E_n [\Delta y_{n+h-1}]$ by averaging over the simulations. However, a consequence of using simulations would be that distributional assumptions about the errors have to be made which can have an impact on the inferences drawn from this simulation exercise. Taking this into consideration we have decided to leave the simulation procedure for future research.

constructed in a straightforward manner using the first element of $\Delta\hat{y}_{n+1}$ as follows

$$\hat{s}_{n+1} = s_n + \Delta\hat{s}_{n+1}$$

and

$$\hat{s}_{n+h} = \hat{s}_{n+h-1} + \Delta\hat{s}_{n+h}$$

for $h = \{2, 3, ..., 52\}$. The prediction error is defined for each of the 52 weeks by the difference between the predicted and real log exchange rate

$$PE(h) = \hat{s}_{n+h} - s_{n+h}$$

If we have obtained the forecasts induced by the model estimated over the estimation period, this estimation window is extended by one observation and the whole procedure (estimation and forecasting) is repeated again. This is done K times resulting in a series of prediction errors $PE_k(h)$ for $k = 1, ..., K$ for the log exchange rates. These prediction errors are used in the calculation of the forecasting measures Root Mean Squared Error (RMSE) and Mean Absolute Error (MAE) which are also used by CT97 to measure the forecasting performance of the models. These well-known measures are defined by

$$RMSE(h) = \left\{ \frac{1}{K} \sum_{k=1}^{K} (PE_k(h)) \right\}^{0.5}$$

$$MAE(h) = \frac{1}{K} \sum_{k=1}^{K} |PE_k(h)|$$

For comparability reasons we keep K equal over all forecasting horizons.

Because we use the 4-, 13-, 26-, and 52-week forward information, the forecasting measures RMSE and MAE are reported for these horizons. Using these measures the forecasting performance of the models can be assessed. To compare the relative forecasting performance, all these measures are benchmarked to the forecasting performance of the RW model. In the next section the data used for estimation and forecasting of the models are discussed.

4.3 Data

We collected exchange rate data (WM/Reuters Closing Spot Rates) on a weekly basis from Datastream for the period 1978:31-2004:42 of U.S. dollar nominated exchange rates for Germany, Japan and the United Kingdom. For the period from 1999:01

the euro-dollar exchange rate is used to obtain a synthetic DMark-dollar rate. Furthermore, for the same period, frequency and corresponding countries we collected the Financial Times Eurocurrency Rates (annualized forward rates with maturities of 1 month, 3 months, 6 months and 1 year). These series are used to construct forward (dollar nominated) exchange rates $F_{h,t}$ with maturity h, $h \in \{4, 13, 26, 52\}$ using the arbitrage relationship.[11] In the further analysis the logarithm of the spot and forward rates, $s_t = \log(S_t)$ and $f_{h,t} = \log(F_{h,t})$ respectively, are used. The period 1978:31-1990:26 is used for estimation of the model (622 observations) while the period 1990:27-2004:42 is used in the forecasting analysis (746 observations).

The spot rates s_t and the forward rates $f_{h,t}$ for all maturities $h \in \{4, 13, 26, 52\}$ for all countries are displayed in Figure 4.1. First of all we note that for each currency the spot and forward exchange rates are almost indistinguishable from each other and seem to follow the same pattern. It also appears that the series are not mean reverting and therefore might contain a unit root. This is confirmed by the results in Table 4.1 where the p-values of the standard Augmented Dickey-Fuller unit root test (see Dickey and Fuller (1979)) for the level and return of s_t and $f_{h,t}$ for all maturities $h \in \{4, 13, 26, 52\}$ for all countries are reported. The level of all series contain a unit root while the return is stationary which means that the spot and forward exchange rates are all I(1) which corresponds to the general finding in the literature.

To visualize the possible relationship between the spot exchange rate return and the forward premium, these series are displayed in Figure 4.2.[12] We see that there is much more variability in the spot exchange rate return than in the forward pre-

[11] For each currency and time to maturity h in weeks, $h \in \{4, 13, 26, 52\}$, we have by arbitrage the following relationship

$$F_{h,t} = S_t \frac{1 + \frac{hr_{h,t}}{5200}}{1 + \frac{hr^{U.S.}_{h,t}}{5200}}$$

where S_t denotes the U.S. dollar denominated spot exchange rate, $r_{h,t}$ the annualized h-week nominal interest related to the non-U.S. dollar currency underlying the exchange rate, and $r^{U.S.}_{h,t}$ the h-week U.S. nominal interest rate. This relation is called the Covered Interest rate Parity (CIP) which holds predominantly according to Taylor (1995). To see why this relation must hold imagine that a U.S.-based investor puts \$ 1 on a local bank account for a certain horizon h against the appropriate interest rate which is $r^{U.S.}_{h,t}$. After h weeks he will have $1 + hr^{U.S.}_{h,t}/5200$. He could also put the money on a foreign bank account. Then he first would convert the \$ 1 by using the exchange rate for which he will get S_t units of the foreign currency. That money will grow to $S_t(1 + hr^{U.S.}_{h,t}/5200)$ which he has to convert back to dollars with the forward rate $F_{h,t}$ so he will obtain $S_t(1 + hr^{U.S.}_{h,t}/5200)/F_{h,t}$. Both options are without risk so they should have the same proceeds, meaning that $1 + hr^{U.S.}_{h,t}/5200 = S_t(1 + hr^{U.S.}_{h,t}/5200)/F_{h,t}$ which after rearranging gives the above equation. Note that for UIP no arbitrage arguments can be used because the future spot exchange rate is unknown ex ante.

[12] Because the graphs are similar for different times to maturity and just meant for illustrative purposes, we only use the 4-week forward rate.

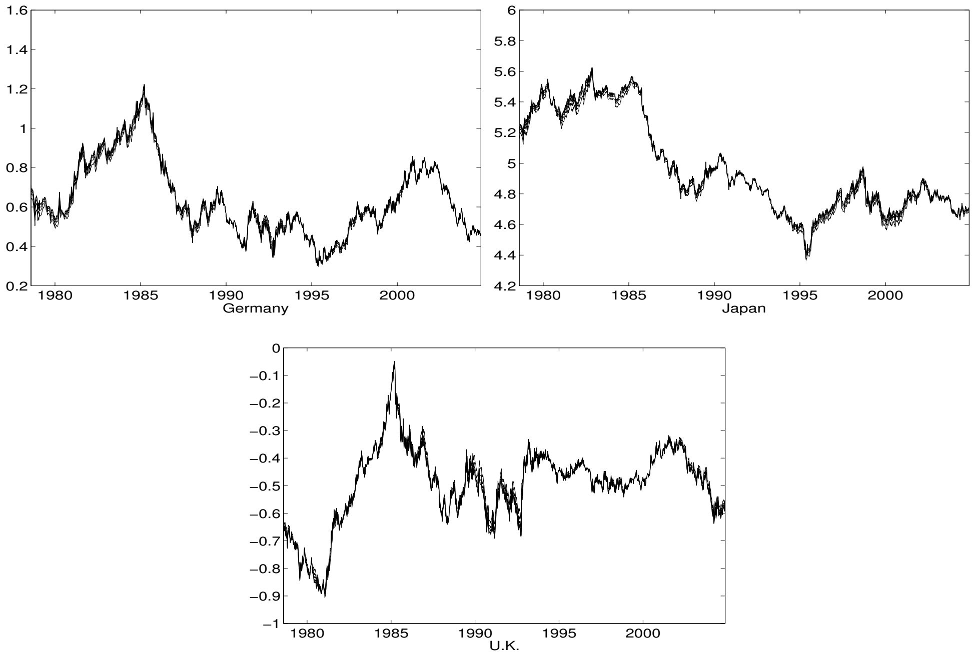


Figure 4.1: Spot and forward exchange rates. U.S. dollar denominated spot exchange rates s_t and forward exchange rates $f_{h,t}$ for all maturities $h \in \{4, 13, 26, 52\}$ for the period 1978:31 to 2004:42 for Germany, Japan, and the U.K. are displayed.

mium. The exchange rate return indeed looks stationary (see again also Table 4.1), while the forward premium seems to contain a unit root. To formally investigate the stationarity of the forward premium the ADF unit root test is conducted. To determine the lag length needed for the ADF unit root test both the Schwarz and Akaike info criteria are used.[13] The p-values for the ADF unit root test are reported in Table 4.2. Although most of the times the unit root is rejected at the 5% level so that there is proof that the forward premium is stationary, the results seem sensitive to the choice of the criterion used to determine the lag length. This finding is in line with the more general statement of Engel (1996) as already discussed in Section 4.2 that evidence of a stationary forward premium is likely to depend on the applied test statistic.

[13] Both the Schwarz and Akaike criteria were also used when the stationarity of the level and the return for the spot exchange rate and the forward exchange rates were investigated (see Table 4.1). Because both unambiguously led to the same conclusion of stationarity of the series only the results for the Schwarz criterion were reported.

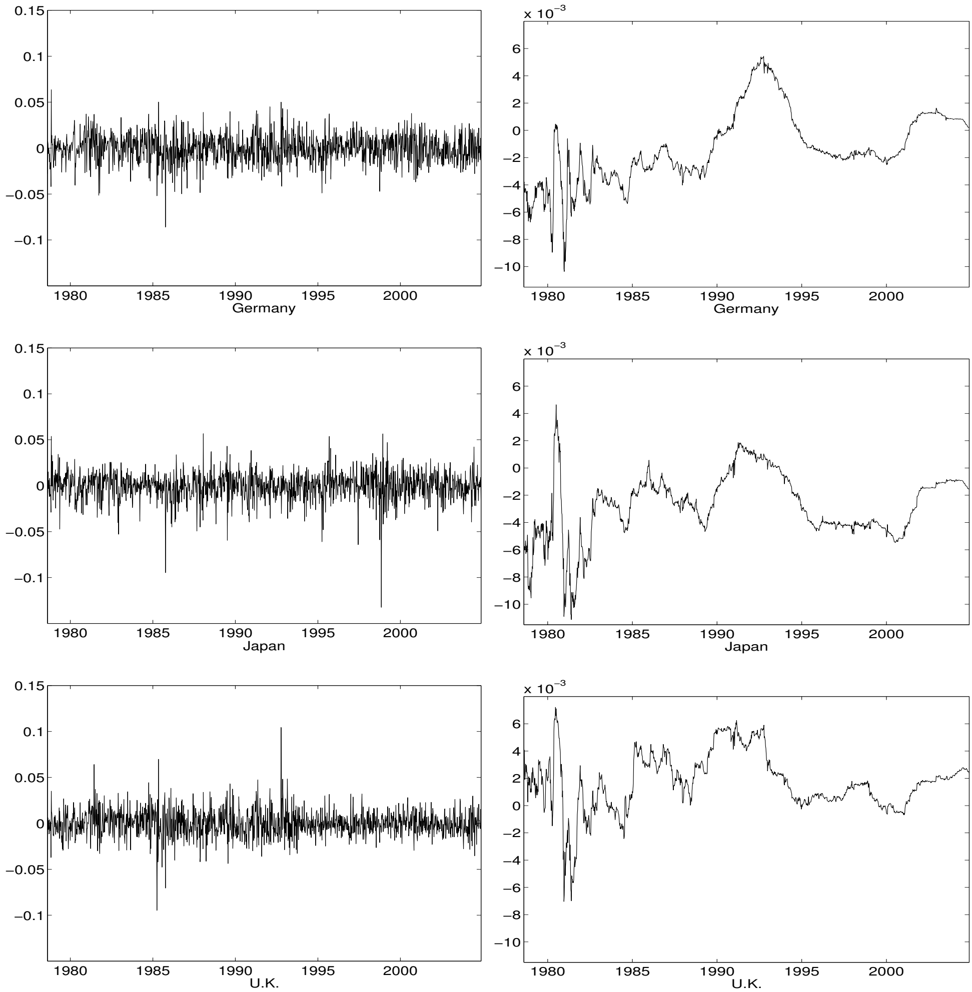


Figure 4.2: Spot exchange rate return and forward premium. Spot exchange rate return (left) and forward premium (right) for the 4-week forward rate for the period 1978:31 to 2004:42 for Germany, Japan, and the U.K. are displayed.

	Germany				
	s_t	$f_{4,t}$	$f_{13,t}$	$f_{26,t}$	$f_{52,t}$
p-value ADF: level	0.365	0.368	0.373	0.381	0.393
p-value ADF: 1st diff.	0.000***	0.000***	0.000***	0.000***	0.000***
	Japan				
	s_t	$f_{4,t}$	$f_{13,t}$	$f_{26,t}$	$f_{52,t}$
p-value ADF: level	0.279	0.283	0.287	0.293	0.305
p-value ADF: 1st diff.	0.000***	0.000***	0.000***	0.000***	0.000***
	U.K.				
	s_t	$f_{4,t}$	$f_{13,t}$	$f_{26,t}$	$f_{52,t}$
p-value ADF: level	0.427	0.426	0.423	0.419	0.415
p-value ADF: 1st diff.	0.000***	0.000***	0.000***	0.000***	0.000***

Table 4.1: Unit root tests on the level and first difference of spot and forward exchange rates. The p-values for the ADF unit root test (H_0: unit root) on the level and return of the spot s_t and forward $f_{h,t}$ exchange rates for all maturities $h \in \{4, 13, 26, 52\}$ for the period 1978:31 to 2004:42 for Germany, Japan, and the U.K. are reported. *, ** and *** denote significance at the 10%, 5% and 1% level, respectively.

Finally, for each currency we test how many cointegrating relations there are in the system of spot and the four forward exchange rates using the trace test and maximum eigenvalue test (see Johansen (1991)). We see from Table 4.3 that Germany has three cointegrating relations and both Japan and the U.K. have the maximum of four independent cointegrating relations. Together with the earlier findings of the non-stationarity and stationarity of respectively the level and first difference of the spot and forward exchange rates and the stationarity of the forward premium, we conclude that for Japan and U.K. the VECM with the forward premia as cointegrating relations as proposed by Clarida and Taylor (1997) is valid. It remains a puzzle why for Germany three cointegrating relations are found instead of four as you would expect. Also from Figure 4.1 it is expected that there is one force that drives the spot and forward exchange rates implying that this system should contain four cointegrating relations. Therefore, and for comparison reasons, we will also use the VECM for Germany.

	Germany			
	$f_{4,t} - s_t$	$f_{13,t} - s_t$	$f_{26,t} - s_t$	$f_{52,t} - s_t$
***p*-value ADF: Schwarz**	0.009***	0.009***	0.018**	0.061*
***p*-value ADF: Akaike**	0.019**	0.013**	0.018**	0.016**
	Japan			
	$f_{4,t} - s_t$	$f_{13,t} - s_t$	$f_{26,t} - s_t$	$f_{52,t} - s_t$
***p*-value ADF: Schwarz**	0.022**	0.044**	0.076*	0.104
***p*-value ADF: Akaike**	0.016**	0.021**	0.030**	0.046**
	U.K.			
	$f_{4,t} - s_t$	$f_{13,t} - s_t$	$f_{26,t} - s_t$	$f_{52,t} - s_t$
***p*-value ADF: Schwarz**	0.009***	0.007***	0.008***	0.033**
***p*-value ADF: Akaike**	0.002***	0.013**	0.025**	0.037**

Table 4.2: Unit root tests on the forward premium. The p-values for the ADF unit root test (H_0: unit root) on the forward premium for the 4-week forward rate for the period 1978:31 to 2004:42 for Germany, Japan, and the U.K. are reported. Both the Schwarz and Akaike criteria are used to determine the lag length needed in the calculation of the unit root test statistic. *, ** and *** denote significance at the 10%, 5% and 1% level, respectively.

	Germany	Japan	U.K.
trace test	3**	4***	4***
maximum eigenvalue	3**	4***	4***

Table 4.3: Number of cointegrating relations in the system of spot and forward exchange rates. The number of cointegrating relations as indicated by the trace test and maximum eigenvalue test in the system of spot and forward exchange rates for the period 1978:31 to 2004:42 for Germany, Japan, and the U.K. are reported. *, ** and *** denote significance at the 10%, 5% and 1% level, respectively.

Germany

	Δs_t	$\Delta f_{4,t}$	$\Delta f_{13,t}$	$\Delta f_{26,t}$	$\Delta f_{52,t}$
Δs_{t-1}	-5.444***	-5.393***	-5.655***	-5.974***	-6.166***
	(1.942)	(1.933)	(1.918)	(1.902)	(1.908)
$\Delta f_{4,t-1}$	8.566***	8.408***	8.671***	9.006***	8.934***
	(2.810)	(2.797)	(2.776)	(2.753)	(2.761)
$\Delta f_{13,t-1}$	-2.526	-2.436	-2.420	-2.293	-1.995
	(1.678)	(1.670)	(1.658)	(1.643)	(1.649)
$\Delta f_{26,t-1}$	-1.234	-1.232	-1.294	-1.492*	-1.354
	(0.895)	(0.891)	(0.884)	(0.877)	(0.880)
$\Delta f_{52,t-1}$	0.717*	0.729**	0.773**	0.829**	0.660*
	(0.366)	(0.365)	(0.362)	(0.359)	(0.360)
$(s - f_4)_{t-1}$	5.492*	5.762*	4.798	5.600*	4.833
	(3.165)	(3.151)	(3.127)	(3.100)	(3.110)
$(s - f_{13})_{t-1}$	-1.937	-1.976	-1.230	-1.952	-0.832
	(2.122)	(2.113)	(2.097)	(2.079)	(2.085)
$(s - f_{26})_{t-1}$	-0.418	-0.462	-0.688	-0.353	-1.081
	(0.958)	(0.953)	(0.946)	(0.938)	(0.941)
$(s - f_{52})_{t-1}$	0.368	0.380	0.388	0.337	0.504**
	(0.243)	(0.242)	(0.240)	(0.238)	(0.239)
c	-0.005***	-0.005***	-0.005***	-0.005***	-0.005***
	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
R^2	0.0373	0.0377	0.0412	0.0443	0.0471

Table 4.4: This table reports the estimated parameters (standard errors between brackets) of the VECM of Clarida and Taylor (1997) over the period 1978:31-1990:26 for Germany. *, ** and *** denote significance at the 10%, 5% and 1% level, respectively.

4.4 Results

In this section we will report the main findings of our analysis. First we will look at some estimation results, followed by the forecasting analysis.

The estimation results of the VECM (see also Equation (4.1)) for Germany, Japan and the U.K. over the period 1978:31-1990:26 are reported in Table 4.4 to 4.6 respectively.

We see that the *lagged* spot and the 4-week forward exchange rate return for Germany and the U.K. are significant at the 1% level for the spot and forward exchange rate return. Also some other lagged forward premia have informational content in describing the spot and forward exchange rate returns for these two countries while for Japan non of these variables are significant. This is in sharp contrast with the results of CT97: for all three countries they find that a lot more lagged endogenous variables are significant. The values of the parameter estimates differ from the estimated

Japan

	Δs_t	$\Delta f_{4,t}$	$\Delta f_{13,t}$	$\Delta f_{26,t}$	$\Delta f_{52,t}$
Δs_{t-1}	-0.230	-0.212	-0.408	-0.642	-1.114
	(1.427)	(1.423)	(1.419)	(1.417)	(1.435)
$\Delta f_{4,t-1}$	1.971	1.960	2.293	2.523	3.169
	(2.089)	(2.084)	(2.077)	(2.074)	(2.100)
$\Delta f_{13,t-1}$	-1.236	-1.249	-1.430	-1.368	-1.826
	(1.282)	(1.279)	(1.274)	(1.273)	(1.289)
$\Delta f_{26,t-1}$	-0.840	-0.857	-0.870	-0.941	-0.439
	(0.770)	(0.768)	(0.765)	(0.764)	(0.774)
$\Delta f_{52,t-1}$	0.399	0.422	0.482	0.494	0.274
	(0.309)	(0.308)	(0.307)	(0.306)	(0.310)
$(s-f_4)_{t-1}$	-0.225	0.090	-0.816	-0.362	-0.440
	(2.334)	(2.327)	(2.319)	(2.316)	(2.345)
$(s-f_{13})_{t-1}$	0.197	0.133	0.698	0.120	0.508
	(1.497)	(1.493)	(1.488)	(1.486)	(1.505)
$(s-f_{26})_{t-1}$	-0.214	-0.279	-0.388	-0.091	-0.547
	(0.729)	(0.727)	(0.725)	(0.724)	(0.733)
$(s-f_{52})_{t-1}$	0.194	0.219	0.208	0.171	0.332
	(0.219)	(0.219)	(0.218)	(0.218)	(0.220)
c	-0.005***	-0.005***	-0.005***	-0.005***	-0.006***
	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
R^2	0.0441	0.0445	0.0501	0.050	0.0528

Table 4.5: This table reports the estimated parameters (standard errors between brackets) of the VECM of Clarida and Taylor (1997) over the period 1978:31-1990:26 for Japan. *, ** and *** denote significance at the 10%, 5% and 1% level, respectively.

values in CT97. Although we are using a slightly shorter sample (1978:31-1990:26 instead of 1977:01-1990:26) this finding is remarkable.

Furthermore CT97 find evidence for long-run UIP for the dollar-mark and the dollar-yen currencies because all the forward premia are negative and significantly different from zero. Note from footnote 3 that the forward premia enter the model with an opposite sign as is usual in this strand of literature such that in this case this means that there is evidence for long-run UIP. For the dollar-sterling they find that the signs are mixed. We do not find evidence in favor of long-run UIP: for Japan and the U.K. none of the error correction terms are significant and although for Germany one out of the four forward premia is significant in explaining the spot exchange rate return, the sign is in contradiction for proof of long-run UIP. In short, we are unable to replicate the in-sample evidence of CT97 that many of the endogenous variables are significant and that the uncovered interest rate parity holds in the long-run.

U.K.

	Δs_t	$\Delta f_{4,t}$	$\Delta f_{13,t}$	$\Delta f_{26,t}$	$\Delta f_{52,t}$
Δs_{t-1}	-4.282***	-4.254***	-4.533***	-4.645***	-5.173***
	(1.585)	(1.583)	(1.590)	(1.615)	(1.657)
$\Delta f_{4,t-1}$	6.714***	6.539***	6.728***	6.676***	6.925***
	(2.407)	(2.404)	(2.414)	(2.452)	(2.516)
$\Delta f_{13,t-1}$	-0.556	-0.388	-0.213	0.177	0.616
	(1.509)	(1.507)	(1.513)	(1.537)	(1.577)
$\Delta f_{26,t-1}$	-2.439***	-2.472***	-2.586***	-2.870***	-2.826***
	(0.808)	(0.807)	(0.810)	(0.823)	(0.845)
$\Delta f_{52,t-1}$	0.604*	0.614*	0.642*	0.698*	0.494
	(0.356)	(0.356)	(0.357)	(0.363)	(0.372)
$(s - f_4)_{t-1}$	2.756	3.114	2.344	2.695	2.151
	(2.624)	(2.621)	(2.631)	(2.673)	(2.743)
$(s - f_{13})_{t-1}$	-0.284	-0.423	0.172	-0.433	0.555
	(1.765)	(1.763)	(1.770)	(1.798)	(1.845)
$(s - f_{26})_{t-1}$	-0.757	-0.761	-0.932	-0.500	-1.284
	(0.946)	(0.945)	(0.948)	(0.963)	(0.989)
$(s - f_{52})_{t-1}$	0.342	0.351	0.352	0.255	0.467
	(0.283)	(0.282)	(0.284)	(0.288)	(0.296)
c	0.002***	0.002***	0.002***	0.002***	0.003***
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
R^2	0.0595	0.0595	0.0614	0.0660	0.0652

Table 4.6: This table reports the estimated parameters (standard errors between brackets) of the VECM of Clarida and Taylor (1997) over the period 1978:31-1990:26 for the U.K. *, ** and *** denote significance at the 10%, 5% and 1% level, respectively.

Trying to improve upon the in-sample performance of the model we also estimate the models in Equation (4.5) to (4.9). We assess the sign and significance of the lagged forward premia and their asymmetric and nonlinear variants in explaining the exchange rate return. The results for all countries are reported in Table 4.7. The first column denotes the results for the lagged forward premia in the original VECM (see also Equation (4.1)). For the other columns these lagged forward premia are replaced by the several asymmetric and nonlinear variants thereof. For the VAR(4) model we add the forward premia and the same variants to assess the importance of these variables in this model.

For the VAR(4) model with the added forward premia we have the same conclusions for Japan and the U.K. as we had for the VECM: none of them are significant. For Germany there are more significant occurrences, however, the forward premia have different signs for different horizons such that nothing can be said about the validity of long-run UIP. For the VECM and VAR(4) where the asymmetric or non-

Germany

		Model						
	horizon h	CT	A		B	C	D	E
VECM	4	pos*	pos*	-	pos*	-	-	
	13	-	-	-	-	-	-	-
	26	-	-	-	-	-	-	neg*
	52	-	-	-	-	-	pos**	pos***
VAR(4)	4	pos***	pos***	-	pos***	-	pos**	-
	13	neg***	neg***	-	neg***	-	-	-
	26	pos*	pos*	-	pos*	-	-	-
	52	-	-	-	-	-	-	pos**

Japan

		Model						
	horizon h	CT	A		B	C	D	E
VECM	4	-	-	-	-	-	-	-
	13	-	-	-	-	-	-	-
	26	-	-	-	-	-	-	-
	52	-	-	-	-	-	pos*	pos***
VAR(4)	4	-	-	-	-	-	-	pos*
	13	-	-	-	-	-	-	-
	26	-	-	-	-	-	-	-
	52	-	-	-	-	-	pos*	pos***

U.K.

		Model						
	horizon h	CT	A		B	C	D	E
VECM	4	-	-	pos*	-	pos**	neg*	-
	13	-	-	-	-	-	-	-
	26	-	-	-	-	-	-	-
	52	-	-	-	pos*	-	-	-
VAR(4)	4	-	-	pos**	-	pos**	-	-
	13	-	-	-	-	-	-	-
	26	-	-	-	-	-	-	-
	52	-	-	-	-	-	-	-

Table 4.7: This table reports the sign, positive (pos) or negative (neg), and significance of (asymmetric and nonlinear variants of) the forward premium in the VECM and the VAR(4) model over all countries, models, and horizons where CT is the model of Equation (4.4), VAR of Equation (4.11) and A to E of respectively Equation (4.5) to (4.9). Estimation of the models takes place over the period 1978:31-1990:26. With *, **, and *** we denote the significance at a level of 10%, 5%, and 1% respectively.

linear forward premia are added, we see for Japan and the U.K. that for the few cases where these variables are significant, they have the wrong sign (all except one). For Germany there are more occurrences compared to the other two countries where the asymmetric or nonlinear forward premia are significant but for each model we have either that the sign is wrong or that the signs are not the same. Summarizing, the transformations of the forward premia as advocated by Bansal (1997) and Wu and Zhang (1996) do not help in finding evidence for long-run UIP.

To measure the forecasting performance for each model we apply the methodology as described in Section 4.2. Table 4.8 reports the RMSE for the RW model for $h = \{4, 13, 26, 52\}$. These numbers are in general increasing for higher horizons because the volatility of the exchange rate increases as well. The relative forecasting performance of the VECM model with respect to the RW model is reported in the first column by the ratio RMSE(VECM)/RMSE(RW) which means the RW would outperform the VECM model for values larger than one. The other columns denote the relative performance of the several variants of the VECM model. The relative performance of the VAR(4) model (last columns) and its variants (first six columns) are also reported in this table. Table 4.9 gives the results for the MAE forecasting measure. For some combinations the RMSE and/or MAE are too large which is an indication that the model might be misspecified. In these cases the RMSE and/or MAE is not interpretable and are left out of the tables.

In contrast with the conclusions of CT97 we do not find evidence that the VECM has a superior performance relative to the RW model, on the contrary, our results are much more in favor of the RW model because the ratios between the RMSE/MAE of the RW model and the VECM are often larger than one. In general we see that the differences in the RMSE/MAE of the RW model and the VECM or VAR(4) model are much smaller than reported in Clarida and Taylor (1997). This could be due to the fact that their parameter estimates of the model are different from the ones we found which might have consequences for the forecasting performance as well.

Furthermore, if we compare the forecast results of the VECM (Equation 4.4) and VAR(4) (Equation (4.11)) we see that the VAR(4) is doing a better job most of the times. The performance of the VAR(4) model is in this respect similar to the RW model. Then, if we compare the variants of the VECM model to variants of the VAR(4) model for both the RMSE and MAE we see that for Germany and Japan the VECM has a better forecasting performance while for the U.K. both models give similar results but still the models are both defeated by the RW model.

In total this leads us to conclude that the VECM of CT97 does a poor job in describing the spot and forward exchange rate returns because relative to CT97's findings there are not that many significant lagged endogenous variables and/or er-

Germany

	horizon h	RMSE RW						
RW	4	0.030						
	13	0.057						
	26	0.079						
	52	0.113						
		RMSE VECM/RMSE RW						
		Model						
	horizon h	CT	A	B	C	D	E	
VECM	4	1.027	1.060	1.024	1.058	1.014	1.009	
	13	1.081	1.253	1.074	1.313	NA	1.040	
	26	1.154	1.770	1.145	2.130	NA	1.079	
	52	1.197	3.174	1.186	9.924	NA	1.105	
		RMSE VAR(4)/RMSE RW						
	horizon h	CT	A	B	C	D	E	VAR
VAR(4)	4	1.034	1.076	1.037	1.068	1.024	1.012	1.002
	13	1.075	1.268	1.088	1.342	NA	1.038	1.002
	26	1.145	2.044	1.130	2.677	NA	1.070	1.004
	52	1.181	6.460	1.204	NA	NA	1.084	0.999

Japan

	horizon h	RMSE RW						
RW	4	0.032						
	13	0.062						
	26	0.086						
	52	0.114						
		RMSE VECM/RMSE RW						
	horizon h	CT	A	B	C	D	E	
VECM	4	0.998	0.997	0.993	1.005	1.002	1.001	
	13	1.012	1.000	0.995	1.012	1.010	1.005	
	26	1.011	1.018	0.994	1.025	0.998	0.977	
	52	0.939	1.003	0.932	1.017	NA	0.905	
		RMSE VAR(4)/RMSE RW						
	horizon h	CT	A	B	C	D	E	VAR
VAR(4)	4	1.006	1.006	0.998	1.004	1.008	1.006	0.996
	13	1.023	1.012	1.003	1.010	1.032	1.027	0.996
	26	1.038	1.043	1.019	1.035	1.032	1.011	1.005
	52	0.971	1.040	0.965	1.026	0.978	0.924	0.999

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U.K.

	horizon h	**RMSE RW**						
RW	4	0.027						
	13	0.048						
	26	0.067						
	52	0.086						
		RMSE VECM/RMSE RW						
	horizon h	CT	A	B	C	D	E	
VECM	4	1.042	1.034	1.003	1.052	1.029	1.030	
	13	1.134	1.101	1.018	1.153	1.114	1.122	
	26	1.237	1.186	1.044	1.262	1.220	1.200	
	52	1.410	1.329	1.079	1.438	1.469	1.338	
		RMSE VAR(4)/RMSE RW						
	horizon h	CT	A	B	C	D	E	VAR
VAR(4)	4	1.044	1.038	1.003	1.052	1.027	1.026	0.996
	13	1.140	1.103	1.017	1.155	1.111	1.109	1.001
	26	1.239	1.183	1.043	1.255	1.212	1.186	1.004
	52	1.403	1.316	1.076	1.402	1.443	1.340	1.004

Table 4.8: This table reports the RMSE for the RW model. The RMSE for the VECM and the VAR(4) model are reported relative to the RMSE of the random walk model. A relative RMSE higher than one means that the RW has a better forecasting performance. CT is the model of Equation (4.4), VAR of Equation (4.11) and A to E of respectively Equation (4.5) to (4.9). The methodology used to calculate these numbers is described in Section 4.2.

ror correction terms. Furthermore, we can not find any evidence whatsoever that supports the validity of (long-run) UIP. Also the VECM does not perform superiorly out-of-sample with respect to the RW model as reported by CT97, in fact the RW model beats the VECM model in producing better forecasts of the spot exchange rate in terms of RMSE and MAE. Furthermore, the inclusion of the asymmetric or nonlinear variants of the forward premium does not improve the in-sample and out-of-sample performance of the VECM.

A seemingly obvious explanation for the difference between our findings and the findings of CT97 is that our data set is similar to theirs but it is not the same with respect to source and sample period; while CT97 obtained the spot and forward exchange rates from the Harris Bank database for the period 1977:01 through 1993:52, our data comes from Datastream covering the period 1978:31-2004:42 (estimation of the model takes place through 1990:26 in both CT97 and this chapter). This could imply that our forward premia are less stationary than the forward premia in CT97 such that they are less appropriate to act as long-run equilibria in the VECM. However, van Tol and Wolff (2005) and van Tol (2005), who use a multivariate

Germany

	horizon h	MAE RW						
RW	4	0.024						
	13	0.047						
	26	0.065						
	52	0.098						
		MAE VECM/MAE RW						
	horizon h	CT	A	B	C	D	E	
VECM	4	1.002	1.048	1.008	1.042	1.014	0.999	
	13	1.028	1.199	1.030	1.230	NA	1.021	
	26	1.082	1.432	1.080	1.543	NA	1.056	
	52	1.138	1.844	1.115	2.626	NA	1.086	
		MAE VAR(4)/MAE RW						
	horizon h	CT	A	B	C	D	E	VAR
VAR(4)	4	1.013	1.063	1.017	1.052	1.017	0.998	0.997
	13	1.039	1.193	1.035	1.244	NA	1.021	1.000
	26	1.075	1.431	1.083	1.681	NA	1.042	1.002
	52	1.131	2.037	1.150	NA	NA	1.064	0.998

Japan

	horizon h	MAE RW						
RW	4	0.024						
	13	0.049						
	26	0.069						
	52	0.094						
		MAE VECM/MAE RW						
	horizon h	CT	A	B	C	D	E	
VECM	4	0.991	0.992	0.988	1.012	0.990	0.989	
	13	0.980	0.974	0.964	1.021	0.980	0.981	
	26	0.977	0.984	0.961	1.031	0.981	0.957	
	52	0.907	1.006	0.905	1.030	NA	0.872	
		MAE VAR(4)/MAE RW						
	horizon h	CT	A	B	C	D	E	VAR
VAR(4)	4	1.004	1.006	0.996	1.016	0.997	0.994	1.001
	13	0.992	0.983	0.974	1.016	0.995	0.993	0.993
	26	1.001	1.003	0.987	1.041	1.009	0.983	1.006
	52	0.932	1.039	0.932	1.037	0.951	0.880	0.997

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U.K.

	horizon h	MAE RW						
RW	4	0.020						
	13	0.034						
	26	0.046						
	52	0.067						
		MAE VECM/MAE RW						
	horizon h	CT	A	B	C	D	E	
VECM	4	1.022	1.013	0.991	1.039	1.011	1.006	
	13	1.068	1.046	1.011	1.100	1.061	1.055	
	26	1.174	1.083	0.994	1.225	1.164	1.128	
	52	1.299	1.186	1.025	1.362	1.362	1.245	
		MAE VAR(4)/MAE RW						
	horizon h	CT	A	B	C	D	E	VAR
VAR(4)	4	1.026	1.021	0.995	1.042	1.015	1.007	1.000
	13	1.073	1.049	1.008	1.102	1.063	1.052	1.004
	26	1.174	1.081	0.991	1.213	1.161	1.109	1.004
	52	1.289	1.183	1.021	1.330	1.346	1.220	1.002

Table 4.9: This table reports the MAE for the random walk model. The MAE for the VECM and the VAR(4) model are reported relative to the MAE of the random walk model. A relative MAE higher than one means that the random walk has a better forecasting performance. CT is the model of Equation (4.4), VAR of Equation (4.11) and A to E of respectively Equation (4.5) to (4.9). The methodology used to calculate these numbers is described in Section 4.2.

threshold vector error correction model, are not able to replicate the results of CT97 either while using the exact same CT97 data set.

4.5 Conclusion

In this chapter we used the insights of Clarida and Taylor (1997), Bansal (1997), and Wu and Zhang (1996) in trying to explain and predict several U.S. dollar exchange rates in a linear framework. Both modeling and predicting exchange rates have become more important over the recent years as a result of an increasing perception by multinational companies and international investors of the foreign exchange rate risks that they run.

Clarida and Taylor (1997) find that the forecasting performance of their vector error correction model is superior to the random walk model which is a striking result because to our knowledge they are the first who present results that seem to beat the RW benchmark model for exchange rates using a linear model. Bansal (1997) and Wu and Zhang (1996) find an in-sample significance of the short term interest rate differential. In this chapter we extend the VECM model by including the short- and longer-term interest rate differential (with longer-term meaning maturities up to and including one year). Furthermore, motivated by the findings of Bansal (1997), we include nonlinear transformations of forward premia as explanatory variables in the linear model.

After having verified that the conditions for using the particular VECM of Clarida and Taylor (1997) are met (the spot and forward exchange rates are difference stationary, the four forward premia are stationary thereby allowed to act as the independent cointegrating relations in the system of these spot and forward exchange rates for two out of three currencies from our data set), the performance in-sample and out-of-sample is being investigated.

Surprisingly, contrary to Clarida and Taylor (1997) we do not find that with respect to the lagged forward premia the VECM is useful in describing the change in the exchange rate. Furthermore, replacing the lagged forward premia by asymmetric or nonlinear variants thereof does not lead to improvements in the fit. van Tol and Wolff (2005) and van Tol (2005) are not able to replicate the results of Clarida and Taylor (1997) either.

But the most striking result is that we do not find the superior outperformance of the VECM model as reported in Clarida and Taylor (1997). Also the absolute differences with respect to the RW benchmark model as measured by the forecast error statistics RMSE and MAE are much smaller for both the VECM and VAR(4) models. The inclusion of asymmetric and nonlinear variants of the forward premia does not seem to have any effect on the forecasting performance of the models. The main conclusion from this study has to be that the random walk model still remains the model that has to be beaten.

Chapter 5

Uncovered Interest Rate Parity in a Nonlinear Model

5.1 Introduction

In the previous chapter we focused on describing exchange rates in a linear framework. The models were motivated by the commonly held view that exchange rate changes should have a relationship with the forward premium. The linear models were inspired by the UIP theory. In this chapter the focus will be on the identification of nonlinear models.

Since the linear model as proposed by Fama (1984) to test the UIP hypothesis has in general produced counterintuitive results (see e.g Engel (1996) and Sarno (2005) and the references therein), other studies have considered that the relationship between the spot exchange rate return and the forward premium might be nonlinear (see e.g. Chinn (1991)).

Several authors have proposed theoretical motivations to justify possible nonlinearities. For example, Baldwin (1990) showed that small transaction costs and the uncertainty in future exchange rates can lead to a so-called hysteresis band: speculation only occurs when deviations from UIP are large enough. Dumas (1992) and Hollifield and Uppal (1997) also focus on transaction costs as a reason for the apparent nonlinearity. Lyons (2001) considers the "limitations to speculation hypothesis" which implies that exploitation of deviations from UIP only takes place when expected excess returns are higher than some threshold value. The role of interventions at the foreign exchange rate markets of central banks is discussed by Mark and Moh

(2004) and they find that deviations from UIP are more severe when central banks intervene.

One possibility to model the apparent nonlinearity between the FX return and the forward premium is to use a linear model where the dependent variable (the spot exchange rate return) is a linear function of nonlinear transformations of the independent variable (the forward premium). We have already seen a few examples of such models in the previous chapter: Bansal (1997) and Wu and Zhang (1996) assume asymmetry in such a way that a positive forward premium is allowed to have a different impact on the exchange rate return than a negative forward premium. Using a linear model they find that these variables are significant in explaining the exchange rate return. Additionally, Bansal (1997) finds that nonlinear transformations of the forward premium capture some of the dynamics present in the exchange rate. Another example of measuring nonlinearity in a linear framework is the panel approach of Huisman *et al.* (1998) who find that UIP holds when the forward premia are large.

Another possibility to model the potential nonlinearity between the spot exchange rate return and the forward premium is to propose models which are nonlinear in both parameters and variables. Baillie and Kiliç (2006) are one of the few studies where such a nonlinear model is introduced. Following the theories of for example Baldwin (1990) and Lyons (2001), they consider two cases when deviations from UIP occur. In the first case market participants are not willing to trade when UIP does not hold due to e.g. transaction costs or limitations to speculation thereby preventing that adjustments towards UIP do take place. In the second case the deviations from UIP are large enough to turn to speculation so that adjustment of the spot and forward exchange rates towards UIP takes place. To capture this in a nonlinear model Baillie and Kiliç (2006) use a dynamic logistic smooth transition regression model. In their model there are two regimes that are equivalent with the aforementioned cases. The transition between these two regimes occurs in a smooth way rather than discrete. In this way they hope to be able to pick up possible nonlinear and asymmetric relationships between the spot exchange rate return and the forward premium. They find that their model contributes to a better explanation of exchange rate behavior.

This chapter follows the approach of Baillie and Kiliç (2006) (henceforth denoted by BK06) as a starting point. First we rewrite their model by assuming that UIP holds and measure the deviations from this hypothesis by the nonlinear part of the model. Subsequently we construct confidence bounds for the nonlinear part to test for which time periods these deviations are significant.

Furthermore, just as the framework of Bansal (1997) and Wu and Zhang (1996) (henceforth jointly denoted as BWZ) seems to lead to better results with respect to trying to understand the UIP puzzle than the framework of Fama (1984), it can be tested whether the inclusion of the asymmetric and nonlinear transformations of the forward premium in the model of Baillie and Kiliç (2006) (this extension is from now on denoted as BK06-BWZ) leads to a better description of the spot exchange rate than when the original BK06 model is used. The BK06-BWZ model can be seen as an extension to the BWZ model. Consequently we can test whether the asymmetric and nonlinear transformations of the forward premium are still informative in explaining the spot exchange rate in the BK06-BWZ model or that the nonlinearity of the BK06 model is better able to describe the behavior of the spot exchange rate in the sense that the BWZ variables become redundant.

Finally, as another extension we present a multivariate case of our version of the BK06 model by collecting the different spot exchange rate returns into a panel such that we make use of the possible cross-sectional information present in the exchange rate returns. For example, we want to make use of the fact that the U.S. dollar nominal interest rate is a common factor in all univariate relationships.[1]

The outline of the chapter is as follows: in the next section the methodology we use is discussed. Section 5.3 and Section 5.4 deal respectively with the data and the estimation results. Finally some concluding remarks are given.

5.2 Models

In this section we will explain in detail the methodology we use in the chapter. Our contribution to the extant literature is threefold. First we measure the extent to which UIP holds. We will pursue this by fitting a nonlinear model that consists of two parts: a part where UIP is explicitly imposed and a part that measures the deviation from this restriction. Subsequently confidence bounds are constructed for these deviations such that the UIP hypothesis can be tested. The second contribution is the extension of the nonlinear model to include the asymmetric and nonlinear transformations of the forward premium as advocated by Bansal (1997) and Wu and Zhang (1996). We test whether the nonlinearity of the BK06-model causes the informativeness of the variables used in the model of BWZ to disappear. The last contribution of the chapter is that we extend the BK06 model to a multivariate case where interdependencies between the variables are exploited. Particularly, we can exploit the fact that there is a common U.S. dollar component in the interest rate differential.

[1] This can easily be seen when the forward premium is rewritten as the interest rate differential, i.e. the difference between the nominal interest rates for the currencies underlying the exchange rate (see also footnote 6 in the previous chapter).

To motivate the nonlinear model we start from the linear regression of the change in the exchange rate on the forward premium or equivalently the nominal interest rate differential which is a common way to test UIP using a linear model (see Fama (1984)).

$$\Delta s_{i,t+1} = \alpha + \beta(r_{i,t} - r_{US,t}) + u_{i,t} \tag{5.1}$$

If UIP holds α is equal to zero and β is equal to one. However, many empirical studies (e.g. Engel (1996) and Sarno (2005) and the references therein) find a negative β, significantly different from one leading to the conclusion that UIP should be rejected.

Recently, more sophisticated models are introduced to disclose evidence in favor of UIP. These are not only linear models as we have already seen in the previous chapter but also nonlinear models are used to explain the relationship between the exchange rate return and the forward premium. BK06 suggest a dynamic logistic UIP regression model by estimating a Logistic Smooth Transition Regression model to uncover UIP. With this model BK06 try to pick up possible nonlinear and asymmetric relationships between the spot exchange rate return and the forward premium by using a smooth transition between two regimes that are equivalent with the cases that UIP does and does not hold. The model looks as follows.

$$\Delta s_{i,t+1} = \alpha_1 + \beta_1 z_{i,t} + [\alpha_2 + \beta_2 z_{i,t}]F(z_{i,t}; \gamma, c) + u_{i,t} \tag{5.2}$$

where

$$F(z_{i,t}; \gamma, c) = \frac{1}{1 + \exp(-\gamma(z_{i,t} - c)/\sigma_z)} \tag{5.3}$$

with $z_{i,t} \equiv r_{i,t} - r_{US,t}$ the interest rate differential. The two regimes are connected by the transition function $F(z_{i,t}; \gamma, c)$. In the case that the transition function is close to 0 (arising for example when $z_t << 0$ and $\gamma > 0$ or vice versa) UIP holds when $\alpha_1 = 0$ and $\beta_1 = 1$ while in case that the transition function is close to 1 (arising for example when $z_t >> 0$ and $\gamma > 0$ or vice versa) UIP holds when $\alpha_1 + \alpha_2 = 0$ and $\beta_1 + \beta_2 = 1$. BK06 estimate the model in such a way that the latter restrictions are true meaning that UIP holds for the second regime. We, however, rewrite the model such that UIP holds for the first regime and that possible deviations from this assumption are measured by the nonlinear part of the model. Furthermore we fix the c parameter by the mean of the interest rate differential so that our base model is as follows[2]

$$\Delta s_{i,t+1} = z_{i,t} + [\alpha + \beta z_{i,t}]F(z_{i,t}; \gamma) + u_{i,t} \tag{5.4}$$

where

$$F(z_{i,t}; \gamma) = \frac{1}{1 + \exp(-\gamma(z_{i,t} - \mu_z)/\sigma_z)} \tag{5.5}$$

[2]By fixing c to the obvious benchmark (unconditional mean of interest rate differential) we limit the number of nuisance parameters in the model.

In this formulation we assume that UIP holds and deviations from this parity are measured by the nonlinear part of the model. Note that for UIP to hold, there is no a priori theoretical value where γ should be equal to. In theory only a γ "equal" to plus (minus) infinity and the interest rate differential below (above) its mean results in a value of zero for the transition function which is impossible in practice. For other cases UIP only holds when the parameters α and β are simultaneously equal to zero. Because it is impossible to statistically test whether γ is equal to plus or minus infinity, inferences on the validity of UIP must be done by formally testing whether the nonlinear deviations from UIP are statistically significantly different from zero. For that reason dynamic confidence bounds for the nonlinear part of the model are constructed using the familiar Delta method (see for example Oehlert (1992)) which is explained in more detail in the Appendix. In time periods where the deviations fall outside the confidence bounds we can not uphold the UIP hypothesis while for the time periods where they are not significantly different from zero UIP can not be rejected.

As a first extension to the base model in Equation (5.4) we add asymmetric and nonlinear variants of the interest rate differential to the nonlinear part of the model because these variables have proven to be valuable in linear models (see Bansal (1997) and Wu and Zhang (1996)). More specifically, we replace the interest rate differential by the positive and negative interest rate differential so that the model becomes

$$\Delta s_{i,t+1} = z_{i,t} + [\alpha + \beta_1 z_{i,t}^{+} + \beta_2 z_{i,t}^{-}]F(z_{i,t};\gamma) + u_{i,t} \tag{5.6}$$

where $F(z_{i,t};\gamma)$ as in (5.5).[3] Moreover, we will also add the square of the interest rate differential to the nonlinear part of Equation (5.4) so that we get the following model[4]

$$\Delta s_{i,t+1} = z_{i,t} + [\alpha + \beta_1 z_{i,t} + \beta_2 (\frac{z_{i,t}}{\sigma_z})^2]F(z_{i,t};\gamma) + u_{i,t} \tag{5.7}$$

[3] $z_{i,t}^{+}$ and $z_{i,t}^{-}$ are defined respectively by

$$z_{i,t}^{+} = \begin{cases} z_{i,t}, & \text{if } z_{i,t} > 0 \\ 0, & \text{if } z_{i,t} \leq 0 \end{cases}$$

and

$$z_{i,t}^{-} = \begin{cases} 0, & \text{if } z_{i,t} > 0 \\ z_{i,t}, & \text{if } z_{i,t} \leq 0. \end{cases}$$

[4] We will add the asymmetric and the squared variables separately to limit the proliferation of parameters. Furthermore, without loss of generality the squared interest rate differential is scaled by the unconditional variance of the interest rate differential to obtain more readable values of the estimated parameter β_2.

where $F(z_{i,t};\gamma)$ as in (5.5). If the presence of the nonlinear specification makes these variables redundant then their parameters should be zero. Therefore we will test if the variables as used by BWZ are statistically different from zero.

Finally, we perform a multivariate analysis where the interdependencies between the returns can be accounted for. We do this by gathering the univariate equations for the several exchange rates (see Equation (5.4)) into the following panel

$$\begin{pmatrix} \Delta s_{GE,t+1} \\ \Delta s_{JP,t+1} \\ \Delta s_{UK,t+1} \end{pmatrix} = \begin{pmatrix} z_{GE,t} \\ z_{JP,t} \\ z_{UK,t} \end{pmatrix} + \begin{pmatrix} \alpha_{GE} + \beta_{1,GE} z_{GE,t} \\ \alpha_{JP} + \beta_{1,JP} z_{JP,t} \\ \alpha_{UK} + \beta_{1,UK} z_{UK,t} \end{pmatrix}$$

$$\circ \begin{pmatrix} F(z_{GE,t};\gamma_{GE}) \\ F(z_{JP,t};\gamma_{JP}) \\ F(z_{UK,t};\gamma_{UK}) \end{pmatrix} + \begin{pmatrix} u_{GE,t} \\ u_{JP,t} \\ u_{UK,t} \end{pmatrix} \tag{5.8}$$

where $F(z_{i,t};\gamma_i)$ as in (5.5) and $\circ$ is the standard Hadamard product where multiplication occurs entry-wise. The errors are assumed to follow a normal distribution with mean zero and covariance Ω

$$\begin{pmatrix} u_{GE,t} \\ u_{JP,t} \\ u_{UK,t} \end{pmatrix} \sim N(0,\Omega)$$

with Ω a 3x3 covariance matrix. Motivation behind this model is that the parameters are estimated more efficiently because the correlation structure between the spot exchange rate returns is taken into account. This could lead e.g. to narrower bands of the confidence bounds for the nonlinear part of the model.

We also try to take advantage of the fact that across the equations in the panel the U.S. dollar interest rate is a common factor in the interest rate differentials. Therefore we will divide the interest rate differential into the local interest rate and the U.S. dollar interest rate. First we allow the parameter for the U.S. dollar interest rate to differ from the parameter of the local interest rate so that we have the following panel

$$\begin{pmatrix} \Delta s_{GE,t+1} \\ \Delta s_{JP,t+1} \\ \Delta s_{UK,t+1} \end{pmatrix} = \begin{pmatrix} z_{GE,t} \\ z_{JP,t} \\ z_{UK,t} \end{pmatrix} + \begin{pmatrix} \alpha_{GE} + \beta_{1,GE} r_{GE,t} - \beta_{2,US(GE)} r_{US,t} \\ \alpha_{JP} + \beta_{1,JP} r_{JP,t} - \beta_{2,US(JP)} r_{US,t} \\ \alpha_{UK} + \beta_{1,UK} r_{UK,t} - \beta_{2,US(UK)} r_{US,t} \end{pmatrix}$$

$$\circ \begin{pmatrix} F(z_{GE,t};\gamma_{GE}) \\ F(z_{JP,t};\gamma_{JP}) \\ F(z_{UK,t};\gamma_{UK}) \end{pmatrix} + \begin{pmatrix} u_{GE,t} \\ u_{JP,t} \\ u_{UK,t} \end{pmatrix} \tag{5.9}$$

where $F(z_{i,t};\gamma_i)$ as in (5.5) and $\beta_{2,US(i)}$ is the parameter for the U.S. dollar interest rate which depends on the foreign currency in country i. Next we will restrict the coefficient of the U.S. dollar interest rate to be the same across the equations in the panel (meaning that $\beta_{2,US(i)} = \beta_{2,US}$ for all $i \in \{\text{Germany}, \text{Japan}, \text{UK}\}$) trying to exploit the interdependencies of the equations caused by this factor such that the panel becomes

$$\begin{pmatrix} \Delta s_{GE,t+1} \\ \Delta s_{JP,t+1} \\ \Delta s_{UK,t+1} \end{pmatrix} = \begin{pmatrix} z_{GE,t} \\ z_{JP,t} \\ z_{UK,t} \end{pmatrix} + \begin{pmatrix} \alpha_{GE} + \beta_{1,GE} r_{GE,t} - \beta_{2,US} r_{US,t} \\ \alpha_{JP} + \beta_{1,JP} r_{JP,t} - \beta_{2,US} r_{US,t} \\ \alpha_{UK} + \beta_{1,UK} r_{UK,t} - \beta_{2,US} r_{US,t} \end{pmatrix}$$

$$\circ \begin{pmatrix} F(z_{GE,t};\gamma_{GE}) \\ F(z_{JP,t};\gamma_{JP}) \\ F(z_{UK,t};\gamma_{UK}) \end{pmatrix} + \begin{pmatrix} u_{GE,t} \\ u_{JP,t} \\ u_{UK,t} \end{pmatrix} \tag{5.10}$$

where $F(z_{i,t};\gamma_i)$ as in (5.5). All models are estimated using Full Information Maximum Likelihood.

5.3 Data

We collected U.S. dollar denominated exchange rates for Germany, Japan, and the U.K. (end-of-month observations) for the period January 1990 to January 2005. For the period from January 1999 the euro-dollar exchange rate is used to obtain a synthetic DMark-dollar rate. 1-Month Eurocurrency rates for the same countries are gathered over the same period (end-of month as well). All data are acquired from Datastream. Because the interest rates are reported on a yearly basis we divide by 12 to obtain the monthly rates as is common practice.

The exchange rates are expressed as the foreign currency values of one U.S. dollar. We let $S_{i,t}$ be the nominal exchange rate for currency i, with $i \in \{\text{Germany}, \text{Japan}, \text{U.K.}\}$.[5] Figure 5.1 shows the exchange rates $S_{i,t}$ and the corresponding returns $\Delta \log S_{i,t}$. Summary statistics for the series are reported in Table 5.1. There does not seem to be a common discernible pattern in the exchange rates, except maybe for the period 2001-2005. Intuitively, the U.S. dollar which is the common currency of these exchange rates can be responsible for this common pattern.

The unconditional correlations are not of the same magnitude across the returns. The correlation between Germany and the U.K. is high while the correlations involving Japan are moderate. The Jarque-Bera statistic is high enough to reject the null of normality.

[5] For notational convenience we don't specify each time that $i \in \{\text{Germany}, \text{Japan}, \text{U.K.}\}$.

	$\Delta \log S_{i,t}$		
	Germany	**Japan**	**U.K.**
Mean	-0.001	-0.002	-0.001
Median	-0.002	0.000	-0.001
Maximum	0.110	0.107	0.131
Minimum	-0.076	-0.155	-0.064
Std. Dev.	0.030	0.033	0.027
Skewness	0.281	-0.601	1.048
Kurtosis	3.756	5.625	6.840
Jarque-Bera	6.644	62.521	143.524
Probability	0.036	0.000	0.000
Observations	180	180	180
	Correlations		
Germany	1	0.436	0.705
Japan		1	0.301
U.K.			1

Table 5.1: Summary statistics for the exchange rate returns $\Delta \log S_{i,t}$ for the period February 1990 to January 2005 for Germany, Japan, and the U.K. (monthly frequency).

We let $r_{i,t}$ and $r_{U.S.,t}$ be the foreign nominal 1-month interest rate and the U.S. nominal 1-month interest rate. Figure 5.2 displays the interest rates, and Figure 5.3 the interest rate differentials with respect to the U.S. dollar interest rate. Table 5.2 reports the summary statistics of the interest rate differentials.

We see a gradual decline of the interest rates over the sample. For the interest rate differentials we note that in the beginning of the nineties relative low interest rates in the U.S. lead to relatively high interest rate differentials. The interest rate differentials decline in the second half of the decade. A second sort of bubble can be seen in the beginning of the new millennium where again the relative value of the U.S. interest rate is low. The correlations between the series are substantially high partly due to possible non-stationarity. Also, when we look at the figures, we see that the interest rate differentials are very persistent which is in contrast with the spot exchange rate returns.

If we compare the nominal exchange rates from Figure 5.1 with the interest rate differentials from Figure 5.3 we can not distinguish a clear pattern for the whole period that would visibly support the UIP hypothesis. For the new millennium it looks like an increase (decrease) in the higher interest rate differential corresponds

	$r_i - r_{US}$		
	Germany	**Japan**	**U.K.**
Mean	0.000	-0.002	0.002
Median	0.000	-0.002	0.002
Maximum	0.005	0.002	0.006
Minimum	-0.002	-0.005	-0.001
Std. Dev.	0.002	0.002	0.002
Skewness	0.688	0.184	0.868
Kurtosis	2.501	1.734	2.719
Jarque-Bera	16.080	13.032	23.190
Probability	0.000	0.001	0.000
Observations	180	180	180
	Correlations		
Germany	1	0.894	0.675
Japan		1	0.844
U.K.			1

Table 5.2: Summary statistics for the interest rate differentials with respect to the U.S. dollar interest rate for the period February 1990 to January 2005 for Germany, Japan, and the U.K. (monthly frequency).

with an increase (decrease) in the exchange rate. However, from the graphs we can not observe a relationship between the exchange rate and the sign of the interest rate differential.

5.4 Results

In this section we report the results of the several models. The results for the model based on Baillie and Kiliç (2006) in Equation (5.4) are reported in Table 5.3. Figure 5.4 shows the value of the transition function $F(.)$ as function of time (left side) and as function of the interest rate differential (right side).

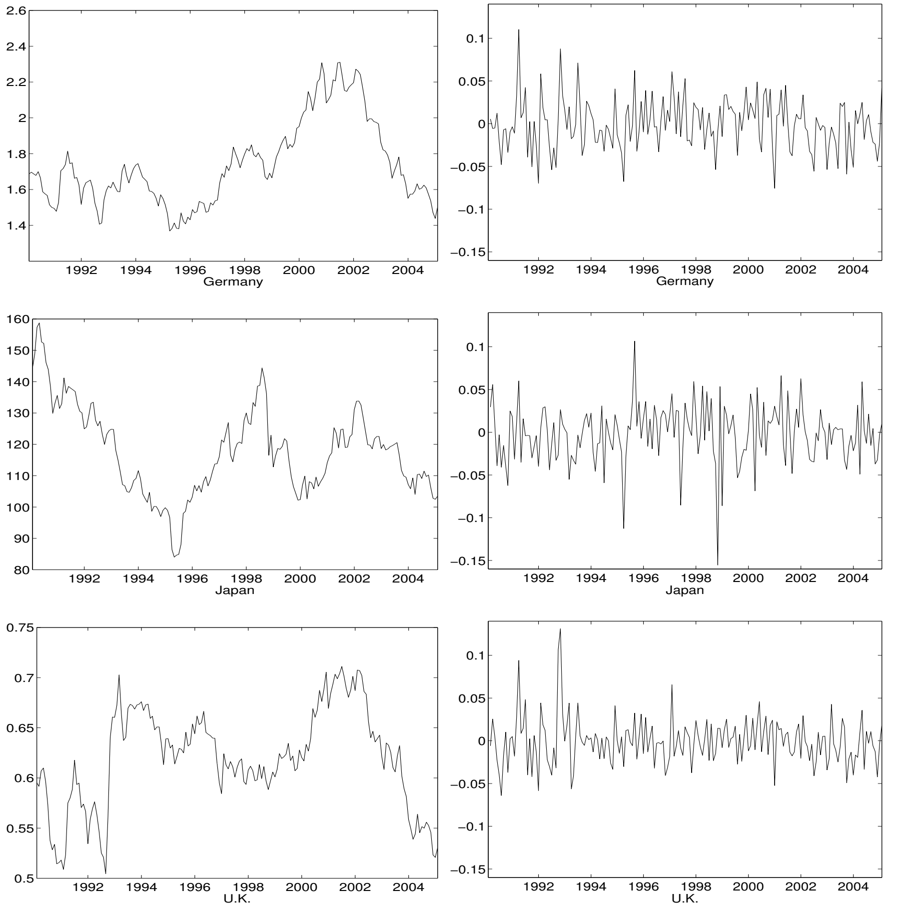

Figure 5.1: Nominal exchange rates and returns. U.S. dollar denominated exchange rates $S_{i,t}$ for the period January 1990 to January 2005 (left) and the corresponding returns $\Delta \log S_{i,t}$ (right) for Germany, Japan, and the U.K. (monthly frequency).

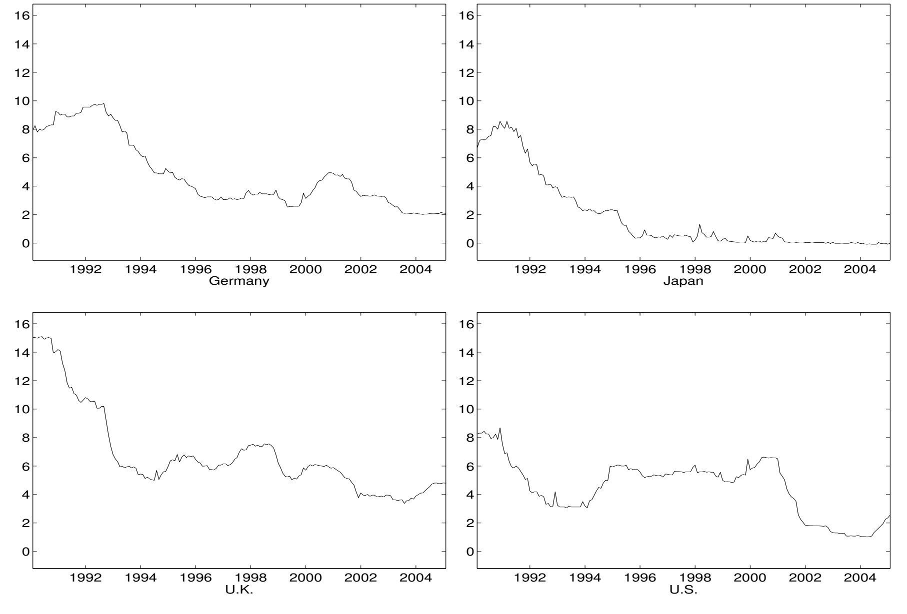

Figure 5.2: Nominal interest rates. Interest rates (annual %) for the period January 1990 to January 2005 for Germany, Japan, the U.K. and the U.S.

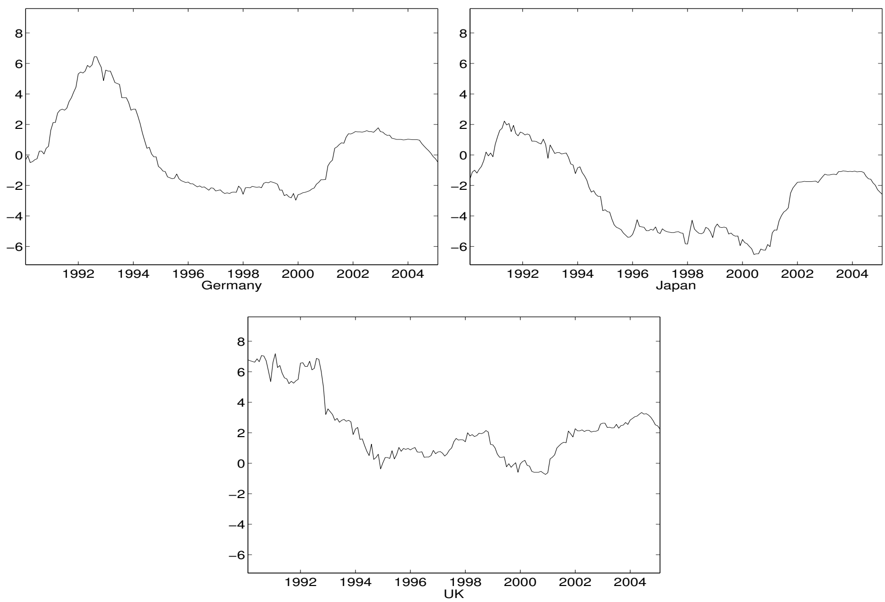

Figure 5.3: Interest rate differentials. Interest rate differentials (annual %) $r_i - r_{US}$ for the period February 1990 to January 2005 for Germany, Japan, and the U.K.

	model Eq. (5.4)			model Eq. (5.6)			model Eq. (5.7)		
	Germany	Japan	U.K.	Germany	Japan	U.K.	Germany	Japan	U.K.
α_1	-0.013*	-0.019	-0.015	-0.018*	-0.020*	-0.015	-0.013*	-0.029	0.002
	(0.008)	(0.016)	(0.011)	(0.011)	(0.010)	(0.012)	(0.008)	(0.040)	(0.005)
β_1	-11.846**	-7.110**	2.570	2.218	11.222*	2.577	-13.369	-15.261	0.638
	(4.725)	(3.507)	(2.389)	(10.179)	(13.613)	(2.393)	(13.433)	(24.858)	(9.534)
β_2				-16.786	-9.768	-931.525	-0.004	-0.006	-0.014
				(12.457)	(5.087)	(9880.393)	(0.031)	(0.016)	(0.020)
γ	-2.217	-0.584	3.937	-0.417	0.586	4.093	-2.544	-1.063	-3.692
	(1.399)	(1.055)	(7.178)	(2.998)	(1.233)	(7.589)	(2.457)	(1.477)	(4.643)
LL	283.46	-497.47	481.28	283.50	-496.94	481.31	283.47	-497.39	481.50
R^2	0.038	0.020	0.008	0.038	0.026	0.009	0.038	0.021	0.011

Table 5.3: The models from Equations (5.4), (5.6) and (5.7) are estimated for Germany, Japan, and the U.K. for the period January 1990 to January 2005. The standard errors are in parentheses and *, ** and *** denote significance at the 10%, 5% and 1% level, respectively.

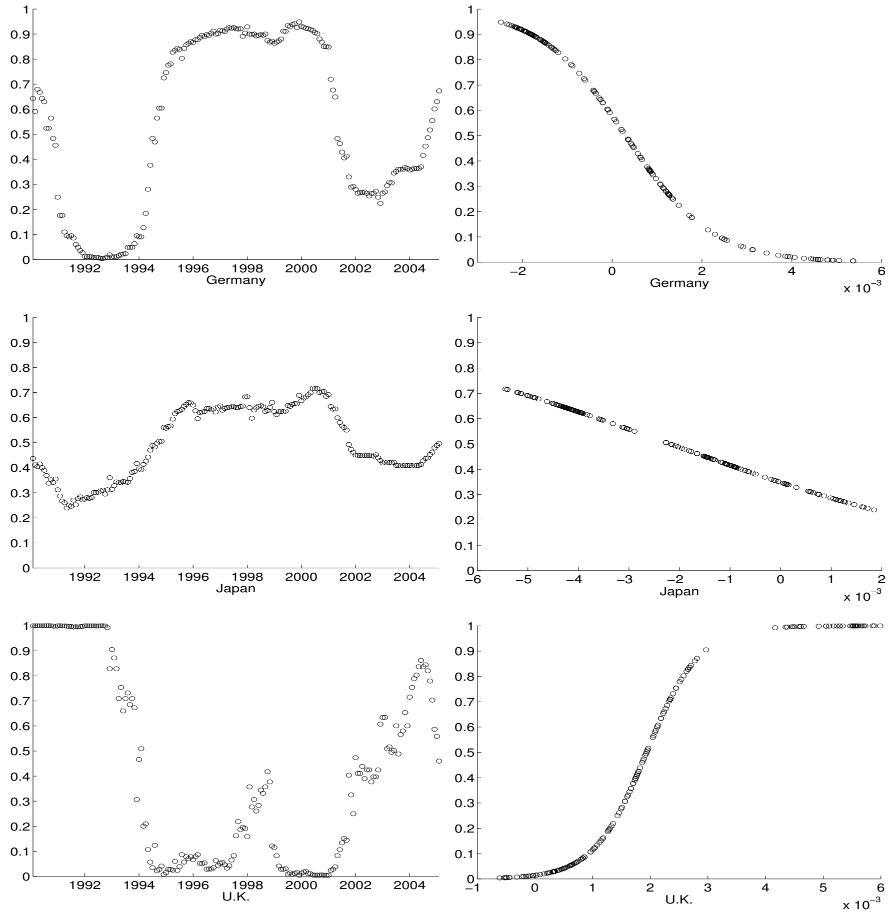

Figure 5.4: Transition function. This figure displays the transition function $F(.)$ as a function of time (left graphs) and of the interest rate differential (right graphs) for Germany (top graphs), Japan (middle graphs) and the U.K. (bottom graphs) of the BK06 model in Equation (5.4).

The intercept and slope coefficient in the nonlinear part are significantly different from zero at the 10% and 5% level respectively for Germany. For Japan only the slope parameter is significant (at 5%) while for the U.K. no parameters are significant. So mixed evidence for the constant and the interest rate differential being important in the nonlinear part is found. From the graphs in Figure 5.4 we see that the transition function recognizes different states, especially for Germany and the U.K.. For Japan no clear states can be identified. To test for which periods UIP holds, the significance of the nonlinear part of the model is determined by calculating confidence bounds using the Delta method (see the Appendix for a derivation of these bounds) which are displayed in Figure 5.5. Here we see that for Germany and Japan for a few periods the UIP hypothesis can be rejected. For the larger part UIP can not be rejected. For the U.K. we do not reject UIP at all. So although it seems that there is some evidence that the interest rate differential is important in the nonlinear part, hardly any evidence for rejection of the UIP exists.

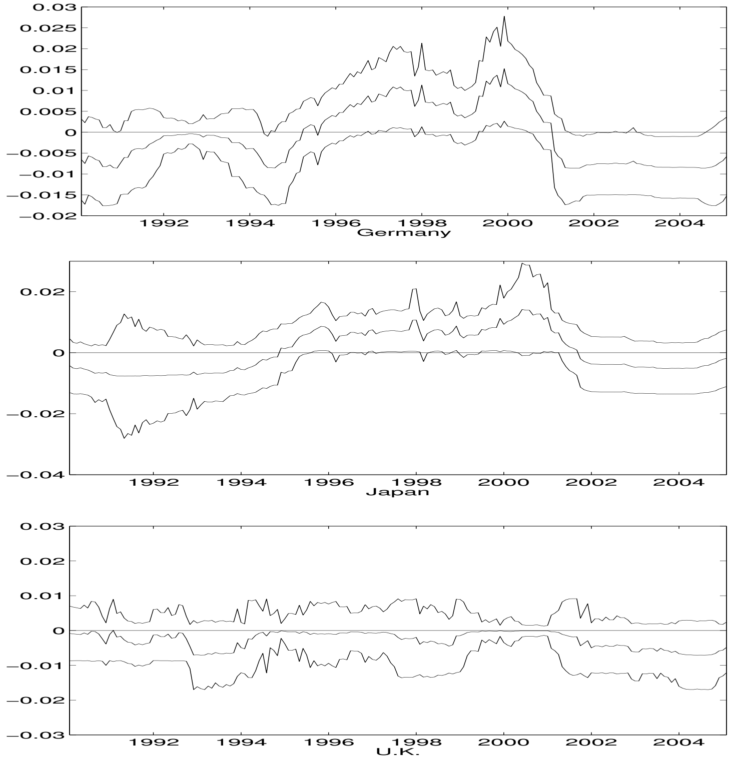

Figure 5.5: Deviations from UIP This figure displays the deviations from UIP (monthly %) which are measured by the nonlinear part of Equation (5.4) in the univariate model. The 95% confidence are calculated as described in the Appendix.

We have also estimated the model with the positive and negative interest rate differentials (see Equation (5.6)) and with the second moment (see Equation (5.7)) added in the nonlinear specification. The results are reported in Table 5.3). We see that β is only significantly different from zero at the 10% level for Japan. Also the second order moment of the interest rate differential is for none of the countries significant so this variable does not improve the performance of our model either. So in general we can not find the added value for these variables in the nonlinear part as opposed to their significance in the linear model according to Bansal (1997) and Wu and Zhang (1996). Therefore it seems that the use of the model of BK06 makes the variables as advocated by BWZ redundant but we note here that the model of BWZ is not nested in the model of BK06-BWZ so we are not able to test this proposition.

Next we gather the equations into a panel. An unrestricted panel is estimated first (see Equation (5.8)). In the second panel the interest rate differential is split so that the influence of the underlying interest rates can be assessed (see Equation (5.9)). In the last panel we estimate a common coefficient for the U.S. interest rate across the equations in the panel (see Equation (5.10)). The results for these panels are displayed in Table 5.4. For the nonlinear part of the model of each equation of the last panel we also calculated the confidence bounds which are displayed in Figure 5.6.

	model Eq. (5.8)			model Eq. (5.9)			model Eq. (5.10)		
	Germany	Japan	U.K.	Germany	Japan	U.K.	Germany	Japan	U.K.
α_1	-0.004	-0.013	-0.009	-0.019	-0.020	-0.032	-0.015	-0.019	-0.019
	(0.005)	(0.011)	(0.008)	(0.015)	(0.017)	(0.022)	(0.014)	(0.014)	(0.019)
β_1	-8.151***	-6.620**	1.218	-8.712**	-6.290**	0.767	-5.766**	-6.285**	-4.136
	(2.770)	(2.774)	(1.735)	(3.754)	(2.986)	(5.528)	(2.863)	(2.713)	(3.720)
β_2				-11.926***	-8.368*	-6.240	-8.744***		
				(4.114)	(4.328)	(3.954)	(3.080)		
γ	-0.736*	0.046	17.627	-1.016**	0.009	-2.606	-0.561	0.370	-2.399
	(0.413)	(1.090)	(61.465)	(0.493)	(1.093)	(3.367)	(0.427)	(0.944)	(2.475)
LL	346.89			349.31			348.04		
R^2	0.025	0.017	0.005	0.026	0.017	0.018	0.022	0.014	0.014

Table 5.4: The models from Equation (5.8), (5.9) and (5.10) are estimated for the period January 1990 to January 2005. Results are reported in the left, middle and right panel respectively. The standard errors are in parentheses and *, ** and *** denote significance at the 10%, 5% and 1% level, respectively.

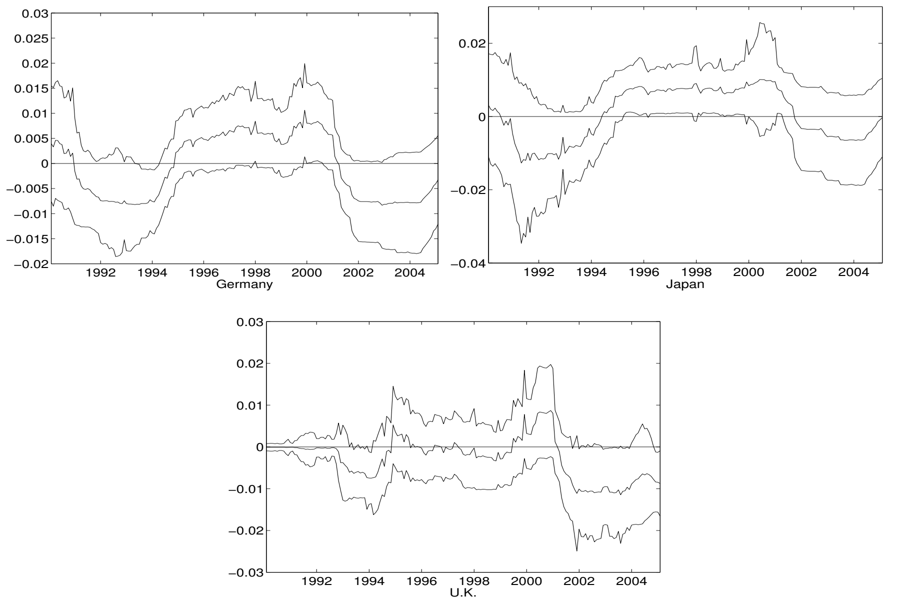

Figure 5.6: Deviations from UIP This figure displays the deviations from UIP (monthly %) which are measured by the nonlinear part of Equation (5.4) in the multivariate model. The 95% confidence are calculated as described in the Appendix.

For the unrestricted panel in Equation (5.8) the β's for Germany and Japan are significant different from 0 at the 1% and 5% respectively. Also the β for Germany and Japan are significant (all at 5% level) for the non-U.S. interest rate for the model in Equation (5.9) and Equation (5.10). For the U.K. we find insignificant β's for all models. The β's for the U.S. interest rate in Equation (5.9) are significant for Germany (1%) and Japan (10%). We find a significant effect of the U.S. interest rate when its parameter is restricted to be the same across the equations (Equation (5.10)). After determination of the confidence bounds of the nonlinear part of the model (see Figure 5.6) we come to the conclusion that the nonlinear part is not significantly different from zero most of the time so that the UIP hypothesis can not be rejected when we apply this model which is similar to the conclusion based on Figure 5.5.

5.5 Concluding remarks

After the general finding by many studies that the linear model of Fama (1984) is unable to validate the UIP theory and the theoretical motivation that the relationship between the spot exchange rate return and the forward premium (or equivalently, the interest rate differential) might not be linear due to e.g. transaction costs or limitations to speculation, some academics considered the possibility of nonlinear models.

In this chapter our contribution to the UIP literature is threefold. First, rewriting the nonlinear model of Baillie and Kiliç (2006) we are able to measure and test deviations from UIP. In this model we impose UIP while we allow for possible nonlinear deviations. Confidence bounds for these deviations are constructed so that the validity of the UIP hypothesis can formally be tested. Second, this model is extended to incorporate the findings of Bansal (1997) and Wu and Zhang (1996) who use a linear model to find that asymmetric and nonlinear transformations of the interest rate differential are of importance in explaining the behavior of the spot exchange rate return. The question is then if the nonlinearity of the model dominates these variables or that they remain important. Third, we apply a multivariate approach to test the UIP hypothesis so that possible interdependencies between the exchange rates and interest rates can be accounted for. Especially we try to exploit the fact that there is a common U.S. dollar component in the interest rate differential.

The main findings of this chapter are that the confidence bounds for the nonlinear part are too broad such that the deviations from UIP are not statistically different from zero. This means that the UIP hypothesis can not be rejected most of the time. Exploiting the common U.S. interest rate across the exchange rates in the multivariate approach does not lead to different conclusions. Further research has to be done to investigate whether this result is due to the validity of the UIP hypothesis or to model misspecification or parameter uncertainty.

The variables used by Bansal (1997) and Wu and Zhang (1996) are not significant in our model which leads to the conclusion that it seems that the nonlinear specification of the model makes the former variables redundant. Stated differently: the nonlinear model seems able to absorb the effects that are picked up by the linear models of Bansal (1997) and Wu and Zhang (1996) but this is not tested more formally because the model by Bansal (1997) is not nested in our adaptation of the model of Baillie and Kiliç (2006).

Although we find that a nonlinear relationship between the spot exchange rate return and the forward premium is supported theoretically and to a lesser extent also empirically, much has to be done in this area of research. One suggestion would be to turn to e.g. behavioral finance or introduce learning by using a Bayesian approach to explain the possible nonlinearity. Another suggestion would be to extend the nonlinear framework to incorporate macro-economic information.

Appendix

In this appendix we will give a derivation of the confidence bounds for the nonlinear part of Equation 5.4. We make use of the Delta method (see for example Oehlert (1992)) where a Taylor expansion is constructed for the nonlinear part around its mean for which the variance can be calculated. The confidence bounds are then calculated by adding and subtracting 2 times the standard deviation to the estimated function.

The nonlinear part of the model is as follows

$$H(\alpha, \beta, \gamma; z_{i,t}) = [\alpha + \beta z_{i,t}]F(z_{i,t}; \gamma)$$

Next a first-order Taylor expansion is made about the estimated value of the function

$$H(\alpha, \beta, \gamma; z_{i,t}) = H(\hat{\alpha}, \hat{\beta}, \hat{\gamma}; z_{i,t}) + \begin{pmatrix} \alpha - \hat{\alpha} \\ \beta - \hat{\beta} \\ \gamma - \hat{\gamma} \end{pmatrix}^T H'(\hat{\alpha}, \hat{\beta}, \hat{\gamma}; z_{i,t}),$$

where H' contain the partial first derivatives of H with respect to the parameters. This vector is evaluated at the estimated parameters

$$H'(\hat{\alpha}, \hat{\beta}, \hat{\gamma}; z_{i,t}) = \begin{pmatrix} F(z_{i,t}; \hat{\gamma}) \\ z_{i,t} F(z_{i,t}; \hat{\gamma}) \\ \frac{(z_t - \mu_z)\exp(-\hat{\gamma}(z_t - \mu_z)/\sigma_z)}{\sigma_z[1+\exp(-\hat{\gamma}(z_t - \mu_z)/\sigma_z)]^2} \end{pmatrix}.$$

Subsequently using the variance-covariance matrix of the estimated parameters, $Var([\hat{\alpha}, \hat{\beta}, \hat{\gamma}])$, we can approximate the scalar variance of $H(.)$ by

$$Var(H(\hat{\alpha}, \hat{\beta}, \hat{\gamma}; z_{i,t})) = H'(\hat{\alpha}, \hat{\beta}, \hat{\gamma}; z_{i,t})^T Var([\hat{\alpha}, \hat{\beta}, \hat{\gamma}]) H'(\hat{\alpha}, \hat{\beta}, \hat{\gamma}; z_{i,t})$$

so that the 95% - confidence bounds are approximated by

$$H(\hat{\alpha}, \hat{\beta}, \hat{\gamma}; z_{i,t}) \pm 1.965\sqrt{Var(H(\hat{\alpha}, \hat{\beta}, \hat{\gamma}; z_{i,t}))}$$

Part III

Exchange Rate Volatility

Chapter 6

A Range-Based Multivariate Stochastic Volatility Model for Exchange Rates[1]

6.1 Introduction

The analysis of exchange rate volatility is an important topic in international economics. For example, exchange rate volatility plays a crucial role in measuring the risks in international trade and portfolio flows. In the extant volatility literature there are two main streams of modelling methodologies for measuring the dynamics of volatility: the (G)ARCH and the Stochastic Volatility (SV) approaches. The former is based on the seminal articles by Engle (1982) and Bollerslev (1986) and the SV approach started with Taylor (1986)). From a risk management point of view it is often warranted to take a multivariate perspective on volatilities. Both for the GARCH and SV approaches multivariate versions exist as presented in, among many other papers, Engle and Kroner (1995), Harvey *et al.* (1994), Liesenfeld and Richard (2003), Chib *et al.* (2005) and McAleer (2005).[2]

In this chapter we develop a multivariate SV model to measure exchange rate volatilities of the main traded currencies. Although GARCH models are typically easier to estimate than SV models, we choose for the latter model. It is well-documented in the literature, see for example Alizadeh *et al.* (2002), that suitable transforma-

[1] This chapter is based on Tims and Mahieu (2006).

[2] See also the special issue of Econometric Reviews (Volume 25, Issues 2 and 3) on multivariate stochastic volatility for recent advances in this field.

tions of the difference between the high and the low result in stochastic variables that are normally distributed. This observation implies that we can circumvent the typical estimation difficulties of SV models and allows us to efficiently estimate a multivariate model for exchange rate volatilities.[3]

Many variables have been suggested to measure volatility.[4] For example, in the GARCH approach the conditional volatility process typically contains variables based on squared or absolute returns. When high-frequency data are available Realized Volatility (RV) seems to be an appropriate measure of volatility, as recent studies show (see for example, Andersen *et al.* (2001b) and Andersen *et al.* (2001a)). Aït-Sahalia (2002), Aït-Sahalia and Mykland (2003) and Aït-Sahalia *et al.* (2005) use the temporal aggregation methods of Andersen *et al.* (2001b) and Andersen *et al.* (2003) to reduce the bias in estimating RV models from high frequency data. However, when high-frequency data is not available other volatility measures can be considered. In this chapter we use the logarithm of the range as a proxy for volatility, following Parkinson (1980) and Alizadeh *et al.* (2002). The logarithmic range is defined as the logarithm of the difference between the logarithmic high and logarithmic low measured over a particular time horizon. This measure is an accurate proxy for the volatility (see Andersen and Bollerslev (1998), and the discussion in Alizadeh *et al.* (2002)). The main advantage of using the logarithmic range as a (log) volatility proxy is that its empirical unconditional distribution for exchange rates is approximately normal.

In order to build a multivariate model for exchange rate volatilities based on logarithmic ranges, the issue of covariance measurement and estimation needs to be resolved. There is no straightforward way to construct a covariance estimator using information on the highs and lows from the two series. One solution is to apply the no-arbitrage methods in Brandt and Diebold (2006). They use the triangular relationship between exchange rates denominated in a particular currency and their cross rates. In this chapter we follow another route and decompose exchange rates into independent currency-specific factors following Mahieu and Schotman (1994). As a result the variance of the exchange rate returns is the sum of the two currency-specific variances, which leads to a parsimonious multivariate volatility model for exchange rates.[5]

[3]Standard SV models are difficult to estimate due to the fact that the latent volatility series need to be integrated out from the objective function. See for example Broto and Ruiz (2004).

[4]See for example Andersen (1992), and Bai *et al.* (2001) for an exposition on measuring volatilities.

[5]We also estimated a model that allowed for a common currency component for all exchange rates that can pick up for example the consequences of international risk sharing (see e.g. Brandt *et al.* (2005)) in Section 6.6.

The remainder of this chapter is structured as follows. The exchange rate data is discussed in Section 6.2. In Section 6.3 the multivariate stochastic volatility model is presented and Section 6.4 presents the news decomposition method. Section 6.5 documents the results of our empirical studies followed by Section 6.6 where some robustness checks on the model are performed. Section 6.7 concludes.

6.2 Exchange rate data

The daily high, low and close prices of six exchange rates were collected from Moneyline Telerate. The high and low values are computed over a 24-hour period which starts at 10pm GMT, 5pm US Eastern Time. The exchange rate set consists of all possible combinations of the following currencies: US dollar (USD), UK sterling (GBP), Japanese yen (JPY) and euro (EUR).[6] The sample runs from September 1, 1989 until September 16, 2003. Before January 1, 1999 the Deutsche mark is used as a proxy for the euro. For illustrative purposes, the log range and the QQ plot for the USD/EUR are presented in Figure 6.1.[7] The time-varying nature of the log range can be directly observed from the figure. From the QQ plot it is apparent that the unconditional distribution of the log range is very close to a normal distribution. The unconditional covariance and correlation matrices for the log range series (so the volatility of the volatility) appear in Table 6.1. As was already noticed in Mahieu and Schotman (1994) the covariance structure of exchange rates expressed in a common numeraire has a very distinct feature. The covariances of all dollar exchange rate returns are all in the same order of magnitude. This observation motivated Mahieu and Schotman (1994) to construct a multivariate exchange rate model exploiting this feature. Building upon their framework, we will model the volatility directly using the range.[8]

[6] The FX trade flows (volumes) are highest for these currencies and are responsible for more than 80% of the total foreign exchange market turnover (see Bank for International Settlements (2001)). Extending the data set with the fifth most important FX rate (the Swiss franc, which contributes for about 3% to the total turnover) does not lead to significantly different results in our analysis which will be shown in Section 6.6.

[7] For the other exchange rates these figures are available upon request.

[8] Note that the model of Mahieu and Schotman (1994) is estimated using squared returns as a proxy for volatility.

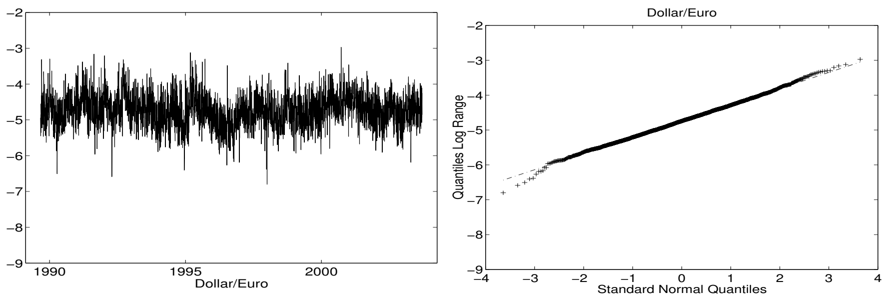

Figure 6.1: Log range and corresponding QQ-plot. Log range of the dollar/euro exchange rate (left) and QQ-plot for the log range of the dollar/euro exchange rate. See Section 6.2 for further details. The period is September 1st, 1989 - September 16th, 2003 (3652 observations).

Table 6.1: Unconditional covariance and correlation matrix. Covariance and correlation matrix (unconditional) of the log range of the exchange rates dollar/pound, dollar/yen, dollar/euro, pound/yen, pound/euro and yen/euro are reported for the period September 1st, 1989 - September the 16th, 2003 (3652 observations).

		covariance matrix					
		(USD/GBP)	(USD/JPY)	(USD/EUR)	(GBP/JPY)	(GBP/EUR)	(JPY/EUR)
dollar/pound	(USD/GBP)	0.356	0.066	0.169	0.108	0.090	0.091
dollar/yen	(USD/JPY)		0.279	0.099	0.165	0.071	0.160
dollar/euro	(USD/EUR)			0.217	0.082	0.104	0.121
pound/yen	(GBP/JPY)				0.220	0.107	0.169
pound/euro	(GBP/EUR)					0.202	0.100
yen/euro	(JPY/EUR)						0.265
		correlation matrix					
		(USD/GBP)	(USD/JPY)	(USD/EUR)	(GBP/JPY)	(GBP/EUR)	(JPY/EUR)
dollar/pound	(USD/GBP)	1	0.209	0.610	0.387	0.336	0.298
dollar/yen	(USD/JPY)		1	0.401	0.666	0.301	0.589
dollar/euro	(USD/EUR)			1	0.378	0.497	0.504
pound/yen	(GBP/JPY)				1	0.510	0.698
pound/euro	(GBP/EUR)					1	0.433
yen/euro	(JPY/EUR)						1

6.3 Multivariate stochastic volatility model

In this section we present our multivariate stochastic volatility model for exchange rates. We impose a similar factor structure on the exchange rate volatility as is done for exchange rate returns in Mahieu and Schotman (1994). In that paper the model is structured as follows. Let $s_{ij}(t)$ is the exchange rate return between currencies i and j. We assume that it can be decomposed into two currency-specific components

$$s_{ij}(t) = e_i(t) - e_j(t). \tag{6.1}$$

The components $e_i(t)$ and $e_j(t)$ can be interpreted as being a representation of the currency-specific news of currencies i and j, respectively. We assume that the news components are independent.[9] If we define $\lambda_i(t)$ as the variance of the news factor $e_i(t)$, $\lambda_i(t) \equiv var[e_i(t)]$ then the variance of the exchange rate can be written as

$$var\left[s_{ij}(t)\right] = var\left[e_i(t)\right] + var\left[e_j(t)\right] = \lambda_i(t) + \lambda_j(t), \tag{6.2}$$

applying the independence assumption between news factors. We use the idea of this setup as the basis for our multivariate volatility model. The range-based volatility measure applies to the logarithmic volatility. Let $y_{ij,t}$ be the logarithmic range for the exchange rate between currencies i and j. We assume that we can decompose the logarithmic range into two independent factors that relate to the two currencies. Consequently, we assume that

$$y_{ij,t} = \alpha_{it} + \alpha_{jt}, \tag{6.3}$$

with α_{it} and α_{jt} two latent factors. In our empirical application we focus on 4 currencies, the U.S. Dollar (USD), the British Pound Sterling (GBP), the Japanese Yen (JPY) and the euro (EUR). This implies that we can construct 6 exchange rates for which we can compute the logarithmic ranges. Note again that a larger number of currencies or a larger number of factors (e.g. a world effect, a region effect, and extreme events) can be accommodated straightforwardly.

The model with the currency-specific factors is presented below where the log ranges of all possible exchange rates of the 4 currencies under consideration are used. For notational purposes we collect the log range at time t for all exchange rates in

[9] We could introduce a more elaborate factor structure here. For example, it is relatively straightforward to include a "world" component that would represent news that applies to all exchange rates (see Section 6.6), or components that would apply to various subsets of exchange rates. For notational reasons we restrict ourselves to the decomposition of exchange rate returns as described in the text.

the vector y_t $(t = 1, \ldots, n)$. The model that we estimate is given by the following state space equations.

$$y_t = c + Z\alpha_t + \varepsilon_t, \qquad \varepsilon_t \sim N(0, H), \qquad t = 1, \ldots, n \tag{6.4}$$

$$\alpha_{t+1} = T\alpha_t + \eta_t, \qquad \eta_t \sim N(0, Q), \qquad t = 1, \ldots, n \tag{6.5}$$

where

$$y_t = \begin{pmatrix} y_{USD,GBP_t} \\ y_{USD,JPY_t} \\ y_{USD,EUR_t} \\ y_{GBP,JPY_t} \\ y_{GBP,EUR_t} \\ y_{JPY,EUR_t} \end{pmatrix},$$

and

$$Z = \begin{pmatrix} 1 & 1 & 0 & 0 \\ 1 & 0 & 1 & 0 \\ 1 & 0 & 0 & 1 \\ 0 & 1 & 1 & 0 \\ 0 & 1 & 0 & 1 \\ 0 & 0 & 1 & 1 \end{pmatrix},$$

$$\alpha_t = \begin{pmatrix} \alpha_{USD_t} \\ \alpha_{GBP_t} \\ \alpha_{JPY_t} \\ \alpha_{EUR_t} \end{pmatrix}.$$

The idea behind the model is to decompose the log range of each exchange rate into a constant, the corresponding currency-specific factors (using Z as a selection matrix) and an error term. There are no restrictions on H, the covariance matrix of the measurement errors. The latent factors evolve over time according to 4 univariate AR(1) processes which are assumed to be independent. Consequently, both T (the autoregressive parameter matrix) and Q (covariance matrix of the error terms) in the transition equations are diagonal.

In general it is not straightforward to estimate stochastic volatility models (see for example Jacquier *et al.* (1994) and Broto and Ruiz (2004)). A problem with standard Maximum Likelihood (ML) approaches lies in the fact that the state-space model cannot be estimated directly with standard Kalman filter methods as either the measurement or the latent state equation exhibits nonlinearities. Also the joint distribution of the return and the latent volatility is highly dimensional. Altogether this means that the likelihood function is not tractable in many cases and therefore cannot be optimized directly. Estimation of the parameters by maximum likelihood implies that the latent volatility needs to be integrated out of the log likelihood by using numerical simulation techniques. See, for example, Jacquier *et al.* (1994), Shephard (1993), and Chib *et al.* (2005). Another obvious solution to deal with the nonlinearity in the state space model is to linearize the SV model to obtain a standard linear state space model and apply Quasi Maximum Likelihood (QML) together with standard Kalman filter techniques to estimate the model (see e.g. Harvey *et al.* (1994) and Mahieu and Schotman (1998)). Alternative procedures for estimating SV models can be found in Shephard (1993), Jacquier *et al.* (1994), and Andersen and Sørensen (1996).

The main contribution of our new approach to estimate exchange rate volatilities is that we do not have to resort to either numerical integration or linearization to estimate the parameters and the latent volatilities. This is caused by the fact that we specify our multivariate model directly on the logarithmic ranges, which have empirical distributions that are shown to be very close to the normal distribution. As a result we can maintain a linear state space model and, consequently we can construct the likelihood function by applying the standard Kalman filter.

Instead of using a numerical procedure like gradient methods we apply the EM algorithm (Dempster *et al.* (1977)) in order to optimize the likelihood function. The main reason for this choice is that the EM algorithm allows us to estimate the parameters in the measurement error covariance matrix H in a straightforward way, without having to resort to adding restrictions in the optimization procedure in order to assure that H remains positive definite.

We estimate the model (6.4)-(6.5) by iterating between estimation of the latent factors given the parameters using the Kalman filter and maximization of the likelihood function given the latent factors using analytical expressions for the optimal parameters. By iterating these steps Dempster *et al.* (1977) have shown that the likelihood function increases. This iterative procedure can be continued until convergence takes place. In the Appendix we present a detailed derivation of the algorithm.

6.4 News decomposition of exchange rates

In this section we demonstrate how to estimate currency-specific news factors based on the estimation results of our multivariate model for exchange rate volatilities. Remember that every exchange rate can be decomposed into two currency specific factors as stated in Equation (6.1). Mahieu and Schotman (1994) have shown that such a currency-specific news factor is a weighted average over the exchange rate returns containing that currency. The weights are constructed from the conditional variances (λ's) of the news factors. The intuition behind this is that the currency specific news for one particular currency can be extracted from all exchange rates that can contain that currency. The higher the conditional volatility of the exchange rate, the less informative this exchange rate is, and therefore this exchange rate has less weight in the news factor. Following Mahieu and Schotman (1994), the news factor at time t, $\hat{e}(t)$, for currency i with $i \in I = \{USD, GBP, JPY, EUR\}$ can be written as

$$\hat{e}_i(t) = \frac{\sum\limits_{j \in I, j \neq i} \lambda_j(t)^{-1} s_{ij}(t)}{\sum\limits_{j \in I} \lambda_j(t)^{-1}} \tag{6.6}$$

where $s_{ij}(t)$ is the continuously compounded return of the exchange rate. Also note that $s_{ji}(t) = -s_{ij}(t)$. To apply this we again note that in our model the logarithmic volatility is modelled instead of the volatility itself. So to obtain the equivalent quantity of the $\lambda(t)$'s in Equation (6.6) we have to transform the estimated time-varying $\hat{\alpha}_t$ that we obtained from the estimation procedure of the multivariate volatility model. Note that $\hat{\alpha}_t = E[\alpha_t | y]$, i.e. the expected logarithmic volatilities given all the data y. This can be computed by applying the Kalman smoother equations (see Appendix). Then we obtain[10]

$$\lambda_j(t) \propto E(\exp(\alpha_{j,t})^2 | y) = \exp(2\hat{\alpha}_{j,t} + 2V_{j,t}) \tag{6.7}$$

where

$$V_{j,t} = Var(\alpha_{j,t} | y).$$

To investigate the relative impact of each individual news series from (6.6) on the underlying exchange rate, we construct an index for each of the news series $\hat{e}_i(t)$. The news indices are defined by

$$I_{i,news}(t) = 100(1 + \sum_{i=1}^{t} \hat{e}_i(t))$$

In the next section we present the news series computed from the multivariate volatility model on our data set of exchange rates.

[10] Here we use the fact that when $X \sim N(\mu, \sigma^2)$ the moment generating function equals $E[\exp(tX)] = \exp(t\mu + t^2\sigma^2/2)$ and that the Kalman smoother equations also deliver the conditional variances of the state vector α_t (see the Appendix for more details).

6.5 Results

In this section the results of the model will be presented.[11] The estimated parameters are given in Table 6.2. We see that part of the variability in the log range series is picked up by the model as the estimated covariance matrix of the errors ($\hat{H}$) has uniformly smaller elements than the covariance matrix of the data itself, which was presented in Table 6.1. Additionally, the correlation matrix constructed from $\hat{H}$ in Table 6.2 shows that, on average, the correlations have decreased with respect to the original data. The latent variables are assumed to follow an AR(1) process and we see that the autoregressive parameters ($\hat{T}$) are close to one. This points to persistent latent logarithmic volatilities, which is a well-known feature of financial time series measured on a daily frequency.

[11] All calculations needed for construction of the graphs and tables are done with Matlab, version 6.

Table 6.2: Parameter estimates of the model. Estimated parameter values from the model (6.4)-(6.5). Standard errors are in parentheses. The correlation matrix of the measurement errors (printed in italics) is constructed from $\hat{H}$. Estimation was performed over the period September 1st, 1989 - September 16th, 2003 (3652 observations).

Exchange rate		$\hat{c}$	$\hat{H}$					
			(USD/GBP)	(USD/JPY)	(USD/EUR)	(GBP/JPY)	(GBP/EUR)	(JPY/EUR)
dollar/pound	(USD/GBP)	-5.0521	0.2570	0.0318	0.1057	0.0830	0.0510	0.0798
		(0.0233)	(0.0073)	(0.0045)	(0.0042)	(0.0061)	(0.0042)	(0.0042)
			1	*0.1469*	*0.5304*	*0.4457*	*0.2965*	*0.3810*
dollar/yen	(USD/JPY)	-4.7732		0.1827	0.0594	0.0855	0.0385	0.0729
		(0.0239)		(0.0044)	(0.0035)	(0.0036)	(0.0041)	(0.0025)
				1	*0.3533*	*0.5451*	*0.2655*	*0.4128*
dollar/euro	(USD/EUR)	-4.7442			0.1547	0.0500	0.0570	0.0766
		(0.0230)			(0.0044)	(0.0041)	(0.0028)	(0.0039)
					1	*0.3462*	*0.4274*	*0.4712*
pound/yen	(GBP/JPY)	-4.6125				0.1348	0.0423	0.0877
		(0.0241)				(0.0029)	(0.0029)	(0.0034)
						1	*0.3398*	*0.5780*
pound/euro	(GBP/EUR)	-4.9199					0.1151	0.0280
		(0.0222)					(0.0036)	(0.0031)
							1	*0.1999*
yen/euro	(JPY/EUR)	-4.6913						0.1709
		(0.0237)						(0.0050)
								1
Currency	$diag(\hat{T})$	$diag(\hat{Q})$						
dollar	0.9639	0.0021						
	(0.0119)	(0.0007)						
pound	0.9675	0.0015						
	(0.0075)	(0.0003)						
yen	0.9609	0.0047						
	(0.0121)	(0.0016)						
euro	0.9415	0.0035						
	(0.0192)	(0.0013)						
$\log L$	-9167.173							

The graphs of the variances of the news factors $\lambda_j(t)$ $(t = 1, \ldots, n)$ are shown in Figure 6.2. It is clear that the yen component is the most volatile of all. Interestingly, the period in the late nineties where the yen is very volatile corresponds with the Asian crisis. The volatility in the other currencies remains relatively stable, implying that the volatility in exchange rates containing the yen are mostly due to the yen-specific volatility. For the pound volatility component the largest increase in the volatility corresponds to the departure of the UK from the European Monetary System (EMS) on September 17, 1992. Surprisingly the dollar volatility component does not show a distinct pattern.

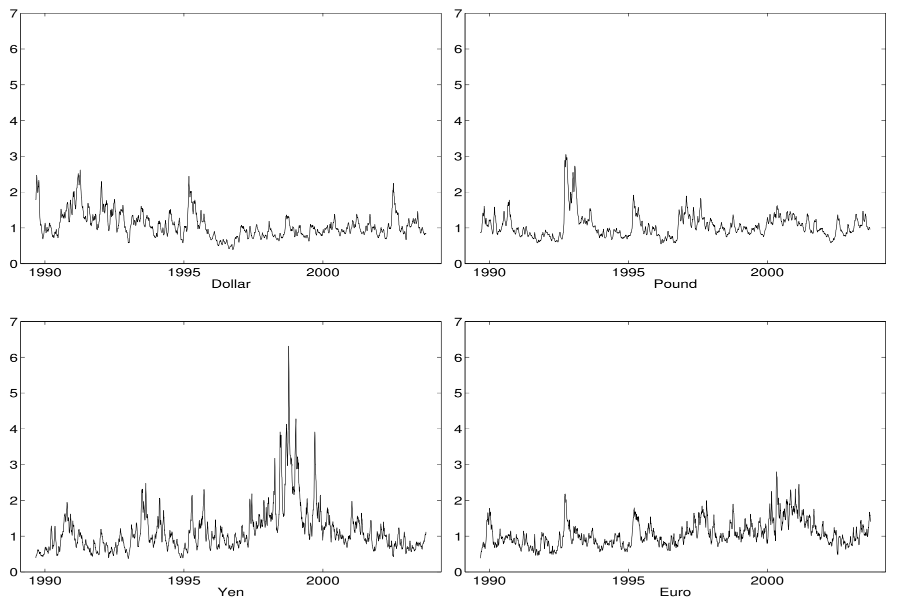

Figure 6.2: Estimated λ's. Graphs of estimated λ's measured in squared daily percentages as defined in Equation (6.7) over the period September 1st, 1989 - September 16th, 2003 (3652 observations).

To focus on the influence of the news components after the introduction of the euro Figure 6.3 show the indexed exchange rates together with the corresponding news indices from January 1, 1999 onwards. To accommodate comparison we have shifted the lines of the news factors downwards. From the figure it can be seen that particular movements in the exchange rates can be assigned to specific currencies. This is most apparent for the period January 1999 until about January 2001 where we see that the euro exchange rates (USD/EUR, GBP/EUR, and JPY/EUR) first rise

Table 6.3: Correlations between exchange rates returns and returns on news components. Correlations between returns on the euro exchange rates and on the news components of the underlying currencies are given over the period January 1st, 1999 - December 31st, 2000.

		News components			
		Dollar	Pound	Yen	Euro
dollar/euro	(USD/EUR)	0.762			0.914
pound/euro	(GBP/EUR)		0.541		0.804
yen/euro	(JPY/EUR)			0.858	0.806

quickly and then fall quickly which is for the USD/EUR and GBP/EUR exchange rates almost totally due to the euro news component (see Table 6.3). We also see that movements in the exchange rates occur simultaneously in both currency-specific news factors.

6.6 Sensitivity analysis

In this section we perform some sensitivity analysis of the results on the model specification.

6.6.1 Adding a world factor

The orthogonality restriction we imposed on the news factors in Equation (6.1) may be too strict. To test this we will estimate another specification and add a so-called world factor to the transition equation of the model. The world factor is assumed to impact all currency pairs simultaneously. Therefore the log range of each exchange rate is now decomposed into two currency-specific components and one world component which is the same across all exchange rates. The new selection matrix Z and the state vector α_t can then be written as

$$Z = \begin{pmatrix} 1 & 1 & 0 & 0 & 1 \\ 1 & 0 & 1 & 0 & 1 \\ 1 & 0 & 0 & 1 & 1 \\ 0 & 1 & 1 & 0 & 1 \\ 0 & 1 & 0 & 1 & 1 \\ 0 & 0 & 1 & 1 & 1 \end{pmatrix},$$

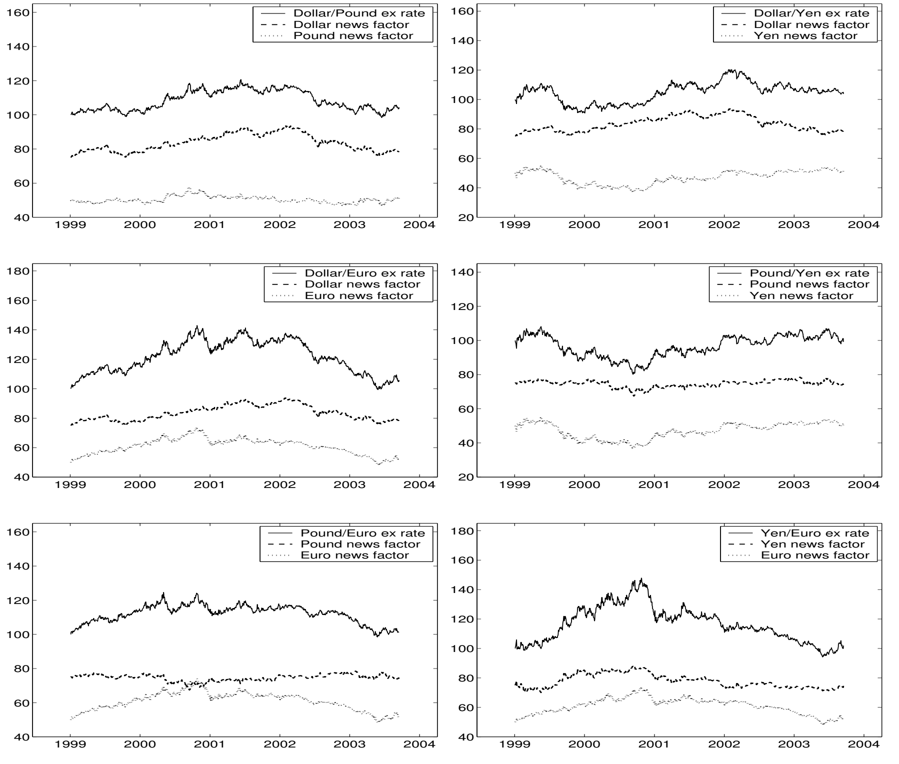

Figure 6.3: Exchange rates with their corresponding news series. Exchange rates and their respective currency-specific news series over the period January 1st, 1999 - September 16th, 2003. All series are presented as an index. News indices are shifted to accommodate comparison. The exchange rate (upper line) starts at 100 and the news indices (middle and lower line respectively) start at respectively 75 and 50.

$$\alpha_t = \begin{pmatrix} \alpha_{USD_t} \\ \alpha_{GBP_t} \\ \alpha_{JPY_t} \\ \alpha_{DEM_t} \\ \alpha_{WORLD_t} \end{pmatrix}$$

The rest of the model remains unchanged. To reserve space the graphs and tables for this model will be omitted except for the graphs involving the euro component.[12]

It is interesting to see if the change in the model leads to differences in the λ_{euro} component. The middle graph of Figure 6.4 shows the λ_{euro} when a world factor

[12]These graphs and tables are, however, available upon request.

is added. Comparing this with the λ_{euro} of the base model in the top graph of the figure, we see that they are very similar. Figure 6.5 shows a scatterplot of the λ_{euro} of the base model against the model including the world factor (left graph). The graph shows that the euro variances tend to be only a little bit lower when a world factor is included. From a practical point of view, this implies that the multivariate model is quite robust against adding a world factor. The same results appear for the other currency-specific components, which are not displayed here.

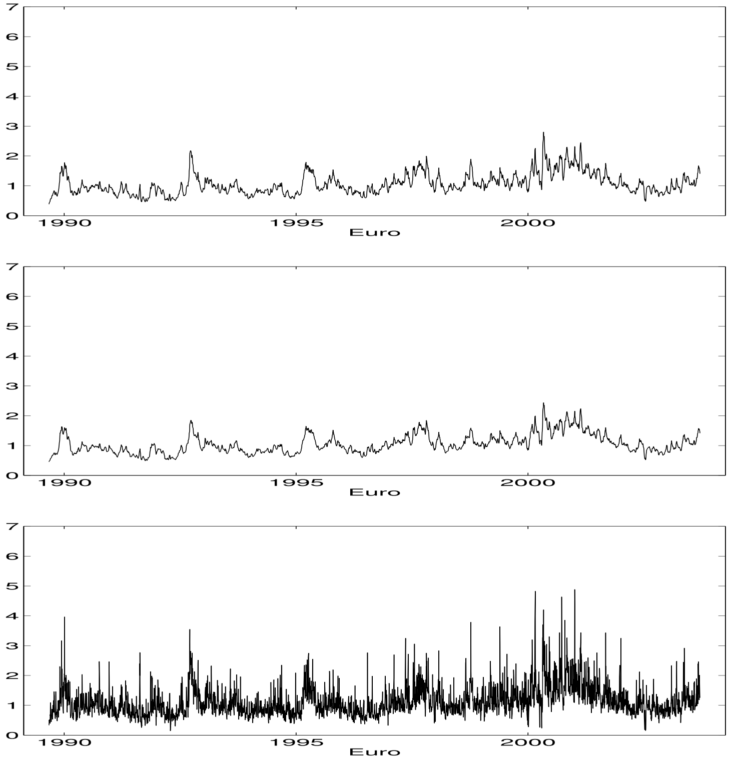

Figure 6.4: Estimated λ's. Euro components for the different models as described in Section 6.6. The top graph shows the euro components for the base model, the middle graph for the model where a world factor is included and the bottom graph for the case where H is a diagonal matrix instead of a full matrix. Estimation of the several models takes place over the period September the 1st, 1989 until September the 16th, 2003 (3652 observations).

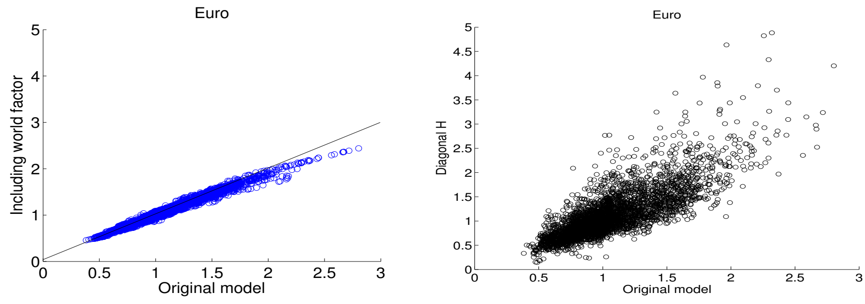

Figure 6.5: Comparison of λ's. Euro components for the model including a world factor (left graph) and for the model where H is a diagonal matrix (right graph) are displayed against the Euro component for the base model. Estimation of the several models takes place over the period September the 1st, 1989 until September the 16th, 2003 (3652 observations).

Another way to analyze this intuition is to calculate the correlation between the news factors of both models.[13] Table 6.4 shows the correlation between the euro news component of the base and the world factor model ((ORIG)-(WRLD)). We see that this correlation is practically 1.

Because the four-state model is a restricted version of the model including a world factor, a Likelihood Ratio (LR) test can be performed to evaluate whether the former model can be rejected in favor of the extended model. There are 2 parameter restrictions on the unrestricted model (the fifth element of both T and Q are set to 0) therefore the LR test statistic has an asymptotic distribution of χ^2_2. The test statistic is equal to LR $= 2(\log L_{unrestricted}$-$\log L_{restricted}) = 2$(-9097.937-(-9167.173)) $= 138.47$ which has a p-value of 0.000. From this perspective the four state model is too restrictive compared to the five state model although the world-factor model does not seem to have a qualitative impact on the λ's and the news factors.

[13]Note that the formulation of the news factors in Equation (6.6) is the same for both the models with and without a world factor and that there is no separate world news factor. The news factor is calculated using a numeraire currency. Therefore it is impossible to identify a world news component which has the same impact on each exchange rate because it is indistinguishable with the numeraire currency. However, adding a world component has an impact on the estimated parameters and latent currency components and therefore also has an indirect influence on the news factors for the several currencies.

		(ORIG)	(WRLD)	(H diag)
Base model	(ORIG)	1.0000	0.9999	0.9883
Inc. world factor	(WRLD)		1.0000	0.9872
H diagonal	(H diag)			1.0000

Table 6.4: Correlation matrix of euro news factors λ_{euro} across several models.

6.6.2 Another model specification

Instead of extending the model, we can also test if restricting the model leads to a model that is still able to capture the dynamics of the exchange rates but is more parsimonious. In the base model we have a full covariance matrix H for the measurement error terms and we like to test whether a diagonal matrix H is too restrictive. Figure 6.4 displays the λ_{euro} for the model with a diagonal H. We see that the euro factor is more volatile. This can also be seen from the right graph in Figure 6.5 which shows that the variances from both models are correlated.

Table 6.4 shows the unconditional correlations of the euro news factors between the model with the restricted H matrix and the base model presented in Section 6.3. We see that the correlation is very high.

The model with the diagonal H is a restricted version of the base model. Again we can test statistically if these models are the same or not. There are 20 parameter restrictions on the unrestricted model (the off-diagonal elements of the symmetric H matrix of the base model are set to 0). Therefore the LR test statistic has an asymptotic distribution of χ^2_{20}. The test statistic is equal to LR $= 2(\log L_{unrestricted}$-$\log L_{restricted}) =$ 2(-9167.173-(-10797.550)) = 3260.754 which has a p-value of 0.000 implying that the model where the H matrix is restricted to be diagonal is rejected in favor of the base model.

6.7 Concluding remarks

In this chapter we have presented a new approach for estimating multivariate stochastic volatility models for exchange rate volatilities. The model draws upon the decomposition of exchange rates into currency-specific factors as advocated by Mahieu and Schotman (1994). By using the range as a volatility measure we are able to estimate a parsimonious multivariate model for exchange rate volatility in a very efficient way.

Our results show that the model can be estimated efficiently through standard Kalman filter techniques using the EM-algorithm. We find that the currency-specific volatilities are substantially different from each other. The model is able to pick up some of the most saliant events in exchange rates that happened during the last decade. Notably, the Asian crisis and the fall of the pound sterling from the European Monetary System can be traced back to the yen and sterling currency components, respectively. We also find that euro-denominated exchange rates are predominantly determined by a euro-specific component during the first years after the introduction of the euro in 1999.

Furthermore, we find that the model is doing well compared to the alternative specifications we tested. Extending the model with a world factor is a statistical improvement, however produces results that are very similar to the results obtained from the base model. On the other hand, restricting the model leads to results which imply that the base model does a good job in describing the dynamics of exchange rate volatilities.

The analysis in this chapter can be extended in several ways. First, the specific factor structure that we imposed can be extended further. For example, each currency-factor could be split into a persistent and a stationary component.[14] Secondly, it would be interesting to analyze whether the news series can be related to economic variables, like interest rates, and monetary variables. Lastly, the multivariate model could be used for analyzing the prices of options and it could be used in analyzing the risks in international investment and trade portfolios.

[14]See also Alizadeh *et al.* (2002), who perform this analysis for univariate stochastic volatility models.

Appendix

The Kalman filter and smoother recursions

The Kalman filter and smoother are given for the sake of completeness. These equations are taken from Durbin and Koopman (2001) and the same notation will be used. Because there appears a constant in the measurement equation which is not present in the standard filter, a straightforward adjustment to the recursions is made. Note however that the smoother does not need to be changed because the effect of the constant is fully captured by the prediction error v_t. Our model is:

$$y_t = c + Z\alpha_t + \varepsilon_t,\ \varepsilon_t \sim N(0, H),\ t = 1, ..., n$$

$$\alpha_{t+1} = T\alpha_t + \eta_t,\ \eta_t \sim N(0, Q),\ t = 1, ..., n$$

$$\alpha_1 \sim N(a_1, P_1)$$

Then the Kalman filter is given as follows: If there are m states, $a_1 = 0_m$ and $P_1 = I_m$ then for $t = 1, ..., n$

$$v_t = y_t - c - Za_t$$

$$F_t = ZP_tZ' + H$$

$$K_t = TP_tZ'F_t^{-1}$$

$$L_t = T - K_tZ$$

$$a_{t+1} = E(\alpha_{t+1}|Y_t) = Ta_t + K_tv_t$$

$$P_{t+1} = cov(\alpha_{t+1}|Y_t) = TP_tL_t' + Q$$

Next the output of the Kalman filter is used in the smoother to construct a proxy for the latent factors. The smoother equations are as follows: If $r_n = 0_m$ and $N_n = 0_{m\times m}$ then for $t = 1, ..., n$

$$r_{t-1} = Z'F_t^{-1}v_t + L_t'r_t$$

$$N_{t-1} = Z'F_t^{-1}Z + L_t'N_tL_t$$

$$\hat{\alpha}_t = E(\alpha_t|y) = a_t + P_tr_{t-1}$$

$$V_t = P_t - P_tN_{t-1}P_t$$

Given the parameters of the model an estimation for the latent factors is then $\hat{\alpha}_t$. Furthermore the disturbance smoother equations from Durbin and Koopman (2001) are used to obtain the estimated errors for the measurement and transition equation, $\hat{\varepsilon}_t$ and $\hat{\eta}_t$.

The disturbance smoother

$$u_t = F_t^{-1} v_t - K_t' r_t$$

$$D_t = F_t^{-1} + K_t' N_t K_t$$

$$\hat{\varepsilon}_t = E(\varepsilon_t | y) = H u_t$$

$$cov(\varepsilon_t | y) = H - H D_t H$$

$$\hat{\eta}_t = E(\eta_t | y) = Q r_t$$

$$cov(\eta_t | y) = Q - Q N_t Q$$

Also in the estimation of the parameters of the model, the (inter temporal) covariance between smoothed states is needed. These expressions (see also Table 4.4 of Durbin and Koopman (2001)) are given by

$$cov(\alpha_t | y) = P_t (I_m - N_{t-1} P_t)$$

$$cov(\alpha_t, \alpha_{t+1} | y) = P_t L_t' (I_m - N_t P_{t+1)}$$

The EM-algorithm (Dempster *et al.* (1977)) consists of an estimation step and a maximization step. We follow the discussion in Shumway and Stoffer (2000) (Chapter 4, paragraph 3).

The Expectation-step of EM

In this step the expectation of the log likelihood is taken given the data and the parameters. The log likelihood in this case is

$$-2 \ln L = \ln |Q| + \sum_{t=1}^{n} (\alpha_{t+1} - T\alpha_t)' Q^{-1} (\alpha_{t+1} - T\alpha_t) +$$

$$+ \ln |H| + \sum_{t=1}^{n} (y_t - c - Z\alpha_t)' H^{-1} (y_t - c - Z\alpha_t)$$

which is similar to equation (4.69) in Shumway and Stoffer (2000). Taking the expectation of this expression gives something similar to equation (4.71) of this reference.

$$-2 \ln L = \ln |Q| + trace \left\{ Q^{-1} \left[S_{11} - S_{10} T' - T S_{10}' + T S_{00} T' \right] \right\} +$$

$$+ \ln |H| + trace \left\{ H^{-1} \sum_{t=1}^{n} \{ (y_t - c - Z\hat{\alpha}_t)(y_t - c - Z\hat{\alpha}_t)' + Z cov(\alpha_t | y) Z' \} \right\}$$

with

$$S_{11} = \sum_{t=1}^{n} \left\{\hat{\alpha}_{t+1}\hat{\alpha}'_{t+1} + cov(\alpha_{t+1}|y)\right\} = \sum_{t=1}^{n} \left\{\hat{\alpha}_{t+1}\hat{\alpha}'_{t+1} + P_{t+1}(I_m - N_t P_{t+1})\right\}$$

$$S_{10} = \sum_{t=1}^{n} \{\hat{\alpha}_{t+1}\hat{\alpha}'_{t} + cov(\alpha_{t+1}, \alpha_t|y)\} = \sum_{t=1}^{n} \left\{\hat{\alpha}_{t+1}\hat{\alpha}'_{t} + P_t L'_t(I_m - N_t P_{t+1)}\right\}$$

$$S_{00} = \sum_{t=1}^{n} \{\hat{\alpha}_{t}\hat{\alpha}'_{t} + cov(\alpha_t|y)\} = \sum_{t=1}^{n} \{\hat{\alpha}_{t}\hat{\alpha}'_{t} + P_t(I_m - N_{t-1}P_t)\}$$

In the above analysis some of the properties of the trace-operator are used. Also in the derivation they use that

$$E(\alpha_t\alpha'_t|y) = \hat{\alpha}_t\hat{\alpha}'_t + cov(\alpha_t|y)$$

$$E(\alpha_{t+1}\alpha'_t|y) = \hat{\alpha}_{t+1}\hat{\alpha}'_t + cov(\alpha_{t+1}, \alpha_t|y)$$

The P_t^n and $P_{t,t-1}^n$ defined in Shumway and Stoffer (2000) are equivalent to respectively $cov(\alpha_t|y)$ and $cov(\alpha_t, \alpha_{t-1}|y)$. The Kalman filter and smoother are applied to obtain $\hat{\alpha}_t$.

The maximization step of EM

Here an estimate of the parameters are given. The log likelihood function after the expectation step is given above. Because T and Q are assumed diagonal this expression can be simplified.

$$\ln|Q| + trace\left\{Q^{-1}\left[S_{11} - S_{10}T' - TS'_{10} + TS_{00}T'\right]\right\} =$$

$$\sum_{k=1}^{m} \ln Q_{kk} + \sum_{k=1}^{m} Q_{kk}^{-1}\left[(S_{11})_{kk} - 2(S_{10})_{kk}T_{kk} + (S_{00})_{kk}T_{kk}^2\right]$$

$$+ \ln|H| + trace\left\{H^{-1}\sum_{t=1}^{n}\{(y_t - c - Z\hat{\alpha}_t)(y_t - c - Z\hat{\alpha}_t)' + Zcov(\alpha_t|y)Z'\}\right\}$$

where is used that

1. if matrices A and B are diagonal then so is AB,
2. if matrix A is diagonal then $trace(AB) = \sum a_{ii}b_{ii}$,
3. $trace(AB) = trace(BA)$,
4. and $trace(A + B) = trace(A) + trace(B)$.

Then the estimates of the parameters are given by

$$c = n^{-1} \sum_{t=1}^{n} \{y_t - Z\hat{\alpha}_t\}$$

$$H = n^{-1} \sum_{t=1}^{n} \{(y_t - c - Z\hat{\alpha}_t)(y_t - c - Z\hat{\alpha}_t)' + Zcov(\alpha_t|y)Z'\}$$

$$T_{kk} = (S_{10})_{kk}/(S_{00})_{kk} \text{for } k = 1, ..., 4$$

$$Q_{kk} = n^{-1} \left[(S_{11})_{kk} - (S_{10})^2_{kk}/(S_{00})_{kk}\right] \text{ for } k = 1, ..., 4$$

Chapter 7

Summary and Concluding Remarks

7.1 Summary

There is a vast literature on exchange rates discussing various topics ranging from exchange rate puzzles to measuring volatility. This thesis sheds more light on the Purchasing Parity Puzzle (PPP), the Uncovered Interest rate Parity puzzle and exchange rate volatility.

In Chapter 2 the PPP hypothesis within the euro area and for the euro versus other major currencies is examined. Furthermore, the effect of the European economic integration process (represented by the Maastricht Treaty in 1992 and introduction of the euro in 1999) on the inferences on PPP is investigated. Evidence for the PPP hypothesis is found for the euro area with the DMark as numeraire for the period 1973-2003 when the mean reversion for all exchange rates in the multivariate panel is restricted to be the same. For the panel without this restriction evidence for PPP is found for some currency pairs but not for others so that can be argued that the homogeneous setting is too restrictive. This is more apparent for the case where the real exchange rate of the euro area as a whole (using "synthetic" data for the pre-euro period) is compared with those of other major economies for the period 1979-2003. Here the unit root is again rejected for the panel as a whole but when heterogeneous mean reversion is accounted for there is only evidence for PPP between the euro area and Switzerland.

For the analysis of the influence of the European economic integration process for PPP the main results are that the adoption of the Maastricht Treaty in 1992 seems to have an effect on the integration process due to the strong evidence for PPP after 1992 for the homogenous panel of exchange rates. When the homogeneous restriction is abandoned no clear pattern of convergence across currency pairs emerges.

Although the power of the heterogeneous multivariate unit root test is generally higher than the power of the univariate unit root test, the differences are small and similar. Furthermore, the power is relatively low suggesting that the research on the validity of the PPP hypothesis to improve the power of the unit root test by using a multivariate approach seems back to square one.

Chapter 3 analyzes the finite-sample properties of the unit root test to test the PPP hypothesis under three different multivariate methodologies. It is shown that the first methodology, where both estimation and testing is done homogeneously, suffers from important adverse effects. Not only does the estimated mean reversion display serious biases, but also estimation uncertainties arise which implies that the statistical power of the unit root test is very limited (in some cases, the power is even lower than the size of the test). Moreover, the power function is not monotonically increasing when the mean reversion parameters under the alternative hypothesis move away from the unit root null. All this can lead to misleading inferences about the validity of the PPP hypothesis.

The power function of the other two methodologies are monotonically increasing functions. Furthermore, it depends on the number of unit roots in the panel. For the panel in this chapter, consisting of real exchange rates for Canada, the euro area, Japan, and the U.K. with the U.S. dollar as numeraire, it holds that the methodology, where heterogeneous testing is performed, has a higher power when there at least three unit roots in the panel. When there is at most one unit root in the panel the methodology with homogeneous testing performs better with respect to the power. The two methodologies perform similar when two unit roots are present in the panel. Note however that these differences in power performance are due to the model where testing is done homogeneously (but estimation still heterogeneously). For this model the power depends on the number of unit roots in the panel while for the model where estimation and testing takes place heterogeneously the power does not so much depends on the number of stationary series in the panel; however, it does depend on the country under study. The power of the non-stationary series are low and close to the size of the test and almost constant even when the degree of mean reversion under the alternative move away from the unit root null hypothesis which is a desirable feature.

Chapter 4 combines the insights of Clarida and Taylor (1997), Bansal (1997) and Wu and Zhang (1996) to try to improve upon describing and forecasting the exchange rate return using possible (nonlinear) information present in the interest rate term structure. The in-sample and out-of-sample evidence reported by Clarida and Taylor (1997) is not validated by the results in this chapter. The lagged forward premia in the VECM model which are extracted from the differential interest rate term structure are not significant and thus not able to explain the in-sample behavior of the exchange rate return. Incorporating these premia into the model does not help to improve the forecast of the exchange rate return compared to the Random Walk either, so that the superior performance as found by Clarida and Taylor (1997) is not supported.

Allowing for asymmetric or nonlinear variants of the forward premia does not give other outcomes. These variables are still unable to capture (part of) the dynamics of the exchange rates, both in-sample and out-of-sample.

In Chapter 5 an adaptation of the model of Baillie and Kiliç (2006) is employed to examine the possible nonlinearity between the exchange rate return and the forward premia/interest rate differential. The deviations from UIP are measured and tested using the nonlinear part of the model. The findings for this model are that the confidence bounds are too broad to reject UIP for large parts of the sample.

This model is extended to the multivariate case where the nonlinear equations are gathered into one panel such that interdependencies between the series are accounted for. Additionally, the U.S. interest rate is a common factor across the equations but exploiting this feature does not lead to different conclusions so that UIP still can not be rejected.

Chapter 6 discusses the parsimonious multivariate Stochastic Volatility model that measures exchange rate volatility and that is estimated efficiently by using the distributional properties of the range-based volatility measure, which makes use of high and low prices. The estimated currency-specific volatilities that are extracted from the exchange rate volatilities are substantially different from each other. The source of the volatility in the exchange rates due to for example the Asian crises and the fall of the sterling pound from the European Monetary System can be found in the yen and sterling components respectively.

7.2 Concluding remarks

A clear evolution in the PPP literature can be recognized: because univariate analyzes are in general unable to find evidence in favor of PPP due to a lack of power of the unit root test, attempts are made to raise the PPP analysis to a multivariate level using panel models to increase the power. After two restrictions of this panel model have been successfully contested in the literature, a third important restriction, homogeneity of the degree of mean reversion across the real exchange rates in the panel, is discussed in this thesis. It is found that estimation of the panel should be done heterogeneously. Furthermore, when testing is also done heterogeneously, the results of the multivariate approach are similar to the univariate approach bringing the literature back to square one which leads us to conclude that there just is not enough data available to find evidence in favor or against PPP. We expect that the PPP literature will focus more on data issues (e.g. heterogeneity in the aggregation of price data) and nonlinear reversion of the real exchange rates to their mean due to business cycles.

Many studies find that UIP does not hold; the slope parameter in the linear (Fama) regression, where the exchange rate return is explained by the forward premium, is significant different from unity and often negative. There are various paths that can be considered to try to solve this puzzle. Exploiting information present in the differential interest rate term structure to explain the exchange rate return is according to Clarida and Taylor (1997) very successful, however, this thesis does not confirm their results. Another way to contest the problem is to have a nonlinear approach to the UIP puzzle. However, this thesis does not find any evidence for nonlinearity either. There is a wide variety of other approaches and no clear main road can be recognized to solve the puzzle, although very promising is the work of Chinn and Meredith (2004) where long-horizon data is used instead of short-horizon data. Long-term data seem to deliver estimates of the slope parameter which are more consistent to UIP.

Measuring volatility using range data (based on high and low prices) is not only efficient (see Alizadeh *et al.* (2002) but also easy due to the availability of this kind of data. Although realized volatility might be more efficient, high-frequency data is only available fairly recently and for a limited number of assets while this is not the case for high and low prices. Future research has to be conducted to see whether the range based volatility measure is able to improve the estimation of other (multivariate) models where volatility plays a crucial role (e.g. option pricing models).

Furthermore, future research is needed to investigate the informational content in the currency-specific volatilities and news factors that are extracted from the multivariate model as discussed in Chapter 6 which might then be used in for example international portfolio choice models.

Bibliography

Abuaf, N. and P. Jorion (1990), Purchasing Power Parity in the Long Run, *Journal of Finance*, **45**, 157–174.

Aït-Sahalia, Y. (2002), Maximum Likelihood Estimation of Discretely Sampled Diffusions: A Closed-Form Approximation Approach, *Econometrica*, **70**, 223–262.

Aït-Sahalia, Y. and P. A. Mykland (2003), The Effects of Random and Discrete Sampling When Estimating Continuous-Time Diffusions, *Econometrica*, **71**, 483–549.

Aït-Sahalia, Y., P. A. Mykland, and L. Zhang (2005), How Often to Sample a Continuous-Time Process in the Presence of Market Microstructure Noise, *Review of Financial Studies*, **18**, 351–416.

Alizadeh, S., M. W. Brandt, and F. X. Diebold (2002), Range-Based Estimation of Stochastic Volatility Models, *Journal of Finance*, **57**, 1047–1091.

Andersen, T. G. (1992), Volatility, working paper 144, Northwestern University.

Andersen, T. G. and T. Bollerslev (1998), Answering the Skeptics: Yes, Standard Volatility Models Do Provide Accurate Forecasts, *International Economic Review*, **39**, 885–905.

Andersen, T. G. and B. E. Sørensen (1996), GMM Estimation of a Stochastic Volatility Model: A Monte Carlo Study, *Journal of Business and Economic Statistics*, **14**, 328–352.

Andersen, T. G., T. Bollerslev, F. X. Diebold, and H. Ebens (2001a), The Distribution of Realized Stock Return Volatility, *Journal of Financial Economics*, **61**, 43–76.

Andersen, T. G., T. Bollerslev, F. X. Diebold, and P. Labys (2001b), The Distribution of Realized Exchange Rate Volatility, *Journal of the American Statistical Association*, **96**, 42–55.

Andersen, T. G., T. Bollerslev, F. X. Diebold, and P. Labys (2003), Modeling and Forecasting Realized Volatility, *Econometrica*, **71**, 579–625.

Andersen, T. G., T. Bollerslev, and F. X. Diebold (2006), Parametric and Nonparametric Volatility Measurement, in L. P. Hansen and Y. Aït-Sahalia (eds.), *forthcoming in Handbook of Financial Econometrics*, Amsterdam: North-Holland.

Bai, X., J. R. Russell, and G. C. Tiao (2001), Beyond Merton's Utopia (I): Effects of Non-Normality and Dependence on the Precision of Variance Estimates Using High-Frequency Financial Data.

Baillie, R. T. and T. Bollerslev (1994), Cointegration, Fractional Cointegration, and Exchange Rate Dynamics, *Journal of Finance*, **49**, 737–745.

Baillie, R. T. and R. Kiliç (2006), Do Asymmetric and Nonlinear Adjustments Explain the Forward Premium Anomaly?, *Journal of International Money and Finance*, **25**, 22–47.

Baldwin, R. (1990), Re-Interpreting the Failure of Foreign Exchange Market Efficiency Tests: Small Transaction Costs, Big Hysterisis Bands, nBER Working Paper No. 3319.

Bank for International Settlements (2001), Why Has Global FX Turnover Declined? Explaining the 2001 Triennial Survey, *BIS Quarterly Review*.

Bank for International Settlements (2003), International Banking and Financial Market Developments, *BIS Quarterly Review*.

Bansal, R. (1997), An Exploration of the Forward Premium Puzzle in Currency Markets, *Review of Financial Studies*, **10**, 369–403.

Bollerslev, T. (1986), Generalized Autoregressive Conditional Heteroskedasticity, *Journal of Econometrics*, **31**, 307–327.

Brandt, M. W. and F. X. Diebold (2006), A No-Arbitrage Approach to Range-Based Estimation of Return Covariances and Correlations, *Journal of Business*, **79**, forthcoming.

Brandt, M. W., J. H. Cochrane, and P. Santa-Clara (2005), International Risk Sharing is Better Than You Think, or Exchange Rates are Too Smooth, *forthcoming in Journal of Monetary Economics*.

Breuer, J. B., R. McNown, and M. Wallace (2002), Series-Specific Unit Root Tests with Panel Data, *Oxford Bulletin of Economics and Statistics*, **64**, 527–546.

Broto, C. and E. Ruiz (2004), Estimation Methods for Stochastic Volatility Models: A Survey, *Journal of Economic Surveys*, **18**, 613–650.

Campbell, J. Y. and P. Perron (1991), Pitfalls and Opportunities: What Macroeconomists Should Know About Unit Roots, nBER Technical Working Paper No. 100.

Caramazza, F. and J. Aziz (1998), Fixed or Flexible? Getting the Exchange Rate Right in the 1990s, Tech. Rep. Economic Issues 13, International Monetary Fund.

Chib, S., F. Nardari, and N. Shephard (2005), Analysis of High Dimensional Multivariate Stochastic Volatility Models, *forthcoming in Journal of Econometrics*.

Chinn, M. D. (1991), Some Linear and Nonlinear Thoughts on Exchange Rates, *Journal of International Money and Finance*, **10**, 214–230.

Chinn, M. D. (2002), Productivity and the Euro-Dollar Exchange Rate Puzzle, nBER Working Paper No. 8824.

Chinn, M. D. (2006), The (Partial) Rehabilitation of Interest Rate Parity in the Floating Rate Era: Longer Horizons, Alternative Expectations, and Emerging Markets, *Journal of International Money and Finance*, **25**, 7–21.

Chinn, M. D. and G. Meredith (2004), Monetary Policy and Long-Horizon Uncovered Interest Parity, *IMF Staff Papers*, **51**, 409–430, international Monetary Fund.

Chinn, M. D. and G. Meredith (2005), Testing Uncovered Interest Parity at Short and Long Horizons During the Post-Bretton Woods Era, nBER Working Paper No. 11077.

Choi, C.-Y. (2004), Searching for Evidence of Long-Run PPP from a Post-Bretton Woods Panel: Separating the Wheat from the Chaff, *Journal of International Money and Finance*, **23**, 1159–1186.

Clarida, R. H. and M. P. Taylor (1997), The Term Structure of Forward Exchange Premiums and the Forecastability of Spot Exchange Rates: Correcting the Errors, *Review of Economics and Statistics*, **79**, 353–361.

Clostermann, J. and B. Schnatz (2000), The Determinants of the Euro-Dollar Exchange Rate: Synthetic Fundamentals an a Non-Existing Currency, discussion Paper, 2/00, Deutsche Bundesbank.

Cornell, B. (1977), Spot Rates, Forward Rates and Exchange Market Efficiency, *Journal of Financial Economics*, **5**, 55–65.

Crowder, W. J. (1994), Foreign Exchange Market Efficiency and Common Stochastic Trends, *Journal of International Money and Finance*, **13**, 551–564.

DeJong, D. N. and C. H. Whiteman (1991a), The Case for Trend-Stationarity is Stronger Than We Thought, *Journal of Applied Econometrics*, **6**, 413–421.

DeJong, D. N. and C. H. Whiteman (1991b), Reconsidering 'Trends and Random Walks in Macroeconomic Time Series', *Journal of Monetary Economics*, **28**, 221–254.

DeJong, D. N., J. C. Nankervis, N. E. Savin, and C. H. Whiteman (1992), Integration versus Trend Stationarity in Time Series, *Econometrica*, **60**, 423–433.

Dempster, A. P., N. M. Laird, and D. B. Rubin (1977), Maximum Likelihood from Incomplete Data Via the EM Algorithm, *Journal of the Royal Statistical Society. Series B (Methodological)*, **39**, 1–38.

Dickey, D. A. and W. A. Fuller (1979), Distribution of the Estimators for Autoregressive Time Series with a Unit Root, *Journal of the American Statistical Association*, **74**, 427–431.

Dumas, B. (1992), Dynamic Equilibrium and the Real Exchange Rate in a Spatially Separated World, *Review of Financial Studies*, **5**, 153–180.

Durbin, J. and S. J. Koopman (2001), *Time Series Analysis by State Space Methods*, Oxford University Press.

Edison, H. J. (1987), Purchasing Power Parity in the Long Run: A Test of the Dollar/Pound Exchange Rate (1890-1978), *Journal of Money, Credit, and Banking*, **19**, 376–387.

Engel, C. (1996), The Forward Discount Anomaly and the Risk Premium: A Survey of Recent Evidence, *Journal of Empirical Finance*, **3**, 123–192.

Engel, C., M. K. Hendrickson, and J. H. Rogers (1997), Intranational, Intracontinental, and Intraplanetary PPP, *Journal of the Japanese and International Economies*, **11**, 480–501.

Engle, R. F. (1982), Autoregressive Conditional Heteroscedasticity with Estimates of the Variance of United Kingdom Inflation, *Econometrica*, **50**, 987–1007.

Engle, R. F. and C. W. J. Granger (1987), Co-Integration and Error Correction: Representation, Estimation, and Testing, *Econometrica*, **55**, 251–276.

Engle, R. F. and K. F. Kroner (1995), Multivariate Simultaneous Generalized ARCH, *Econometric Theory*, **11**, 122–150.

European Central Bank (1999), Inflation Differentials in a Monetary Union, *ECB Monthly Bulletin*, 36–44.

Fama, E. F. (1984), Forward and Spot Exchange Rates, *Journal of Monetary Economics*, **14**, 319–338.

Fleissig, A. R. and J. Strauss (2000), Panel Unit Root Tests of Purchasing Power Parity for Price Indices, *Journal of International Money and Finance*, **19**, 489–506.

Flôres, R., P.-Y. Preumont, and A. Szafarz (1995), Multivariate Unit Root Tests, mimeo.

Flôres, R., P. Jorion, P.-Y. Preumont, and A. Szafarz (1999), Multivariate Unit Root Tests of the PPP Hypothesis, *Journal of Empirical Finance*, **6**, 335–353.

Frankel, J. A. and A. K. Rose (1996), A Panel Project on Purchasing Power Parity: Mean Reversion Within and Between Countries, *Journal of International Economics*, **40**, 209–224.

Froot, K. A. and K. Rogoff (1995), *Perspectives on PPP and Long-Run Real Exchange Rates*, vol. 3 of *Handbook of International Economics*, chap. 32, pp. 1647–1688, north-Holland, Amsterdam.

Froot, K. A. and R. H. Thaler (1990), Anomalies: Foreign Exchange, *Journal of Economic Perspectives*, **4**, 179–192.

Hai, W., N. C. Mark, and Y. Wu (1997), Understanding Spot and Forward Exchange Rate Regressions, *Journal of Applied Econometrics*, **12**, 715–734.

Hall, A. D., H. M. Anderson, and C. W. J. Granger (1992), A Cointegration Analysis of Treasury Bill Yields, *Review of Economics and Statistics*, **74**, 116–126.

Harvey, A. C., E. Ruiz, and N. Shephard (1994), Multivariate Stochastic Variance Models, *Review of Economic Studies*, **61**, 247–264.

Hodrick, R. J. (1987), *The Empirical Evidence on the Efficiency of Forward and Futures Foreign Exchange Markets*, London: Harwood.

Hollifield, B. and R. Uppal (1997), An Examination of Uncovered Interest Rate Parity in Segmented International Commodity Markets, *Journal of Finance*, **52**, 2145–2170.

Huisman, R., K. Koedijk, C. Kool, and F. Nissen (1998), Extreme Support for Uncovered Interest Parity, *Journal of International Money and Finance*, **17**, 211–228.

Im, K. S., M. H. Pesaran, and Y. Shin (2003), Testing for Unit Roots in Heterogeneous Panels, *Journal of Econometrics*, **115**, 53–74.

Imbs, J., H. Mumtaz, M. O. Ravn, and H. Rey (2005), PPP Strikes Back: Aggregation and the Real Exchange Rate, *Quarterly Journal of Economics*, **120**, 1–43.

Jacquier, E., N. G. Polson, and P. E. Rossi (1994), Bayesian Analysis of Stochastic Volatility Models, *Journal of Business and Economic Statistics*, **12**, 371–389.

Johansen, S. (1991), Estimation and Hypothesis Testing of Cointegration Vectors in Gaussian Vector Autoregressive Models, *Econometrica*, **59**, 1551–1580.

Jorion, P. and R. J. Sweeney (1996), Mean Reversion in Real Exchange Rates: Evidence and Implications for Forecasting, *Journal of International Money and Finance*, **15**, 535–550.

Koedijk, C. G., P. C. Schotman, and M. A. Van Dijk (1998), The Reemergence of PPP in the 1990s, *Journal of International Money and Finance*, **17**, 51–61.

Koedijk, K. G., B. Tims, and M. A. Van Dijk (2004), Purchasing Power Parity and the Euro Area, *Journal of International Money and Finance*, **23**, 1081–1107.

Koop, G. (1992), 'Objective' Bayesian Unit Root Tests, *Journal of Applied Econometrics*, **7**, 65–82.

Levin, A., C.-F. Lin, and C.-S. J. Chu (2002), Unit Root Tests in Panel Data: Asymptotic and Finite-Sample Properties, *Journal of Econometrics*, **108**, 1–24.

Liesenfeld, R. and J.-F. Richard (2003), Univariate and Multivariate Stochastic Volatility Models: Estimation and Diagnostics, *Journal of Empirical Finance*, **10**, 505–531.

Lopez, C. and D. H. Papell (2004), Convergence to Purchasing Power Parity at the Commencement of the Euro, *Review of International Economics*, forthcoming.

Lothian, J. R. (1997), Multi-Country Evidence on the Behavior of Purchasing Power Parity under the Current Float, *Journal of International Money and Finance*, **16**, 19–35.

Lothian, J. R. and M. P. Taylor (1996), Real Exchange Rate Behavior: The Recent Float from the Perspective of the Past Two Centuries, *Journal of Political Economy*, **104**, 488–509.

Lütkepohl, H. (1993), *Introduction to Multiple Time Series Analysis*, Springer-Verlag.

Lutz, M. (2002), Price Convergence under EMU? First Estimates, working paper, University of St. Gallen.

Lyons, R. K. (2001), *The Microstructure Approach to Exchange Rates*, Cambridge and London: MIT Press.

Maddala, G. S. and S. Wu (1999), A Comparative Study of Unit Root Tests with Panel Data and a New Simple Test, *Oxford Bulletin of Economics and Statistics*, **61**, 631–652.

Maeso-Fernandez, F., C. Osbat, and B. Schnatz (2001), Determinants of the Euro Real Effective Exchange Rate: A Beer/Peer Approach.

Mahieu, R. J. and P. C. Schotman (1994), Neglected Common Factors in Exchange Rate Volatility, *Journal of Empirical Finance*, **1**, 279–311.

Mahieu, R. J. and P. C. Schotman (1998), An Empirical Application of Stochastic Volatility Models, *Journal of Applied Econometrics*, **13**, 333–360.

Mark, N. C. and Y.-K. Moh (2004), Official Interventions and Occasional Violations of Uncovered Interest Parity in the Dollar-DM Market.

McAleer, M. (2005), Automated Inference and Learning in Modeling Financial Volatility, *Econometric Theory*, **21**, 232–261.

Meese, R. and K. Rogoff (1983a), *The Out-of-Sample Failure of Empirical Exchange Rates: Sampling Error or Misspecification?*, in J. Frenkel (ed.) Exchange Rates and International Economics, 67-105.

Meese, R. A. and K. Rogoff (1983b), Empirical Exchange Rate Models of the Seventies. Do They Fit Out of Sample?, *Journal of International Economics*, **14**, 3–24.

Meese, R. A. and K. J. Singleton (1982), On Unit Roots and the Empirical Modeling of Exchange Rates, *Journal of Finance*, **37**, 1029–1035.

Murray, C. J. and D. H. Papell (2002), The Purchasing Power Parity Persistence Paradigm, *Journal of International Economics*, **56**, 1–19.

Obstfeld, M. and K. Rogoff (2005), The Unsustainable US Current Account Position Revisited, working Paper.

O'Connell, P. G. J. (1998), The Overvaluation of Purchasing Power Parity, *Journal of International Economics*, **44**, 1–19.

Oehlert, G. W. (1992), A Note on the Delta Method, *American Statistician*, **46**, 27–29.

Papell, D. H. (1997), Searching for Stationarity: Purchasing Power Parity under the Current Float, *Journal of International Economics*, **43**, 313–332.

Papell, D. H. (2002), The Great Appreciation, the Great Depreciation, and the Purchasing Power Parity Hypothesis, *Journal of International Economics*, **57**, 51–82.

Papell, D. H. and H. Theodoridis (2001), The Choice of Numeraire Currency in Panel Tests of Purchasing Power Parity, *Journal of Money, Credit, and Banking*, **33**, 790–803.

Parkinson, M. (1980), The Extreme Value Method for Estimating the Variance of the Rate of Return, *Journal of Business*, **53**, 61–65.

Pesaran, M. H. and R. Smith (1995), Estimating Long-Run Relationships from Dynamic Heterogeneous Panels, *Journal of Econometrics*, **68**, 79–113.

Phillips, P. C. B. (1991), To Criticize the Critics: An Objective Bayesian Analysis of Stochastic Trends, *Journal of Applied Econometrics*, **6**, 333–364.

Pollard, P. S. (1998), The Role of the Euro as an International Currency, *Columbia Journal of European Law*, 394–420.

Portes, R. (2001), The Euro and the International Financial System, cEPR Working Paper, 2955.

Portes, R. and H. Rey (1998), The Emergence of the Euro as an International Currency, *Economic Policy*, **13**, 307–343.

Record, N. (2004), *Currency Overlay*, Wiley-Interscience.

Robertson, D. and J. Symons (1992), Some Strange Properties of Panel Data Estimators, *Journal of Applied Econometrics*, **7**, 175–189.

Rogers, J. H. (2001), Price Level Convergence, Relative Prices, and Inflation in Europe, international Finance Discussion Papers, 699, Board of Governors of the Federal Reserve System.

Rogoff, K. (1996), The Purchasing Power Parity Puzzle, *Journal of Economic Literature*, **34**, 647–668.

Sarno, L. (2005), Viewpoint: Towards a Solution to the Puzzles in Exchange Rate Economics: Where Do We Stand?, *Canadian Journal of Economics*, **38**, 673–708.

Sarno, L. and M. P. Taylor (2003), *The Economics of Exchange Rates*, Cambridge and New York: Cambridge University Press.

Sarno, L., G. Valente, and H. Leon (2005), The Forward Bias Puzzle and Nonlinearity in Deviations from Uncovered Interest Parity: A New Perspective, working paper.

Schotman, P. and H. K. Van Dijk (1991), A Bayesian Analysis of the Unit Root in Real Exchange Rates, *Journal of Econometrics*, **49**, 195–238.

Shephard, N. (1993), Fitting Nonlinear Time-Series Models with Applications to Stochastic Variance Models, *Journal of Applied Econometrics*, **8**, 135–152, special issue.

Shumway, R. H. and D. S. Stoffer (2000), *Time Series Analysis and its Applications*, Springer-Verlag New York, Inc.

Solnik, B. H. (2004), *International Investments*, Addison-Wesley.

Taylor, A. M. and M. P. Taylor (2004), The Purchasing Power Parity Debate, *Journal of Economic Perspectives*, **18**, 135–158.

Taylor, M. P. (1995), The Economics of Exchange Rates, *Journal of Economic Literature*, **33**, 13–47.

Taylor, M. P. and L. Sarno (1998), The Behavior of Real Exchange Rates During the Post-Bretton Woods Period, *Journal of International Economics*, **46**, 281–312.

Taylor, S. J. (1986), *Modelling Financial Time Series*, New York: John Wiley.

Tims, B. and R. Mahieu (2006), A Range-Based Multivariate Stochastic Volatility Model for Exchange Rates, *Econometric Reviews*, **25**, 409–424.

van Tol, M. R. (2005), *Beyond the Threshold: Theoretical and Empirical Nonlinear Time-Series Econometrics of Foreign Exchange Markets*, Ph.D. thesis, Universiteit Maastricht.

van Tol, M. R. and C. C. Wolff (2005), Forecasting the Spot Exchange Rate with the Term Structure of Forward Premia: Multivariate Threshold Cointegration, cEPR Working Paper, 4958.

Verbeek, M. (2000), *A Guide to Modern Econometrics*, John Wiley and Sons.

Wu, J.-L. and S. Wu (2001), Is Purchasing Power Parity Overvalued?, *Journal of Money, Credit, and Banking*, **33**, 804–812.

Wu, Y. and H. Zhang (1996), Asymmetry in Forward Exchange Rate Bias: A Puzzling Result, *Economic Letters*, **50**, 407–411.

Nederlandse samenvatting (Summary in Dutch)

Dit proefschrift bestaat uit vijf empirische studies over wisselkoersen. De eerste twee studies hebben als onderwerp de koopkrachtpariteit (hoofdstukken 2 en 3), de derde en vierde studie gaan over de ongedekte interest pariteit (hoofdstukken 4 en 5) en de laatste behandelt de volatiliteit van wisselkoersen (hoofdstuk 6).

De koopkrachtpariteit (KKP) veronderstelt dat prijzen van goederen gelijk zouden moeten zijn in verschillende landen als deze in dezelfde munteenheid zijn weergegeven. Als deze theorie opgaat dan zou de rëele wisselkoers tussen twee landen (de nominale wisselkoers gecorrigeerd voor de inflatie in beide landen) constant over de tijd moeten zijn. Door inefficiënties zoals transactiekosten, mate van verhandelbaarheid etc, gaat de pariteit in praktijk (vaak) niet op. Maar de wetenschappelijke opvatting is dat afwijkingen van deze pariteit tijdelijk zijn zodat reële prijzen naar elkaar convergeren. Statistisch betekent dit dat de reële wisselkoers stationair is en geen eenheidswortel bevat: na een verandering van de rëele wisselkoers door een exogene schok zal deze convergeren naar het lange termijn gemiddelde. Hoe snel dit gaat hangt af van de halfwaardetijd (de tijd nodig om de helft van een stijging/daling in de wisselkoers te absorberen). Als de halfwaardetijd statistisch kleiner dan één is, dan beschouwt men de rëele wisselkoers als stationair en kan de KKP hypothese niet worden verworpen. Als de nulhypothese van een halfwaardetijd gelijk aan één niet verworpen kan worden en dus een eenheidswortel bevat dan gaat de KKP niet op.

De KKP literatuur heeft een duidelijke evolutie ondergaan: eerst werd de aanwezigheid van een eenheidswortel voor elke reële wisselkoers afzonderlijk getoetst. Het nadeel van deze univariate analyse is dat het onderscheidingsvermogen van de toets laag is en er dus nauwelijks bewijs voor de KKP hypothese is te vinden ook al zou deze opgaan. Een oplossing zou zijn om de dataset uit te breiden over een langere periode (zie bij. Edison (1987) en Lothian en Taylor (1996)) maar een nadeel is dat de wisselkoers dan vaak zowel een vast als een variabel wisselkoers regime heeft gevolgd dat de uitkomst van de toets beinvloedt. Een andere oplossing is om niet de tijdsdimensie maar de cross-sectionele dimensie te vergroten zodat de koopkrachtpariteit wordt onderzocht voor het panel bestaande uit reële wisselkoersen als geheel i.p.v. voor individuele rëele wisselkoersen. Het gebruik van dit multivariate model leidde tot meer bewijs voor het opgaan van de pariteit (zie o.a. Abuaf en Jorion (1990)).

De kritiek van O'Connell (1998) op dit model is dat het negeren van mogelijke afhankelijkheid tussen de reële wisselkoersen in het panel kan leiden tot overacceptatie van de koopkracht pariteit hypothese. Een tweede punt van kritiek aangestipt door Papell en Theodoridis (2001) en Wu en Wu (2001) is dat er rekening dient worden gehouden met de heterogene serie correlatie in het panel anders wordt er minder snel bewijs gevonden voor het opgaan van de pariteit.

Een derde, veelvuldig gebruikte restrictie die in dit proefschrift onderzocht wordt is de restrictie van een gezamenlijke halfwaardetijd voor alle reële wisselkoersen in het panel. Omdat men zou verwachten dat de halfwaardetijd van de reële wiselkoers tussen twee landen afhankelijk zou kunnen zijn van afstand, handelsovereenkomsten etc lijkt de homogeniteit van de halfwaardetijd een te strikte beperking van het model.

In hoofdstuk 2 wordt het model met en zonder de restrictie van een homogene halfwaardetijd toegepast om te toetsen of de koopkrachtpariteit opgaat tussen de eurolanden en tussen het euro-gebied en andere belangrijke valuta. Verder wordt de invloed van het Europese integratieproces op de stationariteit van het panel met reële wisselkoersen onderzocht.In het bijzonder wordt bekeken wat het effect is van het Verdrag van Maastricht in 1992 en de introductie van de euro in 1999 op de koopkrachtpariteit.

Bewijs voor het opgaan van de koopkrachtpariteit is gevonden voor het euro-gebied met de DMark als numeraire voor de periode 1973-2003 als een homogene halfwaardetijd voor het panel wordt verondersteld. Zonder deze restrictie geldt KKP voor sommige wisselkoersen maar niet voor anderen zodat de restrictie van een homogene halfwaardetijd discutabel is. Als het euro-gebied als geheel wordt vergeleken met andere valuta dan gaat de KKP hypothese op voor het gehele panel als een homogene halfwaardetijd wordt verondersteld maar als een heterogene halfwaardetijd wordt toegelaten dan wordt er alleen bewijs voor KKP gevonden voor de wisselkoers tussen het euro-gebied en Zwitserland.

Het verdrag van Maastricht blijkt effect te hebben op het integratieproces doordat er sterk bewijs is voor KKP na 1992 voor het panel met een homogene halfwaardetijd.

Echter, als er heterogeniteit wordt verondersteld is er geen zichtbaar patroon van convergentie in de wisselkoersen te ontdekken.

In hoofdstuk 3 worden de eigenschappen van de eenheidsworteltoets onder de loep genomen. Er wordt onderscheid gemaakt tussen drie verschillende methodologieën op basis van het homogeen en heterogeen schatten en toetsen van de modellen. In het eerste model wordt zowel homogeen geschat als getoetst wat in de literatuur een gebruikelijke manier is om de KKP hypothese te evalueren. Bij het tweede model wordt de halfwaardetijd voor iedere wisselkoers in het panel apart geschat maar nog steeds wordt de eenheidsworteltoets voor het gehele panel uitgevoerd. Beide modellen geven geen uitsluitsel over hoeveel en welke wisselkoersen stationair zijn. In het derde model wordt voor elke wisselkoers niet alleen de halfwaardetijd apart geschat maar ook apart getoetst of deze kleiner is dan één.

De belangrijkste conclusie van dit hoofdstuk is dat het eerste model (zowel homogeen schatten als toetsen) nadelige effecten heeft. Niet alleen bevat de geschatte halfwaardetijd flinke meetfouten, ook de onzekerheid van deze schatting is hoog waardoor het onderscheidingsvermogen van de toets al erg beperkt is. Ook verloopt het onderscheidingsvermogen niet monotoon als de halfwaardetijd onder de alternatieve hypothese meer afstand neemt van de nulhypothese dat tot misleidende resultaten voor de validiteit van de KKP hypothese zou kunnen leiden.

Voor de andere twee modellen die heterogeen worden geschat, is het onderscheidingsvermogen van de eenheidsworteltoets wel een stijgende functie. Hoe groot het onderscheidinsvermogen is, hangt ook af van het aantal stationaire wisselkoersen in het panel. Voor het panel in dit hoofdstuk, dat bestaat uit de reële wisselkoersen voor Canada, het euro-gebied, Japan en Groot-Brittannië met de Amerikaanse dollar als numeraire, geldt dat het onderscheidingsvermogen van de heterogene toets het grootst als er drie of vier niet-stationaire wisselkoersen zijn. Als er maximaal één wisselkoers is die niet stationair is dan doet het model dat heterogeen schat maar wel homogeen toetst qua onderscheidingsvermogen het het beste. Als er twee wisselkoersen stationair zijn dan is er nauwelijks verschil. Het onderscheidingsvermogen van het model dat zowel heterogeen schat als toetst hangt niet af van het aantal stationaire wisselkoersen in het panel maar wel van de valuta die als numeraire wordt gebruikt.

De ongedekte interest pariteit (OIP) die in hoofdstukken 4 en 5 behandeld wordt veronderstelt dat de verwachte verandering in de wisselkoers gelijk is aan de termijn premie (verschil tussen de termijnwisselkoers en de spot wisselkoers). Deze termijn premie is gelijk aan het verschil in de rentevoeten van de onderliggende landen.

Het meest gebruikte model om deze pariteit te onderzoeken is het lineaire regressie model van Fama (1984) waar het gerealiseerde rendement van de wisselkoers geregresseerd wordt op de termijn premie. De ongedekte interest pariteit gaat op als de richtingscoëfficiënt statistisch gelijk is aan één. Echter, de literatuur vindt dat deze vaak significant kleiner is dan één en vaak ook negatief zodat bewijs voor de pariteit niet wordt gevonden (zie Hodrick (1987) en Engel (1996) voor een literatuuroverzicht).

Verscheidene pogingen om deze puzzel op te lossen zijn gedaan (zie Sarno en Taylor (2003) en Sarno (2005)). Bijvoorbeeld Bansal (1997) en Wu en Zhang (1996) onderzoeken of een positieve termijn premie een ander effect heeft op het verwachte wisselkoers rendement dan een negatieve termijn premie. Ook bekijkt Bansal (1997) of er een linaire relatie bestaat tussen een niet-lineaire transformatie van de termijn premie en de verwachte verandering in de wisselkoers. Andere studies onderzoeken of er een niet-lineaire relatie bestaat tussen het toekomstige wisselkoers rendement en de termijn premie. Voorbeelden van deze studies zijn Baillie en Kiliç (2006) en Sarno *et al.* (2005) die gebruikmaken van een niet-lineair model dat verschillende regimes onderscheidt waarin de ongedekte interest pariteit wel danwel niet opgaat.

In hoofdstuk 4 wordt de prestatie van lineaire modellen onderzocht. Het model van Clarida en Taylor (1997) vormt de basis voor de analyse. Zij tonen aan dat hun model, dat informatie van de rente termijn structuur gebruikt, het veel beter doet dan het standaard Random Walk model wat opmerkelijk is omdat veel andere studies niet in staat zijn om dit te doen. Het model van Clarida en Taylor (1997) wordt in dit hoofdstuk uitgebreid zodat rekening kan worden gehouden met de bevindingen van bovengenoemde Bansal (1997) en Wu en Zhang (1996). De bevindingen van dit hoofdstuk bevestigen de resultaten van Clarida en Taylor (1997) niet. Ook de uitbreiding van het model met asymmetrische en niet-lineaire transformaties van de termijn premie helpen niet om een betere beschrijving van de verandering in de wisselkoers te geven.

In hoofdstuk 5 wordt bekeken of er een mogelijk niet-lineair verband tussen de termijn premie en de verandering in de wisselkoers bestaat. Het niet-lineaire model van Baillie en Kiliç (2006), dat twee verschillende regimes onderscheidt, wordt als uitgangspunt genomen en aangepast. In dit hoofdstuk wordt in het ene regime de ongedekte interest pariteit als waar aangenomen terwijl mogelijke niet-lineariteiten door het tweede regime kunnen worden opgepikt. Om te toetsen of OIP opgaat worden er betrouwbaarheidsintervallen voor het niet-lineaire gedeelte van het model geconstrueerd. De bevindingen zijn dat de geschatte betrouwbaarheidsintervallen te breed zijn om OIP te kunnen verwerpen.

Om te onderzoeken of de mogelijke afhankelijkheid tussen de wisselkoersen effect heeft op deze resultaten wordt ook een multivariate analyse uitgevoerd door de niet-lineaire vergelijkingen in een panel op te nemen. Ook wordt bekeken wat de invloed is van de gemeenschappelijke Amerikaanse rente in de vergelijkingen van dit panel. De bevindingen op basis van de geschatte modellen zijn dat deze uitbreidingen geen effect hebben op de conclusies zodat OIP nog steeds niet verworpen kan worden.

Hoofdstuk 6 zal het meten van volatiliteit van wisselkoersen behandelen. In de huidige literatuur wordt van drie populaire modellen gebruik gemaakt om volatiliteit te meten. Deze zijn de zogenaamde realized volatility (RV) modellen (zie o.a. Andersen *et al.* (2001b) en Andersen *et al.* (2001a)), de GARCH modellen (zie Engel (1982) en Bollerslev (1986)) en de Stochastic Volatility (SV) modellen (zie Taylor (1986)). De twee laatstgenoemde modellen zijn parametrisch waarvan de GARCH modellen over het algemeen makkelijker te schatten zijn dan de SV modellen. Van beide modellen bestaan multivariate versies. Deze modellen zijn nog moeilijker te schatten, vooral omdat het aantal parameters toeneemt.

In dit hoofdstuk van dit proefschrift wordt een multivariaat SV model geponeerd dat gebruik maakt van de minimum en maximum wisselkoers over een bepaalde periode in plaats van de normaliter gebruikte rendement data. Met behulp van deze data wordt de zogenaamde "range"volatiliteitsschatter geconstrueerd. De range is niet alleen een nauwkeurige schatter (zie Andersen en Bollerslev (1998) en Alizadeh *et al.* (2002)) maar de verdeling van de range heeft voordelige eigenschappen. Deze eigenschappen kunnen gebruikt worden bij het schatten van het multivariate SV model.

Verder worden valuta-specifieke nieuws factoren onttrokken van de wisselkoersen zoals gedaan is in Mahieu en Schotman (1994). De bevindingen zijn dat deze geschatte factoren substantieel van elkaar verschillen en ze in staat zijn om valuta-specifieke gebeurtenissen op te pikken. Bijvoorbeeld, de bron van de volatiliteit van de wisselkoersen bij de Azië-crisis en bij de val van de pond sterling van het EMS is terug te vinden in de componenten van respectievelijk de yen en de pond.

Biography

Ben Tims was born on 9 November 1975 in Rotterdam, the Netherlands. From 1994 till 1999 he studied econometrics at the Erasmus University Rotterdam. From September 1999 to December 2004 he has been a PhD student at the Financial Management department of the Rotterdam School of Management at Erasmus University Rotterdam. Since January 2005 he has been assistant professor at the same department. His research interests focus primarily on the dynamics on exchange rates.

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