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Real Equilibrium Exchange Rates. A Panel Data Approach*

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Abstract

Based on an equilibrium exchange rate model, this paper examines the determinants of the real effective exchange rate and evaluates the degree of misalignment of the Group of Twenty (G-20) currencies over the floating exchange rate period. Within a panel cointegration setting, we estimate the relationship between exchange rate and a set of economic fundamentals, such as traded-nontraded productivity differentials and the stock of foreign assets. Having ascertained the variables are integrated and cointegrated, the long-run equilibrium value of the fundamentals are estimated and used to derive equilibrium exchange rate and misalignments.

Keywords: Equilibrium exchange rates, Group of Twenty, panel data, cointegration, misalignments.

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Introduction

Characterizing exchange rate dynamics has been a topic of intense research as well as one of the focus of attention in economic policy debates for a long time. Indeed, it is largely accepted that the evolution of the real exchange rate is a leading indicator of the strength of an economy. For instance, foreign investment, capital flows or international trade are all deeply influenced by the modifications that the real exchange rate can bring upon the goods and capital markets. Even more, given its close link with competitiveness, it turns out to be a fundamental issue when dealing with economic crises and booms.

In this sense, up until the eighties, the widespread intuition was that, given the mean-reversion mechanism, nominal and real exchange rates could not be misaligned over long periods. However, the observed movements of the exchange rates -long periods of high fluctuations, long-lasting misalignments with economic fundamentals, etc- during the 1980s and 1990s, not only suggested that the real exchange rate may not be mean reverting, but also raised doubts on the performance of the traditional theories of exchange rate determination and/or the econometric models used so far.

In addition, increasing external imbalances have been taking place in the world since the late 1990s. On the one hand, the U.S has been running chronically large deficits; on the other, Japan and some other Asian countries, Russia and the Middle East, have undergone large surpluses. The euro area countries, somehow in the middle, have also experienced important, although not very impressive, current account surpluses[1]. Given the close link of the external position and the exchange rate we should expect these facts to influence future developments in the equilibrium exchange rates.

Under this scenario, is then clear why the debate on exchange rates dynamics, in general, and exchange rate misalignments, in particular, has increasingly been including emerging market and developing countries. Nowadays, the inclusion of these countries is even more evident first, because despite the persistent demands made by the major countries (i.e. Group of Seven) to make exchange rates more flexible, some Asian countries -especially China- have continued *de facto* or *de jure* to run fixed pegs on the US dollar (see Bénassy-Quéré, Duran-Vigneron, Lahréche-Révil, and Mignon (2004)). Second, to the fact that Asian central banks are starting to diversify their huge currency reserves, a move that will probably produce realignments on major exchange rates[2].

It is, precisely, within the Group of Twenty (G-20), which includes, apart from the major industrial nations, emerging market economies, where the debate on exchange rates misalignments should take place[3]. Indeed, as it has been noticed, as far as persistent

[1]For instance, according to IMF data from 2005, while in the major advance economies the current account balance as percentage of GDP was -2.1%, in emerging markets and developing countries it was around 4.1%. In particular, the individual imbalances grew to -6.1% percent of GDP in the U.S,-5.7% in Australia, +4.3% in Taiwan, +10.3% in Hong Kong, +2.0% in South Korea, +13.2% in Russia, and +32.4% in Saudi Arabia. In the euro area growth of the current account was a moderated 0.5% the same year

[2]In 2004, Asia had more than $2.46 trillion in official foreign reserves accumulated, or around 65 %percent of the world's total reserves. These reserves, which are mainly invested in US$ denominated assets, constitute an important source of finance for the US current account deficit.

[3]The G-20 was created as a response both to the financial crises of the late 1990s and a growing recognition that key emerging-market countries were not adequately included in the core of global economic discussion and governance. It is basically, an informal forum that promotes discussion between industrial and emerging-market countries on key issues related to global economic stability. The countries included in the G-20 are: Argentina, Australia, Brazil, Canada, China, France, Germany, India, Indonesia, Italy, Japan, Mexico, Rep. of Korea, Russia, Saudi Arabia, South Africa, Turkey, U.K and the U.S. Another member is the European Union represented by the European Central Bank.

exchange-rate misalignments could be the source of misallocation of resources, this should be an issue discussed in the G-20 meetings (Bénassy-Queré, et. al., 2004).

Given these facts, one is left with some questions concerning the real exchange rate evolution. First, is the real exchange mean reverting, or does it have a unit root?. Second, if it is not mean reverting, what are the economic fundamentals that determine its evolution in the long-run?. Finally, how significant are the misalignments of the effective exchange rates of the countries?

In order to address these questions, we employed data from medium income-level and emerging market economies together with high-income economies, arguing that these countries share common characteristics that justify the use of a panel data approach. From an empirical point of view, we are particularly interested on these techniques to estimate equilibrium exchange rates because, by exploiting the cross-section as well as time series variability, we increase the power over univariate methods[4].

This paper is organized as follows. In the next section we give a brief review of the equilibrium exchange rates theories. Section 2 describes the data. The third section discusses the panel data techniques implemented and the empirical findings. Based on these results, in the fourth section we investigate the size of the effective misalignments. Section five concludes.

1 Modelling Real Equilibrium Exchange Rates

The purchasing power parity (PPP) is a simple and popular concept of long-run or equilibrium exchange rate in the literature, implying a constant equilibrium exchange rate. However, the extensive literature on PPP has reached some consensus regarding some concrete facts. First, most of the empirical studies generally cannot reject the random walk hypothesis for the real exchange rates of the currencies of all the major industrialized countries against one another, suggesting either that deviations from PPP are permanent or that the real exchange rate tends to the purchasing power parity in the very long run. The estimations show, however, that this convergence is extremely slow. Second, the short run deviations from the PPP are high and volatile, even for homogenous goods. Finally, the evidence in favor of PPP is stronger for developing countries.

Hence, the lack of empirical evidence supporting the PPP gave rise to several alternative approaches. In particular, the macroeconomic approach focused in the determinants of the real exchange rate based on the economic fundamentals.

In this sense, the notion of "equilibrium exchange rate" meaning an exchange rate consistent with long-run macroeconomic fundamentals, was introduced as a way to obtain a benchmark to check the possible misalignments of the actual exchange rate.

Thus, besides the PPP, several other ways of calculating equilibrium exchange rates have been suggested. However, two main lines of research are distinguished. The first one, known as the balance of payments approach, highlights the underlying net foreign asset position of a country. The second one, based on the Harrod-Balassa-Samuelson approach, centers its attention in the sectoral (tradable-nontradable) balance.

The balance of payment approach, introduced by Nurkse (1945), is based on the adequacy of the current account to sustain notional or equilibrium capital flows and keep in check saving-investment balances. In this framework, several notions of equilibrium exchange rates have been suggested. For instance, Williamson (1985) suggested the notion

[4] However, we are aware that inference can potentially be affected by the mix of series in a panel.

of Fundamental Equilibrium Exchange Rate (FEER), which is defined as the exchange rate "which is expected to generate a current account surplus or deficit equal to the underlying capital flow over the cycle, given that the country is pursuing internal balance as best it can and not restricting trade for balance of payments reasons" (Williamson, p.14). A variant of the FEER is the Desirable Equilibrium Exchange Rate (DEER), proposed by Bayoumi, Clark, Symansky, and Taylor (1994) which assumes target values for the macroeconomic objectives, such as targeted current account surplus for each country.

Another approach to equilibrium exchange rate is the Behavioral Equilibrium Exchange Rate (BEER) by Clark and MacDonald (1998). The BEER relies on observed long-run relationships between the real exchange rate and a set of (long-run and medium term) economic fundamentals -derived from the determinants of saving, investment and the current account- and a set of transitory factors affecting the real exchange rate in the short run.

The Macroeconomic Balance Framework for estimating equilibrium exchange rates, suggested by Faruquee, Isard, and Masson (1998) can be seen as a specification of the BEER approach since the macroeconomic fundamentals are derived from the determinants of saving, investment and current account. Similarly, the Natural Real Exchange Rate (NATREX), introduced by Stein (1994), falls in between the FEER and the BEER approaches. The NATREX is the exchange rate that permits the attainment of both internal and external equilibrium. However, the current account is modeled as the result of saving and investment behavior, as in the BEER approach

The second line of research, based on the works by Balassa (1964) and Samuelson (1964) relates the long run behavior of the real exchange rate with the productivity performance of traded relative to nontraded goods. Basically, the argument is that as the productivity of the traded goods rises relative to that of nontraded goods, there will be a tendency for the real exchange rate to appreciate.

Following these propositions, Alberola, Cervero, López, and Ubide (1999) and Alberola (2003) propose a theoretical model that encompasses both, the balance of payments and the Balassa-Samuelson approach to real exchange rate determination[5]. The starting point of the model is the decomposition of the exchange rate into two different relative prices. The first one, the price of domestic relative to foreign tradables, captures the competitiveness of the economy and determines the evolution of the foreign asset position. On the contrary, the second one, which is the relative price of non-tradables relative to tradables within each country, plays a central role in adjusting excess demand across sector in the economy.

Thus, based on Alberola, Cervero, López, and Ubide (1999) and Alberola (2003), we identify two main fundamentals for the evolution of the real exchange rate: the level of net foreign assets and a measure of relative sectoral productivity.

2 Data sources and construction of the variables

We consider data from fourteen currencies corresponding to most of the G-20 country members: Argentina, Australia, Brazil, Canada, Indonesia, India, Japan, South Korea, Mexico, Turkey, South Africa, the United States and the United Kingdom. France, Germany and Italy are grouped in the Euro-zone, together with the rest of the twelve initial

[5]We do not discuss this model in detail here but the reader may referred to Alberola, Cervero, López, and Ubide (1999) and Alberola (2003).

members countries (EU12)[6]. The data are quarterly and cover the period 1980 Q1-2004 Q4[7].

The main sources we have used in this papers are the IMF's *International Financial Statistics*, the OECD's *Analytical Database* and the European Commission's *AMECO Database*. Based on these information, we constructed the following variables:

- **Real effective exchange rate:**

  The real effective exchange rate (REER), based on consumer prices, measures movements in the nominal exchange rate adjusted for the differential between the domestic price index and trade-weighted foreign price index. The CPI-based REER indicator, of a country i is:

  $$REER_{it} = \frac{P_{it}S_{it}}{\prod_{j \neq i}^{N}(P_{jt}S_{jt})^{\omega_{ij}}} \tag{2.1}$$

  where j is an index of country i's trade partners; N is the number of countries, ω_{ij} is the competitiveness weight put by country i on country j, P_i and P_j are consumer price indices in countries i and j; and S_i and S_j represent the nominal exchange rates of countries i and j's currencies in US dollars[8]. An increase (decrease) in a country's index indicates an appreciation (depreciation).

  In (2.1) i's N partners are exclusively the G-20 countries which had a competitiveness weight, ω, in i's trade equal or bigger than 2 percent, normalizing these weights to sum to one. The same criteria was used to compute the rest of the variables in multilateral terms[9].

  We work with the logarithm of the real effective exchange rate, denoting this variable as q_t.

- **Productivity differentials**. The impact of productivity differentials is expected to follow the Balassa-Samuelson doctrine, which states that relatively larger increases in productivity in the traded goods sector are associated with a real appreciation of the currency of a country. Given the fact that more of the components of the Producer Price Index (PPI) are tradable than those in the Consumer Price Index (CPI), to capture the diverging productivity trends we use the ratio of CPI to the PPI. We obtained the ratio of this index for each country to its equivalent weighted foreign average and work with its logarithm (pd):

[6] The remaining members of the Euro-zone are: Austria, Belgium, Finland, Greece, Ireland, Luxembourg, Netherlands, Portugal and Spain. OECD provides data for the euro aggregate, except for the terms of trade. The weights used in aggregating the individual country series are based on 1995 purchasing power parity's GDP of the initial members.

[7] Since data is not complete or reliable, we have not included China, Russia and Saudi Arabia in the sample.

[8] For the weights we used data from Durand, Madaschi, and Terribile (1998). This calculation of competitiveness uses a system of weights based on a double-weighting principle, which takes into account the structure of competition in both export and import markets. The weights also consider the influence of emerging market economies, which is an advantage over other sources. Thus, countries like Korea, Indonesia, India, Argentina, Brazil and Mexico, are considered in the weighting matrices, contrary to those presented in Zanello and Desruelle (1997). The weights correspond to the 1995 matrix.

[9] If for some years, or for the whole period, a variable was missing for one or several of the N countries, we distributed the weights between the rest of the partners.

$$PD_{it} = \frac{CPI_{it}/PPI_{it}}{\prod_{j\neq i}^{N}(CPI_{jt}/PPI_{jt})^{\omega_{i,j}}} \tag{2.2}$$

- **Stock of foreign assets**. There are several channels through which the stock of foreign assets can influence the real exchange rate[10]. For instance, portfolio-balances considerations suggest that a deficit in the current account creates an increase in the net foreign debt of a country, which has to be financed by internationally investors which, in order to adjust their portfolio, demand a higher yield. At given interest rates, this can only be achieve through a depreciation of the currency of the debtor country

  Also, the balance of payments channel assumes that a current deficit accumulates net foreign debts, for which interest have to be paid. To service these higher interest payments, the debtor country needs to strength its international price competitiveness. Thus, to increase the attractiveness of its exports, the country needs to depreciate its currency.

  The stock of net foreign assets ACA_t was computed by adding the accumulating past current account balances (CA_{t-j}) to the current one (CA_t):

$$ACA_{it} = \sum_{j=0}^{t} CA_{i,t-j} + CA_{i,t} \tag{2.3}$$

  where we accumulated the current account for each country since 1970, except for the euro-zone where we started in 1975. We use the ratio of net foreign assets over GDP and denoted this variable as NFA.

3 Panel unit root and cointegration tests

There are several reasons that justify the use of panel data unit root and cointegration techniques. First of all, it is well known that time series unit root and cointegration tests have limited power in finite samples against alternative hypothesis with highly persistent deviations from equilibrium. Thus, the use of panel data is usually seen as a mean of generating more powerful tests.

Second, most of the time, countries share common similarities that are usually lost in time series analysis. For instance, non-stationarity appears more as a general feature for macroeconomic series, characterizing an important number of developed and developing countries. Hence, it is interesting to test the homogeneity of the unit root in a panel framework.

Even more, not only panel unit root tests allow to take into account the international dimension and the similarities of a panel of countries, but also some specifications take into consideration the possible cross-sectional dependence among them. Also, a major advantage of panel cointegration techniques is that they allow one to selectively pool the

[10]For example, Lane and Milessi-Ferreti (1999) explore the theoretical link between the real exchange rate and the net foreign assets, and provide evidence that the net foreign asset position is an important determinant of the real exchange rate for developing as well as developed countries.

long run information contained in the panel while permitting the short-run dynamics and heterogeneity among different members.

However, as Breitung and Pesaran (2005) noticed, testing the unit root and cointegration hypothesis by using panel data instead of individual time series involves at least five complications. First, panel data generally introduce a substantial amount of unobserved heterogeneity, rendering the parameters of the model cross section specific. Second, a difficulty that second generation panel unit root tests have doped with, in many empirical applications, is that it is inappropriate to assume that the cross-section units are independent. Third, the panel test outcomes are often difficult to interpreted if the null of the unit root or no cointegration is rejected as the test do not provide explicit guidance as to identify the cross-section units that are stationary or cointegrated. Fourth, with unobserved integrated variables common factors affecting some or all the variables in the panel, it is also necessary to consider the possibility of cointegration between the variables across the groups (cross-section cointegration) as well as within group integration. Finally, the asymptotic theory is considerably more complicated.

Thus, the purpose of this section is to employ panel unit root tests and cointegration methods to, first, evaluate the PPP doctrine and second, to assess the long-run relationship between the real effective exchange rate and the underlying macroeconomic fundamentals in the context of a set of somewhat homogenous countries.

3.1 Methodology

For this purpose we examine our data by performing several unit root tests in a panel framework: Levin and Lin (1992); Im, Pesaran, and Shin (1997); Hadri (2000); Maddala and Wu (1999); Choi (2002); Bai and Ng (2004); Pesaran (2003) and Chang (2002). In addition, to properly investigate the stationarity of the data, we employ two robust tests, suggested by Im, Lee, and Tieslau (2005) and by Carrion-i Silvestre, Del Barrio-Castro, and López-Bazo (2005), that allow for heterogenous structural breaks in a panel framework.

However, it must be added that the rejection (acceptance) of a unit root (stationarity) in the real exchange rate when the test incorporates structural breaks does not necessarily give empirical support for the long-run PPP. Indeed, under PPP, the rate of currency depreciation is approximately equal to the difference between the domestic and foreign inflation rates, a fact that is reflected in a stable real exchange rate in the long-run. Suppose we find evidence of stationarity in the real exchange rate over a certain period. Imagine also that there is a devaluation that is not reflected in equal proportion to the inflation rate. In this case, after the break, the real exchange rate moves up to another level. If we test again for unit root in the period after the break, we can find evidence of stationarity. Still during the whole period the real exchange rate has jump from one level to another without returning to its original value[11].

Since the real exchange rate and the fundamentals are commonly found to be non-stationary variables, they must be modelled in a suitable econometric framework in order to avoid drawing conclusions based on spurious results. Therefore, cointegration panel data analysis is also proposed by performing Pedroni's cointegration tests (Pedroni (1999)). Finally, we estimate the long-run relationship using two alternative methodologies: the Dynamic OLS (DOLS) estimator suggested by Kao and Chiang (2000) and the

[11]The simple fact that a structural break causes that the real exchange rate shifts without going back to its original level can be taken as evidence against the PPP.

Pooled Mean Group (PMG) estimator as proposed by Pesaran, Smith, and Shin (1999).

3.1.1 Panel unit root tests

In the last years, there has been a lot of research on non-stationary panels, in particular in the context of testing for the presence of a unit root. In this literature, test can be classified into three groups depending on whether they allow for structural breaks and cross-sectional correlation, which are features very often found in time series and panel data models.

The first one is based on a cross-sectional independence and no structural break assumptions[12]. Examples of tests in this group are the Levin and Lin (1992), Im, Pesaran, and Shin (1997), Hadri (2000) and Maddala and Wu (1999).

The second category of tests is characterized by not imposing the cross-sectional independence assumption but do not admit breaks. Here, two main approaches are distinguished. On the one side, the contributions of Choi (2002); Bai and Ng (2004) and Pesaran (2003), among others, rely on the factor structure approach. On the other side, Chang (2002) suggested imposing few or none restrictions on the residual covariance matrix and the use of nonlinear intrumental variables method to solve the nuisance parameter problem due to the cross-sectional dependency[13].

Still, it is well known that structural breaks can be mistaken for non-stationary (see Perron (1989)). Hence, suitable panel unit root tests allowing for structural breaks have been suggested recently in the literature. The first of this tests, the panel Lagrange Multiplier (LM) unit root test, $L\bar{M}_{NT}$, suggested by Im, Lee, and Tieslau (2005), permits for up to two heterogenous level shifts. The second one, the $LM(\lambda)$ by Carrion-i Silvestre, Del Barrio-Castro, and López-Bazo (2005) allows for an unknown number of breaks. Both tests allow for common time effects.

3.1.2 Panel cointegration

The literature on panel data cointegration has followed two broad practices. The first one, which is the base of Pedroni's (Pedroni (1999)) and Kao's (Kao (1999)) tests, consists on taking as a null hypothesis that of no cointegration and using the residual derived from the panel analogue of an Engle and Granger statistic regression to construct the test statistics and tabulate the distribution. The second practice, also a residual-based test, takes as a null that of cointegration and is the basis of the test proposed by McCoskey and Kao (1998)[14].

This paper focuses on the method proposed by Pedroni (1999), which has the advantage over others that it allows for considerable heterogeneous variances across countries at each point in time. Thus, it allows one to selectively pool the long-run information contained in the panel, while permitting the short-run dynamics and heterogeneity among different members. This "pooling" can be done across either the "within" or the "between" dimension.

In Pedroni's panel cointegration tests, the null hypothesis is taken to be that for each member of the panel the variables of interest are not cointegrated. The alternative

[12]Actually, these tests only allow for a very limited amount of cross-sectional dependence through the inclusion of time effects

[13]For a complete description on these panel unit root tests see Hurlin (2004), Hurlin and Mignon (2004), Breitung and Pesaran (2005) and Baltagi and Kao (2000).

[14]For a complete discussion on panel cointegration tests see Baltagi (2002) and Banerjee (1999).

hypothesis is taken to be that for each member of the panel there exists a single cointegrating vector, although this cointegrating vector needs not be the same for each member. Thus, the tests not only allow for the dynamics and fixed effects to differ across members of the panel, but also, in the case of the between-dimension statistics, they allow the cointegrating vector to differ across members under the alternative hypothesis.

3.1.3 The long-run relationship

Having established that a cointegration relation indeed exists, the long-run parameters can be estimated efficiently using techniques similar to the ones proposed for individual time series models. Most approaches employ a homogeneous framework, that is, the cointegration vectors are assumed to be identical for all panel units, whereas the short-run parameters are panel specific.

However, such an assumption seems overly restricted for some economic relationships. On the other hand, allowing all the parameters to be individual specific would substantially reduce the attraction of a panel data study: it is therefore important to identify parameters that are likely to be similar across panel units whilst at the same time allowing for sufficient heterogeneity of other parameters.

In this paper, we consider two main approaches to estimating a long-run relationship among integrated variables in a panel framework: (1) the dynamic OLS (DOLS) estimators associated with Pedroni (2000), Mark and Sul (2001) and Kao and Chiang (2000) and (2) the error correction pooled mean-group (PMG) estimators proposed by Pesaran, Smith, and Shin (1999).

The starting point of the DOLS set-up is a long-run relationship as the following equation:

$$y_{it} = \alpha_i + \delta_i t + \theta_t + \beta_1 x_{1it} + ... + \beta_k x_{k,i,t} + e_{it} \tag{3.1}$$

where k is the number of regressors, β_k are elasticities, the index i indicates the section of the panel and the time index t the length of the sample. As it can be seen, fixed effects, individual specific deterministic trends and different error variances are all permitted.

In order to obtain an unbiased estimator of the long-run parameters, this procedure involves a parametric adjustment to the errors of the static relation. The correction is achieved by assuming that there is a relationship between the residuals from the static regression, ε_{it} and the first differences of the leads, lags and contemporaneous values of the regressors in first differences:

$$\varepsilon_{it} = \sum_{j=-q}^{q} c_{ij} \Delta x_{i,t-j} + \varepsilon_{it}^{\star} \tag{3.2}$$

where ε_{it} is orthogonal to all leads and lags of Δx_{it}. A simple OLS regression provides a super-consistent estimate of the long-run parameters. The t-statistic is based on the long-run variance of the residuals instead of the contemporaneous variance, which is commonly used in OLS regressions.

On the other side, Pesaran, Smith, and Shin (1999), PSS hereafter, noticed that there are two procedures commonly used for dynamic panel data models. The first one, known as the Mean Group (MG) estimate, consists of estimating separate regressions for each group (country) and calculating averages of the group specific coefficients. They show that the MG estimator will produce consistent estimates of the average of the parameters. This estimator, however, does not take account of the fact that certain parameters may be the same across groups. On the other extreme are the traditional pooled estimators, such as the fixed or random effects estimators, where the intercepts are allowed to differ freely across groups while all other coefficients and error variances are constraint to be the same. PSS suggest an intermediate estimator, called the Pooled Mean Group (PMG) estimator because it involves both pooling and averaging.

The starting point of the procedures is an autoregressive distributed lag (ARDL) model of order $(p, q, q, ..., q)$ for $t = 1, 2...T$ and $i = 1, 2, ...N$ that reparameterised as a vector error correction mechanism (VECM) system can be expressed as:

$$\Delta y_{it} = \theta_i y_{i,t-1} - \beta_i' x_{i,t-1} + \sum_{j=1}^{p-1} \lambda_{ij} \Delta y_{i,t-j} + \sum_{j=1}^{q-1} \delta_{ij} \Delta x_{i,t-j} + \mu_i + \varepsilon_{it} \tag{3.3}$$

where $x_{i,t-j}$ is the $(k \times 1)$ vector of explanatory variables for group (country) i, β_i are the long-run parameters, the θ_i are the equilibrium (or error) correction parameters, λ_{ij} and δ_{ij} include the country-specific coefficients of the short-term dynamics, μ_i represents the fixed effects and ε_{it} is a white noise process. In principle, the panel can be unbalanced and p and q may vary across countries. The pooled mean group restriction is that the parameters of β_i are common across countries; that is $\beta_i = \beta$ in equation (3.3).

Hence, if θ_i is significantly negative, there exists a long-run relationship between y_{it} and x_{it}. Estimation of (3.3) could proceed by OLS, imposing and testing the cross country restrictions on β. However, this will be inefficient as it ignores the contemporaneous residual covariance. Therefore, PSS suggest a maximum likelihood estimator.

Instead of assuming common long-run parameters, the estimation can also be done with individual specific β_i, which are then averaged over N to obtain the MG estimator. As it can be expected, this is a natural background to test for the presence of slope homogeneity based on a Hausman test. This test is based on the results that an estimate of the long-run parameters in the model can be derived from an average (MG estimation) of the country regressions and this is consistent even under heterogeneity. However, if the parameters are in fact homogenous, the PMG estimates are more efficient. That is, if the poolability assumption is invalid then the PMG estimates are no longer consistent.

Finally, an important assumption of the panel cointegration tests is the absence of cross-sectional correlation. The possible presence of contemporaneous correlation can be addressed either by using time dummies or by subtracting the cross-sectional mean from the data. If the slopes are homogeneous, the MG and the PMG methodologies are equivalent.

3.2 Estimation results

The first step in the panel cointegration approach is to ascertain that the series are non-stationary in levels[15]. The unit root tests were implemented including fixed-effects in the specifications because, *a priori*, there is no reason to believe that all the intercepts are the same across countries. Also, since the test results are very sensitive to the specification of the individual dynamic structures, we allowed the models to have heterogeneous dynamic

[15]Except for the real exchange rate, we did not succeed to get complete series for some countries, then panels are not balanced. Since the MP and BNG tests require balanced panels, we reduced the samples in order to consider the same balanced database. Thus, for the ratio CPI to PPI we used data from 1994Q1 upwards, for terms of trade, from 1993Q3 to 2003Q3 and we eliminated India from the data corresponding to the net foreign assets.

Table 3.1: **First and second generation panel unit root tests on the variables in levels**

	Cross-sectional independence							Cross-sectional dependence			
	LL	**IPS**		**Hadri**		**MW**	**Choi**	**BNg**		**P**	**Chang**
	t^*	t_{bar}	W_{bar}	Z	HCZ	P_{MW}	Z_{MW}	Z_{ε^c}	$Z_{\varepsilon^{mw}}$	$CIPS$	S_N
q	-0.937	1.641	1.066	13.14**	10.20**	36.887	1.188	-0.901	21.257	-2.059	-0.083
	(0.174)		(0.143)	(0.000)	(0.000)	(0.121)	(0.118)	(0.816)	(0.815)	(0.165)	(0.567)
pd	-2.04**	-1.581	-0.306	18.779**	15.864**	34.783	0.907	0.720	33.386	-1.452	0.529
	(0.020)		(0.380)	(0.000)	(0.000)	(0.176)	(0.182)	(0.236)	(0.222)	(0.850)	(0.702)
NFA	0.679	-0.897	2.670	16.46**	16.22**	12.678	-2.048	-1.414	15.807	-1.784	2.225
	(0.752)		(0.996)	(0.000)	(0.000)	(0.994)	(0.980)	(0.921)	(0.941)	(0.575)	(0.987)

Note: 1) *,**,*** indicates the rejection of the unit root null hypothesis at the 1%,5% and 10% significance level, respectively; (2) lags selected according to the the Ljung-Box and LM statistics to ensure no serial correlation in the residuals, with a maximum lag length of 10; (3) All the specifications include fixed effects; (4) p-values are given in parentheses; (5) For the Chang IV test, the IGF used is: $F(x) = x\exp(-c_i|x|)$ where $c_i \in \Re$ is determined by $c_i = 3T^{1/2}s^{-1}(\Delta y_{it})$ where $s^2(\Delta y_{it})$ is the sample standard error of Δy_{it}.

structure, i.e., the models have different autoregressive orders for individual cross-sectional units, ensuring uncorrelated residuals[16]. Table (3.2) summarizes the first generation (LL, IPS, and HA) and second generation (MW, Choi, PE and Bai-Ng) panel unit root tests.

As it can be seen, the results from the LL test oppose with the rest in the case of the ratio CPI to PPI, since the presence of a unit root is rejected. Yet, it is important to remember that the restrictive homogenous assumption used in this test implies that, under the null hypothesis, all the members in the panel contain a unit root, while the alternative hypothesis is that all the series are stationary. Then with as few as one stationary series, the rejection rate rises above the nominal size of the test, and continues to increase with the number of stationary series in the panel. Not only that, the assumption of cross-sectional independence has the consequence that the LL test suffers from significant size distortion in the presence of correlation among contemporaneous cross-sectional error terms.

In the rest of the cases, even though for some tests we found contradictory results, globally most of them failed to reject the null hypothesis of unit root for the variables in levels. These results are corroborated by the Hadri test, in which case we rejected the null hypothesis of stationarity for the (log) real exchange rate, the (log) CPI to PPI ratio and the net foreign assets to GDP.

In order to obtain a confirmatory analysis, we performed panel unit root tests allowing for the presence of structural breaks on the variables in levels[17]. For this, we implemented two complementary tests. The first one, the panel $\bar{LM}_{NT}$ by Im, Lee, and Tieslau (2005), has as a null hypothesis the non-stationarity of the variable. The second one, the $LM(\lambda)$ by Carrion-i Silvestre, Del Barrio-Castro, and López-Bazo (2005), on the contrary, has stationarity as the null hypothesis.

Both tests assume no cross-correlation in the error terms across countries. However, controlling for time-specific effects common to all countries could, in principle, mitigate any cross-sectional dependencies in the data. Thus, to allow for time-specific fixed effects,

[16] In order to evaluate the sensitivity of the results to the lag length, we also present the results of the t-bar and W-bar statistics based on the Dickey-Fuller statistics, denoted $t - bar_{DF}$ and $W - bar_{DF}$. The results that we present turned up to be quite robust to the different lag selection criteria.

[17] We do not present the results for the PPI to CPI ratio and the NFA but are available upon request from the authors

Table 3.2: **Panel unit root tests on the REER allowing for structural breaks**

Test	Test Statistic	Critical Values 5%	10%
LM_{NT}	-3.302	3.842	3.504
$LM(\lambda)$			
homogenous	5.250	3.488	2.811
heterogenous	4.641	4.956	4.256

Notes: (1) Both tests allow for time fixed effects and all regressions include intercept and time trend; (2) Panel $L\bar{M}_{NT}$ test statistic allows for up to two break points and the grid search is performed over the interval $[0.1T, 0.9T]$; (3) In the panel $LM(\lambda)$ the number of break points has been estimated using the SIC, allowing for a maximum of 5 structural breaks; (5) The long-run variance in the panel $LM(\lambda)$ was estimated using the Bartlett spectral kernel with automatic spectral window bandwidth selection; (6) The boostrap distribution is based on 2,000 replications.

we demeaned the data by subtracting the average value of the series from each country in each country[18]. This is especially important in our context, since it is likely that real exchange rates in each country is contemporaneously correlated.

To compute the $L\bar{M}_{NT}$, we allow for a maximum of two structural shifts for each time series, determining the number of augmentations at each 0, 1 and 2 shifts. From the outcome of these results, we determine the optimal number of break location(s) and the optimal number of augmentations jointly and endogenous from the data. The augmentation terms were selected using the "general to specific" procedure. Thus, beginning with a maximum number of eight lags, the last augmentation term is examined to see if it is significant at the 5% level in an asymptotic normal distribution. Then, the shift locations are determined through a grid search over the time interval $[0.1T, 0.9T]$ to eliminate end points. The grid search is repeated at each possible combination of two break points to locate the minimum test statistic. After endogenously determining the location of the break(s), we determined the optimal number of of structural changes for each country by examining the significance of dummy coefficients based on t-tests. like this, the location of breaks, the number of breaks, and the number of lagged augmentation terms are jointly determined for each country.

In the case of the $LM(\lambda)$ the maximum number of time breaks allow are five. We first estimate the dates for all possible breaks in each country and then, following Bai and Perron (1998), we used the Schwarz Information Criteria to select the suitable number of structural breaks[19]. The order of the autoregressive process that was used in the prewithening stage was fixed using the BIC with a maximum of 8 lags.

Table (3.2) displays the unit root tests statistics of the panel LM_{NT} and panel $LM(\lambda)$. From these results we can outline the following facts. First, structural breaks were found to exist in all the countries. Even though, the location of them varies along countries, several breaks were found in the first half of the 1980's and then in 2001[20]. Second, we again did not find evidence to reject (accept) the null hypothesis of a unit root (stationarity) in

[18]Notice that the inclusion of time-specific effects in the panel setting is equivalent to allowing for structural break in each year that is common to all the countries considered. After including common time effects, the remaining structural breaks are specific (heterogenous) to individual countries

[19]In case no structural break is detected to affect a particular time series, we will proceed to compute the test statistic without structural breaks.

[20]The location of the breaks obtained from both methodologies are presented in the appendix, table (C.1).

the level of the (log) real effective exchange rate at the 5% level of significance. The only exception is when using the heterogenous estimate of the long-run variance, although the test is significant at the 10% level.

Notice that our findings are robust regardless of which assumption is made, first, about the cross-sectional dependence and, second, about the existence of structural breaks. However, we want to insist that this conclusion must be moderated since the results strongly depend on the specification of the model and the method used to eliminate the cross-sectional correlations. Also, it is important to keep in mind that by adding the cross-section to the time series dimension to the analysis, one has to be very careful, as the time series characteristics have to be taken properly into account.

Hence, in view of the non-stationarity characteristics of the data and, in order to avoid spurious regression problems, we now turn to a panel cointegration framework[21]. Therefore, first the cointegration tests proposed by Pedroni (1999) were used to verify whether there exists a long-run relationship between the real exchange rate, the productivity differential measure and the net foreign assets. This results are summarized in table (3.3).

Table 3.3: **Pedroni's panel cointegration tests**

	Test statistic	p-value
Panel cointegration		
ν-Stat	2.1566	0.0155**
ρ-Stat	-2.7458	0.0030**
t-Stat(non parametric)	-2.7879	0.0027**
t-Stat(parametric)	-1.1927	0.1165
Group mean cointegration		
ρ-Stat	-3.2434	0.0006**
t-Stat(non parametric)	-3.7273	0.0001**
t-Stat(parametric)	-1.6248	0.0521***

Note: (1) *,**,*** indicates the rejection of the null hypothesis of no cointegration at the 1%,5% and 10% significance level, respectively.

With the exception of the parametric panel t-statistic, the results point to cointegration between the variables, since the null hypothesis of no cointegration is easily rejected at the 5% significance level (10% in the case of the parametric group t-statistic). With this preliminary findings in mind, we now turn to the estimation of the long-run relationship between the variables. For this purpose, the coefficients of the long-term relationship were derived using: (1) Dynamic OLS (DOLS) estimates and (2) Pooled Mean Group (PMG) estimates. Results from these tests are summarized in table (3.4).

In the case of the PMG we have the particular advantage that they provide estimates not only of the long-run parameters, but also of the short-run dynamics and the speed of adjustment to equilibrium[22].

For the PMG estimates, diagnostic tests were applied to the ARDL(4,4,4) with the Schwarz Bayesian Criterium (SBC) used to select lags and imposing long-run homogeneity

[21] Entorf (1997) studied spurious fixed effects regressions when the true model involves independent random walks, finding that for $T \to \infty$ and N finite, the nonsense regression phenomenon holds for spurious fixed effects models and inference based on t-values can be highly misleading.

[22] We do not present the full specifications here, but are available from the authors upon request

to both variables[23]. According to the Hausman test, imposing this restriction (equality of slopes), cannot be rejected[24].

We can get some conclusions from the results of the PMG and MG estimations. First, we have found a negative and statistically significant error correction term, fact that is taken as evidence for the presence of cointegration. The (average) error correction coefficients reported show that adjustment is substantial, about 9% ocurring within a quarter of a year.

Second, in general, it seems to be some homogeneity across countries. Indeed, according to the Hausman tests, the parameters are homogeneous, implying that pooling leads to more efficient estimates than simple averaging the coefficients. Yet, we should consider that, even though the Hausman tests support pooling across all countries, it is highly plausible that developing and emergent economies have some structural differences with developed countries, even if the test is insufficiently powerful to pick this up[25].

Under both methodologies the variables are correctly signed: a rise in the NFA position (an improvement in the current account) or in the PPI to CPI ratio (a rise in the relative productivity) leads to a real exchange rate appreciation. Notice, however, that the net foreign assets parameter is not significant under the DOLS estimation[26].

Table 3.4: **Dynamic OLS Estimates and Pooled Mean Group Estimates**

	DOLS				**PMG**			
	Ordinary		**Time effect**		**Ordinary**		**Time effect**	
	Coef.	*t*-value	Coef.	*t*-value	Coef.	*t*-value	Coef.	*t*-value
pr	1.079	2.850	0.906	2.455	0.863	4.046	0.776	2.981
NFA	0.055	1.14	0.065	1.139	0.068	2.241	0.113	3.049
Error correction coef.					-0.089	-3.834	-0.092	-3.772

Notes: (1) Data are demeaned; (2) DOLS estimation includes 1 lead and lag in DOLS and standard error are based on Andrews and Monahan's Pre-whitening method; (3) We used the static fixed effects OLS estimates as initial estimates for the pooled maximum likelihood estimation; (4) The Newton-Raphson method was used the compute the log-likelihood function.

4 Real Equilibrium Exchange Rates and Misalignments

One of the reasons for finding the determinants of the real exchange rates is to be able to estimate their degree of misalignment. Hence, in this section, we use the long-run estimates of the economic fundamentals obtained from the DOLS and PMG estimations to compute the fitted values of the equilibrium real exchange rates (eq).

[23]We have used the Schwarz criterion to select the lags since, according to PSS, it performs slightly better than the Akaike, due to the fact the further one is a consistent model selection criteria while Akaike is not

[24]The joint Hausman test statistic is 0.52 (with a p-value of 0.77), the individual test of imposing equality on the CPI to PPI ratio is 0.25 (p value of 0.62) and on the NFA is 0.1 (with a p-value of 0.75)

[25]A fact that would be present in a *country-by-country* analysis.

[26]In the case of the PMG estimation, we also imposed homogeneity only in the PPI to CPI ratio, leaving the long-run estimate of the net foreign assets to differ across countries. According to the Hausman test, neither this restriction can be rejected. Yet, neither in this case the net foreign asset coefficient was significant at the 5% significance level.

However, computing equilibrium exchange rates is not straightforward. Indeed, as mentioned by Alberola (2003), finding a long-run cointegration relationship between the real exchange rate and its fundamentals would yield an estimate of the equilibrium rate if we were able to observe the equilibrium level of the fundamentals. Unfortunately, we can observe only the actual values of the variables.

A way to cope with this is to decompose the fundamentals into two components: the first one, when they are at their steady state, would be their equilibrium values. The second component, when the fundamentals are away from their respective steady state, would correspond to the deviations of their equilibrium value. Hence, we used the Hodrick-Prescott filter to decompose the fitted value of the two variables (CPI/PPI ratio and NFA) into their temporary and permanent movement.

To obtain the equilibrium exchange rate values ($EREER - DOLS$ and $EREER - PMG$) we used the permanent movement in the filtered series of long-run estimates of the fundamentals. Finally, we calculated the misalignment in the real exchange rate as the difference between the observed and the equilibrium exchange rate. The graphs of the series can be seen in the appendix D[27].

Several interesting facts come upon the inspection of the equilibrium exchange rates and misalignments between the G-20 currencies. First, the real equilibrium exchange rate tended to depreciate strongly in Argentina, Australia, Brazil, Canada, Indonesia and the UK. On the contrary, it appreciated, in Korea and, slightly, in the Euro-zone and Japan. In the rest of the countries. i.e. India, Mexico, South Africa, Turkey and the USA, the equilibrium values seemed quite stable during the period[28].

In this respect we could say that the depreciation in Argentina, Australia, Brazil, Canada and the UK was most likely due to both, a deterioration in their external position and a fall in the CPI to PPI ratio. In contrast, in Indonesia, the rather stable external position was more than overcome by a weakening in the relative productivity, leading to slight depreciation of the equilibrium exchange rate.

At the same time, the equilibrium value of the euro, the Japanese yen and the Korean won appreciated considerably in the last quarter of the century. The reason was a rising productivity *vis-à-vis* their trade partners and a strong improvement in the net asset position.

On the contrary, in a group of countries the equilibrium exchange rate remained relatively stable for different reasons. For instance, in India and South Africa the decline in the CPI/PPI ratio was counteracted by a fairly stable external position. Quite the opposite, regardless the deterioration in the stock of foreign assets, the equilibrium exchange rate remained relatively stable in Mexico and Turkey.

In the USA the effect of the fall in the net foreign assets position (i.e. the persistent deterioration of the current account) that should induce a depreciation, was offset by a stable or rising productivity. The result, then was a stable equilibrium exchange rate during the period.

Second, with respect to the deviations of the currencies from their equilibrium values, we have found that most of the countries went trough important period of over/under

[27]Values above (below) zero imply an overvaluation (undervaluation) of the real effective exchange rates. Notice that we should expect that the mean of the deviations in each country (misalignments) equals zero in the very long run.

[28]Since we do not have a long series either for the CPI to PPI ratio or on the net foreign assets or on both, we could not compute the equilibrium exchange rate for the whole period in the cases of Argentina, Brazil, India, Indonesia and Turkey. Thus, our estimates for these countries do not start in 1980.

valuations[29]. The only exception seems to be Canada, where the real exchange rate appears quite stable for the whole period, with a relatively small estimated overvaluation at the end of 2004 (between 3 and 5 percent).

In the USA, the currency was overvalued from 1980 to the end of 1986, peaking to most than 15% at the beginning of 1985. A slight undervaluation followed since 1987 for about ten years and then, again, the currency was overvalued from 1997 to 2001. However, in the last three years, there seems to be an evolution to balance.

In the euro zone, after an initial period of a minor overvaluation, that was followed by a few years of undervaluation, the deviations from the equilibrium were moderate until the year 2000. At the end of this year, the currency again went trough a slight phase of undervaluation that seemed to coincide with the start of the EMU. Yet, the euro has approximate its equilibrium value in the last few years.

Albeit the strong appreciation of the Japanese yen over the period, the crises incited persistent misalignments, placing the yen far below its long-run estimated value. For instance, at the end of the year 2004, the estimated undervaluation of the yen was around 12 percent.

For most of the period, the pound sterling was strongly misaligned, the undervaluation reaching levels of almost 16 percent at the end of 1986. However, the multilateral real exchange rate has approached gradually its equilibrium value, placing the estimated undervaluation in less than 2 percent at the end of the period.

The Latin American exchange rates also deviated considerably from their equilibrium values. For instance, in Argentina, after a relative stability, the small period of overvaluation was over with the peso crisis and the devaluation of the currency, that brought about a huge undervaluation of the real exchange rate at the beginning of 2002. In Brazil, the currency crisis at the end of 1999 was not enough to return the currency to its equilibrium level and the huge overvaluation was estimated to reached more 25 percent at the end of 2004. On the contrary, in Mexico, the peso displayed periods of strong undervaluation until 1996, when apparently it started to approach its equilibrium level.

In Asia, particularly in Indonesia and Korea, periods of overvaluation apparently preceded the currency crisis at the end of 1997. Nonetheless, in both cases, there seems to be a reduction in the misalignment and a convergence of the real exchange rate to equilibrium.

Finally, Australia, India, South Africa and Turkey also display persistent misalignments of the currencies, that do not seem to approach their equilibrium value.

5 Summary and conclusions

This paper provides a discussion on the methodological issues of applying different techniques to estimate, first, the long-run characteristics of the real exchange rates and, second, the relationship between real exchange rates and a set of selected macroeconomic fundamentals. Once we found the long-run determinants of the real exchange rate, we estimated it degree of misalignment in a group of countries.

We were particularly interested on the interaction of the major countries with emerging markets and developing countries, arguing not only that the countries share some common characteristics but also that the debate on exchange rate misalignment cannot

[29] Notice that the estimated misalignments are quite robust to the estimation method, i.e. they do not vary considerably if we use the pooled mean group estimation or the fully modified OLS.

be fully understand if emerging market economies are not considered[30]. Also, from the methodological point of view, we were interested on circumvent the problem of low-power display by conventional unit root tests by adding the cross-sectional to the time series dimension.

The main findings can be summarized as follows: (1) the real exchange rate prove to be non-stationary, not even when we increased the power of the tests by adding the cross-section to the time series data, but also allowing for the possibility of heterogenous structural breaks in unit root tests. Thus, our results strongly support that deviations from purchasing power parity are permanent and persistent, a finding that is in sharp contrast to those of Flood and Taylor (1996) or Frankel and Rose (1996); (2) in the long-term the real exchange rate depends on the progress of the relative productivity and on the evolution of the external position of the countries in relation to their trade partners and (3) based on the panel data approach, we calculated real equilibrium exchange rates and multilateral misalignments, finding that while for some countries there is a tendency of the real exchange rate to approach an equilibrium value, the misalignment is persistent for most of them during long periods of time.

Even though the findings of this paper are quite robust, it should be recognized that modelling real exchange rates raises many difficult issues not captured by a model that involves only the Balassa-Samuelson effect and the net foreign asset position.

Also, in this paper, we have maintained the hypothesis of a lineal autoregressive process for the real exchange rate, implying both, a continuous and a constant speed of adjustment towards equilibrium. Yet, the real exchange rates may behave more like a unit root process the closer they are from long-run equilibrium and, in opposition, become more mean reverting the further they are from equilibrium. If that was the case, standard unit root tests will have very low power. Thus, a natural extension of this work would be to investigate the role of nonlinearities in the real exchange rate adjustment toward long-run equilibrium.

[30]This is perhaps our main difference with some previous works on equilibrium exchange rates (as Alberola, Cervero, López, and Ubide (1999)).

Appendix

A Panel unit root tests with individual trend and intercept

Table A.1: **Panel unit root tests in the reer allowing for individual trend and intercept**

Test	$t*$	t_{bar}	W_{bar}	Z	HCZ	P_{MW}	Z_{MW}	$CIPS$	SN
Statistc	0.376	-1.833	1.606	8.789**	7.789**	26.268	-0.232	-2.112	1.516
Prob	0.354	0.609	0.946	0.000	0.000	0.558	0.592	0.840	0.935

Note: (1)*,**,*** indicates the rejection of the unit root null hypothesis at the 1%,5% and 10% significance level, respectively;

B Panel unit root tests with structural breaks

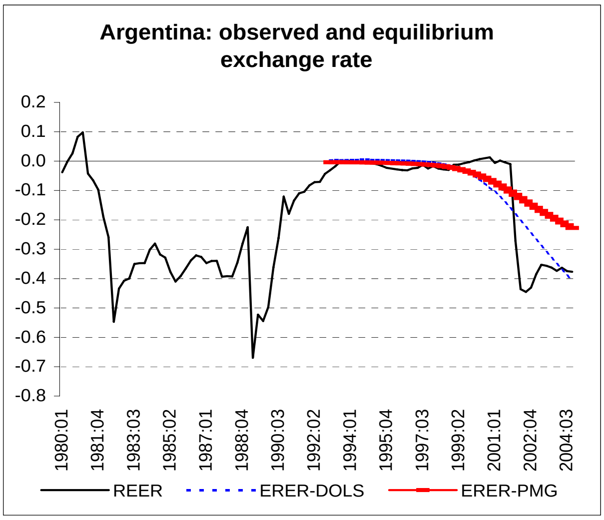
Table B.1: **Panel unit root tests on the PPI/CPI ratio and the NFA allowing for structural breaks**

	pd			NFA		
Test	Test Statistic	Critical Values		Test Statistic	Critical Values	
		5%	10%		5%	10%
LM_{NT}	-3.402	3.842	3.504	-3.186	3.842	3.504
$LM(\lambda)$						
homogenous	15.334	13.698	12.198	38.634	31.152	26.583
heterogenous	8.925	19.506	17.269	42.763	60.701	53.187

Notes: IDEM table (3.2)

C Break locations in the real exchange rate

D Real equilibrium exchange rates. An increase (decrease) indicates appreciation (depreciation)


Argentina: observed and equilibrium exchange rate
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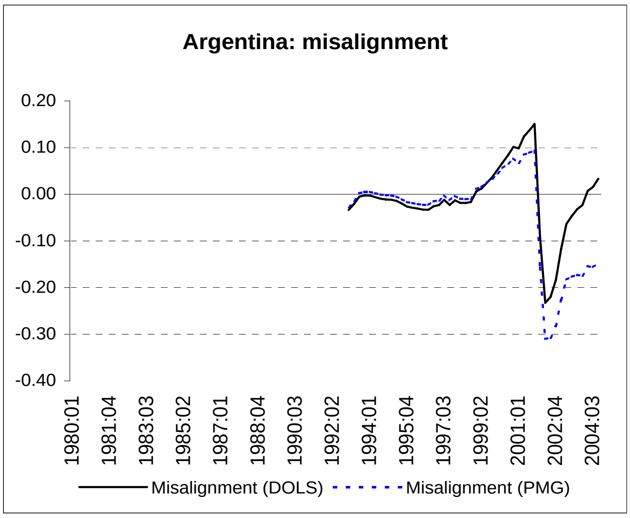
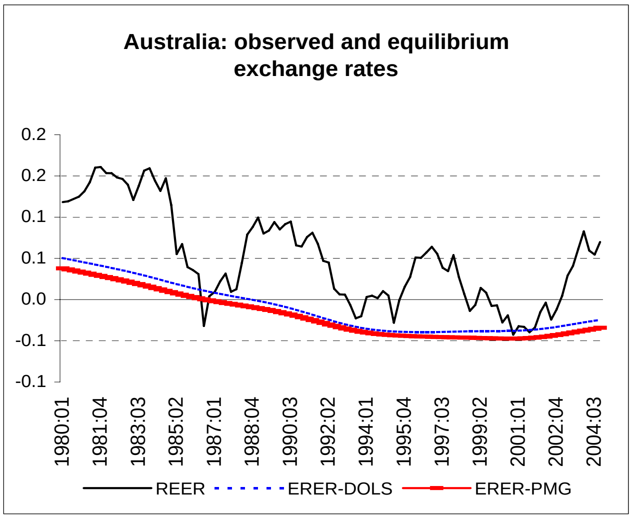
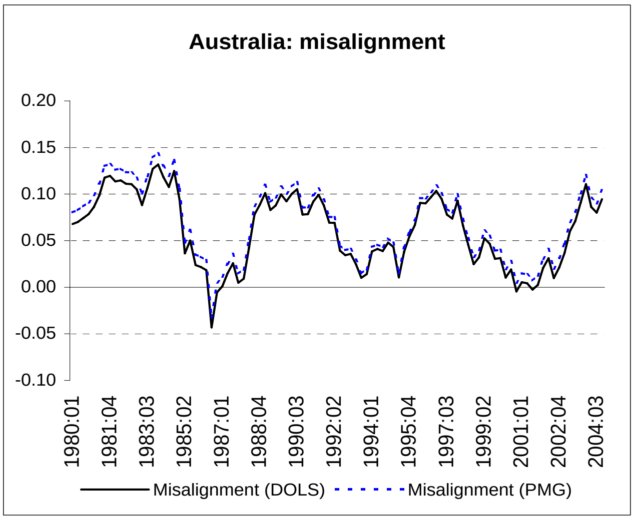
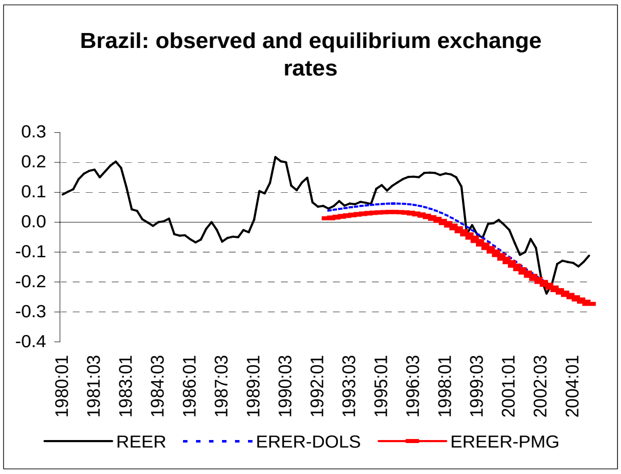
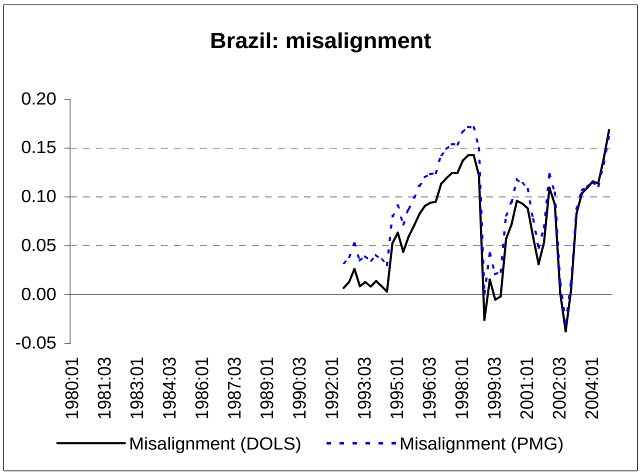
0.1
0.0
-0.1
-0.2
-0.3
-0.4
-0.5
-0.6
-0.7
-0.8
1980:01
1981:04
1983:03
1985:02
1987:01
1988:04
1990:03
1992:02
1994:01
1995:04
1997:03
1999:02
2001:01
2002:04
2004:03
REER
ERER-DOLS
ERER-PMG


Argentina: misalignment
0.20
0.10
0.00
-0.10
-0.20
-0.30
-0.40
1980:01
1981:04
1983:03
1985:02
1987:01
1988:04
1990:03
1992:02
1994:01
1995:04
1997:03
1999:02
2001:01
2002:04
2004:03
Misalignment (DOLS)
Misalignment (PMG)


Australia: observed and equilibrium exchange rates
0.2
0.2
0.1
0.1
0.0
-0.1
-0.1
1980:01
1981:04
1983:03
1985:02
1987:01
1988:04
1990:03
1992:02
1994:01
1995:04
1997:03
1999:02
2001:01
2002:04
2004:03
REER
ERER-DOLS
ERER-PMG


Australia: misalignment
0.20
0.15
0.10
0.05
0.00
-0.05
-0.10
1980:01
1981:04
1983:03
1985:02
1987:01
1988:04
1990:03
1992:02
1994:01
1995:04
1997:03
1999:02
2001:01
2002:04
2004:03
Misalignment (DOLS)
Misalignment (PMG)


Brazil: observed and equilibrium exchange rates
0.3
0.2
0.1
0.0
-0.1
-0.2
-0.3
-0.4
1980:01
1981:03
1983:01
1984:03
1986:01
1987:03
1989:01
1990:03
1992:01
1993:03
1995:01
1996:03
1998:01
1999:03
2001:01
2002:03
2004:01
REER
ERER-DOLS
EREER-PMG


Brazil: misalignment
0.20
0.15
0.10
0.05
0.00
-0.05
1980:01
1981:03
1983:01
1984:03
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1987:03
1989:01
1990:03
1992:01
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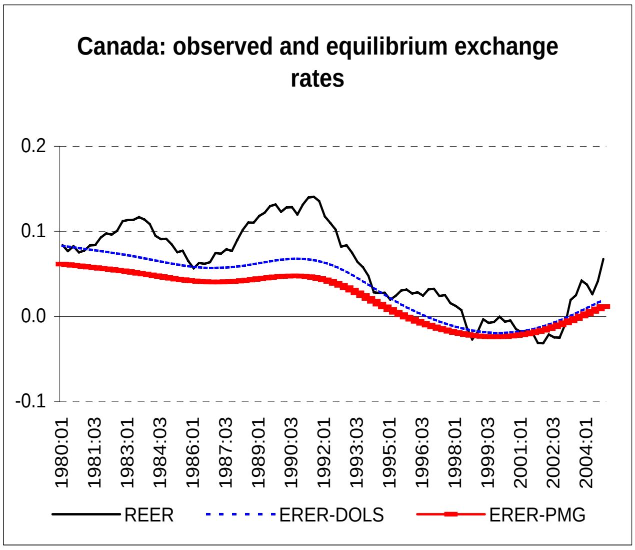
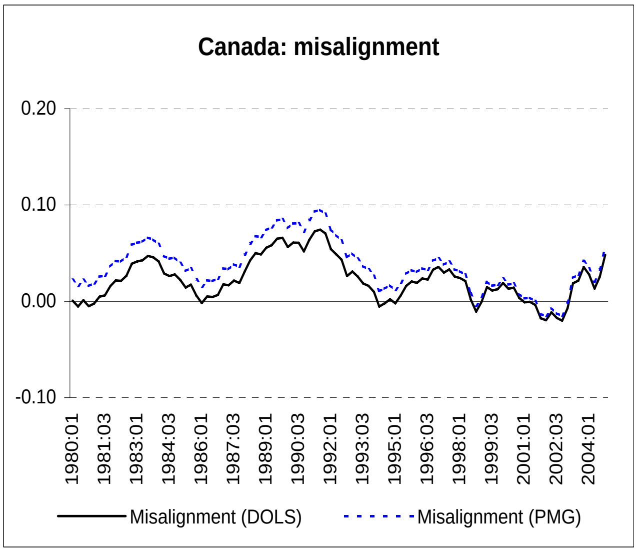
Misalignment (DOLS)
Misalignment (PMG)


Canada: observed and equilibrium exchange rates
0.2
0.1
0.0
-0.1
1980:01
1981:03
1983:01
1984:03
1986:01
1987:03
1989:01
1990:03
1992:01
1993:03
1995:01
1996:03
1998:01
1999:03
2001:01
2002:03
2004:01
REER
ERER-DOLS
ERER-PMG


Canada: misalignment
0.20
0.10
0.00
-0.10
1980:01
1981:03
1983:01
1984:03
1986:01
1987:03
1989:01
1990:03
1992:01
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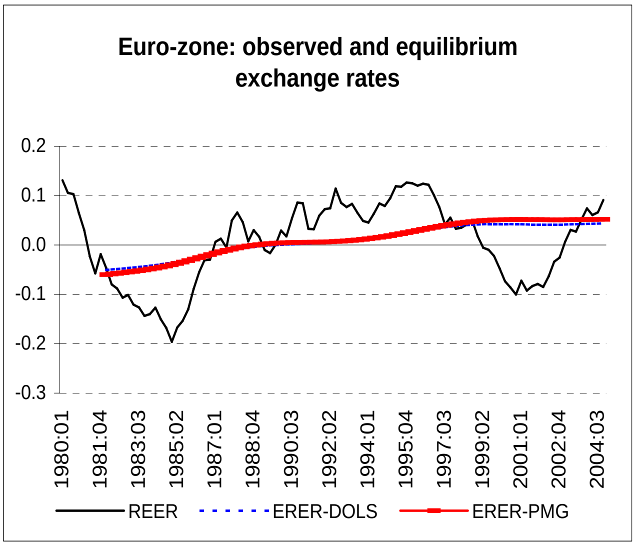
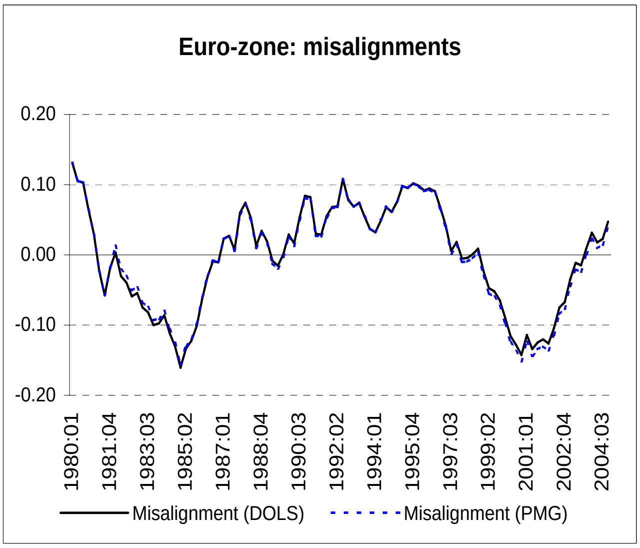
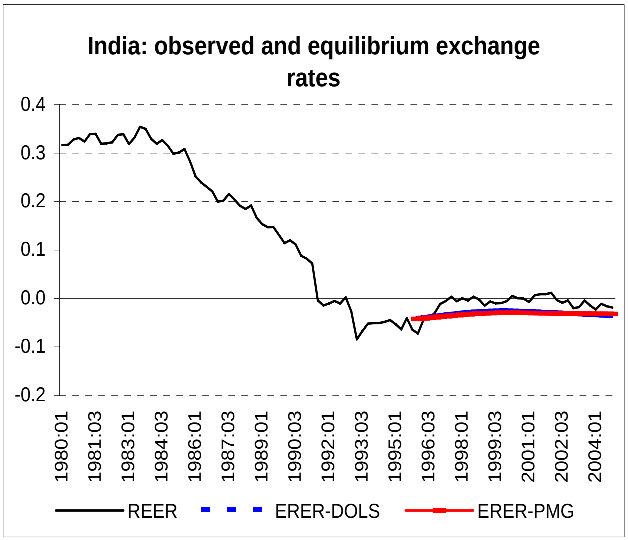
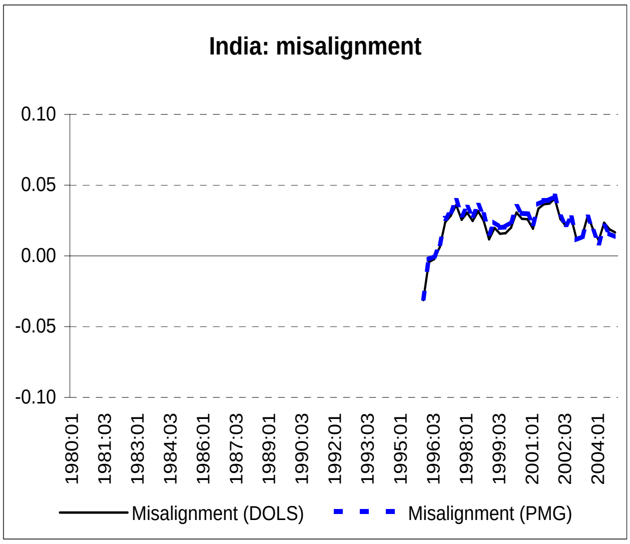
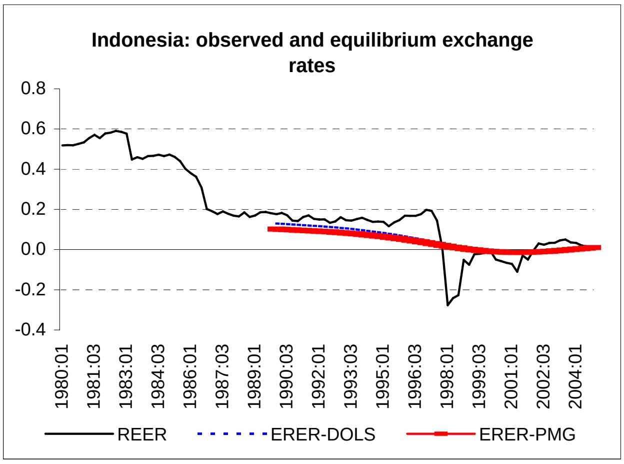
2004:01
Misalignment (DOLS)
Misalignment (PMG)

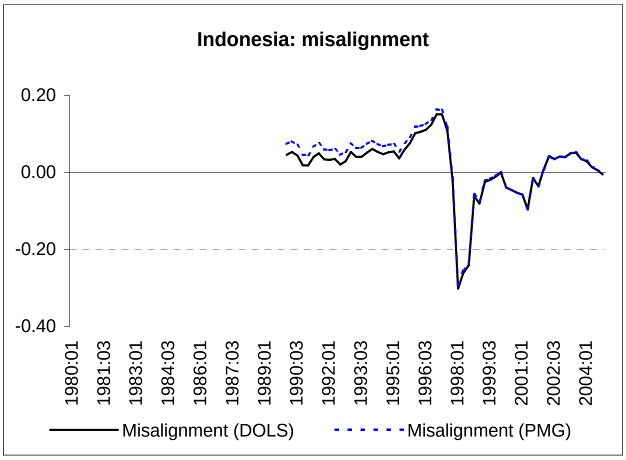

Euro-zone: observed and equilibrium exchange rates
0.2
0.1
0.0
-0.1
-0.2
-0.3
1980:01
1981:04
1983:03
1985:02
1987:01
1988:04
1990:03
1992:02
1994:01
1995:04
1997:03
1999:02
2001:01
2002:04
2004:03
REER
ERER-DOLS
ERER-PMG

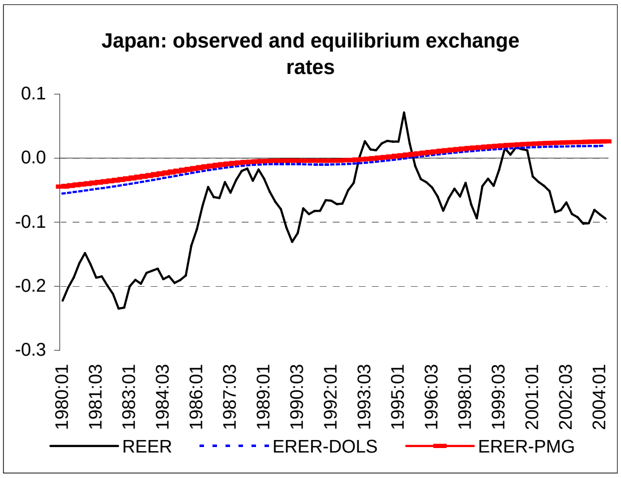

Euro-zone: misalignments
0.20
0.10
0.00
-0.10
-0.20
1980:01
1981:04
1983:03
1985:02
1987:01
1988:04
1990:03
1992:02
1994:01
1995:04
1997:03
1999:02
2001:01
2002:04
2004:03
Misalignment (DOLS)
Misalignment (PMG)

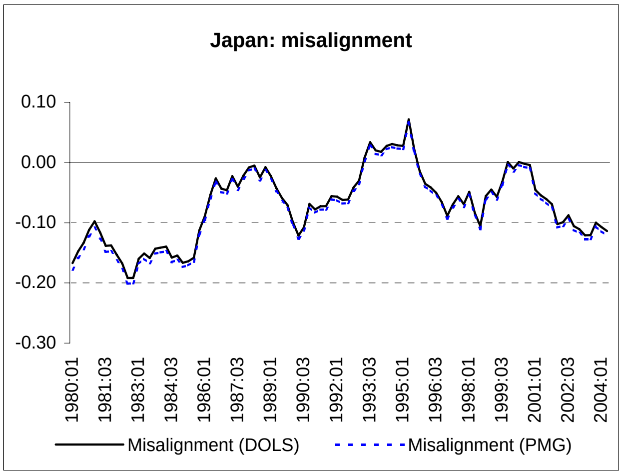

India: observed and equilibrium exchange rates
0.4
0.3
0.2
0.1
0.0
-0.1
-0.2
1980:01
1981:03
1983:01
1984:03
1986:01
1987:03
1989:01
1990:03
1992:01
1993:03
1995:01
1996:03
1998:01
1999:03
2001:01
2002:03
2004:01
REER
ERER-DOLS
ERER-PMG

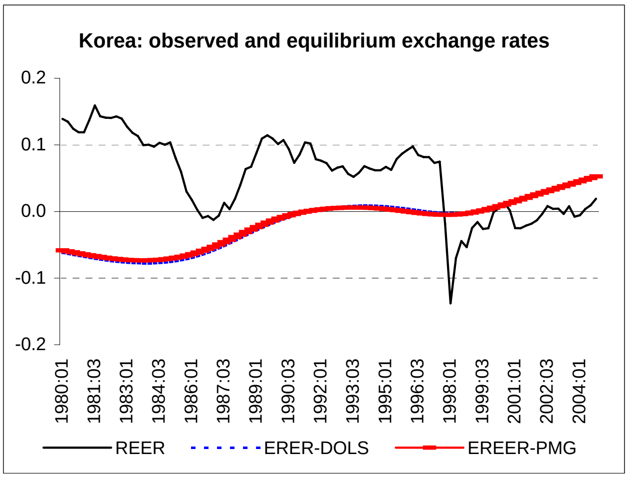

India: misalignment
0.10
0.05
0.00
-0.05
-0.10
1980:01
1981:03
1983:01
1984:03
1986:01
1987:03
1989:01
1990:03
1992:01
1993:03
1995:01
1996:03
1998:01
1999:03
2001:01
2002:03
2004:01
Misalignment (DOLS)
Misalignment (PMG)

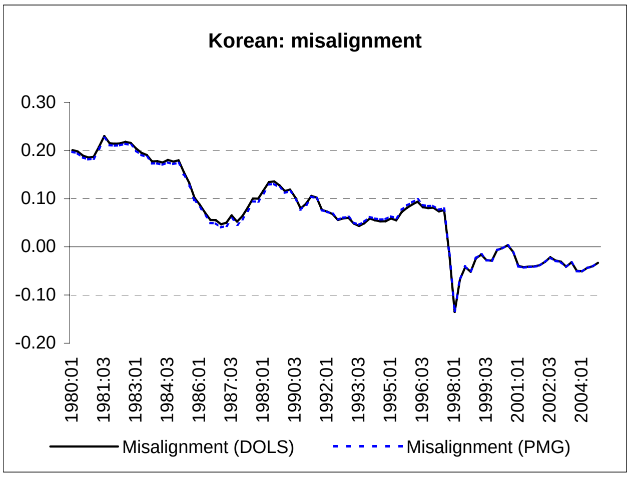

Indonesia: observed and equilibrium exchange rates
0.8
0.6
0.4
0.2
0.0
-0.2
-0.4
1980:01 1981:03 1983:01 1984:03 1986:01 1987:03 1989:01 1990:03 1992:01 1993:03 1995:01 1996:03 1998:01 1999:03 2001:01 2002:03 2004:01
REER
ERER-DOLS
ERER-PMG

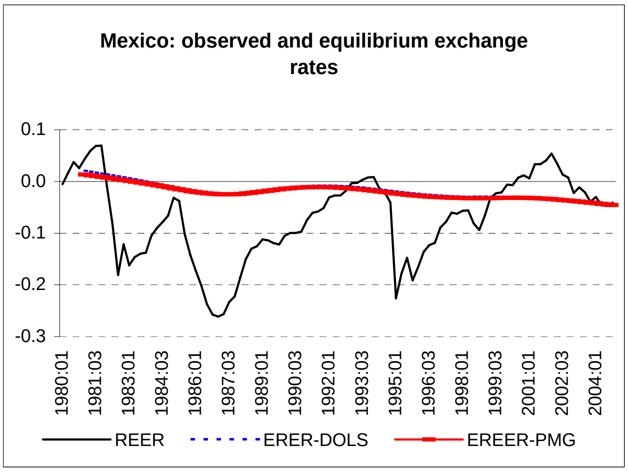

Indonesia: misalignment
0.20
0.00
-0.20
-0.40
1980:01 1981:03 1983:01 1984:03 1986:01 1987:03 1989:01 1990:03 1992:01 1993:03 1995:01 1996:03 1998:01 1999:03 2001:01 2002:03 2004:01
Misalignment (DOLS)
Misalignment (PMG)

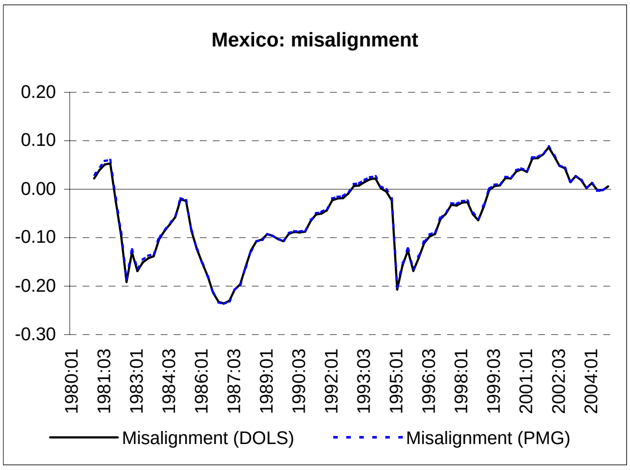

Japan: observed and equilibrium exchange rates
0.1
0.0
-0.1
-0.2
-0.3
1980:01 1981:03 1983:01 1984:03 1986:01 1987:03 1989:01 1990:03 1992:01 1993:03 1995:01 1996:03 1998:01 1999:03 2001:01 2002:03 2004:01
REER
ERER-DOLS
ERER-PMG

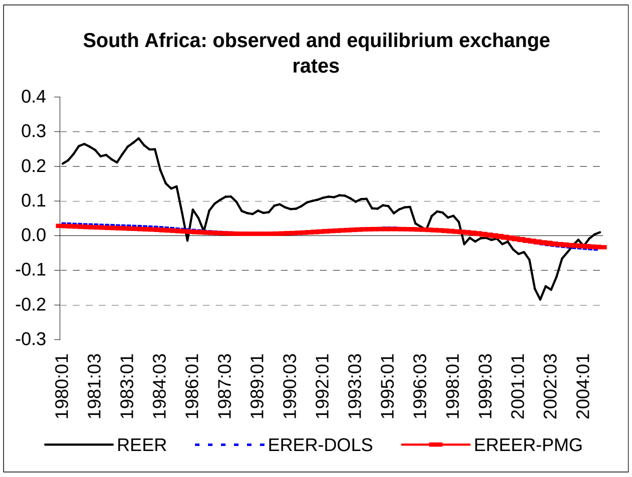

Japan: misalignment
0.10
0.00
-0.10
-0.20
-0.30
1980:01 1981:03 1983:01 1984:03 1986:01 1987:03 1989:01 1990:03 1992:01 1993:03 1995:01 1996:03 1998:01 1999:03 2001:01 2002:03 2004:01
Misalignment (DOLS)
Misalignment (PMG)

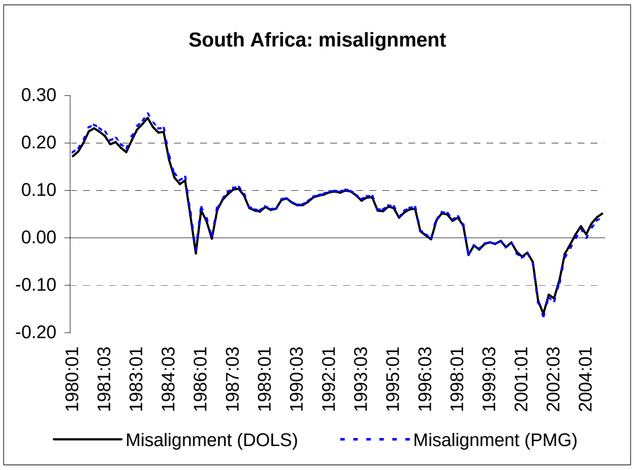

Korea: observed and equilibrium exchange rates
0.2
0.1
0.0
-0.1
-0.2
1980:01 1981:03 1983:01 1984:03 1986:01 1987:03 1989:01 1990:03 1992:01 1993:03 1995:01 1996:03 1998:01 1999:03 2001:01 2002:03 2004:01
REER
ERER-DOLS
EREER-PMG

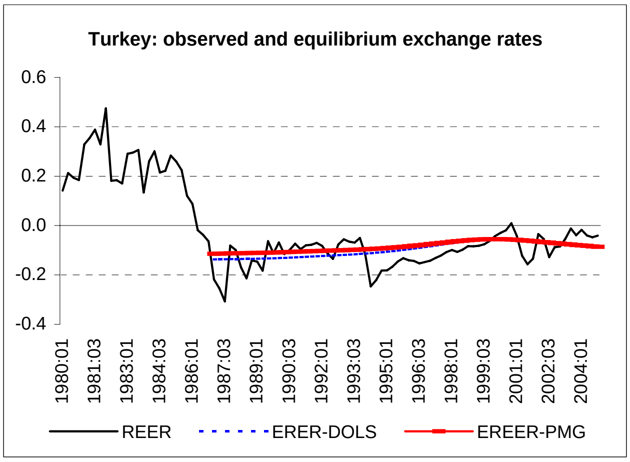

Korean: misalignment
0.30
0.20
0.10
0.00
-0.10
-0.20
1980:01 1981:03 1983:01 1984:03 1986:01 1987:03 1989:01 1990:03 1992:01 1993:03 1995:01 1996:03 1998:01 1999:03 2001:01 2002:03 2004:01
Misalignment (DOLS)
Misalignment (PMG)

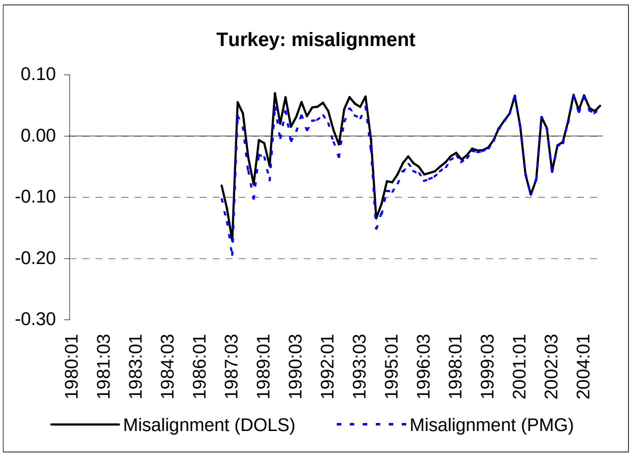

Mexico: observed and equilibrium exchange rates
0.1
0.0
-0.1
-0.2
-0.3
1980:01 1981:03 1983:01 1984:03 1986:01 1987:03 1989:01 1990:03 1992:01 1993:03 1995:01 1996:03 1998:01 1999:03 2001:01 2002:03 2004:01
REER
ERER-DOLS
EREER-PMG

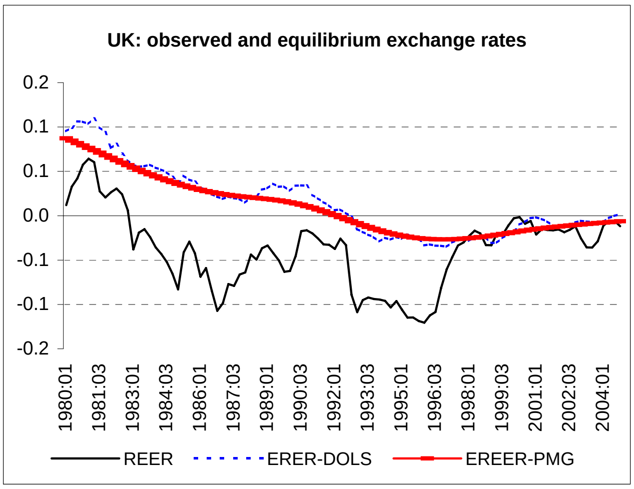

Mexico: misalignment
0.20
0.10
0.00
-0.10
-0.20
-0.30
1980:01 1981:03 1983:01 1984:03 1986:01 1987:03 1989:01 1990:03 1992:01 1993:03 1995:01 1996:03 1998:01 1999:03 2001:01 2002:03 2004:01
Misalignment (DOLS)
Misalignment (PMG)

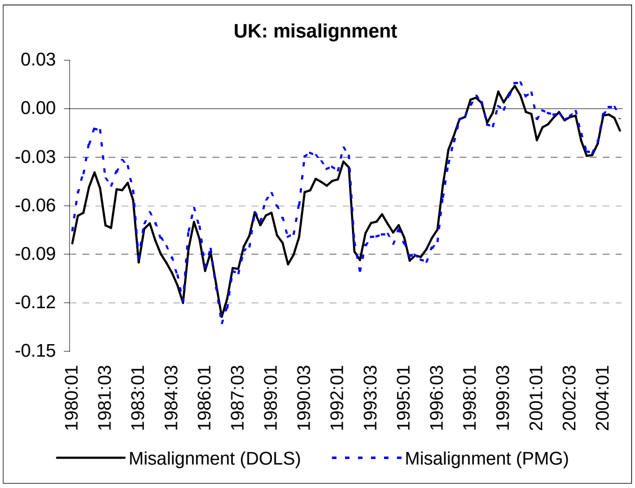

South Africa: observed and equilibrium exchange rates
REER
ERER-DOLS
EREER-PMG

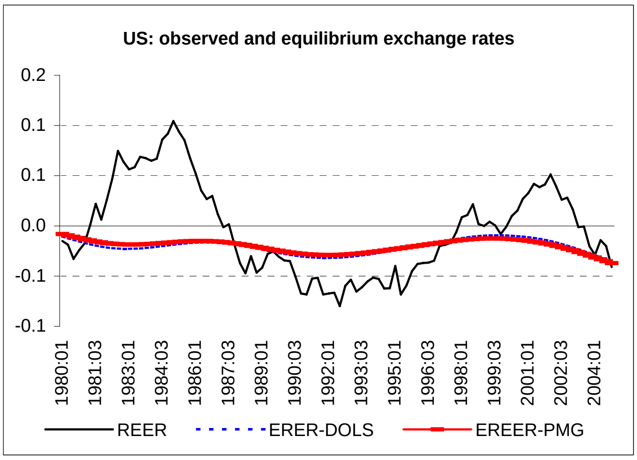

South Africa: misalignment
Misalignment (DOLS)
Misalignment (PMG)


Turkey: observed and equilibrium exchange rates
REER
ERER-DOLS
EREER-PMG


Turkey: misalignment
Misalignment (DOLS)
Misalignment (PMG)


UK: observed and equilibrium exchange rates
REER
ERER-DOLS
EREER-PMG


UK: misalignment
Misalignment (DOLS)
Misalignment (PMG)


US: observed and equilibrium exchange rates
REER
ERER-DOLS
EREER-PMG

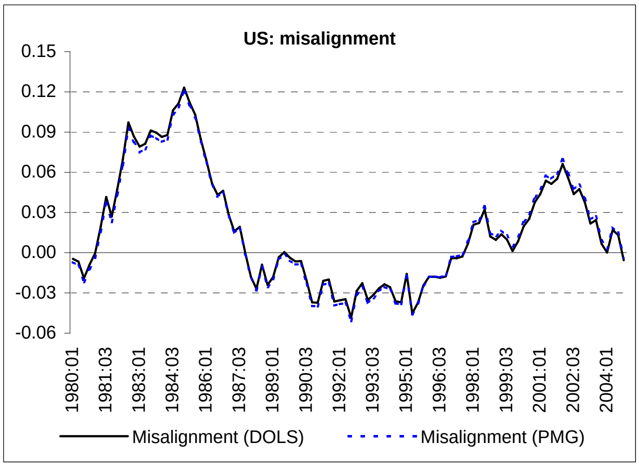

US: misalignment
Misalignment (DOLS)
Misalignment (PMG)

Table C.1: **Break locations in the real exchange rate according to $\bar{LM}_{NT}$ and $LM(\lambda)$ unit root tests**

Country	**Break location**	
	$\bar{LM}_{NT}$	$LM(\lambda)$
Argentina	1990q1,1990q3	1983q3,1990q2, 2001q1
Australia	1985q1, 1992q2	1985q1, 1988q4, 1995q4, 2001q1
Brazil	1989q1,1989q4	1988q4, 1998q4
Canada	1992q3, 1994q1	1985q1, 1991q2, 1995q1, 2001t1
Euro	1987q3,1990q2	1984t3, 1988t2, 1996q4, 2001t1
India	1988q3,1991q2	1983q3, 1991q2, 1996q3
Indonesia	1986q3, 1998q2	1986q3, 1997q4
Japan	1985q3, 1986q1	1985q4, 1989q4, 1995q3, 1999q2
Korea	1985q3, 1997q3	1985q4, 1990q2, 1997q3
Mexico	1982q4, 1985q2	1984q1, 1987q4, 1994q4, 2001q1
S. Africa	1984q2, 1998q2	1984q3, 1991q4, 2001q1
Turkey	1986q1, 1986q4	1986q1
UK	1982q4, 1996q3	1987q1, 1992q3, 1996q4,
USA	1986q4, 1987q3	1985q1, 1988q4, 1992q3, 2001q1

Notes:

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