Real Exchange Rate Dynamics and Monetary Integration in Crisis-Affected Regions

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Abstract

This paper examines Generalised Purchasing Power Parity (G-PPP) across three regions whose constituent countries are closely linked economically and have been affected by a financial crisis. G-PPP is particularly appropriate when interdependence is high and this is examined using the Johansen multivariate cointegration procedure. The extent of convergence towards G-PPP is assessed in the pre- and 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. This includes assessment of whether a long-run equilibrium relationship between the real exchange rates exists within the three groups and quantifying the speed of adjustment to this equilibrium. The econometric results help to indicate how regional exchange rate policy may have evolved following a major financial shock. In addition, economic implications are set out from a monetary integration perspective.

Keywords: G-PPP, financial crisis, cointegration.

JEL classification: C32, F30, F42.

I. INTRODUCTION

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. This paper focuses on a specific version of the PPP theory which is appropriate for countries have high degrees of economic interdependence, namely Generalised PPP (G-PPP). According to Enders and Hurn (1994), G-PPP can provide an explanation to the non-stationarity in 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 stationary trend in the real exchange rates of a group of countries. Thus, G-PPP permits a test of PPP that goes beyond the traditional two-country test. 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).

The theory is based on the premise that traditional PPP may fail to hold due to non-stationarity in economic fundamentals[1]. This follows Ahn et al (2002) who note that while PPP is useful in terms of explaining competitiveness fluctuations among countries, it is perhaps of limited use for groups of countries that are closely aligned from an economic fundamentals perspective. As drivers of real exchange rates, fundamentals that are non-stationary can cause the real exchange rates to be non-stationary. Therefore, 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 when considered on an individual basis, they may be stationary when considered within a system framework across countries[2].

This paper 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,

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[1] This is consistent with canonical models of exchange rate behaviour such as the Dornbusch (1976) overshooting model.

[2] 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.

the Latin American crisis of December 1994, and the South East Asian crisis of July 1997. Specifically, G-PPP is examined by assessing whether or not a long-run equilibrium relationship exists between the real exchange rates of the respective regions. In order to carry this out, the Johansen cointegration technique is employed to test G-PPP in both pre-crisis and post-crisis scenarios.

This paper, to the knowledge of the author, is the first paper to consider G-PPP in pre- and post-crisis scenarios for the countries of the EMS crisis and those of the Latin American crisis 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 more 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 paper explores these issues in the context of the impact made by the financial crisis.

The paper follows the following structure. Section II provides a brief overview of the context for this paper. Section III describes the econometric methodology to be employed. Section IV provides details on the data used and some preliminary analysis is carried out. Section V sets out the cointegration results. Section VI provides an economic interpretation of the results. Section VII concludes.

II. CONTEXT

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. The context against which this paper 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. 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, the Generalised PPP (G-PPP) is employed as the object test[3]. As will be described in Section III, G-PPP is perhaps a more appropriate exchange rate benchmark for countries that have a high degree of economic interdependence. The original G-PPP theory by Enders and Hurn (1994) 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[4]. The interpretation of this is that this group considered as a whole may be suitable for monetary integration.

Essentially, univariate tests of PPP may indicate the presence of a unit root in the data generating process. Similar outcomes can be observed with bi-variate tests. 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, cointegration techniques are the most appropriate means by which to undertake tests of G-PPP. The fact that this occurs is indicative of a high degree of inter-linkage between the economic fundamentals of the respective countries. In addition, the cointegration of real exchange rates can be viewed as being indicative of a similarity in the fundamentals driving the exchange rates and the overarching exchange rate policy.

Whether or not G-PPP holds has important implications for regional exchange rate policy. If G-PPP is found to hold, this would suggest that some form of monetary integration would be

[3] 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), however, 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 cointegration techniques is felt to provide more comprehensive information on the monetary integration decision.

[4] 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.

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 cited[5]. 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 has 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 studies done previously that examine G-PPP in the context of the euro area. Bernstein (2000) assesses the cointegration of the real exchange rates of a range of European economies over the period 1979 to 1996[6]. He then splits 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 (2000) then tests for cointegration for the US dollar real exchange rates of Germany, the UK, and each other European country, i.e. trivariate cointegration tests. He finds that a long-run equilibrium relationship is found between Germany, the UK, and a number of smaller EU countries (each considered individually)[7]. While these results are interesting, Bernstein (2000) perhaps does not fully exploit the

[5] 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.

[6] 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.

[7] 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.

Johansen procedure for cointegration by limiting each test to just three real exchange rates. In addition, there is no account made for the turbulence of 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 of the cointegrating equation. This would clearly affect the robustness and validity of the conclusions drawn.

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

With regard to Latin America, there has been one recent working paper carried out that incorporates G-PPP. Neves et al (2007) 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). They find that the null of cointegration cannot be rejected. However, their results are difficult to interpret in terms of endorsing monetary integration due to a lack of significance on a number of the coefficients of the cointegrating equation. They also do not present an analysis to show how the dynamic may have changed following a major structural break such as a financial crisis.

Some work has previously been carried out on testing G-PPP for South East Asia, notably Ogawa and Kawasaki (2003) who look at the pre-crisis period only and Choudry (2005) who considers both the pre-crisis and post-crisis scenarios. Using the US dollar as the numeraire, Ogawa and Kawasaki (2003) find cointegration in the case of the 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 12 potential common monetary zones are found to be feasible. Choudry (2005) finds no cointegration of the real exchange rates of crisis affected countries in South East Asia prior to 1997, but evidence of a long-run equilibrium relationship in the system in the period following the crisis. Other examples of where currencies were found to be interdependent include Aggarwal and Mougoue (1993) in a study of Japan, Hong Kong, Malaysia, the Philippines and Singapore. Tse and Ng (1997) also find interdependence but only when the system includes Korea and Taiwan. Other empirical studies done 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 paper builds upon the previous literature on G-PPP in an exploration of real exchange rate behaviour amongst systems of interdependent currencies under 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 against the context that many empirical studies 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 fundamentals are the drivers behind the real rates). G-PPP postulates that a sufficiently high degree of interdependence can result in real exchange rate stationarity (Enders, 1995). Essentially, a bivariate real exchange rate may be deemed to be non-stationary, which would imply a failure of traditional PPP. However, it may very well be the case that changes in the bilateral rate depend on relative prices both in the respective two countries 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 notationally 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 is r_{1it} 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 comovement 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 high degrees 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 is observed. Thus, even though the bilateral real exchange rate in the traditional two-country test of PPP is non-stationary, there can exist a linear combination of a number of non-stationary real exchange rates that is itself stationary in the long-run. The implication is that the real exchange rate, although non-stationary, have a common stochastic trend[8]. Such a common trend could be explained by the fact that output shocks have a symmetrical effect on the real exchange rates (Ogawa and Kawasaki, 2001). This can be described notationally as follows:

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

where the residual term, $\varepsilon_{GPPP,t}$ is stationary.

The β parameters reflect the economic interdependencies within the region. As noted by Enders and Hurn (1994), the more similar the aggregate demand functions in each country of the region, the lower the β's in magnitude.

G-PPP holds when at least one linear combination of bilateral real exchange rates can be found. 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. The specific econometric procedure to be employed to examine G-PPP is based on the Johansen cointegration methodology[9]. As already described, the test is based on assessing 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:

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

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

[8] 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.

[9] See Johansen (1988) and Johansen and Juselius (1990) for more detail on the procedure.

Equation (3) can be denoted as a VEC equation as follows (in first-differenced form):

$$\Delta z_t = \Gamma_1 \Delta z_{t-1} + ... + \Gamma_{k-1} \Delta z_{t-k+1} + \Pi z_{t-k} + \varepsilon_t \tag{4}$$

where Γ_i represents – (I - A_1 - .... – A_i), (i = 1,...., k-1), and Π = - (I – A_1 - .... – 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 $\alpha$ represents the speed of adjustment to equilbrium, and $\beta$ represents the cointegrating vectors (i.e. it is the matrix of long-run coefficients such that the term $\beta$`z_{t-k} represents up to (n-1) cointegration relationships in the multivariate model). Hence, the cointegration test amounts to assessing how many r ≤ (n-1) cointegration vectors exist in β (this is equivalent to testing whether has reduced rank). Equation (4) can also be augmented to include a constant term to capture trending behaviour.

Using the Johansen cointegration procedure, two specific test statistics are provided; one relating to the trace test and the other to 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, i.e. 0 ≤ r ≤ n). The trace test statistic computed as follows:

$$\lambda_{trace} = -T\, \Sigma_{i=r+1}^{n} \ln(1- \lambda_i) \tag{5}$$

where λ_i are the (n-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:

$$\lambda_{max} = -T \ln(1- \lambda_{r+1}) \tag{6}$$

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.

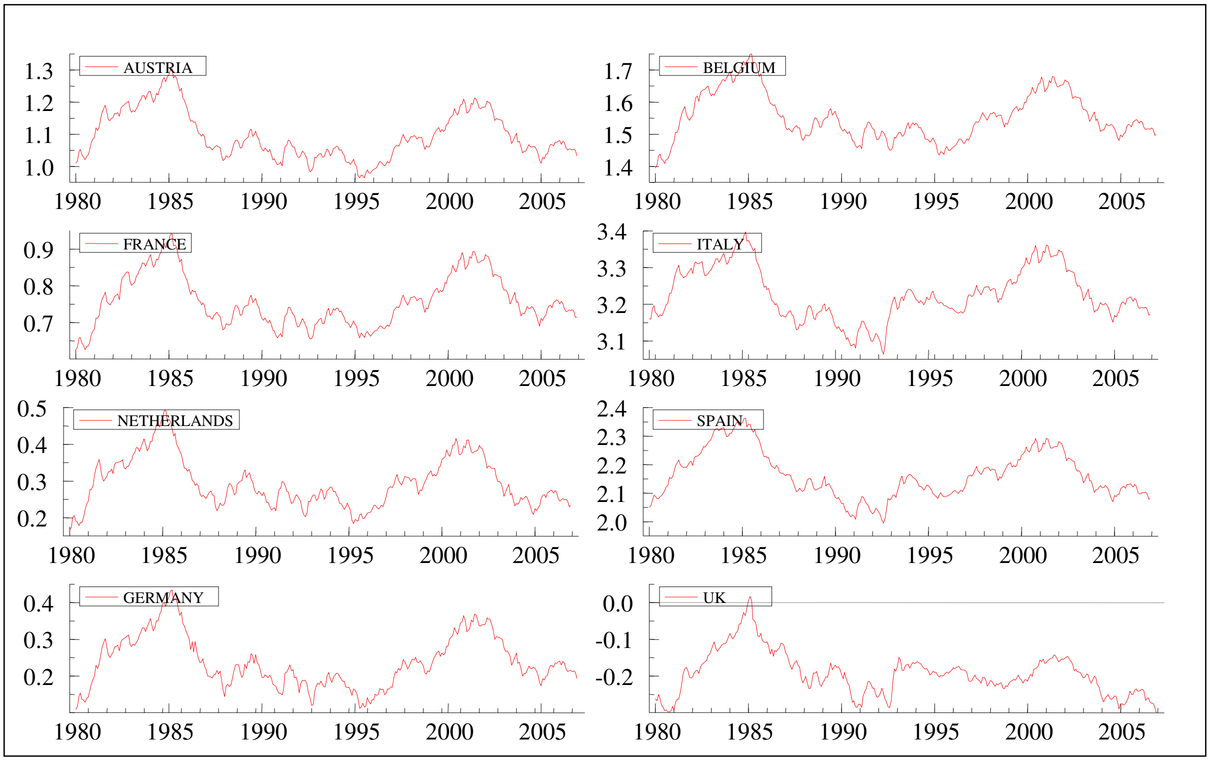
IV. DATA AND PRELIMINARY ANALYSIS

The IMF International Financial Statistics CD-ROM is the source of the data for this study. 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). The time periods were selected so that the crisis occurs at around the mid-point. 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.

Figures 1 to 3 show how the real exchange rates vis-á-vis the US dollar have fluctuated across the currencies of each of the three regions under consideration[10].

[10] 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.

Figure 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[11]. 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 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 is remarkably similar since 1980.

[11] 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.

Figure 2 *(Log) Real Exchange Rates – Latin American Crisis Countries, 1983-2006*

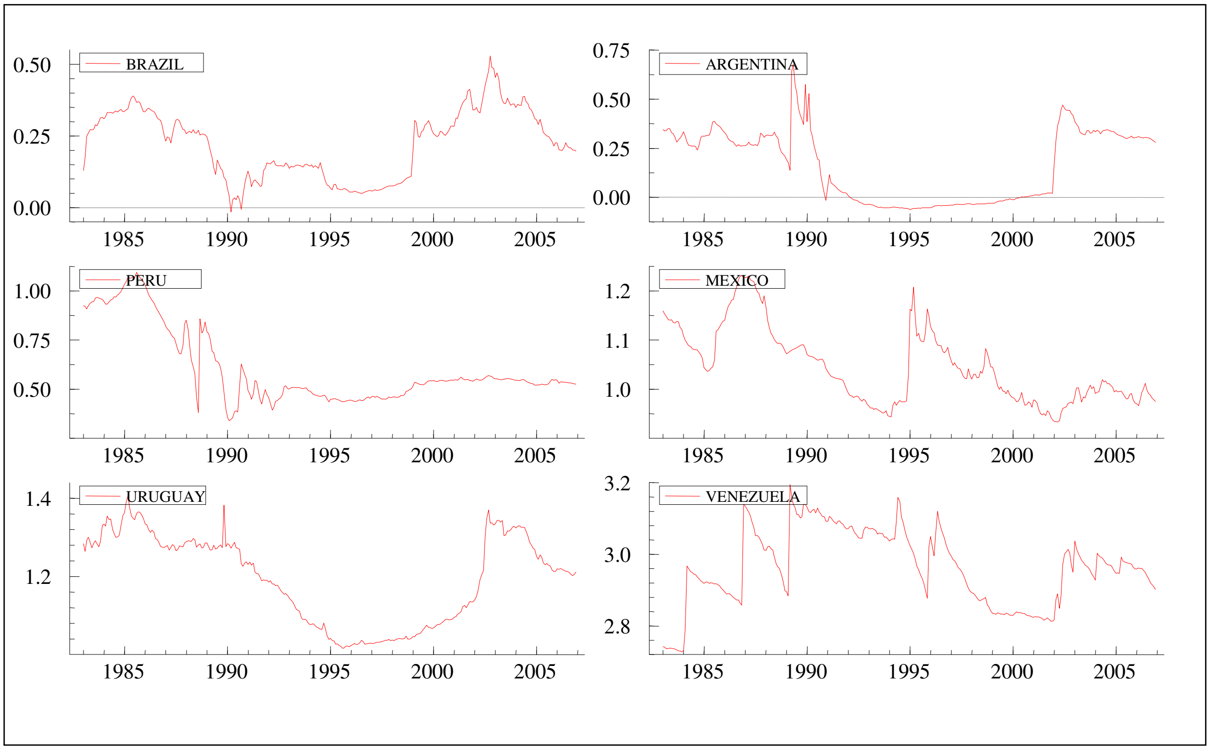


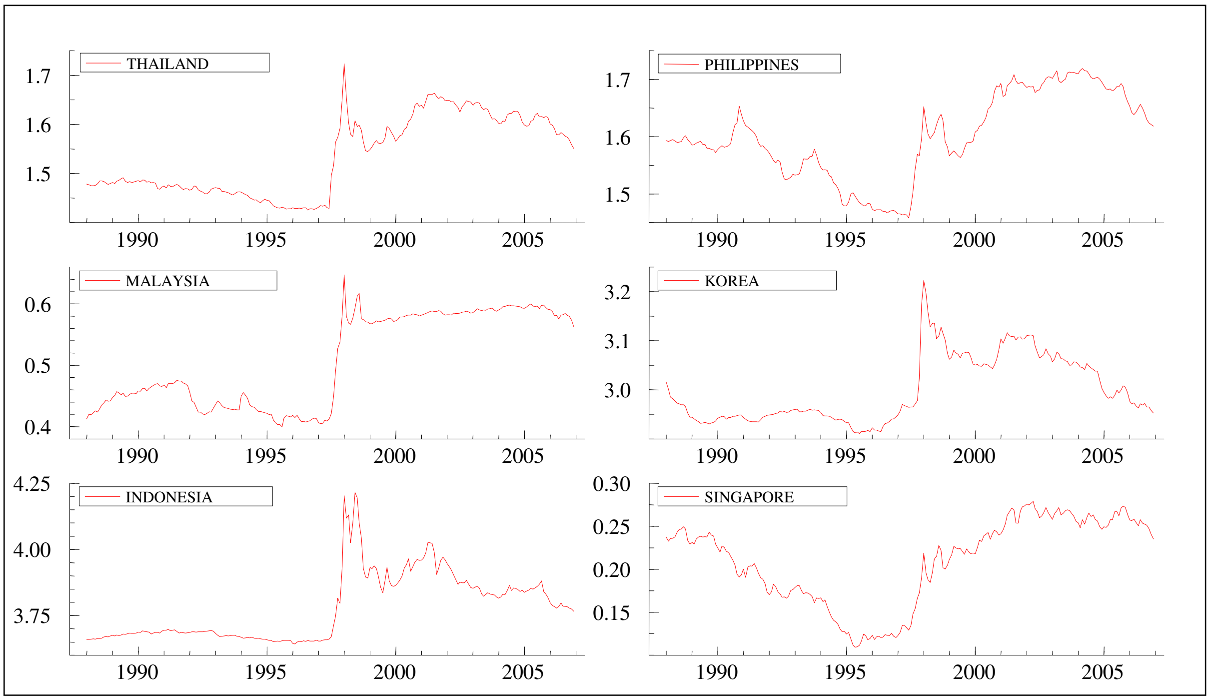
Figure 2 shows a degree of high volatility in the Latin American currencies prior to 1991 in particular. There are a variety of reasons for this. For example, in Brazil exchange rate policy was typically used to control inflation. This was not always successful, however, and when inflation became uncontrollable to the extent that the domestic currency was useless, the Brazilian authorities tended to replace the currency[12]. 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[13]. 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, crawling pegs and crawling bands. They have now

[12] 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.

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

all shifted to floating rate regimes[14]. Previous work done on whether a monetary union is appropriate for Latin America tends to suggest that it is not (e.g. Bayoumi and Eichengreen, 1994; Hallwood et al, 2004; Foresti, 2007; Neves et al, 2007). The common reasons cited for this include a low level of trade integration, asymmetric comovements to shocks, differences in speed of adjustment and size of shocks. The consensus appears to be that more policy co-ordination is necessary before any economic integration in Latin America can proceed.

Figure 3 ***(Log) Real Exchange Rates – Asian Crisis Countries, 1988-2006***



It is clear to see evidence of the de facto peg against the US dollar prior to the crisis events of July 1997, as well as the sharp real devaluation of all of the currencies in question at the crisis point[15]. 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 variants thereof) regimes after the crisis (excluding Malaysia who maintained a fully fixed rate against the US dollar). The countries

[14] 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.

[15] See Reinhart and Rogoff (2004) for more detail on the nature of the exchange rate regimes evident for these and other countries.

selected were considered to be those most affected by the crisis and 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 multivariate cointegration method is employed. In order to proceed with the Johansen technique, a necessary condition is that all of the variables in the system are integrated of the same order. Prior to running the cointegration tests, therefore, a further element of preliminary work must be undertaken, namely performing unit root tests on the real exchange rates of the crisis countries across the three regions. Of course, the nature of the approach for this study means that three sets of tests must be carried out for each region: total period, pre-crisis period, and post-crisis period. The standard Augmented Dickey-Fuller tests show that all of the real exchange rates are non-stationary in levels across the three periods and have been generated via an I(1) process[16]. 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[17]. The ultimate 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 is 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[18]. In selecting the most appropriate

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[16] These results are set out in the Appendix.

[17] Beirne (2007) has shown that unit root procedures employed to test for PPP may have been somewhat complacent in that an intercept term automatically enters the model equation, even in the absence of testing the significance of the constant. A more comprehensive procedure would be to employ the Dolado, Jenkinson and Sosvilla-Rivero (1990) approach and should insignificance be found in both deterministic elements then recursively de-meaned series should be utilized in the model without trend or intercept (as in Beirne, 2007; and Beirne, Hunter and Simpson, 2007).

[18] 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

(continued)

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, r[19]. These results are set out below.

Table 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 is thus 1 for all three regions. The trace and maximum eigenvalue test results are provided below for convenience.

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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.

[19] 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 the maximum eignevalue 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.

Table 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 trace statistics, the analysis indicates the presence of one cointegrating vector for all regions. Therefore, over 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 below.

Table 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 4.

Table 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.

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 between real exchange rates in Asia. There is a dramatic shift, however, in the post-crisis scenario, suggesting some form of co-ordinated policy action supportive of stability. 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 cointegration equations is necessary.

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 has become more pronounced following the crisis. Firstly, the long-run cointegration equation is examined. In this case, monetary integration would appear more appropriate where 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 monetary integration implications. 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 on exchange rate policy.

(i) Long-Run Elasticity

In order to interpret the cointegration results, a necessary first step is to normalize the cointegrating vector on one of the dollar real exchange rates[20]. For Europe the normalization is on USDATS (i.e. Austrian schilling per US \$); for the Latin American sample, the vector is normalized on USDARS (i.e. Argentine peso per US \$); and for the South East Asian sample, normalization is undertaken according to USDIDR (i.e. Indonesian rupiah per US \$). All of the equations are set out in Table 5. Prior to explaining the economic meaning of the cointegrating vectors based on the normalized 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. 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 USDATS increases USDNLG by 1%; and a 0.32% increase in USDATS is associated with a 1% increase in USDITL. The pre- and post-crisis results are all significant for the purposes of comparisons 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[21].

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[20] The results are not dependent upon the normalization base.

[21] 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).

The Latin American full sample results are all statistically significant in the cases of Peru and Venezuela only. The coefficients, however, are somewhat mixed across the 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 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 there is no consistency across the countries 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 monetary integration for the region[22].

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 normalized equations reported for this sample. For the post-crisis period, all of the coefficients are significant, and greater than unity. Korea, Malaysia and 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 countries

[22] 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.

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 parametrically. 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 there. Monetary union may be suitable within sub-groupings of the countries, but certainly not all due to the differences described in the beta coefficients.

(ii) Speed of Short-Run Adjustment

Table 6 provides 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 extent of weak exogeneity. Specifically, 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 cointegrating relationship remains external to the 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. If a variable is weakly exogenous, then it has no explanatory power with respect to 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. This may cast an element of scepticism regarding the suitability of the UK for a monetary integration arrangement with the current euro area members. 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 due to the observed weak exogeneity.

No weak exogeneity is observed for the Asian post-crisis period. 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.

Table 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 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. CONCLUSIONS

This study examines the G-PPP relationship 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 G-PPP relationships may differ following a major crisis. Following the approach of Choudry (2005), the period of the crisis is separated from the analysis given the inherent economic turbulence during that time.

The key conclusion from the analysis appears to be that following a crisis, the movement towards long-run stationarity in real exchange rates 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.

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.

Going forward, it would be of interest to analyse the exchange rate relationships in a less restrictive framework. The preceding empirical work was carried out on real exchange rates, which of course implicitly imposes the proportionality and symmetry conditions associated with PPP. A less restrictive framework would either impose just one of the conditions (e.g. a bi-variate test of the relationship between the nominal exchange rate and the ratio of domestic and foreign prices would impose just the symmetry condition, whereas the most general case – both proportionality and symmetry imposed - would be provided by a tri-variate model incorporating the nominal exchange rate, the domestic price series and the foreign price series). Both of these scenarios, however, are fraught with difficulty due to the large number of variables in the system.

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Appendix

Table A.1 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 A.2 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 A.3 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 A.4 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 A.5 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 A.6 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 A.7 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 A.8 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 A.9 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.

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
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