International Exchange Rate Dynamics and Purchasing Power Parity

A Thesis submitted for the Degree of Doctor of Philosophy

by

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ABSTRACT

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This thesis provides evidence in favour of the long-run validity of Purchasing Power Parity (PPP) using primarily a linear error correction framework. Through an examination of PPP where proportionality and symmetry are implicitly imposed, it is shown that a selection of twelve EU real exchange rates is stationary on a univariate basis. The contribution here is based on the reconciliation of unit root test outcomes across univariate and panel tests. Following this analysis, the Johansen cointegration procedure is employed to examine whether long-run equilibrium relationships can be identified in systems of real exchange rates. The implications of results found are set out in terms of regional exchange rate policy co-ordination, exchange rate regime appropriateness, and monetary integration. By focussing on interdependent regions that were affected by a major financial shock (Europe: EMS crisis; Latin America: Mexican crisis; South East Asia: 1997 crisis), the real exchange rate dynamics are compared in pre- and post-crisis scenarios.

This thesis also presents evidence in favour of PPP by examining the less restrictive scenario where neither proportionality nor symmetry is imposed. Given the fact that most developed economies have highly integrated goods and capital markets and liberalised capital accounts, the failure to find evidence for PPP in previous studies may be due to the exclusion of factors that might reflect the behaviour of capital markets and their influence on the exchange rate. To test this, the traditional nominal exchange rate and domestic/foreign price based system is augmented with an interest rate component. In a tripolar specification, the joint test of PPP and Uncovered Interest Parity (UIP) is found to hold in a system comprising Germany, Denmark and the UK, suggesting well-integrated goods and capital markets and the long-run convergence evident suggests that Denmark and the UK might be suitable for membership of the euro area. This convergence appears to be stronger when short-term interest rates are used as opposed to long-term rates (perhaps since they are not subject to distortions such as taxation and maturity levels). Furthermore, long-rates have been associated recently with an inversion of the yield curve, while evidence to support the yield curve in non-crisis times is mixed. Finally, multivariate and panel cointegration procedures are employed to provide evidence for the suitability of potential future euro area entrants from Central and Eastern Europe in tri-variate systems comprising the euro nominal exchange rate and two price series.

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DECLARATION

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CHAPTER ONE

INTRODUCTION

This thesis is centred on the issue of exchange rate behaviour across developed, transition and developing economies. The broad context for the thesis lies in determining the appropriateness of Purchasing Power Parity (PPP) as a theory of exchange rate determination. The validity of PPP is a key underlying assumption in the majority of exchange rate and open economy macro models and one of the primary doctrines in the international finance literature. PPP, however, has received at best mixed support in empirical studies of the past (Breuer, 1994; Froot and Rogoff, 1995; Sarno and Taylor, 2002).[1] Clearly, a failure of PPP would have significant implications for the range of subsequent theories in international economics where PPP is assumed to be a valid proposition.

The importance of determining whether or not PPP holds is at the heart of this thesis; this is so since PPP is often a key component of both theoretical (see Dornbusch, 1976 and Taylor, 1995) and large-scale macro-models (for example see Wallis et al, 1984). Often this proposition is imposed in a linear or log-linear form rather than tested or estimated empirically. However, the validity of these types of models depends crucially on whether the PPP equation in the models is appropriate. If PPP is not a valid assumption, these models would require a different explanation of exchange rates. In addition, when PPP holds with augmentation, then the nature of appropriate specification and augmentation ought to be well understood, because this may impact on the speed at which prices adjust and payments problems are corrected for. This thesis intends to provide a comprehensive study of PPP in the context of primarily linear error correction models (ECMs).

This thesis presents evidence to show that PPP is, in fact, a valid proposition, albeit in a long-run sense. To do so, empirical analyses are carried out to show that real exchange

[1] This refers to the empirical work undertaken in macro studies. By contrast, a large segment of microeconomic studies on PPP for tradable goods suggest that it holds.

rates are stationary; that co-integration can be found in systems comprising the nominal exchange rate, domestic prices, and foreign prices; that the joint modelling of PPP and Uncovered Interest Parity (UIP) produces a stationary outcome; and that the stationarity in systems of real exchange rates is more prevalent for crisis-affected regions in post-crisis periods. Furthermore, we also endeavour to explain why PPP may have been found not to hold in previous studies. The policy implications of the thesis pertain to regional exchange rate policy co-ordination, exchange rate regime appropriateness, and monetary integration.

A number of recent studies have suggested that non-linearity might be key to explaining the failure to find PPP. Here, it is shown that misspecification of the short-run and long-run relations might be a more pertinent reason for this failure. It should be noticed that error correction terms are embedded in dynamic systems that are differenced and there is a long tradition in economics of using the first difference as an approximation or linearization of non-linear equations (Tobin, 1950, Barten, 1969, Deaton, 1975 and Deaton and Muellbauer, 1980). In this light the error correction term, when it defines a stationary relation, provides inference about PPP and a higher order explanation of the phenomenon. Further, a primary condition given in Granger and Terisverta (1990) for balance is stationarity.

The linear approach is rationalised as an approach, because it can approximate non-linear models, has well defined systems representations and appropriate inferential methods to test hypotheses and model specification (see Hendry, 1995 and Johansen, 1995). In particular, both single equations and systems can be tested for misspecification that is caused by non-linearities. Thus, such non-linear models are not considered in this thesis. A paradigm of parameter linearity is pursued to show that PPP holds. Such an approach enables a coherent framework to be adopted as one progresses through each Chapter, from single through to multiple equation estimations. As a result, a non-linear approach is not undertaken since there currently exist no coherent or consistent multi-equation non-linear estimators. Specifically, while it would be possible to undertake the estimation of univariate or single equation non-linear models, this could not be carried out at the system level. In addition, by following a linear approach, a wide range of conventional techniques can be applied to test for exogeneity, causality and forecast performance. Impulse response function analysis at the system level is a further rationale for the linear as opposed to non-linear approach followed.

Nonetheless, the issue of non-linearity is not ignored. Implicitly, error correction models take account of non-linearity as they provide first-order Taylor series approximations to any non-linear model. Furthermore, linear regression based estimators are often relatively robust and there are a number of reasons why this might not be the case for non-linear models. First, there is a tendency for non-linear models of unknown parametric form such as neural networks that define generalized semi and non-parametric estimators (White, 1988) to over-fit the data. Second, non-linear models are often sensitive to a small number of observations. Third, inference is often not well defined for neural networks or for STAR and ESTAR models (Nielsen, 2000). With the potential for over-fitting data and the sensitivity to a small number of observations, these anomalous observations might be better handled by dummy variables. Diebold (1986) suggests the use of dummies to correct for volatility prior to the application of ARCH and GARCH corrections (for example see Gregoriou, Hunter and Wu, 2009). Over-modelling these unpredictable observations may well yield relations that forecast poorly out of sample.

A feature that is apparent from the literature on PPP that arises from testing stationarity of the real exchange rate, is that panel tests appear to show that PPP holds while on average this is often not true for equivalent tests based on a sequence of univariate analyses based on the same data. However, typical panel tests provide a mean correction to standardise the data. This transformation corrects for the impact of initial conditions and, as a result, the tests are no longer similar. There is also a strong presumption of heterogeneity when a panel approach is applied. Conventional panel estimation provides corrections for fixed and random effects, but these constant or mean zero corrections might be difficult to sustain across a country specific sample, let alone a panel of diverse country data. More specifically, in panel tests, to enhance the heterogeneity that underlies the analysis, the country specific data stacked in the panel is analysed relative to group means (Im et al, 2003), while in univariate cases such tests have often not applied any type of mean correction.[2]

Here, a sequence of mean-adjusted real exchange rates are calculated for developed economies in the EU. We apply conventional Augmented Dickey Fuller tests on data that is recursively mean adjusted. As a result the intercept is not statistically significant and this gives rise to a test that is similar in the univariate and panel case. The conclusion of both forms of analysis is that on average the exchange rate is stationary across the twelve countries considered. Furthermore, the mean corrections in both cases give rise to tests statistics that based on Monte-Carlo simulation have a similar size. In the univariate case all of the models are well defined in terms of serial correlation, and non-IID disturbances associated with excess kurtosis and heteroskedasticity are corrected by the use of alternative robust standard errors. In addition, the models are corrected for ARCH behaviour where volatility is observed. To understand standard panel unit root results on

[2] The rationale for de-meaning is to eliminate the influence of the intercept, thereby correcting for initial conditions (Tremayne, 2006; Haldrup and Jansson, 2006).

the real exchange rate, the pooling procedure implies that PPP holds on average. However, we have shown that by using unadjusted real exchange rate series in a univariate unit root test, PPP holds in only three out of twelve cases. Based on the unadjusted tests it is not feasible that the panel of twelve real exchange rates could be stationary on average. However, using appropriately transformed data, nine of the series are stationary when the univariate analysis is applied, which makes the panel results on PPP appear much more feasible. The univariate and panel tests are coherent when a similar sized test is applied in both cases and the outcomes of the tests are equivalent.

Model specification is often ignored or handled by an appeal to fit via some information criterion, but information criteria are not tests of model specification. In the univariate case, misspecification is often ignored or it is assumed that the specification can be improved by extending the dynamics. In the multivariate case, model misspecification is a key aspect of well defined tests. The tests will not perform well, because the mean or the variance equations are not well defined. Here, we mainly concentrate on the specification of the mean equation as multivariate ARCH and GARCH models under cointegration often perform poorly (Wickens, 1996, MacKinnon *et al,* 1999). While, Rahbek et al (2002) suggests that samples of 600-1000 observations are required to provide sensible inference on cointegration when there is a moderate amount of persistence.

If one concentrates on the mean equation then the specification of the system and the order of the dynamic are important. Clearly, univariate tests of the stationarity of real exchange rates are likely to omit a wide range of factors that might influence the exchange rate. The test equations, though not often treated this way define restricted error correction models, and there is much evidence that the test result is impacted upon by this type of misspecification (Kremers et al, 1992, Hansen, 1995 and Hendry, 1995). The very univariate or restricted bivariate nature of the tests is a rationale for the failure of previous attempts to test the proposition that PPP holds via studies that rely on an exchange rate and the relative prices of two counties alone.

Clearly, tests based on data for two countries alone fail to recognise the interactions and spillovers that can exist amongst the trading relations between countries that would suggests a study of systems of real exchange rates. Such a system amounts to a multi-country test of PPP and is, arguably, more appropriate in the current globalized international economy where the exchange rate of one particular economy may be driven by more than one country interaction. Even when one considers a two-country analysis based on prices and the exchange rate, this implies that the exchange rate is driven purely by trade relations. It ignores the importance of capital flows that are a key factor in determining daily currency flows. Hence, any limited information method used to explain PPP in the long-run may only be well defined when the exchange rate is cointegrating exogenous (Hunter, 1990) for interest rates. Hence, a test of cointegration based on the Philips modified approach (Phillips, 1994), the spectral approach of Marinucci and Robinson (2001) or the Engle-Granger approach (Engle and Granger, 1987) will normally only be appropriate in this instance.

Tests of cointegration using a trivariate system including exchange rates and prices alone require the sub-system involving prices to be invariant in the long-run to the behaviour of long-run interest rates and to a sequential cut in the long-run parameter space between the long-run interest rate relations and the equation associated with PPP (Hunter, 1992, and Ericsson and Irons, 1994). The prevailing evidence on cointegration based on such systems using the methodology of Johansen (1991) would suggest that this is not the case. For evidence on cointegrating and weak exogeneity for the UK see Burke and Hunter (2005), and Hunter (1992), while for studies that suggest a long-run co-dependence between PPP and interest rates see Johansen and Juselius (1992) for the UK, Juselius (1995) for Denmark and Simpson (2002) for Germany and Italy. This leads us in this study to analyse such inter-relations for a number of developed economies. The joint hypotheses of PPP and UIP are tested in bipolar and tripolar models.

Chapter 2 comprises a number of elements. A detailed outline of the theoretical considerations surrounding PPP is described. Such theory forms the basis for the empirical tests that are carried out in subsequent chapters 3, 4 and 5. This chapter also presents a major review of the literature surrounding PPP and real exchange rate stationarity. Econometric analysis is undertaken to critique the types of testing procedures that were undertaken in the past, thereby casting some doubt on earlier findings. This chapter provides an outline of the previous literature on PPP and exchange rate movements. The overall aim of the chapter is to provide a comprehensive assessment of previous studies and to provide some innovative analysis as to whether PPP holds by using univariate and panel methods in a coherent framework. The analysis is carried out to identify the existence of PPP across twelve real exchange rates. Specifically, the univariate series are recursively de-meaned and careful consideration is given to the dynamic forcing the real exchange rate. The study tests for stationarity using a similarly sized approach for both univariate and panel-based tests. [3]

Chapter 3 examines the prevalence of PPP in a multi-country context. Here the analysis is divided into sub-country blocks that have experienced similar economic conditions.

[3] Some of the findings presented here are published in Beirne, Hunter and Simpson (2007) and were presented at the 3rd International Conference on Advances in Financial Economics at the Research and Training Institute of East Aegean (INEAG), Samos, Greece, in July 2006.

Also it would seem reasonable to analyse their behaviour relative to the $ and via an analysis that is implicitly conditioned on US prices. Hence, three economic regions are considered: Europe over the period 1980 to 2006, Latin America over the period 1983 to 2006, and South East Asia over the period 1988 to 2006. This chapter provides an illustration of PPP, considered in a multi-country setting to determine whether the proposition can be accepted when the tests give rise to long-run relations between relative real exchange rates that are as a result stationary. The economic significance of such relations is assessed in terms of the different dynamics in the systems associated with pre- and post-crisis scenarios across the different regions. The empirical work is based on the Generalised PPP theory of Enders and Hurn (1994). The aim is to examine whether the crisis has led to any change in the PPP relation in the post-crisis period. In addition, while many studies have been carried out on the causes of crises, there has not been a substantial amount of research completed on the post-crisis effects. This aspect of the thesis utilises information from the real exchange rate dynamic of the relevant regions to extrapolate the degree of exchange rate policy co-ordination and monetary integration. The empirical work undertaken is supplemented by impulse response function analysis to assess the adjustment of the real exchange rate in the crisis-affected economies to unexpected temporary shocks to real output growth.[4]

Chapter 4 concentrates on the sub-system information required to tests a range of dollar and cross rate parity conditions to provide further evidence that might draw a veil over the PPP puzzle and show why PPP has not been observed in so many studies. In this case, the price based system is developed and augmented by interest rates to jointly test PPP

[4] An academic paper-length version of this chapter has been presented at the 6th Annual INFINITI Conference in Trinity College Dublin, June 9-10th 2008, and at the 12th International Conference on Macroeconomic Analysis and International Finance, University of Crete, 29th-31st May, 2008. The paper appears in the December 2008 issue of the *Journal of International and Global Economic Studies* (Beirne, 2008).

and UIP. This is based on the idea that most developed economies have highly integrated goods and capital markets, and liberalised capital accounts and thus failure of PPP could arise in the long-run by virtue of the exclusion of factors that might reflect the behaviour of capital markets and their influence on the exchange rate. This chapter utilises multivariate cointegration analysis to test for PPP in the long-run, using a five-variable vector autoregression (VAR) containing a bilateral nominal exchange rate, two price series, and two interest rate series. The Johansen procedure is employed to test for PPP and UIP in this system. Following Juselius (1996), the case of Denmark and Germany is explored using what Juselius and Macdonald (2004) has termed a bipolar specification. The analysis is extended to the tripolar case by the addition of UK data in an eight-variable VAR.

Often, macroeconomic models are closed by imposing a PPP or UIP condition to explain the exchange rate. However, the theoretical specification due to Frydman and Goldberg (2002, 2006) indicates that PPP and UIP cannot be assessed independently of each other, since the balance of payments requires a trade off between the current and capital accounts that relies on the interplay between PPP and UIP in the long-run. The empirical models estimated in this chapter provide evidence that for PPP to hold, augmentation by an interest rate component is essential. In the tripolar case, the parity conditions hold jointly when short-term interest rates are used. This may indicate that short-term interest rates are a more appropriate proxy for capital flows than long-term rates (perhaps since they are not subject to distortions such as taxation and maturity levels). Furthermore, long-rates have been associated recently with an inversion of yield curve, while evidence to support the yield curve during non-crisis times is mixed. This chapter not only empirically demonstrates the apparent interdependence of PPP and UIP, but the tripolar model estimated suggests that Germany, Denmark and the UK have well-integrated goods and capital markets. Such long-run convergence provides some indication that Denmark and UK might be financially well integrated into the European Monetary System.

Chapter 5 provides an analysis of the exchange rate and price dynamic in an enlarging euro area in an attempt to determine the suitability of entry into the single currency zone. Tests of PPP are carried out using multivariate and panel cointegration procedures for the period since the introduction of the euro currency. The data considered consists of a selection of developed countries in the European Union (EU) but not in the euro area, and a selection of transition economies that have been gearing themselves up for entry into the euro area since its introduction. Here, the systems approach is comprised of Johansen cointegration tests for a panel of fifteen countries each in turn represented by tri-variate VARs. Each country VAR comprises the nominal exchange rate, the domestic price level and the foreign price level. The panel method is based on the approach of Larsson et al (2001). Of the ten cases where long-run weak-form PPP is found to hold, the driver of the convergence to exchange rate equilibrium varies by country depending on the nature of the exchange rate regime in place. Where the nominal exchange rate is fixed, domestic prices drive the adjustment to PPP. Where a floating regime is in place, the exchange rate is the variable that forces convergence to stationarity. In recognition of the potential problems that arise from an analysis of cross rates due triangular arbitrage, results are provided using both the euro and the US dollar as the numeraire.[5]

[5] This research has been presented at the Money Macro and Finance Research Group 39th Annual Conference, University of Birmingham, UK, 12-14th September, 2007; and the 6th Annual Meeting of the European Economics and Finance Society on *European and Global Integration: Underlying Causes, Issues Arising and Formulating Economic Policies*, Sofia, Bulgaria, 31st May-3rd June, 2007. It was also presented at the 66th Annual Conference of the International Atlantic Economic Society in Montreal, Canada (October 9-12th 2008). The paper has recently been re-submitted to the *International Review of Applied Economics* journal. In addition, a further element of Chapter 5 is appears as a research note in *International Advances in Economic Research* (Beirne, 2009).

Overall, this thesis provides evidence for the validity of PPP across a range of different economies (developed, transition, developing) using a linear approach through the estimation of various forms ECMs. The thesis contributes to the literature on exchange rate dynamics in three key respects. Firstly, using an innovative statistical adjustment to univariate real exchange rate series, it is shown that the correct conclusion of our unit root test results from univariate and panel procedures is that on average the real exchange rate can be viewed as stationary, challenging the view that univariate and panel tests do not produce similar results. Previously, panel estimators could indicate stationarity even when the majority of the individual series appeared to contain a unit root. Thus, when appropriately adjusted to compare with panel estimators, univariate tests of stationarity in real exchange rates may on their own validate traditional PPP.

Secondly, further evidence is found in support of PPP when interactions are permitted amongst the real exchange rates of interdependent economies. The implication here is that for highly interdependent economic regions, the traditional two-country test of PPP may not be sufficient given that any particular bilateral real exchange rate will be affected by more than the domestic and foreign prices of one trading partner. An application of this 'multi-country' test of PPP shows that convergence to PPP is more apparent in post-crisis periods. The implication is that crisis-affected economies learn that some form of co-ordination on exchange rate policy is desirable between trading partners to reduce exchange rate volatility and risk. This analysis also shows that the existence of a long-run equilibrium amongst real exchange rates is not sufficient to endorse monetary integration; this conclusion is in line previous work done in this area. While the real exchange rates cointegrate, recommendations on monetary integration require a deeper analysis of the cointegrating equations in terms of the degree to which symmetry is observed in the long run response to shocks, and the width within which the regime operates. For example, it is possible to observe a long-run equilibrium amongst real exchange rates, while the individual exchange rates in the long run relations have different signs indicating asymmetry in the long run responses to shocks. Differences in the size of coefficients, subject to a particular normalisation[6] indicate that the regime operates within different bands. In this case, while a fixed regime may appear to be suitable (given that a long-run equilibrium exists), the level of variability permitted and the size of the band may mean that the regime operates as a de facto float.

Thirdly, evidence for PPP base on traditional tests based on univariate studies of real exchange rates or price and exchange based systems might not be sufficient to test this proposition in a global economy where international capital flows influence the exchange rate as well as trade flows. While there has been some work done on this in the past, the empirical work carried out in this chapter is differentiated from previous work given the extensive nature of the analysis across short and long-term interest rate scenarios. In particular, it appears from the tripolar model estimated that short-term rates are a more suitable proxy for capital flows. In addition, in the context of the potentially enlarging euro area in the future, the results from the tripolar model estimated have useful implications for the suitability of Denmark and the UK for EMU membership.

The final empirical chapter validates PPP for the forthcoming euro area economies using both multivariate and panel cointegration procedures. This also has important policy implications for the appropriateness of euro area membership for these economies. It is pertinent to remember that the euro area has a single interest rate. There might be certain limitations on this, such as the operation of Landers Banks in Germany and the extent to which the Irish government has been able to deviate from the Maastricht conditions.

[6] It should also be decided whether certain effects are identified subject to the normalization. Excessively large coefficients may indicate that the normalization chosen is incorrect, Boswijk (1996) and Burke and Hunter (2005).

CHAPTER TWO

PURCHASING POWER PARITY AND REAL EXCHANGE RATE STATIONARITY

I. Introduction

This chapter comprises two core components. Firstly, a description and discussion is provided on theoretical aspects relating to Purchasing Power Parity (PPP), upon which hypotheses in subsequent chapters are tested empirically. This includes a discussion on the theoretical foundations of traditional PPP, an explanation of Generalised PPP theory (which permits PPP to be tested on a multi-country basis), and a discussion on theoretical considerations associated with the joint modelling of PPP with Uncovered Interest Parity. Secondly, I describe and critique various econometric techniques that have been used in the past to test PPP, and provide econometric evidence to resolve the discrepancy perceived in the literature between univariate and panel tests of real exchange rates.

Regarding the theoretical discussion, the theory of traditional PPP is described in the first instance, i.e. the two-country case where the bilateral nominal exchange rate equates to domestic and foreign prices. Following this, I present a form of PPP that is perhaps more appropriate in a more globalised economy, where the exchange rate of one economy is affected not only by one other trading partner, but by a number of trading partners. This 'multi-country' test of PPP is based on the Generalised PPP theory of Enders and Hurn (1994). This theory is described, as the basis for empirical testing in Chapter 3. Following this I present theoretical evidence to show that in certain circumstances, traditional PPP may require augmentation with an interest rate component as capital markets become more developed and the need develops to take into account not only trade flows, but also capital flows, in modelling the exchange rate. This theory is tested empirically in Chapter 4.

Secondly, following the theoretical discussion, a further basis for the remainder of the thesis is provided through a comprehensive description of the types of tests that have been used in the past to examine the validity of PPP empirically. Developments in econometric techniques and advances in econometric computing throughout the 1990s have led to arguably more robust results with regard to the issue of whether or not PPP holds. There still remains an absence of consensus however. In particular, there is no consensus as to whether the real exchange rate contains a unit root. Some of the more recent literature has explored the non-linear behaviour of real exchange rates and a discussion is provided of some non-linear estimation techniques which have been growing in popularity over the past five years or so as a possible avenue to solving the PPP puzzle, i.e. that the real exchange rate exhibits substantial volatility in the short-term but shocks appear to subside very slowly.[7]

As well as reviewing the types of tests employed in the past to measure PPP, this chapter explores two key issues that have not received much, if any, attention in the literature. Both issues are explored empirically using real exchange rate data from a selection of developed economies over the past quarter of a century. The first issue pays particular attention to the type of estimation conducted to inform the stationarity outcome. Through assessment of the econometric technique used and specification of the model, an attempt is made to construct a simple procedure that is more closely reflective of the overarching PPP theory. This relates specifically to the appropriateness of the intercept in estimation procedures from both a theoretical and statistical standpoint.

The second issue seeks to provide some explanation for why long-run purchasing power parity is not always observed, looking specifically at a statistical rationale for different stationarity results that can arise from a set of univariate series as compared with when a panel unit root test is conducted across all series. To this end, the coherence between univariate and panel-based tests of stationarity is examined using a time series of the real

[7] This PPP puzzle was first identified by Rogoff (1996). A further more general PPP puzzle exists as to why there exists no strong evidence to support the view that exchange rates actually converge to PPP.

exchange rates of twelve EU countries. Using recursively demeaned data for the univariate tests, it is shown that the univariate results are not largely different to those of the panel.[8]

II. Theoretical Considerations

II.A Theoretical Overview of PPP and Extensions

Purchasing Power Parity (PPP) is one of the key doctrines in the field of international economics. In simple terms, PPP states that the nominal exchange rate between two countries should be equivalent to the ratio of the two relevant price series. The idea has its underpinnings in the proposition that a common basket of goods should cost the same in both countries when prices are quoted in the same currency. As well as being an equilibrium condition and a theory of exchange rate determination, PPP also provides a means by which output across countries can be compared.

PPP has been the subject of a substantial amount of academic research over the past thirty years or so, in the period following the shift to the flexible exchange rate regime. Whilst the general view is that PPP holds in the long-run, there is no unanimity across the economics profession that this is the case. Some studies have found PPP to hold, while others have rejected the concept. Clarity on this issue could have a range of economic implications. For example, the main impulses driving exchange rates can be determined by the degree of persistence in the real exchange rate. When the real exchange exhibits random walk behaviour, shocks are likely to be supply-side based. On the other hand, a low level of persistence would be more related to demand-based factors (Rogoff, 1996). In addition, from a more practical perspective, knowledge of whether or not PPP holds has strong policy implications. As is commonly understood, PPP exchange rates are used to identify nominal

[8] The empirical work carried out on this issue has been published in *Quantitative and Qualitative Analysis in Social Sciences* as "Is the Real Exchange Rate Stationary? – A Similar Sized Test Approach for the Univariate and Panel Cases" (Vol. 1(2), pp. 55-70, 2007). This paper was carried out in collaboration with John Hunter and Mark Simpson.

exchange rate misalignments and to compare income levels across regions and countries. Clearly, if the real exchange rate is found to contain a unit root (i.e. non mean-reverting), then this would greatly affect the usefulness of the PPP concept.

There exists a large literature criticising PPP as a theory of exchange rate determination.[9] It is understood that PPP must hold if the prices of all goods are equalized internationally by arbitrage (i.e. the exploitation of price differences to yield a riskless profit), leading to the Law of One Price. However, some goods are not traded at all. Even when goods are freely traded, the price to final consumer is often affected by local distribution costs. A large segment of the micro literature on PPP for tradable goods would suggest that it holds (Gilbert and Kravis, 1954; Marris, 1984). However, the macro literature is more mixed as regards the validity of PPP as a long-run equilibrium concept.

Perhaps of particular significance as a rationale for PPP not holding was work undertaken by Balassa (1964). Using some empirical analysis completed by Gilbert and Kravis (1954), Balassa (1964) noted that the price of services tended to be relatively higher in developed countries[10], and subsequently, that differences in productivity growth between developing and developed countries could generate a deviation from PPP. Subsequent work by Kravis and Lipsey (1983) showed that the ratio of prices of non-traded goods to traded goods was higher in developed than developing countries, thereby adding some weight to previous arguments made regarding productivity.

Up until recently the evidence in favour of the PPP has been far from conclusive, with results being especially mixed for the post Bretton-Woods period. In **Section III** we find that this failure might be due to the specification of the univariate models used and as a result of a more careful analysis of the dynamics we concur with Luintel (2001) that, on average across

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[9] For example, De Grauwe (1989); Isard (1977, 1987); Katseli-Papaefstratiou (1979); Krueger (1983); MacDonald (1988); McKinnon (1979); and Officer (1976).

[10] Marris (1984) provides further evidence in support of this point.

developed countries, PPP would seem to hold without the need for non-linear estimators. However, the univariate analysis is limited and provides some support to the proposition that PPP holds, but both the conclusions of single equation tests of stationarity and multi-country panel studies must be treated with a degree of circumspection. Firstly, the univariate ADF test has an implicit common factor restriction that implies a potential misspecification that may affect the size and power of the test (Hansen, 1995). Secondly, breaking the common factor restriction reveals the implicit multivariate nature of the structure. Error correction models might be estimated or the ADF test augmented to correct for the common factor problem (Burke and Hunter, 2005, Chapter 2), but beyond an exact bi-variate case (Engle and Granger, 1987), tests of cointegration are different when the problem is tri-variate. However, as shown in **Chapter 3** and **4**, the multivariate dimension that is relevant for investigation becomes a matter of debate. In **Chapter 4**, the primary extension is via the introduction of financial variables and the consideration of the existence of further parity conditions. Failure to observe PPP clearly has consequences for the practical application of theories of exchange rate determination. This would also have implications for other fixed and flexible monetary models of exchange rate determination that assume PPP to hold. The former case is considered by Dornbusch (1976) and the latter Bilson (1978).

II.A1 Traditional PPP Theory

The concept of PPP has its foundations in the work of Cassell (1922). This idea holds that the exchange rate for any two currencies equals the ratio of the prices of goods in the countries. In other words, the purchasing power of money should be the same across countries when the price of goods is measured in the same currency. The reason for this is arbitrage, i.e. the exploitation of price differences to yield a riskless profit. The PPP condition only holds in the long-run. As was pointed by Cassell (1928), factors such as capital flows are likely to keep the exchange rate away from its equilibrium level in the short to medium run. There are two main forms of PPP – Absolute PPP and Relative PPP.

Absolute PPP refers to the idea that a basket of goods in the domestic country should be equal in price to an identical basket of goods in a foreign country when the price of the foreign basket has been adjusted using the exchange rate. Therefore, the absolute version asserts that the nominal exchange rate is equivalent to the ratio of the domestic and foreign price levels, or the real exchange rate equals one (i.e. deviations from parity are stationary whereby the real exchange rate does not contain a unit root). This could be denoted as:

$$S = P/P^* \tag{1}$$

where S is the exchange rate (domestic currency per unit of foreign currency), P is the domestic price level, and P* is the foreign price level.

The version of absolute PPP described above has become known as 'strong-form' PPP. In reality, it is generally acknowledged that this version of PPP is not likely to hold across countries due to hindrances such as transport costs and trade tariffs. More likely to hold, it is argued, is a weaker form of parity arrangement, namely 'weak-form' PPP. In this case, symmetry across the domestic and foreign price levels is sufficient for PPP to hold (as opposed to the strong-form case where domestic and foreign prices should be equal to 1 and -1 respectively). More detail is provided on the distinction between strong form and weak form PPP below in Section II.B3, including a discussion on how the theories can be tested econometrically.

An alternative version of the PPP theory is termed 'Relative' PPP. This states that the exchange rate adjusts by the amount of the inflation differential between countries, i.e. that changes in the exchange rate are equal to changes in the relative national prices. The implication here is that the movement in the domestic price is proportional to the movement in the foreign price adjusted by the exchange rate, or simply that the real exchange rate is constant.[11] This can be denoted as follows:

$$\%\Delta S = \%\Delta P - \%\Delta P^* \tag{2}$$

where $\%\Delta S$ is the percentage change in the exchange rate (defined as the amount of domestic currency per one unit of foreign currency), $\%\Delta P$ is the percentage change in the domestic rate of inflation, and $\%\Delta P^*$ is the percentage change in the foreign rate of inflation.

The empirical work carried out in this thesis focuses on the strong and weak versions of absolute PPP. The time series properties of the variables examined means that the cointegration analysis undertaken (e.g. in **Chapters 3** to **6**) seems appropriate. The relative version of PPP, however, is based on first-differenced transformations and thus a cointegration analysis is not the most natural methodology to employ. Therefore, to maintain consistency in the methodological approach across the chapters, relative PPP is not considered in this thesis. Moreover, one could argue that important 'levels' information is lost in tests of relative PPP.

II.A2 Generalised PPP Theory

Generalised PPP (G-PPP) provides a means by which to test for PPP in a multi-country setting, whereby a bilateral real exchange rate can be affected by the real exchange rates of a number of trading partners. The theory described in the following paragraphs provides the basis for the hypotheses examined in Chapter 3. Following Enders and Hurn (1994) and Ogawa and Kawasaki (2001), the initial supposition is that we have a group of countries that

[11] Like the case of 'absolute' PPP, 'relative' PPP can have stronger and weaker versions, the former requiring that changes in domestic and foreign prices have a proportional and symmetric effect on exchange rate changes, while the latter requires that only the symmetry condition needs to be satisfied.

are geographically close to each other. This group of countries can be classed as an economic area. Furthermore, it is assumed that there is a large country outside of this economic area that impacts upon trade and capital flows in the economic area (this is represented as the US in the context of the empirical work carried out in the chapter). A market clearing scenario also applies whereby aggregate demand and aggregate supply are equal to each other. Since global flows of trade and capital impact upon the aggregate demand of members of the economic area, aggregate demand in one member is closely linked to output in other members. Similarly, the real exchange rate and real interest rate also affect aggregate demand. Following the notation of Ogawa and Kawasaki (2001), the dependence of aggregate demand in each country upon output, the real exchange rate, and the real interest rate of other countries in the economic area can be described as follows:

$$y_{jt} = \Sigma_{i=1}^{m+1}\theta_{ji}y_{it} + \Sigma_{i=1,i\neq j}^{m+1}\eta_{ji}r_{jit} - \tau_j i_t \qquad j = 1,\ldots,m+1 \qquad (3)$$

where is y_{it} the logarithm of GDP in country i, r_{ji} is the logarithm of the real exchange rate of country j's currency relative to that of country i, θ represents the propensity to import from country j, η is the price elasticity of demand, and τ reflects the sensitivity of aggregate demand to the interest rate.

In this framework, the real exchange rate is assumed to be non-stationary. This can be rationalised on the basis that even if one accepts that PPP holds, shocks that affect the real GDP of countries asymmetrically will cause deviations from PPP. Where these shocks follow a stochastic process, the real exchange rate should contain a unit root. Following the notation of Enders and Hurn (1994), G-PPP can be described as follows:

$$r_{12t} = \alpha + \beta_{13}r_{13t} + \beta_{14}r_{14t} + \beta_{15}r_{15t} + \ldots.. + \beta_{1m}r_{1mt} + \varepsilon_t \qquad (4)$$

where r_{1it} is the log of the bilateral real exchange rate in period t between country i and country 1; α is the intercept term; β_{1i} are the parameters of the cointegrating vector representing the degree of co-movement of the real exchange rate and ε_t is a stationary stochastic disturbance term.

Equation (4) describes the interdependence that exists between the real exchange rates of members of the economic area. For example, a real economy shock that affects Germany and the US asymmetrically will lead to movement in the real exchange rate (DM/US$). This fluctuation then spills over to other members of the economic area. This most pronounced where trade linkages are particularly high between countries of the region.

Ogawa and Kawasaki (2001) proceed to transform equation (3) into a vector representation. This is transformed in terms of the effect on the real exchange rate. Following this, it is shown that the components of (3) are cointegrated. Citing Stock and Watson (1988), the authors convert equation (3) into an equation that contains m+1 common trends. This is represented as follows:

$$Y_t = \delta\varphi_t \tag{5}$$

where Y is a vector of aggregate demands of each country, δ is an (m+1)(m+1) matrix (non-stationary),φ is an m+1 vector containing non-stationary stochastic trends. The transformation of A3.1 (to an equation like $R_t = AY_t$) clearly shows that the real exchange rate is affected by not only aggregate demand in each country of the region, but also propensity to import, price elasticity of demand, and the interest rate in other countries. From equation (5), we can infer that the real exchange rates of the countries are affected by common output trends.

II.A3 PPP Augmented with an Interest Rate Component

Purchasing Power Parity

PPP has already been described at length as one of the best-known parity relationships in international finance, equating the nominal exchange rate and the ratio of domestic and foreign prices. To inform the construction of the joint PPP-UIP framework, let us firstly recall a standard PPP relationship:

$$e_t = p_t - p_t^* \qquad (6)$$

where e_t is the nominal exchange rate, p_t is the domestic price level, and p_t^* is the foreign price level. Equation (6) represents the strong form version of PPP, implying that when measured in the same currency denomination, the price of a common basket of goods is the same. Of course, this version of the theory assumes that there are no market frictions such as transaction costs and that arbitrage is costless and instantaneous.[12]

Uncovered Interest Parity

An alternative to the exchange rate being determined by relative prices is that it is explained by the interest rate spread. The UIP theory claims that the interest differential between two financial assets should be exactly offset by the expected rate of change in the exchange rate between the currencies over the period to maturity (Cumby and Obstfeld, 1988). This can be denoted as follows:

$$i_t - i_t^* = E(e_{t+1}) - e_t \qquad (7)$$

where i and i* are the nominal rates of interest denominated in domestic and foreign currencies. Camarero and Tamarit (1996) noted the importance of testing UIP, because when

[12] Refer to Chapter 2 for details on empirical evidence on strong form PPP.

there is complete diversification of foreign risk and market participants are risk-neutral, then a failure of UIP would be indicative of market inefficiency. The other main reason for the importance of UIP is due to the fact that the theory is a crucial building block in the international finance literature, underpinning the monetary model of the exchange rate, the Dornbusch (1976) overshooting model, and the Krugman (1991) target zone model. This would include the models cited above in the context of PPP, and also other models such as the Frankel (1979) real interest differential model.

UIP has received a lot of comment in the academic and financial press at the time of writing. This was due to the currency carry trade that re-emerged in the latter half of 2006. Carry trade occurs when investors sell or borrow low-yielding currencies to purchase higher yielding currencies. Essentially, the investor converts the funds from a low yielding currency into a high interest rate currency, and then lends this amount at the high interest rate. Of course, economic theory suggests that there should be no opportunity to profit from exploiting interest rate differentials. This is because when UIP holds the interest differential exactly reflects the rate at which investors expect the high yielding currency to depreciate vis-à-vis the low yielding currency. After the depreciation, the amount lent in the high yielding currency is less valuable with the capital loss cancelling out any interest gain. Therefore, when UIP holds the return from lending in the high yielding currency should be the same as the cost of borrowing in the low yielding currency. However, even if markets were efficient, arbitrage requires time and in this light the observation of any profit from the carry trade might simply be a result of arbitrage. Nonetheless, one might imagine that efficiency should be reflected in UIP holding in a long-run sense.

A range of studies have been undertaken to test the empirical validity of UIP without much success.[13] These studies include Baillie and McMahon (1990), Froot and Thaler (1990), Lewis (1995) and Engle (1996), and the suggestion of the research is that the failure of UIP is due to a number of reasons including excessive exchange rate instability (e.g. in the case of Mexico), expectations errors, and the presence of a time-varying risk premium that is correlated with the interest rate differential (this is defined as the ex-ante expected profit from carry trade). Testing for UIP often incorporates an auxiliary assumption associated with the formation of expectations (Hallwood and MacDonald, 2000). Generally it is assumed that expectations are Rational in the sense of Muth (1962). Previous empirical analysis undertaken by example Hacche and Townsend (1981), Davidson (1985) and Loopesko (1984) all make this assumption and strongly reject UIP. Cumby and Obstfeld (1988) arrive at the same conclusion by applying a test of the randomness to the deviations from UIP. These authors rationalize their negative findings by virtue of the fact that empirical tests of UIP require observations of expectations and the simplification often used is to assume that the forward rate equals the expected future spot rate. Hence, rejection of UIP arises, because this simplification can give rise to a time-varying risk premium. An alternative approach to using the forward rate as a measure of the rational expectations is to find direct measures of the expectations based on analysts' forecasts or survey data.[14] MacDonald and Torrance (1990) used this type of approach, but again could not find evidence in support of UIP. This leads Hallwood and MacDonald (2000) to conclude that the failure of UIP is due to the inaccuracy of the empirical measures of the expectations used. A more recent study by Chaboud and Wright (2005) examines UIP using a large dataset of 5-minute US dollar exchange rate data for the period 1988 to 2002 for: Japan, Germany/Euro area, Switzerland and the UK. They

[13] It should be noted that an interest rate parity condition does apply to the model analysed by Hunter (1992), but as was shown in Burke and Hunter (2005) the model used as an example of the notion of Cointegrating Exogeneity is not well specified. Furthermore, some of the series used by both Hunter (1992), and Johansen and Juselius (1992) might well be I(2), which has implications for the conventional Johansen methodology (e.g. Burke and Hunter (2005), Chapter 6).

[14] Such measures can be sourced from a range of financial services institutions.

find UIP holds when the tests are applied to very short windows of data (i.e. shorter than one day).[15] While with time horizons longer than one day, UIP fails. The authors argue that this occurs because the risk premium shrinks as that time interval becomes small thus allowing UIP to hold. It is also the case that UIP is often pre-programmed in dealing desks and used directly by dealers to determine their expectations of movements in the exchange rate. However, when the data are aggregated across trades and time, the exact correspondence is likely to break down. An earlier study by Taylor (1985) found support for UIP using trade by trade data directly observed from dealer transactions.

It is understood from a strict adherence to the efficient markets, rational expectations based theory of the exchange rate (Begg and Haque, 1985), that UIP suggests that countries that offer high interest rates should be compensating investors for the risk that their currency will depreciate (or that the forward rate should be a good predictor of the future spot rate). Burnside et al (2006) show that empirically the reverse occurs, i.e. that the high yielding currencies have tended to appreciate while the low yielding currencies have tended to depreciate. This point was also made by Chaboud and Wright (2005), who noted that a test of UIP involves regressing exchange rate returns on the interest differential with a prior that the intercept should be zero and the slope coefficient should be unity. Empirical tests, however, reveal the slope coefficient to be negative (thus implying that the higher-yielding currency appreciates).

Such findings have existed in the literature for some time (Hunter and Smith, 1982), but as was stated above, have often been put down to the assumptions that are made with respect to the replacements of expectations by forward values or lagged values of the actual exchange rate. This type of evidence suggests a more fundamental basis for the failure of UIP and

[15] This would be consistent with the evidence that UIP is programmed into dealer trading programmes and the evidence drawn from Taylor (1985) similarly found UIP with data samples drawn from observations within the day trades.

explains how carry trade can be profitable. It is difficult to provide a rationale for why this takes place. Academic research has struggled to find an acceptable solution.[16]

However, let us assume that investors require a further risk premium that is in addition to the higher interest rate to lead them to invest in the foreign currency. Such a risk premium may dominate interest rate differentials when investors become concerned about the credit worthiness of a state or the credibility of the Central Bank. Many financial institutions do devote a lot of resources to country risk analysis and the application of this study might yield significant risk premia when comparison is made between exchange rates in the developed and developing world. However, this would seem less important for comparisons of rates across developed economies, which is pertinent as many of the studies already referred to relate primarily to OECD and the EU economies.

It is also important to reflect that UIP is observed to hold when the analysis considers the market at a micro level. The studies by Taylor (1985), and Hallwood and MacDonald (2000) relate to an analysis that centres on the agent either in terms of the transactions data or the expectations, while Chaboud and Wright (2005) are looking at what can be seen as market micro-structure effects. These within-the-day or short-frequency data analyses can be assumed to be conditioned on current inflation rates and macro-conditions. Hence, they may be invariant to inflation, macro policy and global conditions. By contrast, medium or long frequency time series analyses cannot be seen in this way unless the underlying exchange rate equations can treat interest rates as being in some sense exogenous and/or invariant to the behaviour of other macro factors.

This leads to the possibility that macro-variables will interact with both the exchange rate and interest rate equations in the system suggesting that in a time series context the exchange rate

[16] It should be noted that aggregation might be a primary concern. For example, in the Consumer Demand literature, even when individual agent responses to relative price rises are negative, it is possible after aggregation for the market response to be positive and seemingly contrary to the law of demand (see Hardle, Hildenbrand and Jerison, 1991).

needs to be considered as part of a macro model that at least considers the primary parity conditions and risk factors influencing the exchange rate and inflation. Once one is asked to analyse a system, then the dimension of the data and the extent of the underlying dynamics become a concern. Here, we have restricted the extent to which the system is expanded to consider the impact of prices through PPP and inflation, and interest rates. Clearly, there are a range of studies that also consider PSBR, relative GDP and other measures that might influence country specific risk and risk premia.

To this end, both econometrics and economics suggest that inflation and PPP may also help to shed some light on the failure of UIP. It follows for PPP to hold in a long-run sense that the currencies of high inflation countries should depreciate to ensure that the real exchange rate is stationary. According to the Fisher equation, the nominal interest rate should be higher in countries with high inflation, allowing UIP to hold and indicating that carry trade should not be profitable. As has already been discussed, the evidence in favour of PPP or the stationarity of real exchange rates is not conclusive. The evidence obtained here is suggestive that PPP might hold, but this conclusion is sensitive to the efficiency, consistency and specification of the univariate and panel studies considered. Further, even were the tests consistent and the specifications reasonable characterisations of the data, there may be issues related to power and size that are sensitive to the underlying distributional assumptions and the potential for time varying volatility when prices and exchange rates are analysed.

The failure to find that PPP holds would appear to indicate that UIP cannot be empirically validated and this is manifested in the observation that there is significant carry trade. This does not imply that UIP is not important, but that the failure to observe that the theory holds is as a result of the many limitations to the analysis considered above. In principle, it is still a key building block in the finance literature and the failure may well be used as an indication that the foreign exchange market may not be in some sense efficient. It does suggest caution when assessing UIP in relation to currencies that may be subject to significant carry trade. While writing this chapter, the yen has been subject to substantial carry trade. This has taken the form of investors borrowing yen at extremely low interest rates and investing (i.e. carrying) the funds in high yielding foreign currency assets in order to yield a profit.[17] High-yielding currencies that are frequent 'targets' of carry trade include the UK Pound Sterling, the Australian dollar, and the New Zealand dollar.[18] As the theory explains above, in normal market conditions carry trade is something of an anomaly, as one would obviously expect arbitrage to bid up the Japanese interest rate and bid down the interest rate of the target currency to the point at which the profit is zero.

It is important to consider what is meant by market efficiency. Often in Finance Theory (Copeland et al, 2003), the notion of market efficiency is equated with technical or product efficiency. However, it is only the notion of informational efficiency that might be well analysed using data on asset market prices alone. Hence, a clearing asset market does not lead to full equilibrium of the system. In fact asset market clearing is almost axiomatic, while goods market clearing is almost unobservable. Hence, our capacity to observe UIP or PPP in the aggregate may be limited in the short-run and the long-run. Firstly, UIP can be observed with high frequency data, because prices are given or essentially fixed, while PPP cannot hold on this timeframe by construction. While in the medium term when prices and other macro factors move, then UIP in aggregate data is sensitive to risk factors and the macro-economy.

[17] It is important to recall that the Japanese economy has suffered a considerable recession that the economy has not been able to shake off. In particular, a strong monetary disequilibria has been caused by deflation that both depresses purchases by firms and consumers as prices are falling and limits the extent to which the monetary authorities can impact real interest rates. Given the lower limit of zero on the nominal interest rate either via conventional IS/LM analysis or disequilibrium theory (Barro and Grossman, 1977), monetary policy on its own is not able to move the economy towards a goods market equilibrium (Keynes (1936), Hansen (1949), Leijonhufvud (1968) and Krugman (1998, 1999)).

[18] As can be observed by the credit crunch and LIBOR rates, the capacity of the monetary authorities to bring effective nominal rates down has been stemmed in the US and Europe. In this case it is falling asset prices that have impacted on the implicit riskiness of debt across the banking sector that has caused banks not to lend to each other. As it is explained in Hunter and Isachenkova (2006), Basel II and the associated credit rules may in a perverse way intensify the nature of the downturn, as firms in crisis are unable to avoid failure as their short loan position and credit rating impact on their capacity to borrow and roll over debt. The two prongs of downgraded credit ratings and the high cost of bank borrowing is leading to bank led failure, and the difficulty that the monetary authorities face in solving the crisis via conventional injections of liquidity alone.

Hence, observing meaningful parity relations might be possible only in the long-run and then in combination (Juselius, 1995) or certainly as conjoined long-run relations (Burke and Hunter, 2005).

A further issue that is of interest relates to covered interest parity (CIP). When CIP is considered, then the interest rate differential should be reflected in the difference between the spot and forward exchange rate (low-yielding currencies are at premium in the forward market, while high-yielding currencies are at a discount). As has been explained, the evidence in favour of UIP is slim, to the extent that it would seem that the problem is unlikely to be dependent on the failure of the expectations assumption alone. Hence, it would seem that the evidence on UIP would suggest that CIP also fails. Ultimately the profit that can be made from carry trade is subject to exchange rate risk. In the absence of exchange rate fluctuations, the profit is proportional to the interest rate differential and the forward premium between the currencies. Since exchange rates can be volatile, any appreciation of the lower-yielding currency or depreciation of the high-yielding currency can lead to potential losses, and this ultimately leads to the unwinding of carry trade positions in order to minimize losses. This type of scenario helps to explain the yen carry trade. When the Bank of Japan increased interest rates at the beginning of 2007, positions started to unwind as investors feared excessive losses. As the yen appreciated, carry trade investors needed to sell more foreign currency than was originally planned in order to purchase the same quantity of yen. Further yen appreciation can eliminate the entire profit and lead to losses.[19]

The absence of a significant carry trade provides a rationale for selecting the most appropriate countries to examine the incidence of PPP and UIP, because the UIP relationship can be distorted by this phenomenon. This justifies the omission of Japan from an analysis that

[19] This unwinding of carry trade positions can have implications for global financial market stability. A similar event took place in October 1998, when the yen started to appreciate against the US dollar after a period of three years when it was depreciating. Yen 'carry traders' unwound at the same time, causing a substantial yen appreciation and a drop in the price of high-yielding assets, causing the Federal Reserve to reduce the Fed funds rate twice (25 basis points each).

involves testing for UIP, and results more closely aligned to UIP would be more likely in relation to countries not subject to carry trade. Studies of UIP that incorporate Japan have been undertaken in the past (e.g. MacDonald and Marsh, 2004). Such studies that exist that suggest some form of UIP holds tend to be modelled in conjunction with PPP. However, and to the knowledge of this author, the sample period selected for the countries analysed here, does not include a period where carry trade would be a popular trading strategy for investors.

Consideration of UIP in conjunction with PPP

As well as being careful with regard to the selection of countries to examine in terms of UIP, it has also become reasonably popular in the literature to examine UIP in conjunction with PPP. Juselius and MacDonald (2004b) note that the joint modelling of UIP and PPP has suggested that UIP and PPP might appear more credible. Hence, there exists some evidence to suggest that deviations from the parities are stationary. These authors make the assumption that deviations from PPP are the result of real factors such as productivity differences or fiscal imbalances. These factors are manifested in the current account, and under a floating exchange rate regime (where it is assumed that the central bank does not alter its foreign reserves), the imbalance in the current account must be financed by the capital account. On theoretical grounds, therefore, there would appear to be a strong linkage between PPP and UIP. In addition, Juselius (1995) makes the point that as the forecast horizon for exchange rates increases, then deviations from PPP should have an influence in the formation of expectations. Thus, there exists an implicit link between the goods market and the capital market. Notationally, equation (6) can be adjusted to describe the long-run expected exchange rate. Thus:

$$E(e_{t+1}) = p_t - p_t^*. \tag{8}$$

Then, PPP and UIP can be combined via the insertion of equation (8) into equation (7):

$$i_t - i_t^* = p_t - p_t^* - e_t. \tag{9}$$

The theoretical and empirical work undertaken to date would appear to suggest, therefore, that the consideration of PPP and UIP in a joint modelling framework indicates that the parity deviations are stationary. Clearly, the PPP hypothesis cannot be considered as a complete model of exchange rate determination given the absence of factors able to explain the capital market. It is appropriate to undertake an analysis that incorporates such a component via a system that can handle UIP through the inclusion of home and overseas interest rates. Intuitively, it makes sense that any model that seeks to explain exchange rate determination should consider both the goods and capital markets. In addition, in the current highly globalised world economy, international financial capital flows have become more widespread compared to, say, the 1950s and 1960s when perhaps trade flows were more important in influencing the exchange rate (and indeed where capital controls may have been in place). This may help to explain why some studies of PPP have found that it does not hold (i.e. because PPP was tested in isolation with no account taken of capital flows and the interest rate spread). This view is endorsed by Johansen and Juselius (1992), who suggest that some studies of PPP or UIP that failed to validate the conditions may have been misspecified, given that the links between goods and capital markets were not taken into account. Thus, the omission of variables from the cointegrating vector was the source of the failure to find PPP or UIP (i.e. tests of PPP excluded interest rates, and tests of UIP excluded prices).

Where PPP was found not to hold in previous studies, the common reasons put forward were market imperfections, price index selection, transport costs and trade barriers. The failure of UIP was typically rationalized by the presence of a time varying risk premium or limited capital mobility. These reasons, of course, are all entirely reasonable. However, it could also be the case that the exchange rate is not being modelled correctly. Indeed, the rejection of the conditions on an individual basis could be due to the implicit link that exists between the conditions. For example, it is understood that PPP is an arbitrage condition pertaining to the goods market, and thus refers to adjustment in the current account alone. UIP, on the other hand, is an arbitrage condition that drives efficient futures market prices and ensures equilibrium in the capital account through changes in domestic and foreign interest rates. As a result, disequilibria in the goods market may spill over into capital market, and vice versa (given that the balance of payments equals the sum of the current and capital accounts). Thus, it may be necessary for PPP and UIP to hold simultaneously for the appropriate balance to be observed in the long-run for the exchange rate associated with a stable balance of payments position. In addition, Camerero and Tamarit (1996) make the point that there can be deviations from PPP and UIP in the short-run due to price stickiness and the presence of a risk premium. These authors note that it could be possible that, in the long run, these types of deviations could balance each other out, when both the goods and capital markets were considered jointly. Otherwise, implicit in the long-run relationship is the notion the imbalance in the current can be balanced against a capacity to roll over debt on the capital account.

II.B Econometric Testing Procedures

Following Froot and Rogoff (1995), the empirical work done on PPP can be categorised as follows: (i) correlation-based tests (performed in the early 1970s and early 1980s); (ii) stationarity tests of the real exchange rate (performed from the 1980s onwards); and (iii) cointegration-based studies of nominal exchange rates and relative prices (performed from the 1980s onwards). Some studies also focused on the extent of mean reversion in real exchange rates through estimation of half-lives. More recently, panel estimators have been applied as researchers seek to utilise available data in a more informative way and increase the power of testing procedures (compared to a univariate framework). Longer time spans of data have also been used to increase the power of tests. Even more recently, there has been a growing literature on the non-linear nature of the PPP relationship. This literature is discussed in terms of how it can provide some explanation for the failure for long-run PPP to hold.

Largely, PPP was not found with respect to the first category of testing procedures (with some exceptions however), while mixed results were found using unit root analyses and cointegration studies. Froot and Rogoff (1995) make the additional distinction amongst the three categorisations described based on the nature of the null hypothesis in each case. With the simple correlation-based studies, the null hypothesis is that PPP holds. With unit root-based tests, the null is that the real exchange rate is characterised by a random walk and the alternative hypothesis is that PPP holds in the long-run. Cointegration tests are based on the null hypothesis that *"deviations away from any linear combination of prices and exchange rates are permanent"*.

II.B1 Correlation-Based Tests

Following Sarno and Taylor (2002), the correlation-based tests of PPP were based on estimating equations of the form (logarithmic):

$$s_t = \alpha + \beta p_t + \beta^* p_t^* + \omega_t \tag{10}$$

Testing the restrictions that $\beta=1$, $\beta^*=-1$, and the constant equals zero would be a test of absolute PPP, while relative PPP could be tested by imposing the same restrictions on the same equation in first differences, i.e.

$$\Delta s_t = \alpha + \beta \Delta p_t + \beta^* \Delta p_t^* + v_t \tag{11}$$

For example, Frenkel and Johnson (1978) carried out tests based on the equations above for the dollar-pound, franc-dollar, and franc-pound exchange rates for the period February 1921 to May 1925. PPP was found in these cases in both absolute and relative form. Frenkel

(1981a) used the same regression-based technique on the floating rate era from 1973 for the dollar-pound, dollar-franc, and dollar-deutschmark. Here PPP was not found. Frenkel argued that this could be due to temporary real shocks and sticky goods prices. Krugman (1978) finds similar results, and makes the point that *"There is some evidence then that there is more to exchange rates than PPP. This evidence is that the deviations of exchange rates from PPP are large, fairly persistent, and seem to be larger in countries with unstable monetary policies"* (p.407). Krugman's methodology helped to address one of the problems with the models being estimated at that time, namely that of endogeneity in the price level (which resulted in a downward bias in OLS estimates of β). Krugman resolved this problem by introducing a time trend as an instrumental variable. While this produced estimates of closer to 1, the failure of PPP was still evident however.

The result by Frenkel and Johnson (1978) was driven by the fact that hyperinflation was prevalent for the countries under consideration. The limitations of the correlation-based tests in finding PPP was highlighted by Froot and Rogoff (1995), who noted that with the exception of hyperinflations, PPP is likely to be strongly rejected using these types of approaches.

Sarno and Taylor (2002) suggest that many of the tests based on equations (10) and (11) tended to reject PPP. A major common problem with the early literature carried out was that the data involved was not tested for stationarity in the residuals. Clearly, conventional inference is not valid for series that are non-stationary as their variances do not converge to a constant and estimated standard errors will be underestimated (Breuer, 1994). Of course, had such testing been undertaken and non-stationarity in the results found, then this would imply that part of the shocks on the real exchange rate were permanent (i.e. PPP does not hold). As time moved on and econometric technology improved, it became possible to test for stationarity in the data series (Engle and Granger, 1987).

II.B2 Tests of Stationarity

Tests based on stationarity testing were of the form:

$$q_t \equiv s_t - p_t + p_t^* \qquad (12)$$

where q_t is the log of the real exchange rate, s_t is the nominal exchange rate, p_t is the domestic price series and p_t^* is the foreign price series.

Tests based on equation (12) are usually structured under the null hypothesis that the real exchange rate follows a random walk, i.e. the null of non-stationarity.[20] The alternative hypothesis is thus of real exchange rate stationarity, thereby implying that PPP holds in the long-run. The rationale for the null being defined as a unit root in the real exchange rate has been provided by a number of researchers. The earliest work in this regard was perhaps by Roll (1979) who based his argument on the efficiency of the foreign exchange markets, i.e. efficient markets would imply unpredictability in the real exchange rate. As noted by Froot and Rogoff (1995), however, this particular rationale is not entirely appropriate since real exchange rates are not traded. An alternative rationale was provided by Froot and Rogoff (1985) based on the Balassa-Samuelson effect whereby permanent shocks to productivity across countries can lead to real exchange rate non-stationarity. In addition, Obstfeld and Rogoff (1995) describe a model whereby productivity is affected by any economic shock that causes a transfer of wealth across countries. This, in turn, can induce a unit root in the real exchange rate.

The first type of unit root test procedure was introduced by Dickey and Fuller (1979) – the Dickey-Fuller test. This was later adapted to account for higher order autoregressions, and models whose residuals exhibit signs of autocorrelation, with the Augmented Dickey-Fuller

[20] Early tests of this sort include Hakkio (1984), Frankel (1986), Huizinga (1987), and Meese and Rogoff (1988).

(ADF) test (the standard DF test would be sufficient where the series is classed as AR(1)). The most general form of the ADF test can be notated as follows:

$$\Delta q_t = \alpha_0 + \alpha_1 t + \alpha_2 q_{t-1} + \Sigma\beta_i\Delta q_{t-i+1} + \varepsilon_t \qquad (13)$$

where Δq_t is the one-period change in the real exchange rate, α_0 is an intercept, $\alpha_1 t$ is a time trend, and $\Sigma\beta_i\Delta q_{t-i+1}$ represents higher-order autoregressive components. This model can also be specified without the deterministic elements α_0 and $\alpha_1 t$.

The real exchange rate is deemed to have a unit root where α_2 equals 1 (i.e. the null of non-stationarity). In this case, therefore, the real exchange rate, q_t, would be deemed to be represented as a random walk with drift. The alternative hypothesis is defined as a stationary process, i.e. that PPP holds. Under the alternative hypothesis, therefore, α_1 equals 0 and α_2 is less than 1. In determining rejection or acceptance of the null, Dickey and Fuller (1979) indicated that the conventional confidence intervals are not valid for unit root tests based on equation (13). Conventional inference is only valid for stationary series. Wider intervals, calculated by Dickey and Fuller (1979) using Monte Carlo simulations, determined the outcome of the test. The critical values determining acceptance or rejection of the null hypothesis have become known as the *tau* statistics (the decision rule is applied in the same manner as a conventional t-test: if the computed tau is greater than the critical tau, then the null is rejected, implying a stationary series). The relevant Dickey-Fuller critical values across a range of specifications and hypotheses are provided in **Table 2.1** below.

Table 2.1 Dickey-Fuller Critical Values

Model	Hypothesis	Test Statistic	Critical Values for 95% and 99% Confidence Intervals
$\Delta q_t = \alpha_0 + \alpha_1 t + \alpha_2 q_{t-1} + \varepsilon_t$	$\alpha_2=0$	τ_τ	-3.45 and –4.04
	$\alpha_0=0$ given $\alpha_2=0$	$\tau_{\alpha\tau}$	3.11 and 3.78
	$\alpha_1=0$ given $\alpha_2=0$	$\tau_{\beta\tau}$	2.79 and 3.53
	$\alpha_2=\alpha_1=0$	ϕ_3	6.49 and 8.73
	$\alpha_0=\alpha_2=\alpha_1=0$	ϕ_2	4.88 and 6.50
$\Delta q_t = \alpha_0 + \alpha_2 q_{t-1} + \varepsilon_t$	$\alpha_2=0$	τ_μ	-2.89 and –3.51
	$\alpha_0=0$ given $\alpha_2=0$	$\tau_{\alpha\mu}$	2.54 and 3.22
	$\alpha_0=\alpha_2=0$	ϕ_1	4.71 and 6.70
$\Delta q_t = \alpha_2 q_{t-1} + \varepsilon_t$	$\alpha_2=0$	τ	-1.95 and –2.60

(Source: Enders, 1995)

Following the work of Dickey and Fuller, a further early unit-root test procedure was developed by Phillips and Perron (1988). They noted that a degree of sensitivity was evident with the Dickey-Fuller tests since they make the assumption that the errors are Gaussian, i.e. that there is no heteroskedasticity or autocorrelation. The approach of Phillips and Perron (1988) relaxes these assumptions somewhat. Specifically, they devise a model that allows for weak dependence and heterogenous distribution of the errors. Their model is based on the following model:

$$q_{it} = \theta_{0i} + \theta_{1i}(t\text{-}2^{-1}T) + \theta_{2i}q_{it-1} + \varepsilon_{it} \qquad (14)$$

where the null hypothesis of a unit root is θ_{1i}.

The lag selection with the Phillips-Perron test is determined automatically due to the Newey & West Bandwidth using the Bartlett Kernal spectral estimation procedure.

The early literature on unit-root testing developed further following Perron (1989) who devised a technique to deal with the problem of bias towards non-rejection of a unit root in the presence of a structural break. Perron (1989) argued that a series that is stationary with a one-time break can appear non-stationary using Dickey-Fuller type testing procedures. Perron's unit root test in the presence of a break is based on detrending the series allowing for a level shift and/or a shift in the trend. The breakpoint is imposed either via *a priori* information or estimated endogenously using the approach of Perron and Vogelsang (1992). Dummy variables are used to reflect the shifts in the series and an ADF-type test is run on the adjusted series. A number of researchers have argued for incorporating structural breaks into the modelling of real exchange rate dynamics, notably Baum et al (1999). A significant amount of work has been done over the past decade or so to improve the power of standard univariate unit root tests such as the ADF and PP tests. Of these, two worth discussing are the ADF-GLS test and the KPSS test as these tests have had widespread application in the literature to testing for PPP. Developed by Elliott, Rothenberg and Stock (1996) the ADF-GLS test is based on testing the null of non-stationarity using the following model:

$$\Delta q^{d}_{it} = \alpha_i + \gamma_i t + \rho_i q^{d}_{it-1} + \Sigma^{k}_{j=1}\delta_t \Delta q^{d}_{it-1} + \varepsilon_{it} \qquad (15)$$

where q^{d}_{it} refers to the demeaned real exchange rate for country i and time t.

The number of lags, k, in the ADF-GLS test is determined by the Modified AIC following Ng and Perron (2001). The null of non-stationarity is that $\rho_i = 0$. In contrast to the ADF-GLS (and the majority of other unit root tests), the KPSS test (developed by Kwiatkowski et al, 1992) is constructed under the null of stationarity. This test is performed by regressing the series, q_t, on an intercept and a trend, i.e. of the form:

$$q_t = \alpha_t + \beta t + e_t \qquad (16)$$

where e_t follows a stationary process and α_t is a random walk (i.e. $\alpha_t = \alpha_{t-1} + u_t$).

Following this, the partial sum of the residuals is calculated as $s_t = \Sigma^t_{i=1} e_i$. The KPSS test statistic can then be calculated as: $LM = T^{-2} \Sigma^T_{t=1} s_t^2/\sigma_T(l)$; where $\sigma_T(l)$ is an estimate of the long-run variance of the residuals. The null of stationarity is rejected when the KPSS LM is large (i.e. greater than its critical value), implying that the series wanders from its mean.

Of course, different unit root tests, and indeed different lag selection criteria, can sometimes provide different results, and thus it is important to make sure that the outcome identified is the most feasible. For example, it is well known that the AIC is not consistent for large samples. On other hand, the modified AIC (MAIC) has good size and power properties in finite samples.[21] Thus, this type of information should be borne in mind when undertaking unit root tests. As an example of how different tests can provide different results, consider the yen real exchange rate. Studies of the yen real exchange rate appear to be most pronounced in terms of rejecting PPP and this has been a mystery for international finance researchers. In fact, Chortareas and Kapetanios (2004) notes that the yen real exchange rate has emerged as a prime example of a major currency that refutes PPP. In this short empirical exercise, three bilateral yen exchange rates are tested for a unit root using six different unit root test procedures and three different lag selection criteria. The data comprises a rich source of euro-yen exchange rates from 1991 to 2006, and pound-yen and dollar-yen rates from 1984 to 2006 (all monthly frequency).

[21] As suggested by Ng and Perron (2001), the MAIC criterion avoids the size problems typically associated with the AIC and SIC criteria.

Table 2.2 Unit Root Tests for 3 Yen Real Exchange Rates

Unit Root Test Method	R^{Euro}				R^{US}				R^{UK}			
	Intercept		**Intercept and Trend**		**Intercept**		**Intercept and Trend**		**Intercept**		**Intercept and Trend**	
	Lags	*t-stat*	*Lags*	*t-stat*	*Lags*	*t-stat*	*Lags*	*t-stat*	*Lags*	*t-stat*	*Lags*	*t-stat*
ADF	1^{SIC}	-1.06	1^{SIC}	-1.74	1^{SIC}	-2.43	1^{SIC}	-2.48	$1^{SIC, AIC}$	-0.79	$1^{SIC, AIC}$	-1.59
	2^{AIC}	-1.12	2^{AIC}	-1.71	2^{AIC}	-2.54	2^{AIC}	-2.60	2^{MAIC}	-0.89	2^{MAIC}	-1.66
	4^{MAIC}	0.99	4^{MAIC}	-1.54	3^{MAIC}	-2.68	3^{MAIC}	-2.77	4	-1.11	4	-1.77
PP	1^{SIC}	-1.09	1^{SIC}	-1.78	1^{SIC}	-2.18	1^{SIC}	-2.20	$1^{SIC, AIC}$	-0.72	$1^{SIC, AIC}$	-1.52
	2^{AIC}	-1.17	2^{AIC}	-1.83	2^{AIC}	-2.24	2^{AIC}	-2.25	2^{MAIC}	-0.79	2^{MAIC}	-1.57
	4^{MAIC}	-1.22	4^{MAIC}	-1.86	3^{MAIC}	-2.26	3^{MAIC}	-2.27	4	-0.91	4	-1.66
DF-GLS	1^{SIC}	-0.98	1^{SIC}	-0.90	1^{SIC}	-0.80	1^{SIC}	-0.94	$1^{SIC, AIC}$	-0.98	$1^{SIC, AIC}$	-1.10
	2^{AIC}	-1.08	2^{AIC}	-1.02	2^{AIC}	-0.90	2^{AIC}	-1.12	2^{MAIC}	-1.06	2^{MAIC}	-1.18
	4^{MAIC}	-1.05	4^{MAIC}	-0.97	3^{MAIC}	-0.90	3^{MAIC}	-1.13	4	-1.28	4	-1.39
ERS	1^{SIC}	10.86*	1^{SIC}	33.04*	1^{SIC}	22.38*	1^{SIC}	39.55*	$1^{SIC, AIC}$	7.57*	$1^{SIC, AIC}$	23.49*
	2^{AIC}	9.09*	2^{AIC}	28.13*	2^{AIC}	16.93*	2^{AIC}	30.56*	2^{MAIC}	6.83*	2^{MAIC}	21.39*
	4^{MAIC}	9.37*	4^{MAIC}	29.63*	3^{MAIC}	16.80*	3^{MAIC}	31.27*	4	5.30*	4	16.71*
KPSS	1^{SIC}	1.98*	1^{SIC}	1.10*	1^{SIC}	1.96*	1^{SIC}	1.94*	$1^{SIC, AIC}$	3.51*	$1^{SIC, AIC}$	1.54*
	2^{AIC}	1.34*	2^{AIC}	0.75*	2^{AIC}	1.32*	2^{AIC}	1.31*	2^{MAIC}	2.37*	2^{MAIC}	1.04*
	4^{MAIC}	0.83*	4^{MAIC}	0.46*	3^{MAIC}	1.01*	3^{MAIC}	0.99*	4	1.46*	4	0.64*

Table 2.3 **Ng and Perron Unit Root Tests**

	Intercept					**Intercept and Trend**				
$\mathbf{R}^{\mathbf{Euro}}$	*Lags*	MZ_α	MZ_t	MS_B	MP_T	*Lags*	MZ_α	MZ_t	MS_B	MP_T
	1^{SIC}	-2.29	-0.99	0.43*	10.18*	1^{SIC}	-2.54	-0.91	0.36*	28.54*
	2^{AIC}	-2.77	-1.11	0.40*	8.61*	2^{AIC}	-3.27	-1.08	0.33*	23.97*
	4^{MAIC}	-2.64	-1.07	0.41*	8.99*	4^{MAIC}	-3.13	-1.05	0.33*	24.72*
$\mathbf{R}^{\mathbf{US}}$	1^{SIC}	-1.24	-0.77	0.63*	19.39*	1^{SIC}	-2.19	-0.90	0.41*	34.45*
	2^{AIC}	-1.62	-0.89	0.55*	14.87*	2^{AIC}	-3.13	-1.12	0.36*	26.19*
	3^{MAIC}	-1.60	-0.88	0.55*	15.05*	3^{MAIC}	-3.11	-1.11	0.36*	26.33*
$\mathbf{R}^{\mathbf{UK}}$	$1^{SIC,\ AIC}$	-3.11	-1.00	0.32*	7.55*	$1^{SIC,\ AIC}$	-3.62	-1.12	0.31*	21.80*
	$2^{\ MAIC}$	-3.63	-1.11	0.31*	6.82*	$2^{\ MAIC}$	-4.20	-1.23	0.29*	19.67*
	4	-5.16	-1.39	0.27	5.31*	4	-5.89	-1.52	0.26*	15.25*

For the purposes of the exercise at hand, **Tables 2.2** and **2.3** report the test results in levels only. In interpreting the results, it is important to bear in mind that the interpretation is governed by the null hypothesis. All tests are defined under the null of a unit root, with the exception of the KPSS test which is defined under the null of stationarity. It is clear to see the stark differences in some of the results. For example, the ADF, PP and DF-GLS, KPSS, and Ng-Perron tests all show that the yen real exchange rate is non-stationary in levels (recall that the KPSS test is structured under the null of stationarity). Only the ERS point-optimal test rejects the null of a unit root in yen real exchange rate levels. Thus, it would appear that there is very strong evidence to suggest that the yen real exchange rate is non-stationary in levels. This is robust to five alternative unit root tests and three different information criteria used to inform the most appropriate lag length. Therefore, in undertaking stationarity tests of a variable, it would help to ensure robustness if more than one test yielded the same result. The bulk of the literature tends to present results for the presence of a unit root based on both a test structured under the null of non-stationarity and a test structured under the null of stationarity.

Other tests of PPP based on estimating the stationarity of the real exchange rate include variance ratio tests and fractional integration tests. With variance ratio tests, the null is defined as a unit root process whereby the variance of the real exchange rate increases in a linear fashion over time. This procedure was developed by Cochrane (1988), and measures the persistence of the real exchange rate, z(k) as follows:

$$z(k) = k^{-1} [var(q_t - q_{t-k}) / var(q_t - q_{t-1})] \tag{17}$$

where k is a positive integer and var represents the variance.

The exchange rate is deemed to contain a unit root if $z(k) = 1$. If $0 < z(k) < 1$, then the random walk hypothesis is rejected in favour of mean reversion. This type of approach was applied by Huizinga (1987) who was able to reject the random walk hypothesis using dollar exchange rates in the post-1973 era, concluding that real exchange rates are mean-reverting over the time horizons of 48 months or longer. Glen (1992) also uses this approach in a test of nine dollar rates and two cross rates over the 1900-1987 period. He finds that mean reversion is evident for the post-Bretton Woods period as the time period approaches 32 months.

With regard to fractional integration, there exists a higher range of alternative hypotheses since d can lie anywhere between 0 and 1. As noted by Breuer (1994), a series that is fractionally integrated will revert to the mean but at a slower rate than a stationary series. This type of approach is not the most widespread test applied to PPP, although some studies have been carried out in the past. For example, Cheung and Lai (1993b) find greater power with fractional integration compared to ADF tests in detecting mean reversion when d lies between 0.35 and 0.65. These results are in contrast to those of Diebold, Husted and Rush (1991) who find little additional power with fractional integration using a sample of over 100 years. They reject the random walk and posit that real exchange rates are best represented by an ARMA framework.

While unit-root tests are more sophisticated than the earlier simple correlation-based tests, they have a problem in relation to power. Specifically, the tests are recognised to have low power to reject the null hypothesis. Froot and Rogoff (1995) make the point that *"Because slow, albeit positive, rates of reversion toward PPP are plausible in many models, random walk tests may provide little information against relevant alternative hypotheses"*. Froot and Rogoff (1995) make the additional point that the post-1973 floating rate era is not sufficiently long to provide a reliable rejection of the null of non-stationarity.

II.B3 Cointegration-Based Tests

As well as correlation-based tests of PPP and tests based on assessing real exchange rate stationarity, the most recent innovative approach to testing for PPP is based on the concept of cointegration. Cointegration holds that the combination of two or more non-stationary series can yield a long-run stationary relationship as long as the series are integrated of the same order. Essentially the non-stationarity evident in each of the series is cancelled out and a long-run stationary relationship can be observed. The early cointegration-based work was carried out by Engle and Granger (1987). This approach takes places in two stages: firstly, the series are tested for unit roots using the ADF procedure; then if we find that both series are integrated of the same order, we run the OLS regression, save the residuals, run an autoregression of the residuals and test the significance of the coefficient on the lagged residual. Formally, having established that the series y_t and x_t have the same order of integration, we run $y_t = \alpha_t + \beta x_t + e_t$, then we run $\Delta\hat{e}_t = \gamma + \hat{e}_{t-1} + \nu_t$. The series are deemed to be cointegrated if we can reject the null hypothesis that γ=0 (such a test is the same as testing for a unit root in $\hat{e}_t$).

Based on their two-step approach, and the notion of superconsistency, it follows that in the bivariate case, equation (10) represents a typical equation for testing for cointegration between the nominal exchange rate and relative prices. Recall equation (10): $s_t = \alpha + \beta p_t + \beta^* p_t^* + \omega_t$. In this case, should s_t, p_t, and p_t^* be classified as being I(1), i.e. integrated of order one, then weak-form PPP is deemed to hold as long as the residual error process does not have a unit root (i.e. that it is stationary). It follows that the outcome defined by the relation (s_t-α-βp_t-$\beta^* p_t^*$) defines a stationary series and the regression error has a well-defined variance. Strong-form PPP would hold if (in addition to weak-form) β=1 and β^*=-1, i.e. the proportionality restriction. As has already been mentioned, however, the strong-form version of the PPP hypothesis is unlikely to hold due to market disturbances such as transportation costs and trade barriers. Weak-form PPP, thus, is more likely to hold in practice and this weak form of PPP is implied by the symmetry restriction that β = - β^*.

A precursor to Engle and Granger (1987) was Granger (1983) and Granger and Weiss (1983), who presented a test for cointegration based on an error correction model framework. The 'Granger Representation Theorem' holds that there exists an error correction representation of any cointegrated series. This representation can be notated as follows:

$$\phi(1\text{-}L)X_t = \eta v_{t-1} + e_t \tag{18}$$

where (1-L) is the lag operator, ϕ are the coefficients of the endogenous and exogenous variables (for the endogenous variable Φ=1), and is the error correction mechanism (where $\eta \neq 0$).

For example, assuming that a cointegrating relationship exists between s, p and p* from equation (10), this could be re-written as an error correction model as follows:

$$\Delta s_t = \mu + \phi_1 \Delta p_t + \phi_2 \Delta p_t^* + \eta v_{t-1} + \varepsilon_t \quad (19)$$

From equation (19), s, p, and p* are deemed to be cointegrated if the hypothesis that η=0 can be rejected. Within this framework, information is provided on both the long-run and short-run dynamics, the former evident in v_{t-1} and the latter evident in the Δ variables. Clearly, since the error correction framework comprises stationary series, conventional inference is valid.

An early application of the Engle and Granger approach was undertaken by Kim (1990). Using a bivariate model, he uses this method to test for non-stationarity across five real exchange rates. Following this, he estimates an error correction model and determines whether or not PPP exists in the long-run in the manner described above. Kim concludes that PPP in the long-run is more apparent when the wholesale price index (WPI) instead of the consumer price index (CPI) is used as the price series in the system (cointegration was identified 80% of the time using the WPI compared to 40% of the time with the CPI).

The Engle and Granger approach is limited in that it only allows us to test for the presence of one cointegrating vector. As a result, its application to PPP testing is limited to bivariate frameworks. There may also be problems in terms of choosing the appropriate dependent variable. As well as this endogeneity problem, tests based on Engle and Granger can be prone to autocorrelation.

In order to circumvent these problems, a more powerful cointegration test procedure was developed by Johansen (1988) and Johansen and Juselius (1990). This approach to testing for cointegration has become widely used in the context of PPP, particularly since the mid-1990s. In addition, the Johansen procedure is a standard estimation technique that can be performed in a wide variety of econometric software packages. Unlike the Engle and Granger approach, the Johansen approach is more efficient in that there is only one step required. For example, the coefficients from equation (10) can be estimated and the test for a unit root can be performed simultaneously using the maximum-likelihood estimator proposed by Johansen. This feature is also important in terms of accuracy since it avoids a carry-over of potential misspecification to the latter stage of the cointegration test (this can be a problem with the Engle and Granger method whereby any mistakes made in estimating the residuals can be transmitted to the final step of the test). The Johansen approach also has the benefit it parametrically corrects for autocorrelation and endogeneity, and it can be applied to a multivariate framework.

In the context of testing for PPP, the Johansen method is based on assessing whether or not a long-run equilibrium relationship can be identified between a bilateral nominal exchange rate and two price series (domestic and foreign). These tests differ from tests of unit roots in real exchange rates since the symmetry and proportionality restrictions that are implicitly imposed in real exchange rate tests are not imposed in this case. Rather, these conditions can be explicitly tested by imposing appropriate restrictions on the parameters of the long-run matrix. As well as this, cointegration can be used as the basis of a form of multi-country test of PPP, as opposed to the traditional two-country test. This type of test can help to recognise the dependence of stationarity of real exchange rate dynamics of more than one trading partner for example. Such a test would examine whether or not the real exchange rates of the relevant countries are cointegrated. In order to describe this, consider firstly the following VAR(k) model:

$$z_t = A_1 z_{t-1} + \ldots.. + A_k z_{t-k} + \varepsilon_t \qquad \varepsilon_t \sim IN\,(0, \Sigma) \tag{20}$$

where z_t is the log of the real exchange rate in the form (n x 1) and A_i represents a matrix of parameters (n x n).

Equation (20) can be denoted as a VEC equation as follows (in first-differenced form):

$$\Delta z_t = \Gamma_1 \Delta z_{t-1} + \ldots + \Gamma_{k-1} \Delta z_{t-k+1} + \Pi z_{t-k} + \varepsilon_t \tag{21}$$

where Γ_i represents $-(I - A_1 - \ldots. - A_i)$, $(i = 1,\ldots., k-1)$, and $\Pi = -(I - A_1 - \ldots. - A_k)$

By notating the system in this way, information is provided on the long-run and short-run relationships, i.e. an indication is provided of how the system responds in both the long-run and the short-run to changes in the z_t. Short-run information is given by the estimates of Γ_i, while long-run information is provided by estimates of Π. The series Π is decomposed as $\Pi = \alpha\beta`$, where the matrix α represents the speed of adjustment to equilibrium, and β represents the cointegrating vectors. Equation (21) can also be augmented to include a constant term to capture trending behaviour. Of course, cointegration tests of real exchange rates implicitly impose symmetry and proportionality restrictions. A multi-country test of PPP that relaxes these restrictions would be computationally very demanding given the number of variables that would have to enter the system. This issue is addressed further in **Chapter 3**.

Using the Johansen cointegration procedure, two specific test statistics are provided; one relating to the trace test and the other to the maximum eigenvalue test. Both tests yield the number of cointegrating vectors in the system. In performing the Johansen test, it is important to appropriately set the lag length, since there exists a high degree of sensitivity to the number of lags used (Breuer, 1994). A summary of the steps involved in performing the Johansen test are as follows:

Table 2.4 The Johansen Cointegration Procedure

Step	Activity
1	Perform unit root tests.
2	Set the appropriate lag length.
3	Select appropriate Johansen model.
4	Perform cointegration rank tests.
5	Weak exogeneity tests.
6	Analysis of normalised cointegrating vector(s).
7	Analysis of speed of adjustment coefficients.

Step 1 is performed in the usual way using the ADF tests to establish the presence of absence of a unit root. In order to proceed with cointegration it is necessary that all of the variables are integrated of the same order. The appropriate lag length is determined through analysis of the residuals. It is common practice to commence with a long lag length and systematically reduce this length in stages until the optimal lag is reached. The optimal lag is the one that yields the minimum number of lags in conjunction with well-behaved residuals, i.e. where the residuals can be classed as white noise. It is also useful to examine information criteria such as the Akaike Information Criterion (AIC) or the Schwartz Bayesian Criterion (SBC) in selecting the appropriate lag length. This is indicated by the VAR specification yielding the lowest AIC or SBC. The selection of the appropriate Johansen model can be guided by the Pantula (1989) principle. With the Johansen procedure, a range of possible models are possible, based on the presence and nature of the deterministic components in the VAR and the cointegrating space. There are five possible model types:

- **Model 1:** Here there are no deterministic components (i.e. no intercepts and no trends) in the VAR or in the cointegrating relationships.

- **Model 2:** In this model, there are restricted intercepts (in the cointegrating relationship(s) only) and no trends.

- **Model 3:** This model contains unrestricted intercepts (i.e. in the VAR and the cointegrating equation) and no trends.

- **Model 4:** Here, unrestricted intercepts and restricted trends are observed.

- **Model 5:** This model contains unrestricted intercepts and trends.

Step 4 of the procedure is crucial. It is vital to ensure that the correct cointegration rank is attained. The selection of a rank that is either too low or too high may affect the inference that can be drawn from the long-run normalised vector(s). Cointegration rank is determined on analysis of the trace statistics reported in the Johansen output. Following this, weak exogeneity can be assessed if the speed of adjustment coefficients are found not to be significant. Clearly, under such a scenario, no information can be provided on the long-run.[22]

Early studies using the Johansen approach to test for PPP include Kugler and Lenz (1993) and Liu (1993). Kugler and Lenz (1993) looked at the 1973-1990 period with data comprising 15 deutschmark exchange rates. Cointegration is found for 10 of the 15 currencies in the trivariate system. These results were found to be robust to the lag length.

[22] Burke and Hunter (2005; pp.116-118) provide a useful review of the performance of the Johansen test, including a discussion on how this is determined by the degree to which the model exhibits the required Gaussian errors.

Liu (1993) examines PPP for nine Latin American currencies over the forty year period up to 1989. Using the Johansen trace statistics, cointegration is found amongst all of the currencies. This result holds regardless of which price index is used (WPI or CPI). The results are somewhat mixed using the maximum eigenvalue test however. Liu's paper highlights the power of the Johansen test in that it does not depend on the normalisation condition. This is shown as he also performs the Engle and Granger method and finds cointegration in less than half the cases.

Essentially, evidence of stationarity with regard to the real exchange rate would support the idea that PPP holds in the long-run. Cointegration emerged towards the end of the 1980s (Engle and Granger, 1987; Johansen, 1988) and was utilised by many empirical researchers of PPP. These cointegration tests, however, largely failed to observe long-run PPP (e.g. Taylor, 1988; Mark, 1990). Generally, cointegration tests of PPP appear to suggest a greater likelihood of observing PPP under fixed exchange rate regimes and also when producer prices are used rather than consumer prices as a measure of the price level (McKinnon, 1971). This latter point refers to the idea that producer prices are comprised of a greater share of tradable goods.[23]

II.B4 Mean-Reversion and 'Half-Lives'

As well as cointegration testing, real exchange rates were also analysed in terms of mean reversion, i.e. how quickly the real exchange rate reverts to its mean level. Previous studies have looked at this issue in terms of the 'half-life', i.e. the length of time it takes for the elimination of one half of shock to the real exchange rate. An early example of work done on half-life estimation was carried out by Frankel (1986), who modelled a first-order autoregressive process for the real exchange rate of the form:

[23] Advocates of this view include Angell (1926), Heckscher (1930, and Ohlin (1967). An alternative view exists however to suggest that the most appropriate price index should be as broad as possible, such as the GNP deflator or CPI (e.g. Hawtrey, 1919; Cassel, 1928). This view is based on the idea that the exchange rate is an asset price and, therefore, the relative price of two monies.

$$(q_t - q^*) = \varphi(q_{t-1} - q^*) + \varepsilon_t \qquad (22)$$

where q is the real exchange rate, q^* is the constant equilibrium level of q, ε_t is a random noise term, and φ is the autocorrelation coefficient. This φ term is an unknown parameter which determines the speed of mean reversion.

As pointed out by Taylor and Taylor (2004), the real exchange rate reverts to its mean at the rate of $(1- \varphi)$ per period. Under the scenario where $\varphi=1$, the real exchange rate would never revert to its mean and would thus follow a random walk. Frankel (1986) estimated that $\varphi=0.86$, and the null hypothesis of a random walk is rejected.

An important contribution to the literature in terms of providing a possible rationale for the slow reversion to mean was made by Balassa (1964) and Samuelson (1964). This has become known as the Balassa-Samuelson effect, which essentially takes into account productivity effects within a traded goods/non-traded goods environment. Some of these studies have been fundamentally flawed, however, from an econometric perspective. Chinn and Johnston (1996) highlighted the fact that the Balassa-Samuelson effect measures the relationship between the real exchange rate and productivity levels. Studies done prior to 1995 tended to find Balassa-Samuelson effects although there were clear specification problems with the estimated models. This was due to the fact that the regressions were carried out in first-differences, and not in levels.

Research on Balassa-Samuelson effects using cointegration tests based on levels have been carried out largely since 1995, and these have had none of the prior mis-specification problems. The Johansen test has been used (e.g. Faruqee, 1995; Chinn and Johnston, 1996) and cointegrating relationships were found between the real exchange rate and productivity.

The slow mean reversion of real exchange rates in conjunction with the lack of overwhelming evidence for PPP holding in the long-run have become known as the 'PPP puzzles'. As noted by Huizinga (1987), even when PPP appears to holds, the mean reversion speed remains slow. These puzzles were summarised by Rogoff (1996) as follows: *"The purchasing power parity puzzle then is this: How can one reconcile the enormous short-term volatility of real exchange rates with the extremely slow rate at which shocks appear to damp out?"* (p.647). Rogoff (1996) noted that studies of the estimates of mean reversion speeds of real exchange rates towards their PPP levels tended to lie in the range of three to five years (based on estimates of the half-life adjustments).

II.B5 Longer Time Spans and Panel Tests of Stationarity

A key issue which has been identified in the literature is to examine whether the real exchange rate is non-stationary or whether it is mean-reverting. The limitations of testing for non-stationarity, however, have been explained in terms of the potential for the power of the unit root tests to be low, i.e. the null hypothesis to reject non-stationarity of the real exchange rate is low. Subsequently, researchers have sought to increase the power of the tests either by lengthening the span of the data or using panel data. The former approach entails its own problems, however, in that there can be spans across both fixed and floating exchange rate regimes. This type of approach was first carried out by Frankel (1986), as mentioned previously, using 116 years of data for the US dollar/UK pound real exchange rate. Fitting an AR(1) produced a coefficient of 0.86, indicating an annual decay in the deviation from PPP of 14% and a half-life of 4.6 years. In this case, the null of a unit root was rejected at the 5% level. Other researchers adopting long spans of data include Edison (1987); Johnson (1990); Abuaf and Jorion (1990); and Glen (1992). All of these studies rejected the random walk hypothesis.

An important issue was raised by Froot and Rogoff (1995) in relation to these types of studies adopting long spans of data across different types of exchange rate regimes: the inclusion of fixed rate periods in the sample may colour the outcome somewhat in favour of rejecting non-stationarity (since PPP is more likely to hold for fixed rates). Thus, it is uncertain whether mean reversion would be observed for a long span of floating rate data. Given that such data is not available, it is not possible to provide a conclusive answer to this issue. Some work was carried out, however, by Lothian and Taylor (1996) who looked at the dollar/pound over the 1791-1990 period and the franc/pound over the 1803-1990 period. For the post-1973 period, they cannot reject the null of a unit root. This hypothesis is rejected, however, when the entire period is considered across both exchange rates. Using a Chow test, they find no structural change in the AR(1) coefficients of the sample before Bretton Woods with that of after 1973. Lothian and Taylor (1996) thus refute the suggestion by Froot and Rogoff (1995) by showing that long time spans that incorporate fixed rate and floating periods are not biased in favour of stationarity due to the exchange rate relationship prevalent during the fixed rate period.

Due to the homogenous nature of the tests to find stationary real exchange rates, various panel estimators have been used to improve the power of the tests applied to single series.[24] The first panel-type study of exchange rates was carried out by Hakkio (1984), where a random walk was identified for a system of four exchange rates against the US dollar. Specifically, in order to gain power, he used GLS to permit cross-exchange rate correlation in the residuals of the series. In spite of this innovative approach, however, the null of a unit root was not rejected.

An influential early paper to the panel data study of PPP was undertaken by Abuaf and Jorion (1990). Their analysis used the Seemingly Unrelated Regression (SUR) procedure

[24] Panel estimation assumes that there are benefits to pooling real exchange rate series across a number of countries.

to test for PPP across the exchange rates of a panel of the G-10 countries over the period 1973 to 1987. The evidence was supportive of PPP, albeit at only the 10% level of significance. Subsequent studies also found evidence of PPP for the floating rate period from 1973 (Wu, 1996; Oh, 1996).

Advances in econometric technology enabled greater scrutiny to be made to the robustness of results found. In particular, the issue of cross-sectional dependence was a key problem associated with panel tests of PPP, as noted by O'Connell (1998); Papell (2003); Taylor (2003); and, Wu and Wu (2001). In fact, by factoring cross-sectional dependence into the estimation procedure can actually result in the rejection of PPP. Adjustment to account for this issue can make it more difficult to reject the null hypothesis of a random walk (Rogoff, 1996).

The early work on testing for PPP using panel data was based on performing panel unit root tests. There are currently six commonly used panel unit root tests: the Im, Pesaran and Shin (2003) IPS test, the Maddala and Wu (1999) Fisher-type test based on the ADF, the Choi (2001) Fisher-type test based on the Phillips-Perron test, the Levin, Lin and Chu (2002) LLC test, the Breitung (2000) test, and the Hadri (2000) test. The first five tests listed are structured under the null of non-stationarity, while the Hadri test is based on the null of stationarity. In addition, heterogeneous unit roots are allowed in the IPS, Fisher-ADF, and Fisher-PP tests, while homogenous unit roots are assumed in the LLC, Breitung and Hadri tests. This has important implications for the interpretation of the results from the tests. For example, in tests where heterogeneous unit roots are permitted, the implication is that rejection of the null hypothesis does not necessarily mean that all of the panel members are stationary. On the other hand, for the tests with the assumption of homogenous unit roots, the implication is that all panel members are stationary. One of the most recent papers carried out using panel unit root tests in respect of PPP was carried out by Kalyoncu and Kalyoncu (2008). They use the IPS test across a panel of real exchange rates of 25 OECD countries over the period 1980 to 2005 and find evidence in favour of long-run PPP. The finding that panel unit root tests of real exchange rates appear to support PPP in the long-run, as opposed to univariate tests, has been made in a range of other papers.[25]

An alternative approach to utilising panels was devised by Pedroni (2004): panel cointegration. His approach is powerful in that it allows for the maximum amount of heterogeneity that is possible within the panel. He uses this method to provide evidence of PPP for the post-Bretton Woods era. Larsson, Lyhagen and Lothgren (2001) devised an alternative approach to panel cointegration. Their method is based on the Johansen cointegration procedure. The Larsson et al (op. cit.) test is similar to the Pedroni procedure in the sense that it allows for heterogeneity in both the parameters and the dynamics. It is computationally less complex, however, whereby the trace statistic for the panel is calculated as the average of the trace statistics attained for the individual series using Johansen (these results are then standardised). The number of cointegrating vectors evident in the panel follows the normal distribution (see, for example, Sideris, 2006, for recent applications to PPP).

In evaluating panel-based approaches, it is important to bear in mind the important concern raised by Taylor and Sarno (1998). They noted that these tests were based in testing the null hypothesis of non-stationarity across all members of the panel. Rejection of the null therefore implies that at least one of the real exchange rates is stationary. This point has tended to be largely ignored in the literature, certainly up until the end of the 1990s, where studies that rejected the null inferred stationarity across all of the panel

[25] For further details, please see Wu (1996), Oh (1996), Frankel and Rose (1996), Papell (1997), Lothian (1997), O'Connell (1998), Taylor and Sarno (1998), Papell and Theodoridis (1998), Fleissig and Strauss (2000), and Wu and Wu (2001).

members. Taylor and Sarno (1998) attempted to resolve this problem by developing a new type of test that poses the hypothesis that at least one real exchange rate within a panel contains a unit root. Rejection of this hypothesis can be interpreted as indicative of stationarity across all panel members. This type of test suffers from severe power problems, however, and this concern over panel results continues to remain (Taylor and Taylor, 2004).

II.B6 Non-Linear Behaviour of the Exchange Rate

In recent years, there has been a movement towards examination of the non-linear dynamics of exchange rate behaviour. This has its foundations in Meese and Rogoff (1983) who noted that random walk models were more successful than linear nominal exchange rate models in describing exchange rate behaviour. Subsequent studies have also provided additional weight to this argument.[26] The economic justification for non-linear models is based on the idea that while exchange rate models are typically comprised of appropriate fundamental variables, the relationship between the exchange rate and market fundamentals is not linear (Creedy and Martin, 1997). In other words, the speed of adjustment towards PPP may be greater as the deviation from PPP increases in absolute terms. This, of course, contrasts with the linear approach whereby the speed of adjustment is assumed to be constant across all sizes of deviation (Taylor and Taylor, 2004).

Non-linear estimation has become very popular over the past number of years in applied time series research (Potter, 1999). Some recent work has shown that using a non-linear framework PPP holds in the long-run, and that deviations of the real exchange rate from PPP levels tend to fade over time (e.g. Michael, Nobay and Peel, 1997; Sarno, 2000;

[26] For example, Sarno (2000); Choi and Saikkonen (2004); Haug and Basher (2005); and Dufrenot et al (2006).

Taylor, Peel and Sarno, 2001; Sarno and Taylor, 2001). This would imply that the real exchange rate is stationary.

The reasoning behind the non-linearity in such studies is due to the transactions costs associated with arbitrage. In relation to the goods market, the strict version of PPP assumes perfect arbitrage across the commodities between countries. Due to transactions costs and other market frictions, however, arbitrage is affected whereby a virtual zone of inactivity exists around the real exchange rate. Here, there is no arbitrage due to its excessive cost. Once outside the inactivity zone, arbitrage resumes and the real exchange rate moves towards its long-run equilibrium (until it swings back into the zone of inactivity). The idea is that the larger the deviation of the real exchange rate from the PPP level, the more rapid the mean-reversion process (Taylor et al, 2001). This is because greater deviations from PPP are associated with higher arbitrage profits. An asset-market based rationale for non-linearity in the real exchange rate could be based on the heterogeneous actions of foreign exchange traders (Schnatz, 2006). When the exchange rate is close to its equilibrium level, then its dynamic is driven primarily by the technical analysis of so-called 'chartists'. Here the random walk hypothesis prevails. In the scenario where the exchange rate is far away from its equilibrium level, the dynamic is then driven by economic fundamentals, whereby the so-called 'fundamentalists' drive the wandering exchange rate back to its long-run equilibrium. Thus, the non-linearity argument is founded upon the idea that the real exchange rate appears to revert to its mean only when it is deemed to be misaligned from its equilibrium level. While it is close to this level, but not deemed to be misaligned, the behaviour of the real exchange rate is more akin to a random walk.

As has already been outlined, a puzzle exists in relation to why the speed of adjustment to PPP is slow in spite of high short-term volatility. The non-linear approach has been cited as one potential means by which to explain this. Within a linear framework, by definition, the speed of the deviation from PPP is uniform. Under such circumstances, the standard estimation procedure is based on estimating the half-life, i.e. how long it takes for one half of the deviation from PPP to be eliminated. In reality, however, the speed of adjustment may not always be uniform. This may be due to the presence of 'market frictions', such as transport costs and trade tariffs. Such frictions may enable a deviation from PPP without an associated rise in goods arbitrage (O'Connell, 1998).

This non-linear idea goes back almost 100 years when Heckscher (1916) made this same point. Here, the idea is that if PPP fails and the price of two goods are not the same, it is futile to engage in arbitrage unless the expected profit is greater than the associated transaction costs (such as shipping costs). The transaction cost issue in particular has been the focus of the work of a number of researchers in this area. For example, Benninga and Protopapadakis (1998) and Sercu, Uppal, and Van Hulle (1995) have provided some evidence that the slow speed of adjustment to PPP could be explained by the presence of transaction costs. Other work carried out by Dumas (1992), Williams and Wright (1991), Uppal (1993), and Coleman (1995) has indicated that with transaction costs present, the adjustment process to PPP levels is not necessarily linear. Non-linearity in exchange rate behaviour could also be explained by the activity of chartists in international markets, which cause movements in the exchange rate which are not necessarily related to underlying fundamentals.[27]

Non-linear tests of PPP are based on models that allow variation in the autoregressive parameter. For example, threshold autoregressive (TAR) models allow the mean reversion process to take place as long as the real exchange rate lies outside a particular threshold level that is equal to the transaction cost. Once outside this level, the arbitrage

[27] For example, Taylor and Allen (1992) and McMillan (2005).

process induces mean-reversion. This occurs until the real exchange rate falls below the threshold level, where arbitrage is no longer profitable and random walk behaviour is observed. Empirical work done in this area includes Canjels, Prakash-Canjels and Taylor (2004), who found that the US dollar/sterling exchange rate was consistent with a TAR model over the period 1879 to 1913. Other work identifying non-linear real exchange rate behaviour using TAR models include Obstfeld and Taylor (1997), Sarno, Taylor and Chowdhury (2004), and Zussman (2003).

A more sophisticated class of non-linear models take into account the fact that transaction costs are different across goods, thus implying a difference in the speed at which goods price arbitrage occurs. Under such a scenario, there may exist a range of threshold levels. Extended versions of the TAR to take this into account include smooth transition autoregressive models (STARs) and the exponential STAR (ESTAR). The STAR model was developed by Granger and Terasvirta (1993), and allows smooth adjustment of the real exchange rate whereby the adjustment speed rises as the deviation from PPP rises. These models are structured under the null hypothesis of a unit root versus the alternative of non-linear mean-reversion. Taylor, Peel and Sarno (2001) used a STAR approach to find evidence for non-linear mean reversion of the dollar real exchange rates of the G-5 countries. They also find that the half-life is smaller the larger the shock to the real exchange rate. The ESTAR model is an extension of the TAR model that takes into account an infinite number of regimes and a continuously varying and bounded adjustment speed (Taylor and Taylor, 2004).

The growing literature in the non-linear field is essentially based on dissatisfaction among applied econometricians with the standard linear autoregressive moving average (ARMA) framework to test for unit roots (Kapetanios and Shin, 2002). As will be discussed in more detail in Section III, standard unit root test procedures can suffer from power problems. In fact, previous research carried out has shown that if standard unit root tests have no power within test of time series which exhibit non-linearity.[28] Consideration of a non-linear framework represents one method to overcome this issue. Clearly, finding the presence of a unit root in the real exchange can have ramifications for exchange rate policy. For example, under this scenario there would be some serious concerns about using PPP as a model of exchange rate determination. Therefore, it follows that as well as examining the nature of the equation used to estimate the presence of a unit root, it is also imperative to take into account potential inherent non-linearity.

II.C Concerns Regarding the Tests Employed in the Literature

The main concern in relation to PPP tests relates to the specification of the equation used. The dynamic can involve a tri-variate test of the relationship between the nominal exchange rate, the domestic price series and the foreign price series, a bi-variate test of the relationship between the nominal exchange rate and the ratio of domestic and foreign prices, or a uni-variate test of the real exchange rate. The most general case is provided by the former model, whereby there are no restrictions imposed on either proportionality (price coefficients restricted to be [1, -1] or symmetry (price coefficients have the same magnitude but opposite sign). The bi-variate test imposes the symmetry condition, while both proportionality and symmetry and imposed with the univariate case. While traditional PPP requires that the proportionality and symmetry conditions must hold (and therefore largely must be imposed), Cheung and Lai (1993a) have noted that the imposition of these restrictions can influence the cointegrating relationship. Thus, their preference in testing for PPP is for the tri-variate test, which places no restrictions on the price coefficients.

[28] For example, Kilian and Taylor (2001) and Taylor et al (2001) provided evidence to show that many time series are in fact non-linear in nature and that standard ADF unit root tests do not have power within this framework.

A key point to note from the literature is that the form of the equations used to estimate PPP relationships has tended to include an intercept term. Whilst the tests impose a zero restriction on the intercept, the inclusion of this term in the estimation affects the size of the critical values for testing stationarity. In the following section, an application is performed to illustrate that perhaps there should not be an intercept in estimated PPP equations. This is consistent with the 'purist' view of PPP, whereby there should be no drift in the PPP relationship. At the very least, the significance of the intercept should be examined in the first instance prior to its inclusion into models (the literature does not appear to have examined such significance). This same situation should be in place for all deterministic components.

The next section considers the stationarity of the real exchange rate across a number of currencies over the period 1982 to 2005 across specifications with varying deterministic components.[29] It is shown that careful consideration must be given to these components and that they must be tested appropriately prior to inclusion in models.

III. Application Highlighting Importance of Deterministic Components in Unit Root Tests

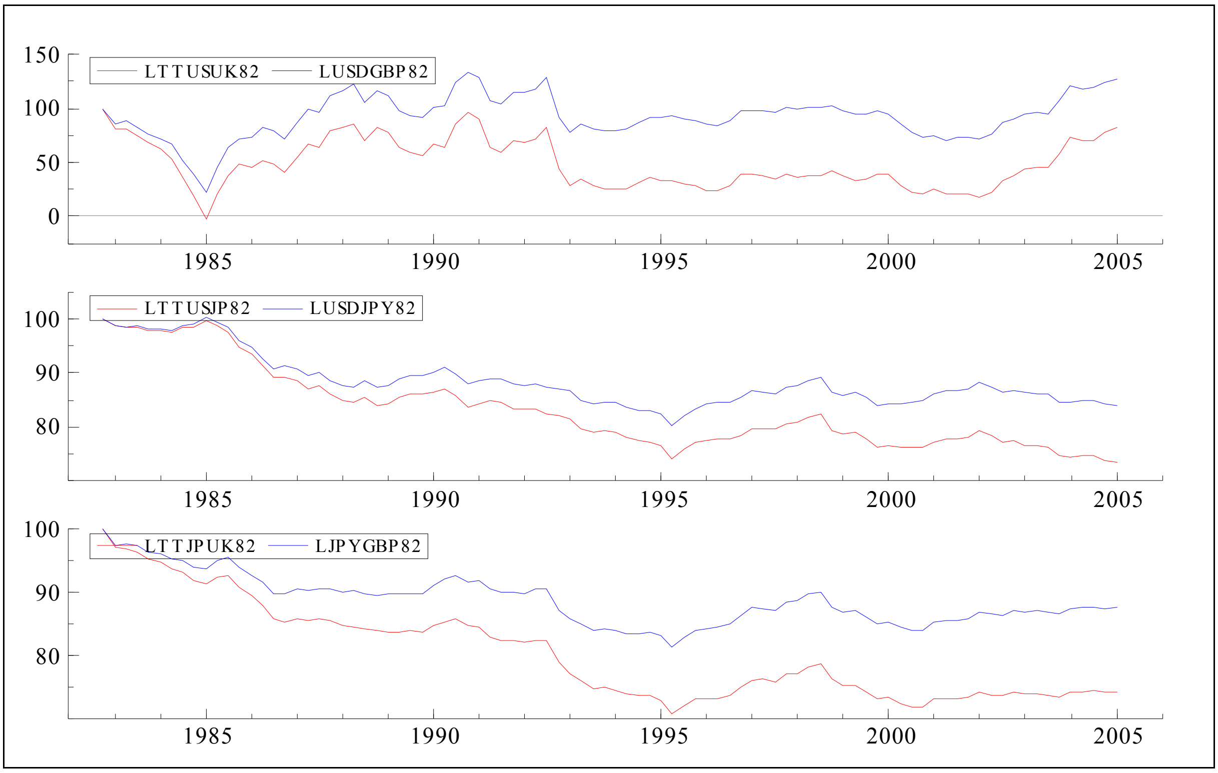
III.A Data and Preliminary Analysis

The data (quarterly frequency) examined consists of the following variables: US Dollar/UK Pound nominal exchange rate (USDGBP); US Dollar/Japanese Yen nominal exchange rate (USDJPY); Japanese Yen/UK Pound nominal exchange rate (JPYGBP); US/UK Terms of Trade; US/JP Terms of Trade; and, JP/UK Terms of Trade.

[29] The year 1982 was selected as the start of the period due to the some of the price series exhibiting I(2) behaviour in years prior to 1982. The analysis runs to 2005 since that was the latest data point available when the analysis was carried out.

As a starting point to analysing the time series structure of the data, it is useful and informative to observe the relative fluctuations of the nominal exchange rate and terms of trade over the period 1982 to 2005 for the data in question. **Figure 2.1** below plots the logarithms of the nominal exchange rate and terms of trade for two dollar rates and a cross rate over the period 1982 to 2005.

Figure 2.1 Logs of the Nominal Exchange Rate and Terms of Trade (1982=100)



(Note: the terms of trade is defined as the inverse of the real exchange rate, or 1÷SP*/P, where S is the nominal exchange rate, P* is the foreign price, and P is the domestic price).

It is clear from the plots that the nominal exchange rates and the terms of trade have tended to move together over the 1982-2005 period in the case of the US and the UK. Hence, the real exchange rate is not particularly informative in terms of helping to explain movements in the nominal exchange rate for this case, notably with regard to fluctuations from the PPP level. Clearly, other forces are responsible for the deviation in the nominal exchange rate from its PPP level. This could imply that there exists a need to focus on other factors which push the nominal exchange rate from its PPP level, such as the nominal interest rate perhaps.

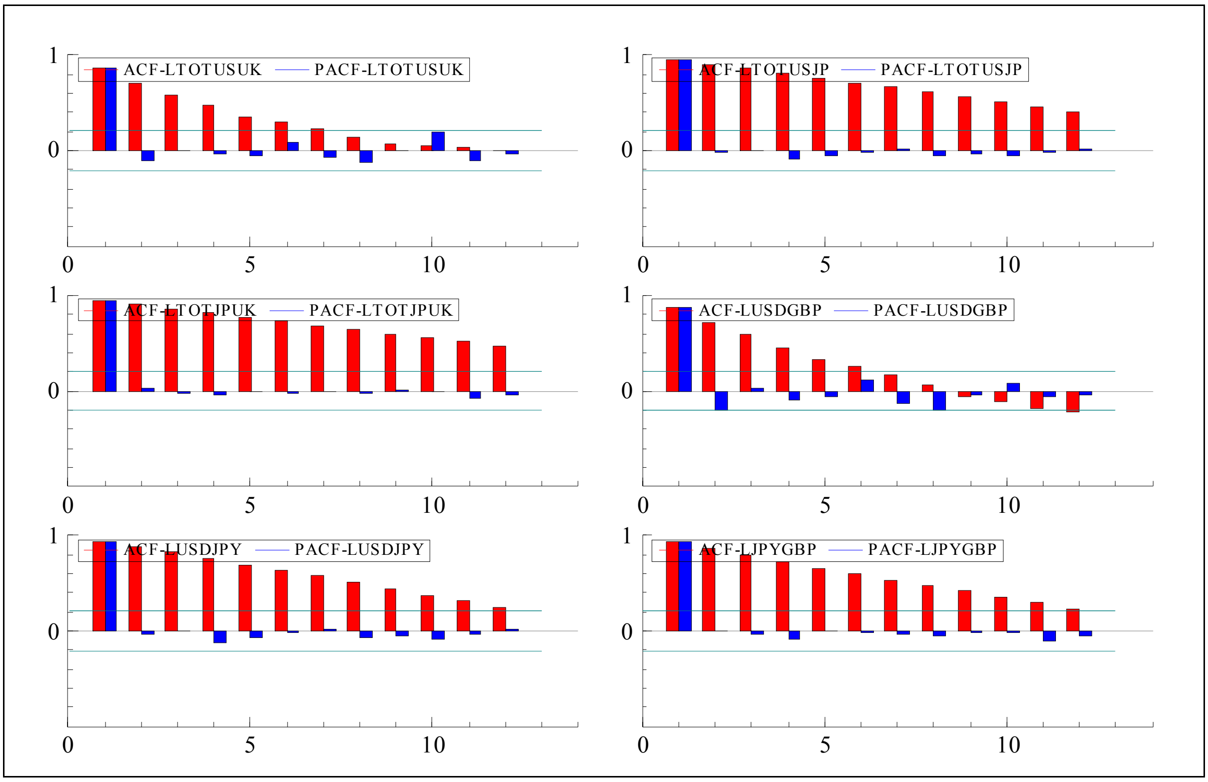
With regard to the US/Japan and Japan/UK data, however, the plots suggest some progressive deviation from between the terms of trade and the nominal exchange rate, particularly in the case of the latter, i.e. the cross rate. Japan is often cited as providing a good example of the Balassa-Samuelson effect.[30] Therefore, the fast rate of productivity in Japan could be one reason for the apparent divergence in the paths of the nominal exchange rate and terms of trade vis-á-vis the dollar. This is not entirely satisfactory, however, given the relative stagnation of the Japanese economy throughout much of the 1990s.

Research carried out in the 1990s, however, helps to provide a further rationale for the divergence in spite of low economic growth in Japan. De Gregorio, Giovannini, and Wolf (1994) have shown that demand factors (such as real government consumption) may generate deviations from PPP if there is a bias towards the service sector (since this will increase the price of non-tradables.

The autocorrelation and partial autocorrelation functions for each of the six variables considered as illustrated in **Figure 2.2**.

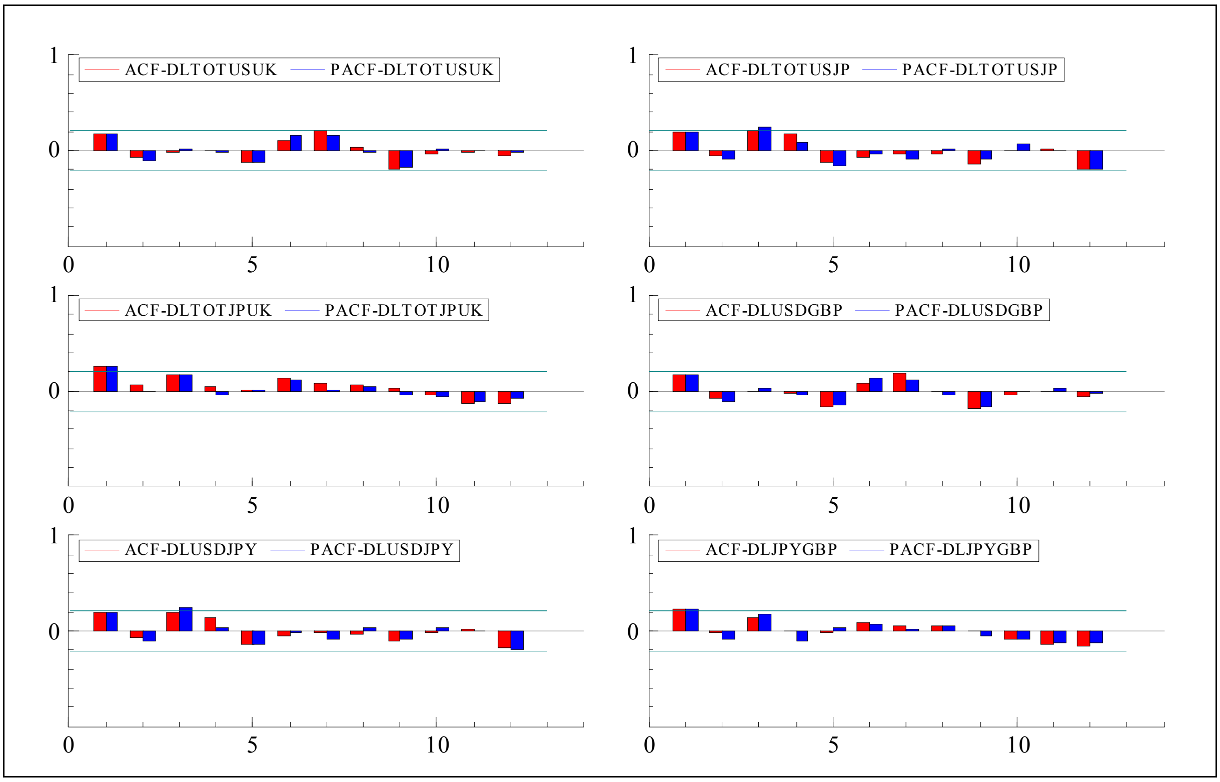
[30] Rogoff (1996) Marston (1986) provide strong evidence for Balassa-Samuelson effect in JPYUSD. The latter provided specific evidence to show how productivity in the traded sector in Japan outstripped that of the US.

Figure 2.2 ACF and PACF for each series (levels)



Regarding the ACF, it is clear that there is a gradual decline across each of the series. In addition, after the first lag the PACF drops dramatically. All of this suggests the series is non-stationary in levels. As is known from theory, a slowly decaying ACF in conjunction with one large spike in the PACF is consistent with AR(1) behaviour. These series would appear to follow random walks, i.e. indicative of non-stationarity.

Figure 2.3 **ACF and PACF for each series (first differences)**



Here the pattern of the first differenced ACF and PACF is smoother than in the levels case. This indicates that after differencing once, stationarity is evident. However, the question remains as to whether non-stationarity can be conclusively proved. In addition, the question remains as to whether the degree of non-stationarity can be improved by amending the standard form of the equation used to conduct unit root tests and other procedures for estimating stationarity.

III.B Influence of the Intercept Term

The particular focus here is on the concept of PPP and stationarity testing in the absence of an intercept term from the estimation equation. The contention is that the absence of the intercept term is not only more consistent with the foundations of PPP, but also that it produces results which are more non-stationary in nature than the alternative (i.e. by including an intercept term). This also implies a coherence between tests of stationarity that can be viewed as a restriction on equation (10) and the tests applied to the residuals of (10).

Recall equation (10) from **Section II.B** that the standard form of equation used to estimate PPP relationships was of the form:

$$s_t = \alpha + \beta p_t + \beta^* p_t^* \omega_t \tag{23}$$

It is important to note the role played by the inclusion of the intercept term in the equation, as this may have consequences in defining the degree to which stationarity (or non-stationarity) is observed. Observation of the logs of the levels would appear to suggest non-stationarity in the terms of trade. This can be formally tested using an Augmented Dickey Fuller (ADF) test on the following model:

$$\Delta z_t = \pi_0 + \gamma z_{t-1} + \Sigma \pi_i \Delta z_{t-i} + \varepsilon_t \tag{24}$$

This is the most common test for a unit root in univariate series, which (generally) regresses the first difference of the variable on an intercept, the lagged level of the variable and k lagged first differences.

There is some evidence to suggest, however, that shortcomings exist regarding the standard testing procedures for stationarity. For example, it is well-documented that ADF tests can suffer from power problems, resulting in an acceptance of the null hypothesis when the alternative is true. It is important to point out that ADF tests are sensitive to initial conditions within a series, or μ_0. This would provide a reason to include an intercept term in the estimated unit root test. In order to proceed with these tests in the absence of an intercept term, then it is appropriate to recursively de-mean the series.[31]

[31] See Taylor (2001; 2002) for more information on the recursive de-meaning procedure as it pertains to unit root testing.

In undertaking these tests, the optimal lag length is ascertained using the AIC and the tests are carried out across three scenarios – with inclusion of an intercept term and time trend, with an intercept and no time trend, and without inclusion of an intercept or trend. In undertaking this analysis, the procedure of Dolado, Jenkinson and Sosvilla-Rivero (1990) is employed. This procedure provides a framework for the elimination of deterministic regressors that are deemed to be inappropriate for a particular model. Typically, the least restrictive model is estimated, i.e. the model including a time trend and an intercept. If the null hypothesis of a unit root is not rejected, we then test the significance of the trend term *given* a unit root. If this is deemed not to be significantly different from zero, we can eliminate it, and re-estimate the model using a more restrictive model, i.e. with an intercept and without a trend. The same procedure is followed to test the significance of the intercept and to ascertain whether or not it should be included in the estimated equation. If not, it is eliminated and we estimate the most restrictive case with no intercept or trend. The procedure outlined is applied to our data as follows.

III.C Results and Discussion

Table 2.5 ADF Tests for Unit Roots*

Series	Model 1 Intercept and Trend		Model 2 Intercept, No Trend		Model 3** No Intercept, No Trend	
	Levels	First Differences	Levels	First Differences	Levels	First Differences
LTOTUSUK	-2.58 (1) [2.18] {-0.41}	-7.86 (0) [-0.79] {0.87}	-2.72 (1) [2.64]	-7.83 (0) [-0.06]	-0.70 (10)	-3.19 (9)
LTOTUSJP	-1.81 (1) [-1.66] {1.10}	-7.66 (0) [1.92] {-1.08}	-1.94 (3) [-1.80]	-3.95 (2) [1.56]	-1.82 (11)	-2.07 (10)
LTOTJPUK	-1.52 (1) [1.34] {-0.58}	-7.61 (0) [-2.46] {1.77}	-2.28 (1) [2.10]	-7.32 (0) [-1.89]	-1.81 (12)	-2.30 (11)
LUSDGBP	-2.68 (1) [-2.40] {-0.82}	-7.84 (0) [0.11] {-0.32}	-2.58 (1) [-2.60]	-7.87 (0) [-0.37]	-1.49 (11)	-1.95 (10)
LUSDJPY	-1.85 (1) [1.74] {-0.67}	-7.64 (0) [-1.50] {1.05}	-2.28 (3) [2.22]	-3.98 (2) [-1.00]	-1.54 (12)	-2.46 (11)
LJPYGBP	-2.24 (3) [-2.15] {-0.87}	-4.31 (2) [1.41] {-1.23}	-2.42 (3) [-2.39]	-4.12 (2) [0.69]	0.91 (10)	-2.48 (9)

* ADF Critical Value (5%) is –3.46 for Model 1, –2.89 for Model 2, and –1.94 for Model 3. Beneath the ADF test statistics, the optimal lag lengths are reported in round brackets, the significance of the intercept term is reported in square brackets, and the significance of the trend is reported below this (both t-statistics).

** Tests carried out on recursively de-meaned series.

Looking at Model 1 in the first instance, it is clear that the ADF test statistic is lower than the ADF critical value of –3.46. This implies that we cannot reject the null of a unit root with the trend and intercept included in the estimated regression. In fact, the inclusion of these terms may be reducing the power of the unit root test. Now we must test the significance of the trend term, given a unit root. At the 5% level, the critical value is 2.79 (Dickey-Fuller tau statistics). All of the computed t-statistics across all series in Model 1 are lower than 2.79, implying that the trend term is not significantly different from zero.

Now we can move on to estimate Model 2, i.e. the more restrictive model with an intercept but without a trend. As can be seen, all of the computed ADF test statistics are below the 5% critical value, implying non-rejection of the null of a unit root (i.e. the presence of non-stationarity). Since this has been established, we can now test the significance of the intercept given a unit root. The appropriate tau statistic at the 5% level is 2.54. Looking at the computed t-statistics for the intercept across each of the series, it would seem that four of the six series should not contain a drift term in the estimated regression, i.e. in these case the intercept is deemed to be not statistically different from zero. At the 5% level, it seems that only LTOTUSUK and LUSDGBP can warrant inclusion of a drift term, based on these results. The relevant tau statistic at the 1% level is 3.22, however.

Based on these results, there would appear to be some evidence to warrant removal of the intercept term from this particular set of series. In four of the series, the intercept is not statistically different from zero at the 5% level, while for the remaining two series it is not statistically significant at the 1% level.

In addition, it seems that the estimations undertaken in the absence of an intercept term appear to be more supportive of stationarity post-differencing once. In order to ensure robustness to the stationarity test, the same procedure is applied using the Phillips-Perron indicator.

Table 2.6 Phillips-Perron Tests for Unit Roots

Series	With intercept		Without intercept*	
	Levels	First Differences	Levels	First Differences
LTOTUSUK (k=3)	-2.93 [0.0112]	-7.79 [0.9527]	-0.37	-16.05
LTOTUSJP (k=3)	-2.03 [0.0622]	-7.58 [0.0491]	-0.25	-83.50
LTOTUKJP (k=3)	-3.05 [0.0026]	-7.35 [0.0619]	-0.24	-88.99
LUSDGBP (k=3)	-2.40 [0.0410]	-7.83 [0.7155]	-0.04	-12.45
LUSDJPY (k=3)	-2.28 [0.0281]	-7.56 [0.2323]	-0.15	-88.91
LJPYGBP (k=3)	-2.89 [0.0041]	-7.50 [0.3931]	-0.11	-96.54

* PP Critical Value for 'with intercept' is –2.89 at 5% level and 3.51 at 1% level; for 'without intercept' is –1.94 at 5% level and –2.59 at 1% level.

* Tests carried out on recursively de-meaned series.

The PP tests appear to indicate that the removal of the intercept from the equation yields more conclusive results in terms of the stationarity of each of the series after differencing occurs once. Based on conventional inferencing, therefore, there would appear to be *prima facie* evidence that the removal of the intercept terms yields results which are more non-stationary than the alternative.

IV. Coherence Between Univariate and Panel Tests of PPP[32]

The coherence between univariate and panel tests of real exchange rate stationarity is now examined. It has already been shown that the literature indicates a greater degree of power with panel unit root tests as compared with univariate tests. Such results are compared with alternative panel estimators where individual series undergo similar mean adjustment. Much of the recent evidence would suggest that PPP does not hold as the hypothesis that real exchange rates are stationary is commonly rejected. Abuaf and Jorion (1990) found that univariate tests based on the null of a unit root when applied to single currencies have tended to accept the null of non-stationarity. However, Hunter and Simpson (1995) found a stationary cointegrating vector accepting the PPP restriction, for a system estimated to explain the UK effective exchange rate.[33] Lothian and Taylor (1996), using long time series averages, observed that as the period increases the strength of correlation between the exchange rate and relative prices increased with the length of period used. They concluded that this evidence supported the proposition that PPP holds in the long run. Empirical evidence also exists that is supportive of stationarity when cross sectional dependence is considered, e.g. Luintel (2001)[34]; Im, Pesaran and Shin (2003). However, Sarno and Taylor (1998) suggest that panel tests of stationarity may often be dominated by a small number of stationary and Caner and Killian (2001) call into question the power of many of the tests used.

By constructing the univariate and panel tests in a coherent and consistent framework may lead to stationarity results that are largely consistent across both types of tests. Specifically, the univariate tests are constructed by recursive mean adjustment and careful consideration of the dynamic forcing the real exchange rate. The analysis considers the

[32] A published version of this analysis appears as Beirne, Hunter and Simpson (2007).

[33] This is distinct from the evidence in Hunter (1992), Johansen and Juselius (1992) and Juselius (1995), for which the PPP has to be augmented by interest rates and similar results for Germany and Italy presented in Simpson (2002).

[34] The work of Luintel (2001) was applied to an earlier unpublished version of the Im, Pesaran, Shin (2003) paper.

logarithm of the real dollar exchange rate series for twelve European countries. All of the series are demeaned prior to testing in order to correct for initial conditions (Tremayne, 2006; and Haldrup & Jansson, 2006). Since the transformed data defines a mean zero series, the univariate and panel tests exhibit similar size.[35] The mean-adjusted univariate results are reported because they support the panel results and refute the Sarno and Taylor (1998) argument that stationary panel results are the result of the presence of a small number of stationary individual series. In addition, Taylor (2002) notes that the recursively demeaned series are used as they represent similar tests across univariate and panel cases, and they would also appear to have greater power than the GLS-corrected Dickey-Fuller test developed by Elliot, Rothenberg and Stock (1996). The models are all well defined in terms of serial correlation, but where appropriate, heteroscedasticity consistent standard errors are used to take some account of the fact the disturbances may not all be identically distributed and correct those models where volatility is observed for ARCH behaviour in the disturbances. The corrected univariate results would suggest that on average the real exchange rate is stationary - a proposition supported by the panel analysis.

In conclusion, this would appear to be a case where the concerns voiced by Sarno and Taylor (1998) do not apply or rather the panel data is on average stationary with white noise disturbances. To support the univariate and panel analysis the proposition that the real exchange is stationary is compared to Hunter and Simpson (2001), who used the test due to Hadri (2000), because this test operates under the null of stationarity and unlike similar tests considered by Caner and Killian (2001) has excellent size and power for the sample size available and is robust to non-normality. This test also takes account of heterogeneous serial dependence using a non-parametric correction. The comparison

[35] The recursive correction appears to be similar in nature to the Generalized Least Squares correction developed by Elliot, Rothenberg and Stock (1996), which yields a similar asymptotic distribution to the model without constant (Davidson and MacKinnon (2005)).

with the Hadri test is also consistent with the argument made in Kwiatowski et al (1992) that the KPSS test should be confirmed or reinforced with a test based on the null of non-stationarity.

In the first instance the recursive mean adjustment suggested by Taylor (2002) is applied to the univariate series and in the case of the t-bar test of Im et al (2003) the data are considered relative to their cross sectional country means. The results from this analysis are then compared to the Hadri test results performed by Hunter and Simpson (2001) on the same set of recursively mean-adjusted univariate data. The sample selected is 1980q1 – 1998q1, thus avoiding the period of the 1970s when US prices were integrated of order two (I(2)) and inflation rates for some countries were considered non-stationary, and ends prior to the formation of the euro area.

The mean adjustment of the single time series removes an initial value problem and it also implies that a similar test of the null of non-stationarity is applied to both the univariate series and the panel as the transformed series have a zero intercept. The Hadri test operates under the null of stationarity and it is shown by Hadri that for 50 time series observations the test is appropriately sized and has excellent power even when the local alternatives are close to the null.

Quarterly observations on dollar real exchange rates[36] were drawn from the Datastream database for the period (1980q1 – 1998q1) for twelve countries: Italy, Spain, Belgium, Denmark, Finland, France, Germany, Ireland, Luxembourg, Holland, Portugal and the UK. As a preliminary stage to the main analysis, Augmented Dickey-Fuller tests (Dickey and Fuller (1979)) were applied to each exchange rate in turn. It is important to note that the Dickey-Fuller test is sensitive to initial conditions, dynamics in both the conditional variance (ARCH) and the mean equation (serial correlation), and non-normality. Thus

[36]The real exchange rates used are based on relative consumer price indices.

much of the recent literature has suggested that there might be some form of non-linear adjustment in exchange rates.

Therefore, prior to specifying the time series auto-regressive model from which the ADF test is derived, the data is transformed by recursively de-meaning the series in turn using the procedure described by Taylor (1999). The correlogram for the series under the null of non-stationarity is assessed to determine the order of dynamic for each model of the real exchange rate, i.e. to determine the maximum lag order (Burke and Hunter, 2005, Chapter 2). This corresponds with the view presented in Said and Dickey (1984) that long order AR models improve size, though the introduction of redundant lagged terms may also lead to a loss of power (Haldrup and Jansson, 2006). To improve the power of the univariate tests, a General-to-Specific approach is adopted (Taylor, 2002) and intermediate lags that are insignificant at the 10% level are discarded based on conventional inference. In addition, on the basis of both conventional inference and simulated critical values the intercept term is excluded.[37] As there is no trend in the original data and the asymptotic distribution of the Dickey Fuller test is not sensitive to the inclusion of differenced series, the test is applied to the following model:

$$\Delta x_{it} = \gamma x_{it-1} + \sum_{j=1}^{l} \pi_j \Delta x_{it-j} + \varepsilon_{it}. \tag{25}$$

where $x_{it}=y_{it}-\sum^{t}_{j=1}y_{it-j}/t$ is the recursively de-meaned real exchange rate for country i.

The results in **Table 2.7** compare critical values for equation (25) calculated under the null of non-stationarity (γ=0) for a sample of 68 observations. Using the 95% critical

[37] As can be observed from **Table 2.7**, the intercepts for the demeaned series are small and when an intercept is included in the regression these are neither significant based on conventional inference or on the basis of critical values simulated under the null associated with the Dickey Fuller in the case with an intercept (1000 replications generated with an intercept in the regression and the standard deviation being 1 and .0387 were using the AR(1) model in first differences in PC-Naïve (Doornik and Hendry, 2006) which gave rise to t-values -3.0345 and -3.4051 respectively for an intercept of -.008).

values simulated by Ox as –1.9714[38], the real exchange rates of Belgium, France, Finland, Germany, Holland, Ireland and the UK[39] are stationary. Based on a similar argument to Phillips and Perron (1988) we provide corrected standard errors to determine the tests for stationarity. Firstly, we apply White's (1982) standard errors to the conventional Dickey-Fuller t-test and this yields a standardised estimate of the residual variance that has no significant impact on the behaviour of the mean equation.[40] This is useful for a number of reasons: alternative estimates of the standard errors are less sensitive to the risk of pre-test bias that may arise due to the initial exclusion of variables in the auto-regressive equations that describe the real exchange rate, and they remove the impact of extreme observations on the estimate of the error variance.

White standard errors are used to correct the error variance for the undue influence of large observations reflected in the Jarque-Bera statistics reported in **Table 2.7**. As a result of this transformation all the Dickey-Fuller test statistics increase except for Spain. Using inference based on White Standard errors, [41] the real exchange rate is stationary for eight out of the ten countries at the nominal 5% level, while in the case of Denmark this is the case when the test is applied at the 10% level, but for Spain stationarity cannot be accepted at any conventional level of significance. It is of particular interest to note that the substantive increases in the test statistic occur when the errors are non-normal.

In the case of two countries, Luxembourg and Portugal, there is evidence from column four of **Table 2.7** of Auto-regressive Conditional Heteroscedasticity (ARCH). Here, the

[38] These simulations, carried out by Hunter, are generated for T=68 and B=10000 replications. The data are calibrated using a typical real exchange rate series with the recursive mean transformation used to remove the initial condition, for a nominal size of the test of 95%, the rejection frequency is 96.3% so the tests are slightly undersized leading to an over acceptance of the null. The true critical value ought to be greater than the value we have simulated.
[39] Luxembourg has the same exchange rate as Belgium, but a different price series.
[40] This is essentially the same as applying the first term in the semi-parametric correction used by Phillips and Perron (1988), but here we are not concerned with serial correlation. However, we would anticipate with some large outliers that the White Standard errors would yield more robust inference, while making little difference to the underlying distributions associated with the test of stationarity.
[41] Based on the same simulated data as before the corrected test has the following critical value, -1.6861 (90%) and -2.0244 (95%) with the empirical size of the test being 92% and somewhat better than the conventional at 0.95885%.

variance estimates are improved by estimating the ADF model using the ARCH estimator to correct for observed volatility. Boswijk (2001) has suggested that by modelling the volatility the power of unit root tests might be greatly improved. In the case of Luxembourg and Portugal the standard errors are derived using the ARCH (1) and restricted ARCH (4) estimator. It is suggested in Boswijk (op. cit.) that for the near integrated case conventional inference should be acceptable as long as the two Brownian Motions driving the process are not correlated, which would appear to be the case for GARCH and any other process explaining the volatility *"as long as they have a continuous-time diffusion limit"* (p.16). In this light we assume for ARCH processes that are not integrated, that the asymptotic theory goes through, which for the case of Luxembourg leads to the acceptance of the alternative of stationarity at the 5% level. However, based on all of the inferential procedures adopted here, it is not possible to accept that the real exchange rate in the case of Portugal is stationary. Given the likely misspecification that arises with ARCH it would appear more important to use correct standard errors, than using those derived from OLS that are biased and in some instances inconsistent. If one were to accept conventional inference, then the models do not suffer from serial correlation, heteroscedasticity and non-linearity. However, in all but two cases the errors are non-normal and in two cases there is significant ARCH behaviour, which leads to adopt a non-standard approach.

Table 2.7 Summary of Augmented Dickey Fuller Tests

Country	t-OLS/t-W	Mean	ARCH(4)	BP(9)	JB(2)	LR(i)
Belgium	-2.36/-2.96	.000517	1.5619	4.5313	18.4357	2.9582(4)
Denmark	-1.62/-1.7	.0010700	1.9120	9.1640	1.6948	1.3307(4)
France	-2.51/-3.02	.1803E-3	2.3684	4.4180	63.9455	1.9710(2)
Finland	-2.07/-2.17	-.0021874	1.7105	1.4099	41.0066	.10537(2)
Germany	-2.33/-2.51	.0010926	2.3639	6.8697	1.1113	.45821(3)
Holland	-3.28/-3.36	.4953E-3	6.0082	4.7406	3.2804	.33688(3)
Italy	-1.86/ -2.3	-.1929E-3	2.1029	9.5322	139.1943	1.4761(3)
Luxembourg	-1.85/-2.60	.1373E-4	22.0553	2.0923	47.2989	.063297(3)
Ireland	-2.26/-2.12	-.6408E-3	2.8921	6.5168	8.5586	3.0704(4)
Portugal	-.93/-1.08	.0011172	12.7212	4.8049	95.3690	1.7532(3)
Spain	-1.54/-1.47	.1536E-5	5.0087	4.8213	10.1260	2.5905(3)
UK	-2.06/-2.19	.0013465	1.3039	7.0078	13.4754	2.0265(4)

(Note: the selection of the sample of countries used was driven by the availability of consistent data over the period 1980 to 1998).

Abuaf and Jorion (1990) question the usefulness of univariate ADF tests to detect stationary real exchange rates when the sample is short. This has led to tests based on the null of non-stationarity being applied to panel data to improve the power of the test by pooling observations across countries. However, O'Connell (1998) has argued that many of these studies *"fail to control for cross-sectional dependence in the data"* (p.1). Luintel (2001) has addressed this issue by applying the demeaned LM-bar and T-bar tests proposed by Im et al (1997) to data for 20 OECD countries. Luintel (2001) suggests that the finding of stationarity is due to a reduction in the order of cross-sectional dependence and cites the study by Wu and Wu (1999) where tests based on Deutsche Mark rates appear more likely to accept stationarity.[42]

However, in the context of real exchange rates the primary interest is in testing the null of stationarity and subject to an appropriate level for the test it is subsequently important to

[42] It should be noticed that data derived from cross rates embodies an implicit sequence of cross arbitrage conditions, which affect the structure of the underlying model and the validity of tests. See Smith and Hunter (1985) for conventional dynamic models, and models that impose PPP and uncovered interest arbitrage, and Hunter and Simpson (2004) for dynamic single equations models, which implies that inference based on conventional Dickey Fuller tests will be distorted as the estimate of the variance-covariance matrix is incorrect under the null and the alternative, and the estimate of γ is biased and inconsistent.

minimise the probability of wrongly rejecting the alternative by selecting a locally most powerful test. Unfortunately, Taylor and Sarno (1998) have voiced concerns about tests based on the null of stationarity, for which the null has been accepted even when a single series in the panel is truly stationary. The issue of the appropriate null for many of the conventionally used tests has also concerned Caner and Kilian (2001), who found significant size distortion for the KPSS (Kwiatkowski et al, 1992) and Leybourne and McCabe (1994) tests for the null of stationarity.

On the basis of the research presented thus far, it can be stated that by careful analysis of each series in turn that on average the series selected here do appear to be stationary. Of course our single equation analysis would suggest that real exchange rates are fairly heterogenous and this would suggest that one ought not to be engaged in any form of pooling. By applying panel methods on the basis that for the panel the pooled series broadly satisfies the appropriate criterion. The test by Im et al (2003) has the benefit of pooling the t-tests for appropriately calibrated country series, where each univariate series is scaled relative to the cross-section means. Hunter and Simpson (2001) apply the Hadri (2000) test to the same recursively demeaned data used in our IPS and univariate tests. The Hadri test offers significant gains in terms of size and power when compared with other tests developed under the null of stationarity. More importantly, given that the sample under consideration is robust to non-normality and the convergence in distribution occurs quickly in small cross sections with quite modest time series dimensions and a broad range of values for the variance ratio being minimized via the test.

Im et al (2003) suggest a test of stationarity that averages the conventional Dickey Fuller tests across the panel, while Hadri (2000) proposes a Lagrange-Multiplier (LM) test of the null that a series is stationary (either around a deterministic level or a trend). An exact small sample correction to the LM test statistic means that the test is asymptotically normal. Furthermore, Hadri (2000) provides evidence that after correction the test has good size properties and is robust to non-normality.

Luintel (2001) addresses this issue by applying the earlier version of the tests proposed by Im et al (2003) to data for 20 OECD countries. Therefore the real exchange rate equation without transformation:

$$\Delta y_{it} = \pi_{i0} + \gamma_i y_{it-1} + \sum \pi_{ij} \Delta_1 y_{it-j} + \varepsilon_{it}$$

when de-meaned is:

$$\Delta \tilde{y}_{it} = \tilde{\delta}_{t\tau} + \beta_i \tilde{y}_{it-1} + \sum \theta_{ij} \Delta \tilde{y}_{it-j} + \xi_{it} \qquad (26)$$

where: $\tilde{y}_{it} = y_{it} - \bar{y}_t$, $\tilde{\delta}_{it} = \delta_{it} - \bar{\delta}_t$, $\tilde{\varepsilon}_{it} = \varepsilon_{it} + \theta_t - \bar{\varepsilon}_t$, θ_t is the time-specific common fixed effect and $\xi_{it} = \tilde{\varepsilon}_{it} + \frac{1}{N}\sum_{j=1}^{N}(\beta_i - \beta_j) y_{jt-1}$. We consider the t-bar test or average Dickey Fuller test based on estimating (26) for each cross section observation by calculating:

$$\bar{t}_{NT} = \frac{1}{N}\sum_{j=1}^{N} t_{iT}$$

The null tested is that all the coefficients are consistent with non-stationarity:

$$H_0 : \beta_i = 0 \ \ for \ \ i = 1 \ldots N$$

$$H_A : \beta_i < 0 \ \ for \ i = 1 \ldots N$$

The test is compared with a critical value simulated by Im et al (2003), with $\bar{t}_{12,70} = -2.0028$.[43] When compared with a 5% critical value of –1.96 with p-value of

[43] From Table 4 in Luintel (2001), this result corresponds quite closely with the findings for the EC with $\bar{t}_{11,100} = -2.128$. .

0.0249 the null is rejected and the joint hypothesis of stationarity is accepted for the panel.

The Hadri (2000) test is derived under the null of stationarity. Following the suggestion of Papell (1997) and Luintel (2001) that real exchange rates associated with developed economies are not trended, the version of the real exchange rate (y_{it}) is assumed to move around a deterministic level:

$$y_{it} = r_{it} + \varepsilon_{it} \tag{27}$$

where $t=1 \ldots T$ time periods and $i=1 \ldots N$ countries. Equation (27) assumes that the series can be decomposed into a random walk and a stationary disturbance term:

$$r_{it} = r_{it-1} + u_{it} \tag{28}$$

where, u_{it} are independently and identically distributed across i and over t with $\sigma^2_u \geq 0$. The test that the real exchange rate is stationarity, considers the following hypotheses:

$$H_0\!: \lambda=0 \qquad against \qquad H_1\!: \lambda>0,$$

where, $\lambda=\sigma^2_u/\sigma^2_\varepsilon$, and $\sigma^2_u=0$ under the null. Each equation in the panel can be presented thus:

$$y_i = X_i B_i + e_i \tag{29}$$

where, $y'_i = [y_{i1} \ldots y_{iT}]$, $e'_i = [e_{i1} \ldots e_{iT}]$ and X_i is a $Tx1$ unit (1) vector. The LM test is:

$$LM = \frac{1}{N}\sum_{t=1}^{N}\frac{1/T^2 \sum_{t=1}^{T} S_{it}^2}{\sigma_i^{*2}} \tag{30}$$

Where, σ_i^{*2} is the variance estimated from each individual sample and the partial sum of the residuals is $S_{it} = \sum_{j=1}^{t} \varepsilon_{ij}$. For comparison with the ADF test, the following non-parametric correction for serial correlation is applied to each variance term in the panel:

$$\sigma_i^{*2}(x) = \gamma_o + 2\sum_{s=1}^{T-1} \kappa(x)\gamma_s. \qquad (31)$$

Where, $\gamma_0 = \sigma_i^{*2}$, the bandwidth $x=s/l+1$, l is the lag truncation and $\gamma_s = \frac{1}{T}\sum_{t=s+1}^{T} e_{it}e_{it-s}$. A number of choices are available for the kernel $[\kappa(x)]$, each with different properties. Initially, we consider the following simple truncation:

$$\text{Truncated (T): } \kappa_T(x) = \begin{Bmatrix} 1 \text{ for x} < 1 \\ 0 \text{ otherwise} \end{Bmatrix}$$

Hadri has suggested that the Quadratic-spectral (QS) kernel might be optimal, but for comparison Bartlett (BT) and Tukey-Haning (TH) kernels are also used. Should the kernel truncation operate too early, then serial correlation might not be well modelled. The speed of decay of each kernel can be observed from **Table 2.8**. Except for the truncated kernel, the QS decays at the slowest rate.

Table 2.8 Kernel Weightings

S	Truncated	Bartlett (BT)	Tukey-Hanning (TH*)	Quadratic-spectral (QS)
1	1	0.9375	0.9904	0.9945
2	1	0.8750	0.9619	0.9780
3	1	0.8125	0.9157	0.9509
4	1	0.7500	0.8536	0.9139
5	1	0.6875	0.7778	0.8679
6	1	0.6250	0.6913	0.8139
7	1	0.5625	0.5975	0.7531
8	1	0.5000	0.5000	0.6869
9	1	0.4375	0.4025	0.6168
10	1	0.3750	0.3087	0.5443
11	1	0.3125	0.2222	0.4708
12	1	0.2500	0.1464	0.3979
13	1	0.1875	0.0843	0.3270
14	1	0.1250	0.0381	0.2592
15	1	0.0625	0.0096	0.1959

The following finite sample correction to the LM statistic is asymptotically normal:

$$Z_u = \frac{\sqrt{N}\;(LM_u - \xi_u)}{\zeta_u} \qquad (32)$$

From Hadri (2000), ξ_u=1/6 and ζ_u^2 = 1/45. Hadri shows for T∃50, that the empirical size of the test is approximately .054 and for λ in the range *[.1,4]* the test has maximum power.[44] Test results for the different kernels are summarised in **Table 2.9**.

[44] For the sample used, the test can distinguish perfectly cases for which the variance of the stochastic trend is greater than one tenth of the variance of the real exchange rate after correction for serial correlation.

Table 2.9 **Non-parametric correction to Hadri test based on alternative Kernels**

Kernel	Test Statistic
Truncated	0.935337
Bartlett	1.662121
Tukey (TH)	1.297038
Quadratic (QS)	0.925401

It should be noted that the test is one sided, which for a test at the 5% level implies a critical value of 1.645. Ordering the tests by speed of decay, the test statistics based on TH, QS and T kernels all accept the null of stationarity, while the test using the BT kernel marginally fails at the 5% level. As Hadri (2000) suggests that the test is slightly undersized the test at the 5% level is actually being undertaken at the 4.5% level, which would suggest that the null of stationarity might be accepted for the BT kernel.

Evidence seems to be being amassed to show that the real exchange rate is stationary. This study along with a number of others, notably Luintel (2001) appears to find support for the proposition when the null of non-stationarity is used. Support follows from careful analysis of the behaviour of the univariate test that after the selection of appropriate lags and mean adjustment reveal models that are well defined and seem to accept the proposition for 9 of the 12 countries analysed that real exchange rates are stationary. If one uses a broader test criterion based on the notion that incorrect rejection is more important than incorrect acceptance of the alternative then we can extend the univariate analysis to include Denmark. This compares with three without such transformations (Simpson (2002)).

The panel results are very similar to the conclusions of Luintel (2001) that on average the series investigated are I(0). However, given the reservations associated with tests of the null of stationarity by Caner and Killian (2001) and of the sensitivity of panel tests found in the simulations of Sarno and Taylor (1998), any test based on either null should be powerful and correctly sized. The test proposed by Hadri (2000) takes account of dynamic heterogeneity, corrects for both serial correlation and heteroscedasticity and has optimal size and power for the sample selected in this study. Furthermore, the test appears not to be sensitive to the underlying distribution of the data and the underlying hypothesis tested is that real exchange rates are stationary. The final test applied to a panel of demeaned real exchange rates would appear to confirm that for the eighties and nineties real exchange rates are predominantly stationary.

Both panel and univariate Dickey Fuller tests have a similar size and the univariate results broadly agree with the conclusions made from the panel that the real exchange rate when appropriately transformed is stationary. There also appears little sign of non-linearity or misspecification in the univariate time series models. The fact that both the univariate and panel analyses (based on the null of non-stationarity) come to the same conclusion would seem to obviate the concern of Sarno and Taylor (1998) that the overall panel outcome can be driven by a small sub-set of stationary series.

V. Conclusions

Purchasing Power Parity is a long established proposition, which dates from well beyond its first technical exposition by Cassel (1922). The theory of PPP is essentially the law of one price applied to a basket of equivalent goods traded internationally. The theory has come under some scrutiny of late in terms of its validity and a number of issues have arisen about the coherency of the PPP theory. Firstly short-run day-to-day trading in exchange is dominated by capital flows, suggesting that the exchange rate may deviate from PPP as long as a country's trade deficit is funded. Secondly, it is well known that the nominal exchange rate can be viewed as following a non-stationary time series process or process with a unit root (Baillie and McMahon, 1990).[45] Furthermore, the notion that the time-series process driving the exchange rate has a unit root in discrete time is quite consistent with the theoretical notion of overshooting considered by Dornbusch (1976). Indeed, as noted by Taylor and Taylor (2004), the overshooting theory helps to explain why PPP can be retained as a long-run proposition, where deviations from it in the short-run are the result of sticky prices.

This chapter has shown that it is highly important for model specification to be appropriate in undertaking tests of stationarity. Different model specifications can lead to very different test results. A particular example given explored the intercept. Using quarterly data over the period 1982 to 2005 for terms of trade and nominal exchange rates for the US, UK and Japan, it was found that the exclusion of the intercept term from the estimating equation led to results that were more supportive of stationarity (using de-meaned data for the series without the intercept). This is an important finding. Stationarity tests of PPP conducted by many researchers previously have tended to include an intercept term in the estimated ADF equation. However, it is not clear whether these intercept terms are statistically different from zero. If they are not, then this would have serious consequences for the resulting determination of stationarity.

The point to be made here is broader than exclusive focus on the intercept term. This constitutes just one example of how stationarity test procedures can be inappropriate and therefore lead to inaccurate results. Stationarity tests must be specified as correctly as is feasibly possible in order to provide results which are as accurate as possible. This involves not only rigorous diagnostic testing at the outset, but also consideration of a range of alternative econometric techniques and procedures.

[45] If a time series (y_t) follows a random walk, then $y_t = \phi_0 + \phi_1 y_{t-1} + \varepsilon_t$ has a unit root or ϕ_1=1 and the series is termed I(1) or integrated of order 1.

Panel data methods have been shown to overcome the low power problems associated with ADF tests. A further important conclusion relates to the coherence between univariate and panel estimators of stationarity. By using recursively de-meaned series in the univariate tests across 12 EU real exchange rates (1980Q1-1998Q1), the test results are not substantially different to those of the panel across the majority of the series. The pooling procedure involved with standard panel unit tests of the real exchange rate imply that PPP holds on average. However, this contrasts to the results for unadjusted data based on univariate unit root tests where only three out of twelve real exchange rates are stationary. Based on this, it would not be feasible that the panel of twelve real exchange rates could be stationary on average if only three are stationary on a univariate basis. Using appropriately transformed series, however, we find that nine of the univariate series are stationary, making the panel outcome feasible. The results that we find are reinforced by Hunter and Simpson (2001) who apply the Hadri test on the same set of recursively de-meaned real exchange rates. In the models that we implemented, as well as recursively de-meaning each of the series, careful consideration was also given to the dynamic forcing the real exchange rate. All of the models are well-defined in terms of serial correlation, and non-IID disturbances associated with excess kurtosis and heteroskedasticity are corrected using alternative robust standard errors. The models are also appropriately corrected for ARCH.

The findings in this chapter add weight to the argument supporting PPP theory and that real exchange rates are stationary in the long-run. The innovation in this chapter lies in the recognition that the misspecification in the univariate test causes the discrepancy in the panel results. The mean-adjustment transformation corrects for initial conditions so that the estimates based on the univariate and panel tests are on a level playing field. The insignificance of the intercept then means that the univariate test becomes coherent in terms of the panel test.

Given the problem associated with cross exchange rate arbitrage, we have focussed the analysis in this section on US dollar-based real exchange rates. Following Smith and Hunter (1985), triangular arbitrage means that there are strong restrictions on exchange rate equations that incorporate cross rates. Hunter and Simpson (2004) provide some univariate empirical support for this the theoretical proposition of Smith and Hunter (1985). They show that in a sample of twelve EU countries, the usual Dickey Fuller test applied to cross rate equations suffers from omitted variable bias. Thus, the use of cross rates is likely to lead to misspecification unless the parameters of the cross rates are the same as those based on dollar rates. In order to circumvent this issue, US dollar based real exchange rates were used to demonstrate how coherence can be achieved between panel and univariate unit root tests of the real exchange rate. The expansion of the Hunter and Simpson (2004) work to the multi-variate system context is the subject of research that I am currently undertaking in collaboration with John Hunter.

From a policy context, the finding in this chapter that relative prices and exchange rates converge is an important proposition for monetary policy. For example, whether real exchange rates are stationary or not has implications for the nature of exchange rate regime that might be viewed as being optimal and on whether, and by how much, governments should correct for significant exchange rate misalignments. In addition, it is reassuring from the perspective of economic theory that PPP would appear to hold given that the majority of exchange rate models and open economy macro models incorporate PPP as an equilibrium condition as regards exchange rate determination. A failure of PPP would imply that macro-models of the economy are founded upon an inappropriate exchange rate assumption and that some alternative explanation of exchange rates would be necessary.

The following chapters continue the broad theme of exploring exchange rate dynamics, using cointegration-based models. All of these models pay close attention to potential problems with the data used and a comprehensive preliminary analysis is undertaken to ensure that the appropriate adjustments are made to the models. As well as examining the time series properties of each variable in the cointegration systems (through unit root testing and assessing the dynamics of autocorrelation and partial autocorrelation functions), excessive noise is dealt with through an examination of scaled residuals from unrestricted VARs and the use of dummy variables accordingly. In addition, seasonality is controlled for. The appropriateness of the inclusion of deterministic components is also assessed across the models. Prior to moving to the cointegration analysis, across all models, the residuals from the VARs in unrestricted form are tested for misspecification, i.e. non-normality, autoregressive behaviour, ARCH behaviour, heteroskedasticity. Due to the extent of the preliminary work done in the cointegration models constructed in subsequent chapters, there is confidence that the results found are robust.

Bearing in mind this discussion, **Chapter 3** examines real exchange rate dynamics in a system context in what amounts to a multi-country test of PPP. In subsequent **Chapters 4** and **5**, the symmetry and proportionality restrictions that are implicitly imposed by testing for unit roots in real exchange rates are relaxed in reduced-form models that assess PPP in cointegrated systems comprised of nominal exchange rates, prices and nominal interest rates.

CHAPTER THREE

REAL EXCHANGE RATE DYNAMICS AND MONETARY INTEGRATION IN CRISIS-AFFECTED REGIONS

I. Introduction

This chapter continues the broad theme from **Chapter 2** as to the reasons why traditional PPP may fail to hold. In that chapter, the controversy surrounding the stationarity of real exchange rates in univariate tests compared to panel tests was addressed by ensuring that the nature of the variables was comparable across both types of test. The evidence presented above would suggest that on average PPP may even hold when the proposition is tested using a univariate analysis of the real exchange rate, whether applied to corrected single variable or to a panel. However, there are still individual cases for which the proposition does not hold.

Now, the proposition is tested using real exchange rate data for three economic regions, but estimated as part of a system for each region. In the current highly globalised world economy (both in terms of finance and trade), it is perhaps more plausible to suggest that real exchange rate stationarity may be a function of a multilateral relation. In other words, the stationarity of a country's real exchange rate may depend on a dynamic that exists between the real exchange rates of a country's trading partners, i.e. where economic interdependence is high.

The theory underlying the analysis is based on the Generalised Purchasing Power Parity (G-PPP) theory of Enders and Hurn (1994). This theory was developed *"to explain the stylized facts of real exchange rate behavior"* (p. 179) given the failure of empirical work to validate the traditional PPP theory. An application is made to three regions whose constituent countries are closely linked economically and have been affected by a financial crisis. As has already been outlined in **Chapter 2**, previous studies on PPP do not provide any consensus on the veracity of the theory, based on a range of alternative testing procedures and data specifications. G-PPP is particularly appropriate when interdependence is high and this is more appropriately examined by testing the proposition of stationarity in a multivariate context using the Johansen procedure. To this end, the extent of convergence towards G-PPP is assessed in the post-crisis period across the EMS crisis in Europe in 1992, the Latin American crisis in 1994, and the South East Asian crisis in 1997. The analysis includes an assessment of whether a long-run equilibrium relationship exists between the real exchange rates within each of the three groups and the speed of adjustment towards equilibrium. The econometric results help to indicate how regional exchange rate policy may have evolved following a major financial shock. In addition, the economic implications are set out from the perspective of monetary integration.[46] As a form of robustness check, the Perron unit root test in the presence of a structural break is implemented to confirm that the real exchange rate series in question are shown to be characterised by a unit root process and not trend stationary around a segmented deterministic component. In addition, the findings of the cointegration analysis are strengthened by a separate empirical study of the impulse responses from the VAR that is used to assess how an unexpected temporary shock to real output might affect the real exchange rates of the three crisis-affected regions.

The chapter is structured in the following way. Section II provides a brief overview of the literature and places the work in context. Section III describes the econometric methodology to be employed. Section IV provides details and a preliminary analysis of the data. Section V sets out the cointegration results. Section VI provides an economic interpretation of the results. Section VII provides the results of the Perron unit root tests. Section VIII sets out the impulse response analysis. Section IX concludes.

[46] A published version of this work appears as Beirne (2008) in the December 2008 edition of the *Journal of International and Global Economic Studies*.

II. Context and Previous Literature

The theoretical foundations underpinning G-PPP have already been described in detail in Section II of Chapter 2 (Section II.A2). The basic premise is that if two of more non-stationary real exchange rates are found to be cointegrated, this indicates that some common trend exists among them that drives the system of real exchange rates to equilibrium in the long-run. In other words, PPP holds on a multi-country level, i.e. G-PPP holds. The evidence in favour of PPP has been mixed for the recent floating exchange rate period. Nonetheless, PPP remains a cornerstone of the international finance literature, and continues to represent a benchmark against which the overvaluation or undervaluation of a currency can be measured.

This chapter focuses on a specific version of the PPP theory appropriate for countries that have a high degree of economic interdependence, namely Generalised PPP (G-PPP). According to Enders and Hurn (1994), G-PPP can provide an explanation for the non-stationarity of real exchange rates. Specifically, even though a real exchange rate may be non-stationary on a univariate basis (implying a failure of traditional PPP), there may exist a long-run tendency for the real exchange rates of a group of countries to be stationary. Thus, G-PPP permits a test of PPP that goes beyond the traditional testing applied to two-countries. Where economic interdependence is high, it makes sense intuitively that a country's bilateral exchange rate may be influenced by the exchange rates of other countries (and ultimately the economic fundamentals of other countries). Economic theory indicates that real fundamental variables such as output and productivity are primary determinants of real exchange rates. In addition, there is a substantial body of evidence to show the strong long-run link that exists between the real exchange rate and various macroeconomic fundamentals (e.g. MacDonald, 1997). Thus, it follows that if the economic fundamentals across economies are inter-related, then their real exchange rates should move together. If such co-movement in real exchange rates is not observed, then one could infer that there exists little reason for the series to follow a common trend.

As indicated, the theory provides some rationale for why traditional tests of PPP using single real exchange rates may fail due to the non-stationarity of the economic fundamentals that drive them.[47] This follows Ahn et al (2002) who note that while PPP is useful in terms of explaining fluctuations in competitiveness between countries, it is perhaps of limited use for groups of countries whose economic fundamentals are closely aligned. If a long-run equilibrium relationship exists between the fundamentals, then G-PPP may hold. In other words, even though real exchange rates may be non-stationary on an individual basis, they may be stationary when their inter-relation is considered within a system.[48]

This chapter seeks to assess G-PPP in the midst of a financial crisis. The three major crisis episodes examined are as follows: the European Monetary System crisis of September 1992, the Latin American crisis of December 1994, and the South East Asian crisis of July 1997.[49] Specifically, G-PPP is examined by assessing whether or not a long-run equilibrium relation exists between the real exchange rates of the respective regions. In order to tests this proposition, the Johansen procedure is employed to test for G-PPP in both pre-crisis and post-crisis periods. By assessing G-PPP in the midst of a crisis, with distinct analyses undertaken before and after the event, insights can be drawn on whether the crisis has caused some change in the real exchange rate dynamic within

[47] This is consistent with canonical models of exchange rate behaviour such as the Dornbusch (1976) overshooting model.
[48] This would of course imply that the real economic fundamentals across the relevant countries are also inter-linked, or that they share some common trends.
[49] The three crisis episodes were selected both on the basis of ensuring a mix of developed and developing economic regions and on the basis of ensuring that a sufficient number of observations was available in the pre- and post-crisis periods to validate the chosen methodology. Of course, the periods in question are not characterised by fully flexible exchange rate regimes. However, it is not felt that this biases the outcome of the analysis since the exchange rate regimes in place across many of the countries operate within a band of fluctuation. Therefore, in these cases, the regimes certainly could not be classified as being fully fixed and may be better described as de facto floating within a band.

economic regions and whether this has implications for regional exchange rate policy as economies seek to protect themselves for future crises.

To the knowledge of the author, this is the first analysis carried out that considers G-PPP in pre- and post-crisis situations for the countries of the EMS and for Latin American in a coherent framework. At a high level, the main issue to be explored is to examine whether a major financial shock has any implications for regional exchange rate policy. For example, is there any evidence that regional exchange rate policies become better co-ordinated following a currency crisis? Should the countries then be considered appropriate for monetary integration?

Assessing the cointegration of the real exchange rates in systems comprising the country members of the EMS, Latin America, and South East Asia before and after the crisis helps to provide an answer to these important policy issues. For example, where a cointegrating relationship is found in a system of real exchange rates, a similarity in the long-run coefficients in terms of sign would be indicative of symmetry in the response to shocks. Also, a similarity across the real exchange rates in terms of the speed of adjustment to the long-run equilibrium would also be indicative that regional exchange rate policy is co-ordinated and that monetary integration may be appropriate. This chapter explores these issues in the context of the impact made by the financial crisis.

A vast amount of empirical research has been undertaken to date with respect to PPP. This is understandable given the importance attached to PPP as a benchmark theory of exchange rate determination. The context against which this chapter is set relates to whether or not PPP is an appropriate theory for groups of countries which have close economic linkages. It may not be surprising to find that PPP is valid for such a set of countries. However, how does the dynamic change when a major structural break occurs, such as a financial crisis? Are there notable differences in the exchange rate relationships before and after the crisis? In tackling these questions, Generalised PPP (G-PPP) is employed as an objective test.[50] The original article by Enders and Hurn (1994) that first considered G-PPP found that the long-run real exchange rates of the industrialized countries did not cointegrate (i.e. G-PPP did not hold). However, when the system was augmented to include both industrialised countries and a selection of countries from the Pacific Rim, G-PPP was found to hold.[51] The interpretation of this result made by Enders and Hurn (1994) was that the extended group considered as a whole may be suitable for monetary integration.

Much of the early analysis based on univariate tests of PPP could not reject the presence of a unit root in the data generating process. Similar outcomes arose when bi-variate tests of cointegration were employed. The G-PPP theory is based on the premise that the combination of the exchange rates of a number of closely linked currencies may in combination be stationary, even though the univariate series are deemed to be non-stationary. Hence, multivariate tests of cointegration are considered the most appropriate means to detect G-PPP. If G-PPP is observed, then it is indicative of a number of common features that characterize the joint behaviour. There are a number of common stochastic trends driving each of the exchange rates that represent the inter-linkage between the economic fundamentals of the respective countries. If the observation of cointegration amongst real exchange rates is seen as an indication of the similarity in the fundamentals that drives them, then so might the overarching aims of their exchange rate policies.

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[50] Common alternative approaches to assessing monetary integration dynamics use VAR-based models due to Bayoumi and Eichengreen (1993) and Bayoumi, Eichengreen and Mauro (2000). These models focus on assessing real exchange rate dynamics amongst countries and analysing the symmetry of supply shocks. As was pointed out by Ogawa and Kawasaki (2001), a lack of symmetry in the response to supply shocks does not always mean that a currency union should not be pursued. In this sense, the G-PPP approach using the notion of cointegration is felt to provide a more comprehensive treatment of the monetary integration decision.

[51] The industrialized countries were comprised of Germany, Japan, the US, and the UK; while the Pacific Rim countries were comprised of Australia, Korea, the Philippines, Thailand, and Singapore.

Whether or not G-PPP holds, has important implications for regional exchange rate policy. If G-PPP does hold, this would suggest that some form of monetary integration would be suitable and that common exchange rate policies or initiatives may be appropriate. This would promote financial stability by reducing the exposure to the negative effects of financial contagion. It could be argued that G-PPP would be more likely to hold following a major financial shock such as a currency crisis, as countries recognize the mutual benefits to some form of monetary co-operation. Indeed, following the South East Asian crisis, a regional approach to the operation of exchange rate and monetary activities became evident for the types of reasons already cited.[52] Thus, G-PPP is important as it can be indicative of greater financial co-operation and stability. Also, there exists some evidence to suggest that greater moves towards regional monetary and exchange rate activity occur following a crisis. In this respect, it is of interest to examine whether G-PPP is more prevalent following a crisis.

While there have been a small number of studies that examine G-PPP in SE Asia, there has been very little previous work done on assessing the behaviour of systems of real exchange rates in pre- and post-crisis scenarios for the crises of Europe and Latin America. To the knowledge of the author, there have been two previous studies that examine G-PPP in the context of the euro area. Bernstein (2000) assesses the cointegration of the real exchange rates for a range of European economies over the period 1979 to 1996.[53] He separates the countries into two groups based on their level of economic development. For example, one of the groups is comprised of countries that have high inflation and a relative lack of currency credibility. Bernstein (op. cit.) then

[52] For example, the ASEAN plus Korea, China and Japan (ASEAN+3) established a regional initiative in 1998 in respect of bilateral repos and currency swaps. The ADB are also engaged in multi-lateral monitoring and surveillance of the East Asian economies (through the Regional Economic Monitoring Unit), thereby indicating the beginnings of a framework towards monetary integration and enhanced financial stability.

[53] Real exchange rates against the US dollar (and CPI) were examined across the following economies: Austria, Belgium, Denmark, Finland, France, Germany, Greece, Ireland, Italy, the Netherlands, Portugal, Spain, Sweden, and the United Kingdom.

tests for cointegration for the US dollar real exchange rates of Germany, the UK, and each of the other European countries, that is a sequence of trivariate cointegration tests. He finds that a long-run equilibrium relation exists between Germany, the UK, and a number of the smaller EU countries.[54] While these results are interesting, Bernstein (op. cit.) would appear not to have fully exploited the multivariate nature of the Johansen procedure by limiting the analysis to trivariate systems in each case. In addition, no account is taken of the turbulence associated with the exchange rate crisis in the EMS in September 1992. In addition, the author appears to overlook the lack of significance evident in a number of the long-run and short-run coefficients in many of the cointegrating relations estimated. Insignificance in the coefficients of the cointegrating vectors has implications for the inclusion of variables in the long-run relation and the appropriateness of any normalisation of the cointegrating vectors affects their identification (see Boswijk, 1996) with clear implications for the robustness and validity of the conclusions drawn. Insignificance of the loadings also has implications for the response to the long-run equilibrium. The smaller the loading coefficient for a specific equation, the slower the responsiveness of the exchange rate equation to disequilibrium and in the limit as the coefficient tends to zero then the relevant exchange rate becomes insensitive to some of the real exchange rates in the system. If all the loadings for a particular equation are zero then the variable related to this equation is weakly exogenous and the long-run can be conditioned on that variable. When this occurs the equation is driven solely by the common trends in the system and not the stationary combinations associated with cointegration.

Antonucci and Girardi (2005) examine G-PPP for the eleven countries that joined EMU in 1999. Monthly German DM real exchange rates over the period 1984 to 2002 are

[54] For example, G-PPP is found to hold between the US dollar real exchange rates of Germany, the UK, and each of Austria, Belgium, Denmark, France, Greece, Italy, Portugal and Sweden. G-PPP fails between Germany, the UK, and each of Finland, Ireland, the Netherlands, and Spain.

assessed. They find that the real exchange rates of Belgium, Luxembourg and the Netherlands are stationary. These authors find that the EMU is suitable as a monetary integration zone for the countries sampled with the exception of Spain and Ireland. Rigidities in the economic structure are provided as a rationale for the poor performance of Spain, while the poor explanation in the case of Ireland is rationalised by the stronger trade linkages that exist with the US and UK relative to other European countries.

With regard to Latin America, there has been one recent working paper that has considered G-PPP. Neves et al (2008) assess the cointegration of non-stationary real exchange rates vis-á-vis the US dollar over the period 1970 to 2006 for the Mercosur countries (Argentina, Brazil, Paraguay, Uruguay, and Venezuela). Using the Johansen procedure, they find that the null of cointegration cannot be rejected. However, their results are difficult to interpret in terms of endorsing monetary integration, because a number of the coefficients in the cointegrating equation are insignificant. They also do not analyse how the dynamic may have changed following a major structural break such as a financial crisis.

A number of studies have already tested for G-PPP in South East Asia. Notably Ogawa and Kawasaki (2003) consider the pre-crisis period, while Choudry (2005) considers both the pre-crisis and post-crisis periods. Using the US dollar as the numeraire, Ogawa and Kawasaki (2003) find cointegration for a system comprising Singapore, Malaysia, Thailand and Indonesia. However, when the numeraire is an equally weighted common basket of three major currencies (US dollar, yen, German DM), then it is found that there may be 12 potential common monetary zones. However, Choudry (2005) finds no cointegration amongst the real exchange rates of the crisis affected countries of South East Asia prior to 1997, while evidence of a long-run equilibrium relationship does appear to exist in the system for the period following the crisis. Other examples of studies where currencies were found to be interdependent include, Aggarwal and Mougoue (1993) who study Japan, Hong Kong, Malaysia, the Philippines and Singapore. Tse and Ng (1997) also found interdependence for the same group of countries, though only when the system included Korea and Taiwan. Other empirical studies using the G-PPP theory include Liang (1999) who found G-PPP to hold in a system comprised of China, Hong Kong, Japan, and the US.

This chapter builds on the previous literature on G-PPP in an exploration of real exchange rate behaviour amongst systems of interdependent currencies for two distinct periods: *before* a crisis, and, *after* a crisis. The econometric results help to indicate how regional exchange rate policy may have evolved following a major financial shock.

III. Methodology

Enders and Hurn (1994) developed the G-PPP theory in the context of empirical studies that failed to find strong evidence in support of traditional PPP. This was based on the premise that non-stationarity in the economic fundamentals causes non-stationarity in real exchange rates and thus a failure of PPP (since the non-stationary fundamentals are the drivers behind the real rates). G-PPP postulates that a sufficiently high degree of interdependence can result in real exchange rate non stationarity (Enders, 1995). Essentially, a bivariate real exchange rate may be deemed to be non-stationary, when another exchange rate that forces it is driven by a non-stationary fundamental and this would ordinarily imply a failure of traditional PPP. However, it may very well be the case that changes in the bilateral rate depend on relative prices not only in the two countries considered, but also in other countries where economic interdependence is high (e.g. trading partners). Following the notation of Enders and Hurn (1994), G-PPP can be described as follows:

$$r_{12t} = \alpha + \beta_{13}r_{13t} + \beta_{14}r_{14t} + \beta_{15}r_{15t} + \ldots.. + \beta_{1m}r_{1mt} + \varepsilon_t \qquad (1)$$

where r_{1it} is the log of the bilateral real exchange rate in period t between country i and country 1; α is the intercept term; β_{1i} are the parameters of the cointegrating vector representing the degree of co-movement of the real exchange rate and ε_t is a stationary stochastic disturbance term.

Equation (1) represents the spillover effects due to real shocks in country i that are transmitted to other economies that have a high degree of economic interdependence with country i. It should be clear to see that if all of the β_{1i} are equal to zero, then the traditional absolute PPP relationship is observed. G-PPP holds when at least one linear combination of bilateral real exchange rates that are non-stationary is observed to cointegrate. Thus, even though the bilateral real exchange rate in the traditional two-country test of PPP is non-stationary, there exists at least one linear combination of a number of non-stationary real exchange rates that is itself stationary. The implication is that the group of real exchange rates, although non-stationary, have at least one common stochastic trend.[55]

Such a common trend could be explained by the fact that output shocks have a symmetric effect on the real exchange rates (Ogawa and Kawasaki, 2001). In other words, a high degree of inter-relation between macroeconomic fundamentals that determine real exchange rates could mean that even when individual real exchange rates are non-stationary, particular groupings of these real exchange rates may in fact be stationary. Using their notation:

$$r_{i0,t} = \Sigma\beta_j r_{j0,t} + \varepsilon_{GPPP,t} \qquad (2)$$

[55] Ogawa and Kawasaki (2001) make the point that countries that have a high degree of factor mobility should have a common trend in their real exchange rates. This corresponds with Mundell (1961) who noted that factor mobility is an important Optimum Currency Area (OCA) criterion.

where the residual term, $\varepsilon_{GPPP,t}$ is stationary. The individual β parameters reflect the size and nature of the economic interdependencies within the region. Enders and Hurn (1994) suggest that when the aggregate demand functions in each country of the region similar, then the magnitude of the β's will be lower. Enders and Hurn (1994) show that the coefficients from equation (1) in the cointegrating vector are closely linked to the aggregate demand functions of a goods market-clearing relationship.[56] In a multi-equation framework the most appropriate econometric procedure to examine G-PPP would seem to be that devised by Johansen (1987, 1991).[57] As already described, the test is based on assessing whether or not the real exchange rates of the relevant countries are cointegrated.

First let us consider the following VAR(k) model:

$$z_t = A_1 z_{t-1} + \ldots.. + A_k z_{t-k} + \varepsilon_t \qquad \varepsilon_t \sim IN\,(0, \Sigma) \qquad (3)$$

where z_t is an $n \times 1$ vector of log real exchange rates and A_i represents an $n \times n$ matrix of parameters. Equation (3.3) has the following VEC representation:

$$\Delta z_t = \Gamma_1 \Delta z_{t-1} + \ldots + \Gamma_{k-1} \Delta z_{t-k+1} + \Pi z_{t-k} + \varepsilon_t \qquad (4)$$

where $\Gamma_i = -(I - A_1 - \ldots. - A_i)$, ($i = 1, \ldots., k-1$) and $\Pi = -(I - A_1 - \ldots. - A_k)$. In this notation the parameters of the system are informative of the long-run and short-run relations, i.e. an indication is provided of how the system responds in both the long-run and the short-run to changes in z_t. Information about the short-run can be derived from the estimates of Γ_i, while for the long-run this is obtained from Π. In general Π is an $n \times n$ matrix which can be decomposed as follows: $\Pi = \alpha\beta'$.

[56] See Appendix 1 for further details.
[57] Best described in Johansen (1995) or for a broader discussion of the non-iid Gaussian case see Burke and Hunter (2005).

The key condition for cointegration being that $rank(\Pi)=r$ and $0\geq r\geq n$. In the full rank case where $r=n$, then following the normalisation rule $\beta'=I$ and $\alpha=\Pi$. As a result all the series are stationary ($I(0)$) and GPPP collapses to some form of PPP. In general, $r<n$ and we have cointegration in the sense that β' is an $r\times n$ dimensioned matrix that transforms $n\sim I(1)$ series into $r\sim I(0)$ series. The matrix α is an $n\times r$ dimensioned matrix that defines the extent to which the column vectors in β enter each equation in the system. β represent the cointegrating vectors (i.e. it is the matrix of long-run coefficients such that:

$$\beta' z_{t-k}=J(L)\varepsilon_t. \tag{5}$$

Where J(L) is a matrix polynomial of possibly infinite order containing the moving average coefficients and ε are white noise disturbances. It follows by definition that $\beta' z_{t-k} \sim I(0)$. If $r=0$, then $\alpha=\beta=\Pi=0$ and we are unable to detect the existence of cointegration for the n dimensioned system that has been estimated. In the n dimensioned system the eigen values of the matrix α characterize the average speed of adjustment to equilibrium. Hence, the cointegration test amounts to assessing how many $r \leq (n-1)$ cointegration vectors exist in β (this is equivalent to testing whether Π has reduced rank).

Using the Johansen methodology (Johansen, 1995), there are two test statistics the trace test and the maximum eignevalue test. Both tests yield the number of cointegrating vectors in the system. The null hypothesis is that there are at most r cointegrating vectors. The trace test statistic computed as follows:

$$\lambda_{trace} = -T\, \Sigma_{i=r+1}{}^{n}\, ln(1- \lambda_i\,) \tag{6}$$

where λ_i are the $(n\text{-}r)$ smallest squared canonical correlations of z_{t-1} with respect to Δz_t, corrected for lagged differences and T is the sample size.

The maximum eigenvalue test is computed as follows:

$$\lambda_{max} = -T \ ln(1- \lambda_{r+1}) \qquad (7)$$

With the maximum eigenvalue test, the null hypothesis is that there are *r* cointegrating vectors against the alternative that *r+1* exist. Thus, rejection of the hypothesis implies that a maximum of *r* cointegrating vectors exist.[58]

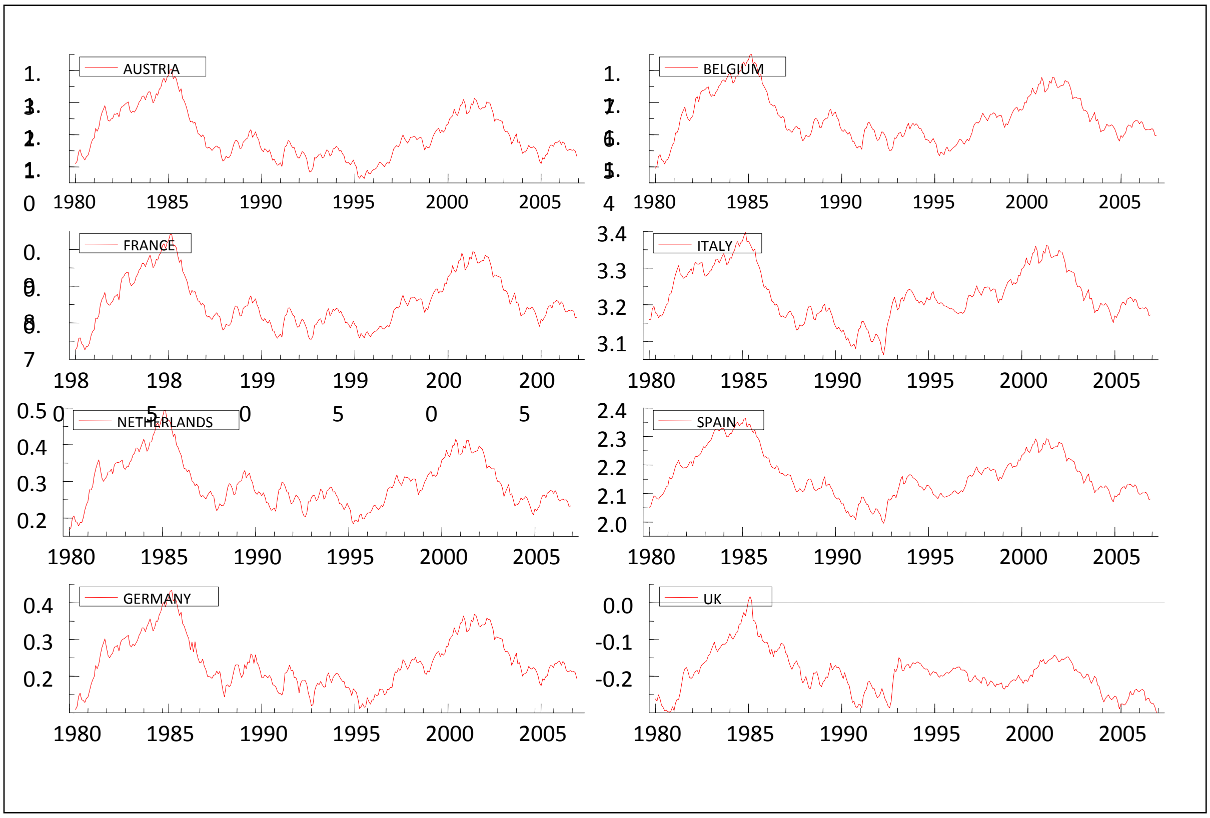
IV. Data and Preliminary Analysis

The IMF International Financial Statistics CD-ROM is the source of the data for this study.[59] The EMS countries considered are the United Kingdom, Germany, Italy, France, Spain, Belgium, Austria and the Netherlands over the period 1980 to 2006 (monthly frequency). The Latin American dataset is comprised of Mexico, Brazil, Argentina, Uruguay, Venezuela, and Peru over the 1983 to 2006 period (monthly frequency). The Asian crisis countries considered are Thailand, Indonesia, Korea, Malaysia and the Philippines and the data period is 1988 to 2006 (monthly frequency).[60] The time periods were selected so that the crisis occurs at around the mid-point.[61] With regard to the European sample, a time series was constructed for the currencies of the Eurozone members (i.e. all countries except the UK) from 1999:1 to 2006:12 using the rate at which the pre-EMU currency was converted to the Euro and the Euro/US dollar rate.

[58] Johansen (1995) shows that an optimal ordering only arises in the case of the trace test.
[59] The real exchange rate is defined as the product of the nominal bilateral exchange rate relative to the US and the ratio of US to domestic CPI. Bilateral real exchange rates have been used as opposed to effective exchange rates due to the theoretical argument that cross exchange rates are likely to be misspecified (see Smith and Hunter, 1985). In addition, the use of bilateral rates follows other studies that examine G-PPP, e.g. Enders and Hurn (1994).
[60] A monthly data frequency was used to ensure that there were sufficient observations available to validate the cointegration analysis and ensure robustness. Quarterly data would not be sufficient, particularly given the necessity for lags and dummy variables with the systems estimated.
[61] Since the analysis was undertaken in 2006, this is the end point of the sample period. The start date of the analysis for each region was selected to ensure comparability across pre- and post-crisis scenarios in terms of each scenario being comprised of a similar number of observations. Having a start date for each region prior to those used in the analysis would present comparability problems with the post-crisis analysis. Moreover, given the economic turbulence in the mid-late 1970s, an extension of the analysis back to then would most likely introduce substantial noise into the analysis.

Figures 3.1 to **3.3** show how the real exchange rates vis-á-vis the US dollar has fluctuated across the currencies of each of the three regions under consideration.[62]

Figure 3.1 **(Log) Real Exchange Rates – EMS Crisis Countries, 1980-2006**



A notable feature of the movement in the EMS currencies is the real appreciation against the dollar over the 1985 to 1987 period. This was due to concerted action by the then G5 (France, Germany, Japan, the US, and the UK) to intervene in the currency markets to devalue the US dollar.[63] The fall in the dollar affected the real exchange rates of the selected EMS countries above in a similar fashion, as can be seen in **Figure 3.1**. The effects of the EMS crisis in September 1992 are also evident, as all of the currencies experienced a real devaluation against the dollar. It is notable that across all currencies, the pattern of fluctuation has been remarkably similar since 1980.

[62] For each series, the log of the real exchange rate per unit of US dollar is provided using the standard international currency abbreviations. These abbreviations are set out in full in the Appendix.

[63] The co-called 'Plaza Accord' involved a $10 billion sell-off of US dollars by the G5 central banks in order to reduce the US current account deficit and to stimulate economic growth in the US. This action, in conjunction with currency market speculation caused a dramatic fall in the dollar in the two years following the signing of the agreement on September 22nd 1985. The continued fall in the dollar was halted in 1987 with the Louvre accord.

The behaviour of the Latin American countries can be observed in **Figure 3.2** below.

Figure 3.2 **(Log) Real Exchange Rates – Latin American Crisis Countries, 1983-2006**

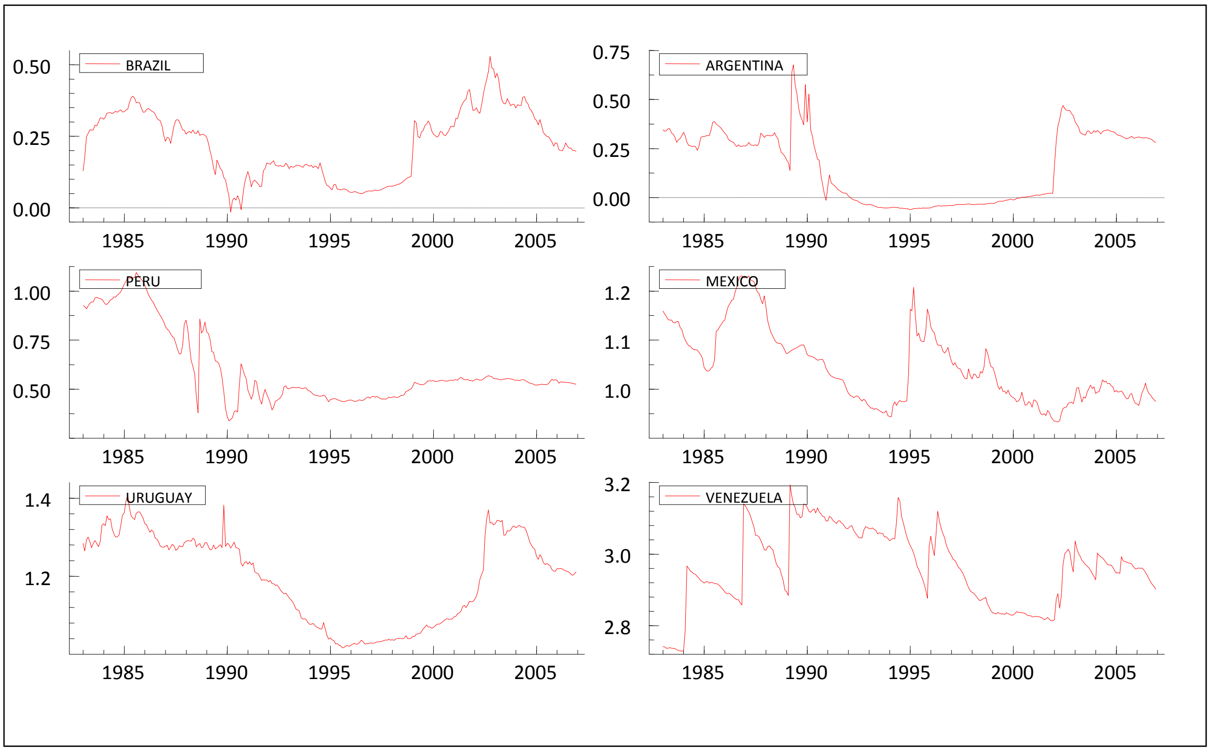


Figure 3.2 shows a degree of high volatility in the Latin American currencies prior to 1991 in particular. There are a variety of reasons for this. For example, in Brazil exchange rate policy was typically used to control inflation. This was not always successful, however, and when inflation became uncontrollable to the extent that the domestic currency was demonetized, the Brazilian authorities replaced the currency.[64] Chronic inflation also affected other Latin American countries in the late 1980s and early 1990s and similar measures as those adopted by Brazil were employed.[65] It is also important to bear in mind the exchange rate regime in place over the period under consideration. In the 1980s and early 1990s, all of the Latin American economies in the sample (with the exception of Uruguay) employed intermediate regimes such as soft pegs,

[64] Since 1986, Brazil has had five different currencies: the Cruzado (1986-1989), the Novo Cruzado (1989-1990), the Cruzeiro (1990-1993), the Cruzeiro Real (1993-1994) and the Real which was introduced in 1994 and remains in place today.

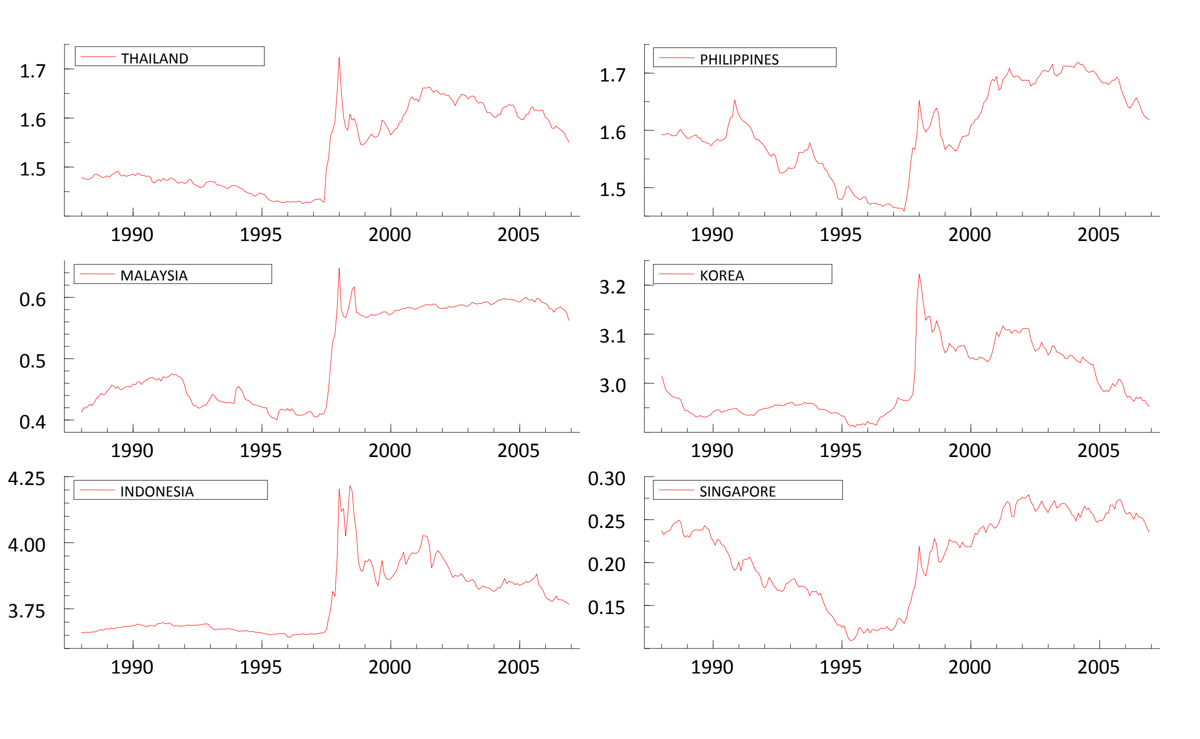
[65] Argentina replaced the Peso temporarily between June 1985 and January 1992 with the Austral. Also Peru replaced its currency in 1990 to combat hyperinflation.

crawling pegs and crawling bands relative to the dollar. They have now all shifted to floating rate regimes.[66] Thus far, the previous studies undertaken to determine whether a monetary union was appropriate for Latin America tend not to support this proposition (e.g. Bayoumi and Eichengreen, 1994; Hallwood et al, 2004; Foresti, 2007; Neves et al, 2008). The commonly cited reasons for this negativity include a low level of trade integration, asymmetric co-movements to shocks, differences in speed of adjustment and the relative size of shocks. The consensus appears to be that more policy co-ordination is necessary before any economic integration in Latin America can occur.

In **Figure 3.3** below the behaviour of the South East Asian countries currencies can be observed around the crisis in 1997.

[66] Intermediate exchange rate regimes tended to be more prone to crises due to the vulnerability to speculative attack that accompanied a lack of full commitment to the peg. In addition, a shift away from intermediate regimes was also related to the global drop in inflation.

Figure 3.3 (Log) Real Exchange Rates – Asian Crisis Countries, 1988-2006



Regarding the South East Asian economies, there is clear evidence, prior to the crisis events of July 1997 of a de facto peg against the US dollar and after the crisis of a sharp real devaluation of all of the currencies in question.[67] After July 1997, these quasi-fixed arrangements were abandoned following a failure to defend against speculative attack by raising interest rates and reducing reserves. The economies generally moved to fully floating or similar regimes after the crisis with the exception of Malaysia who maintained a fully fixed rate against the US dollar. The countries selected were considered to be those most affected by the crisis and they are also closely linked in terms of trade relations.

As mentioned earlier, in order to test for a change in the exchange rate relationship for the crisis-affected countries, the Johansen procedure is employed.[68] In order to proceed with

[67] See Reinhart and Rogoff (2004) for more detail on the nature of the exchange rate regimes evident for these and other countries.
[68] Subsequent Ramsey RESET tests across all regions and sub-periods are performed on the ECMs with squared terms. The results for the pre- and post- samples across all regions fail to reject the null hypothesis of linearity (see Table

the Johansen technique, a necessary condition is that all of the variables in the system are integrated of the same order. To determine this unit root tests are applied to the real exchange rates of the crisis countries across the three regions. Of course, the nature of the study means that three sets of tests must be carried out for each region: for the total period, pre-crisis period, and post-crisis period. Conventional Augmented Dickey-Fuller tests show that all of the real exchange rates are non-stationary in levels across the three periods and would appear to have been generated by I(1) processes.[69] Selection of the optimal model is determined in the course of applying the Johansen test, via the Pantula principle (Johansen, 1992). The finding that all of the real exchange rate series are I(1) is consistent with the G-PPP theory, whereby interdependence in economic fundamentals is reflected in the behaviour of the real exchange rates.

V. Cointegration Results

This stage involves testing for cointegration among the real exchange rates using the Johansen procedure. The appropriate lag length for the EMS model, the Latin American model and the Asian model is 6, 6 and 12 respectively.[70] In selecting the most appropriate model as regards the VAR deterministic components, the Pantula principle (Johansen, 1992) is applied whereby three specifications are estimated and assessed. The Pantula principle selects both the correct specification of the deterministic components as well as the order of the cointegration rank.[71] These results are set out below.[72]

A3.16 in Appendix I). This vindicates the choice of methodology chosen, allaying concerns regarding functional form and possible non-linearity.

[69] These results are set out in the Appendix I.

[70] The lag length selection was based on analysis of an unrestricted VAR yielding Gaussian error terms and the lowest AIC (the results are also robust across two alternative information criteria – the Schwarz Bayesian IC and the Modified AIC). The resulting VAR estimated for the Asian group does exhibit some signs of non-normality. However, due to Gonzalo (1994), the Johansen procedure remains robust in the presence of non-normality. This is shown to be due to excess kurtosis rather than skewness. Full misspecification test results are shown in Tables A3.11-A3.14 in Appendix I. In addition, the standardized residuals of the VAR in unrestricted form revealed a number of outliers that were dealt with using intervention dummies (see Appendix I).

[71] The Pantula principle involves estimating the three alternative models (i.e. no intercept or trend, intercept and no trend, intercept and trend) and moving from the most restrictive to the least restrictive model. The trace test statistic or

Table 3.1 Pantula Principle Test Results for Full Sample

R	n-r	Model 2	Model 3	Model 4
EMS (k=6)				
0	3	106.8*	104.6*	116.8
1	2	69.2	67.0	75.0
2	1	35.4	33.6	41.6
Latin America (k=6)				
0	3	106.5*	101.7*	147.1*
1	2	57.1	52.5	64.4
2	1	30.4	35.8	31.2
Asia (k=12)				
0	3	92.0*	86.9*	109.2*
1	2	52.1	48.4	40.4
2	1	29.8	26.1	32.4

Notes: * denotes rejection of the null hypothesis of no cointegration. Model 2 assumes no intercept or trend in either the cointegrating equation (CE) or the VAR; Model 3 allows for an intercept but no trend in the CE or VAR; and Model 4 allows for an intercept and trend in the CE and VAR.

The Pantula methodology would suggest proceeding with Model 3 for the EMS, Model 4 for Latin America, and Model 4 for the Asian countries. The cointegration rank (*r*) is determined by assessing the trace and maximum eigenvalue test statistics.

the maximum eigenvalue test statistic is compared to the critical value in each case, and the most appropriate model is deemed to be the one where the null hypothesis is not rejected for the first time.

[72] It is pertinent to notice at this point that these tests are not similar, Nielsen and Rahbek (2000) suggest the test without intercept is not used and that a similar testing strategy applies the test with restricted intercept and restricted time trend for each of the cases above as alternative similar test procedure. As a check, therefore, cointegration rank tests were performed for each of the three regions where the deterministic components are comprised of both a restricted intercept and trend. A rank of one is identified in all three cases.

Table 3.2 LR Trace and Max tests: Full Sample

H_0: rank=p	λ_{trace}	95%	λ_{max}	95%
EMS (Sample 1980M01: 2006M12) – Model 3; 6 lags				
p=0	104.6*	0.011	37.6	0.093
p≤1	67.0	0.082	33.4	0.056
p≤2	33.6	0.524	14.3	0.804
p≤3	19.3	0.469	10.5	0.696
Latin America (Sample: 1983M01: 2006M12) – Model 4; 6 lags				
p=0	147.1*	0.000	119.2*	0.000
p≤1	64.4	0.074	66.7	0.082
p≤2	31.2	0.077	18.0	0.244
p≤3	17.8	0.577	11.8	0.567
Asia (Sample: 1988M01: 2006M12) – Model 4; 12 lags				
p=0	109.2*	0.001	38.8*	0.044
p≤1	40.4	0.058	28.1	0.144
p≤2	32.4	0.077	19.3	0.286
p≤3	23.1	0.108	14.1	0.245

Note: * denotes rejection of the null hypothesis of no cointegration. Model selection was based on the Pantula principle as in the case of the full sample results, and lag selection was based on the lowest AIC in conjunction with the observation of Gaussian errors.

Based on the more robust trace statistics[73], the analysis indicates the presence of one cointegrating vector for all regions. Therefore, over the full sample period for all regions, there is evidence to suggest a long-run stationary relationship in the real exchange rates between the countries within their respective regional groups. This is indicative of a close interdependence in the long-run between the real exchange rates analysed. In this sense, the results are supportive of evidence in favour of G-PPP over the entire period. This finding confirms the results of Choudry (2005), who made a similar finding in relation to the South East Asian economies. Analysis of the extent of cointegration in the relevant pre-crisis periods is provided in **Table 3.3** below.

[73] See Gonzalo (1994) and Cheung and Lai (1993) for details on the robustness of the trace test compared to the maximum eigenvalue test.

Table 3.3 **LR Trace and Max tests: Pre-Crisis Sample**

H_0: rank=p	λ_{trace}	95%	λ_{max}	95%
EMS (Sample 1980M01: 1992M08) – Model 4; 3 lags				
p=0	136.2*	0.000	63.9*	0.000
p≤1	62.3	0.227	30.7	0.127
p≤2	51.7	0.343	20.1	0.643
p≤3	31.5	0.414	17.4	0.422
Latin America (Sample 1983M01: 1994M11) – Model 2; 2 lags				
p=0	100.8*	0.000	67.5*	0.000
p≤1	38.3	0.070	32.2	0.055
p≤2	31.1	0.081	27.8	0.062
p≤3	23.2	0.110	21.4	0.067
Asia (Sample: 1988M01: 1997M06) – Model 3; 3 lags				
p=0	66.6	0.088	33.4	0.057
p≤1	33.2	0.547	19.9	0.347
p≤2	13.3	0.897	10.5	0.699
p≤3	2.8	0.976	2.8	0.961

Notes: * denotes rejection of the null hypothesis of no cointegration.

Based on the trace statistics, the analysis indicates that the null hypothesis of no cointegration cannot be rejected for the Asian pre-crisis period. With the Asian pre-crisis sample, the optimal lag structure now becomes 3 and the most appropriate model includes an intercept but no trend in the CE and VAR (i.e. Model 3). On the other hand, evidence of cointegration can be seen in the case of the EMS and Latin American countries during the pre-crisis period.

Cointegration results for the post-crisis period are set out in **Table 3.4**. The post-crisis results indicate evidence for one cointegrating relationship for each region. In the pre-crisis period, there was no long-run stationary equilibrium relationship among the real exchange rates in Asia. There is a dramatic shift, however, in the post-crisis period, suggesting some form of co-ordinated policy action to improve stability. All of the cointegrating relationships identified would appear to be stable, as shown from the 60-month recursive estimates of the eigen values for each vector, because they seem to be generally constant (see **Figure A3.1** in Appendix I).

Table 3.4 LR Trace and Max tests: Post-Crisis Sample

H_0: rank=p	λ_{trace}	95%	λ_{max}	95%
EMS (Sample 1993M01: 2006M12) – Model 4; 3 lags				
p=0	154.5*	0.000	50.2*	0.011
p≤1	49.4	0.103	38.3	0.112
p≤2	40.2	0.157	32.1	0.337
p≤3	36.1	0.203	25.8	0.630
Latin America (Sample 1995M09: 2006M12) – Model 4; 2 lags				
p=0	109.3*	0.000	59.9*	0.001
p≤1	52.4	0.051	31.4	0.067
p≤2	38.7	0.124	21.3	0.163
p≤3	28.7	0.225	11.3	0.277
Asia (Sample: 1998M06: 2006M12) – Model 4; 5 lags				
p=0	136.5*	0.000	45.3*	0.007
p≤1	51.2	0.057	24.0	0.054
p≤2	31.1	0.116	21.6	0.064
p≤3	16.6	0.255	11.3	0.112

Notes: * denotes rejection of the null hypothesis of no cointegration.

Unlike the case of SE Asia, the European and Latin American countries appear to have always exhibited G-PPP. This finding, however, is not sufficient to provide answers to the policy questions to be addressed, namely has regional exchange rate policy become more co-ordinated following the crisis, and what are the implications for monetary integration? In order to provide answers to these questions, a more detailed analysis of the cointegrating equations is required.

VI. Interpretation of the Results

In interpreting the results, the focus is placed on comparing the cointegrating relationships in the pre- and post-crisis scenarios to assess whether regional exchange rate policy co-ordination and the scope for monetary integration have become more pronounced following the crisis. Firstly, the long-run cointegration equation is examined. In this case, monetary integration would appear more appropriate when the long-run coefficients in the systems have the same sign, i.e. the variables move in the same direction. The magnitude of these coefficients is also important in assessing the implications for monetary integration. As noted by Enders and Hurn (1994) in their original model, very large coefficients can be indicative of a lack of similarity in the demand parameters across the countries. Secondly, the speed of adjustment coefficients is assessed to identify how quickly the real exchange rates move towards the long-run equilibrium. Clearly, similar speeds of adjustment would be indicative of co-ordination in exchange rate policy.

VI.A Long-Run Elasticity

In order to interpret the cointegration results, a necessary first step is to normalise the cointegrating vector on one of the dollar real exchange rates.[74] For Europe the normalisation is on USDATS (i.e. Austrian schilling per US $); for the Latin American sample, the vector is normalised on USDARS (i.e. Argentine peso per US $); and for the South East Asian sample, normalisation is undertaken according to USDIDR (i.e. Indonesian rupiah per US $). All of the equations are set out in **Table 3.5**. Prior to

[74] The results are not dependent upon the normalization base. Moreover, in order to check the appropriateness of the normalization, long-run exclusion tests were carried out for the three scenarios (i.e. full period, pre-crisis, and post-crisis). In all cases, the long-run exclusion restriction was not accepted, confirming firstly that all of the variables in the systems are relevant for the long-run equilibrium. This, in conjunction with the absence of weak exogeneity in the normalised variables (excluding possible weak exogeneity in the post-crisis case for the Latin American system normalised on Argentina), ensures confidence in the normalization scheme. The long-run exclusion tests are provided in Table A3.17 in the Appendix.

explaining the economic meaning of the cointegrating vectors based on the normalised equations, it is necessary firstly to consider the statistical significance of the results.

For the European full sample, only the coefficients of Italy, the Netherlands and Spain are significant (this is relative to a normalisation). There is some variability across the European countries in relation to size effects, although all of the coefficients are less than unity. This is encouraging in terms of the economic relationship between the countries in terms of their real exchange rates. The results show that a 0.90% decrease in the US $ Austrian Schilling exchange rate (USDATS) increases the US $ Netherlands exchange rate (USDNLG) by 1%; and a 0.32% increase in USDATS is associated with a 1% increase in the US $ Italian Lira (USDITL). For the purpose of comparison, the pre- and post-crisis results are all significant for all countries excluding Spain and France (although France is significant in the post-crisis cohort). The coefficient is lower in the post-crisis period (compared to the pre-crisis period) in the cases of Belgium, the Netherlands and Italy. A further feature of the European case is that there appears to be some asymmetry in the response to shocks with regard to the UK and Italy in the post-crisis period (which has a positively signed coefficient, while for the other countries the sign is negative). Concerns regarding this, however, are allayed by the very low magnitude of the coefficients (in the range -0.283 to 0.133). Thus, while there is some asymmetry, the extremely narrow range within which the coefficients lies means that the European case can still be considered consistent with monetary integration in the post-crisis case.[75]

The Latin American full sample results are all statistically significant in the cases of Peru and Venezuela. However, the sizes of the coefficients are somewhat mixed across the

[75] Indeed, differences in consumption patterns and inflation mean that the movement of the real exchange rates can be volatile and not always symmetric in response to shocks. Where this is the case, as long as the fluctuations take place within narrow bands, then monetary integration can still be considered appropriate (e.g. Gros and Lane, 1992).

real exchange rates, with Peru showing a value greater than one, while Venezuela has a value less than one. These coefficients can be interpreted as follows: a 0.76% rise in USDARP increases USDVEB by 1%; and a 2.23% decrease in USDPEN increases USDARS by 1%. The pre-crisis results are significant in the cases of Brazil, Mexico, Peru and Venezuela, while the post-crisis results are significant for Brazil, Peru, Uruguay and Venezuela. There are greater size effects evident with the post-crisis sample with the majority of the significant coefficients greater than unity. Brazil, Peru, and Uruguay have notably higher coefficients post-crisis compared to pre-crisis. For the post-crisis results, the only significant coefficient below one is that of Venezuela. Considering Latin America as a whole, it is evident that the results are not coherent in terms of the sign of the coefficients, thus indicating an asymmetric exchange rate adjustment process across the countries. Also, there are some signs of large coefficients, indicating a dis-similarity in the demand parameters across countries, although this becomes much more muted in the post-crisis period (with the exception of Uruguay). Overall, the differences evident in the beta coefficients, in terms of magnitude and sign, are not supportive of a proposal for the monetary integration of the region.[76]

The full sample results for the South East Asian currencies are significant only in the cases of Malaysia, the Philippines, and Singapore – a 1% increase in USDMYR is associated with a 1.22% decrease in USDIDR; while a 1% rise in USDPHP is linked with a 0.72% rise in USDIDR; and a 1% increase in USDSGD is linked with a 1.77% rise in USDIDR. There were no cointegrating vectors identified for the pre-crisis sub-sample and thus there are no normalised equations reported for this sample. For the post-crisis period, all of the coefficients are significant, and greater than unity. Korea, Malaysia and

[76] This finding is reflective of previous studies done on assessing monetary integration in Latin America, e.g. Foresti (2007), Neves, Stocco and da Silva (2007), Hallwood, Marsh and Schiebe (2004), Arora (1999), and Bayoumi and Eichengreen (1994). The main reason cited in these studies for the lack of support for monetary union are the low levels of trade integration within Latin America, the asymmetric adjustments to shocks, and the differences in speeds of adjustment and size of shocks. Generally, it is felt that greater policy co-ordination is necessary before any form of economic integration can proceed.

Singapore are positively signed, while the Philippines and Thailand are negatively signed, indicating that while there is no symmetry in the exchange rate adjustment process across all the countries considered together, there is some symmetry within sub-groups. The beta coefficients are large in all countries, notably in Korea and Malaysia, indicating that the underlying economic fundamentals may be different. The symmetry as regards the exchange rate response to shocks suggests that there has been some form of co-ordinated approach to regional exchange rate policy for sub-regions in South East Asia following the crisis. Monetary union may be suitable within sub-groupings of the countries, but certainly not all and this would seem to be caused by the differences in the beta coefficients.

VI.B Speed of Short-Run Adjustment

As there is at most a single cointegrating vector, then the results in **Table 3.6** provide estimates of the speed of adjustment of each of the real exchange rates under consideration towards the long-run equilibrium. These coefficients can be interpreted as a measure of how quickly each of the real exchange rates converges to G-PPP. Analysis of the significance of each of the coefficients reveals information about the weak exogeneity of each of the variables in turn for β (Johansen, 1992, Hunter, 1992). As outlined by Harris (1995), a variable is deemed to be weakly exogenous if its speed of adjustment coefficient is not statistically different from zero. Where weak exogeneity is observed in one of the series, this implies that the cointegrating relationship does not help to explain that short-run equation. Clearly, this is an important issue to consider and in many respects it is perhaps unusual that the Johansen procedure tests for weak exogeneity *after* estimating the long-run coefficients. However, except for the special case of a single cointegrating vector the identification the long-run is not well defined without further restriction and even in this case that is sensitive to the normalisation, which will not be appropriate when the variable might be excluded from the long-run. Hence, although the short-run equation associated with the weakly exogenous variable is not pertinent to the estimation of the cointegrating vector, the long-run equation might be best explained by the weakly exogenous variables that are relevant for the estimation, identification and interpretation of the long-run coefficients.

For the countries of the EMS, there is much less variability in the speed of adjustment to the long-run equilibrium in the post-crisis case. Also, the coefficients are very similar in size. In the post-crisis period, the only sign of weak exogeneity relates to the UK. The dynamic short-run exchange rate equation does not respond to G-PPP and as a result the long-run can be conditioned on the UK real exchange rate. In essence, the UK exchange rate is predominantly driven by the random walk or its own stochastic trend. This may cast doubt on the suitability of the UK for a monetary integration within euro area as currently defined. For the Latin American currencies, the majority of the short-run coefficients are significant for the full sample period and the pre-crisis sub-sample. The largest coefficient is that of Peru (0.183) in the pre-crisis period, implying that USDPEP moves at the rate of 18.3% per month towards the long-run equilibrium. The Uruguayan real exchange rate vis-á-vis the dollar converges about half as fast in the same period (8.1% per month). It is clear that the speed of adjustment appears to be faster in the post-crisis period, although the results are somewhat tainted by the observed weak exogeneity in many of the real exchange rates.

Weak exogeneity is not observed in any of the currencies for the Asian post-crisis period. Hence, the long-run is defined by the inter-action of the real exchange rates of all of the Asian economies The post-crisis period in Asia reveals the largest speed of adjustment in the case of Indonesia (14.2% per month), while Thailand, Korea, Malaysia and the Philippines have broadly similar convergence rates of between 1.8% and 3.5% per month.

This similarity is indicative of a synchronized response to shocks to the real exchange rate.

Overall, the results from the analysis can be interpreted in terms of the policy implications that they infer. The evidence for the European sample of countries suggests a similarity exists in the long-run demand parameters. For example, in the post-crisis period, all of the coefficients are low and similarly signed except in the cases of the UK and Italy. This implies that the reaction of the real exchange rates within the systems is predominantly symmetric. Moreover, the very low level of the coefficients in the case of the UK and Italy are such that the asymmetry evident remains consistent with monetary integration for the system as a whole. Given differing business cycle, trade, and consumption patterns across the countries in question, it is not feasible that all of the coefficients would be identical, or indeed moving in the same direction. This may help to explain, for example, the slight asymmetric effects in the cases of the UK and Italy. However, these effects remain consistent with monetary integration given the very low size effect. For the Latin American sample, this is not the case however. The coefficients are large and differ quite substantially across countries, and there is a lack of similarity in the direction of the effect. This implies that monetary integration is not feasible. This is explained by a low level of trade integration. For the South East Asian economies, while the long-run coefficients are large, after the crisis, the direction of the responses of the real exchange rates to shocks is the same for all of the countries. Therefore, while there are some underlying differences in demand parameters and economic fundamentals, the similar direction of the response to shocks could indicate that monetary integration may be feasible at some point in the future, or indeed for a sub-sample of the countries. A more detailed interpretation on individual real exchange rate effects would require a more comprehensive analysis of business cycle, trade and consumption patterns within regions.

Table 3.5 Normalised Cointegrating Equations and Long-Run Coefficients

Europe								
	Austria	*Germany*	*UK*	*Belgium*	*France*	*Italy*	*Netherlands*	*Spain*
Full Sample	1.000	0.277*** (0.156)	0.723 (0.577)	0.199 (0.196)	-0.289 (0.203)	0.320* (0.038)	-0.900* (0.170)	-0.267* (0.040)
Pre-Crisis	1.000	0.655* (0.237)	-0.343* (0.145)	0.962* (0.225)	-0.268 (0.221)	-0.754* (0.133)	-0.934* (0.100)	-0.100 (0.069)
Post-Crisis	1.000	-0.110* (0.021)	0.128* (0.042)	-0.273** (0.116)	-0.283* (0.119)	0.133* (0.036)	-0.212*** (0.116)	-0.176* (0.044)

Latin America						
	Argentina	*Brazil*	*Mexico*	*Peru*	*Uruguay*	*Venezuela*
Full Sample	1.000	0.383 (0.372)	-0.261 (0.222)	-2.234* (0.220)	0.196 (0.220)	0.755* (0.280)
Pre-Crisis	1.000	-1.163* (0.408)	-3.188* (0.690)	-1.112* (0.150)	-1.528 (1.038)	3.633* (0.965)
Post-Crisis	1.000	1.481* (0.248)	0.021 (0.060)	-1.373* (0.208)	-6.566* (0.771)	0.262* (0.045)
Asia						
	Indonesia	*Korea*	*Malaysia*	*Philippines*	*Thailand*	*Singapore*
Full Sample	1.000	-0.637 (0.419)	-1.223* (0.304)	0.722** (0.316)	-0.440 (0.623)	1.773* (0.461)
Pre-Crisis	n/a	n/a	n/a	n/a	n/a	n/a
Post-Crisis	1.000	4.062* (1.071)	6.779* (2.958)	1.721** (0.946)	2.656* (0.990)	0.942* (0.352)

Notes: Standard errors are reported in parentheses. There are no results for the pre-crisis period of Asia due to the failure to reject the null hypothesis of no cointegration for this region during the pre-crisis time period. *, **, and *** indicates statistical significance at the 1%, 5%, and 10% levels respectively.

Table 3.6 Adjustment Coefficients (vis-á-vis the US dollar)

Europe								
	Austrian schilling	*German DM*	*UK pound*	*Belgian franc*	*French franc*	*Italian lira*	*Dutch guilder*	*Spanish peseta*
Full Sample	-0.921* (0.372)	-0.754* (0.411)	0.734 (0.556)	-0.380* (0.172)	-0.426* (0.145)	-0.870* (0.330)	-0.244* (0.115)	-0.450* (0.260)
Pre-Crisis	-0.965* (0.553)	-0.673* (0.379)	0.876* (0.466)	-0.465 (0.559)	-0.534* (0.225)	0.016 (0.504)	-0.282* (0.165)	0.400 (0.520)
Post-Crisis	-0.912* (0.388)	0.274* (0.117)	0.598 (0.483)	0.178* (0.072)	0.399* (0.131)	-0.305* (0.132)	-0.417* (0.206)	0.298* (0.134)

Latin America						
	Argentine peso	*Brazilian real*	*Mexican peso*	*Peruvian nuevo sol*	*Uruguayan peso*	*Venezuelan bolivar*
Full Sample	-0.286* (0.036)	0.001 (0.016)	-0.188* (0.024)	0.181* (0.073)	-0.039* (0.017)	0.085* (0.036)
Pre-Crisis	-0.160* (0.043)	0.077 (0.058)	0.040*** (0.024)	0.183* (0.075)	0.081* (0.032)	-0.053 (0.052)
Post-Crisis	-0.027 (0.034)	-0.047 (0.052)	0.386 (0.416)	-0.145 (0.138)	0.300* (0.046)	-0.197 (0.737)

Asia						
	Indonesian rupiah	*Korean won*	*Malaysian ringgit*	*Philippine peso*	*Thai baht*	*Singaporean dollar*
Full Sample	-0.319* (0.065)	0.001 (0.034)	-0.065* (0.025)	-0.065* (0.030)	-0.050 (0.036)	-0.078* (0.017)
Pre-Crisis	n/a	n/a	n/a	n/a	n/a	n/a
Post-Crisis	-0.142* (0.030)	-0.026*** (0.014)	-0.018* (0.005)	-0.035* (0.010)	-0.032* (0.011)	-0.029* (0.009)

Notes: Standard errors are reported in parentheses. There are no results for the pre-crisis period of Asia due to the failure to reject the null hypothesis of no cointegration for this region during the pre-crisis time period. *, **, and *** indicates statistical significance at the 1%, 5%, and 10% levels respectively.

VII Robustness Check 1: Stationarity in the Presence of a Structural Break

An alternative initial approach to the preceding preliminary analysis would be to test for a unit root in the series in the presence of a structural break. This could potentially change the deterministic nature of the underlying VAR model. For example, it could be the case that the real exchange rate series are not characterized by a unit root process, and instead that they are trend stationary around a segmented deterministic component.

In order to test for a unit root in the presence of a structural break such as a financial crisis, it has been conjectured that the standard unit root tests procedures may not be appropriate due to a bias towards non-rejection of a unit root (Harris, 1995).[77] With this in mind, the Perron (1989) procedure can be employed to test for a unit root where a break in the time series is observed.[78] In the case of the EU this would be supported by the finding in **Chapter 2** that these exchange rates are predominantly stationary. The Perron (1989) test as with the Dickey-Fuller test is based on the null of a unit root. However, this is tested using the following model:

$$z_t=a_0+a_2t+\mu_2D_L+z\_hat_t \tag{8}$$

where D_L is a level dummy that equals 1 for all $t > \tau$ (i.e. the break date) and zero otherwise.

Specifically, the null is represented as a unit root process with a one-time jump in the level at time period $t = \tau+1$; whereas under the alternative hypothesis, y_t is generated via

[77] With regard to the crises under consideration in this study, the crises dates are commonly understood - EMS crisis: September 1992; Mexican crisis: December 1994; Asian crisis: July 1997. This has been confirmed formally for the series under consideration using recursive estimates of the residuals generated by OLS estimates (this approach to break-point identification was also taken by Goldberg and Frydman, 2001). One could test for a unit root when the break-point in unknown using the approach of Perron and Vogelsang (1992), Zivot and Andrews (2002), Perron (1997) or Vogelsang and Perron (1998).

[78] Such an approach preserves degrees of freedom compared to splitting the sample and running separate regressions.

a trend stationary process with a one-time jump in the intercept (Enders, 1995). Model (7) is estimated for each of the real exchange rates and the residuals are saved. The second stage of the Perron approach involves regressing these saved residuals on their lagged values. More than one lag may need to be applied in cases where autocorrelation is present. This augmented form of the residual regression is of the form:

$$z\_hat_t = a_1 z\_hat_{t-1} + \Sigma a_i \Delta z\_hat_{t-i} + \varepsilon_t \quad (9)$$

As described by Perron (1989), α_1 should equal 1 under the null of non-stationarity. This procedure is carried out in the following paragraphs. As an initial step, however, two commonly used tests are employed to confirm the occurrence of the structural break for each of the three crisis periods. The first test is the Chow (1960) test which uses the F-distribution under the null hypothesis of no break to examine whether a structural break has occurred, when there is a prior in relation to the date of the break. To ensure a degree of robustness in the analysis, a recursive estimator is employed to test for the break in the absence of *a priori* information. To this end, the Cusum and Cusum of Squares tests are performed. All of these tests confirm that the commonly known break-points for each region represent a break point in each of the real exchange rate series under consideration. With regard to the Chow test, the computed F-statistic is greater than the critical value in all cases at or below the 1% level. For the Cusum test, which is based on the cumulative sum of the residuals, the test statistic falls outside the critical boundary at the point where the crisis is initiated. The Cusum of Squares test is based on the squares of the recursive residuals where the statistic lies between 0 and 1. No structural break would be indicated by a random dispersion of the statistic close to zero, the observation of a break arises when the statistic approaches 1. In the case of the Asian crisis countries, the Cusum of Squares test statistic is close to zero up until July 1997, at which point it rises dramatically towards the upper limit of 1. Similar analyses carried out on the European and Latin American groups of countries indicate break dates at September 1992 and December 1994 respectively.

As a result, break-dates have been confirmed to be similar to the dates commonly perceived in the literature for each of the three regions. Of course, this discovery is not ground-breaking. Nonetheless, it is a necessary starting point to the analysis and enables us to proceed with testing for a unit root in the presence of a structural break where the date of the break is known. As is common knowledge, the presence of a structural break in a time series system can lead to a breakdown in economic relationships, model misspecification, and inaccurate statistical inference. It is crucial, therefore, to take these issues on board in examining the PPP relationship between regional economies in the midst of major structural breaks such as those of September 1992, December 1994 and July 1997.

Table 3.7 Perron Unit Root Test with Structural Break

	a_1	Standard Error	Ho: a_1=1 (*t*-statistic)
Europe			
Austria	0.981	0.010	-1.900
Germany	0.977	0.011	-2.091
UK	0.973	0.012	-2.250
Belgium	0.979	0.009	-2.333
France	0.981	0.010	-1.900
Italy	0.981	0.010	-1.900
Netherlands	0.976	0.010	-2.400
Spain	0.984	0.008	-2.000
Latin America			
Argentina	0.971	0.014	-2.071
Brazil	0.988	0.009	-1.333
Mexico	0.975	0.014	-1.786
Peru	0.968	0.015	-2.133
Uruguay	0.987	0.009	-1.444
Venezuela	0.917	0.023	-3.609
Asia			
Indonesia	0.967	0.066	-0.500
Korea	1.140	0.063	-2.222
Malaysia	0.805	0.059	-3.305
Philippines	0.969	0.016	-1.938
Thailand	0.846	0.043	-3.581
Singapore	0.966	0.016	-2.125

Note: Approach based on testing for a unit root across the entire sample due to possible concerns that the real exchange rate series are not characterized by a unit root process, and instead that they are trend stationary around a segmented deterministic component. In all cases, the null of a unit root cannot be rejected at the 5% critical value of -3.76, however, thereby allaying these concerns. The model is based on estimating $z_t=a_0+a_2t+\mu_2D_L$+z-hat$_t$, where D_L is a level dummy such that D_L=1 if $t>\tau$ (i.e. the break date) and zero otherwise and z-hat$_t$ are the residuals. z-hat$_t$=a_1z-hat$_{t-1}+\varepsilon_t$ is then estimated. The null hypothesis of a_1=1 (i.e. null of a unit root) is rejected if the *t*-statistic is greater than the Perron critical value. See Perron (1989) for full details of the procedure.

The results of the regressions based on equations (7) and (8) above are provided in **Tables A3.18** to **A3.37**. Each of the regressions has been adjusted to eliminate serial correlation.[79] The resulting t-statistics under the null hypothesis that α_1=1, determine whether a unit root exists in the presence of the structural break across each crisis-affected region. In all cases, the null of a unit root cannot be rejected at the 5% critical

[79] The Cochrane-Orchutt iterative procedure was employed where evidence of second-order autoregressive behaviour in the residual series was found.

value of -3.76 and thereby confirming that the observed non-stationarity in the series is not purely driven by the break. This reinforces the results obtained from the more conventional ADF-based unit root tests. Using Perron's test, the proposition is that the real exchange rates are characterized by a unit root process, and that they are not trend stationary with a structural break (i.e. they are stationary except for the break). This confirms that the Johansen approach is an appropriate methodology to test for the existence of long-run relations.

VIII. Robustness Check 2: Impulse Responses from a VAR

In order to provide additional empirical support for the results found from the cointegration analysis in pre- and post-crisis periods, simulations of the response in the real exchange rate to shocks imposed on real output are undertaken for each of the three regions. Impulse responses from a bivariate vector autoregressive model (structural VAR in first differences) consisting of real output and the real exchange rate are estimated.[80] The same data period as was used in the cointegration analysis is used for the impulse response functions. However, a quarterly frequency is used instead of monthly due to the lack of availability of monthly GDP data.[81]

A uniform lag of order 2 is applied to the quarterly data, because this seems to correspond quite well with the lag order associated with the series and to maintain a coherent framework across country blocks and samples. Also the parameterization might be seen as excessive in some of the sub-periods when higher order lags are used. Therefore:

[80] A bi-variate assessment is carried out in order to keep the analysis relatively straightforward.
[81] While monthly industrial production data was available for the European sample, such data was not available for Latin America and South East Asia for a sufficiently long time period.

$$\begin{bmatrix} GDP \\ RER \end{bmatrix}_t = c + \sum_{i=1}^{2} A_i \begin{bmatrix} GDP \\ RER \end{bmatrix}_{t-1} + \varepsilon_t \tag{10}$$

where GDP represents a measure of real output (this has been adjusted for inflation using the CPI) and the RER denotes the real exchange rate. The focus is on the adjustment of the real exchange rate to an unexpected temporary shock in real output. In computing the impulse response functions the capacity to detect the contemporaneous relation between the real exchange rate and real output is derived from a Cholesky decomposition of the residual variance-covariance matrix. The VAR equations are estimated in first differences. The responses are estimated as the percentage shares of the accumulated responses of the real exchange rate to cumulative real output shocks that arise from Equation (10). The use of accumulated responses helps to aid interpretation. Since the VARs are estimated with first-differences of the variables, the accumulated responses of the first-differences equate to the responses of the variables in terms of levels. Moreover, the results are not sensitive to the ordering of the variables.[82] The plots of the impulse responses are shown in **Figures A3.2** to **A3.7**. They show that an unexpected temporary shock in real output is followed by a delayed adjustment of the real exchange rate in both pre- and post-crisis cases.

Figure A3.2 shows a small degree of variability across the countries in the pre-crisis period in terms of the response of the real exchange rate to a positive one standard deviation output shock. In the main, a depreciation in the real exchange rate is observed at first. This subsides after 4-5 quarters and then the exchange rate appreciates until the full adjustment is completed after 15-20 quarters. Across the sub-samples of groups of countries under consideration, the pattern of adjustment is remarkably similar for the

[82] A similar approach was undertaken in Beirne and de Bondt (2008).

Netherlands, France, Spain, Austria, Belgium, and Italy. The adjustment process takes considerably longer in the cases of Germany and the UK, although both exhibit a similar gradually appreciating exchange rate after the first 5 periods.

In the post-crisis period, for the majority of the countries, the initial depreciation in the exchange rate reverts to an appreciation after 2 periods. Then a gradual depreciation is observed until full adjustment is achieved after about 15 periods. The pattern of the response to the output shock is remarkably similar across the countries, both in terms of scale and length of time for the shock to subside. Generally the exchange rate adjusts by about 10-15% to the output shocks.

The pattern of adjustment for the Latin American economies in the pre-crisis period appears to be different. While the exchange rate initially depreciates in the cases of Argentina, Peru and Uruguay, the opposite initial effect takes place for Brazil, Mexico and Venezuela. For the post-crisis results, not only are there different initial effects, but also differences in scale and duration of impact. For example, the real output shock is fully transmitted to the real exchange rate for Mexico and Venezuela after approximately 15 quarters. This contrasts with the other countries in the region, where full pass-through is about three times longer than this. Having said that there does appear to be some similarity in the pattern of exchange rate adjustment for sub-samples of countries. For example, the real exchange rates of Argentina and Brazil exhibit similar behaviour, as do Peru and Uruguay (albeit to a lesser extent). Overall, the differences apparent for the countries as a whole suggest a lack of convergence in the exchange rates of said economies.

For the South East Asian economies, the exchange rate adjustment appears to differ across countries in the pre-crisis period. While the adjustment duration is relatively rapid in the cases of Korea, Malaysia, Thailand and Indonesia (10-15 quarters), it is substantially higher for the Philippines and Singapore. In the post-crisis case, the response patterns of the data seem to eventually converge. The transmission of the shock is complete after approximately 10 quarters for all countries. The pattern of adjustment is particularly symmetric in terms of both duration and scale for four economies: Indonesia, the Philippines, Thailand and Singapore.

Overall, the impulse responses indicate that there are notable (and significant) differences in the pre- and post-crisis results. In the pre-crisis case, the response of the real exchange rate to an output shock across the three regions is not the same. This contrasts with the post-crisis case for the European economies and a sub-group of the South East Asian countries, where a symmetry in the real exchange rate response appears to be observed. These results support the cointegration analysis that seems to indicate a greater degree of regional exchange rate policy co-ordination in Europe and a sub-group of the Asian countries. Such an observation is not apparent for the Latin American economies after the crisis.

IX. Conclusions

This study examines the existence of a long-run G-PPP relation between the countries of three regions that were affected by major financial crises, namely Europe (EMS crisis), Latin America (Mexican crisis), and South East Asia (Asian crisis). Using the Johansen multivariate cointegration technique, long-run stationary relationships were identified in the real exchange rates of the countries according to their regional grouping. The study examined three particular sample periods for each of the three regions: the full period (i.e. inclusive of the crisis period), the pre-crisis period, and the post-crisis period. The aim was to assess how the existence of a long-run G-PPP relation may be affected by a major crisis. Given the inherent economic turbulence during that time, the approach of Choudry (2005) is followed such that analysis is distinguished from the period of the crisis.

The key conclusion from the analysis appears to be that following crisis, the identification of a long-run equilibrium relationship amongst systems of real exchange rates in economic regions is more apparent. There also appears to be a faster speed of adjustment towards G-PPP in the post-crisis period, as shown in the short-run coefficient estimates of the Johansen test. Of course, this does not mean that monetary integration should be endorsed. A deeper analysis of the cointegration equations, based on the sign and magnitude of long-run coefficients and the similarity of short-run adjustment coefficients, revealed that monetary integration in the post-crisis case seems appropriate for the former countries of the EMS and a selection of countries in the South East Asian group (specifically Singapore, Indonesia, the Philippines, and Thailand). The Latin American countries do not appear to be suitable for monetary union, primarily due to the implied differences in the demand parameters across these countries.

Thus, after their respective crises, the European countries considered and the relevant South East Asian economies seem to have learned that a co-ordinated approach to regional exchange rate policy is favourable in terms of lower volatility and greater stability. This can help to insulate the economies from crises as such a co-ordinated arrangement helps to increase the credibility of the respective currencies, which reduces the likelihood of speculative attack. Clearly, in terms of monetary integration, the European countries in the sample are now part of a monetary union within EMU. This would appear to vindicate the conclusions made regarding the analysis of the former EMS countries. The South East Asian economies where monetary integration seems appropriate (i.e. Singapore Indonesia, the Philippines, and Thailand) may engage in some form of currency union in the future. The results from this study indicate that they are well placed to do so in terms of the interaction of their real exchange rates. For the Latin American economies, there has been a lot of variation across the real exchange rates, which was exacerbated by inflation and debt problems. It is not surprising, therefore, that a monetary integration of this particular group seems to be inappropriate. The dollarization observed in a number of Latin American countries might also be an issue.

The cointegrated VAR models estimated do not show signs of misspecification associated with error auto-correlation, ARCH, and heteroskedasticity. In addition, the cointegrating rank appears to be stable given the constant nature of the recursive estimates of the eigenvalues derived from a 60-month rolling window.

It was also shown that a key condition for cointegration, that all the series are I(1), is robust even when the Perron test (that tests for a unit root test in the presence of a structural break) is used. Across all of the countries examined and all of the real exchange rate series used there is no evidence that the real exchange rate cannot be characterized by a unit root process. Thus they are not trend stationary around a segmented deterministic component. This analysis was carried out to examine the proposition that a break in the series had led to the finding that they were non-stationary, i.e. that any non-stationarity purely arises by virtue of the break.

Furthermore, an alternative analysis based on impulse response functions from a first-differenced bi-variate VAR was used to determine the response of the real exchange rates to real output shocks in pre- and post-crisis periods. The findings from this analysis were consistent with those of the earlier cointegration analysis of systems of real exchange rates in the respective regions.

In extending the analysis considered thus far, it would be of interest to handle the exchange rate behaviour in a less restrictive framework. The preceding empirical work was carried out on real exchange rates and in the short-run the data implicitly imposes the proportionality and symmetry conditions associated with PPP. A less restrictive framework could be represented by a system comprising the nominal exchange rate and the ratio of domestic and foreign prices for each country. In this case, only the symmetry condition would be imposed. The most general case (with no restrictions imposed for proportionality or symmetry) could be constructed by including the nominal exchange rate, domestic prices, and foreign prices as separate series in the system. With both of these scenarios, however, as the dimension of the VAR expands, then the number of variables in the system increases and the cross country analysis becomes cumbersome and the parameterization burdensome.

Appendix 1

Table A3.1 International Currency Abbreviations

Abbreviation	Description
EMS Currencies	
LRUSDATS	Log of real Austrian schilling per US dollar
LRUSDBEF	Log of real Belgian franc per US dollar
LRUSDFRF	Log of real French franc per US dollar
LRUSDITL	Log of real Italian lira per US dollar
LRUSDNLG	Log of real Dutch guilder per US dollar
LRUSDESP	Log of real Spanish peseta per US dollar
LRUSDDEM	Log of real German DM per US dollar
LRUSDGBP	Log of real UK pound per US dollar
Latin American Currencies	
LRUSDBRL	Log of real Brazilian real per US dollar
LRUSDARS	Log of real Argentine peso per US dollar
LRUSDPEN	Log of real Peruvian nuevo sol per US dollar
LRUSDMXN	Log of real Mexican peso US dollar
LRUSDUYU	Log of real Uruguayan peso US dollar
LRUSDVEB	Log of real Venezuelan bolivar per US dollar
South East Asian Currencies	
LRUSDIDR	Log of real Indonesian rupiah per US dollar
LRUSDKRW	Log of real Korean won per US dollar
LRUSDMYR	Log of real Malaysian ringgit per US dollar
LRUSDPHP	Log of real Philippine peso per US dollar
LRUSDTHB	Log of real Thai baht per US dollar
LRUSDSGD	Log of real Singaporean dollar per US dollar

ADF Unit Root Tests for EMS, Latin American, and South East Asian Regions

Table A3.2 ADF Tests for Unit Roots – European Real Exchange Rates (v.USD): 1980M01-2006M12

(Log) Real Exchange Rate Series	**Model 1 Intercept and Trend**		**Model 2 Intercept, No Trend**		**Model 3 No Intercept, No Trend**	
	Levels	First Differences	Levels	First Differences	Levels	First Differences
LRUSDATS (Austrian schilling)	-2.89 (1) [2.85]* {0.026}	-8.65 (0) [-0.56] {0.55}	-2.56 (2) [2.55]**	-8.35 (1) [-0.21]	-0.26 (1)	-8.68 (0)
LRUSDBEF (Belgian franc)	-2.90 (1) [2.87]* {0.37}	-8.57 (0) [-0.53] {0.55}	-2.59 (2) [2.58]**	-8.11 (1) [-0.13]	-0.16 (1)	-8.61 (0)
LRUSDFRF (French franc)	-3.12 (1) [3.06]* {0.74}	-8.38 (0) [-0.55] {0.59}	-2.62 (1) [2.61]*	-8.56 (0) [-0.06]	-0.19 (2)	-8.40 (1)
LRUSDESP (Spanish peseta)	-2.68 (1) [2.63]* {1.53}	-8.27 (0) [-0.92] {0.95}	-2.39 (1) [2.38]**	-8.22 (0) [-0.19]	-0.24 (1)	-8.25 (0)
LRUSDITL (Italian lira)	-2.97 (3) [2.97]* {1.95}***	-5.79 (2) [-0.21] {0.29}	-1.88 (2) [1.88]***	-8.22 (1) [0.11]	0.08 (2)	-8.24 (1)
LRUSDNLG (Dutch guilder)	-3.12 (1) [2.95]* {0.15}	-8.38 (0) [-0.31] {0.32}	-2.62 (2) [2.59]**	-8.42 (1) [-0.09]	-0.40 (2)	-8.45 (1)

Table A3.3 ADF Tests for Unit Roots – European Real Exchange Rates (v.USD): 1980M01-1992M08

(Log) Real Exchange Rate Series	**Model 1 Intercept and Trend**		**Model 2 Intercept, No Trend**		**Model 3 No Intercept, No Trend**	
	Levels	First Differences	Levels	First Differences	Levels	First Differences
LRUSDATS (Austrian schilling)	-2.52 (1) [2.51]** {-1.19}	-5.73 (0) [0.01] {-0.45}	-2.26 (1) [2.24]**	-5.75 (0) [-0.74]	-0.80 (1)	-5.73 (0)
LRUSDBEF (Belgian franc)	-2.69 (1) [2.69]* {-1.24}	-5.44 (0) [-0.02] {-0.36}	-2.41 (1) [2.40]**	-5.47 (0) [-0.65]	-0.69 (1)	-5.46 (0)
LRUSDFRF (French franc)	-2.68 (1) [2.67]* {-1.14}	-5.51 (0) [-0.03] {-0.34}	-2.44 (1) [2.41]**	-5.54 (0) [-0.63]	-0.73 (1)	-5.53 (0)
LRUSDESP (Spanish peseta)	-2.96 (1) [2.94]* {-2.67}	-5.89 (0) [-0.66] {-0.09}	-1.23 (1) [1.19]	-5.94 (0) [-1.43]	-1.46 (1)	-5.72 (0)
LRUSDITL (Italian lira)	-2.82 (1) [2.82]* {-2.34}**	-5.39 (0) [-0.04] {-0.49}	-1.60 (1) [1.59]	-5.40 (0) [-0.90]	-0.92 (1)	-5.34 (0)
LRUSDNLG (Dutch guilder)	-2.58 (1) [2.56]** {-0.77}	-5.45 (0) [0.17] {-0.50}	-2.52 (1) [2.46]**	-5.47 (0) [-0.52]	-0.76 (1)	-5.47 (0)

Table A3.4 ADF Tests for Unit Roots – European Real Exchange Rates (v.USD): 1993M01-2006M12

(Log) Real Exchange Rate Series	Model 1 Intercept and Trend		Model 2 Intercept, No Trend		Model 3 No Intercept, No Trend	
	Levels	First Differences	Levels	First Differences	Levels	First Differences
LRUSDATS (Austrian schilling)	-1.65 (1) [1.64] {0.87}	-6.21 (0) [-0.02] {0.19}	-1.42 (1) [1.42]	-6.25 (0) [0.29]	0.23 (1)	-6.29 (0)
LRUSDBEF (Belgian franc)	-1.53 (1) [1.53] {0.71}	-6.56 (0) [0.14] {0.08}	-1.36 (1) [1.37]	-6.61 (0) [0.40]	0.49 (1)	-6.64 (0)
LRUSDFRF (French franc)	-1.69 (1) [1.70]*** {0.77}	-6.61 (0) [0.42] {-0.17}	-1.51 (1) [1.53]	-6.66 (0) [0.55]	0.46 (1)	-6.62 (0)
LRUSDESP (Spanish peseta)	-2.08 (1) [2.11]** {0.54}	-6.05 (0) [1.17] {-0.89}	-2.12 (1) [2.13]**	-6.12 (0) [0.64]	0.61 (1)	-6.11 (0)
LRUSDITL (Italian lira)	-2.57 (1) [2.57]** {-0.15}	-6.65 (0) [1.16] {-1.06}	-2.81 (1) [2.81]*	-5.96 (0) [0.30]	0.47 (2)	-6.59 (1)
LRUSDNLG (Dutch guilder)	-1.78 (1) [1.72]*** {0.77}	-6.39 (0) [0.24] {-0.07}	-1.61 (1) [1.62]	-6.44 (0) [0.36]	0.18 (2)	-6.47 (1)

Table A3.5 ADF Tests for Unit Roots – Latin American Real Exchange Rates (v.USD): 1983M01-2006M12

(Log) Real Exchange Rate Series	Model 1 Intercept and Trend		Model 2 Intercept, No Trend		Model 3 No Intercept, No Trend	
	Levels	First Differences	Levels	First Differences	Levels	First Differences
LRUSDARS (Argentine peso)	-2.54 (6) [7.28]* {4.46}*	-15.62 (5) [-7.01]* {6.29}*	-1.52 (3) [-1.78]***	-11.99 (2) [-1.03]	-0.83 (3)	-16.37 (2)
LRUSDBRL (Brazilian real)	-1.31 (7) [1.21] {-0.74}	-5.24 (6) [0.04] {0.02}	-1.52 (1) [1.41]	-8.70 (0) [-0.03]	-0.25 (7)	-5.48 (6)
LRUSDUYU (Uruguayan peso)	-0.23 (1) [-0.05] {1.50}	-10.62 (0) [-4.92]* {3.69}*	-2.47 (4) [2.33]*	-3.36 (3) [-1.78]***	-1.70 (6)	-2.41 (5)
LRUSDVEB (Venezuelan bolivar)	-2.05 (2) [2.06]** {-2.16}**	-8.49 (1) [0.09] {-0.84}	-0.50 (2) [0.47]	-8.46 (1) [-1.23]	-1.24 (2)	-8.35 (1)
LRUSDPEN (Peruvian nuevo sol)	-2.04 (9) [1.86]*** {0.79}	-6.73 (8) [-2.12]** {2.07}**	-2.83 (9) [2.80]*	-6.35 (8) [-0.61]	-0.73 (9)	-6.35 (8)
LRUSDMXN (Mexican peso)	-2.33 (1) [2.31]** {1.15}	-5.76 (0) [-0.10] {0.16}	-2.03 (1) [2.03]**	-7.80 (0) [0.01]	-0.10 (1)	-7.84 (0)

Table A3.6 ADF Tests for Unit Roots – Latin American Real Exchange Rates (v.USD): 1983M01-1994M11

(Log) Real Exchange Rate Series	Model 1 Intercept and Trend		Model 2 Intercept, No Trend		Model 3 No Intercept, No Trend	
	Levels	First Differences	Levels	First Differences	Levels	First Differences
LRUSDARS (Argentine peso)	0.003 (9) [-7.27]* {3.75}*	-11.67 (5) [-4.95]* {3.09}*	-2.62 (6) [-5.33]*	-10.33 (5) [-4.36]*	-0.13 (4)	-11.42 (3)
LRUSDBRL (Brazilian real)	0.02 (7) [1.04] {1.86}***	-4.71 (6) [2.41]** {-2.53}**	-1.68 (1) [1.60]	-6.02 (0) [-0.33]	0.62 (1)	-6.08 (0)
LRUSDUYU (Uruguayan peso)	-2.45 (1) [2.42]** {-2.43}**	-7.58 (0) [-2.35]** {-0.02}	-0.31 (1) [0.07]	-7.68 (0) [-4.27]*	-1.70 (11)	-2.22 (2)
LRUSDVEB (Venezuelan bolivar)	-3.39 (3) [3.39]* {-1.33}**	-5.11 (3) [1.01] {1.07}	-1.94 (4) [1.94]***	-5.05 (3) [-0.13]	-0.14 (4)	-5.11 (3)
LRUSDPEN (Peruvian nuevo sol)	-1.97 (9) [1.86]*** {0.85}	-4.79 (8) [-1.03] {0.62}	-1.89 (9) [1.85]***	-4.82 (8) [-1.02]	-1.10 (9)	-4.73 (8)
LRUSDMXN (Mexican peso)	-0.79 (1) [0.71] {0.44}	-5.74 (0) [-2.39]** {2.04}**	-2.16 (1) [2.12]**	-5.20 (0) [-1.20]	-1.26 (1)	-5.05 (0)

Table A3.7 ADF Tests for Unit Roots – Latin American Real Exchange Rates (v.USD): 1995M09-2006M12

(Log) Real Exchange Rate Series	Model 1 Intercept and Trend		Model 2 Intercept, No Trend		Model 3 No Intercept, No Trend	
	Levels	First Differences	Levels	First Differences	Levels	First Differences
LRUSDARS (Argentine peso)	-2.39 (1) [-2.23]** {2.02}***	-4.35 (0) [1.33] {-0.55}	-1.36 (1) [-1.01]	-4.36 (0) [1.66]	-1.90 (1)	-3.96 (0)
LRUSDBRL (Brazilian real)	-1.15 (4) [-0.50] {4.77}*	-5.98 (2) [-3.43]* {4.88}*	0.94 (1) [-0.68]	-2.96 (0) [1.19]	1.77 (5)	-2.59 (0)
LRUSDUYU (Uruguayan peso)	-3.85 (0) [3.86]* {3.28}*	-4.34 (0) [1.61] {-0.45}	-1.82 (1) [1.84]***	-7.12 (0) [1.54]	1.52 (1)	-6.85 (0)
LRUSDVEB (Venezuelan bolivar)	-2.56 (1) [2.59]** {-2.56}**	-5.54 (0) [0.74] {-1.10}	-1.10 (1) [1.09]	-5.41 (0) [-0.41]	-0.43 (1)	-5.46 (0)
LRUSDPEN (Peruvian nuevo sol)	-0.79 (1) [0.78] {1.39}	-5.86 (0) [-0.13] {1.36}	-0.69 (1) [-0.64]	-5.66 (0) [1.89]***	1.91 (1)	-3.83 (0)
LRUSDMXN (Mexican peso)	-2.32 (1) [2.31]** {-1.77}***	-5.74 (0) [-2.39]** {2.04}**	-1.25 (6) [1.20]	-4.68 (5) [-1.50]	-1.55 (6)	-4.35 (5)

Table A3.8 ADF Tests for Unit Roots –Asian Real Exchange Rates (v.USD): 1988M01-2006M12

(Log) Real Exchange Rate Series	Model 1 Intercept and Trend		Model 2 Intercept, No Trend		Model 3 No Intercept, No Trend	
	Levels	First Differences	Levels	First Differences	Levels	First Differences
LRUSDIDR (Indonesian rupiah)	-2.48 (1) [2.50]** {1.37}	-11.44 (0) [0.37] {-0.29}	-2.08 (1) [2.09]**	-11.46 (0) [0.24]	-0.17 (1)	-11.49 (0)
LRUSDKRW (Korean won)	-1.82 (2) [1.84]*** {0.77}	-9.95 (1) [0.61] {-0.59}	-2.08 (1) [2.71]*	-8.45 (0) [0.12]	0.16 (2)	-9.97 (1)
LRUSDMYR (Malaysian ringgitt	-2.04 (1) [2.03]** {1.78}***	-11.15 (0) [0.19] {0.22}	-1.00 (1) [2.89]*	-11.17 (0) [0.77]	0.61 (1)	-11.16 (0)
LRUSDPHP (Philippine peso)	-1.87 (1) [1.86]*** {1.50}	-9.53 (0) [-0.11] {0.33}	-1.16 (1) [1.18]	-9.55 (0) [0.36]	0.30 (1)	-9.57 (0)
LRUSDTHB (Thai baht)	-2.59 (1) [2.59]** {2.17}**	-9.95 (0) [0.16] {0.12}	-1.39 (1) [1.42]	-9.97 (0) [0.54]	0.46 (1)	-9.98 (0)

Table A3.9 ADF Tests for Unit Roots –Asian Real Exchange Rates (v. USD): 1988M01-1997M06

(Log) Real Exchange Rate Series	Model 1 Intercept and Trend		Model 2 Intercept, No Trend		Model 3 No Intercept, No Trend	
	Levels	First Differences	Levels	First Differences	Levels	First Differences
LRUSDIDR (Indonesian rupiah)	-1.54 (8) [1.54] {-1.43}	-4.33 (7) [-0.85] {0.18}	-0.98 (9) [0.98]	-3.12 (9) [-1.20]	-0.99 (9)	-2.46 (8)
LRUSDKRW (Korean won)	-0.89 (1) [0.88] {0.59}	-7.03 (0) [-0.62] {0.95}	-1.50 (2) [1.50]	-4.80 (1) [0.31]	0.42 (1)	-6.99 (0)
LRUSDMYR (Malaysian ringgitt	-2.79 (1) [2.76]* {-2.40}**	-6.93 (0) [-0.40] {0.05}	-1.37 (1) [1.33]	-6.97 (0) [-0.72]	-0.79 (1)	-6.95 (0)
LRUSDPHP (Philippine peso)	-3.46 (10) [3.42]* {-3.02}*	-4.96 (9) [-2.89]* {1.54}	-0.58 (1) [0.54]	-6.56 (0) [-1.26]	-1.28 (1)	-6.42 (0)
LRUSDTHB (Thai baht)	-3.07 (1) [3.06]* {-2.91}*	-8.04 (0) [-0.89] {0.11}	-1.05 (5) [1.02]	-5.60 (4) [-2.43]**	-1.90 (4)	-4.62 (3)

Table A3.10 ADF Tests for Unit Roots –Asian Real Exchange Rates (v.USD): 1998M06-2006M12

(Log) Real Exchange Rate Series	Model 1 Intercept and Trend		Model 2 Intercept, No Trend		Model 3 No Intercept, No Trend	
	Levels	First Differences	Levels	First Differences	Levels	First Differences
LRUSDIDR (Indonesian rupiah)	-3.44 (7) [3.39]* {-1.57}	-5.63 (6) [-2.02]** {1.45}	-2.37 (11) [2.35]**	-6.39 (10) [-1.46]	-1.50 (11)	-6.20 (10)
LRUSDKRW (Korean won)	-1.41 (7) [1.40] {-1.21}	-4.83 (6) [-0.52] {-0.29}	-1.70 (8) [1.69]***	-3.32 (7) [-1.30]	-1.37 (10)	-3.43 (9)
LRUSDMYR (Malaysian ringgitt	-2.73 (6) [2.72]*** {2.60}**	-6.86 (5) [-0.40] {0.67}	-1.68 (9) [1.69]***	-4.11 (8) [0.07]	0.05 (9)	-4.17 (8)
LRUSDPHP (Philippine peso)	-1.11 (1) [1.15] {0.20}	-6.67 (0) [1.18] {-1.11}	-1.57 (1) [1.57]	-6.57 (0) [0.44]	0.40 (1)	-6.59 (0)
LRUSDTHB (Thai baht)	-1.57 (10) [1.60] {0.13}	-3.52 (9) [0.87] {-0.81}	-1.78 (10) [1.78]***	-3.52 (9) [0.35]	0.32 (10)	-3.54 (9)

Notes: ADF Critical Value (5%) is –3.46 for Model 1, –2.89 for Model 2, and –1.94 for Model 3. Beneath the ADF test statistics, the optimal lag lengths are reported in round brackets, the significance of the intercept term is reported in square brackets, and the significance of the trend is reported below this (both t-statistics). *, **, and *** indicates statistical significance at the 1%, 5%, and 10% levels respectively.

Table A3.11 Misspecification Tests: Full Sample

Region	Variable	$AR_{(1-7)}$	Normality	$ARCH_{(1-7)}$	Hetero
Europe	USDAUS	0.589 [0.765]	18.324 [0.000]**	1.483 [0.187]	0.766 [0.927]
	USDDEM	1.574 [0.144]	23.187 [0.000]**	1.629 [0.124]	1.318 [0.073]
	USDGBP	1.560 [0.148]	26.409 [0.000]**	1.552 [0.151]	0.797 [0.892]
	USDBEF	1.127 [0.347]	0.307 [0.858]	0.219 [0.981]	0.533 [0.999]
	USDFRF	0.897 [0.509]	0.149 [0.928]	0.925 [0.488]	0.507 [0.999]
	USDITL	1.258 [0.272]	0.457 [0.796]	1.149 [0.334]	0.548 [0.999]
	USDNLG	0.753 [0.627]	0.136 [0.935]	0.601 [0.755]	0.490 [1.000]
	USDESP	0.571 [0.779]	0.618 [0.734]	1.037 [0.406]	0.508 [0.999]
Latin America	USDARS	1.322 [0.232]	112.77 [0.000]**	1.876 [0.076]	0.799 [0.855]
	USDBRL	1.296 [0.254]	67.114 [0.000]**	1.488 [0.188]	0.252 [1.000]
	USDMXN	1.258 [0.273]	141.61 [0.000]**	1.533 [0.158]	0.857 [0.769]
	USDPEN	0.989 [0.441]	244.88 [0.000]**	0.764 [0.618]	1.336 [0.105]
	USDUYU	1.206 [0.291]	51.789 [0.000]**	1.944 [0.065]	1.142 [0.287]
	USDVEB	0.684 [0.686]	633.41 [0.000]**	0.064 [0.999]	0.164 [1.000]
Asia	USDIDR	1.590 [0.195]	2.015 [0.365]	0.212 [0.968]	0.862 [0.745]
	USDKRW	1.398 [0.201]	1.313 [0.519]	0.647 [0.692]	0.929 [0.632]
	USDMYR	1.811 [0.141]	3.755 [0.153]	0.417 [0.858]	0.888 [0.701]
	USDPHP	1.723 [0.161]	0.324 [0.850]	0.429 [0.849]	0.296 [1.000]
	USDTHB	1.828 [0.138]	0.547 [0.761]	0.209 [0.969]	0.378 [1.000]
	USDSGD	1.079 [0.404]	2.002 [0.368]	0.197 [0.973]	0.590 [0.989]

Table A3.12 Misspecification Tests: Pre-Crisis Sample

Region	Variable	$AR_{(1-7)}$	Normality	$ARCH_{(1-7)}$	Hetero
Europe	USDAUS	1.037 [0.409]	0.286 [0.8669]	0.534 [0.807]	0.798 [0.799]
	USDDEM	0.595 [0.759]	0.780 [0.6771]	1.928 [0.072]	0.879 [0.682]
	USDGBP	0.575 [0.775]	0.163 [0.9217]	1.704 [0.116]	0.631 [0.957]
	USDBEF	1.483 [0.180]	1.516 [0.4686]	0.649 [0.715]	0.724 [0.886]
	USDFRF	1.309 [0.252]	1.067 [0.5866]	0.541 [0.801]	0.774 [0.830]
	USDITL	1.018 [0.422]	0.475 [0.7885]	0.416 [0.891]	0.833 [0.751]
	USDNLG	0.799 [0.589]	0.099 [0.9517]	0.646 [0.717]	0.828 [0.759]
	USDESP	1.560 [0.154]	2.465 [0.2916]	0.784 [0.602]	0.833 [0.752]
Latin America	USDARS	1.981 [0.066]	262.40 [0.000]**	0.029 [1.000]	0.678 [0.858]
	USDBRL	1.656 [0.130]	18.546 [0.000]**	0.244 [0.973]	0.270 [1.000]
	USDMXN	0.488 [0.841]	69.394 [0.000]**	0.113 [0.997]	0.385 [0.999]
	USDPEN	1.266 [0.276]	98.130 [0.000]**	0.408 [0.895]	0.518 [0.983]
	USDUYU	1.237 [0.263]	80.471 [0.000]**	1.096 [0.373]	0.817 [0.749]
	USDVEB	1.520 [0.164]	392.55 [0.000]**	0.015 [1.000]	0.360 [0.999]
Asia	USDIDR	1.332 [0.247]	0.057 [0.972]	0.850 [0.550]	0.665 [0.876]
	USDKRW	1.329 [0.249]	2.879 [0.237]	0.876 [0.531]	0.483 [0.979]
	USDMYR	0.657 [0.707]	19.368 [0.000]**	0.112 [0.997]	0.184 [1.000]
	USDPHP	2.054 [0.060]	22.607 [0.000]**	0.262 [0.966]	0.278 [0.999]
	USDTHB	1.863 [0.089]	1.137 [0.566]	0.545 [0.797]	0.275 [0.999]
	USDSGD	1.456 [0.197]	2.879 [0.237]	1.257 [0.286]	0.235 [1.000]

Table A3.13 Misspecification Tests: Post-Crisis Sample

Region	Variable	$AR_{(1-7)}$	Normality	$ARCH_{(1-7)}$	Hetero
Europe	USDATS	1.306 [0.254]	3.193 [0.198]	1.392 [0.217]	0.439 [0.995]
	USDDEM	1.803 [0.094]	3.657 [0.161]	1.575 [0.151]	0.477 [0.990]
	USDGBP	1.149 [0.338]	1.812 [0.404]	0.483 [0.845]	0.189 [1.000]
	USDBEF	1.313 [0.251]	1.476 [0.478]	1.276 [0.269]	0.231 [1.000]
	USDFRF	1.139 [0.344]	1.078 [0.583]	1.661 [0.127]	0.219 [1.000]
	USDITL	1.013 [0.426]	1.199 [0.549]	1.970 [0.066]	0.206 [1.000]
	USDNLG	1.261 [0.276]	1.610 [0.447]	1.387 [0.219]	0.229 [1.000]
	USDESP	1.100 [0.368]	1.379 [0.502]	1.195 [0.312]	0.203 [1.000]
Latin America	USDARS	1.289 [0.268]	100.31 [0.000]**	0.0923 [0.999]	123.18 [0.184]
	USDBRL	0.580 [0.769]	35.072 [0.000]**	0.155 [0.993]	110.88 [0.459]
	USDMXN	1.016 [0.427]	34.708 [0.000]**	0.102 [0.998]	98.572 [0.774]
	USDPEN	0.793 [0.596]	3.689 [0.158]	0.935 [0.486]	108.82 [0.514]
	USDUYU	1.343 [0.244]	5.064 [0.079]	1.353 [0.207]	107.39 [0.553]
	USDVEB	1.654 [0.134]	27.913 [0.000]**	0.163 [0.992]	127.08 [0.127]
Asia	USDIDR	0.761 [0.603]	2.502 [0.286]	1.478 [0.201]	0.219 [0.999]
	USDKRW	0.701 [0.650]	1.775 [0.412]	0.256 [0.955]	0.157 [1.000]
	USDMYR	2.173 [0.057]	15.608 [0.000]**	0.126 [0.993]	0.434 [0.969]
	USDPHP	2.115 [0.063]	4.320 [0.115]	0.221 [0.968]	0.119 [1.000]
	USDTHB	0.414 [0.867]	0.846 [0.655]	0.601 [0.729]	0.151 [1.000]
	USDSGD	1.068 [0.391]	0.392 [0.822]	0.514 [0.796]	0.184 [1.000]

Table A3.14 **Skewness and Excess Kurtosis of Systems with signs of Non-normality**

	Variables	Skewness	Excess Kurtosis
Europe All	USDATS	0.142	3.052
	USDDEM	-0.162	4.970
	USDGBP	-0.184	4.122
	USDBEF	1.327	13.530
	USDFRF	0.436	4.474
	USDITL	0.888	7.795
	USDNLG	-0.775	6.427
	USDESP	0.035	4.699
Latin America All	USDARS	0.790	9.285
	USDBRL	1.285	9.405
	USDMXN	3.112	28.683
	USDPEN	4.849	60.232
	USDUYU	0.418	6.092
	USDVEB	4.633	41.958
Latin America Pre-Crisis	USDARS	0.028	3.879
	USDBRL	-0.123	3.902
	USDMXN	2.011	12.876
	USDPEN	1.764	15.269
	USDUYU	0.203	3.7140
	USDVEB	2.334	18.937
Asia Pre-Crisis	USDIDR	-0.058	2.981
	USDKRW	0.055	3.214
	USDMYR	0.689	5.891
	USDPHP	1.132	7.322
	USDTHB	0.185	3.060
	USDSGD	-0.350	3.169
Latin America Post-Crisis	USDARS	3.625	32.345
	USDBRL	0.986	6.884
	USDMXN	1.334	8.417
	USDPEN	0.004	3.168
	USDUYU	0.044	3.797
	USDVEB	1.297	6.638
Asia Post-Crisis	USDIDR	0.070	3.343
	USDKRW	-0.184	2.626
	USDMYR	-0.202	5.082
	USDPHP	-0.357	2.982
	USDTHB	0.065	3.444
	USDSGD	0.133	2.882

Table A3.15 **Intervention Dummy Variables**

Country	Period
Europe	1985M04 (LRUSDGBP)
Latin America	1986M12 (LRUSDVEB), 1988M09 (LRUSDPEN), 1999M01 (LRUSDBRR), 2002M01 (LRUSDARG)
Asia	1990M11 (LRUSDPHP)

Table A3.16 Ramsey RESET Test

Region	Test Statistic [p-value]		
	Full Sample	**Pre-Crisis**	**Post-Crisis**
Europe	1.926 [0.197]	0.646 [0.425]	0.206 [0.651]
Latin America	1.185 [0.278]	2.687 [0.107]	2.270 [0.143]
Asia	0.073 [0.788]	0.074 [0.786]	0.921 [0.341]

Table A3.17 Long-Run Exclusion ($\lambda^2(1)$ statistics)

Europe								
	Austria	*Germany*	*UK*	*Belgium*	*France*	*Italy*	*Netherlands*	*Spain*
Full Sample	13.371*	4.967**	5.107**	19.695*	20.374*	3.503***	1.533	0.445
Pre-Crisis	14.937*	11.318*	12.367*	25.216*	21.720*	0.009	3.886**	0.119
Post-Crisis	18.553*	10.063*	9.044*	7.958*	1.054	9.222*	3.064***	10.330*

Latin America						
	Argentina	*Brazil*	*Mexico*	*Peru*	*Uruguay*	*Venezuela*
Full Sample	25.590*	19.390*	5.118**	9.188*	25.126*	4.073**
Pre-Crisis	9.833*	3.071*	8.447*	8.584*	3.095***	1.278
Post-Crisis	29.399*	12.045*	11.187*	5.564**	5.494**	3.849**
Asia						
	Indonesia	*Korea*	*Malaysia*	*Philippines*	*Thailand*	*Singapore*
Full Sample	12.307*	11.140*	9.088*	2.518	9.404*	11.290*
Pre-Crisis	n/a	n/a	n/a	n/a	n/a	n/a
Post-Crisis	7.259*	11.427*	6.995*	12.215*	6.411*	12.387*

Note: Following Juselius (2006), the long-run exclusion restriction can sometimes be accepted for reasons relating to multicollinearity. Moreover, even where the test is not rejected (indicating that the variable in question is long-run excludable), there can be grounds for the variable to remain in the model if it is significant in the Π matrix. For this reasons, all of the variables in the systems above are deemed non-excludable for the long-run relation.

Table A3.18 Perron Unit Root Test: Austria

Dependent Variable: LRUSDAUS				
	Coefficient	Std. Error	t-Statistic	Prob.
C	1.123346	0.008694	129.2077	0.0000
T	-8.39E-05	8.52E-05	-0.984344	0.3257
DL	-0.026647	0.015966	-1.668950	0.0961
Dependent Variable: AU_RES				
	Coefficient	Std. Error	t-Statistic	Prob.
AU_RES(-1)	0.981837	0.009509	103.2568	0.0000

Table A3.19 Perron Unit Root Test: Belgium

Dependent Variable: LRUSDBEF				
	Coefficient	Std. Error	t-Statistic	Prob.
C	1.558261	0.008886	175.3675	0.0000
T	5.57E-05	8.71E-05	0.639893	0.5227
DL	-0.026473	0.016318	-1.622282	0.1057
Dependent Variable: BEF_RES				
	Coefficient	Std. Error	t-Statistic	Prob.
BE_RES(-1)	0.979209	0.009220	106.2078	0.0000

Table A3.20 Perron Unit Root Test: Germany

Dependent Variable: LRUSDDEM				
	Coefficient	Std. Error	t-Statistic	Prob.
C	0.238369	0.008429	28.28125	0.0000
T	0.000129	8.26E-05	1.556384	0.1206
DL	-0.039227	0.015479	-2.534296	0.0117
Dependent Variable: DEM_RES				
	Coefficient	Std. Error	t-Statistic	Prob.
DEM_RES(-1)	0.976711	0.010689	91.37884	0.0000

Table A3.21 Perron Unit Root Test: France

Dependent Variable: LRUSDFRF				
	Coefficient	Std. Error	t-Statistic	Prob.
C	0.752174	0.008408	89.45991	0.0000
T	8.49E-05	8.24E-05	1.031027	0.3033
DL	-0.016583	0.015441	-1.073982	0.2836
Dependent Variable: FRF_RES				
	Coefficient	Std. Error	t-Statistic	Prob.
FRF_RES(-1)	0.980581	0.009597	102.1710	0.0000

Table A3.22 Perron Unit Root Test: Italy

Dependent Variable: LRUSDITL				
	Coefficient	Std. Error	t-Statistic	Prob.
C	3.252634	0.008178	397.7371	0.0000
T	-0.000433	8.01E-05	-5.400238	0.0000
DL	0.084126	0.015018	5.601583	0.0000
Dependent Variable: ITL_RES				
	Coefficient	Std. Error	t-Statistic	Prob.
ITL_RES(-1)	0.980936	0.010016	97.93476	0.0000

Table A3.23 Perron Unit Root Test: Netherlands

Dependent Variable: LRUSDNEG				
	Coefficient	Std. Error	t-Statistic	Prob.
C	0.309314	0.007954	38.88563	0.0000
T	-2.62E-05	7.79E-05	-0.336317	0.7369
DL	-0.019712	0.014608	-1.349407	0.1782
Dependent Variable: NEG_RES				
	Coefficient	Std. Error	t-Statistic	Prob.
NEG_RES(-1)	0.976374	0.010258	95.18464	0.0000

Table A3.24 Perron Unit Root Test: Spain

Dependent Variable: LRUSDSPP				
	Coefficient	Std. Error	t-Statistic	Prob.
C	2.201916	0.009707	226.8417	0.0000
T	-0.000387	9.51E-05	-4.069092	0.0001
DL	0.047544	0.017826	2.667084	0.0080
Dependent Variable: SP_RES				
	Coefficient	Std. Error	t-Statistic	Prob.
SP_RES(-1)	0.983935	0.008238	119.4420	0.0000

Table A3.25 Perron Unit Root Test: UK

Dependent Variable: LRUSDGBP				
	Coefficient	Std. Error	t-Statistic	Prob.
C	-0.150339	0.006546	-22.96820	0.0000
T	-0.000382	6.41E-05	-5.961479	0.0000
DL	0.035642	0.012021	2.965092	0.0033
Dependent Variable: GBP_RES				
	Coefficient	Std. Error	t-Statistic	Prob.
UK_RES(-1)	0.973276	0.011620	83.75850	0.0000

Table A3.26 Perron Unit Root Test: Argentina

Dependent Variable: LRUSDARG				
	Coefficient	Std. Error	t-Statistic	Prob.
C	0.178371	0.023567	7.568625	0.0000
T	0.000278	0.000255	1.091094	0.2762
DL	-0.112189	0.042334	-2.650068	0.0085
Dependent Variable: ARG_RES				
	Coefficient	Std. Error	t-Statistic	Prob.
ARG_RES(-1)	0.971340	0.013924	69.75867	0.0000

Table A3.27 Perron Unit Root Test: Brazil

Dependent Variable: LRUSDBRR				
	Coefficient	Std. Error	t-Statistic	Prob.
C	0.200256	0.015550	12.87810	0.0000
T	0.000142	0.000168	0.846352	0.3981
DL	0.002006	0.027933	0.071801	0.9428
Dependent Variable: BRA_RES				
	Coefficient	Std. Error	t-Statistic	Prob.
BRA_RES(-1)	0.987602	0.009141	108.0435	0.0000

Table A3.28 Perron Unit Root Test: Uruguay

Dependent Variable: LRUSDURU				
	Coefficient	Std. Error	t-Statistic	Prob.
C	1.229407	0.013323	92.27746	0.0000
T	0.000231	0.000144	1.606389	0.1093
DL	-0.135423	0.023932	-5.658555	0.0000
Dependent Variable: URU_RES				
	Coefficient	Std. Error	t-Statistic	Prob.
URU_RES(-1)	0.986897	0.009515	103.7157	0.0000

Table A3.29 Perron Unit Root Test: Venezuela

Dependent Variable: LRUSDVEN				
	Coefficient	Std. Error	t-Statistic	Prob.
C	2.925887	0.011677	250.5701	0.0000
T	0.001053	0.000126	8.350288	0.0000
DL	-0.226767	0.020976	-10.81095	0.0000
Dependent Variable: VEN_RES				
	Coefficient	Std. Error	t-Statistic	Prob.
VEN_RES(-1)	0.916815	0.022802	40.20692	0.0000

Table A3.30 Perron Unit Root Test: Peru

Dependent Variable: LRUSDPEP				
	Coefficient	Std. Error	t-Statistic	Prob.
C	0.842223	0.018851	44.67776	0.0000
T	-0.002038	0.000204	-10.00476	0.0000
DL	0.109488	0.033863	3.233274	0.0014
Dependent Variable: PER_RES				
	Coefficient	Std. Error	t-Statistic	Prob.
PER_RES(-1)	0.967923	0.015217	63.60748	0.0000

Table A3.31 Perron Unit Root Test: Mexico

Dependent Variable: LRUSDMEP				
	Coefficient	Std. Error	t-Statistic	Prob.
C	1.159140	0.006293	184.1812	0.0000
T	-0.001192	6.80E-05	-17.52742	0.0000
DL	0.112130	0.011305	9.918457	0.0000
Dependent Variable: MEX_RES				
	Coefficient	Std. Error	t-Statistic	Prob.
MEX_RES(-1)	0.975007	0.013555	71.92733	0.0000

Table A3.32 Perron Unit Root Test: Indonesia

Dependent Variable: LRUSDINR				
	Coefficient	Std. Error	t-Statistic	Prob.
C	3.727446	0.008679	429.4924	0.0000
T	-0.000970	0.000118	-8.211747	0.0000
DL	0.330612	0.015552	21.25844	0.0000
Dependent Variable: INR_RES				
	Coefficient	Std. Error	t-Statistic	Prob.
INR_RES(-1)	0.966548	0.065719	14.70728	0.0000
INR_RES(-2)	-0.182674	0.065704	-2.780258	0.0059

Table A3.33 Perron Unit Root Test: Korea

Dependent Variable: LRUSDKOW				
	Coefficient	Std. Error	t-Statistic	Prob.
C	2.981841	0.005054	590.0079	0.0000
T	-0.000641	6.88E-05	-9.313137	0.0000
DL	0.182967	0.009056	20.20301	0.0000
Dependent Variable: KOW_RES				
	Coefficient	Std. Error	t-Statistic	Prob.
KOW_RES(-1)	1.141828	0.063273	18.04605	0.0000
KOW_RES(-2)	-0.325340	0.063610	-5.114641	0.0000

Table A3.34 Perron Unit Root Test: Malaysia

Dependent Variable: LRUSDMAR				
	Coefficient	Std. Error	t-Statistic	Prob.
C	0.441054	0.003049	144.6398	0.0000
T	-9.07E-05	4.15E-05	-2.183843	0.0300
DL	0.157087	0.005464	28.74791	0.0000
Dependent Variable: MAR_RES				
	Coefficient	Std. Error	t-Statistic	Prob.
MAR_RES(-1)	0.805433	0.059085	13.63177	0.0000

Table A3.35 Perron Unit Root Test: Philippines

Dependent Variable: LRUSDPHP				
	Coefficient	Std. Error	t-Statistic	Prob.
C	1.560938	0.007399	210.9763	0.0000
T	-0.000268	0.000101	-2.661018	0.0084
DL	0.141575	0.013258	10.67836	0.0000
Dependent Variable: PHP_RES				
	Coefficient	Std. Error	t-Statistic	Prob.
PHP_RES(-1)	0.969178	0.016271	59.56540	0.0000

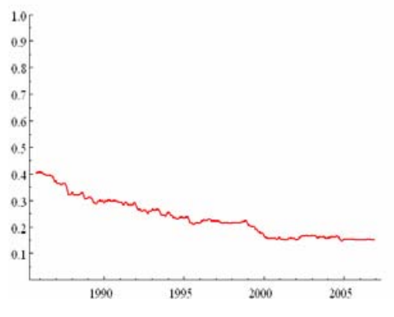
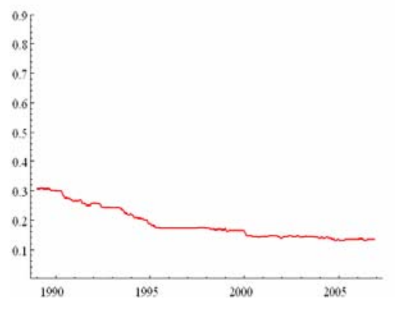
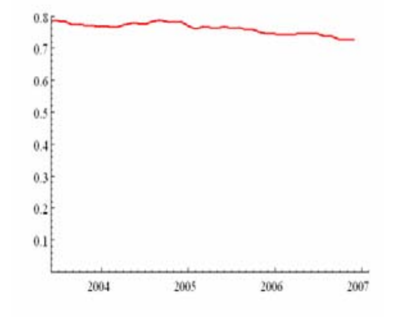
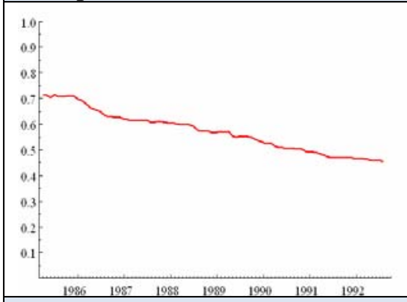
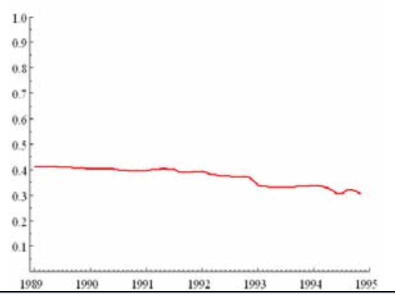
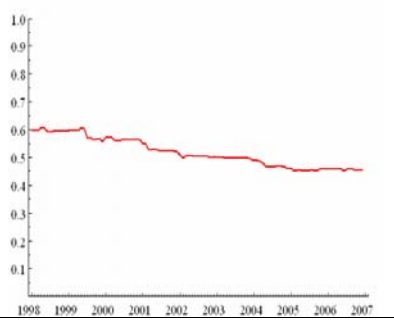
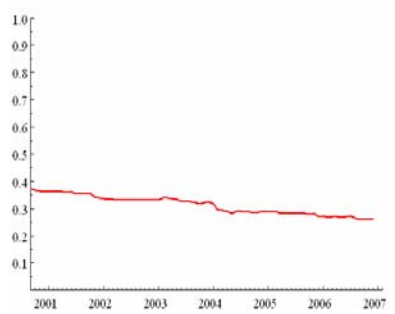
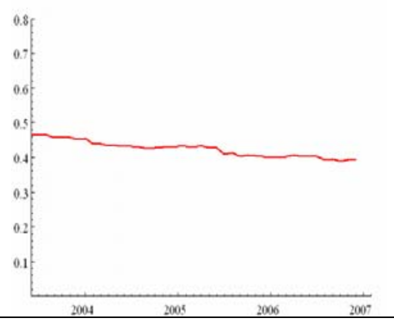
Table A3.36 Perron Unit Root Test: Singapore

Dependent Variable: LRUSDSNG				
	Coefficient	Std. Error	t-Statistic	Prob.
C	0.196657	0.005171	38.02983	0.0000
T	-0.000357	7.04E-05	-5.076547	0.0000
DL	0.108304	0.009266	11.68776	0.0000
Dependent Variable: SNG_RES				
	Coefficient	Std. Error	t-Statistic	Prob.
SNG_RES(-1)	0.966403	0.016376	59.01280	0.0000

Table A3.37 Perron Unit Root Test: Thailand

Dependent Variable: LRUSDTHB				
	Coefficient	Std. Error	t-Statistic	Prob.
C	1.473563	0.004043	364.4461	0.0000
T	-0.000223	5.50E-05	-4.060289	0.0001
DL	0.174220	0.007245	24.04540	0.0000
Dependent Variable: THB_RES				
	Coefficient	Std. Error	t-Statistic	Prob.
THB_RES(-1)	0.845715	0.042839	19.74171	0.0000
AR(2)	0.048313	0.074659	0.647116	0.5182

Figure A3.1 Recursive Analysis of Eigenvalues, 60 month window

Full Sample		
Europe	**Latin America**	**Asia**
		
Pre-Crisis		
Europe	**Latin America**	**Asia**
		(not applicable, no cointegration)
Post-Crisis		
Europe	**Latin America**	**Asia**
		

Appendix 2 Plots of Impulse Response Function Analysis: Adjustment of the Real Exchange Rate to an Unexpected Temporary Shock to Real Output

Figure A3.2 European Economies Pre Crisis: 1980Q1-1992Q2

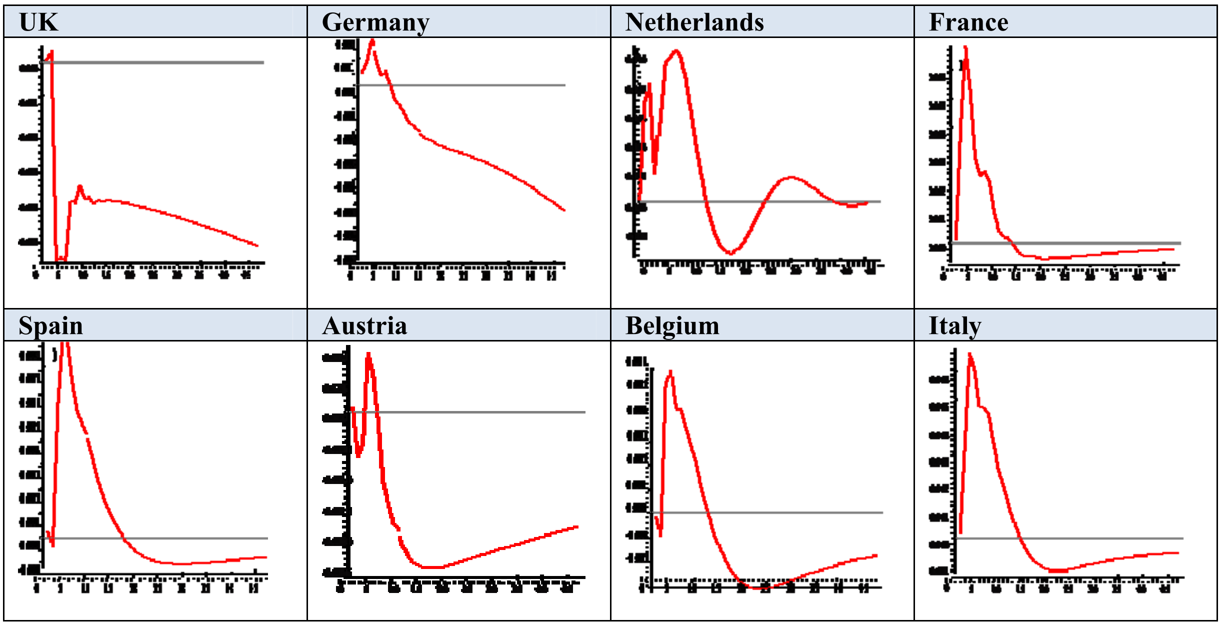


Figure A3.3 European Economies Post Crisis: 1993Q1-2006Q4

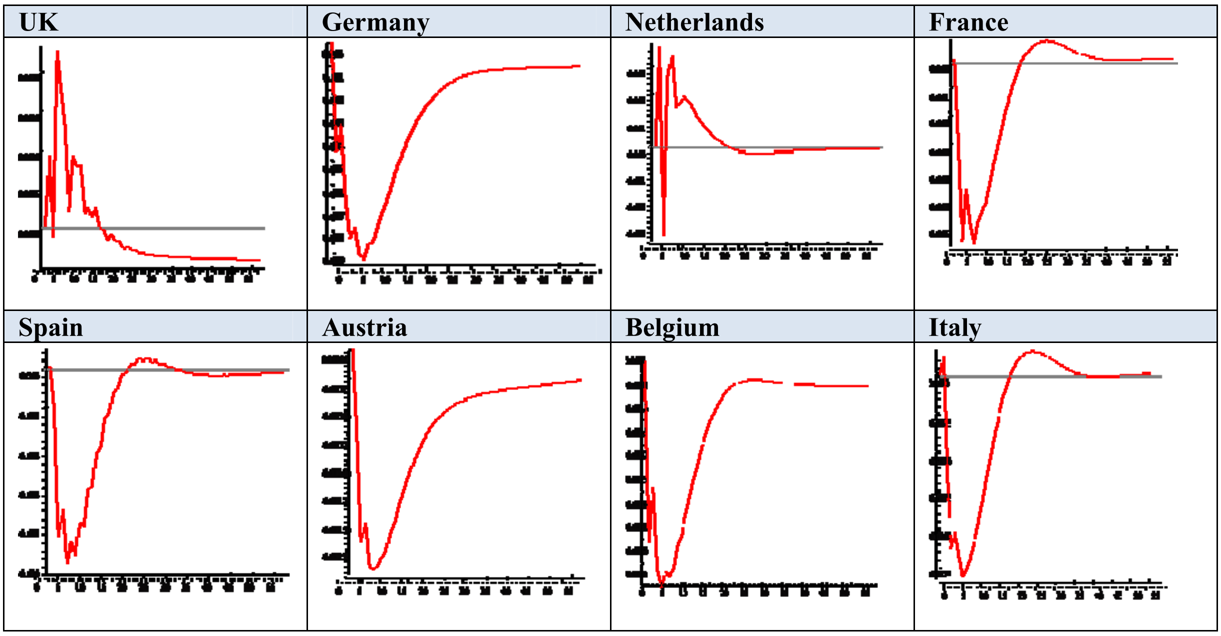


Figure A3.4 Latin American Economies Pre-Crisis: 1983Q1-1994Q3

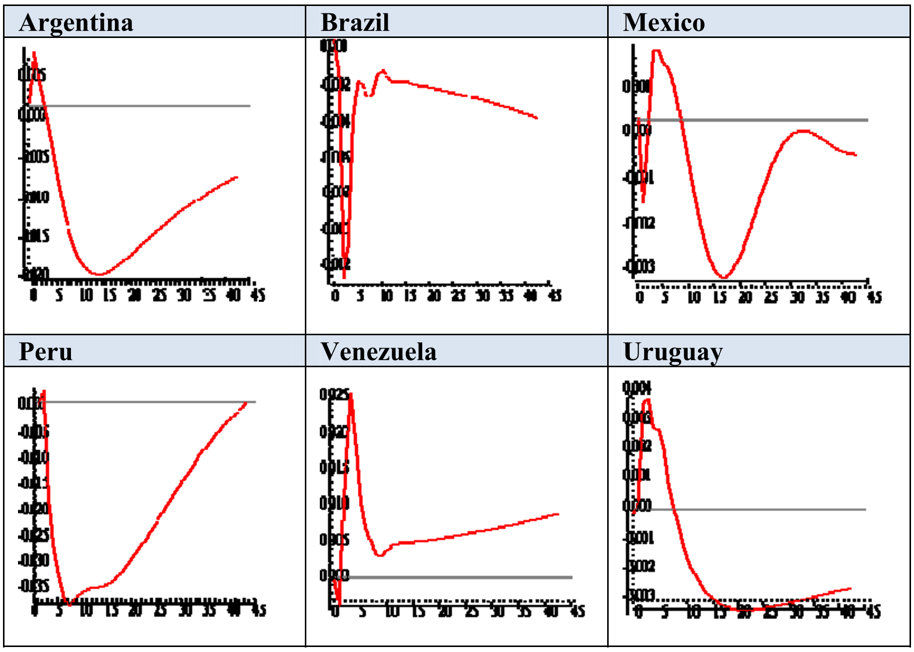


Figure A3.5 Latin American Economies Post-Crisis: 1995Q3-2006Q4

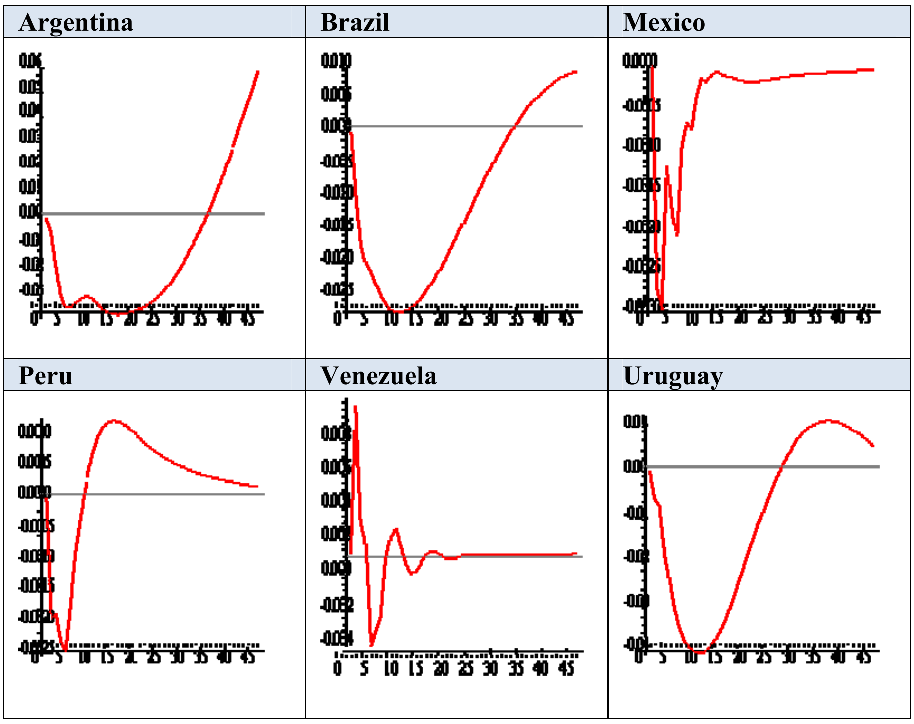


Figure A3.6 South East Asian Economies Pre-Crisis: 1988Q1-1997Q2

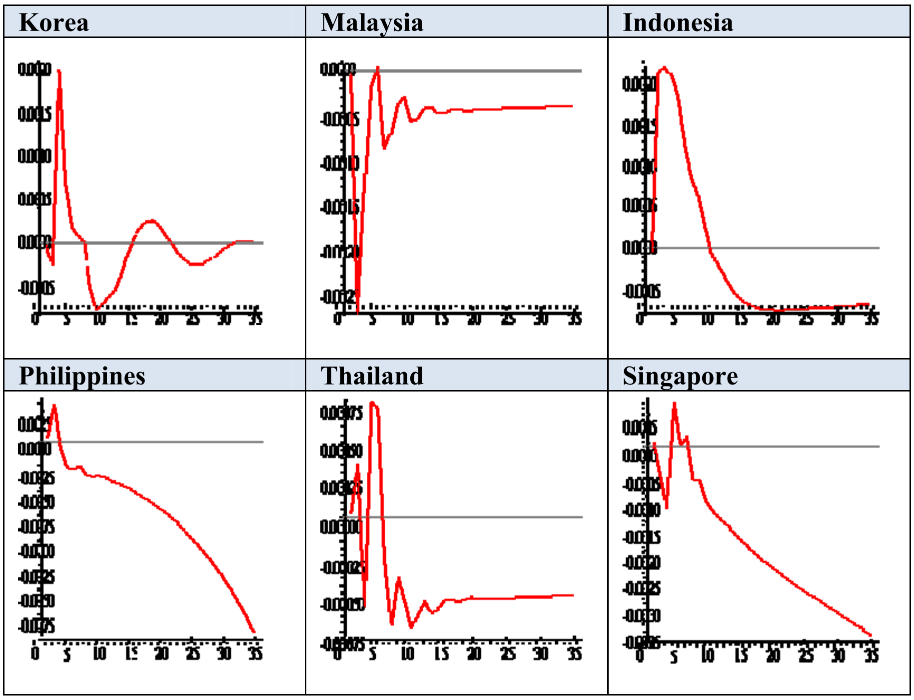
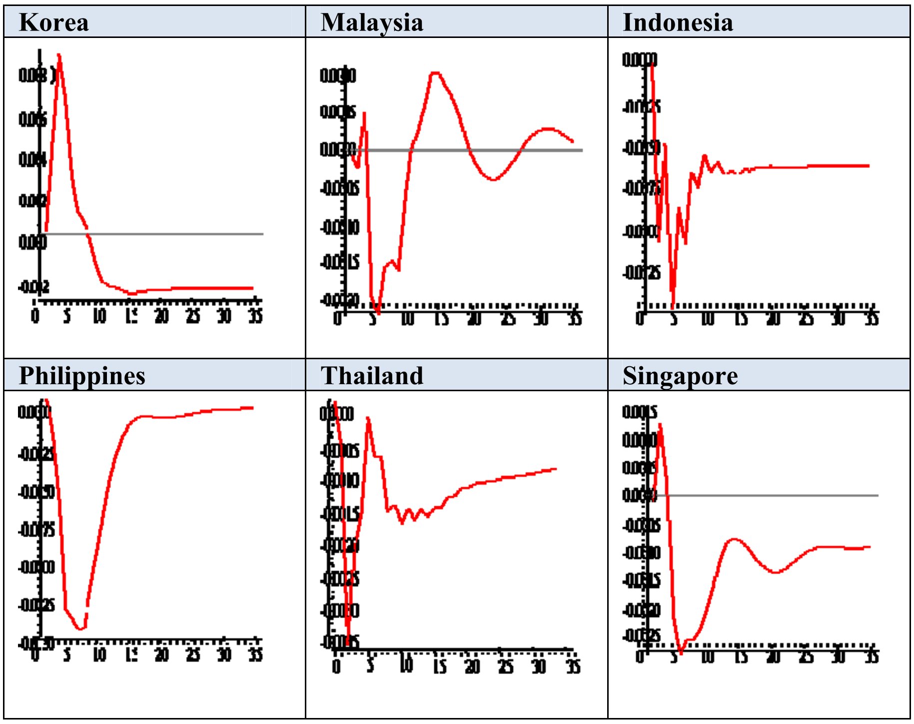


Figure A3.7 South East Asian Economies Post-Crisis: 1998Q3-2006Q4



CHAPTER FOUR

INTERDEPENDENCE BETWEEN PPP AND UIP: EVIDENCE FROM A SELECTION OF EU ECONOMIES IN A BIPOLAR AND TRIPOLAR MODELLING FRAMEWORK

I. Introduction

As noted by Juselius (1995), *"A common feature of most empirical results is the empirical rejection of standard economic hypotheses such as the prevalence of the purchasing power parity (PPP) as a backward adjustment mechanism in the goods market and the uncovered interest parity (UIP) as a forward-looking market clearing mechanism in the capital market"* (p.212).

While the empirical work undertaken in **Chapters 2** and **3** provide evidence that Purchasing Power Parity (PPP) holds in the long-run, it is important to point out that the PPP condition does not take account of international capital flows. The empirical results from **Chapters 2** and **3**, nonetheless, suggest that PPP holds in the long-run, perhaps suggesting that integration in the goods market in the economies considered dominated the long-run path of the exchange rate (as opposed to omitted factors associated more directly with financial markets and their integration). The time period considered also plays a role in the findings cited in previous chapters. In **Chapter 2**, the period of the 1990s was assessed for a sample of European countries, a period where one could argue that trade flows as opposed to financial flows were of key importance in determining the exchange rate. In **Chapter 3**, the sample is comprised primarily of transition economies, where financial flows may not have taken on as much economic importance as yet compared to trade flows.[83]

Clearly, in the context of the current global economy, one could argue that any modelling of the exchange rate must take into account not only trade flows, but also capital flows, certainly with respect to economies that have very liberalized capital accounts. This, of

[83] As is observed in Hunter and Isachenkova (2001), for models of company failure associated with Russia, financial variables such as liquidity have little impact relative to more direct trading related variables such as profits and sales turnover. In particular, economies often engaged in trade such as barter with such trade also operating across borders in a throw back to the old days of COMECON. Furthermore, as financial markets were not well developed there was some degree of credit rationing in Eastern Europe.

course, does not mean that PPP or Uncovered Interest Parity (UIP) can never be validated when tested separately. Rather, the proposition is that when PPP or UIP is found not to hold when each condition is tested individually, one reason could be that the cointegrating vector lacks a key variable (i.e. tests of PPP lack interest rates and tests of UIP lack price levels).

Specifically, given the increased level of both trade and financial integration in the current world economy, modelling the exchange rate should therefore take into account both the goods and capital markets. Thus, it follows that a case exists to suggest that PPP and UIP are implicitly linked. Moreover, from a theoretical point of view, we know that the balance of payments is equal to the sum of the current and capital accounts. If an economy's interest lies in balancing its payments, then exchange rate models should take account of both components of the balance of payments.

This chapter has its foundations in the works of Hunter (1992), Johansen and Juselius (1992), Juselius (1995), Juselius and MacDonald (2004a, 2004b), and MacDonald and Marsh (2004). These studies highlight the lack of empirical support for PPP and UIP when considered individually. Their work is based upon a joint modelling framework for PPP and UIP that shows that when considered jointly, these parity conditions can describe stationary relationships in the long-run. This, of course, underlines the strong linkages that exist between the goods market and the capital market. The rationale for such a linkage has been described by MacDonald (2000). He states that current account deficits are financed through foreign reserves or via the capital and financial account. Under floating exchange rates, the assumption is that the central bank does not engage in intervention, and, thus, there exists an implicit link between PPP and the capital and financial account.

Juselius (1995) examined the long-run foreign transmission effects in a cointegrated VAR comprising the prices, interest rates and exchange rates of Denmark and Germany, while Juselius and MacDonald (2004b) assessed international parity relationships between the USA and Japan. MacDonald and Marsh (2004) present a simultaneous exchange rate model of the US dollar, the German mark and the Japanese yen.

The first element of this chapter involves setting out a context for the rationale for examining PPP and UIP simultaneously. In addition, we provide a review of some of the key pieces of empirical work already undertaken. Following this, a bipolar model of PPP and UIP is set out using the case of Denmark and Germany. To this end, a well-specified five-variable cointegrated VAR model is constructed for the Danish Kroner/German DM exchange rate, the price indices of Denmark and Germany, and the interest rates of Denmark and Germany. This analysis represents an update of the original work of Juselius (1995) which considered similar issues for the same two countries. In the current analysis, the data is of a different frequency and a sequence of different restriction tests are applied to the alpha matrix parameters when compared with the article of Juselius. The model constructed lays the foundation for an extension to the tripolar case (whereby the United Kingdom is added to the Denmark-Germany bipolar model).

The tripolar model comprises Denmark, Germany and the United Kingdom. These economies are of interest in the context of the possible future direction of the euro area. The rationale for choosing the UK was that it represents a large economy outside of the euro area. Analysing simultaneously the exchange rates, prices, and interest rates of the UK in a system which also include Germany (a large economy in the euro area) and Denmark (a small economy outside the euro area) is of interest in terms of a possible expansion of the euro area. Following Camerero and Tamarit (1996), and Haldane and

Pradhan (1992), it is understood that the validation of PPP and UIP relates to two criteria for the long-run convergence of countries that could be deemed suitable for integration (PPP referring to goods market integration and UIP referring to financial market integration). The work undertaken contributes to the literature by providing a coherent extension to a bipolar model that provides economically interpretable results. The results are developed further by the extension associated with a tripolar model by comparing the cointegration tests and the long-run derived from models including short-term interest rates with a model that incorporates longer-term interest rates. Specifically, the results that support both PPP and UIP seem to be clearer in the case of a tripolar model when short-term interest rates enter the system.

The evidence seems to point towards the PPP restriction holding when long-term interest rates are used only when the UIP restriction is not imposed. If there was clear evidence that the yield curve holds, then the underlying stochastic trend associated with interest rates ought to be captured by both short and long rates. If the yield curve holds, then all interest rates have an exact error correction as per Davidson et al (1978) with a single alternate rate. Burke and Hunter (2008) suggest a similar result for sequences of N prices in a competitive economy. When the bond market is fully efficient, then a single common trend drives interest rates. Hence, each rate in turn is representative of the common trend and, as a result, findings for cointegration based on any of the rates ought to be the same. Findings on the yield curve have been mixed (Backus, 1984, and Campbell and Ammer, 1993), suggesting that the information content of long-rates might be different. The recent inversion of the yield curve associated with the credit crunch would be suggestive of this and might provide evidence as to why interest rate parity might not hold when longer rates are used in the analysis. Hence, long-term rates may make a less appropriate proxy for international capital flows than short-term rates.

Section IIA.3 in Chapter 2 has presented in detail the theoretical considerations linked with the joint modelling of PPP and UIP. Regarding previous empirical studies, a number of papers find PPP and UIP do not hold when considered separately, but do hold when these conditions are combined. The first such relation was observed by Johansen and Juselius (1992), who considered the case of the UK effective exchange rate over the period 1972 to 1987. PPP was found not to hold when tested in isolation, but a stationary outcome was found when PPP was combined with two interest rates. However, in Johansen and Juselius (1992), the restriction on the interest rate was not rejected. A similar long-run equation for the exchange rate was found by Hunter (1992) along with a second long-run equation that satisfies UIP. Other studies finding similar outcomes for different countries include Sjoo (1995), Helg and Serati (1996), Pesaran et al (2000), Caporale et al (2001), Bjornland and Hunges (2002), and Bevilacqua (2006).

The empirical analysis in the following sections of this chapter is based on the exchange rate, interest rate and price dynamic across three economies: Denmark, Germany and the United Kingdom. Firstly, a bipolar model for Denmark and Germany is constructed using a five-variable VAR. This model is then systematically augmented by adding the appropriate variables for the United Kingdom, so that the model is based on an eight-variable VAR. The seminal piece of literature that tests for UIP and PPP in a system involving Denmark and Germany was Juselius (1995). In this work, a 5-variable system was constructed comprising Danish and German prices, exchange rates, and interest rates (bond rates) over the period 1972 to 1991. By testing for PPP and UIP separately, there exists weak evidence to support stationarity. Consideration of the parities jointly, however, yields results that appear to be much more stationary. Specifically, Juselius (1995) notes that *"Deviations from the PPP were found to be important for the long-run determination of the exchange rates and the interaction between the goods and capital* *markets was found to be crucial for a full understanding of the movements of interest rates, prices, and exchange rates"* (p.236). In the following section we start our analysis by re-appraising data associated with the original study by Juselius (1995).

II. A Bipolar Model Examining PPP and UIP in Denmark and Germany

In this section, a five-variable bipolar model comprising the prices, interest rates, and exchange rate between Denmark and Germany is constructed in order to examine the validity of PPP and UIP and confirm whether the relations first uncovered by Juselius (1995) still hold. It is considered that a study relating Germany and Denmark is useful in two key respects. Firstly, it provides both a comparison with and a check on the robustness of the analysis in the article by Juselius (1995). Secondly, it is useful in terms of the composition of the EU to examine the inter-relationships between two developed trading partner economies, one of which is a large economy in the euro area and the other a smaller economy outside of the euro area.

Data has been collected for Denmark and Germany over the period 1984M06 to 2007M03 for the following variables: Danish krona/German DM, Danish producer prices, Danish 10-year Government bond yield, German producer prices, German 10-year Government bond yield. The choice of start date was governed by data availability. Specifically, a coherent German exchange rate series is only available from 1984M06. Long-term interest rates are used in the bipolar study for direct comparison with the work of Juselius (1995). In the tripolar work (Sections III and IV), separate analyses are carried out for long and short term interest rates, and a rationale is provided as to why systems comprising long rates give rise to different results to those comprising short rates.

The majority of the data was sourced from the IMF's International Financial Statistics database. This excludes German producer prices which were sourced from the Deutsche Bundesbank, Danish Treasury Bill rates that came from the Danmarks Nationalbank, and the exchange rate series sourced from Tradermade International Limited.[84] All variables have been transformed into logarithmic form and to more appropriately scale the problem, the interest rate series have been divided by 100.

The methodology to be employed follows the Johansen procedure best described in Johansen (1995) and adopted by Juselius (1995).[85] As a first step, the variables are inspected visually (in logarithmic form) via plots of their respective autocorrelation functions (ACFs) and partial autocorrelation functions (PACFs). As already stated, the first element of this empirical analysis is to re-assess the type of five-variable VAR analysed by Juselius (1995) on a data set that includes more recent observations. To this end, the following series have been collected at the monthly frequency over the period 1987M03 to 2007M03: Danish PPI, German PPI, DEMDKK exchange rate (where Denmark is the domestic currency), and the bond yields in Denmark and Germany.[86] Observation of the ACFs and PACFs, as well as formal ADF unit root tests, reveal that all of the variables are integrated of order 1. In terms of a further stage of preliminary analysis, the scaled residuals[87] are used to detect periods of excessive fluctuation or volatility.

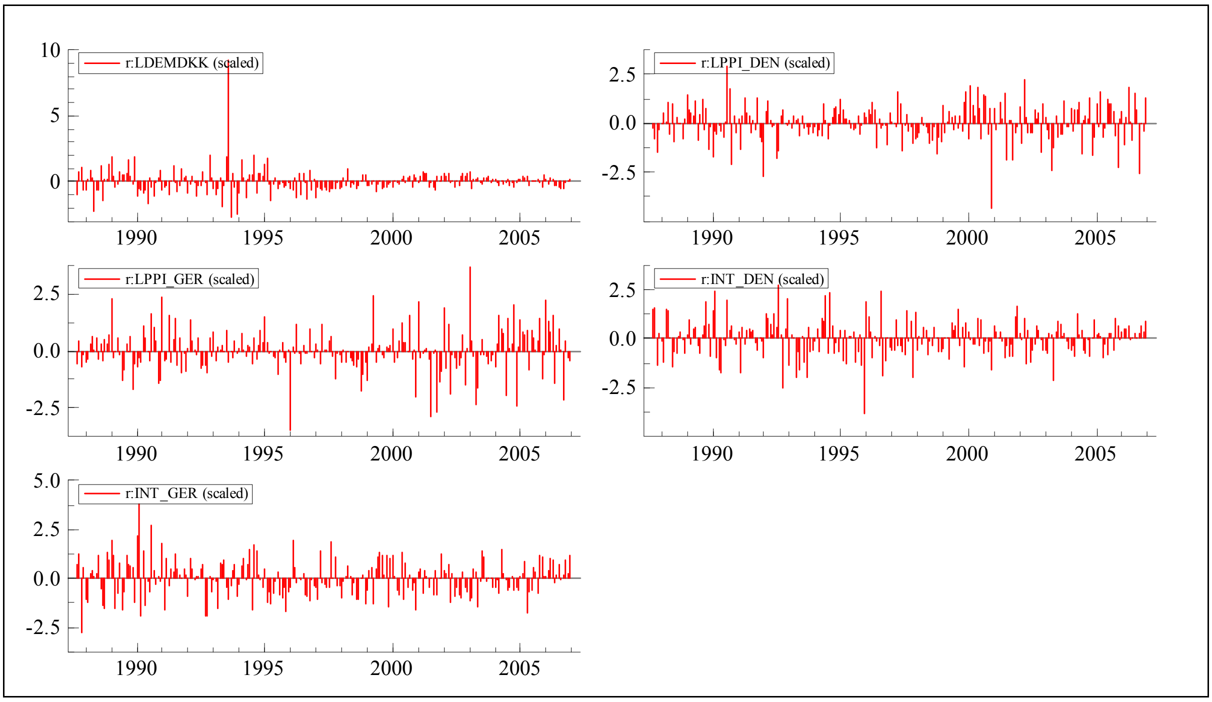
[84] Tradermade International Limited is a specialist firm that produces data for leading international financial indicators. The German exchange rate data was sourced specifically from the acaTrader product. This product produces exchange rate data for the current eurozone members in terms of their pre-EMU currencies, using the rate at which the DM was converted to the euro and the euro/dollar rate.

[85] This methodology has been described in detail in Chapters 2 and 3.

[86] Recall that Juselius (1995) used quarterly data over the period 1972:1 to 1991:4, where prices were represented by the CPI for Denmark and Germany. In this re-examination, a consistent price series is only available for PPI.

[87] Scaled residuals are used to circumvent the well-known problem associated with using unscaled residuals, i.e. that the values of these residuals are dependent upon scale and the unit used. Moreover, there is no cut-off point for a "large" residual. The scaled residuals are computed by dividing the residual by the standard error of the residual.

Figure 4.1 **Scaled Residuals of the Series**



By analysing the graphs in **Figure 4.1** the following observations on the scaled residuals exceed 3: 1990M02 (INT_GER), 1993M08 (DEMDKK), 1995M12 (INT_DEN), 1996M01 (PPI_GER), 2000M12 (PPI_DEN), 2003M01 (PPI_GER). To determine whether dummy variables used to account for these extreme observations a prior regression including the dummies detected by observation is considered.[88]

[88] A similar type of approach was undertaken by Bevilacqua (2006).

Table 4.1 **t-statistics of the Dummy Variables (dependent variables in logged first differences)**

	D199002	D199308	D199512	D199601	D200012	D200301
E	0.58	**5.71**	0.79	0.75	-0.53	-0.76
p_{DEN}	-0.77	-0.79	-0.31	-0.28	0.92	0.94
p_{GER}	-1.03	-0.35	-0.03	-0.22	0.48	0.73
i_{DEN}	1.59	-0.02	-0.42	-0.22	-0.39	-1.01
i_{GER}	1.47	0.34	-0.04	-0.17	-0.25	-1.04

The only significant dummy appears to be for D1993M08, and therefore the analysis proceeds factoring this dummy into the analysis. This period coincides with a sharp strengthening of the DM vis-á-vis the Danish krone, following a well documented period of speculative pressure on the krone that arose as a result of the fall-out, which followed from the ERM crisis. As a result, the krone was devalued at this point in 1993.

Table 4.2 **Unrestricted VAR(7) Misspecification Test Results**

	Autocorrelation AR 1-7: F(7,168)	Normality χ^2 (2)	Skewness	Excess Kurtosis	ARCH 1-7: F(7,161)	Heteroskedasticity F(70,104)
E	0.683 [0.686]	16.826 [0.000]*	0.581	5.213	2.778 [0.794]	0.807 [0.831]
p_{DEN}	1.529 [0.161]	2.356 [0.311]	-0.372	4.685	0.606 [0.750]	0.615 [0.985]
p_{GER}	1.394 [0.211]	24.611 [0.023]**	-0.682	7.662	0.825 [0.568]	0.554 [0.996]
i_{DEN}	0.695 [0.676]	18.265 [0.000]*	0.412	5.725	0.851 [0.547]	0.495 [0.999]
i_{GER}	1.306 [0.250]	3.463 [0.177]	-0.198	3.258	2.109 [0.445]	1.064 [0.388]

Notes: * and ** refer to the 1% and 5% levels of significance respectively. Residual diagnostics on the VAR indicate some signs of autocorrelation at the 5% level: AR 1-7: F(175,679) distribution with p-value of 0.012, and evidence of system-wide non-normality.

Excepting the problem of excess kurtosis, it is apparent from the statistics presented in **Table 4.2** that the system seems to be well formulated. It is important to consider the likely implications of this type of deviation from the normality assumption. Following Juselius and MacDonald (2002), it would appear that these estimates of and tests for cointegration devised from the VAR are more sensitive to deviations from normality that arise from skewness than excess kurtosis. If as is suggested from the results in **Table 4.2**, the non-normality resulted from excess kurtosis, then it follows from the simulations derived in Gonzalo (1994) that in large samples, the Johansen trace test is robust to non-normality.[89] Given that the specification tests in **Table 4.2** suggest that there is no serial correlation or ARCH, then it would appear that k=7 represents an appropriate lag structure for the data generating process, which corresponds well with the model based on 3 lags used with quarterly data by Juselius (1995). Unlike Juselius (1995), who analysed a period that included more volatile price inflation, there appears to be no univariate evidence that any of the series are I(2). Thus, the analysis proceeds by testing for the cointegrating rank to estimate the number of cointegrating relationships in the system.

Table 4.3 Cointegration Rank Test

Rank	**Trace Test (T-nm)**	**Max Eigenvalue Test (T-nm)**
0	80.68 [0.004]**	34.32 [0.040]*
1	46.36 [0.067]	26.93 [0.058]
2	19.43 [0.473]	16.18 [0.223]
3	3.25 [0.947]	2.75 [0.952]

The sequence of LR test statistics adjusted for sample size indicates that there exists one cointegrating vector in the system. A coherent sequence of tests only arises when the trace test is applied (Johansen, 1995). One compares the case with rank zero or the likelihood derived for the model without cointegration with the alternative that there is one or more cointegrating vectors that arises by cumulating the trace tests for the alternative, r=5 to 1. If the test is significant, then the null is rejected (r=0) and the

[89] Paruolo (1995) provides further evidence that non-normality driven mainly by kurtosis is less of a problem than that induced by skewness.

alternative (r≥1) cannot be rejected. As can be seen from the second row in **Table 4.3**, the null is accepted that (r=1), implying that no further testing is required. It can be observed that the subsequent tests also accept the null implying no higher value for r is appropriate. It is shown in Johansen (1995) that the trace test only has optimal size under this ordering and there is no equivalent optimality result for the maximum eigenvalue test, which leads one to prefer the trace test.

Prior to undertaking the tests of PPP and UIP, a range of weak exogeneity tests are carried out on the unrestricted vector in order to provide a natural means by which to condition the equations. These are carried out on an individual basis in the first instance, followed by a selection of joint tests. The results of these tests are set out in **Table 4.4** below:

Table 4.4 Weak Exogeneity Tests on the Unrestricted Vector

Test on α	**Chi-squared Statistic**
$\alpha_e=0$	4.965 (0.030)
$\alpha_p=0$	0.313 (0.576)
$\alpha_{p*}=0$	0.135 (0.713)
$\alpha_i=0$	4.895 (0.027)
$\alpha_{i*}=0$	0.243 (0.622)
$\alpha_p=0$ and $\alpha_{p*}=0$	0.386 (0.825)
$\alpha_{p*}=0$ and $\alpha_{i*}=0$	0.578 (0.749)
$\alpha_e=0$, $\alpha_p=0$, and $\alpha_{p*}=0$	6.369 (0.095)

These results indicate that weak exogeneity is not evident for the exchange rate or the domestic interest rate. Thus there are two possible appropriate normalisation schemes; using either the exchange rate or the home interest rate. In order for the coefficients to be interpreted in terms of an elasticity with respect to the exchange rate, the latter is used to condition the equations. A normalisation of the vector on the exchange rate is represented in **Table 4.5** below. It should be noted that in the r=1 case the normalisation is sufficient to identify the cointegrating vector.

Table 4.5 Long-Run Cointegrating vector normalised on DEMDKK

	β
e	**1.000000**
$\mathbf{p_{DEN}}$	-0.450970
$\mathbf{p_{GER}}$	0.141930
$\mathbf{i_{DEN}}$	-0.014157
$\mathbf{i_{GER}}$	0.011627

From the normalised eigenvectors in **Table 4.5**, it is encouraging to note that the coefficients on the price series are correctly signed (i.e. negative for the domestic price level and positive for the foreign price level). The signs and magnitudes on these eigenvectors are coherent in the sense that the representation meets the signing appropriate for both weak form PPP and UIP to hold simultaneously. This vector can be tested formally for PPP and UIP by imposing restrictions on the coefficients. These restrictions essentially test the hypothesis that the unrestricted estimates are not significantly different from the imposed restricted values.

Table 4.6 Tests of Restrictions for PPP and UIP

	e	p_{DEN}	p_{GER}	i_{DEN}	i_{GER}	χ^2 (df)	p-value
Weak form PPP only							
H_1	1	$-\beta_{21}$	β_{21}	0	0	13.987 (3)	0.003
Strong form PPP only							
H2	1	-1	1	0	0	30.250 (4)	0.000
UIP only							
H3	1	0	0	β_{41}	$-\beta_{41}$	5.106 (3)	**0.164**
Weak form PPP plus UIP							
H4	1	β_{21}	$-\beta_{21}$	β_{41}	$-\beta_{41}$	3.938 (2)	**0.140**
Strong form PPP plus UIP							
H5	1	-1	1	β_{41}	$-\beta_{41}$	5.436 (3)	**0.124**
Test of Normalisation							
H6	0	β_{21}	β_{31}	β_{41}	β_{51}	5.907 (1)	0.015
Pure Interest Rate Parity							
H7	0	0	0	1	-1	14.380 (4)	0.006
Strong form PPP plus Pure Interest Rate Parity							
H8	1	-1	1	1	-1	14.604 (4)	0.006
Weak form PPP plus Pure Interest Rate Parity							
H9	1	β_{21}	$-\beta_{21}$	1	-1	3.852 (3)	**0.278**
Weak form PPP plus Unrestricted Interest Rates							
H10	1	β_{21}	$-\beta_{21}$	γ	η	1.377 (1)	**0.241**
Strong form PPP plus Unrestricted Interest Rates							
H11	1	-1	1	γ	η	2.982 (2)	**0.228**

The tests of restrictions reveal some interesting results. There is clear evidence of a failure of PPP, in either weak form (H1) or strong form (H2), when it is considered in isolation. A joint test of PPP and UIP, however, indicates non-rejection of the restrictions, thereby providing support for the rationale that a stationary long-run exchange rate relation requires factors that explain the influence of both the goods market and capital market. When tested in conjunction with UIP, both the stronger and weaker versions of PPP appear to hold. As well as these tests, H3 tests for UIP alone.

Surprisingly, given the results of Juselius (1995)[90], this hypothesis is not rejected and may be indicative of a greater relative weight attached to the importance of capital flows as opposed to trade flows in the determination of the German/Danish exchange rate over the period 1987M3 to 2007M03. It is now worthwhile considering whether the equations estimated can be empirically identified based on the selected normalisation. This is done by setting the coefficient on the exchange rate to zero, we can observe from the test that the restriction is not accepted, i.e. H6. As a result, potential problems associated with incorrect normalisation described in Hunter and Simpson (1995) are not relevant in this case. Following, Hunter (1992) it is also of interest to examine whether the long-run is simply defined as being dependent on terms in the interest rates or whether we have a pure UIP correction term that is driving exchange rate returns. This can be observed by setting all the price and exchange rate coefficients to zero. It is found, however, that pure interest rate parity in isolation is not accepted (i.e. H7). Pure interest rate parity, in conjunction with a PPP relationship (in weak form only), however, appears to satisfy the relevant joint restrictions (H9). These results suggest that, for the economies under investigation, the deviations from the PPP and UIP parities when considered jointly are stationary and interaction of the goods and capital markets must be taken into account in models of exchange rate determination pertaining to the economies in question.[91]

These results permit estimation of the short-run drivers of stationary relationships across the international parity conditions under consideration. By imposing restrictions on the α matrix, a test of the significance that determines whether the parameter associated with a specific element in this vector is zero can provide information on the short-run dynamic. More pertinently this is a test for weak exogeneity (Johansen, 1992). Unlike **Table 4.4**,

[90] A similar equation was observed by Hunter and Simpson (1995) using UK data on the effective exchange rate and this is reported in Chapter 4 and 5 of Burke and Hunter (2005).
[91] Whilst not a joint parity relationship, PPP in conjunction with unrestricted interest rates also appears to yield a stationary outcome.

however, where the tests were carried out on the unrestricted vector, in this case the response to deviations from the parity conditions is examined for cases where the joint PPP-UIP hypothesis is valid. Specifically, the focus is on the response to (i) the joint Weak-form PPP & UIP combination (**Table 4.7**) and (ii) the joint Strong-form PPP & UIP combination (**Table 4.8**). This is only applicable when the restricted cointegrating relation is stationary and the likelihood ratio test of the β restrictions has not been rejected. This is similar to the type of approach previously undertaken by Akram (2006).[92]

Table 4.7 Short-Run Dynamics: Weak form PPP plus UIP

Weak form PPP plus UIP	α Matrix					Test Statistic
	e	p	p*	i	i*	$\chi^2(3)$
Unrestricted	0.0122 (0.0053)	-0.0038 (0.0017)	-0.0063 (0.0047)	0.0354 (0.0126)	-0.0089 (0.0082)	
e fixed	0	-0.0290 (0.0276)	-0.0201 (0.0183)	0.0141 (0.0041)	-0.0098 (0.0267)	6.8523 (0.0768)**
p fixed	0.0095 (0.0037)	0	-0.0038 (0.0035)	0.0241 (0.0089)	-0.0064 (0.0057)	4.1456 (0.2462)
p* fixed	0.0141 (0.0058)	0.0930 (0.0083)	0	0.0405 (0.0138)	-0.0074 (0.0089)	5.3436 (0.1483)
i fixed	-0.0381 (0.0111)	-0.0035 (0.0183)	0.0135 (0.0120)	0	0.0408 (0.0138)	8.5490 (0.0359)*
i* fixed	0.0177 (0.0098)	-0.0074 (0.0157)	-0.0091 (0.0103)	0.0844 (0.0181)	0	5.1235 (0.1630)

Notes: * indicates rejection of the null at the 5% level or below; ** indicates rejection of the null at the 10% level; standard errors are denoted in brackets, while the p-values associated with the chi-squared test statistics are denoted in squared brackets. The joint restrictions imposed are: β_{11}=1, β(1,2)=-β(1,3), β(1,4)=-β(1,5); β(1,1)=1, β(1,2)=-β(1,3), β(1,4)=-β(1,5), α(1,1)=0; β(1,1)=1, β(1,2)=-β(1,3), β(1,4)=-β(1,5), α(2,1)=0; β(1,1)=1, β(1,2)=-β(1,3), β(1,4)=-β(1,5), α(3,1)=0; β(1,1)=1, β(1,2)=-β(1,3), β(1,4)=-β(1,5), α(4,1)=0; β(1,1)=1, β(1,2)=-β(1,3), β(1,4)=-β(1,5), α(5,1)=0. The 'Unrestricted' row refers to estimates of short-run adjustment with restrictions in place for the β matrix only, i.e. α is unrestricted in this row. All other rows test jointly the β and relevant α restrictions.

[92] As well as providing information on the extent of the response of variables in the system to correct for deviations to the long-run equilibrium, this approach also permits an additional check on the normalisation scheme.

The unrestricted estimates of the adjustment coefficients are provided in the first output row, indicating in this case the extent to which each of the equations in the system respond to eliminate deviations from weak form PPP and UIP. From a statistical significance perspective, the exchange rate, domestic prices and the domestic interest rate reveal the strongest response to disequilibrium. The size of the coefficients indicates the extent to which error correcting or as David Hendry (1995) terms equilibrium correcting adjustment occurs in these equations. A greater weight of importance is attached to the response in the domestic interest rate equation that adjusts at the rate of 3.5% per month to restore equilibrium, compared to 1.2% per month for the exchange rate and 0.4% per month for domestic prices. In the case where these coefficients are zero, then the variable is weakly exogenous for the single cointegrating vector. There is no equilibrium correction, so the level variables describe more subtle random walk behaviour or following Juselius (2006) they are in essence the stochastic trends.

These findings are substantiated through a series of hypothesis tests where the null hypothesis is that each variable in the system does not respond to restore the weak form PPP and UIP equilibrium. At the 10% level of significance or below, this null is not rejected for all cases except the exchange rate and the domestic interest rate.

These results indicate that the main driver of the movement towards the joint stationary equilibrium of weak form PPP and UIP is domestic interest rates. This is an interesting finding, and one which is comparable to that found by Juselius (1995), who noted that in relation to one of the two cointegrating vectors identified in her study that *"the short-run adjustment [to this relation] takes place primarily in the Danish interest rate equation"* (p.235).

Table 4.8 Short-Run Dynamics: Strong form PPP plus UIP

Strong form PPP plus UIP	α Matrix					Test Statistic
	e	p	p*	i	i*	$\chi^2(4)$
Unrestricted	0.0108 (0.0046)	-0.0030 (0.0014)	-0.0054 (0.0029)	0.0296 (0.0108)	-0.0079 (0.0070)	
e fixed	0	-0.0069 (0.0080)	-0.0087 (0.0052)	0.0313 (0.0118)	-0.0823 (0.0769)	10.927 (0.0274)*
p fixed	0.0111 (0.0045)	0	-0.0044 (0.0043)	0.0297 (0.0108)	-0.0074 (0.0069)	4.1546 (0.3855)
p* fixed	0.0106 (0.0041)	0.0005 (0.0059)	0	0.0273 (0.0098)	-0.0057 (0.0062)	5.3720 (0.2512)
i fixed	0.0114 (0.0050)	-0.0043 (0.0082)	-0.0070 (0.0054)	0	-0.0218 (0.0061)	13.242 (0.0102)*
i* fixed	0.0103 (0.0044)	-0.0016 (0.0070)	-0.0041 (0.0046)	0.0356 (0.0082)	0	5.5395 (0.2363)

Notes: * indicates rejection of the null at the 5% level or below; standard errors are denoted in brackets, while the p-values associated with the chi-squared test statistics are denoted in squared brackets. The joint restrictions imposed are: β(1,1)=1, β(1,2)=-1, β(1,3)=1, β(1,4)=-β(1,5); β(1,1)=1, β(1,2)=-1, β(1,3)=1, β(1,4)=-β(1,5), α(1,1)=0; β(1,1)=1, β(1,2)=-1, β(1,3)=1, β(1,4)=-β(1,5), α(2,1)=0; β(1,1)=1, β(1,2)=-1, β(1,3)=1, β(1,4)=-β(1,5), α(3,1)=0; β(1,1)=1, β(1,2)=-1, β(1,3)=1, β(1,4)=-β(1,5), α(4,1)=0; β(1,1)=1, β(1,2)=-1, β(1,3)=1, β(1,4)=-β(1,5), α(5,1)=0. The 'Unrestricted' row refers to estimates of short-run adjustment with restrictions in place for the β matrix only, i.e. α is unrestricted in this row. All other rows test jointly the β and relevant α restrictions.

In this case, unlike the joint Weak form PPP & UIP case, only the adjustment coefficients for the exchange rate and the domestic interest rate are significant. However, the domestic interest rate responds to restore equilibrium at almost three times the rate of the exchange rate. Thus, with regard to the case where long-run strong form PPP and UIP restrictions were found to hold, the driver of the movement towards the joint equilibrium in the short-run appears to be either the exchange rate or domestic interest rates. The greater relative weight attached to the domestic interest rate in driving the convergence to the stationary state is confirmed by the hypothesis tests undertaken. Whilst the null (that the response of each variable in the system is zero) is not rejected for all cases except the exchange rate and the domestic interest, the level of significance associated with the latter outweighs the former.

Across the scenarios where long-run restrictions were imposed to test for PPP and UIP jointly, the econometric evidence would appear to suggest that the domestic interest rate is the primary mechanism for adjustment to the equilibrium level in the short-run. Across the analyses undertaken, there would appear to be signs of weak exogeneity associated with German interest rates, and possibly also German prices (albeit to a much lesser extent). This may reflect the influential role played by German interest rates in international markets in the period prior to the introduction of the euro.

Overall, the results from the bipolar model across the prices, interest rates and exchange rate between Denmark and Germany are interesting in light of the work of Juselius (1995). Using an updated data set, and a different data frequency than that used by Juselius (1995), it is shown that the consideration of PPP and UIP in a joint cointegration framework finds validity in the parity conditions in the long run. A similar finding was reported by Juselius (op. cit.), although she encountered some I(2) behaviour in the price series. As well as finding PPP and UIP to hold, the evidence appears to suggest that the driver of the short-run adjustment to the equilibrium in the system is Danish interest rates, and this also reflects an element of the findings of Juselius (1995).[93]

Sections III and IV below augment this bipolar system to the tripolar case. Two particular scenarios are addressed, the first based on a system using long-term interest

[93] In the period after this work was undertaken, a new book has been published by Juselius (2006, Chapter 21), wherein an analysis is carried out on foreign transmission effects between Denmark and Germany using quarterly data, and the CPI as the price level indicator, over the period 1973 to 2003. My work uses a different time period, a different data frequency, and a different price level indicator. My work also tests alternative restrictions on the eigenvectors. It is encouraging to find similar results across both pieces of empirical work, however, which helps to indicate robustness in the findings (namely, that the Danish exchange rate and interest rate are largely determined by the international level, as well as the equilibrating effect of Danish interest rates on the joint PPP-UIP relation).

rates and the second based on short-term interest rates. To ensure a degree of comparison with the Juselius (op. cit.) work, the bipolar model estimated above focuses on long-term rates. The issue of whether different types of interest rate variables yield different results, as regards the validation of PPP and UIP, is examined in the context of the tripolar models. The United Kingdom is added to the bipolar model of Denmark and Germany in order to illustrate the exchange rate, price, and interest rate dynamic across three EU economies, two of which remain outside the euro area. Thus, instead of assessing PPP and UIP in a joint framework across a five variable system, this augmented system utilises an eight variable system.

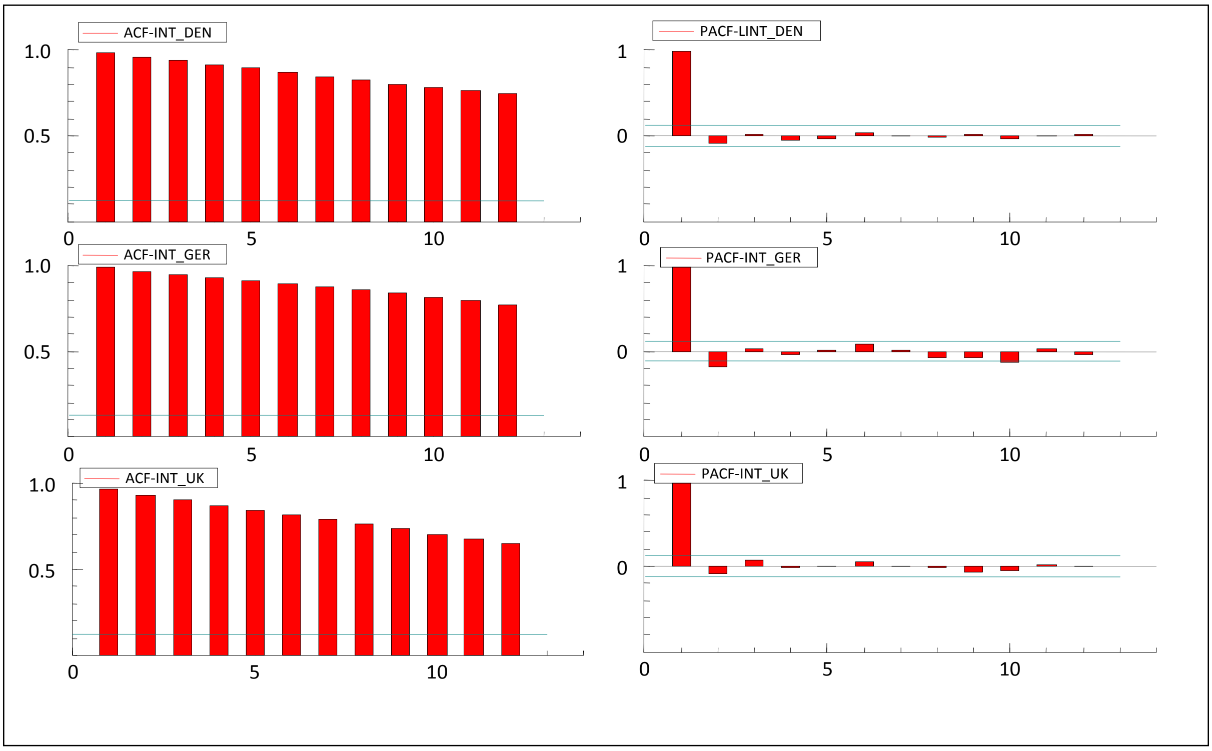
III. Extension to the Tripolar Case: Augmenting the System with the United Kingdom – The Long-term Interest Rate Case

This section augments the bipolar system of Denmark and Germany to the tripolar case, where the system is augmented to include the United Kingdom. In this case, an eight variable system is constructed to incorporate the prices, interest rates and exchange rates across Denmark, Germany and the United Kingdom. The aim is to examine the PPP and UIP relationships in a joint framework comprising a greater range of dynamics than that of the bipolar model. Here, two DM rates are used, namely the £/DM exchange rate, and the Danish krone/DM exchange rate. The price series are given by the respective PPIs of Denmark, Germany and the UK, while 10-year government bond yields represent the measure used for long-term interest rates for each country. Data availability (at the time of writing) permits the analysis to run from mid-1984 to the end of 2006, and the frequency is monthly. As in the case of the bipolar model, all series have been transformed into logarithmic form and the interest rate series have been divided by 100.

The econometric analysis to be pursued, whilst based on the Johansen multivariate cointegration procedure, is consistent with the General-to-Specific methodology (Hendry, 1995). Clearly, the parity conditions described in equations (1) to (4) set out the theoretical equilibrium relationships that would exist in a simplified world. For this purpose, the General-to-Specific approach is adopted to account for various stochastic properties of the data. To this end, an unrestricted VAR is constructed and this represents a reasonable reflection of the joint distribution of the observed series.[94]

The vector under examination can be represented as $x' = (e_{den}, e_{uk}, p_{den}, p_{uk}, p_{ger}, i_{den}, i_{uk}, i_{ger})$. As a starting point, the time series properties of the data are considered on observation of the ACFs and PACFs of each of the series.

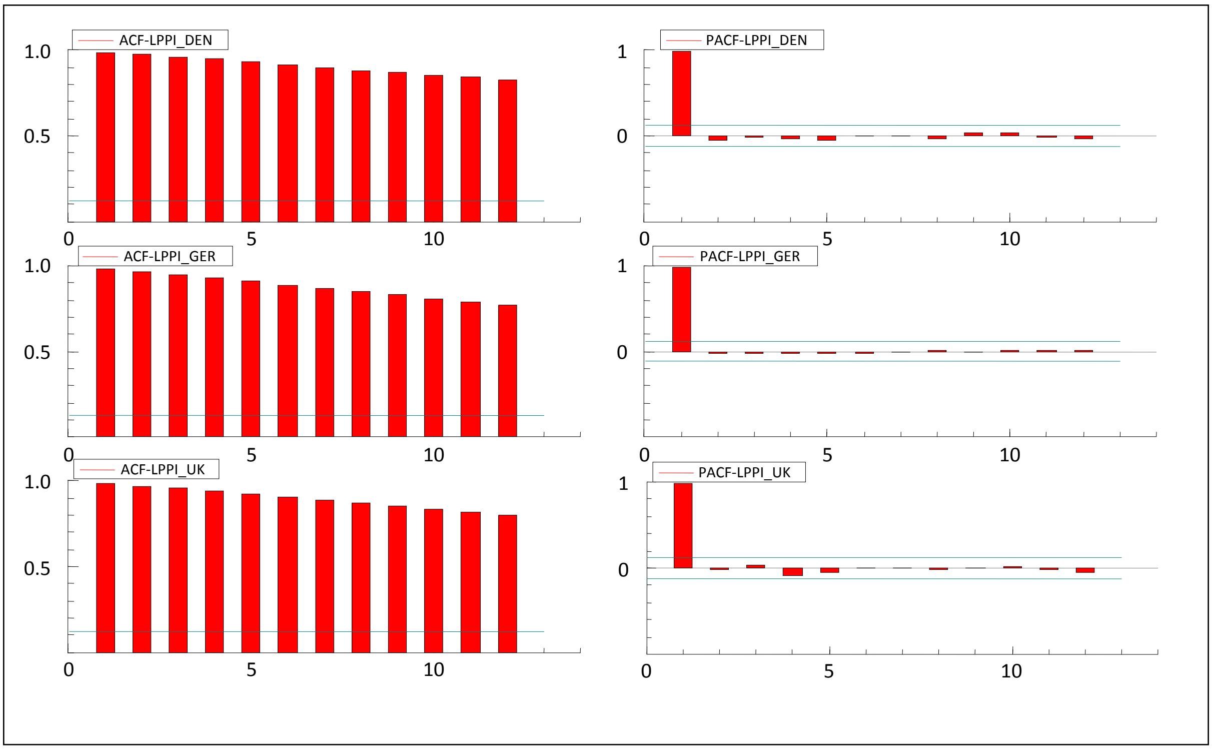
Figure 4.2 ACF and PACF – Long-term Interest Rates, 1984M06-2006M12



[94] For more detail on this, see Hendry and Mizon (1993).

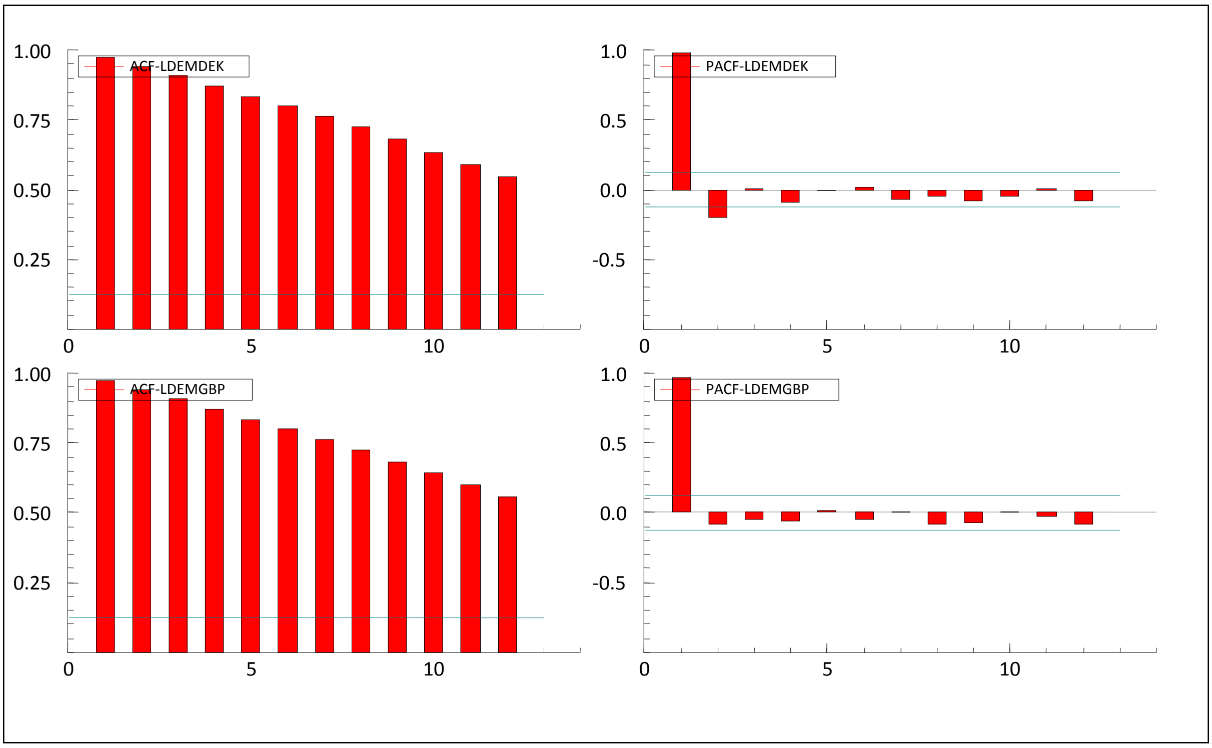
The evidence from the 10-year Government bond series for Denmark, Germany and the United Kingdom appears to suggest that differencing once induces stationarity, i.e. the series are I(1).

Figure 4.3 ACF and PACF – Producer Price Index (log), 1984M06-2006M12



As in the case of long-term interest rates, producer prices would appear to be I(1)

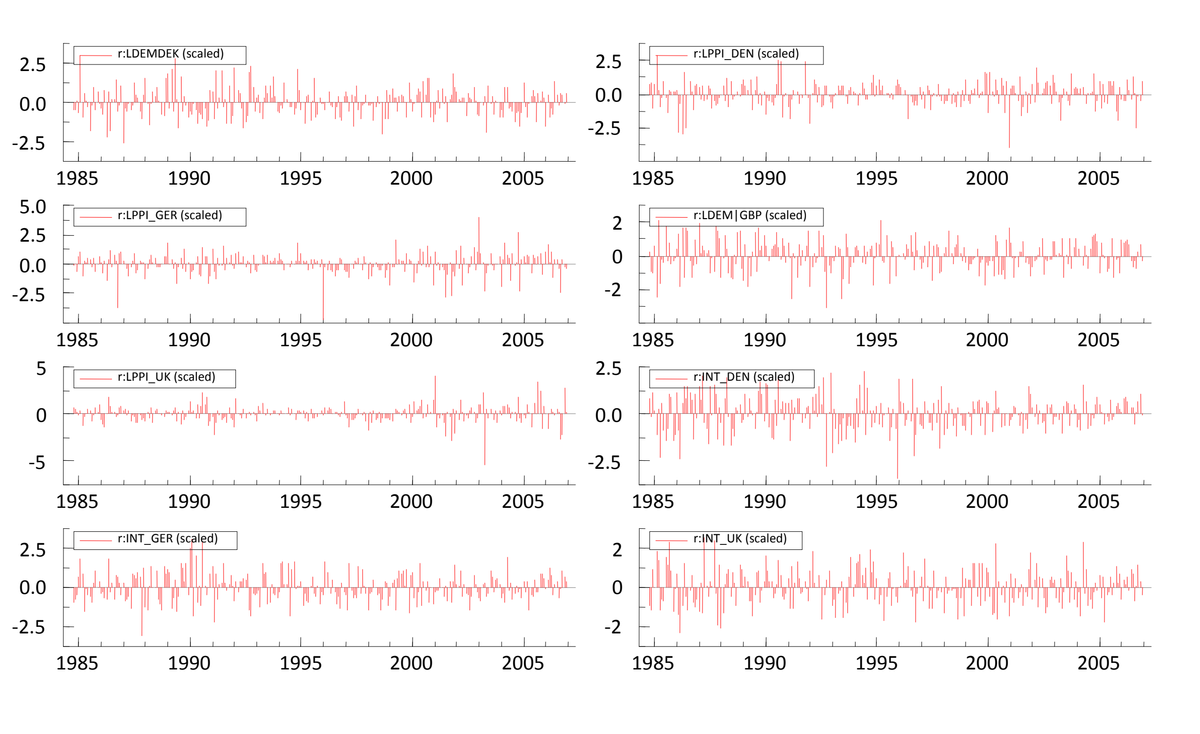
Figure 4.4 ACF and PACF – Nominal Exchange Rates (log), 1984M06-2006M12



The log of the nominal exchange rates of Denmark and the UK vis-á-vis the German DM also appear to be I(1).

Examination of the scaled residuals reveals a number of outliers (these have been defined as having a value greater than 3 in absolute terms). In relation to Danish prices, a dummy is required for 2000M12. For UK prices, dummies are introduced for 2001M01, 2003M04 and 2004M04. German prices reveal extremities at 1996M01 and 2003M01, and Danish interest rates require dummies for 1992M10 and 1995M12.

Figure 4.5 Scaled Residuals of the Series, 1989M01-2006M12



Following Juselius (1995), the relevance of the dummy variables is assessed on examination of their significance across the eight variables under consideration.

Table 4.9 t-statistics of the Dummy Variables (dependent variables in logged first differences)

	D199210	D199512	D199601	D200012	D200101	D200301	D200304	D200404
e_{DEN}	0.77	0.72	0.58	-1.72	-1.64	-1.47	-0.07	0.94
e_{UK}	**-2.91**	0.15	-1.21	**2.39**	-0.26	0.78	0.66	-0.91
p_{DEN}	0.45	0.42	0.20	**-3.73**	-0.03	0.35	-1.55	0.92
p_{UK}	-0.10	0.11	0.43	0.99	**3.49**	**2.23**	**-4.63**	1.38
p_{GER}	-0.98	-0.33	**-4.46**	-1.25	**2.11**	**3.87**	-0.89	0.86
i_{DEN}	**-3.55**	**-4.23**	1.87	-1.12	0.14	-0.64	0.45	0.78
i_{UK}	0.79	-0.72	-0.10	-1.72	-0.18	0.21	0.71	**2.14**
i_{GER}	**-2.50**	-1.17	-1.07	-1.38	-0.71	-0.88	0.79	0.99

Assessment of the significance of the dummy variables based on the logged differences of the variables under consideration (i.e. I(1) transformations), it is clear that all of the dummies are significant for at least one of the variables. In addition, centred seasonal dummies are included in the model.

Table 4.10 Unrestricted VAR(11) Misspecification Test Results

Scalar Residuals	Autocorrelation AR 1-7: F(7,68)	Normality χ^2 (2)	Conditional Heteroskedasticity ARCH 1-7: F(7,61)	Heteroskedasticity χ^2 (176)
e_{DEN}	0.388 [0.907]	9.321 [0.009]*	0.216 [0.981]	174.670 [0.514]
e_{UK}	0.592 [0.761]	8.902 [0.012]**	0.359 [0.922]	189.541[0.229]
p_{DEN}	0.419 [0.888]	5.515 [0.063]	1.093 [0.379]	196.513 [0.138]
p_{UK}	1.632 [0.141]	26.008 [0.000]*	0.214 [0.981]	173.177 [0.546]
p_{GER}	0.887 [0.521]	5.760 [0.059]	0.871 [0.535]	195.128 [0.154]
i_{DEN}	1.481 [0.189]	4.609 [0.099]	0.613 [0.743]	195.522 [0.149]
i_{UK}	0.545 [0.798]	8.204 [0.017]**	1.389 [0.227]	195.469 [0.150]
i_{GER}	0.799 [0.591]	4.666 [0.097]	1.899 [0.085]	184.111 [0.322]

Notes: * and ** refer to the 1% and 5% levels of significance respectively. Residual diagnostics on the VAR indicate no evidence of autocorrelation: AR 1-7: F(448,113) distribution with p-value of 0.712, but there exists evidence of system-wide non-normality.

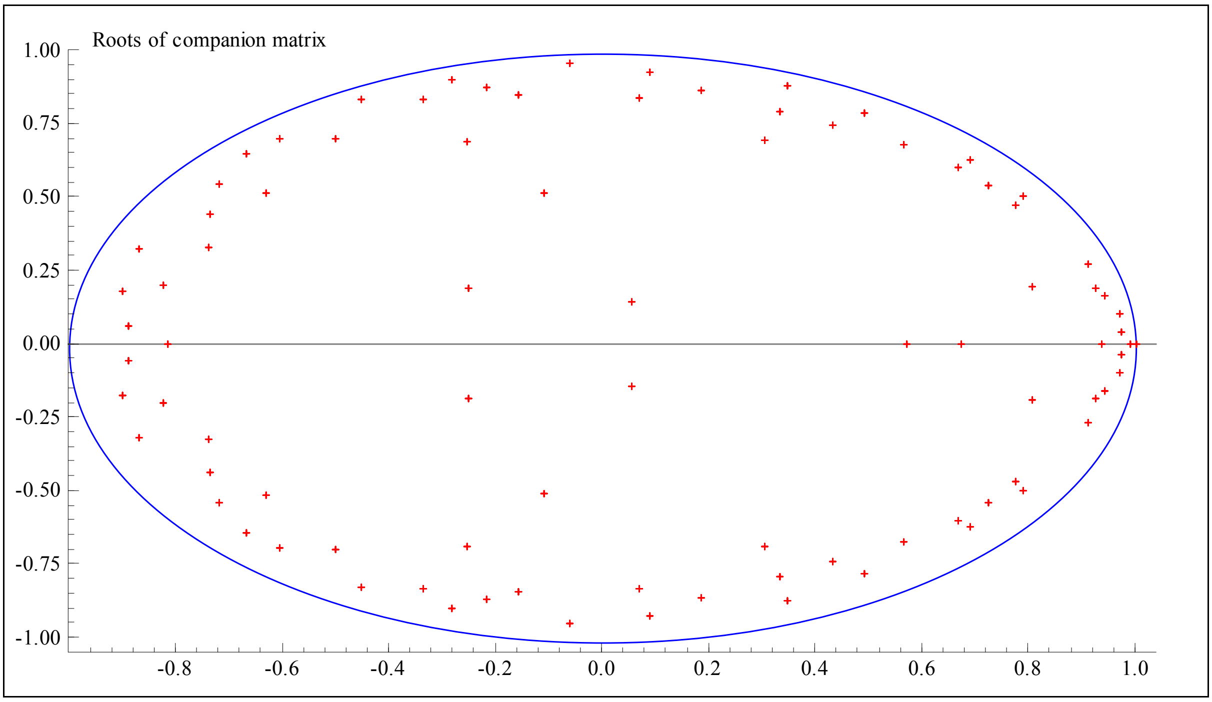
The unrestricted VAR(11) is considered suitable for cointegration analysis based on the Johansen procedure requirements. There are no signs of autocorrelation, ARCH or heteroskedasticity. However, there are some violations with respect to the normality condition. However, based on Gonzalo (1994) who found that the Johansen technique is robust to non-normality, and Cheung and Lai (1993) who reported that the trace test for cointegration rank is fairly robust to normality violations, it is considered that the VAR(11) represents the most appropriate model to proceed for cointegration analysis.

Table 4.11 **Cointegration Rank Test**

Rank	Trace Test (T-nm)	Max Eigenvalue Test (T-nm)
0	212.59 [0.000]**	67.09 [0.000]**
1	145.50 [0.001]**	41.95 [0.135]
2	103.55 [0.012]*	34.44 [0.193]
3	69.11 [0.055]	26.07 [0.328]

The LR trace test indicates the presence of three cointegrating vectors, while the max test shows a cointegration rank of one. Based on the robustness of the trace test to non-normality as described previously, the trace test is preferable to the max test. Further evidence to help to determine the most appropriate cointegration rank is provided on analysis of the roots of the companion matrix (**Figure 4.6** below).

Figure 4.6 **Roots of the Companion Matrix**



The roots or eigenvalues of the companion matrix provide information on the number of roots on the unit circle, thereby providing information on the cointegration rank. The model being estimated contains 88 roots in the companion matrix (i.e. n x k = 88). The six largest roots have moduli in the range 0.9750 to 0.9923. This would suggest that n-r = 6, and thus that the cointegration rank, r, equals two. Based on this information in conjunction with that of the trace test (where the third vector was significant at only the 5% level), it is concluded that there is a rank of two. This would also seem sensible given the presence of two noticeably large eigenvalues on the long-run matrix.

Table 4.12 Eigenvalues of Long-run Matrix

Real	Imaginary	Modulus
-0.1797	0.1121	**0.2118**
-0.1797	-0.1121	**0.2118**
0.1208	0.0000	0.1208
0.02761	0.07921	0.08388
0.02761	-0.07921	0.08388
-0.03947	0.0000	0.03947
-0.005514	0.01061	0.01196
-0.005514	-0.01061	0.01196

Table 4.13 Cointegrating Vectors and Loading Factors

β Vectors							
e_{DEN}	e_{UK}	p_{DEN}	p_{UK}	p_{GER}	i_{DEN}	i_{UK}	i_{GER}
1.0000	1.0325	1.0297	0.2487	-0.8972	0.0030	-0.0102	0.0065
-0.60274	1.0000	-0.4739	-2.4872	1.1879	-0.0109	0.0053	0.0305

α loadings		
e_{DEN}	-0.092 (0.032)	0.198 (0.087)
e_{UK}	0.044 (0.022)	-0.036 (0.018)
p_{DEN}	0.048 (0.021)	-0.127 (0.123)
p_{UK}	0.002 (0.018)	0.012 (0.013)
p_{GER}	0.045 (0.031)	-0.029 (0.020)
i_{DEN}	0.463 (0.188)	-0.223 (0.110)
i_{UK}	0.085 (0.032)	0.009 (0.004)
i_{GER}	-0.012 (0.112)	-0.030 (0.078)

Note: Standard errors shown in parentheses.

The estimated cointegrating vectors are outlined in **Table 4.13**. The vectors are conditioned on the exchange rate variables, both of which are significant in the unrestricted α matrix, indicating an absence of weak exogeneity (moreover, zero restrictions imposed upon the exchange rate parameters in the α matrix are rejected at below the 5% level). These unrestricted estimates are difficult to interpret without undertaking formal restriction tests on the parameters. Nonetheless, it does appear that a number of PPP and UIP relationships might exist. For example, the first cointegrating vector could be representative of a PPP relationship for Denmark (in the strong form case). The adjustment coefficients reveal potential signs of weak exogeneity in the case of German interest rates, and also possibly UK prices. Formal tests are carried out across a range of scenarios to examine whether the long-run relationships evident amongst the variables in the system are consistent with PPP and UIP.

Table 4.14 Restrictions to Test for Weak form and Strong form PPP and UIP

	e_{DEN}	e_{UK}	p_{DEN}	p_{UK}	p_{GER}	i_{DEN}	i_{UK}	i_{GER}	χ^2 (df)	p-value
Weak form PPP tests										
H1: Denmark	1	0	β	0	-β	0	0	0	31.882 (5)	0.000
H2: UK	0	1	0	β	-β	0	0	0	27.432 (5)	0.000
H3: Both	H1 + H2								76.234 (10)	0.000
Strong form PPP tests										
H5: Denmark	1	0	-1	0	1	0	0	0	50.586 (6)	0.000
H6: UK	0	1	0	-1	1	0	0	0	34.476 (6)	0.000
H7: Both	H5 + H6								106.14 (12)	0.000
Weak form PPP plus UIP tests										
H8: Denmark	1	0	β	0	-β	α	0	-α	10.828 (4)	**0.059**
H9: UK	0	1	0	β	-β	0	α	-α	16.530 (4)	0.003
H10: Both	H8 + H9								44.636 (8)	0.007
Strong form PPP plus UIP tests										
H11: Denmark	1	0	-1	0	1	α	0	-α	24.269 (5)	0.002
H12: UK	0	1	0	-1	1	0	α	-α	34.660 (5)	0.000
H13: Both	H11 + H12								68.908 (10)	0.000
Weak form PPP plus Unrestricted Interest Rates										
H14: Denmark	1	0	β	0	-β	μ	γ	η	2.902 (2)	**0.234**
H15: UK	0	1	0	β	-β	μ	γ	η	6.709 (2)	0.035
H16: Both	H14 + H15								23.731 (4)	0.000
Strong form PPP plus Unrestricted Interest Rates										
H17: Denmark	1	0	-1	0	1	μ	γ	η	6.733 (3)	**0.081**
H18: UK	0	1	0	-1	1	μ	γ	η	6.965 (3)	**0.073**
H19: Both	H17 + H18								27.831 (6)	0.000

Based on the restriction tests, it is clear that H8, H14, H17 and H18 are not rejected. H8 identifies weak form PPP and UIP for Denmark. Thus, in the tripolar model estimated, it would appear that PPP augmented by a long-term interest rate is a reasonable description of the first cointegrating vector. The less restrictive case of H14 where long-term interest rates are not restricted indicates a weaker form of the weak form PPP and interest rate dependence in the case of Denmark (H15 is also close to not being rejected with a p-value of 0.04 suggesting that weak form PPP augmented by unrestricted long-term interest rates may hold for the UK, albeit at a much lower level of significance than the case of Denmark).[95] In addition, strong form PPP plus interest rates in unrestricted form appear to hold for the case of both Denmark and the UK. The short-run adjustment coefficients associated with the restricted cointegrating vectors are outlined below. These loadings refer to the case where weak form PPP and unrestricted interests rates appeared to be valid for Denmark in the first cointegrating vector and the UK in the second cointegrating vector (although less significantly). The rationale for examining the restricted loadings is due to the fact that they take into account the variance of the coefficients in the cointegrating vectors (whereas the unrestricted estimates can be reflective of collinearity in the variables entering the VAR for example). The restricted estimates, therefore, represent the movement of the variables towards the identified long run restricted relationship.

Table 4.15 Loadings on Restricted Cointegrating Vectors

	First Cointegrating Vector	**Second Cointegrating Vector**
$\mathbf{e_{DEN}}$	-0.079 (0.025)	0.189 (0.062)
$\mathbf{e_{UK}}$	0.045 (0.014)	-0.035 (0.011)
$\mathbf{p_{DEN}}$	0.019 (0.009)	-0.090 (0.118)
$\mathbf{p_{UK}}$	-0.002 (0.005)	-0.021 (0.034)
$\mathbf{p_{GER}}$	0.017 (0.012)	-0.052 (0.075)
$\mathbf{i_{DEN}}$	0.366 (0.098)	-0.245 (0.089)
$\mathbf{i_{UK}}$	0.083 (0.031)	0.012 (0.006)
$\mathbf{i_{GER}}$	-0.013 (0.011)	-0.036 (0.071)

Note: Standard errors shown in parentheses.

[95] Please note that the hypotheses examined in H14 to H19 are not bound by the theoretical underpinnings linked with the joint testing of PPP and UIP, but only to testing for PPP along with a "raw" measure of capital flows. For this reason, the interest rates for each of the three countries are left in unrestricted form. Thus, for example, H14 examines the hypothesis that weak form PPP holds between Denmark and Germany when the interest rates of Denmark, Germany and the UK enter the system. This enables three countries to enter each of the systems from H14 to H19. Moreover, the co-movement in the interest rates for the three countries over the period in question would justify the inclusion of all three rates. Therefore these restrictions, in effect, test for PPP for two countries in conjunction with interest rate convergence across three countries. It should be noted that the more traditional or expected hypotheses were also carried out, e.g. where H14 and H17 restrict UK interest rates to zero and H15 and H18 restrict Danish interest rates to zero. The results are consistent with those reported, which would confirm the interest rate convergence across the three countries.

It is evident to see that the movement of Danish interest rates in particular is a key force behind the restoration of the equilibrium. This is not a joint PPP-UIP equilibrium, however, since in this case the interest rate variables are in unrestricted form (although it is an equilibrium reflective of trade and capital market interaction). The significance of the results found where PPP is augmented by unrestricted long-term interest rates as opposed to the UIP condition is a point of interest. Indeed, it may very well be the case that restrictions imposed on the long-term rate could fail to hold given the inherent distortions associated with the long-term rate. For example, it is plausible to suggest that issues affecting the long-term rate such as maturity levels and tax distortions could mean that is perhaps not a very appropriate measure for international capital flows. The imposition of restrictions on this rate may magnify the degree of distortion emanating from the long-term rate. However, for the reasons outlined, PPP augmented by unrestricted interest rates is perhaps more likely to hold when a long-term rate is used.[96] Leaving the interest rate variable unrestricted of course does not amount to a test of UIP. Given the lack of distortions on the short-term rate compared to the long-term rate, it is quite possible that the short-term rate reflects international capital flows better than the long-term rate. The following section replicates the analysis for the tripolar model where short-term interest rates are used instead of long-term rates.

IV. Extension to the Tripolar Case: Augmenting the System with the United Kingdom – The Short-term Interest Rate Case

In this section, short-term interest rate indicators replace the long-term interest rate variables for Denmark, Germany and the United Kingdom. The tripolar system estimated by MacDonald and Marsh (2004) across Germany, Japan and the United States

[96] Indeed, in the case of the bipolar model (where long-term interest rates were used), while the evidence suggested that PPP plus UIP restrictions were not rejected, it was also clear that the scale of the non-rejection of the restrictions for PPP plus unrestricted (long-term) interest rates was higher (by a factor of about 2).

incorporated short-term (3 month eurocurrency interest rate) as opposed to long-term interest rates. The authors rationalise this based firstly on the work of others who have undertaken this type of research (e.g. Fisher et al, 1990; Johansen and Juselius, 1992; Lee et al, 1994). In addition, they make the argument that short-term rates are better proxies for capital flows than bond yields, and that bond yields can be subject to distortions such as taxation effects that can differ across countries. For this reason, the analysis is re-estimated using short-term rates. Data availability restrictions alter the start date of the analysis to 1989M01. As in the case of the original system estimated, as first step involves assessing the time series properties of the data by observing the plots of the ACF and PACF for each variable.

Figure 4.7 ACF and PACF – Short-term Interest Rates, 1989M01-2006M12

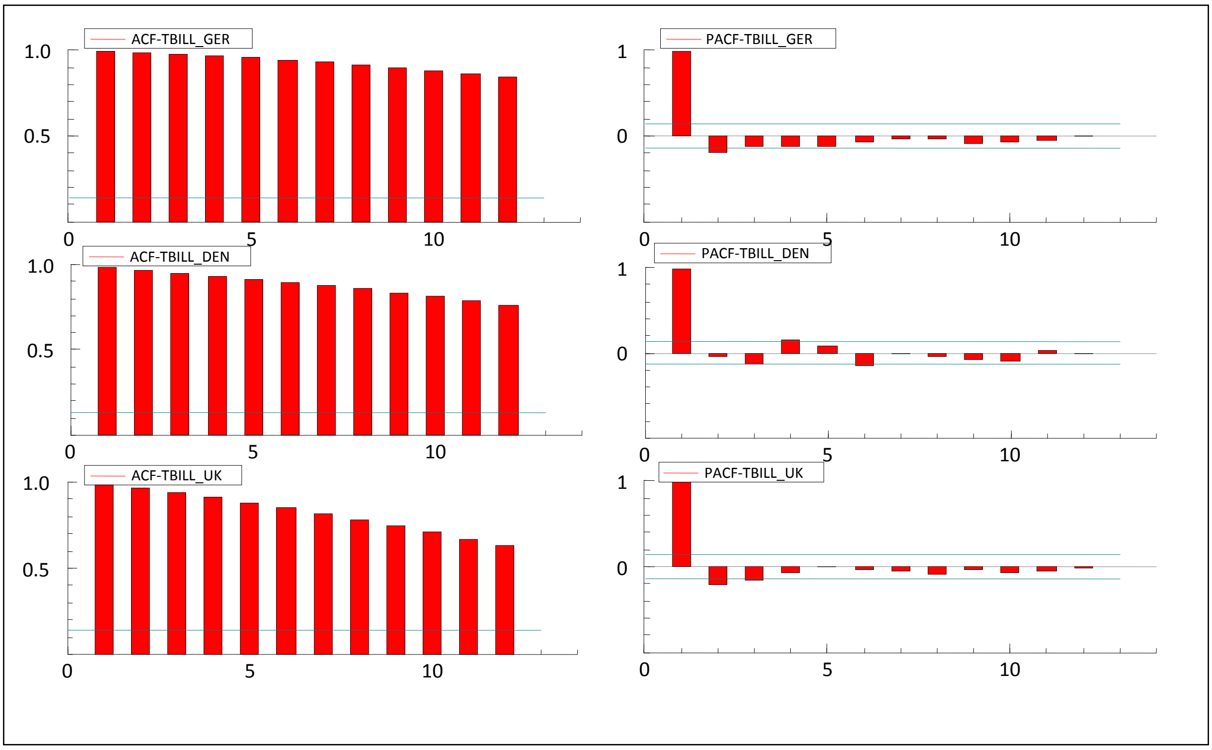


Figure 4.8 ACF and PACF – Producer Price Index (log), 1989M01-2006M12

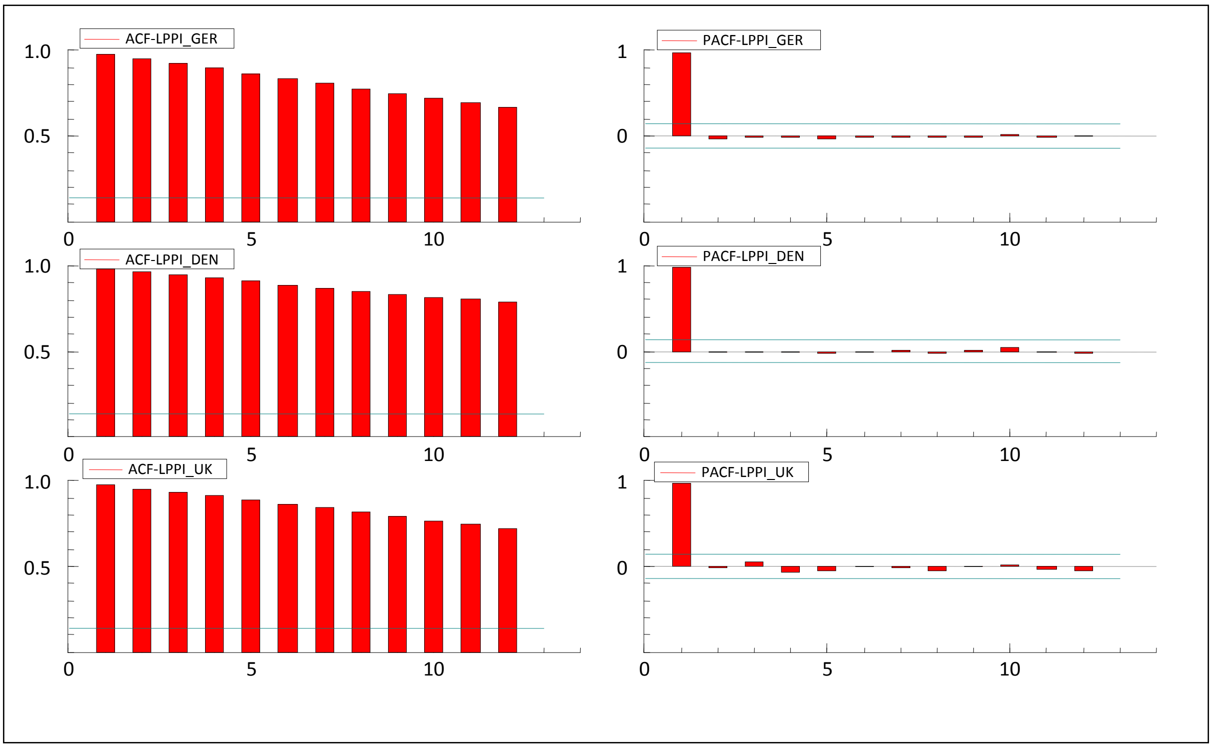
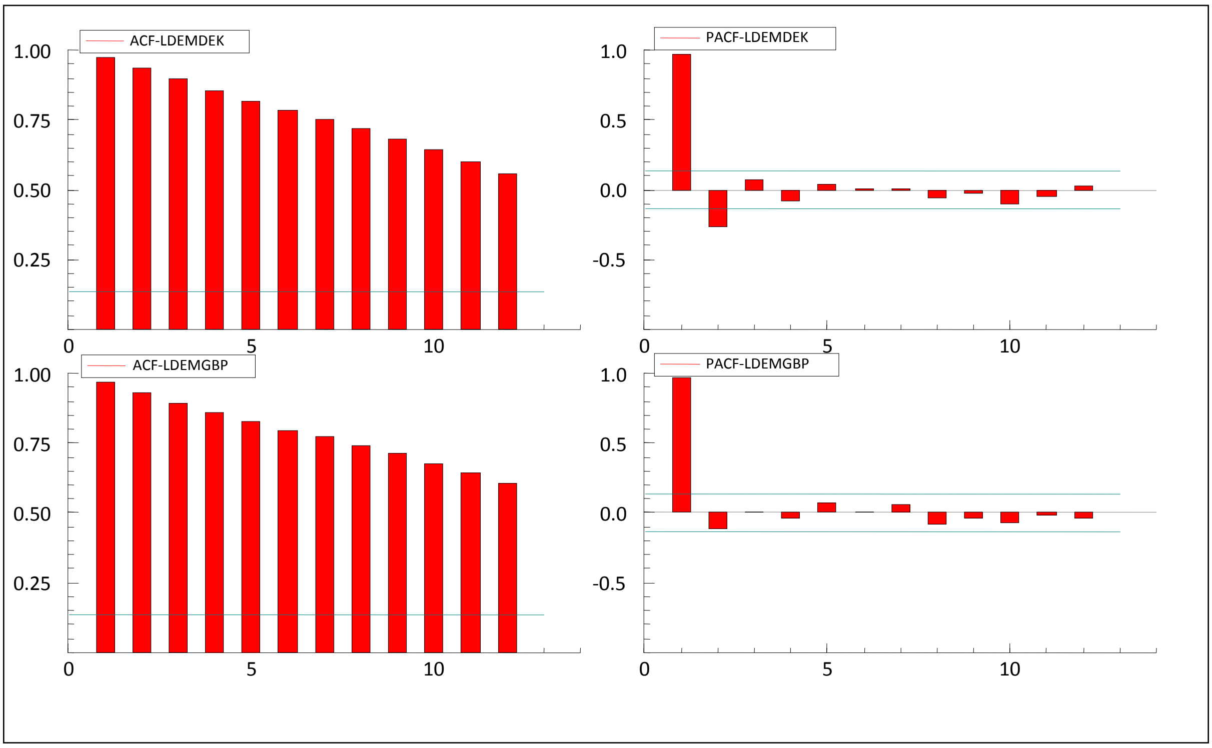
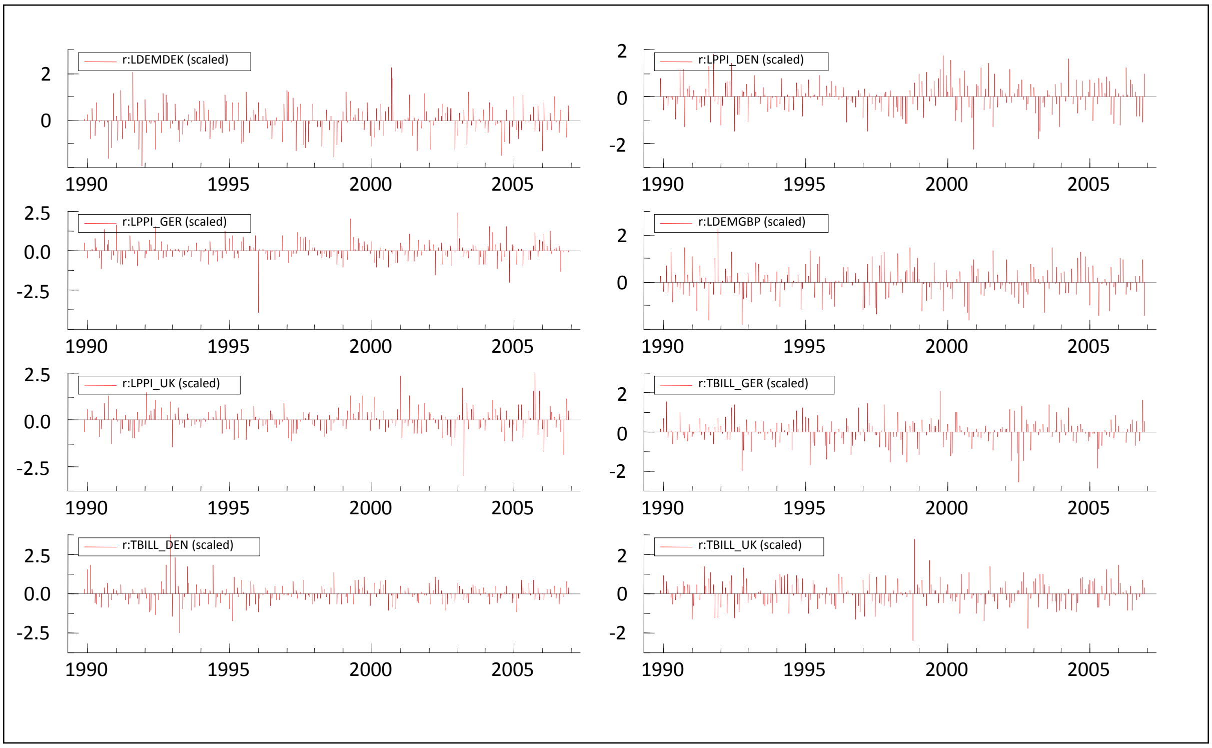


Figure 4.9 ACF and PACF – Nominal Exchange Rate (log), 1989M01-2006M12



The plots of the ACFs and PACFs for each of the eight variables suggest that the first differencing process eliminates the non-stationarity evident in the variables in levels. The vector subject to analysis is the same as the case of the previous analysis, except that in this case the interest rate variable is a short-term variable given by the 90-day short-term interest rate. Thus the vector can be denoted in the same terms: $x' = (e_{den}, e_{uk}, p_{den}, p_{uk}, p_{ger}, i_{den}, i_{uk}, i_{ger})$. A first step involves determining the requirement for dummy variables to correct for outliers and possible non-normality in the system. To this end, the scaled residuals of the series are assessed, and then any dummies identified are tested for significance across the variables in the system.

Figure 4.10 Scaled Residuals of the Series, 1989M01-2006M12



On examination of scaled residuals for each of the series that have an absolute value of 3 or more, there appears to be a requirement for a number of dummy variables: D1992M12 in relation to Danish interest rates, D1996M01 is in relation to German prices, and D2004M04 with regard to UK prices. Prior to making a firm judgment on this, however, an assessment is made of the t-statistics for these dummies across the variables in the system.

Table 4.16 t-statistics of the Dummy Variables (dependent variables in logged first differences)

	e_{DEN}	e_{UK}	p_{DEN}	p_{UK}	p_{GER}	i_{DEN}	i_{UK}	i_{GER}
D199210	0.76	**-3.15**	0.44	-0.11	-1.09	1.51	-0.64	**-4.74**
D199212	0.13	-0.70	-0.78	-0.48	-1.03	**6.28**	0.66	0.02
D199512	0.70	0.21	0.41	0.09	-0.42	**-1.87**	-0.96	-0.03
D199601	0.56	-1.25	0.18	0.36	**-4.67**	-0.83	-0.65	**-1.96**
D200012	**-1.84**	**2.67**	**-4.18**	0.90	-1.37	-0.12	**-1.82**	**-2.39**
D200212	-1.10	**1.86**	0.55	-0.40	0.21	-0.34	-0.21	**-2.12**
D200304	-0.11	0.75	-1.76	**-4.34**	-0.98	0.07	0.19	0.28
D200404	0.94	-0.93	0.95	**1.94**	0.80	0.08	0.08	1.00

At this stage the analysis proceeds with the inclusion of just three dummies in the estimation in unrestricted form, as well as centred seasonal dummies.

Table 4.17 Unrestricted VAR(10) Misspecification Test Results

Scalar Residuals	Autocorrelation AR 1-7: F(7,40)	Normality χ^2 (2)	Conditional Heteroskedasticity ARCH 1-7: F(7,33)	Heteroskedasticity χ^2 (160)
e_{DEN}	0.931 [0.493]	14.206 [0.001]*	0.239 [0.972]	135.726 [0.919]
e_{UK}	1.802 [0.112]	20.984 [0.000]*	0.182 [0.987]	130.239 [0.959]
p_{DEN}	0.846 [0.556]	3.862 [0.162]	0.104 [0.998]	155.241 [0.592]
p_{UK}	1.170 [0.343]	16.577 [0.003]*	0.213 [0.981]	155.743 [0.580]
p_{GER}	0.559 [0.785]	6.622 [0.043]**	1.263 [0.298]	184.897 [0.087]
i_{DEN}	1.048 [0.414]	4.720 [0.089]	0.277 [0.959]	149.777 [0.708]
i_{UK}	0.594 [0.757]	31.006 [0.000]*	0.265 [0.963]	158.073 [0.528]
i_{GER}	1.247 [0.301]	5.376 [0.066]	0.119 [0.997]	140.839 [0.859]

Notes: * and ** refer to the 1% and 5% levels of significance respectively. Some evidence of system-wide non-normality was also evident.

After performing a range of lag length criteria tests for the VAR, k=10 provides the best results in terms of the VAR specification. As in the case where the eight variable system included long-term interest rates, the system that includes short-term rates also exhibits signs of non-normality but no signs of any other residual problems. For this reason, the cointegration rank test is based on the trace test which, as noted earlier, is robust to non-normality (particularly in the case where the non-normality is driven by excess kurtosis, as is the case here).

Table 4.18 Cointegration Rank Test

Rank	Trace Test (T-nm)	Max Eigenvalue Test (T-nm)
0	217.79 [0.000]**	71.20 [0.000]**
1	146.59 [0.001]**	62.10 [0.000]**
2	84.50 [0.231]	31.01 [0.374]
3	53.49 [0.486]	28.54 [0.195]

The evidence from the trace test indicates a cointegration rank of two. The unrestricted β and α matrices are outlined below, normalised on the Danish krone/German DM exchange rate in the first cointegrating vector, and on the UK pound/German DM exchange rate in the second cointegrating vector.

Table 4.19 **Cointegrating Vectors and Loading Factors**

β Vectors							
e_{DEN}	e_{UK}	p_{DEN}	p_{UK}	p_{GER}	i_{DEN}	i_{UK}	i_{GER}
1.0000	1.0428	-0.7510	-0.1211	0.5278	0.0110	0.0064	-0.0007
0.6148	1.0000	0.6959	-0.1819	-0.3972	0.0066	-0.0008	-0.0051

α loadings		
e_{DEN}	-0.130 (0.071)	0.166 (0.018)
e_{UK}	0.267 (0.132)	0.152 (0.067)
p_{DEN}	0.075 (0.023)	0.052 (0.025)
p_{UK}	0.023 (0.027)	0.066 (0.043)
p_{GER}	0.121 (0.048)	0.163 (0.152)
i_{DEN}	-0.152 (0.066)	0.170 (0.059)
i_{UK}	-0.057 (0.049)	-0.022 (0.013)
i_{GER}	0.069 (0.072)	0.059 (0.043)

Note: Standard errors shown in parentheses.

From the two unrestricted cointegrating vectors above, it is evident weak exogeneity is not prevalent in the exchange rate variables, which are used as the basis for conditioning each of the vectors.[97] Interpretation of the unrestricted estimates in terms of possible PPP and UIP relationships is difficult without conducting the appropriate hypothesis tests on the long-run matrix, although it is clear that some of the coefficients appear to be correctly signed in unrestricted form (e.g. the first cointegrating vector could reflect weak form PPP between Germany and Denmark, while the second cointegrating vector might indicate UIP between the UK and Germany). In addition, the short-run adjustment coefficients indicate possible signs of weak exogeneity for UK prices and German interest rates in the first cointegrating vector; and for German prices in the second cointegrating vector. Formal restriction tests on the parameters in each cointegrating vector are considered in the following sub-sections.

[97] A formal test of zero restrictions imposed on the exchange rate parameters in the α matrix is also rejected.

Table 4.20 Restrictions to Test for Weak form and Strong form PPP and UIP

	e_{DEN}	e_{UK}	p_{DEN}	p_{UK}	p_{GER}	i_{DEN}	i_{UK}	i_{GER}	χ^2 (df)	p-value
Weak form PPP tests										
H1: Denmark	1	0	β	0	-β	0	0	0	62.704 (5)	0.000
H2: UK	0	1	0	β	-β	0	0	0	64.790 (5)	0.000
H3: Both	H1 + H2								98.825 (10)	0.000
Strong form PPP tests										
H5: Denmark	1	0	-1	0	1	0	0	0	78.893 (6)	0.000
H6: UK	0	1	0	-1	1	0	0	0	67.155 (6)	0.000
H7: Both	H5 + H6								164.02 (12)	0.000
Weak form PPP plus UIP tests										
H8: Denmark	1	0	β	0	-β	α	0	-α	6.131 (4)	**0.172**
H9: UK	0	1	0	β	-β	0	α	-α	7.128 (4)	**0.071**
H10: Both	H8 + H9								11.369 (8)	**0.053**
Strong form PPP plus UIP tests										
H11: Denmark	1	0	-1	0	1	α	0	-α	43.665 (5)	0.000
H12: UK	0	1	0	-1	1	0	α	-α	48.067 (5)	0.000
H13: Both	H11 + H12								82.63 (10)	0.000
Weak form PPP plus Unrestricted Interest Rates										
H14: Denmark	1	0	β	0	-β	μ	γ	η	6.7664 (2)	0.034
H15: UK	0	1	0	β	-β	μ	γ	η	4.2115 (2)	**0.123**
H16: Both	H14 + H15								21.548 (4)	0.002
Strong form PPP plus Unrestricted Interest Rates										
H17: Denmark	1	0	-1	0	1	μ	γ	η	8.8359 (3)	0.032
H18: UK	0	1	0	-1	1	μ	γ	η	8.0372 (3)	0.045
H19: Both	H17 + H18								38.286 (6)	0.000

With the model that uses the short-term interest rate as the interest rate variable, the results are more conclusive with regard to the parity conditions holding when modelled jointly. It is clear that H8, H9, H10, and H15 are not rejected. This would appear to add some credence to the point made earlier that long-term bond yields can be distortive (due to issues such as the length to maturity and the fact that this means that different taxation policies across countries can be reflected in the yields) and not the best proxies for the interest rate spread and international capital flows. A deeper analysis is carried out for this scenario as regards assessing the adjustment coefficients associated with the restricted cointegrating vectors. This analysis carried out in respect of the weak form PPP plus UIP case (i.e. where the restrictions on the long-run vector were not rejected). **Table 4.21** outlines the loading factors on the restricted long-run vector.

Table 4.21 Loadings on Restricted Cointegrating Vectors

	First Cointegrating Vector	Second Cointegrating Vector
e_{DEN}	-0.125 (0.057)	0.093 (0.040)
e_{UK}	0.286 (0.194)	0.189 (0.082)
p_{DEN}	0.065 (0.059)	0.043 (0.048)
p_{UK}	0.013 (0.016)	0.055 (0.058)
p_{GER}	0.162 (0.039)	0.126 (0.116)
i_{DEN}	-0.216 (0.071)	0.199 (0.077)
i_{UK}	-0.093 (0.067)	-0.033 (0.018)
i_{GER}	0.093 (0.098)	0.073 (0.107)

Note: Standard errors shown in parentheses.

The results from **Table 4.21** show a strong role played by Danish interest rates and the krone/DM exchange rate in correcting deviations from the long-run equilibrium associated with weak form PPP and UIP. It can be see that both of these variables react sharply to deviations from the joint PPP-UIP equilibrium across both cointegrating vectors. This is of interest since it reinforces the MacDonald and Marsh (2004) 'spillover' idea whereby the complex (but interdependent) dynamic evident means that shocks that occur in one particular component of the system can be spread throughout the entire system. Across both cointegrating vectors, Danish interest rates react very sharply to such deviations (21.6% per month regarding the first cointegrating vector, and 19.9% per month for the second cointegrating vector). Both exchange rate variables also appear to react sharply to weak form PPP and UIP deviations. The much slower (and in many cases insignificant) response of prices is not surprising given their inherent stickiness in the short-run (particularly given the monthly data frequency used).

A notable similarity with the bipolar model undertaken in Section II relates to the weak exogeneity associated with German interest rates. This would seem to confirm the dominant role played by German interest rates and the broader monetary policy framework of the Deutsche Bundesbank in influencing financial markets. In this case, this influence is clear in the case of European financial markets, and appears to be robust to different modelling scenarios. A further similarity with the bipolar model relates to the driver of the restoration to the PPP-UIP equilibrium in the face of shocks. Deviations from the equilibrium appear to be largely driven by Danish interest rates, although in the tripolar case the higher degree of interdependence in the dynamics are such that a number of other variables play a role in restoring the equilibrium (notably the krone/DM exchange rate, but also UK interest rates and German prices). The apparent key role played by Danish interest rates can also be seen for the tripolar model undertaken in Section III where long-term rates entered the system instead of short-term rates (although in this case the results are not entirely informative as regards correcting deviations from the joint PPP-UIP equilibrium).

V. Conclusions

The joint modelling of PPP and UIP is rationalized due to the implicit link that exists between the goods market and the capital market. Moreover, theoretical models by Frydman and Goldberg (2002, 2006) suggest that PPP and UIP cannot be assessed independently of each other. Thus, the implication is that since equilibrium in the balance of payments requires both current and capital account equilibria, balance of payments equilibrium can only hold when PPP and UIP hold simultaneously. In other words, it may very well be the case that PPP and UIP are not independent of each other. In addition, the PPP and UIP conditions are fundamental to a number of theoretical exchange rate models, notably the monetary model of the exchange rate. Thus, the finding that both arbitrage conditions hold simultaneously is an important finding in terms of the validity and usefulness of exchange rate models such as the monetary model.

The chapter firstly examines the joint modelling of PPP and UIP for a five-variate VAR for Denmark and Germany (comprising the nominal exchange rate, and the PPI and interest rate for each country). The results find empirical support for stationarity and the validity of the parity conditions when considered in a single system. In this bipolar case, a single cointegrating relationship was found. A subsequent extension to this model by adding the UK to the system (i.e. an eight-variate VAR) showed that this result regarding the joint modelling of PPP and UIP also holds in a tripolar model. The tripolar model validates PPP and UIP simultaneously particularly for the case where short-term rates enter the system. While there is some evidence of weak form PPP plus UIP holding jointly in the long-term rate tripolar model in the case of Denmark, the scale of non-rejection of the restrictions in the short-term rate model is much higher. Overall, it seems that a hypothesis test of PPP plus unrestricted interest rates is most likely to hold for models incorporating long-term interest rates, while a joint restriction of PPP and UIP is most likely to hold when short-terms rates are used in the models. This might suggest that distortions linked with long-term rates are somehow amplified if restrictions for proportionality or symmetry are imposed upon them. Thus short-term interest rates may be a more appropriate proxy for capital flows given that they are not subject to the same types of distortions that can impinge upon long-term bond yields. The tripolar model identified two cointegrating relationships. As regards identifying the drivers of correcting disequilibria and restoring convergence to stationarity, the bipolar model revealed domestic (i.e. Danish) interest rates as being the key variable.

In the tripolar case with short-term interest rates, Danish interest rates would appear to be the primary mechanism for adjustment (certainly in terms of scale and significance). However, the interdependence in the dynamics of the tripolar model suggests that the krone/DM exchange rate, UK interest rates and German prices also react to deviations from the joint PPP-UIP equilibrium. While strong evidence of joint PPP-UIP holding could not be found in the tripolar model with long-term interest rates, an equilibrium for PPP combined with unrestricted interest rates was found to hold strongly for Denmark (and more weakly for the UK). In this case, as with the bipolar model and the tripolar model with short-term rates, a key role appeared to be played by Danish interest rates in reacting to deviations from the identified PPP and interest rate equilibrium relationship. In addition, an interesting finding that holds across both the bipolar and tripolar models was the weak exogeneity of German interest rates. This would seem to make sense intuitively, suggesting high degree of influence by the Deutsche Bundesbank on European financial markets (certainly in the period prior to the formation of the European Central Bank).

The empirical findings outlined have important policy implications. Firstly, it would appear that a comprehensive understanding of the fluctuations of exchange rates, interest rates, and prices can only be attained when the goods and capital markets are modelled simultaneously, i.e. when the model facilitates an interaction between both markets. This type of finding is reflective of Caporale et al (2001), who make the point that the empirical failure of some exchange rate models based on PPP and UIP may have been due to the failure to account for interactions across exchange rates, interest rates and prices, as well as short- and long-run dynamics. Bearing this in mind, it follows that this type of system-based approach can be very informative in terms of exchange rate management for economies. For example, interactions amongst economies can help to provide information on the most appropriate exchange rate regime. Moreover, should a fixed-but-adjustable regime seem appropriate, the system would help to inform the band width within which the regime should operate.

Overall, the main conclusion is that it would appear that for certain groups of economies, the validation of PPP requires interest rate components to be incorporated into the model. This would appear to be particularly intuitive in the current global economy, where trade flows no longer dominate exchange rate fluctuations (certainly in economies with liberalised capital accounts). The increased level of financial globalization, particularly over the past ten years, has also meant that an increasing role exists for international capital flows in the determination of the exchange rate. An empirical example was provided to illustrate the dynamic between the exchange rates, interest rates, and price levels of Germany, Denmark and the UK. These particular economies are of interest in the context of the potential future direction of the euro area. The results provide evidence to suggest that all three economies are suitably integrated in terms of the integration evident across their respective goods and asset markets. Whilst bearing in mind the importance of both the Maastricht criteria and other political concerns, the long-run convergence identified in the analysis is not inconsistent with the suitability, from a monetary perspective, of the UK and Denmark as possible euro area members. Of course, any decision as regards euro area entry would be conditional on adherence to the monetary and fiscal criteria of the Maastricht treaty, as well as country-specific political considerations. Nonetheless, the empirical results in this paper suggest that such a path, from a monetary stand-point, might not be unfeasible.

CHAPTER FIVE

MULTIVARIATE AND PANEL COINTEGRATION:

APPLICATIONS TO THE EURO AREA

I. Introduction

This chapter focuses on the exchange rate behaviour of potential forthcoming euro area entrants, with a view to making a judgement on their suitability for euro area membership.[98] When this chapter was written (in 2006) ten former Central and Eastern European countries, and Cyprus and Malta, were scheduled to enter the euro area by 2014. Of course, since the analysis was undertaken, Slovenia (January 2007), Cyprus and Malta (January 2008), and Slovakia (January 2009) have now joined the euro area. However, since the chapter was written in 2006, Slovenia, Cyprus, Malta, and Slovakia are analysed as prospective euro area entrants.[99]

The expansion of the euro area will increase the heterogeneity within the zone given the differences that exist in the respective stages of development of the countries. Nonetheless, the criteria for euro area entry are clear and should the criteria be satisfied then it follows that a prospective member is applicable for EMU membership. In order to examine this issue more closely (albeit only focussing on the exchange rate and price dynamic), this chapter uses PPP theory and multivariate cointegration techniques (individual and panel) to assess the nature of the relationship between the exchange rate, domestic prices and foreign prices of the countries due to enter the euro area. Given that the focus is on the suitability of the countries for the euro area, the euro exchange rate and euro area prices are incorporated into the models estimated.

Specifically, tests of Purchasing Power Parity (PPP) are implemented using multivariate and panel cointegration procedures for the period since the introduction of the euro

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98 A shortened version of this work was presented at the 6th Annual Meeting of the European Economics and Finance Society (Sofia, Bulgaria, May 2007); the 39th Annual Money Macro and Finance conference (University of Birmingham, September 2007), and the 66th Annual Conference of the International Atlantic Economics Society (Montreal, October 2008). I thank conference participants for all comments received.

99 In fact, given that these four countries have entered EMU after the time of writing, it would be of interest to gauge this against my PPP results for these countries.

currency. The data considered consists of a selection of developed countries in the European Union (EU) but not in the euro area, and a selection of transition economies due to enter the euro area. As a result of the relatively short period since the introduction of the euro, the data selected are at the monthly frequency. The multivariate component relates to Johansen cointegration tests applied to 15 countries individually analysed by a tri-variate VAR comprising the nominal exchange rate, the domestic price level and the foreign price level. The panel component is based on the approach of Larsson et al (2001). Of the ten cases where PPP is found to hold, either the exchange rate or domestic prices drive the short-run adjustment to stationarity. These results have important implications in terms of the suitability of new euro area members. In recognition of the impact of cross arbitrage, results are provided using both the euro and the US dollar as the numeraire currency.[100]

Following this analysis, a short empirical application is provided using the Generalised Purchasing Power Parity (G-PPP) framework and the Johansen multivariate cointegration procedure. The aim is test whether or not a long-run (LR) equilibrium relationship can be identified between the real exchange rates of the euro area and the four largest transition economies that are due to enter the common currency zone by 2014. This helps to assess their suitability as forthcoming members of the euro area. This analysis also acts as a cross-check of findings from the multivariate and panel cointegration tests previously carried out.

The remainder of the chapter is as follows. Section II describes the context of the study. Section III outlines the methodology to be employed. Section IV provides details of the data used and some preliminary analysis. Section V sets out the long-run cointegration

[100] See Smith and Hunter (1985) for more details on the potential misspecification associated with systems that incorporate cross exchange rates.

results by country. Section VI provides an analysis of the short-run adjustment to PPP. Section VII provides the panel cointegration results. Section VIII provides the G-PPP study applied to a selection of the transition economies. Finally, Section IX concludes.

II. Context

The evidence on the validity of the PPP theory has been mixed for the recent floating exchange rate period. Nonetheless, PPP remains a cornerstone in the international finance literature, and continues to represent a benchmark against which the overvaluation or undervaluation of a currency can be measured. A vast amount of empirical research has been undertaken to date on the issue of PPP. This is understandable given the importance attached to PPP as a benchmark theory of exchange rate determination.

Given the planned adoption of the euro by 2014, at the time of writing, by eight former Central and Eastern European countries along with Cyprus and Malta, the exchange rate regimes in these countries has had to align with the euro in preparation for the change. Using the IMF's classification for exchange rate regimes, **Table 5.1** sets out the way in which the exchange rate regimes in the countries concerned has evolved since the 1990s.

Table 5.1 Exchange Rate Regimes in Forthcoming Euro Area Entrants

	1992	1993	1994	1995	1996	1997	1998	1999	2000	2001	2002	2003	2004	2005
Bulgaria	8	8	8	8	8	2	2	2	2	2	2	2	2	2
Czech Rep	3	3	3	3	6	7	7	7	7	8	7	7	7	7
Cyprus	4	4	4	4	4	4	4	4	4	4	4	4	4	4
Hungary	3	3	3	6	6	6	6	6	6	4	4	4	4	4
Malta	3	3	3	3	3	3	3	3	3	3	3	3	3	3
Poland	5	5	5	6	6	6	6	6	8	8	8	8	8	8
Slovakia	3	3	3	3	6	6	7	7	7	7	7	7	4	4
Estonia	2	2	2	2	2	2	2	2	2	2	2	2	2	2
Latvia	8	8	3	3	3	3	3	3	3	3	3	3	3	3
Lithuania	8	8	2	2	2	2	2	2	2	2	2	2	2	2
Romania	7	7	7	7	7	7	7	7	7	6	6	6	6	7

IMF Classification:
1: Dollarization/Euroisation 5: Crawling Pegs
2: Currency Board 6. Crawling Bands
3: Conventional Fixed Pegs 7. Managed Float with no Pre-announced Path for the Exchange Rate
4: Horizontal Bands 8. Independent Float
Source: IMF Reports on Exchange Arrangements and Exchange Restrictions, 1992-2005.

In the early 1990s, there existed a wide variety of regimes in place in these countries and this situation has changed significantly in accordance with ERM II. As noted by Markiewicz (2006), many of the CEECs had high inflation rates in the early 1990s and thus many opted for pegged exchange rate polices against an external anchor. As a result, the economies stabilized throughout the 1990s with inflation coming under control. This attracted a flow of foreign capital, which in a sense challenged the fixed exchange rate arrangements. Thus, towards the mid-1990s, some of the countries opted for more flexibility in their regimes (notably the Czech Republic, Poland, Hungary, and Slovakia).

The fact that different regimes have been in place in these countries implies that the transition paths to the ERM were also different (Coudert and Yanitch, 2001). Due to ERM II and forthcoming Euro Area entry, there have been a number of changes in exchange rate arrangements. Where a currency is pegged, the anchor has largely switched from the dollar to the euro (e.g. Lithuania). Ultimately, where a peg exists, the anchor in preparation for Euro Area entry should be the euro (Bauer and Herz, 2005). For example, Latvia currently operates an SDR peg. Until July 1997, Malta pegged a basket of currencies that gave a large weight to the euro, while Cyprus shadowed the euro within narrow bands. Since July 2007, both countries' currencies were fixed rates against the euro in advance of them adopting the euro in January 2008. As they progress towards entry, the Czech Republic and Slovakia have followed Cyprus by adopting narrow bands vis-á-vis the euro. Some countries that are further away from euro area entry have tended to adopt an arrangement by which their exchange rate shadows the euro, e.g. Hungary shadows the ERM II around the euro within a band of +/- 15%. Many countries, however, have been able to retain some form of floating exchange rate relative the euro, because large capital inflows have required intervention. Bulgaria and Romania are the transition economics that were most recently admitted into the EU, being accepted in January 2007. The Bulgarian lev has been pegged under a currency board arrangement to the DM from 1997-1999 and to the euro thereafter. In contrast, Romania has operated a managed float after a devaluation that followed the beginning of the transition phase.

For those countries that have in place some form of fixed regime, this makes them more vulnerable to a speculative attack, particularly as they are not as developed as most of the industrialized countries. Thus, credibility in relation to the exchange rate regime is of paramount importance. Of the ten countries due to enter the euro area, Bauer and Herz (2005) have highlighted that the entry arrangements of Poland and Hungary, would seem to lack credibility. After the switch from a crawling peg to a free float in 2000, the Polish zloty suffered from a loss of credibility following the financial crisis of 2001. With

Hungary, the switch from a crawling peg to a widely fluctuating peg was slow, causing a degree of exchange rate volatility not consistent with smooth transition to the Euro Area.

It is worth noting that the financial crises in Southeast Asia and Russia in 1997 and 1998 respectively were partly responsible for the decision of the transition economies to alter their exchange rate regimes to make them less vulnerable to speculative attack. In particular, it is well understood that fixed regimes with adjustable pegs are most susceptible to currency crisis-related attacks. After the crisis episodes of the late 1990s, there was a move away to the safer extremes of either currency board arrangements or free floats. This arose, because the capital outflows following the crisis ultimately forced a devaluation of the domestic currencies.

In terms of providing a rationale for the different types of exchange rate regimes operated by the transition countries, Markiewicz (2006) notes that this is largely a function of the extent of financial and trade openness in the countries. Countries that impose capital account restrictions or exhibit more protectionist trade policies are more likely to have a flexible exchange rate arrangement. EU and euro area accession requires, among other things, full financial and trade liberalization in respect of the EU. The implication is that there should be a natural progression towards a fixed exchange rate regime as the economies adopt the euro. As can be seen from **Table 5.1**, all of the forthcoming euro area countries with the exceptions of the Czech Republic and Poland either have had relatively fixed regimes since 1992 or have moved towards increasing degrees of fixity. As described by Lavrac and Zumer (2003), this would appear to be somewhat contradictory and the question remains as to when these countries will alter their regimes to conform to the fixed arrangement that clearly exists within EMU.

The case of the Czech Republic and Poland is interesting as these economies are commonly viewed as being at the forefront of the transition to higher levels of development, with both economies commencing their transition path in the early 1990s.[101] The greater level of reform in Poland, however, through private sector development and exposure to market economy practices perhaps enabled Poland to withstand the financial crisis of 1997. The same could not be said of the Czech Republic, which entered a recession in 1997. These countries share a number of similarities: the transition to a market economy began in the late 1980s and progressed into the 1990s with liberalized currency markets and currency convertibility. As a result, the currencies were undervalued during the early stages of transition. High inflation rates were also prevalent due to price liberalization. As noted by Sideris (2006), the nominal and real exchange rates of these countries were heavily influenced by nominal and real economic shocks. Other researchers such as Dibooglu and Kutan (2001); Suppel (2003); and Borghijs and Kuijs (2004) support this view.[102]

Studies of PPP for the transition economies under consideration are not very prevalent in the literature. Nonetheless, there has been some work done in this area, and the consensus from this appears to be that PPP does not hold for the transition economies. Sideris (2006) examines PPP for 17 European countries in transition[103], finding no evidence for weak form or strong form versions of the theory using the Johansen cointegration procedure. Choudry (1999) used a similar procedure for examining cointegration between relative prices and the nominal exchange rates of Poland, Romania, Slovenia and Russia against the US dollar and he found that relative PPP holds

[101] Barlow (2003) notes that they entered the transition phase from very diverse initial positions. While the former Czechoslovakia was wealthier than Poland at the end of the communist era, it was also less reformed.
[102] Dibooglu and Kutan (2001) examine Poland and Hungary; Suppel (2003) considers Poland, the Czech Republic, and Slovakia; while Borghijs and Kuijs (2004) considers the Czech Republic, Hungary, Poland, Slovakia, and Slovenia.
[103] The countries tested were: Estonia, Latvia, Lithuania, Bulgaria, Croatia, the Czech Republic, Hungary, Macedonia, Poland, Romania, Slovakia, Slovenia, Belarus, Georgia, Moldova, Russia, and the Ukraine.

for Russia and Slovenia. Other examples of previous research include Christev and Noorbakhsh (2000) who reject PPP for Bulgaria, the Czech Republic, Hungary, Poland, Romania and Slovakia. While these authors found some evidence of the existence of a long-run equilibrium relationship, the restrictions imposed for proportionality and symmetry were rejected. An interesting issue raised by them, was that the speed of adjustment towards the long-run equilibrium was higher for economies that had in place floating exchange rate arrangements (i.e. Poland and Hungary). Barlow (2003) tests for PPP for two advanced transition economies (Poland and the Czech Republic) and one less developed transition economy (Romania). He finds that the real exchange rates of Poland and the Czech Republic vis-á-vis a US dollar/German mark geometric average are not stationary, though stationarity is found for Romania on this basis. While an earlier study by Thacker (1995) rejects the hypothesis that real exchange rate stationarity for Hungary and Poland.

Since the end of the Bretton Woods fixed rate period in 1973, a vast amount of research has been carried out on whether PPP holds in the long-run. This chapter considers specifically the period since the introduction of the euro currency in January 1999, with a particular interest in the forthcoming euro area entrants. The dataset effectively encompasses the EU countries that are not in the euro area, that is: Sweden, Denmark, the United Kingdom, Poland, the Czech Republic, Hungary, Slovakia, Slovenia, Latvia, Lithuania, Cyprus, Malta, Estonia, Bulgaria, and Romania.[104]

The latter twelve countries are planning to enter the euro area by 2014. With this in mind, it is appropriate to examine the PPP relationships across these countries to provide some indication of how their exchange rate dynamic would fit in an enlarged euro area.

[104] It is important to note that this refers to the situation in 2006, when the analysis was undertaken. Of course, Slovenia, Cyprus, Malta and Slovakia have all entered to euro area in the years since the analysis was performed.

Denmark, Sweden and the United Kingdom are also assessed as potential euro area entrants over the medium to long term. In addition to assessing whether long-run PPP holds (under both strong form and weak form definitions of the theory), an assessment is made on the adjustment to PPP in the short-run across the regions that do not reject the long-run restrictions. As well as examining the PPP relationships across the regions individually, a panel cointegration test is performed using the Larsson et al (2001) technique which allows for heterogeneity across the groups.

III. Methodology

The methodology to be employed is twofold: the Johansen (1991) multivariate cointegration procedure[105] is used to conduct cointegration tests on each system individually; and the Larsson et al (2001) panel cointegration test is employed to assess cointegration when the country specific systems are considered as a panel. The Johansen test is used to assess whether or not cointegration exists between the nominal exchange rate, domestic price level and foreign price level across fifteen VAR systems. In order to describe this, suppose there is a vector x_t such that $x'_t = (s_t, p_t, p_t^*)$ with a VECM representation:

$$\Delta x_t = \Gamma_1 \Delta x_{t-1} + ... + \Gamma_{k-1} \Delta x_{t-k+1} + \Pi x_{t-k} + \varepsilon_t. \qquad (1)$$

Where the errors $\varepsilon_t \sim IN(0, \Sigma)$, as denoted previously Γ_i represents $-(I - A_1 - .... - A_i)$, $(i = 1,...., k-1)$, and $\Pi = -(I - A_1 - .... - A_k)$. The short-run information is given by the estimates of Γ_i, while long-run information is provided by estimates of Π. The matrix Π has the following decomposition $\Pi = \alpha\beta$\`, where the matrix α captures the speed of adjustment to equilibrium, and β represents the cointegrating vectors. Equation (1) can

[105] See the previous discussion in Chapter 3 or Johansen (1995, 1991) for more detail.

also be augmented to include a constant term to capture trending behaviour, as well as dummy variables to capture seasonal effects or regime shifts.[106]

The Johansen cointegration procedure, two specific test statistics are provided; one relating to the trace test and the other to the maximum eigenvalue test.[107] Both tests determine the number of cointegrating vectors in the system, i.e. the cointegrating rank. Once cointegration is identified, a test for PPP is employed by imposing restrictions on the long-run coefficients in each cointegrating vector as is explained in Johansen and Juselius (1992). Johansen and Juselius have shown that fixing a valid cointegrating vector to satisfy the restrictions associated with PPP gives rise to a likelihood ratio test that is distributed chi-squared under the null. When as might be anticipated here, for each country in turn, there is a single cointegrating vector, then the test in each case has two degrees of freedom given a normalisation. On the basis of an existing literature that is not supportive of PPP, appropriate linear restrictions are tested for both weak form PPP (Peel, 2005) and strong form PPP (Johansen and Juselius, 1992 and Hunter, 1992). The former version of the theory refers to what has been termed the symmetry condition, whereby assuming the rank is one, the β vector x_t when normalised on the exchange rate should give rise to long-run price coefficients of equal and opposite sign, i.e. the vector matrix will not be significantly different from $(1\ -\beta_1\ \beta_1)$. The latter more stringent version of the theory refers to the proportionality condition whereby, for PPP to hold, the β vector should not be significantly different from (1 -1 1). A particular advantage of the Johansen procedure in such studies is that unlike conventional regression there is no requirement to condition the problem on pre-defined exogenous variables.[108] Following

[106] Burke and Hunter (2007) show that the model with intercept is, in general, always used as the base model.
[107] For the reasons given in Chapter 3 above only the trace test results are reported.
[108] However, this is not to say that the identification and interpretation of the long-run equations is not sensitive to the exogeneity status of these variables.

the analysis of the systems individually, the panel test for cointegration due to Larsson et al (2001) is performed. This test corresponds well to the previous analysis since it also relies on the maximum likelihood estimator of Johansen (1988). Unlike the panel test applied by Pedroni (1999, 2004), this does not rely on unit root tests of the residuals and does not assume that there is only a single cointegrating. The technique is applied to a panel of equations based on a sequence of country specific equations:

$$\Delta Y_{i,t} = \Pi_i Y_{i,t-1} + \Sigma \Gamma_{ik} \Delta Y_{i,t-k} + \varepsilon_{it}. \quad (2)$$

Therefore, equation (2) defines the data generating process for each cross section in the panel represented as a VECM. The procedure due to Larsson et al (2001) use the maximum likelihood estimator to produce a trace statistic for each member of the panel in turn. Then the trace statistic for the panel is calculated as an average trace statistic across the members of the panel. The following null and alternative hypotheses due to Larsson et al (op. cit.) are:

$$H_0\text{: rank } (\Pi_i) = r_i \leq r \ \forall \ i = 1, \ldots\ldots, N \quad (3)$$

$$H_a\text{: rank } (\Pi_i) = n \ \forall \ i = 1, \ldots\ldots, N. \quad (4)$$

Where n is the number of variables used to test cointegration in each of the systems in the panel.

The joint hypothesis defined by (3) states that all the countries in the panel have long-run parameters matrices (Π_i) with the same rank. The standardized panel trace statistic is denoted by Larsson et al (2001) as Y_{LR} and is calculated using the following formula:

$$Y_{LR} = \sqrt{N}\ (LR_{NT} - E[Z_k]) / \sqrt{Var(Z_k)} \quad (5)$$

where $E[Z_k]$ and $Var(Z_k)$ represent the mean and variance of the asymptotic trace statistic. In the limit Y_{LR} follows a standard normal distribution. In addition, Larsson et al (op. cit.) indicate that as N and T approach infinity, the power of the test to reject a false hypothesis increases. Specifically, for N≥25 and T≥10, the power of the panel trace statistic is close to unity. Here, the analysis is based on N=15 and T=93, which would appear to suggest that the test ought to be powerful. We have no evidence about the size test.

IV. Data and Preliminary Analysis

The IMF International Financial Statistics CD-ROM is the source of the data for this study. The data includes the euro and US dollar nominal exchange rates for each country and prices (CPI) for the euro area, the US and each of the countries. The set of countries subject to analysis are: the United Kingdom, Sweden, Denmark, Poland, the Czech Republic, Hungary, Slovakia, Slovenia, Latvia, Lithuania, Estonia, Bulgaria, Romania, Cyprus, and Malta over the period 1999M01 to 2006M12 (i.e. monthly frequency). Euro exchange rates were computed using the US dollar nominal exchange rates for each country.[109]

The Johansen multivariate cointegration method entails a number of important pre-conditions, notably that all of the variables in the system are integrated of the same order. Prior to performing the trace test on each of the15 tri-variate VAR systems, the stationarity of each of the series was assessed. The results of the unit root tests are shown in **Tables A5.1** and **A5.2** of the Appendix. Three alternative unit root tests are performed: the ADF test, the DF-GLS test and the KPSS test.[110] In addition, with regard

[109] Euro rates were derived using the following formula: s/s^*, where s denotes the domestic currency per unit of US dollar and s^* denotes the euro per unit of US dollar.
[110] Please refer to Chapter 2 for a more detailed description of the tests.

to the deterministic components, two cases are examined. Firstly, the model with intercept only is examined and then this is augmented to include a linear time trend. The inclusion of the trend is not theoretically consistent with the proposition that PPP holds in the long-run (Zhang and Lowinger, 2006), although it may be for the transition economies as the Balassa-Samuelson effect could be pertinent here. Marcela et al (2003) also noted that a trend may be appropriate since the non-stationarity of the real exchange rate for traded goods can arise as a result of menu costs or pricing-to-market strategies. However, the order of integration of the variables (I(1)) seems to be coherent whether a trend is included or not and as a result it is felt that the time trend is not instrumental in driving the individual series.

Other factors that need to be taken into account are the lag structure of the VAR and the characteristics of the residuals; these issues are dealt with in due course. Based on our analysis, it is clear that all of the series under consideration are non-stationary in levels and have been generated via an I(1) process.[111] Panel unit root tests were also performed using the IPS test, and these results also indicate that the series are all I(1) for the panel as a whole. The ultimate selection of the optimal model is determined in the course of applying the Johansen test, via the Pantula principle and an assessment of the time series properties of the data.

V. Johansen Cointegration Results

This stage involves testing for cointegration among each of the 15 systems, which are comprised of the nominal exchange rate (relative to the euro or US dollar), the domestic price series, and the foreign price series (euro area or US). All of the series are transformed into logarithms. It is likely that the most appropriate lag length for each

[111] This is based on the finding that all the series are considered I(1) at the 10% level or below in at least two of the three unit root tests undertaken.

system will differ across the panel.[112] In selecting the most appropriate model as regards the VAR deterministic components, the Pantula principle (Johansen, 1992)[113] is applied whereby three specifications are estimated and assessed. This principle is not the sole technique applied in selecting the optimal model. Other considerations taken on board include an analysis of the time series properties of the data, e.g. analysis of information provided by the unit root tests, and assessment of the residual diagnostics. The headline cointegration results are set out below. While the focus is on the euro-based results, the US dollar-based results are provided as a means to indicate robustness.[114]

[112] The lag length selection was based on analysis of an unrestricted VAR yielding Gaussian error terms and the lowest AIC. Analysis of the scaled residuals of the series reveals no outliers in the system. The model includes centered seasonal dummies.

[113] Based on the discussion in Burke and Hunter (2007) the tests are applied with intercept even when the simulated DGP excludes an intercept. While Nielson and Rahbek (2000) suggest for similarity of the tests one considers the case with restricted trend even when a trend is not included in the DGP.

[114] It has been shown that data derived from cross rates embodies an implicit sequence of cross arbitrage conditions, which affect the structure of the underlying model and the validity of tests. See Smith and Hunter (1985) for conventional dynamic models, and models that impose PPP and uncovered interest arbitrage, and Hunter and Simpson (2004) for dynamic single equations models. Smith and Hunter (1985) show that in a sample of twelve EU countries, the usual Dickey Fuller test applied to cross rate equations suffers from omitted variable bias. This implies that inference based on conventional Dickey Fuller tests will be distorted as the estimate of the variance-covariance matrix is incorrect under the null and the alternative, and the estimate of $\gamma$is biased and inconsistent. As the Johansen cointegration test is essentially a multivariate version of the ADF test, the use of cross rates in this framework is likely to lead to misspecification unless the parameters of the cross rates are the same as those based on dollar rates. The results from the empirical work undertaken in this chapter indicate only a very marginal difference in the euro-based and US-dollar based tests, thereby removing any concerns regarding cross rate arbitrage. There is a suggestion in Hunter and Simpson (1995) that the effect might be ameliorated by aggregation across a basket of currencies or by the use of independently collected euro exchange rate data.

Table 5.2 LR Trace and Maximum Eigenvalue Test Results (euro as numeraire)

H_0: rank=p	λ_{trace}	95%	λ_{max}	95%
UK (k=9; model includes restricted intercept)				
0	58.0*	0.00	52.2*	0.00
1	10.8	0.19	12.7	0.16
2	2.2	0.56	3.2	0.56
Sweden (k=5; model includes restricted intercept)				
0	54.9*	0.00	39.8*	0.00
1	11.1	0.22	11.5	0.22
2	2.3	0.49	3.6	0.49
Denmark (k=5; model includes restricted intercept)				
0	49.5*	0.00	34.6*	0.00
1	11.3	0.24	10.9	0.27
2	4.0	0.43	4.0	0.43
Poland (k=8; model includes unrestricted intercept)				
0	30.2*	0.04	17.6	0.15
1	11.1	0.13	12.6	0.09
2	0.6	0.88	0.0	0.88
Czech Republic (k=10; model includes unrestricted intercept)				
0	41.0*	0.00	27.9*	0.00
1	10.3	0.11	12.9	0.08
2	0.2	0.63	0.2	0.63
Hungary (k=10; model includes unrestricted intercept)				
0	31.4*	0.03	21.1*	0.05
1	8.4	0.26	9.2	0.27
2	1.7	0.28	1.2	0.28
Slovakia (k=4; model includes restricted intercept)				
0	56.9*	0.00	40.0*	0.00
1	10.3	0.12	15.5	0.06
2	1.7	0.82	1.7	0.82
Slovenia (k=7; model includes unrestricted intercept)				
0	37.6*	0.01	24.8*	0.01
1	11.3	0.13	12.5	0.09
2	0.2	0.65	0.2	0.65
Latvia (k=10; model includes unrestricted intercept)				
0	34.5*	0.01	21.2*	0.05
1	9.3	0.10	12.8	0.08
2	0.5	0.48	0.5	0.48
Lithuania (k=9; model includes restricted intercept)				
0	58.7*	0.00	42.1*	0.00
1	10.8	0.15	10.9	0.27
2	5.7	0.23	5.7	0.23
Cyprus (k=6; model includes restricted intercept)				
0	59.3*	0.00	44.0*	0.00
1	10.9	0.22	13.9	0.10
2	1.4	0.88	1.4	0.88
Malta (k=7; model includes unrestricted intercept)				
0	29.8*	0.05	19.8*	0.08
1	9.3	0.28	9.9	0.22
2	0.1	0.74	0.1	0.74
Estonia (k=10; model includes unrestricted intercept)				
0	35.4*	0.01	28.2*	0.00
1	6.3	0.56	7.0	0.50
2	0.2	0.63	0.2	0.63
Bulgaria (k=8; model includes unrestricted intercept)				
0	54.49*	0.00	38.6*	0.00
1	8.3	0.34	8.8	0.21
2	0.9	0.84	1.1	0.77
Romania (k=8; model includes unrestricted intercept)				
0	43.5*	0.00	29.3*	0.00
1	7.1	0.54	6.9	0.29
2	1.3	0.97	1.3	0.62

Notes: * denotes rejection of the null hypothesis of no cointegration.

Table 5.3 LR Trace and Maximum Eigenvalue Test Results (US dollar as numeraire)

H_0: rank=p	λ_{trace}	95%	λ_{max}	95%
UK (k=11; model includes unrestricted intercept)				
0	33.9*	0.02	22.0	0.04
1	11.9	0.16	8.7	0.32
2	3.2	0.07	3.2	0.07
Sweden (k=9; model includes unrestricted intercept)				
0	37.3*	0.01	30.5*	0.00
1	6.8	0.61	6.2	0.60
2	0.7	0.42	0.7	0.42
Denmark (k=9; model includes unrestricted intercept)				
0	31.2*	0.03	19.6	0.08
1	11.6	0.18	11.5	0.13
2	0.1	0.81	0.1	0.81
Poland (k=6; model includes unrestricted intercept)				
0	32.0*	0.03	19.6	0.08
1	12.4	0.14	10.5	0.19
2	1.9	0.17	1.9	0.17
Czech Republic (k=9; model includes restricted intercept)				
0	51.0*	0.00	31.5*	0.00
1	10.5	0.06	12.9	0.14
2	6.7	0.15	6.7	0.15
Hungary (k=5; model includes unrestricted intercept)				
0	31.0*	0.04	24.2*	0.02
1	6.8	0.61	6.4	0.57
2	0.4	0.55	0.4	0.55
Slovakia (k=9; model includes unrestricted intercept)				
0	34.4*	0.01	26.8*	0.01
1	7.6	0.52	6.9	0.51
2	0.7	0.42	0.7	0.42
Slovenia (k=8; model includes unrestricted intercept)				
0	34.2*	0.01	24.1*	0.02
1	10.1	0.28	9.7	0.24
2	0.5	0.49	0.5	0.49
Latvia (k=5; model includes unrestricted intercept)				
0	30.4*	0.04	16.9	0.18
1	13.4	0.10	11.6	0.13
2	1.9	0.17	1.9	0.17
Lithuania (k=7; model includes restricted intercept)				
0	39.6*	0.01	26.5*	0.01
1	13.1	0.37	10.3	0.32
2	2.8	0.63	2.8	0.63
Cyprus (k=6; model includes restricted intercept)				
0	39.1*	0.02	19.9	0.11
1	12.2	0.07	16.2	0.05
2	3.0	0.59	3.0	0.59
Malta (k=6; model includes restricted intercept)				
0	52.6*	0.00	38.3*	0.00
1	14.3	0.27	9.5	0.40
2	4.8	0.31	4.8	0.31
Estonia (k=3; model includes restricted intercept)				
0	55.7*	0.00	37.7*	0.00
1	13.1	0.10	11.1	0.25
2	7.0	0.13	7.0	0.13
Bulgaria (k=10; model includes unrestricted intercept)				
0	38.2*	0.00	28.2*	0.02
1	11.7	0.18	9.3	0.12
2	2.8	0.09	3.6	0.21
Romania (k=5; model includes unrestricted intercept)				
0	46.7*	0.00	23.8	0.10
1	12.1	0.15	10.6	0.15
2	0.3	0.59	0.0	0.88

Notes: * denotes rejection of the null hypothesis of no cointegration.

Tables 5.2 and **5.3** indicate the optimal VAR specification for each of the tri-variate systems, for different lag lengths and deterministic components (the former table refers to results where the euro is the numeraire currency, while the results of the latter table are based on the US dollar as the numeraire currency).[115]

In the majority of cases (nine out of fifteen for the euro-based results; ten out of fifteen for the dollar-based results), the most appropriate model appears to be that which excludes a trend and permits the intercept to enter both the cointegration space and the VAR, i.e. unrestricted intercept. The remaining models restrict the intercept to the cointegration space. All of the unrestricted VAR models described are well-specified in terms of residual diagnostics (refer to **Tables A5.3** and **A5.4** for details and results of misspecification tests undertaken). There are no signs of autoregressive behaviour, ARCH or heteroskedasticity. There are, however, some signs of non-normality across a very small proportion of the models.[116] However, as described by Gonzalo (1994) and Cheung and Lai (1993), the trace statistic is robust to non-normality in the residuals. The observation that the deterministic time trend seems not to be significant is of theoretical interest from the perspective of observing a Balassa-Samuelson effect. The suggestion is that Balassa-Samuelson effects may not be relevant for the transition economies in question for the time period under consideration. Some empirical support for this proposition can be found in De Broeck e Slok (2001) and Caporale et al (2008). The rationale is that overall competitiveness is not affected by the real exchange rate appreciation as a result of the initial undervaluation of the currencies and the dampening effect on inflation. Otherwise, these economies have developed considerably following

[115] The cointegration analysis to be pursued is driven by the trace test results given the robustness of this test to certain types of non-normality. The maximum eignevalue test results are provided in any case for comparative purposes.
[116] Across the 120 normality test statistics computed for US dollar and euro numeraire cases, in less than 10% of the test statistics computed for normality do we appear to accept the alternative of non-normality. In addition, where non-normality was found, results indicate that it is due to excess kurtosis rather than skewness (See **Table A5.5** in Appendix).

the breakdown of the Eastern block. In particular, there have been considerable market reforms and high levels of investment by German Korean and Japanese companies.

Based on the trace test statistics, there is evidence that the cointegrating rank for the panel is one regardless of the selection of the numeraire currency. All of the models estimated incorporate centred seasonal dummies to capture any seasonal effects. With one cointegrating vector within each of the systems, the implication is that there is some long-run stationary equilibrium relationship between the variables in each system. However, without further tests this cannot be interpreted as indicating that PPP holds for these transition economies.

In order to determine PPP, firstly an analysis is provided of the long-run coefficients. Following this, two hypothesis tests are performed across each system: one tests for the weak form version of PPP by imposing the symmetry condition that the domestic and foreign price series are of equal magnitude but opposite in sign; the other tests for the strong form version of PPP by imposing the proportionality condition that the domestic and foreign prices series are not significantly different from 1 and –1. Both hypothesis tests are performed assuming a normalisation based on the nominal exchange rate.

Table 5.4 PPP Tests of Restrictions on β Matrix (euro as numeraire)

Region	Vectors normalised on s			Weak form PPP	Strong form PPP
	s	p	p*	β_1=-β_2 (χ^2(1))	β_1=-1; β_2=1 (χ^2(2))
UK	1	-57.08	70.86	48.65* [0.00] 2df	50.42* [0.00] 3df
Sweden	1	-5.83	3.39	1.22 [0.54] 2df	1.44 [0.70] 3df
Denmark	1	-0.65	0.54	0.98 [0.61] 2df	1.02 [0.80] 3df
Poland	1	-5.85	8.07	3.70 [0.05]	16.25* [0.00]
Czech Republic	1	16.42	-23.37	11.58* [0.00]	28.01* [0.00]
Hungary	1	23.64	-69.25	8.04* [0.00]	13.15* [0.00]
Slovakia	1	-53.48	53.71	0.38 [0.83] 2df	1.26 [0.74] 3df
Slovenia	1	-74.75	100.42	0.75 [0.39]	8.54* [0.01]
Latvia	1	11.52	-26.37	7.30* [0.01]	7.47* [0.02]
Lithuania	1	-12.95	14.55	5.18 [0.08] 2df	32.54* [0.00] 3df
Cyprus	1	-0.87	0.25	0.12 [0.94] 2df	3.65 [0.30] 3df
Malta	1	-1.05	1.87	3.71 [0.05]	13.69* [0.00]
Estonia	1	13.32	-26.58	11.11* [0.00]	20.47* [0.00]
Bulgaria	1	-0.00	0.01	2.84 [0.09]	43.49* [0.00]
Romania	1	-4.55	8.36	0.72 [0.39]	27.56* [0.00]

(Note: * indicates rejection of the null at the 5% level or below; p-values are denoted in squared brackets).

Tables 5.4 and **5.5** report the long-run cointegrating coefficients and hypothesis tests of PPP as described. Both the weak and strong versions of PPP are tested whereby the matrix [s p p*] is restricted to [1 -β_1 β_1] and [1 –1 1] respectively.

Table 5.5 PPP Tests of Restrictions on β Matrix (US dollar as numeraire)

Region	Vectors normalised on s			Weak form PPP	Strong form PPP
	s	p	p*	β_1=-β_2 (χ^2(1))	β_1=-1; β_2=1 (χ^2(2))
UK	1	-12.13	10.26	10.10* [0.00]	17.37* [0.00]
Sweden	1	-1.41	1.82	0.22 [0.64]	7.80* [0.02]
Denmark	1	-0.03	0.03	0.12 [0.73]	4.83 [0.09]
Poland	1	-5.79	1.85	2.68 [0.10]	13.89* [0.00]
Czech Republic	1	-2.78	1.90	4.20 [0.12] 2df	9.16* [0.03] 3df
Hungary	1	6.55	-11.85	13.55* [0.00]	25.31* [0.00]
Slovakia	1	0.93	0.66	4.72* [0.03]	13.98* [0.00]
Slovenia	1	-10.57	8.77	0.36 [0.55]	9.58* [0.01]
Latvia	1	-13.28	24.06	8.66* [0.00]	9.91* [0.01]
Lithuania	1	-4.51	4.01	5.16 [0.08] 2df	11.75* [0.01] 3df
Cyprus	1	-6.86	-4.90	6.63* [0.04] 2df	6.84 [0.08] 3df
Malta	1	-12.73	12.43	0.69 [0.71] 2df	3.61 [0.31] 3df
Estonia	1	-34.26	41.97	2.04 [0.36] 2df	2.74 [0.43] 3df
Bulgaria	1	-5.83	7.36	1.72 [0.19]	21.31* [0.00]
Romania	1	-6.48	7.64	0.13 [0.72]	27.27* [0.00]

(Note: * indicates rejection of the null at the 5% level or below; p-values are denoted in squared brackets).

Across the fifteen systems for the euro-based results, it can be seen that many of the unrestricted long-run equations appear to be representative of PPP when normalised on the nominal exchange rate. For example, Denmark, the Slovak Republic, Lithuania, and Bulgaria appear to reflect a weak form PPP, given the almost perfect symmetry and correct signs evident for domestic and foreign prices. A range of other countries exhibit the correct signs but there is no symmetry (e.g. the UK, Sweden, Slovenia, Cyprus, Malta). Since the normalisation is on the log of the exchange rate, the interpretation of the price coefficient is as an elasticity with respect to the exchange rate. While for some countries, notably the Slovak Republic and Slovenia, the price coefficients are considerably higher, than say Bulgaria or Cyprus.

While it is clear that some of the magnitudes across countries vary substantially, a plausible rationale for this may relate to the extent of exchange rate pass-through in each of the economies. For example, in the case of the Slovak Republic, it is apparent that a change in prices is associated with a much greater than proportionate change in the exchange rate, thereby suggesting a low rate of exchange rate pass-through to prices. By contrast, a higher pass-through effect is apparent for Bulgaria for example (where a small change in the exchange rate is associated with a greater than proportionate change in prices). There exists some economic intuition behind these findings in terms of the exchange rate regime in place. One would expect a strong relationship to exist between the exchange rate and nominal variables under a fixed exchange rate regime, i.e. high pass-through. A flexible regime would be more likely to exhibit low pass-through as the exchange rate and price linkages weaken[117]. This pass-through effect should be regarded somewhat tentatively as more accurate estimates of pass-through would require a fully-fledged cointegration framework incorporating variables other than the exchange rate and

[117] See Beirne and Bijsterbosch (2009) for empirical evidence to this effect for the new EU Member States.

domestic and foreign price levels. It is difficult to precisely infer the implications for the PPP relationships of the various countries from the unrestricted estimates alone, other than recognizing the sign and symmetry characteristics and whether these accord with PPP. It is clear that some of the price coefficients appear to be somewhat excessive, and this has been a feature of previous PPP studies that use cointegration methods.[118]

Finding a solution to this issue has been the subject of some recent research. Paya and Peel (2006a) advocate a bootstrap analysis associated with an ESTAR model and suggest that the method due to Saikonnen is preferable to the Johansen procedure. However, this does not concur with recent work by Nielsen (2008) and Rahbek et al (2008) for example.[119] An alternative rationale may relate to the conditioning of the equations. For example, following Boswijk (1996), the large coefficients may suggest that the results are sensitive to the normalisation imposed. Testing of the imposition of a zero restriction on the exchange rate in the long-run matrix, however, is rejected in all cases with the exception of Hungary and Latvia, where there is a marginal non-rejection at the 10% level.[120]

[118] A similar issue relating to the magnitude of price coefficients in cointegration tests of PPP was found in MacDonald (1993) and Baum et al (2001). In a study of PPP for the US and Germany, MacDonald reported coefficients of 15.166 and -7.825 for the US and German price levels (CPI) in a cointegrating vector normalised on the US dollar/German Mark exchange rate. Using wholesale prices, coefficients of 65.984 and -37.594 were reported. In the paper by Baum et al, again examining PPP for the US and Germany but over a longer time period, CPI-based price coefficients of 2.7 and -0.668 are reported, while for wholesale prices the coefficients are 62.11 and -36.466. Both of these papers also examine other trivariate systems, where there appears to be some variation in magnitude regarding the price coefficients reported, e.g. for the US and France in the MacDonald analysis (CPI) and the US and the UK in the Baum et al study (CPI). Neither of the papers provides any reasoning for the coefficients estimated.

[119] Paya and Peel (2006a) suggest, via a bootstrap analysis associated with an ESTAR DGP, that the Johansen procedure gives rise to estimates of the long-run that are poorly defined and as a result they suggest that the method due to Saikkonen (1991) is to be preferred as the results are less sensitive. Some care needs to be taken with this interpretation firstly, because there is no theoretical basis to suggest that bootstrap inference gives rise to small sample refinements when the Johansen estimator is considered (Rahbek et al (2008)). Further, Nielson (2008) suggests that inference based on the ESTAR estimator may not be well defined. An alternative interpretation of the results in Paya and Peel is that the observation that the Johansen estimates are ill-defined is quite consistent with the sensitivity of the Johansen estimates. However, poorly defined coefficients estimated from eigenvectors might arise, because the normalization is incorrect that might occur when the underlying problem is in fact non-linear. The error correction model provides only a first order correction to non-linearity, though estimation of non-linear structure is often sensitive to a small number of observations (Hendry, 1995).

[120] See **Table A5.6** for the test results (carried out on the model that includes the euro exchange rate and euro area prices).

While the models estimated are robust to a range of misspecification tests, it could very well be the case that more reasonable coefficients estimates may emerge where PPP relationships are augmented by interest rates. This could be relevant for countries with strong or increasingly developing capital markets, for example, where exchange rate relationships are driven not only by trade flows (represented by PPP) but also by capital flows (which could be represented by Uncovered Interest Parity for example).[121] This issue is not considered here however. The transition economies considered in the analysis are at different stages in terms of capital market development, with some economies having poorly developed capital markets. Augmentation of PPP with interest rates would only be appropriate where the capital markets are developed. Moreover, to aid comparability across all countries and to enable the construction of the panel cointegration test, the tri-variate analysis of the exchange rate and domestic and foreign prices is maintained for all countries.

As regards formal tests of restrictions for weak form PPP, these are not rejected in ten cases, while the restrictions for strong form PPP are not rejected in four cases. This is a remarkable result, though an explicit comparison of these results with any other existing study is limited by the fact the data used relates to euro exchange rates and euro area prices. Of notable interest are the former Eastern European countries that are due to enter the euro area by 2014. The results show that only the Czech Republic, Hungary, Latvia, and Estonia reject the weak form PPP restrictions.

It is encouraging to note that two-thirds of the EU countries not in the euro area appear to exhibit adherence to PPP, implying an element of stability in exchange rate management in markets free from distortion. For the other countries that decided not to enter the euro

[121] Care must be taken when markets are developing to consider the limits on efficiency that may arise via an inversion or time variation in the yield curve. For an alternative discussion of similar issues see Minford and Peel (2007).

area at its inception in 1999, both Denmark and Sweden do not reject weak form PPP restrictions. By contrast, there is strong evidence to suggest that PPP does not hold for the UK. If the numerically smaller test statistics are considered, then this is also true when the tests reported in **Table 5.3** are considered where the US dollar is used as the numeraire currency.[122]

When the numeraire currency is the US dollar, it can be seen that the results are very similar to those of the euro. Overall, the cointegrating vectors associated with ten of the VAR systems do not reject the weak form PPP restrictions, while four systems do not reject the strong form PPP restrictions.[123] Of notable interest are the former CEECs that are due to enter the Euro area by 2014. From **Table 5.3**, of 12 such countries, only Hungary, Slovakia, Latvia, and Cyprus appear to reject the restrictions for weak form PPP with the rejection being very marginal for Cyprus. From **Table 5.2**, for the same set of countries, the Czech Republic, Hungary, Latvia, and Estonia reject the weak form PPP restrictions. Overall, for the weak form PPP tests, it is clear that eight of the countries do not reject the restrictions imposed regardless of the base currency used. This robustness across euro and US dollar based cases would suggest, from the perspective of restrictions on the eigen vectors, that the tests on the restrictions are not greatly affected by cross arbitrage.[124]

The US dollar based results in **Table 5.5** can be compared directly with those of Sideris (2006), who also used the US dollar exchange rate and US prices as the basis of the

[122] The reason for undertaking these tests is to provide a comparison with the euro-based results in light of the previously voiced concerns regarding cross exchange rate arbitrage.

[123] While the total number of countries that do not reject the restrictions is the same across euro and US dollar based results, there is a slight difference in the countries that make up these totals. For the euro results, Slovakia and Cyprus adhere to weak form PPP, whereas for the dollar results the Czech Republic and Estonia do not reject the restrictions. With the tests of strong-form PPP, only Denmark and Cyprus do not reject the restrictions across both numeraire currency scenarios.

[124] Similarly, for strong-form PPP, two of the countries do not reject the restrictions across both euro and US dollar based systems.

analysis. While, the results in **Table 5.3** are similar to Sideris (2006), because they also reject the strong form of PPP. For the ten transition economies that are common to both sets of results, Sideris (2006) rejects strong form PPP in all ten cases, whereas here the proposition is rejected in nine out of ten cases (all but Estonia). However, there are differences in relation to the results that relate to the weak form of PPP. Here it is found that the proposition holds in eight out of ten cases, while Sideris (2006) finds evidence in favour of the proposition only in the case of Estonia. The justification for the differences in these results is difficult to determine. Sideris (2006) uses quarterly data while the data used in this study are monthly.[125] The study by Sideris (2006) runs from 1990Q1 to 2004Q1 and a further basis for the discrepancy could be the exchange rate volatility that arose in Central and Eastern European in the early 1990s, both in terms of exchange rate movements and inflation associated with the behaviour of the real exchange rate. In addition, these economies up until the mid-1990s were still heavily influenced by their initial conditions that derived from their history as centrally planned economies. Though there some significant differences between the more liberal economies such Hungary and very strict regimes such as Romania. For this type of reason, it could be argued that PPP would be more likely to hold once the process of 'catch-up' has begun and this might provide a basis for excluding the extremely different and volatile period of the early 1990s. Indeed, previous studies of PPP that include the period of the early 1990s largely failed to find evidence of its validity (e.g. Christev and Noorbakhsh, 2000; Barlow, 2003; Thacker, 1995).

[125] Pava and Peel (2006b) suggest that temporal aggregation may have ramifications for PPP in the context of ESTAR models and similar considerations might apply here.

VI. Short-Run Adjustment to PPP

Using the Johansen procedure also enables one to determine the speed of adjustment to PPP in the short run. As a first step, the statistical significance of the α loadings is observed from the unrestricted model across each country.[126] Following this, restrictions are imposed on the short-run coefficients in order to determine whether the deviation is corrected through the nominal exchange rate or prices (domestic or foreign). The results set out in **Tables A5.7** to **A5.10** (Appendix) apply such restrictions on the α matrix for each system. In addition joint tests are applied both to α and β. It is then possible to determine whether deviations from PPP are corrected via the exchange rate or prices. Effectively, the hypothesis that the exchange rate, domestic prices, and foreign prices do not correct for PPP is tested. The results considered in this section relate only to the systems associated with those countries where one of the two forms of PPP has been accepted.

Of the fifteen countries tested for weak form PPP, ten did not reject the restrictions imposed. In the context of a single cointegrating vector, when an element of the α vector is set to zero, then this implies that the equation associated with the zero coefficient has no long-run and as a result this variable can be viewed as being weakly exogenous for the cointegrating vector (Johansen, 1992). If a variable is not weakly exogenous and there has to be one equation where the correction occurs, then the adjustment must operate in this equation and the larger the coefficient the more powerful the response. Using the euro as the numeraire currency, subsequent tests indicate that for six of these countries, domestic prices are the main driver behind the adjustment towards PPP in the short-run. The countries where this holds are Slovenia, Malta, Bulgaria, Slovakia, Lithuania, and

[126] The unrestricted model in this case refers to the model that imposes only the β restrictions.

Cyprus. Adjustment towards PPP is driven by the euro exchange rate in the cases of Sweden, Denmark, Poland, and Romania. Four countries did not reject the strong form PPP restrictions that were imposed (Sweden, Denmark, Slovakia, and Cyprus). Across these countries the exchange rate drives the adjustment process for Denmark and Sweden, while domestic prices are the driver for Slovakia and Cyprus.

These findings have some intuitive appeal. Where a fixed or pegged exchange rate regime is in place, it makes sense that domestic prices should be the lever of adjustment. By contrast, where inflation expectations are anchored by a combination of a floating regime and inflation targeting, it makes sense that the exchange rate drives the adjustment.[127]

As a robustness check, additional hypothesis tests are performed by considering whether two of the α coefficients are restricted to zero. The results from these tests appear to reinforce the findings made in relation to the primary driver of convergence towards PPP. For example, in the case of Sweden, the exchange rate appears to relate to the strongest response given a single zero restriction on the α matrix is accepted. While, the joint zero restriction on two of the α factors shows that the exchange rate is conditioned in the long-run the two prices and as such the adjustment to equilibrium operates in the short-run on the exchange rate equation. A joint zero restriction on the exchange rate and domestic prices is strongly rejected meaning that these two variables are in combination not weakly exogenous. The same test applied to both the exchange rate and foreign prices cannot be rejected at the 5% level, but the acceptance is marginal. Hence, the two variables can just be considered weakly exogenous for the cointegrating vector restricted

[127] See McKinnon (1991), Williamson (1994), Rodrik (1994) and Grubacic (2000) for further detail on this. These authors examine exchange rate regime appropriateness and real exchange rate dynamics in the EU transition economies, noting that a fixed exchange rate is beneficial at the beginning of the transition process in order to control inflation. Over the long-term, however, a gradual shift to a more flexible regime is needed in order to ensure competitiveness gains, e.g. export sector restructuring.

to satisfy a form of PPP. By contrast, the same form of restriction applied to the domestic and foreign price cannot be rejected. As a result the long-run can be conditioned on prices so the equilibrium can be seen as relating to the exchange rate in the long-run, while the short-run correction in the short-run operates on the dynamic exchange rate equation or the error correction model for the exchange rate. A similar pattern is evident for the other countries examined.

Overall, the fact that PPP appears to hold for the majority of the economies is encouraging in terms of their suitability for entering EMU. The literature is quite mixed in terms of any consensus over the main driver behind convergence to PPP. For example, it has been suggested that the exchange rate and domestic prices are the main factors by Goldfajn and Valdes (1999).[128] On the other hand, Cheung, Lai and Bergman (2004) suggest that domestic and foreign prices have similar adjustment dynamics in the case of Britain, France, Germany, Italy and Japan for the period 1973M04-1998M12. Akram (2006) suggests that the response of foreign prices is negligible when the case of a small economy, Norway is studied for the period 1972Q1-2003Q4. Clearly, different dynamics can become evident when different time periods and data frequencies are used as the basis for the analysis and this makes it difficult to generalize as to which particular factor dominates the adjustment process. Given the size of the majority of the transition economies in this analysis, it is not feasible from economic theory to view foreign prices (i.e. euro area prices) as adjust to restore equilibrium in the exchange rate.

VII. Panel Cointegration Results

The panel cointegration results provided are based on Larsson et al (2001). **Table 5.6** refers to euro-based results, while **Table 5.7** refers to dollar-based results.

[128] This particular study involved 93 developed and developing economies over the period 1960M01 to 1994M12.

Table 5.6 Panel Cointegration Results (euro as numeraire)

COUNTRY	Cointegration Rank (Johansen Trace Results)		
	r=0	**r=1**	**r=2**
UK	58.0	10.8	2.2
Sweden	54.9	11.1	2.3
Denmark	49.5	11.3	4.0
Poland	30.2	11.1	0.6
Czech Republic	41.0	10.3	0.2
Hungary	31.4	8.4	1.7
Slovak Republic	56.9	10.3	1.7
Slovenia	37.6	11.3	0.2
Latvia	34.5	9.3	0.5
Lithuania	58.7	10.8	5.7
Cyprus	59.3	10.9	1.4
Malta	29.8	9.3	0.1
Estonia	35.4	6.3	0.2
Bulgaria	54.5	8.3	0.9
Romania	43.5	7.1	1.3
	Panel Results*		
$E(Z_k)$	14.955	6.086	1.137
$Var(Z_k)$	24.733	10.535	2.212
$LR_{NT \sim DEVELOPED}$	54.133	11.067	2.083
$Y_{LR \sim DEVELOPED}$	**2.179**	**1.191**	**1.133**
$LR_{NT \sim TRANSITION}$	42.733	9.450	1.208
$Y_{LR \sim TRANSITION}$	**3.671**	**1.957**	**0.622**
$LR_{NT \sim TRANSITION\ WITH\ FIXED\ REGIME}$	42.580	8.800	1.480
$Y_{LR \sim TRANSITION\ WITH\ FIXED\ REGIME}$	**2.363**	**1.135**	**0.881**
$LR_{NT \sim TRANSITION\ WITH\ FLEXIBLE\ REGIME}$	42.843	9.914	1.314
$Y_{LR \sim TRANSITION\ WITH\ FLEXIBLE\ REGIME}$	**2.809**	**1.595**	**0.749**

	PPP Restriction Tests for Panel**	
	Weak form PPP (β_1=-β_2)	**Strong form PPP (β_1=-1; β_2=1)**
PANEL 1: DEVELOPED	**50.85** $\chi^2(6)$	**52.88** $\chi^2(9)$
PANEL 2: TRANSITION	**55.43** $\chi^2(15)$	**216.08** $\chi^2(27)$
PANEL 3: TRANSITION WITH FIXED REGIME	**30.14** $\chi^2(6)$	**117.66** $\chi^2(7)$
PANEL 4: TRANSITION WITH FLEXIBLE REGIME	**25.29** $\chi^2(9)$	**98.42** $\chi^2(11)$

(Note: *Panel cointegration results based on critical values of 2.33, 1.96, and 1.65 at the 10%, 5%, and 1% levels respectively. **The panel restriction tests are based on the standard chi-squared 5% critical values).

Table 5.7 Panel Cointegration Results (US dollar as numeraire)

COUNTRY	Cointegration Rank (Johansen Trace Results)		
	r=0	r=1	r=2
UK	33.9	10.6	1.8
Sweden	37.3	6.8	0.7
Denmark	31.2	9.4	0.1
Poland	32	9.1	1.9
Czech Republic	51	9.1	1.7
Hungary	31	6.1	0.4
Slovak Republic	34.4	6.7	0.7
Slovenia	34.2	7.3	0.5
Latvia	30.4	11.2	1.9
Lithuania	34.6	11.3	1.7
Cyprus	33.1	10.2	3
Malta	52.6	10.6	2.3
Estonia	52.7	10.9	3.1
Bulgaria	38.2	10.5	1.8
Romania	36.7	9.3	0.3
	Panel Results*		
$E(Z_k)$	14.955	6.086	1.137
$Var(Z_k)$	24.733	10.535	2.212
$LR_{NT \sim DEVELOPED}$	34.133	8.933	1.267
$Y_{LR \sim DEVELOPED}$	**1.525**	**0.588**	**0.125**
$LR_{NT \sim TRANSITION}$	42.733	9.450	1.208
$Y_{LR \sim TRANSITION}$	**3.671**	**1.957**	**0.622**
$LR_{NT \sim TRANSITION\ WITH\ FIXED\ REGIME}$	41.700	10.900	2.160
$Y_{LR \sim TRANSITION\ WITH\ FIXED\ REGIME}$	**2.325**	**1.512**	**1.510**
$LR_{NT \sim TRANSITION\ WITH\ FLEXIBLE\ REGIME}$	36.057	8.257	1.214
$Y_{LR \sim TRANSITION\ WITH\ FLEXIBLE\ REGIME}$	**2.444**	**1.201**	**0.495**
	PPP Restriction Tests for Panel**		
	Weak form PPP (β_1=-β_2)	**Strong form PPP (β_1=-1; β_2=1)**	
PANEL 1: DEVELOPED	**10.44** $\chi^2(3)$	**30.00** $\chi^2(6)$	
PANEL 2: TRANSITION	**50.54** $\chi^2(17)$	**155.35** $\chi^2(29)$	
PANEL 3: TRANSITION WITH FIXED REGIME	**18.27** $\chi^2(8)$	**49.32** $\chi^2(13)$	
PANEL 4: TRANSITION WITH FLEXIBLE REGIME	**32.27** $\chi^2(9)$	**106.03** $\chi^2(16)$	

(Note: *Panel cointegration results based on critical values of 2.33, 1.96, and 1.65 at the 10%, 5%, and 1% levels respectively. **The panel restriction tests are based on the standard chi-squared 5% critical values).

As well as being consistent with the Johansen procedure, the Larsson et al (op. cit.) approach also overcomes the weakness of other panel cointegration techniques where the estimated cointegrating vector for each panel member may not be unique when multiple regressors are considered as is the case with McCoskey and Kao (1998). PPP is tested across four panel groupings: the panel of developed economies in the sample (UK, Sweden, Denmark), the transition countries, the transition economies operating fixed exchange rate regimes (Bulgaria, Estonia, Latvia, Lithuania, Malta) and the transition economies with flexible exchange rate regimes (Czech Republic, Hungary, Poland, Romania, Slovakia, Slovenia, Cyprus).[129] These results are constructed under the assumption that the individual Johansen trace statistics are independent. The panel trace statistics, Y_{LR}, indicate a maximum of one cointegrating vector for all four panel cohorts (5% level). The subsequent panel PPP tests are carried out on this identified cointegrating vector (see Burke and Hunter, 2005, Chapter 5).

Regarding the PPP tests, the symmetry and proportionality restrictions are rejected for all four panel groupings at conventional significance levels. There is little weak evidence that the weak form PPP may hold for the transition economies with flexible exchange rates when one considers the euro-based results, because non-rejection is at the 0.1% level - and this is too small to draw any robust inference. Moreover, satisfaction of these restrictions is counter intuitive given the expectation that PPP would be more likely to hold in countries with fixed regimes since there ought to be a stronger link between the exchange rate and the prices. For the US-dollar based results, there is stronger evidence (non-rejection at the 1% level) in favour of the weak form PPP for the panel of developed economies and the one for transition economies with fixed exchange rates. Inferring any

[129] Please note that the selection of the regime for the countries is based on the type of regime that was in place over the period 1999 to 2006.

economic implications is avoided due to the marginal significance of the test statistics significance. In recognizing the results of the panel tests, it is important to bear in mind the strong influence of a small number of countries in driving the overall panel outcome (see Beirne, 2009). A similar result is observed in the simulation evidence provided by Taylor and Sarno (1998) who show that panel unit root tests can be sensitive to the introduction of a single stationary series. Here, the reverse seems to be observed as the presence of a single country that strongly rejects the PPP restrictions seems to give rise to this hypothesis being rejected by the panel. These results hold seem to be coherent whether the analysis is undertaken using the euro or the US dollar as the numeraire currency.

VIII. An Alternative Approach to Determine Whether GPPP holds for the Enlarging Euro Area

This short section aims to make the conclusions drawn from earlier sections more robust by examining the real exchange rate dynamic of a selection of the forthcoming euro area entrants over the same period and frequency, i.e. 1999M01 to 2006M12. The methodology is based on the Generalised PPP theory of Enders and Hurn (1994).[130] The focus here is to test whether or not such an OCA exists for the euro area plus the four largest countries due to enter by 2010 (i.e. Poland, the Czech Republic, Hungary, and Slovakia).[131]

Not withstanding the results in **Chapter 2**, many univariate tests of PPP in the past indicate the presence of a unit root in the data generating process. Similar outcomes can be observed with bi-variate tests based on cointegration, while Paya and Peel (2006a)

[130] Please refer to Chapter 3 for a full description of the theory.
[131] Analysis of GDP data for 2006 indicates that of the ten countries due to enter the euro area by 2010, Poland, the Czech Republic, Hungary, and Slovakia together account for 84% of the total GDP for all countries forthcoming. The analysis is thus focused on these countries in particular.

suggest that this may be a consequence of the underlying DGP being derived from an ESTAR model. The G-PPP theory is based on the premise that the combination of the exchange rates of a number of closely linked currencies may exhibit a stationary trend, even though they may be non-stationary individually. Hence, the cointegration approach would seem an appropriate means to undertake tests of G-PPP. The fact that is occurs relies on a high degree of inter-linkage between the economic fundamentals of the respective countries.

The context against which this application is set relates to whether or not PPP is an appropriate theory for groups of countries which have close economic linkages. It may not be surprising to make the finding that PPP is valid for such a set of countries. This application focuses on whether or not G-PPP exists between the euro area and forthcoming adopters of the euro. Since the establishment of the theory, the introduction of the euro represents the first real case of currency union across a number of diverse economies. As the EU expands with the entry of the former Eastern European countries, the euro will eventually become the national currency for about 1 billion people. Of course, entry into the zone is dictated by adherence to the Maastricht criteria. However, such criteria have been viewed as weak and thus easy to attain in a relatively short timescales, while economic fundamentals may not be in alignment over a time periods that exceed 2-3 years. Further, some economies have been observed to flaunt the conditions once they are part of the zone.

Plots of the US$ real exchange rates indicate a similar pattern of fluctuation over the period in question (shown in **Figure A5.1** in Appendix 1). In addition, the unit roots tests shown in **Table A5.11**, indicate that all of the variables are I(1).[132] While for testing cointegration among the real exchange rates, the appropriate lag length for the model is set at five.[133] In selecting the most appropriate model as regards the VAR deterministic components, the Pantula principle is applied whereby three specifications are estimated and assessed. This principle is not the sole technique applied in selecting the optimal model. Other considerations include an analysis of the time series properties of the data, e.g. analysis of information provided by the unit root tests, and assessment of the residual diagnostics. Misspecification tests of the unrestricted VAR are reported in **Table A5.13**, indicating no signs of any error auto-correlation, non-normality, ARCH or heteroskedasticity. The Pantula and cointegration rank test results are also set out in the Appendix (**Tables A5.12** and **A5.14**).

Evidence is found of a long-run stationary relationship in the real exchange rates. **Figure A5.2** provides an indication of the stability of the cointegrating vector. This is indicative of a close interdependence in the long-run between the real exchange rates analysed. In this sense, the results are supportive of evidence in favour of G-PPP over the entire period. As a result it would appear that the four countries that comprise 85% of the GDP of the ten forthcoming entrants to the euro area by 2010 are suitable for entry. **Table 5.8** normalises the long-run matrix on the US/euro real exchange rate, and also sets out the short-run coefficients.

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[132] There is a degree of inconclusiveness, however, in relation to the deterministic components of the models. Using the Dickey-Fuller *tau* statistics, it seems appropriate to conclude that none of the series exhibit trend behaviour. The presence of an intercept, however, is largely open to debate and sensitive to the type of test employed. The ultimate selection of the optimal model is determined in the course of applying the Johansen test, via the Pantula principle. The finding that all of the real exchange rate series are I(1) is consistent with the G-PPP theory, whereby interdependence in economic fundamentals is reflected in the behaviour of the real exchange rates.

[133] The lag length selection was based on analysis of an unrestricted VAR yielding Gaussian error terms and the lowest AIC. Analysis of the scaled residuals of the series reveals no outliers in the system. The model includes centered seasonal dummies.

Table 5.8 Normalised Cointegrating Equations (Long-Run Elasticity) and Short-Run Coefficients

Normalised Cointegrating Equations (LR)				
Euro area	*Poland*	*Slovakia*	*Hungary*	*Czech Republic*
1.000	-0.168** (0.084)	0.024 (0.142)	-0.047 (0.109)	-0.948* (0.074)
Short-Run Speed of Adjustment Coefficients				
Euro	*Polish Zloty*	*Slovak Koruna*	*Hungarian Forint*	*Czech Koruna*
0.391* (0.146)	0.224*** (0.164)	0.438* (0.167)	0.323* (0.154)	0.813* (0.147)

Notes: Results are reported for the most significant cointegrating vector in each region. Standard errors are reported in parentheses. *, **, and *** indicates statistical significance at the 1%, 5%, and 10% levels respectively.

As detailed in **Chapter 3**, an optimum currency area could be considered appropriate if a number of conditions are met: (i) that a long-run equilibrium relationship can be identified; (ii) that the long-run coefficients are low in magnitude and similarly signed; and (iii) that similar speeds of adjustment are evident. Of course, statistical significance of each parameter is also a crucial factor. A pre-condition to carrying out such an analysis is also conditional on attaining a VAR in unrestricted form that exhibits appropriate levels of specification. As already described, tests of autoregressive behaviour, non-normality, ARCH behaviour, and heteroskedasticity are all rejected at the 5% level or below (refer to **Table A5.13** in Appendix).

Slovakia has a different sign and is not significant, while the long-run coefficients for Poland and the Czech Republic have the same sign and are significant at the 5% level. All of the long-run coefficients are less than unity, although the one for the Czech Republic is fractionally less than 1. The size of the coefficient is important in the context of G-PPP, because according to Enders and Hurn (1994) when real output processes are the drivers of the real exchange rates across each of the countries the normalised vector will be smaller when countries are deemed to have similar output parameters. The results show that, for example, a 1% decrease in the Polish/US real exchange rate increases the real euro/dollar rate by 0.17%, i.e. a weakening of USDPLN is associated with a strengthening of USDEUR. The Czech koruna real exchange rate has a similar effect on the euro real exchange rate in terms of sign, although the magnitude of the effect is much larger in this case.[134] These normalised vectors are representative of the long-run elasticity of the real exchange rates for each of the countries in a panel relative to the euro.

The insignificance of the long-run coefficients for Hungary and Slovakia is an interesting finding. This could provide some support to the suggestion of Brissimis, Sideris, and Voumvaki (2005) that long-run PPP is not likely to be observed when economies operate strict exchange rate targeting via intervention. In these cases, exchange rate policy is focused on short-run adjustment to PPP and this observation can be accepted for Hungary and Slovakia, because the short-run coefficients are significantly different from zero. Therefore, while it is the case that all of the transition economies under consideration operated some form of managed exchange rate regime, these results suggest that the policy was perhaps more stringent for Hungary and Slovakia, as compared with Poland and the Czech Republic. Certainly, Hungary has operated a policy for some considerable time of shadowing a basket of currencies including the ECU/euro and the $. And this occurred for external transactions even prior to 1990.

[134] The larger magnitude of the Czech/US real exchange rate on the euro/dollar rate compared to Poland may be related to two issues. Firstly, the Czech koruna is traded in large amounts in London relative to other CEE currencies. Secondly, despite maintaining a floating exchange rate regime, the Czech Republic maintains a wide peg, the basket of which is comprised of about 50% euro. By contrast, the Polish currency is not traded at even close to that of the Czech koruna. In addition, Poland has less scope to maintain a peg dominated by euro due to its high share of dollar-denominated debt and higher relative share of trade with Russia (Egert and Lahreche-Revil, 2003).

Overall, this analysis provides some support for findings identified in the trivariate models and panel cointegration results of the earlier Chapters. More specifically, in no case is a linear time trend found to be significant, suggesting little evidence of a Balassa-Samuelson effect. Otherwise, the drift that occurs by virtue of the Balassa-Samuelson might not be observed, because of the nature of the exchange rate management discussed above.

When G-PPP is considered, then a unique long-run cointegrating equation is detected from the system of real exchange rates. The parameters within this equation seem to make theoretical sense except for the case of Slovakia where the sign is incorrect, but highly insignificant. However, based on the significance of a one sided t-test, the cointegrating vector would seem to feature in each of the short-run equations supporting the idea that the currencies of the Czech Republic, Hungary, Poland and Slovakia were ready to enter the Euro zone, because each was correcting in the short-run to the euro. Further, the conclusion from the tri-variate cointegrated systems using the US dollar as the numeraire currency that neither Hungary nor Slovakia accepts either form of PPP is consistent with the finding from the cointegrating vector from the G-PPP analysis, that the long-run coefficients for Hungary and Slovakia are not significant.

IX. Conclusions

This chapter examines PPP across the EU-27 countries outside of the euro area over the period 1999 to 2006 using the Johansen multivariate cointegration procedure and the Larsson et al (2001) panel cointegration test to assess both the weak and strong versions of PPP. Across the countries considered, the Johansen tests on individual tri-variate systems indicate that weak form PPP holds in ten out of fifteen cases, while strong form PPP holds in four out of fifteen cases. For eight of the countries, these results are not sensitive to the use of either euro or the US dollar as the numeraire and this suggests that for these countries cross rate arbitrage may not be an issue.

With the euro as the numeraire currency, weak form PPP restrictions are not rejected for Sweden, Denmark, Poland, Slovakia, Slovenia, Lithuania, Bulgaria, Romania, Cyprus and Malta. These results are encouraging in terms of an enlarging euro area, whereby some form of stable exchange rate dynamic in relation to the forthcoming entrants is a favourable monetary characteristic. The latter transition economies are scheduled to enter the euro area at some point over the next decade or so.[135] It should be borne in mind that the results found should be interpreted as recommending euro area entry. Rather, they should be viewed amongst a broader body of research on the issue and assessed accordingly in relation only to the PPP relationships evident.

Thus, these results would appear to add weight to the argument that these countries in particular are well suited to membership of EMU based on their exchange rate and price dynamics relative to the euro and euro area prices. Analysis of the short-run adjustment mechanism reveals the driver of the convergence to exchange rate equilibrium. The 'driving' variables vary by country depending on the nature of exchange rate regime in place. Where the nominal exchange rate is fixed (or some variant of a fixed regime), domestic prices cause the adjustment to PPP. Where a floating regime is in place (in conjunction with inflation targeting), changes in the exchange rate account for PPP convergence. These findings conform to what one would expect to drive the PPP process.

[135] Indeed, Slovenia joined as of January 2007, Cyprus and Malta joined in January 2008, and Slovakia joined in January 2009. Indeed, it is worth pointing out that the euro-based PPP results in Table 5.4 (which were conducted before the entry of these countries to the euro area) suggest that these countries were in fact suitable for entry.

The panel cointegration results yield evidence of one cointegrating relationship across all four of the panel groupings tested. These relationships, however, do not appear to be suggestive of PPP holding in the panel context. Considering these results, it could be the case that the countries where PPP was rejected on an individual basis dominate the panel outcome. In other words, the presence of a few countries where the rejection of PPP is strong drives the overall rejection of PPP at the level of the panel (see Beirne, 2009, for further details).

Of the former CEECs examined, perhaps the most surprising results are for Poland and Romania (where weak form PPP holds) and Estonia (where weak form PPP was rejected). Poland and Romania currently operate alternative forms of floating exchange rate regimes[136], while Estonia operates a currency board arrangement pegged to the euro. Under a fixed exchange rate regime, one would expect a strong relationship between the nominal exchange rate and prices, and indeed a high exchange rate pass-through to prices (e.g. Beirne and Bijsterbosch, 2009). It follows that PPP would be expected to hold under a fixed regime.[137] Thus, on this basis at least, it would perhaps be expected that PPP would be rejected for Poland and Romania, and not rejected for Estonia.[138]

Of the other CEECs that did not reject weak form PPP, all operate some form of fixed peg against the euro. For the remaining transition economies where PPP was rejected, the Czech Republic operates a managed float, Hungary maintains a horizontal band, though this has suffered a loss of credibility due to a period of excessive exchange rate volatility, while Latvia maintains a fixed peg where the anchor is the SDR as opposed to

[136] Poland operates a fully independent float while Romania operates a managed float.

[137] However, the precise nature of the regime in place and whether intervention in foreign exchange markets takes place can impact upon this. See Brissimis, Sideris and Voukvaki (2005) for some empirical evidence that long-run PPP is more likely to fail when countries intervene in the market to support an exchange rate rule.

[138] A possible explanation for the validation of PPP in the cases of Poland and Romania may be related to the effectiveness of inflation targeting policies.

the euro. Thus, the results for these countries are as one might expect. Nonetheless, there are of course other reasons beyond the form of exchange rate regime that may cause deviations from long-run PPP, notably shocks to productivity and a strong non-tradable sector, although these issues lie outside the scope of this chapter, and have not been pursued at this point due to the apparent insignificance of a time trend variable included in the VARs.

The results for Sweden and Denmark are noteworthy, because they provide some form of *prima facie* evidence that these countries may be suitable members of the euro area should they decide to make a commitment to join at some point in the future. This is particularly interesting in the Swedish case where its currency freely floats and no attempt has been made by Sweden to fulfil the exchange rate stability criteria that governs euro zone entry.

It also appears that the results presented here suggest that the entry in 2007 of a selection of recent EU members appears to be warranted on the basis of the G-PPP theory. Using the Johansen multivariate cointegration technique, it is shown that one cointegrating vector exists in the system comprising the US real exchange rates of the euro area, Poland, Slovakia, Hungary, and the Czech Republic. Though these results need to be interpreted with some caution, because the long-run coefficients for Hungary and Slovakia are not significant and that for Slovakia has the wrong sign. This is perhaps indicative of the strictly managed exchange rate regimes in place in these countries.

Appendix 1

Table A5.1 Unit Root Tests with Intercept only

Country	ADF		DFGLS		KPSS	
	Level	1st Diff.	Level	1st Diff.	Level	1st Diff.
Euro Exchange Rates						
Bulgaria	-8.34	-10.44	-2.50	-10.32	0.43	0.13
Cyprus	-6.58	-9.73	-2.47	-9.77	0.48	0.14
Czech Republic	-1.77	-10.68	0.96	-0.97	1.16	0.07
Denmark	-0.35	-9.68	-1.74	-2.74	0.44	0.07
Estonia	-8.16	-9.63	-1.76	-9.00	0.51	0.19
Hungary	-2.67	-9.31	-2.39	-6.78	0.16	0.06
Latvia	-0.99	-11.86	-1.04	-9.56	0.98	0.37
Lithuania	-3.48	-11.23	-0.42	-0.49	0.84	0.44
Malta	-2.32	-14.11	-2.02	-13.33	0.72	0.16
Poland	-1.69	-8.28	-1.66	-2.60	0.56	0.08
Romania	-3.73	-9.53	0.49	-1.13	1.07	0.23
Slovakia	0.83	-7.99	0.99	-1.82	1.22	0.30
Slovenia	-2.27	-9.89	0.19	-15.38	1.18	0.46
Sweden	-1.49	-8.86	-1.48	-1.96	0.61	0.09
UK	-1.70	-8.49	-1.06	-3.37	0.73	0.23
Dollar Exchange Rates						
Bulgaria	-0.64	-6.93	-0.83	-4.56	0.89	0.44
Cyprus	-0.93	-6.55	-0.97	-4.69	0.71	0.44
Czech Republic	0.44	-8.07	-0.03	-0.54	1.16	0.43
Denmark	-0.66	-6.86	-0.84	-4.52	0.88	0.45
Estonia	-0.64	-6.88	-0.83	-4.56	0.89	0.45
Hungary	-0.67	-7.39	-0.84	-5.29	0.85	0.53
Latvia	-1.08	-7.12	-1.22	-5.43	0.67	0.23
Lithuania	-0.35	-6.58	0.55	-6.56	1.19	0.15
Malta	-0.50	-7.28	-0.66	-5.07	0.95	0.39
Poland	0.12	-7.84	-0.41	-0.65	1.03	0.43
Romania	-1.62	-5.02	-0.22	-6.55	0.74	0.24
Slovakia	0.13	-6.93	-0.15	-4.47	1.04	0.55
Slovenia	-1.72	-6.79	-0.87	-4.57	0.37	0.67
Sweden	-0.75	-6.96	-0.89	-6.31	0.71	0.40
UK	-0.08	-8.20	-0.20	-7.05	0.89	0.33
CPI						
Bulgaria	-0.75	-7.01	1.40	-5.49	1.26	0.03
Cyprus	-0.58	-4.89	0.92	-0.26	1.16	0.31
Czech Republic	-1.51	-8.59	1.71	-8.16	1.22	0.18
Denmark	-0.57	-6.99	2.37	0.25	1.29	0.45
Estonia	-0.23	-7.79	2.50	-7.81	1.26	0.08
Hungary	-2.24	-7.07	1.95	-0.16	1.27	0.34
Latvia	2.71	-7.63	2.39	-7.33	1.24	0.74
Lithuania	1.65	-8.50	2.41	-8.39	0.93	0.48
Malta	-0.64	-8.47	1.85	-7.61	1.27	0.38
Poland	-3.71	-5.51	0.77	-5.12	1.16	0.41
Romania	-1.96	-6.66	-0.20	-3.12	1.22	0.62
Slovakia	-2.37	-8.32	1.56	-8.07	1.28	0.34
Slovenia	-3.07	-8.32	1.82	-8.27	1.25	1.22
Sweden	-1.12	-8.72	0.88	-0.64	1.25	0.27
UK	1.01	-1.05	5.01	-1.23	1.30	0.33
Euro area	-0.60	-6.85	2.30	-6.89	1.30	0.04
US	-0.63	-8.23	2.51	-7.95	1.28	0.09

(Critical Values at 1%, 5%, and 10% respectively: ADF: -3.50, -2.89, -2.58. DFGLS: -2.59, -1.94, -1.61. KPSS: 0.74, 0.46, 0.35).

Table A5.2 Unit Root Tests with Intercept and Trend

Country	ADF		DFGLS		KPSS	
	Level	1st Diff.	Level	1st Diff.	Level	1st Diff.
Euro Exchange Rates						
Bulgaria	-2.60	-10.38	-2.04	-12.34	0.59	0.09
Cyprus	-2.86	-9.68	-2.87	-9.81	0.56	0.12
Czech Republic	-2.24	-10.62	-2.19	-2.31	0.11	0.06
Denmark	-1.91	-9.63	-1.92	-8.35	0.18	0.06
Estonia	-8.14	-9.59	-7.87	-9.32	0.16	0.15
Hungary	-2.71	-9.26	-2.54	-8.19	0.13	0.05
Latvia	-3.09	-11.95	-1.63	-11.46	0.19	0.20
Lithuania	-3.49	-11.41	-1.72	-10.41	0.29	0.13
Malta	-4.54	-14.07	-2.35	-14.07	0.17	0.11
Poland	-1.67	-8.23	-1.72	-6.82	0.14	0.07
Romania	-0.97	-11.34	-0.01	-5.87	0.33	0.05
Slovakia	-2.62	-8.13	-1.32	-5.41	0.29	0.04
Slovenia	-2.42	-10.09	-1.69	-15.21	0.32	0.11
Sweden	-1.94	-8.81	-1.66	-6.75	0.09	0.09
UK	-3.29	-8.51	-1.38	-7.50	0.14	0.17
Dollar Exchange Rates						
Bulgaria	-2.54	-7.09	-1.21	-6.38	0.19	0.17
Cyprus	-2.38	-6.63	-1.22	-6.20	0.25	0.19
Czech Republic	-2.71	-8.16	-1.12	-4.68	0.18	0.15
Denmark	-2.51	-7.01	-1.21	-6.32	0.19	0.17
Estonia	-2.53	-7.03	-1.21	-6.36	0.19	0.17
Hungary	-2.53	-7.55	-1.17	-6.92	0.17	0.18
Latvia	-2.14	-7.19	-1.56	-6.58	0.15	0.09
Lithuania	-2.12	-6.56	-1.75	-6.56	0.15	0.13
Malta	-2.38	-7.40	-1.26	-6.90	0.19	0.15
Poland	-2.52	-8.10	-0.89	-5.42	0.28	0.06
Romania	-2.09	-6.54	-0.03	-5.17	0.32	0.12
Slovakia	-2.58	-7.25	-1.00	-6.46	0.22	0.17
Slovenia	-2.53	-7.12	-0.95	-6.58	0.23	0.16
Sweden	-2.12	-7.13	-1.21	-7.09	0.19	0.13
UK	-1.98	-8.36	-1.09	-8.13	0.20	0.10
CPI						
Bulgaria	-3.11	-6.97	-3.10	-6.39	0.17	0.03
Cyprus	-4.55	-2.87	-4.19	-1.65	0.08	0.30
Czech Republic	-1.71	-8.67	-1.50	-8.47	0.19	0.06
Denmark	-3.37	-2.21	-1.77	-1.42	0.32	0.09
Estonia	-2.05	-7.74	-2.02	-7.82	0.15	0.08
Hungary	-2.17	-7.41	-0.91	-6.84	0.27	0.07
Latvia	0.97	-8.20	-0.57	-8.24	0.32	0.04
Lithuania	-0.22	-8.86	-0.43	-8.93	0.22	0.12
Malta	-3.29	-8.42	-3.30	-8.24	0.10	0.35
Poland	-2.59	-6.37	-0.66	-6.42	0.24	0.18
Romania	-4.84	-7.58	-1.98	-7.25	0.32	0.29
Slovakia	-2.58	-8.57	-1.14	-8.66	0.20	0.05
Slovenia	-0.38	-9.76	0.21	-8.93	0.33	0.08
Sweden	-1.44	-8.75	-1.42	-1.20	0.26	0.40
UK	-1.46	-1.41	-1.68	-1.22	0.28	0.09
Euro area	-4.86	-6.78	-4.79	-7.49	0.07	0.02
US	-2.44	-8.19	-2.37	-8.17	0.19	0.09

(Critical Values at 1%, 5%, and 10% respectively: ADF: -4.06, -3.46, -3.15. DFGLS: -3.60, -3.05, -2.76. KPSS: 0.22, 0.15, 0.12).

Table A5.3 Misspecification Tests (euro as numeraire)

Country	Variable	$AR_{(1-5)}$	Normality	$ARCH_{(1-5)}$	Hetero
Bulgaria	s	0.69 (0.63)	9.53 (0.01)*	0.06 (0.99)	37.70 (0.86)
	p	0.14 (0.98)	0.27 (0.87)	0.33 (0.89)	41.87 (0.72)
	p*	1.84 (0.12)	1.15 (0.56)	0.52 (0.76)	52.65 (0.29)
	System	0.84 (0.74)	16.14 (0.02)*		259.90 (0.88)
Cyprus	s	2.12 (0.08)	1.58 (0.45)	0.14 (0.98)	0.65 (0.86)
	p	2.44 (0.05)	0.51 (0.78)	0.62 (0.69)	0.32 (0.99)
	p*	1.72 (0.15)	0.48 (0.79)	1.11 (0.37)	0.29 (0.99)
	System	1.39 (0.08)	2.32 (0.89)		0.36 (1.00)
Czech Republic	s	0.18 (0.97)	1.18 (0.55)	0.86 (0.52)	64.41 (0.33)
	p	0.51 (0.77)	1.29 (0.52)	0.55 (0.74)	58.19 (0.54)
	p*	1.71 (0.16)	1.29 (0.52)	0.32 (0.90)	51.10 (0.79)
	System	0.71 (0.89)	5.28 (0.51)		347.14 (0.68)
Denmark	s	0.667 (0.65)	0.64 (0.73)	0.48 (0.79)	0.25 (0.99)
	p	1.51 (0.18)	2.07 (0.35)	0.24 (0.94)	0.48 (0.97)
	p*	1.64 (0.16)	0.45 (0.80)	0.38 (0.86)	0.41 (0.99)
	System	1.30 (0.13)	1.73 (0.94)		0.37 (1.00)
Estonia	s	0.23 (0.95)	3.12 (0.21)	0.31 (0.91)	52.39 (0.75)
	p	2.35 (0.06)	5.36 (0.07)	0.27 (0.93)	58.12 (0.54)
	p*	0.50 (0.77)	4.24 (0.12)	0.46 (0.81)	51.85 (0.76)
	System	0.86 (0.71)	11.29 (0.08)		337.98 (0.79)
Hungary	s	0.36 (0.87)	3.29 (0.19)	0.45 (0.81)	60.57 (0.46)
	p	2.09 (0.09)	1.71 (0.96)	0.06 (0.99)	59.33 (0.50)
	p*	0.59 (0.70)	0.32 (0.85)	0.52 (0.76)	58.42 (0.53)
	System	0.83 (0.75)	2.87 (0.81)		367.80 (0.38)
Latvia	s	0.59 (0.73)	0.18 (0.94)	0.63 (0.68)	50.34 (0.81)
	p	0.40 (0.84)	0.92 (0.63)	0.25 (0.94)	64.44 (0.32)
	p*	1.77 (0.14)	0.71 (0.70)	0.26 (0.93)	52.36 (0.75)
	System	0.69 (0.91)	1.62 (0.95)		327.16 (0.89)
Lithuania	s	1.62 (0.18)	1.10 (0.58)	0.34 (0.89)	58.79 (0.33)
	p	1.50 (0.21)	5.19 (0.07)	0.25 (0.94)	49.77 (0.64)
	p*	0.49 (0.78)	0.22 (0.89)	0.11 (0.99)	49.79 (0.64)
	System	1.31 (0.14)	8.68 (0.19)		315.48 (0.62)
Malta	s	0.63 (0.68)	1.81 (0.40)	0.42 (0.83)	0.31 (0.99)
	p	1.73 (0.15)	2.81 (0.25)	1.34 (0.27)	0.10 (1.00)
	p*	1.80 (0.13)	0.95 (0.62)	0.46 (0.81)	0.13 (1.00)
	System	0.84 (0.74)	5.04 (0.54)		0.09 (1.00)
Poland	s	1.67 (0.15)	0.99 (0.61)	0.49 (0.78)	34.85 (0.92)
	p	0.98 (0.44)	1.29 (0.52)	0.37 (0.869	39.64 (0.79)
	p*	0.24 (0.94)	3.29 (0.19)	1.04 (0.40)	35.87 (0.90)
	System	1.17 (0.26)	8.52 (0.20)		242.87 (0.98)
Romania	s	0.53 (0.75)	3.73 (0.16)	0.17 (0.97)	54.79 (0.23)
	p	2.49 (0.05)	0.14 (0.93)	0.41 (0.84)	51.09 (0.35)
	p*	1.95 (0.11)	0.07 (0.97)	0.37 (0.87)	50.45 (0.38)
	System	1.24 (0.19)	3.68 (0.72)		304.37 (0.24)
Slovakia	s	0.57 (0.72)	4.46 (0.11)	0.87 (0.51)	0.15 (1.80)
	p	0.93 (0.47)	1.37 (0.50)	0.55 (0.74)	0.59 (0.86)
	p*	1.49 (0.21)	1.31 (0.52)	0.42 (0.83)	0.11 (1.00)
	System	0.93 (0.61)	4.63 (0.59)		0.11 (1.00)
Slovenia	s	1.76 (0.14)	4.65 (0.10)	0.60 (0.70)	1.09 (0.40)
	p	0.57 (0.73)	11.71 (0.01)*	0.21 (0.96)	0.63 (0.88)
	p*	1.75 (0.14)	1.22 (0.54)	0.30 (0.91)	0.31 (0.99)
	System	1.19 (0.23)	14.94 (0.03)*		0.06 (1.50)
Sweden	s	1.08 (0.38)	4.66 (0.10)	0.04 (0.99)	0.49 (0.97)
	p	1.09 (0.37)	11.22 (0.01)*	0.41 (0.84)	0.46 (0.98)
	p*	1.65 (0.16)	0.70 (0.70)	0.48 (0.79)	0.55 (0.94)
	System	1.22 (0.19)	16.04 (0.02)*		0.53 (1.00)
UK	s	1.76 (0.14)	1.54 (0.46)	0.37 (0.87)	62.08 (0.21)
	p	0.92 (0.48)	0.36 (0.84)	0.37 (0.86)	61.29 (0.23)
	p*	1.99 (0.10)	0.39 (0.82)	0.41 (0.84)	52.65 (0.53)
	System	1.02 (0.46)	1.71 (0.94)		351.57 (0.14)

Table A5.4 Misspecification Tests (US dollar as numeraire)

Country	Variable	$AR_{(1-5)}$	Normality	$ARCH_{(1-5)}$	Hetero
Bulgaria	s	0.63 (0.68)	3.42 (0.18)	0.46 (0.81)	58.82 (0.52)
	p	0.16 (0.98)	4.70 (0.10)	0.11 (0.99)	53.51 (0.71)
	p*	0.59 (0.71)	3.78 (0.15)	0.39 (0.85)	66.77 (0.26)
	System	0.63 (0.95)	11.36 (0.08)		373.24 (0.30)
Cyprus	s	1.52 (0.20)	0.50 (0.78)	0.64 (0.67)	0.68 (0.83)
	p	1.34 (0.27)	2.82 (0.24)	0.52 (0.76)	0.30 (0.99)
	p*	1.42 (0.24)	20.76 (0.00)*	0.65 (0.66)	0.65 (0.85)
	System	1.16 (0.26)	24.79 (0.00)*		0.32 (1.50)
Czech Republic	s	0.59 (0.71)	1.01 (0.60)	1.13 (0.36)	56.95 (0.37)
	p	1.34 (0.27)	2.34 (0.31)	0.03 (0.99)	52.40 (0.54)
	p*	1.74 (0.15)	3.75 (0.15)	0.92 (0.48)	47.31 (0.73)
	System	1.17 (0.26)	5.90 (0.43)		329.17 (0.41)
Denmark	s	0.43 (0.82)	4.35 (0.11)	0.42 (0.83)	41.37 (0.90)
	p	0.55 (0.74)	3.36 (0.19)	0.42 (0.83)	49.83 (0.64)
	p*	0.36 (0.87)	2.70 (0.26)	0.46 (0.80)	55.60 (0.41)
	System	0.73 (0.87)	6.94 (0.33)		324.78 (0.48)
Estonia	s	0.43 (0.82)	5.09 (0.08)	0.51 (0.77)	0.84 (0.65)
	p	0.82 (0.54)	5.77 (0.06)	0.19 (0.96)	0.56 (0.91)
	p*	1.13 (0.36)	5.25 (0.07)	1.23 (0.31)	0.62 (0.86)
	System	0.80 (0.80)	6.91 (0.37)		0.07 (1.50)
Hungary	s	0.70 (0.63)	2.50 (0.29)	0.92 (0.48)	0.71 (0.81)
	p	1.90 (0.11)	9.08 (0.01)*	0.25 (0.94)	0.55 (0.94)
	p*	0.55 (0.73)	1.97 (0.36)	1.05 (0.40)	0.48 (0.97)
	System	0.94 (0.58)	21.48 (0.01)*		0.45 (1.00)
Latvia	s	0.57 (0.72)	2.11 (0.35)	0.98 (0.44)	0.48 (0.97)
	p	1.31 (0.28)	0.51 (0.77)	1.33 (0.27)	0.87 (0.64)
	p*	1.40 (0.24)	2.82 (0.23)	1.10 (0.37)	0.78 (0.74)
	System	0.80 (0.80)	6.27 (0.39)		0.49 (1.00)
Lithuania	s	1.64 (0.17)	0.86 (0.65)	1.01 (0.43)	0.15 (0.99)
	p	1.18 (0.32)	9.32 (0.01)*	0.51 (0.77)	0.24 (0.99)
	p*	0.44 (0.82)	6.05 (0.05)	0.71 (0.62)	0.20 (0.99)
	System	1.15 (0.28)	18.67 (0.01)*		0.08 (1.50)
Malta	s	0.99 (0.43)	0.84 (0.66)	0.16 (0.97)	0.27 (0.99)
	p	1.20 (0.32)	0.33 (0.85)	0.68 (0.64)	0.30 (0.99)
	p*	2.44 (0.05)	4.43 (0.11)	1.22 (0.32)	0.32 (0.99)
	System	1.08 (0.37)	5.63 (0.47)		0.28 (1.00)
Poland	s	0.38 (0.86)	2.75 (0.25)	0.30 (0.91)	0.23 (0.99)
	p	0.89 (0.50)	1.24 (0.54)	0.62 (0.68)	0.35 (0.99)
	p*	0.79 (0.56)	1.67 (0.44)	0.69 (0.64)	0.41 (0.99)
	System	1.42 (0.07)	10.18 (0.12)		0.28 (1.00)
Romania	s	0.36 (0.87)	6.11 (0.05)	0.16 (0.98)	0.43 (0.98)
	p	1.09 (0.38)	0.20 (0.90)	0.40 (0.84)	0.62 (0.89)
	p*	1.19 (0.32)	5.51 (0.06)	0.95 (0.46)	0.35 (0.99)
	System	0.98 (0.51)	8.60 (0.20)		0.35 (1.00)
Slovakia	s	1.05 (0.41)	1.61 (0.45)	0.56 (0.739	68.47 (0.09)
	p	0.68 (0.64)	1.03 (0.57)	0.24 (0.94)	30.72 (0.99)
	p*	1.55 (0.20)	5.92 (0.05)	0.61 (0.70)	56.28 (0.39)
	System	1.44 (0.08)	5.25 (0.51)		311.57 (0.68)
Slovenia	s	0.41 (0.84)	1.91 (0.38)	0.78 (0.57)	54.33 (0.25)
	p	0.81 (0.55)	1.18 (0.55)	0.33 (0.89)	52.57 (0.30)
	p*	1.55 (0.19)	2.76 (0.25)	0.76 (0.59)	62.05 (0.08)
	System	0.91 (0.63)	4.38 (0.63)		289.77 (0.46)
Sweden	s	0.67 (0.65)	0.16 (0.92)	0.17 (0.97)	52.75 (0.52)
	p	1.28 (0.29)	4.58 (0.10)	1.08 (0.39)	56.50 (0.38)
	p*	0.90 (0.49)	5.20 (0.07)	0.34 (0.88)	60.71 (0.25)
	System	1.42 (0.09)	5.42 (0.49)		338.60 (0.28)
UK	s	0.46 (0.80)	1.34 (0.51)	0.29 (0.91)	67.51 (0.43)
	p	0.64 (0.67)	1.99 (0.37)	0.43 (0.82)	61.91 (0.62)
	p*	1.80 (0.14)	0.65 (0.72)	0.59 (0.71)	71.75 (0.29)
	System	1.19 (0.25)	7.69 (0.26)		400.85 (0.42)

Table A5.5 Skewness and Excess Kurtosis of Systems with signs of Non-normality

Euro numeraire Systems	Variables	Skewness	Excess Kurtosis
Bulgaria	s p p*	0.397 0.140 -0.120	16.176 3.120 2.277
Slovenia	s p p*	0.468 0.062 -0.095	3.530 2.902 2.473
Sweden	s p p*	0.447 0.373 0.018	3.979 5.220 3.297
US Dollar numeraire Systems	**Variables**	**Skewness**	**Excess Kurtosis**
Cyprus	s p p*	-0.092 -0.298 0.187	2.530 3.384 5.757
Hungary	s p p*	-0.185 0.509 0.189	3.108 4.886 4.531
Lithuania	s p p*	-0.198 0.630 0.337	2.774 3.727 4.176

Table A5.6 Normalisation Test

Euro numeraire Systems	LR Test: Zero Restriction on Exchange Rate in Long-Run Matrix
Bulgaria	8.023 (0.005)
Cyprus	8.086 (0.004)
Czech Republic	11.784 (0.001)
Denmark	4.886 (0.027)
Estonia	13.500 (0.000)
Hungary	2.377 (0.123)
Latvia	2.425 (0.119)
Lithuania	7.496 (0.006)
Malta	3.901 (0.048)
Poland	4.241 (0.039)
Romania	2.738 (0.098)
Slovakia	11.653 (0.001)
Slovenia	4.206 (0.040)
Sweden	7.063 (0.008)
UK	4.886 (0.027)

Note: p-values in parentheses.

Table A5.7 PPP Tests of Joint Restrictions (Weak form PPP) on α Matrix (euro numeraire)

Country	Model	α Matrix			Test Statistic
		s	p	p*	$\chi^2(2)/\chi^2(3)/\chi^2(4)$
Sweden	Unrestricted	-0.0013 (0.0016)	0.0006 (0.0004)	0.0014 (0.0002)	
	s fixed	0	0.0006 (0.0004)	0.0013 (0.0002)	41.59* [0.00]
	p fixed	-0.0014 (0.0016)	0	0.0012 (0.0002)	5.73 [0.13]
	p* fixed	-0.0005 (0.0016)	-0.0002 (0.0003)	0	2.58 [0.46]
	s and p fixed	0	0	0.0012 (0.0002)	41.93* [0.00]
	s and p* fixed	0	-0.0002 (0.0003)	0	7.57 [0.06]
	p and p* fixed	-0.0004 (0.0012)	0	0	6.86 [0.14]
Denmark	Unrestricted	0.0002 (0.0001)	0.0010 (0.0003)	0.0015 (0.0003)	
	s fixed	0	0.0010 (0.0003)	0.0014 (0.0003)	32.84* [0.00]
	p fixed	0.0001 (0.0001)	0	0.0010 (0.0002)	14.13* [0.00]
	p* fixed	0.0001 (0.0000)	0.0000 (0.0002)	0	4.09 [0.25]
	s and p fixed	0	0	0.0007 (0.0002)	29.26* [0.00]
	s and p* fixed	0	0.0006 (0.0003)	0	13.69* [0.01]
	p and p* fixed	0.0001 (0.0000)	0	0	4.84 [0.18]
Poland	Unrestricted	-0.0027 (0.0028)	0.0003 (0.0002)	-0.0000 (0.0000)	
	s fixed	0	0.0001 (0.0000)	-0.0000 (0.0000)	9.27* [0.01]
	p fixed	-0.0047 (0.0035)	0	-0.0002 (0.0000)	6.09 [0.05]
	p* fixed	-0.0036 (0.0033)	0.0004 (0.0002)	0	5.33 [0.07]
	s and p fixed	0	0	0.0000 (0.0000)	18.29* [0.00]
	s and p* fixed	0	0.0002 (0.0000)	0	12.06* [0.01]
	p and p* fixed	-0.0143 (0.0062)	0	0	7.54 [0.06]
Slovakia	Unrestricted	-0.0017 (0.0032)	-0.0011 (0.0006)	-0.0007 (0.0001)	
	s fixed	0	-0.0012 (0.0006)	-0.0007 (0.0001)	1.40 [0.70]
	p fixed	-0.0023 (0.0029)	0	-0.0006 (0.0000)	30.73* [0.00]
	p* fixed	0.0018 (0.0013)	-0.0006 (0.0002)	0	4.59 [0.20]
	s and p fixed	0	0	-0.0006 (0.0001)	39.59* [0.00]
	s and p* fixed	0	-0.0003 (0.0009)	0	3.59 [0.46]
	p and p* fixed	-0.0012 (0.0009)	0	0	36.98* [0.00]
Slovenia	Unrestricted	0.0191 (0.0169)	-0.0038 (0.0012)	0.0004 (0.0005)	
	s fixed	0	-0.0050 (0.0015)	0.0004 (0.0006)	1.91 [0.38]
	p fixed	0.0033 (0.0016)	0	0.0000 (0.0000)	11.66* [0.00]
	p* fixed	0.0200 (0.0181)	-0.0047 (0.0011)	0	1.29 [0.53]
	s and p fixed	0	0	0.0009 (0.0004)	17.37* [0.00]
	s and p* fixed	0	-0.0059 (0.0015)	0	4.58 [0.21]
	p and p* fixed	-0.0311 (0.0099)	0	0	12.35* [0.01]
Lithuania	Unrestricted	-0.0192 (0.0184)	-0.0026 (0.0024)	-0.0037 (0.0007)	
	s fixed	0	-0.0028 (0.0024)	-0.0039 (0.0007)	8.37* [0.04]
	p fixed	-0.0208 (0.0185)	0	-0.0035 (0.0007)	39.99* [0.00]
	p* fixed	-0.0053 (0.0115)	0.0031 (0.0013)	0	8.23* [0.04]
	s and p fixed	0	0	-0.0036 (0.0007)	42.29* [0.00]
	s and p* fixed	0	0.0022 (0.0009)	0	5.08 [0.19]
	p and p* fixed	-0.0403 (0.0167)	0	0	10.34* [0.04]
Cyprus	Unrestricted	0.0030 (0.0029)	-0.0071 (0.0035)	-0.0077 (0.0012)	
	s fixed	0	-0.0066 (0.0034)	-0.0078 (0.0012)	1.69 [0.64]
	p fixed	0.0023 (0.0029)	0	-0.0074 (0.0012)	41.32* [0.00]
	p* fixed	0.0026 (0.0025)	-0.0061 (0.0029)	0	5.84 [0.12]
	s and p fixed	0	0	-0.0075 (0.0012)	45.35* [0.00]
	s and p* fixed	0	-0.0048 (0.0025)	0	6.71 [0.15]
	p and p* fixed	0.0039 (0.0029)	0	0	42.33* [0.00]
Malta	Unrestricted	-0.0101 (0.0098)	0.0142 (0.0065)	-0.0025 (0.0019)	
	s fixed	0	0.0121 (0.0056)	-0.0024 (0.0016)	1.89 [0.39]
	p fixed	-0.0094 (0.0099)	0	-0.0031 (0.0019)	7.61* [0.02]
	p* fixed	-0.0134 (0.0108)	0.0171 (0.0071)	0	2.83 [0.24]
	s and p fixed	0	0	-0.0023 (0.0013)	11.07* [0.01]
	s and p* fixed	0	0.0144 (0.0060)	0	5.10 [0.16]
	p and p* fixed	-0.0203 (0.0166)	0	0	8.86* [0.03]

Table A5.7 PPP Tests of Joint Restrictions (Weak form PPP) on α Matrix (euro numeraire) (contd.)

Country	Model	α Matrix			Test Statistic
		s	p	p*	$\chi^2(2)/\ \chi^2(3)/\ \chi^2(4)$
Bulgaria	Unrestricted	-2.0594 (0.4986)	3.9519 (3.1451)	0.4709 (0.5493)	
	s fixed	0	8.7342 (2.5191)	0.6747 (0.4885)	5.98 [0.05]
	p fixed	-2.3620 (0.4553)	0	0.3370 (0.5460)	31.23* [0.00]
	p* fixed	2.1251 (0.4947)	3.3985 (3.0902)	0	4.38 [0.11]
	s and p fixed	0	0	0.1181 (0.2247)	15.56* [0.00]
	s and p* fixed	0	1.6975 (0.9345)	0	4.32 [0.23]
	p and p* fixed	-1.1001 (0.3946)	0	0	10.99* [0.01]
Romania	Unrestricted	0.0022 (0.0063)	0.0074 (0.0015)	-0.0005 (0.0005)	
	s fixed	0	0.0076 (0.0016)	-0.0006 (0.0006)	33.42* [0.00]
	p fixed	-0.0179 (0.0177)	0	-0.0016 (0.0015)	0.93 [0.63]
	p* fixed	0.0022 (0.0061)	0.0072 (0.0015)	0	2.40 [0.30]
	s and p fixed	0	0	-0.0008 (0.0011)	21.71* [0.00]
	s and p* fixed	0	0.0059 (0.0016)	0	15.75* [0.00]
	p and p* fixed	-0.0312 (0.0141)	0	0	2.20 [0.53]

Notes: * indicates rejection of the null at the 5% level or below; standard errors are denoted in brackets, while the p-values associated with the chi-squared test statistics are denoted in squared brackets. The joint restrictions imposed are (Rows 2 to 4): $\beta(1,1)=1$, $\beta(1,2)=-\beta(1,3)$; $\beta(1,1)=1$, $\beta(1,2)=-\beta(1,3)$, $\alpha(1,1)=0$; $\beta(1,1)=1$, $\beta(1,2)=-\beta(1,3)$, $\alpha(2,1)=0$; $\beta(1,1)=1$, $\beta(1,2)=-\beta(1,3)$, $\alpha(3,1)=0$. The 'Unrestricted' row refers to estimates of short-run adjustment with restrictions in place for the β matrix only, i.e. α is unrestricted in this row. Rows 2 to 4 test jointly the β and relevant α restrictions, while rows 5 to 7 impose joint restrictions on the α coefficients.

Table A5.8 PPP Tests of Joint Restrictions (Strong form PPP) on α Matrix (euro numeraire)

Country	Model	α Matrix			Test Statistic
		s	p	p*	$\chi^2(3)$/ $\chi^2(4)$/ $\chi^2(5)$
Sweden	Unrestricted	-0.0012 (0.0016)	0.0006 (0.0004)	0.0013 (0.0002)	
	s fixed	0	0.0006 (0.0004)	0.0013 (0.0002)	41.61* [0.00]
	p fixed	-0.0014 (0.0015)	0	0.0012 (0.0002)	2.91 [0.57]
	p* fixed	-0.0004 (0.0015)	-0.0002 (0.0003)	0	6.06 [0.19]
	s and p fixed	0	0	0.0012 (0.0002)	41.99* [0.00]
	s and p* fixed	0	-0.0002 (0.0003)	0	41.68* [0.00]
	p and p* fixed	-0.0003 (0.0015)	0	0	7.14 [0.21]
Denmark	Unrestricted	0.0002 (0.0001)	0.0010 (0.0003)	0.0015 (0.0003)	
	s fixed	0	0.0010 (0.0003)	0.0015 (0.0003)	33.68* [0.00]
	p fixed	0.0001 (0.0001)	0	0.0010 (0.0002)	14.15* [0.01]
	p* fixed	0.0001 (0.0001)	0.0000 (0.0003)	0	4.26 [0.37]
	s and p fixed	0	0	0.0010 (0.0002)	35.50* [0.00]
	s and p* fixed	0	0.0000 (0.0003)	0	16.22* [0.01]
	p and p* fixed	0.0001 (0.0001)	0	0	5.34 [0.38]
Slovakia	Unrestricted	-0.0013 (0.0034)	-0.0013 (0.0006)	-0.0007 (0.0001)	
	s fixed	0	-0.0013 (0.0006)	-0.0007 (0.0001)	1.67 [0.80]
	p fixed	-0.0015 (0.0033)	0	-0.0006 (0.0001)	42.47* [0.00]
	p* fixed	-0.0009 (0.0034)	-0.0003 (0.0006)	0	7.39 [0.12]
	s and p fixed	0	0	-0.0007 (0.0001)	38.86* [0.00]
	s and p* fixed	0	0.0000 (0.0005)	0	5.48 [0.36]
	p and p* fixed	-0.0030 (0.0037)	0	0	12.02* [0.01]
Cyprus	Unrestricted	0.0026 (0.0027)	-0.0054 (0.0033)	-0.0071 (0.0012)	
	s fixed	0	-0.0049 (0.0032)	-0.0072 (0.0012)	5.29 [0.26]
	p fixed	0.0021 (0.0027)	0	-0.0069 (0.0012)	44.84* [0.00]
	p* fixed	0.0033 (0.0027)	-0.0027 (0.0032)	0	7.74 [0.10]
	s and p fixed	0	0	-0.0069 [0.0012]	46.13* [0.00]
	s and p* fixed	0	-0.0022 [0.0032]	0	8.54 [0.13]
	p and p* fixed	0.0030 [0.0027]	0	0	25.46* [0.00]

Notes: * indicates rejection of the null at the 5% level or below; standard errors are denoted in brackets, while the p-values associated with the chi-squared test statistics are denoted in squared brackets. The joint restrictions imposed are (Rows 2 to 4): $\beta(1,1)=1$, $\beta(1,2)=-1$, $\beta(1,3)=1$; $\beta(1,1)=1$, $\beta(1,2)=-1$, $\beta(1,3)=1$, $\alpha(1,1)=0$; $\beta(1,1)=1$, $\beta(1,2)=-1$, $\beta(1,3)=1$, $\alpha(2,1)=0$; $\beta(1,1)=1$, $\beta(1,2)=-1$, $\beta(1,3)=1$, $\alpha(3,1)=0$. The 'Unrestricted' row refers to estimates of short-run adjustment with restrictions in place for the β matrix only, i.e. α is unrestricted in this row. Rows 2 to 4 test jointly the β and relevant α restrictions, while rows 5 to 7 impose joint restrictions on the α coefficients.

Table A5.9 PPP Tests of Joint Restrictions (Weak form PPP) on α Matrix (US dollar numeraire)

Region	Model	α Matrix			Test Statistic
		s	p	p*	$\chi^2(2)$
Sweden	Unrestricted	-0.0023 (0.0560)	0.0184 (0.0055)	-0.0059 (0.0054)	
	s fixed	0	0.0185 (0.0053)	-0.0059 (0.0054)	17.25* [0.00]
	p fixed	-0.0673 (0.0491)	0	-0.0115 (0.0046)	0.01 [0.99]
	p* fixed	-0.0008 (0.0560)	0.0209 (0.0050)	0	2.13 [0.34]
Denmark	Unrestricted	-0.0274 (0.0342)	0.0027 (0.0026)	-0.0076 (0.0037)	
	s fixed	0	0.0032 (0.0025)	-0.0074 (0.0037)	7.37* [0.03]
	p fixed	-0.0368 (0.0330)	0	-0.0094 (0.0032)	1.20 [0.55]
	p* fixed	-0.0215 (0.0339)	0.0052 (0.0022)	0	2.05 [0.36]
Poland	Unrestricted	0.0271 (0.0157)	0.0044 (0.0017)	0.0008 (0.0018)	
	s fixed	0	0.0058 (0.0020)	0.0016 (0.0021)	10.53* [0.01]
	p fixed	0.0017 (0.0056)	0	0.0000 (0.0006)	15.35* [0.00]
	p* fixed	0.0316 (0.0172)	0.0047 (0.0018)	0	6.53* [0.04]
Czech Republic	Unrestricted	-0.0094 (0.0047)	0.0006 (0.0004)	0.0014 (0.0004)	
	s fixed	0	0.0007 (0.0005)	0.0015 (0.0004)	12.17* [0.01]
	p fixed	-0.0103 (0.0046)	0	0.0013 (0.0004)	8.68* [0.03]
	p* fixed	-0.0095 (0.0039)	0.0000 (0.0004)	0	23.82* [0.00]
Slovenia	Unrestricted	0.0288 (0.0129)	0.0066 (0.0020)	0.0009 (0.0015)	
	s fixed	0	0.0152 (0.0043)	-0.0002 (0.0033)	5.34 [0.07]
	p fixed	-0.0623 (0.0201)	0	-0.0023 (0.0022)	11.41* [0.00]
	p* fixed	0.0334 (0.0162)	0.0080 (0.0023)	0	0.82 [0.66]
Lithuania	Unrestricted	-0.0070 (0.0038)	0.0000 (0.0010)	0.0019 (0.0006)	
	s fixed	0	0.0004 (0.0010)	0.0020 (0.0006)	10.76* [0.01]
	p fixed	-0.0070 (0.0037)	0	0.0019 (0.0005)	20.49* [0.00]
	p* fixed	-0.0084 (0.0042)	-0.0015 (0.0010)	0	5.59 [0.13]
Malta	Unrestricted	0.0127 (0.0071)	0.0011 (0.0003)	0.0007 (0.0002)	
	s fixed	0	0.0012 (0.0003)	0.0007 (0.0002)	5.94 [0.11]
	p fixed	0.0127 (0.0068)	0	0.0006 (0.0002)	15.77* [0.00]
	p* fixed	0.0103 (0.0066)	0.0010 (0.0003)	0	5.10 [0.13]
Estonia	Unrestricted	-0.0024 (0.0014)	0.0009 (0.0002)	0.0009 (0.0002)	
	s fixed	0	0.0011 (0.0002)	0.0010 (0.0002)	5.60 [0.13]
	p fixed	-0.0027 (0.0012)	0	0.0004 (0.0001)	22.23* [0.00]
	p* fixed	-0.0006 (0.0002)	0.0000 (0.0000)	0	5.12 [0.12]

Notes: * indicates rejection of the null at the 5% level or below; standard errors are denoted in brackets, while the p-values associated with the chi-squared test statistics are denoted in squared brackets. The joint restrictions imposed are: β(1,1)=1, β (1,2)=-β(1,3); β(1,1)=1, β(1,2)=-β(1,3), α(1,1)=0; β(1,1)=1, β(1,2)=-β(1,3), α(2,1)=0; β(1,1)=1, β(1,2)=-β(1,3), α(3,1)=0. The 'Unrestricted' row refers to estimates of short-run adjustment with restrictions in place for the β matrix only, i.e. α is unrestricted in this row. All other rows test jointly the β and relevant α restrictions.

Table A5.10 PPP Tests of Joint Restrictions (Strong form PPP) on α Matrix (US dollar numeraire)

Region	Model	α Matrix			Test Statistic
		s	p	p*	$\chi^2(3)$
Denmark	Unrestricted	-0.0172 (0.0326)	0.0024 (0.0024)	-0.0067 (0.0035)	
	s fixed	0	0.0027 (0.0023)	-0.0065 (0.0036)	9.93* [0.02]
	p fixed	-0.0252 (0.0313)	0	-0.0083 (0.0031)	4.24 [0.24]
	p* fixed	-0.0109 (0.0321)	0.0046 (0.0021)	0	5.43 [0.14]
Cyprus	Unrestricted	0.0051 (0.0123)	0.0000 (0.0024)	-0.0049 (0.0014)	
	s fixed	0	0.0003 (0.0024)	-0.0049 (0.0014)	5.61 [0.23]
	p fixed	0.0151 (0.0122)	0	-0.0049 (0.0014)	19.54* [0.00]
	p* fixed	0.0139 (0.0122)	0.0011 (0.0024)	0	3.46 [0.48]
Malta	Unrestricted	0.0122 (0.0068)	0.0010 (0.0003)	0.0007 (0.0002)	
	s fixed	0	0.0010 (0.0003)	0.0007 (0.0002)	8.01 [0.09]
	p fixed	0.0131 (0.0068)	0	0.0006 (0.0002)	15.82* [0.00]
	p* fixed	0.0124 (0.0068)	0.0008 (0.0003)	0	5.11 [0.14]
Estonia	Unrestricted	-0.0023 (0.0016)	0.0010 (0.0002)	0.0010 (0.0002)	
	s fixed	0	0.0011 (0.0002)	0.0010 (0.0002)	5.61 [0.23]
	p fixed	-0.0027 (0.0016)	0	0.0006 (0.0002)	23.10* [0.00]
	p* fixed	-0.0021 (0.0016)	0.0004 (0.0002)	0	5.78 [0.22]

Notes: * indicates rejection of the null at the 5% level or below; standard errors are denoted in brackets, while the p-values associated with the chi-squared test statistics are denoted in squared brackets. The joint restrictions imposed are: β(1,1)=1, β(1,2)=-1, β(1,3)=1; β(1,1)=1, β(1,2)=-1, β(1,3)=1, α(1,1)=0; β(1,1)=1, β(1,2)=-1, β(1,3)=1, α(2,1)=0; β(1,1)=1, β(1,2)=-1, β(1,3)=1, α(3,1)=0. The 'Unrestricted' row refers to estimates of short-run adjustment with restrictions in place for the β matrix only, i.e. α is unrestricted in this row. All other rows test jointly the β and relevant α restrictions.

Figure A5.1 (Log) Real Exchange Rates – Euro area and Four Forthcoming Entrants, 1999-2006

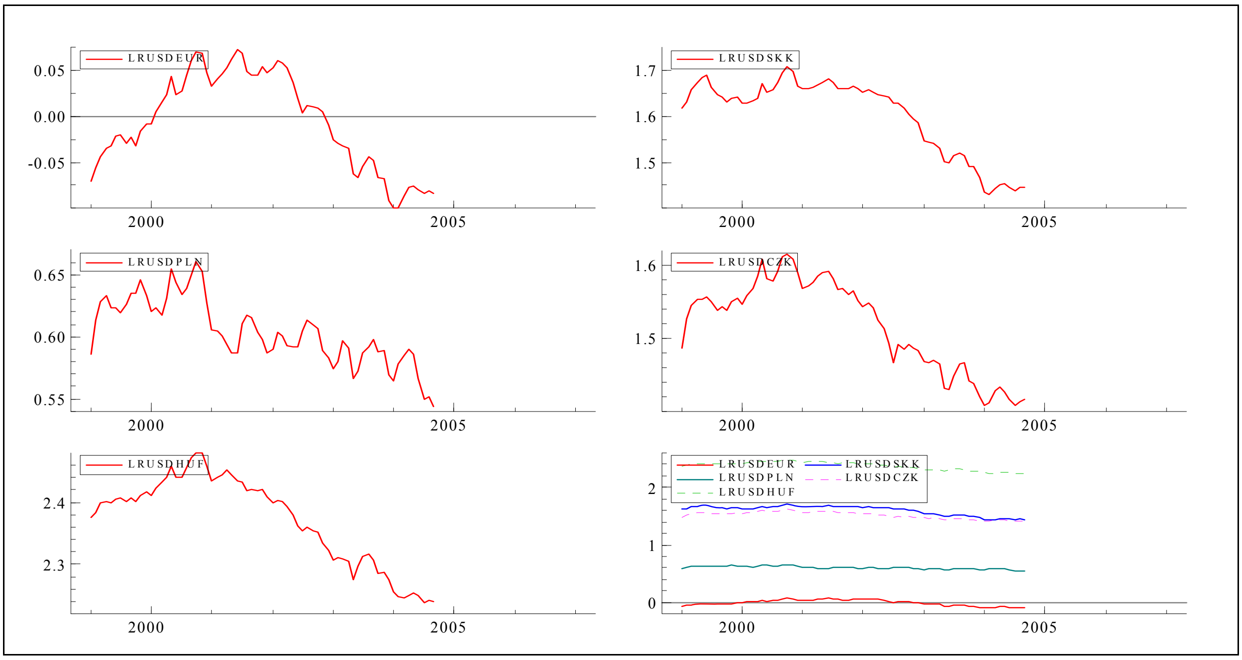


Table A5.11 ADF Tests for Unit Roots

(Log) Real Exchange Rate Series	Model 1 Intercept and Trend		Model 2 Intercept, No Trend		Model 3 No Intercept, No Trend	
	Levels	First Differences	Levels	First Differences	Levels	First Differences
LRUSDEUR (euro)	-2.47 (1) [1.91]*** {2.52}**	-6.89 (0) [0.82] {-1.09}	-0.98 (1) [3.39]*	-6.81 (0) [-0.29]	-0.78 (1)	-6.84 (0)
LRUSDPLN (Polish zloty)	-2.97 (5) [2.96]* {-2.92}*	-7.81 (1) [-0.06] {-0.52}	-0.73 (2) [0.64]	-7.84 (1) [-1.07]	-1.13 (2)	-7.76 (1)
LRUSDSKK (Slovak koruna)	-2.47 (1) [2.46]** {-2.54}**	-6.89 (0) [-0.09] {-0.69}	-0.33 (1) [0.24]	-6.88 (0) [-1.39]	-1.41 (1)	-6.71 (0)
LRUSDHUF (Hungarian forint)	-1.88 (5) [1.86]*** {-1.66}	-6.99 (1) [-0.36] {-0.29}	-0.49 (2) [0.44]	-7.03 (1) [-1.28]	-1.29 (2)	-7.60 (0)
LRUSDCZK (Czech koruna)	-2.89 (1) [2.89]* {2.96}*	-6.70 (2) [0.02] {-0.94}	-0.23 (2) [0.16]	-6.66 (2) [-1.67]***	-1.41 (2)	-6.92 (1)

Notes: ADF Critical Value (5%) is –3.46 for Model 1, –2.89 for Model 2, and –1.94 for Model 3. Beneath the ADF test statistics, the optimal lag lengths are reported in round brackets, the significance of the intercept term is reported in square brackets, and the significance of the trend is reported below this (both t-statistics). *, **, and *** indicates statistical significance at the 1%, 5%, and 10% levels respectively.

Table A5.12 Pantula Principle Test Results

r	n-r	Model 2	Model 3	Model 4
k=5				
0	3	93.6*	73.0*	104.8*
1	2	52.8	38.6	53.7
2	1	30.9	19.3	33.2

Notes: * denotes rejection of the null hypothesis of no cointegration. Model 2 assumes no intercept or trend in either the cointegrating equation (CE) or the VAR; Model 3 allows for an intercept but no trend in the CE or VAR; and Model 4 allows for an intercept and trend in the CE and VAR.

The Pantula methodology would suggest proceeding with Model 4. Thus, the model estimated contains an intercept and trend in the CE and VAR, and the cointegrating rank is 1.[139] The trace and maximum eigenvalue test results are provided below for convenience.

Table A5.13 Misspecification Tests

Variable	$AR_{(1-4)}$	Normality	$ARCH_{(1-4)}$	Hetero
LRUSDEUR	1.40 (0.25)	2.78 (0.25)	0.68 (0.61)	49.57 (0.49)
LRUSDPLN	0.76 (0.56)	2.01 (0.37)	0.62 (0.65)	52.09 (0.39)
LRUSDCZK	2.42 (0.07)	0.98 (0.61)	0.91 (0.47)	48.59 (0.53)
LRUSDHUF	1.87 (0.14)	0.97 (0.62)	1.99 (0.12)	55.44 (0.28)
LRUSDSKK	2.32 (0.08)	3.41 (0.18)	2.12 (0.10)	53.10 (0.36)

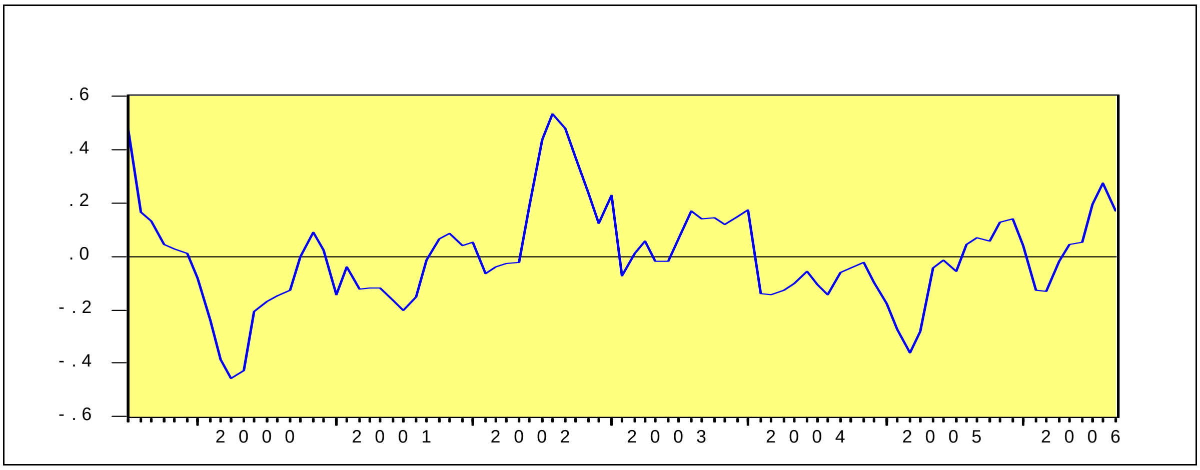
Table A5.14 LR Trace and Max Tests

H_0: rank=p	λ_{trace}	95%	λ_{max}	95%
5 lags				
p=0	104.8*	0.002	51.2*	0.001
p≤1	53.7	0.267	20.4	0.619
p≤2	33.2	0.324	14.9	0.646
p≤3	18.4	0.319	12.9	0.339

Note: * denotes rejection of the null hypothesis of no cointegration.

[139] This is in concurrence with observation of the plots and the unit root tests, which were significant for all series where the model contained an intercept and trend.

Figure A5.2 Cointegrating Relation of the VAR(5) System



CHAPTER SIX

CONCLUSIONS, POLICY IMPLICATIONS AND FUTURE WORK

I. Conclusions

This thesis set out to explore the validity of PPP as a theory of exchange rate determination. Given the eminent role played by the PPP theory in both theoretical and empirical exchange rate and open economy macro models, a failure to endorse that PPP holds would have serious ramifications on the range of models that it underpins. Using primarily linear error correction models, this thesis shows that PPP is valid as a long-run proposition. Specifically, empirical analyses are carried out to show that real exchange rates are stationary; that co-integration can be found in systems comprising the nominal exchange rate, domestic prices, and foreign prices; that the joint modelling of PPP and Uncovered Interest Parity (UIP) produces a stationary outcome; and that the stationarity in systems of real exchange rates is more prevalent for crisis-affected regions in post-crisis periods. Thus, specific empirical scenarios are set out in each chapter to provide a rationale for why PPP may have been found not to hold in previous studies.

PPP has come under some scrutiny of late in terms of its validity and a number of issues have arisen about the coherency of the PPP theory. Firstly short-run day-to-day trading in exchange is dominated by capital flows, suggesting that the exchange rate may deviate from PPP as long as a country's trade deficit is funded. Secondly, it is well known that the nominal exchange rate can be viewed as following a non-stationary time series process or process with a unit root (Baillie and McMahon, 1990).[140] Furthermore, the notion that the time-series process driving the exchange rate has a unit root in discrete time is quite consistent with the theoretical notion of overshooting considered by Dornbusch (1976). Indeed, as noted by Taylor and Taylor (2004), the overshooting

[140] If a time series (y_t) follows a random walk, then $y_t = \phi_0 + \phi_1 y_{t-1} + \varepsilon_t$ has a unit root or ϕ_1=1 and the series is termed I(1) or integrated of order 1.

theory helps to explain why PPP can be retained as a long-run proposition, where deviations from it in the short-run are the result of sticky prices.

Chapter 2 presents a comprehensive review of the literature on PPP and also provides empirical evidence to indicate that real exchange rates are in fact stationary on a univariate basis. The aim of the empirical work in this chapter is to demonstrate that stationarity tests of univariate real exchange rates are in fact consistent with results gleaned from a panel unit root test of the same set of real exchange rates. Panel data methods have been shown to overcome the low power problems associated with ADF tests. A further important conclusion relates to the coherence between univariate and panel estimators of stationarity. By using recursively de-meaned series in the univariate tests across 12 EU real exchange rates (1980Q1-1998Q1), the test results are not substantially different to those of the panel across the majority of the series. The pooling procedure involved with standard panel unit tests of the real exchange rate implies that PPP holds on average. However, this contrasts to the results for unadjusted data based on univariate unit root tests where only three out of twelve real exchange rates are stationary. Based on this, it would not be feasible that the panel of twelve real exchange rates could be stationary on average if only three are stationary on a univariate basis. Using appropriately transformed series, however, we find that nine of the univariate series are stationary, making the panel outcome feasible. The results that we find are reinforced by Hunter and Simpson (2001) who apply the Hadri test on the same set of recursively de-meaned real exchange rates. In the models that we implemented, as well as recursively de-meaning each of the series, careful consideration was also given to the dynamic forcing the real exchange rate. All of the models are well-defined in terms of serial correlation, and non-IID disturbances associated with excess kurtosis and heteroskedasticity are corrected using alternative robust standard errors. The models are also appropriately corrected for ARCH.[141]

The findings in **Chapter 2** add weight to the argument supporting PPP theory and that real exchange rates are stationary in the long-run. The innovation in this chapter lies in the recognition that the misspecification in the univariate test causes the discrepancy in the panel results. The mean-adjustment transformation corrects for initial conditions so that the estimates based on the univariate and panel tests are on a level playing field. The insignificance of the intercept then means that the univariate test becomes coherent in terms of the panel test.

Chapter 3 continues the analysis of real exchange rates, although in this case in a multi-country context. Using the Generalised PPP (G-PPP) theory of Enders and Hurn (1994), the stationarity of blocks of real exchange rates in economic regions affected by a financial crisis is examined. Hence, three economic regions are considered: Europe over the period 1980 to 2006, Latin America over the period 1983 to 2006, and South East Asia over the period 1988 to 2006. Using the Johansen multivariate cointegration technique, long-run stationary relationships were identified in the real exchange rates of the countries according to their regional grouping. The empirical work focussed particularly on the regional exchange rate dynamic in pre- compared to post-crisis periods. The aim was to assess how G-PPP relationships may differ following a major crisis.

The key conclusion from the analysis appears to be that following a crisis, the identification of a long-run equilibrium relationship amongst systems of real exchange

[141] The results from this empirical work were presented at the 3rd International Conference on Advances in Financial Economics at the Research and Training Institute of East Aegean (INEAG), Samos, Greece, in July 2006. The work is published as Beirne et al (2007).

rates in economic regions is more apparent. There also appears to be a faster speed of adjustment towards G-PPP in the post-crisis period, as shown in the short-run coefficient estimates of the Johansen test. Of course, this does not mean that monetary integration should be endorsed. A deeper analysis of the cointegration equations, based on the sign and magnitude of long-run coefficients and the similarity of short-run adjustment coefficients, revealed that monetary integration in the post-crisis case seems appropriate for the former countries of the EMS and a selection of countries in the South East Asian group (specifically Singapore, Indonesia, the Philippines, and Thailand). The Latin American countries do not appear to be suitable for a monetary union arrangement, primarily due to the differences evident in the demand parameters in the countries. In addition, an alternative analysis based on impulse response function analysis to assess how the real exchange rate responds to real output shocks in pre- and post-crisis periods is consistent with the results found from the cointegration analysis.[142]

Chapter 4 argues that it is possible that PPP may require augmentation with an interest rate component in cases where capital markets have become increasingly important. The joint modelling of PPP and UIP is rationalized due to the implicit link that exists between the goods market and the capital market. Moreover, theoretical models by Frydman and Goldberg (2002, 2006) suggest that PPP and UIP cannot be assessed independently of each other. Thus, the implication is that since equilibrium in the balance of payments requires both current and capital account equilibria, balance of payments equilibrium can only hold when PPP and UIP hold simultaneously. In other words, it may very well be the case that PPP and UIP are not independent of each other. In addition, the PPP and

[142] The work undertaken in this chapter was presented at the 6th Annual INFINITI Conference in Trinity College Dublin, June 9-10th 2008, and at the 12th International Conference on Macroeconomic Analysis and International Finance, University of Crete, 29th-31st May, 2008. The paper appears in the December 2008 issue of the *Journal of International and Global Economic Studies* (Beirne, 2008).

UIP conditions are fundamental to a number of theoretical exchange rate models, notably the monetary model of the exchange rate. Thus, the finding that both arbitrage conditions hold simultaneously is an important finding in terms of the validity and usefulness of exchange rate models such as the monetary model.

The chapter firstly examines the joint modelling of PPP and UIP for a five-variate VAR for Germany and Denmark (comprising the nominal exchange rate, and the CPI and interest rate for each country). The results find empirical support for stationarity and the validity of the parity conditions when considered in a single system. In this bipolar case, a single cointegrating relationship was found. A subsequent extension to this model by adding the UK to the system (i.e. an eight-variate VAR) showed that this result regarding the joint modelling of PPP and UIP also holds in a tripolar model. The tripolar model validates PPP and UIP simultaneously particularly for the case where short-term rates enter the system. While there is some evidence of relative PPP plus UIP holding jointly in the long-term rate tripolar model in the case of Denmark, the scale of non-rejection of the restrictions in the short-term rate model is much higher. Overall, it seems that a hypothesis test of PPP plus unrestricted interest rates is most likely to hold for models incorporating long-term interest rates, while a joint restriction of PPP and UIP is most likely to hold when short-terms rates are used in the models. This might suggest that distortions linked with long-term rates are somehow amplified if restrictions for proportionality or symmetry are imposed upon them.

Thus short-term interest rates may be a more appropriate proxy for capital flows given that they are not subject to the same types of distortions that can impinge upon long-term bond yields. The tripolar model identified two cointegrating relationships. As regards identifying the drivers of correcting disequilibria and restoring convergence to stationarity, the bipolar model revealed domestic (i.e. Danish) interest rates as being the key variable. In the tripolar case with short-term interest rates, Danish interest rates would appear to be the primary mechanism for adjustment (certainly in terms of scale and significance). However, the interdependence in the dynamics of the tripolar model suggests that the krone/DM exchange rate, UK interest rates and German prices also react to deviations from the joint PPP-UIP equilibrium. While strong evidence of joint PPP-UIP holding could not be found in the tripolar model with long-term interest rates, an equilibrium for PPP combined with unrestricted interest rates was found to hold strongly for Denmark (and more weakly for the UK). In this case, as with the bipolar model and the tripolar model with short-term rates, a key role appeared to be played by Danish interest rates in reacting to deviations from the identified PPP and interest rate equilibrium relationship. In addition, an interesting finding that holds across both the bipolar and tripolar models was the weak exogeneity of German interest rates. This would seem to make sense intuitively, suggesting high degree of influence by the Deutsche Bundesbank on European financial markets (certainly in the period prior to the formation of the European Central Bank).

Chapter 5 examines PPP across the EU-27 countries outside of the euro area over the period 1999 to 2006 using the Johansen multivariate cointegration procedure and the Larsson et al panel cointegration test to assess both the weak form and strong form versions of PPP. Across the countries considered, the Johansen tests on individual tri-variate systems indicate that weak form PPP holds in ten out of fifteen cases, while strong form PPP holds in four out of fifteen cases. These results hold across both euro and US dollar numeraire currency scenarios in that eight of the countries are fully robust to the choice of numeraire currency. This helps to allay some of the initial concerns regarding cross rate arbitrage.

With the euro as the numeraire currency, relative PPP restrictions are not rejected for Sweden, Denmark, Poland, Slovakia, Slovenia, Lithuania, Bulgaria, Romania, Cyprus and Malta. These results are encouraging in terms of an enlarging euro area, whereby some form of stable exchange rate dynamic in relation to the forthcoming entrants is a favourable monetary characteristic. The latter eight transition economies are scheduled to enter the eurozone at some point over the next decade or so.[143] The panel cointegration results yield evidence of one cointegrating relationship across three panel groupings tested (although there may possibly be two relationships for the panel of all countries). There is evidence to suggest that the panel comprising countries where PPP was not rejected individually validates panel restrictions for weak form PPP and strong form PPP. For the panel of all countries and the panel of transition countries, PPP restrictions are rejected. Considering the results of the panel tests of PPP, it could be the case that the countries where PPP was rejected on an individual basis dominate the panel outcome. In other words, the presence of a few countries where PPP does not hold drives the overall panel rejection of PPP.[144]

The overarching conclusion from the thesis as a whole is that PPP is a valid proposition in a long-run sense. The previous empirical literature on PPP has lacked consensus on whether the theory holds. The extensive analyses carried out in this thesis address this issue, bearing in mind the nature of empirical testing undertaken in previous studies. The

[143] Indeed, Slovenia joined as of January 2007.

[144] A paper-length version of this chapter has been presented at the Money Macro and Finance Research Group 39th Annual Conference, University of Birmingham, UK, 12-14th September, 2007; and the 6th Annual Meeting of the European Economics and Finance Society on *European and Global Integration: Underlying Causes, Issues Arising and Formulating Economic Policies*, Sofia, Bulgaria, 31st May-3rd June, 2007. It was also presented at the 66th Annual Conference of the International Atlantic Economic Society in Montreal, Canada (October 9-12th 2008). The paper has been re-submitted to the *International Review of Applied Economics* journal (second round). In addition, a rationale for the results gleaned from the panel cointegration analysis is published in *International Advances in Economic Research* (Beirne, 2009).

overall conclusion that PPP is valid has a number of implications from a policy perspective, and these are discussed in detail in the following section.

II. Policy Implications

From a policy context, the finding from **Chapter 2** that relative prices and exchange rates converge is an important proposition for monetary policy. For example, whether real exchange rates are stationary or not has implications for the nature of exchange rate regime that might be viewed as being optimal and on whether, and by how much, governments should correct for significant exchange rate misalignments. In addition, it is reassuring from the perspective of economic theory that PPP would appear to hold given that the majority of exchange rate models and open economy macro models incorporate PPP as an equilibrium condition as regards exchange rate determination. A failure of PPP would imply that macro-models of the economy are founded upon an inappropriate exchange rate assumption and that some alternative explanation of exchange rates would be necessary.

Chapter 3 provides a means by which to extrapolate the extent to which exchange rate policy co-ordination and monetary integration is evident. There appears to be a high extent of both for the European economies and a sub-sample of the South East Asian economies. Thus, after their respective crises, the European countries considered and the relevant South East Asian economies seem to have learned that a co-ordinated approach to regional exchange rate policy is favourable in terms of lower volatility and greater stability. This can help to insulate the economies from crises as such a co-ordinated arrangement helps to increase the credibility of the respective currencies, which reduces the likelihood of speculative attack on the currencies. Clearly, in terms of monetary integration, the European countries in the sample are now part of a monetary union within EMU. This would appear to vindicate the conclusions made regarding the analysis of the former EMS countries. The South East Asian economies where monetary integration seems appropriate (i.e. Singapore Indonesia, the Philippines, and Thailand) may engage in some form of currency union in the future. The results from this study indicate that they are well placed to do so in terms of the interaction of their real exchange rates. For the Latin American economies, there has been a lot of variation across the real exchange rates, which was exacerbated by inflation and debt problems. It is not surprising, therefore, that a monetary integration this particular group seems to be inappropriate. The findings from the cointegration analysis are also reinforced by a separate analysis that assesses the response of the real exchange rate across the regions to an unexpected temporary shock to real output using impulse responses from a VAR.

Chapter 4 also infers important policy implications. Firstly, it would appear that a comprehensive understanding of the fluctuations of exchange rates, interest rates, and prices can only be attained when the goods and capital markets are modelled simultaneously, i.e. when the model facilitates an interaction between both markets. This type of finding is reflective of Caporale et al (2001), who make the point that the empirical failure of some exchange rate models based on PPP and UIP may have been due to the failure to account for interactions across exchange rates, interest rates and prices, as well as short- and long-run dynamics. Bearing this in mind, it follows that this type of system-based approach can be very informative in terms of exchange rate management for economies. For example, interactions amongst economies can help to provide information on the most appropriate exchange rate regime. Moreover, should a fixed-but-adjustable regime seem appropriate, the system would help to inform the band width within which the regime should operate.

Overall, the main conclusion from **Chapter 4** is that it would appear that for certain groups of economies, the validation of PPP requires interest rate components to be incorporated into the model. This would appear to be particularly intuitive in the current global economy, where trade flows no longer dominate exchange rate fluctuations (certainly in economies with liberalised capital accounts). The increased level of financial globalization, particularly over the past ten years, has also meant that an increasing role exists for international capital flows in the determination of the exchange rate. An empirical example was provided to illustrate the dynamic between the exchange rates, interest rates, and price levels of Germany, Denmark and the UK. These particular economies are of interest in the context of the potential future direction of the euro area. The results provide evidence to suggest that all three economies are suitably integrated in terms of the integration evident across their respective goods and asset markets. The long-run convergence identified in this system provides some indication that the UK and Denmark may be appropriate for membership of EMU.

Chapter 5 provides clear policy implications as regards whether membership of EMU is feasible for the EU countries that currently remain outside of the euro area. Where PPP holds, this would appear to add weight to the argument that these countries in particular are well suited to membership of EMU based on their exchange rate and price dynamics relative to the euro and euro area prices. Analysis of the short-run adjustment mechanism reveals the driver of the convergence to exchange rate equilibrium. The 'driving' variables vary by country depending on the nature of exchange rate regime in place. Where the nominal exchange rate is fixed (or some variant of a fixed regime), domestic prices cause the adjustment to PPP. Where a floating regime is in place (in conjunction with inflation targeting), changes in the exchange rate account for PPP convergence. These findings conform to what one would expect to drive the PPP process. The results for Sweden and Denmark are noteworthy, providing some form of *prima facie* evidence that these countries may be suitable members of the euro area should they decide to make a commitment to join at some point in the future. This is particularly interesting in the Swedish case where its currency freely floats and no attempt has been made by Sweden to fulfil the exchange rate stability criteria that governs eurozone entry.

III. Future Work

The work that I have completed as part of this thesis has led to a number of research avenues that I am currently pursuing. The empirical work undertaken in the majority of the thesis focuses on US-dollar based exchange rates. This is due to the problem associated with cross exchange rate arbitrage. Following Smith and Hunter (1985), triangular arbitrage means that there are strong restrictions on exchange rate equations that incorporate cross rates. Hunter and Simpson (2004) provide some univariate empirical support for this the theoretical proposition of Smith and Hunter (1985). They show that in a sample of twelve EU countries, the usual Dickey Fuller test applied to cross rate equations suffers from omitted variable bias. Thus, the use of cross rates is likely to lead to misspecification unless the parameters of the cross rates are the same as those based on dollar rates. In order to circumvent this issue, US dollar based real exchange rates were used to demonstrate how coherence can be achieved between panel and univariate unit root tests of the real exchange rate. The expansion of the Hunter and Simpson (2004) work to the multi-variate system context is the subject of research that I am currently undertaking in collaboration with John Hunter.

The findings from the analyses carried out in **Chapter 5** have led to a number of further research avenues which I am now currently undertaking. Firstly, a branch of the literature has cited the role played by productivity gains for the euro area accession economies (e.g. Taylor and Sarno, 2001; Egert, 2002; and Crespo-Cuaresma et al, 2005). Thus, the exchange rate behaviour of these economies may be sensitive to the Balassa-Samuelson effect. Taylor and Sarno (2001) have taken account of this in their work by using a time trend variable to proxy relative productivity gains. It should be noted that the results presented in **Chapters 2-5** appear to give little role to deterministic trends, though an analysis of a broader range of countries might give rise to different results. The analysis of the new and prospective euro area entrants is also hampered by the fact that many of these countries have shadowed the euro and in the past some of these economies have done the same for the ECU. Studies by Engle (1999), Paya et al (2003), Groen and Lombardelli (200), and Betts and Kehoe (2006) have used the ratio of CPI/PPI as a proxy for relative prices.[145]

It is the author's intention to further explore the existence of Balassa-Samuelson effects by extending his current research. As the systems also need extension, then this new work needs a longer span of data based on synthetic euro exchange rates. This is beyond the scope of the current thesis, but it might help to further resolve the PPP and GPPP puzzles observed in particular in **Chapter 5**. This new research would involve exploring alternative panel cointegration tests. In particular when the analysis is limited to a single cointegrating vector, then it is possible that the error correction based tests due to Pedroni (1999 and 2004) might be applied. Though this is less likely when the models need to be augmented by a further variable, such as the interest rates considered in **Chapter 4**, or relative prices as discussed above (alternatively, relative wages might better capture productivity).

[145] Using the CPI/PPI ratio, the assumption is that all non-goods components of the CPI are non-tradables and thus not captured in the PPI. Therefore the ratio reflects the sum of non-tradables plus tradables divided by tradables.

Regarding the work on G-PPP, it would be of interest to analyse the exchange rate relationships in a less restrictive framework. The empirical work in **Chapters 3** and **5** was carried out on real exchange rates, which of course implicitly imposes the proportionality and symmetry conditions associated with PPP. A less restrictive framework could be represented by a system comprising the nominal exchange rate and the ratio of domestic and foreign prices for each country. In this case, only the symmetry condition would be imposed. The most general case (with no restrictions imposed for proportionality or symmetry) could be constructed by including the nominal exchange rate, domestic prices, and foreign prices as separate series in the system. Both of these scenarios, however, are fraught with difficulty due to the large number of variables in the system.

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