Nonlinear Adjustment of the Real Exchange Rate Towards its Equilibrium Value: a Panel Smooth Transition Error Correction Modelling*

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Abstract

We study the nonlinear dynamics of the real exchange rate towards its behavioral equilibrium value (BEER) using a Panel Smooth Transition Regression model framework. We show that the real exchange rate convergence process in the long run is characterized by nonlinearities for emerging economies, whereas industrialized countries exhibit a linear pattern. Moreover, there exists an asymmetric behavior of the real exchange rate when facing an over- or an undervaluation of the domestic currency. Finally, our results suggest that the real exchange rate may be unable to unwind alone global imbalances.

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1 Introduction

The assessment of equilibrium values for the real exchange rate has always been an important issue in international macroeconomics, especially in the current context of global imbalances. Between the short-run market view and the PPP attractor supposed to

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hold at a remote time horizon, a wide range of intermediate approaches have been developed.[1] Among them, there is the BEER or "Behavioral Equilibrium Exchange Rate" model which has been introduced by Clark and MacDonald (1998) and has proved to be a consistent framework to derive equilibrium exchange rate values.[2] This approach consists in the estimation of a long-run (cointegrating) relationship between the real effective exchange rate and a set of economic fundamentals. The BEER value is then calculated by predicting the real effective exchange rate from the estimated long-run equation. Vector error correction models (VECM) are subsequently perfectly accurate to assess the speed at which the real exchange rate converges towards its equilibrium value.[3]

In this context, according to the standard macroeconomic view, any deviation from the equilibrium level is considered as temporary since there are forces ensuring quickly mean-reverting dynamics. However, in many countries, the experience of real exchange rates over the last two decades has been characterized by substantial misalignments, with time lengths much higher than suggested by the theoretical models (Dufrénot, Lardic, Mathieu, Mignon and Péguin-Feissolle, 2008). The fact that exchange rates can spend long periods away from their fundamental values implied a revival of interest in the study of exchange rate misalignments. Our aim is to contribute to this literature by investigating the dynamics of the adjustment process of the exchange rate towards its equilibrium value in a nonlinear panel framework.

The nonlinear cointegration support allows us to investigate the slowness of the adjustment process towards the long-run equilibrium. Numerous factors may explain such a nonlinear dynamics: transaction costs (Dumas, 1992; Sercu, Uppal and Van Hulle, 1995; O'Connell and Wei, 1997; Obstfeld and Taylor, 1997; Imbs, Mumtaz, Ravn and Rey, 2003), heterogeneity of buyers and sellers (Taylor and Allen, 1992), speculative attacks on currencies (Flood and Marion, 1999), presence of target zones (Krugman, 1991; Tronzano, Psaradakis and Sola, 2003), noisy traders causing abrupt changes (De Long, Shleifer, Summers and Waldmann, 1988), heterogeneity of central banks' interventions

[1] For recent surveys, see MacDonald (2000), Driver and Westaway (2004) and MacDonald and Dias (2007).

[2] See Bénassy-Quéré, Béreau and Mignon (2009) for a detailed study on the robustness of the BEER approach.

[3] An interesting study on this topic is provided by MacDonald and Dias (2007) who investigate exchange rate adjustments which are necessary to achieve equilibrium values, corresponding to Williamson (2006)'s scenarios.

(Dominguez, 1998). All these factors imply, either a nonlinear relationship between the exchange rates and the economic fundamentals, or a nonlinear adjustment mechanism with time-dependence properties. We consider here a smooth transition model for the adjustment process which can be viewed as a reduced form of structural models of fundamental exchange rate accounting for nonlinearities such as transaction costs, changing-regimes fluctuations,... Moreover, such models help at modelling asymmetries inherent to the adjustment process. This is particularly interesting since these asymmetries may explain, for instance, the unequal durations of undervaluations and over-valuations.

While numerous contributions have applied this nonlinear cointegration methodology in time series[4], this has not be done so far in the panel context. This constitutes a lack since we think that, to derive consistent equilibrium values of exchange rates, it seems important to work with a large panel of countries. Indeed, as noticed by Bénassy-Quéré, Duran-Vigneron, Lahrèche-Révil and Mignon (2004) among others, the large literature on equilibrium exchange rates has typically focused on country-by-country estimations of equilibrium exchange rates (Clark and MacDonald, 1998) or on consistent estimations of equilibrium exchange rates for a set of industrial economies (Williamson, 1994; Wren-Lewis and Driver, 1998). Until the mid-1990s, this approach was in line with a two-tier international monetary system, the first tier consisting in a small number of key currencies (the dollar, the Deutschemark, the yen and the British pound) and the second tier consisting in all other currencies. Since the mid-1990s, the rising share of emerging countries in global imbalances has made such divide no longer adequate and calls for the estimation of consistent sets of equilibrium exchange rates for a large number of currencies. To account for this evolution, we consider the G-20 in deriving our estimates of equilibrium exchange rates, a group that covers both industrial and emerging economies.

To sum up, the goal of this paper is to investigate the nonlinear behavior of the real exchange rate adjustment process towards its equilibrium value in a panel framework by estimating a Panel Smooth Transition Error Correction Model. To this end, the rest of the paper is organized as follows. Section 2 briefly sketches out methodological aspects relating to panel nonlinear models. Section 3 discusses our approach, data and their

[4]See, for instance, Michael, Nobay and Peel (1997); Ma and Kanas (2000); Chen and Wu (2000); Taylor, Peel and Sarno (2001); Baum, Barkoulas and Caglayan (2001); Dufrénot and Mignon (2002); Dufrénot, Mathieu, Mignon and Péguin-Feissolle (2006); Dufrénot et al. (2008); López Villavicencio (2008).

properties. Section 4 contains the estimation results and related comments, as well as robustness checks. Section 5 concludes.

2 Panel nonlinear models

2.1 PTR and PSTR models

In his seminal paper, Hansen (1999) introduced the panel threshold regression (PTR) model to allow regression coefficients to vary over time.
Let $\{y_{i,t}, s_{i,t}, x_{i,t}; t = 1, ..., T; i = 1, \ldots, N\}$ be a balanced panel with t denoting time and i the individual. Denoting $y_{i,t}$ the dependent variable, μ_i the individual fixed effects, $s_{i,t}$ the threshold variable and $x_{i,t}$ a vector of k exogenous variables, the PTR model can be written as follows:

$$y_{i,t} = \left\{ \begin{array}{ll} \mu_i + \beta_1' x_{i,t} + \varepsilon_{i,t}, & s_{i,t} \leq c \\ \mu_i + \beta_2' x_{i,t} + \varepsilon_{i,t}, & s_{i,t} > c \end{array} \right\} \quad (1)$$

In this model, the observations in the panel are divided into two regimes depending on whether the threshold variable is lower or larger than the threshold c. The error term $\varepsilon_{i,t}$ is independent and identically distributed. As in the time series context, the transition from one regime to another is abrupt and the model implicitly assumes that the two groups of observations are clearly identified and distinguished, which is not always feasible in practice.

To account for possible smooth and gradual transitions, González, Teräsvirta and van Dijk (2005) have introduced the panel smooth transition regression (PSTR) model, which is given by:[5]

$$y_{i,t} = \mu_i + \beta_0' x_{i,t} + \sum_{j=1}^{r} \beta_j' x_{i,t} g_j \left(s_{i,t}^{(j)}; \gamma_j, c_j \right) + \varepsilon_{i,t} \quad (2)$$

where $r + 1$ is the number of regimes, the $g_j \left(s_{i,t}^{(j)}; \gamma_j, c_j \right)$, $j = 1, ..., r$, are the transition functions, normalized and bounded between 0 and 1, $s_{i,t}^{(j)}$ the threshold variables which may be exogenous variables or a combination of the lagged endogenous one[6] (see van

[5] See also He and Sandberg (2004) and Fok, van Dijk and Franses (2005) who have introduced dynamic nonlinear panel models through the development of PLSTAR (panel logistic smooth transition autoregressive) models.

[6] As Fouquau (2008) reminds us, the endogenous variable must be lagged to avoid simultaneity problems.

Dijk, Teräsvirta and Franses, 2002), γ_j the speeds of transition and c_j the threshold parameters. Following Granger and Teräsvirta (1993) and Teräsvirta (1994) in the time series context or González et al. (2005) in a panel framework, the logistic specification can be used for the transition function:[7]

$$g(s_{i,t};\gamma,c) = \left[1+\exp\left(-\gamma\prod_{l=1}^{m}(s_{i,t}-c_l)\right)\right]^{-1} \tag{3}$$

with $\gamma > 0$ and $c_1 \leq c_2 \leq ... \leq c_m$. When $m = 1$ and $\gamma \to \infty$, the PSTR model reduces to a PTR model. González et al. (2005) mention that from an empirical point of view, it is sufficient to consider only the cases of $m = 1$ or $m = 2$ to capture the nonlinearities due to regime switching.[8]

2.2 Methodology

Following the methodology used in the time series context, González et al. (2005) suggest a three step strategy to apply PSTR models: (i) specification, (ii) estimation, (iii) evaluation and choice of the number of regimes (choice of r). Let us give some explanations about each of these steps.

The aim of the **identification** step is to test for homogeneity against the PSTR alternative. This can be done by testing the null hypothesis $\gamma = 0$. Due to the presence of unidentified nuisance parameters under the null, a first-order Taylor expansion around zero is used for the function g (see Lüükkonen, Saikkonen and Teräsvirta, 1988, or González et al., 2005):

$$y_{i,t} = \mu_i + \beta_0'^{*} x_{i,t} + \beta_1'^{*} x_{i,t}s_{i,t} + ... + \beta_m'^{*} x_{i,t}s_{i,t}^m + \varepsilon_{i,t}^{*} \tag{4}$$

where $\beta_1'^{*}, ...\beta_m'^{*}$ are multiple of γ and $\varepsilon_{i,t}^{*} = \varepsilon_{i,t} + r_m\beta_1' x_{i,t}$, r_m being the remainder of the Taylor expansion. Testing the null hypothesis of linearity is then equivalent to test $\beta_1'^{*} = ... = \beta_m'^{*} = 0$ in Equation (4). To this end, González et al. (2005) provide a LM-test statistic that is asymptotically distributed as a $\chi^2(mk)$ under the null.

[7]To simplify notations, we drop the j in the equation.

[8] Note that the case $m = 1$ corresponds to a logistic PSTR model and $m = 2$ refers to a logistic quadratic PSTR specification.

As in the time series context, this test can be used to select (i) the appropriate transition variable as the one that minimizes the associated p-value and (ii) the appropriate order m in Equation (3) in a sequential manner.

Turning to the **estimation** step, nonlinear least squares are used to obtain the parameter estimates, once the data have been demeaned. It should be noted that demeaning the data is not straightforward in a panel context (see Hansen, 1999, and González et al., 2005 for details).

The **evaluation** step consists in (i) applying misspecification tests in order to check the validity of the estimated PSTR model and (ii) determining the number of regimes. González et al. (2005) propose to adapt the tests of parameter constancy over time and of no remaining nonlinearity introduced by Eitrheim and Teräsvirta (1996) in the time series context. The test of no remaining nonlinearity, which is interpreted as a test of no remaining heterogeneity in panel data context, can be useful for determining the number of regimes of the PSTR model. To this end, González et al. (2005) suggest a sequential procedure starting by estimating a linear model, then a PSTR model if the homogeneity hypothesis is rejected, a PSTR model with 3 regimes if the no remaining heterogeneity hypothesis is rejected in the PSTR 2 regimes model, and so on.

3 Data and their properties

3.1 The model

As mentioned in the introduction, our aim is to study the possible nonlinear convergence process of the real exchange rate towards its long-run equilibrium value given by a BEER specification. Combining the BEER approach with the modelling of the short term dynamics and using the former notations for the PSTR model in Section 2, our complete model with two regimes can be written as follows:

$$\begin{aligned}\Delta q_{i,t} =& \mu_i + \underbrace{\theta z_{i,t-1} + \beta' \Delta X_{i,t}}_{\text{Regime 1}} + \\ & \Big[\underbrace{\theta^* z_{i,t-1} + \beta^{*\prime} \Delta X_{i,t}}_{\text{Regime } 2_{-1}}\Big] g\left(s_{i,t}; \gamma, c\right) + \varepsilon_{i,t}\end{aligned} \tag{5}$$

with:

$$g\left(s_{i,t};\gamma,c\right)=\Big[1+exp\Big(-\gamma\prod_{l=1}^{m}\left(s_{i,t}-c_l\right)\Big)\Big]^{-1}\text{for } m=1,2 \tag{6}$$

and:

$$z_{i,t}=q_{i,t}-\underbrace{\widehat{c_i}-\widehat{\beta}^{LT\prime}X_{i,t}}_{BEER_{i,t}} \tag{7}$$

where $q_{i,t}$ is the logarithm of the real effective exchange rate of country i (an increase in $q_{i,t}$ corresponds to a real depreciation of currency i), and $X_{i,t}$ is a vector of n fundamentals which are expected to influence the long run exchange rate (see the following sub-section for further details). $\widehat{c_i}$ and $\widehat{\beta}^{LT\prime}$ respectively stand for the estimated long-run fixed effect and coefficients from the linear cointegrating relationship between the real effective exchange rate and the explanatory variables (namely the linear panel BEER equation).

It has to be noticed that at any time, the coefficients of the explanatory variables in Equation (5) are given by: $c_x=\beta_x+\beta_x^{*}g(s_{i,t};\gamma,c)$ with $\beta_x=\theta,\beta_1,\beta_2,...,\beta_n$. When $g(s_{i,t};\gamma,c)=0$, then $c_x=\beta_x$ and the estimated coefficients correspond to those of the linear regime (Regime 1). At the other extreme (Regime 2 = Regime 1 + Regime 2_{-1}), i.e. when $g(s_{i,t};\gamma,c)=1$, then $c_x=\beta_x+\beta_x^{*}$. Between those two points, c_x takes a continuum of values depending on the realization of $g(s_{i,t};\gamma,c)$. As Fouquau (2008) mentions, there may exist some cases where the dynamics will never be described by those extreme regimes. It is then preferable to interpret the signs of the coefficients or their variations regarding those of the threshold variable rather than their magnitude.

3.2 Data and exchange rate determinants

Following Bénassy-Quéré et al. (2009), we concentrate on 15 countries or areas belonging to the Group of Twenty (G-20).[9] Regarding the long-run BEER equation, the dependent variable is the real effective exchange rate (q). Various explanatory variables may be used[10], but as noticed by MacDonald and Dias (2007) among others, usual determinants are: the net foreign asset position (nfa), the productivity differential (rpi), terms of trade (tot), and the interest rate differential (int); the later being rather a short or medium

[9] The exact composition of the G-20 sample is given in Appendix A.

[10] See among others Faruqee (1994) and MacDonald (1998a) for a general review of the real exchange rate determinants, or Egert, Halpern and MacDonald (2006) for a survey on equilibrium exchange rate models applied to transition economies.

run determinant of the exchange rate. We consider these four determinants and, as a measure of the productivity differential we retain the relative CPI-to-PPI ratio.[11] rpi and tot are in logarithms, nfa is expressed as share of GDP in percentage points and int is in percentage points. Data are annual and cover the period 1980 to 2005. We expect a positive link between the real effective exchange rate and all those potential determinants. Indeed, the real effective exchange rate is expected to appreciate (q to decrease) if (i) the net foreign asset position increases, due to implied net interest receipts, (ii) productivity in the tradable sector increases relative to the rest of the world, according to the Balassa-Samuelson effect, (iii) terms of trade follow an increasing trend, leading to an improvement of the trade balance, (iv) the foreign real interest rate decreases relative to the domestic one.

The real effective exchange rate $q_{i,t}$ for each country i is calculated as a weighted average of real bilateral exchange rates against each j trade partner. Denoting the variables in logarithms in lower cases, we can write:

$$q_{i,t} = \sum_{j \neq i} \omega_{ij} \left(e_{i,t} - e_{j,t} \right) = \sum_{j \neq i} \omega_{ij} e_{ij,t} \text{ where } \sum_{j \neq i} \omega_{ij} = 1 \tag{8}$$

where $e_{i,t}$ denotes the real bilateral exchange rate of currency i vis-à-vis the USD, $e_{ij,t}$ the one against the j currencies[12] and ω_{ij} the trade weights. Note that trade weights are given by the share of each partner in imports and exports of goods and services in 2005.[13]

Let us now turn to the explanatory variables. Concerning the net foreign asset position, we rely on the Lane and Milesi-Ferretti database from 1980 to 2004[14], the 2005 data being calculated by adding the current account position to the 2004 NFA value.[15] Regarding the

[11] To check the robustness of our results, we also consider other proxies of the productivity differential (see Section 4.4).

[12] Bilateral real exchange rates are based in 2000 and derived from nominal rates and CPI. The database is World Bank *World Development Indicators* (WDI) for nominal exchange rates and CPI data, except for the EUR/USD exchange rate which was extracted from Datastream and China's real exchange rate which was calculated with GDP deflator (WDI).

[13] Source: IMF *Direction of Trade Statistics* (DOTS). Intra-Eurozone flows have been excluded and trade weights have been normalized to sum to one across the partners included in the sample.

[14] Source: `http://www.imf.org/external/pubs/cat/longres.cfm?sk=18942.0`, see Lane and Milesi-Ferretti (2007).

[15] Source: IMF *International Financial Statistics* (IFS), March 2007. Unfortunately, valuation effects

CPI-to-PPI ratio, data were extracted from WDI and IFS (IMF *International Financial Statistics*) databases. We take the difference between the value for country i and the weighted average of its j partners' values as follows:

$$rpi_{i,t} = ln\left(\frac{CPI}{PPI}\right)_{i,t} - \sum_{j\neq i}\omega_{ij} ln\left(\frac{CPI}{PPI}\right)_{j,t} \tag{9}$$

Terms of trade are extracted from WDI and are given by:

$$tot_{i,t} = ln(TOT)_{i,t} - \sum_{j\neq i}\omega_{ij} ln(TOT)_{j,t} \tag{10}$$

The interest rate differential is obtained using a similar calculation:

$$int_{i,t} = r_{i,t} - \sum_{j\neq i}\omega_{ij} r_{j,t} \tag{11}$$

where r_i (resp. r_j) denotes country's i (resp. j) real interest rate (source: WDI).

Finally, note that, in addition to lagged values of exchange rate variations, misalignments and nfa, we will also consider three other potential transition variables in our PSTR estimation: the observed current account value ca of country i, the gap cag between the observed value of the current account of country i and its target value, and the gap $nfag$ between the observed value of the net foreign asset position of country i and its target value. Current account data were extracted from the WDI database. They are also expressed in proportion of GDP in absolute terms (the sum is supposed to be equal to zero, which is not the case in practice due to a large world discrepancy). The current account and net foreign asset position gaps are defined as follows:

$$cag_{i,t} = ca_{i,t} - \overline{ca}_{i,t} \text{ with } \overline{ca}_{i,t} = \phi(\overline{nfa}_{i,t})$$
$$nfag_{i,t} = nfa_{i,t} - \overline{nfa}_{i,t} \text{ with } \overline{nfa}_{i,t} = \psi(dem_{i,t}, gdebt_{i,t}, gdppc_{i,t})$$

where $dem_{i,t}$, $gdebt_{i,t}$ and $gdppc_{i,t}$ respectively stand for the demographic structure, the public debt-to-GDP ratio and the logarithm of GDP per capita; ϕ and ψ being linear functions.
$\overline{ca}_{i,t}$ and $\overline{nfa}_{i,t}$ denote the target values of the current account and the net foreign

cannot be included in the 2005 figure because the composition of gross assets and liabilities was not available.

asset position-to-GDP ratio respectively. To assess those values, we rely on the long-run net foreign asset position model proposed by Lane and Milesi-Ferretti (2001) and derive target values for the current account that are consistent with the reach of the equilibrium net foreign asset positions in 5 years.[16]

4 Estimation results

4.1 Estimation of the BEER equation

Before estimating the long term BEER relationship, we have to check that our series are cointegrated. To this end, we first proceed to the application of panel unit root tests. Five tests have been considered: Levin and Lin (1992, 1993), Maddala and Wu (1999), Breitung (2000), Im, Pesaran and Shin (2003), and Hadri (2000). Whereas the first four tests are based on the null hypothesis of a unit root, Hadri's test considers the null of stationarity. Our findings show that the only series for which the unit root hypothesis is always rejected is the interest rate differential.[17] This result is not surprising regarding the literature which generally concludes that the interest rate differential is not a key determinant of the real exchange rate in the long run, while being more important in the short run (for a survey relating to the link between the exchange rate and the interest rate differential, see MacDonald (1998b)). As a consequence, we drop this variable from the long-term relationship.

We then implement the seven panel cointegration tests proposed by Pedroni (1999, 2004). The results—not reported here, but available upon request—indicate that two cointegrating relationships exist: (i) a relationship between $q_{i,t}$, $nfa_{i,t}$ and $rpi_{i,t}$, and (ii) a relationship between $q_{i,t}$, $nfa_{i,t}$, $rpi_{i,t}$ and $tot_{i,t}$. In that follows, we retain the parsimonious specification given by the first relationship, which corresponds to the model studied by Alberola, Cervero, Lopez and Ubide (1999) that has proved to be consistent to numerous robustness checks as shown by Bénassy-Quéré et al. (2009). However, to check the robustness of our nonlinear estimations, we will also consider the second relationship for deriving a second set of misalignments (see Section 4.4). This will allow us to study the sensitivity of our results to the definition of the misalignments and, hence, to exchange rate fundamentals.

[16]See Bénassy-Quéré, Béreau and Mignon (2008) for further details on the specification and estimation of ϕ and ψ.

[17] All the results are available upon request to the authors.

The long-run relationship between the real exchange rate and the two considered explanatory variables, estimated using the panel Dynamic OLS procedure, is given by:

$$\widehat{q}_{i,t} = \widehat{\mu}_i - 0.331nfa_{i,t} - 0.829rpi_{i,t} \tag{12}$$

The results from the panel cointegration estimation appear consistent with the theory: the real exchange rate appreciates (q falls) in the long run if the net foreign asset position rises and if the tradable-to-non-tradable productivity ratio increases compared to the rest of the world (as a Balassa-Samuelson effect would suggest[18]).

4.2 The linear error correction model

As a first approximation, and for comparative purposes, we have estimated linear error correction models (ECM) for the whole panel (G-20) and for different groups of countries. Four sub-groups of countries are considered: the G-7 group, emerging countries (non G-7 group), Asian developing countries (Asia group) and countries that have overcome a recent financial crisis (denoted as 'Crisis' in the following tables).[19] The estimated model is the following:

$$\Delta q_{i,t} = \mu_i + \rho \Delta q_{i,t-1} + \theta z_{i,t-1} + \beta_1 \Delta nfa_{i,t} + \beta_2 \Delta rpi_{i,t} + \varepsilon_{i,t} \tag{13}$$

where $z_{i,t-1}$ corresponds to the past deviation of the real exchange rate from its equilibrium value as calculated in Equation (7) (i.e. the misalignment of the real exchange rate at year t-1). Given that Equation (13) is a dynamic panel data model, we have estimated it by the Generalized Method of Moments (GMM), which provides a convenient framework for obtaining efficient estimators in this context.[20] The results show that the dynamic term (ρ) was not significant in the specification in any of the different panels. Therefore, we have dropped the lagged exchange rate variations in our final estimation, keeping only the short-run fundamentals and the error correction term.[21]

[18] An alternative interpretation of this effect is that a positive shock on productivity in the tradable sector leads to a rise in intertemporal income, hence on the demand for both tradables and non-tradables. Because non-tradables cannot be imported, their relative price rises, which amounts to an exchange-rate appreciation. See, e.g., Schnatz and Osbat (2003).

[19] The composition of each country group is detailed in Appendix A.

[20] See among others, the seminal papers of Anderson and Hsiao (1982) and Arellano and Bond (1991).

[21] We have also estimated Equation (13) by Instrumental Variables (IV), finding similar results. To avoid too many tables, IV specifications are not presented here, but are available upon request to the authors.

As mentioned before, we are particularly interested in the characteristics of the adjustment speed of the real effective exchange rate towards its long-run equilibrium value (i.e. θ in Equation (13)). The theory of cointegration predicts that, if the real exchange rate and its fundamental determinants are cointegrated, we may expect a later reversal in case of a misalignment. Indeed, if the error correction coefficient is significantly negative, then a past undervaluation of currency i (resp. over-valuation) will generate a current real appreciation (resp. depreciation) of currency i vis-à-vis the j currencies. In other words, if $z_{i,t-1}$ is positive (resp. negative), meaning that currency i is undervalued (resp. over-valued), a negative sign of θ will guaranty a current appreciation (resp. depreciation) of the current real exchange rate corresponding to a decrease (resp. increase) in $q_{i,t}$. Table 1 reports the GMM estimates of the error correction coefficient in our final linear specification for the whole G-20 panel and the different sub-groups of countries.

Table 1: **GMM estimates of the error correction coefficient - linear specification**

	G-20	G-7	Non G-7	Asia	Crisis
θ	-0.155	-0.156	-0.132	-0.129	-0.089
T-stat	-4.23	-5.27	-3.19	-3.78	-2.0

As expected, we find a negative and statistically significant error correction term in each case, implying that if the fundamentals in the last period dictate a lower (resp. upper) real exchange rate than that observed, then the real exchange rate will strictly depreciate (resp. appreciate) in the current period. The (average) error correction coefficients reported here show that the adjustment is relatively important (between 9% and 16% taking place within a year).

4.3 Nonlinear error correction model

The linear ECM implicitly assumes that the adjustment speed towards equilibrium is both continuous and constant, regardless of the extend of the real misalignment. However, as mentioned before, we may imagine that the convergence speed increases with the size of the deviation from equilibrium, a feature that the previous linear model would not be able to capture. In that case, Equation (13) could be better approximated by a panel nonlinear model.

To formally analyze this possibility, we have tested linearity in model (13)[22] using the González et al. (2005) test with different possible transition variables. First, we use the lagged estimated cointegrating vector ($z_{i,t-1}$) as the appropriate threshold variable. This model is particularly attractive from an economic point of view as it implies the existence of a lower threshold (whether a logistic function is used in Equation (5) with $m = 1$) or a band (whether the function is a logistic quadratic one, i.e. $m = 2$ in Equation (5)) above or outside which there is a strong tendency for the real exchange rate to revert to its equilibrium value.[23] In addition, we have tested for nonlinearity using $\Delta q_{i,t-1}$ as the threshold variable. This specification is also attractive, since it allows the adjustment speed to vary whether the real exchange rate appreciates (when $\Delta q_{i,t-1}$ is below a threshold, c) or depreciates (when $\Delta q_{i,t-1}$ is above c).

The results are summed up in Tables 2 and 3. They show that, when the past misalignment is used as the threshold variable (Table 2), linearity is strongly rejected for all groups of countries, except for the panel composed of industrialized countries alone (namely the G-7 countries), where linearity seems to be a pattern. Therefore, we estimated the corresponding panel smooth transition regression models for the G-20, the emerging markets (non G-7 countries), Asian emerging markets (Asia) and countries having overcome a recent financial crisis (crisis).

Table 2: **PSTR model with $z_{i,t-1}$ as the threshold variable**

	Regime 1		Regime 2		Transition	
	θ	T-stat	$\theta + \theta^*$	T-stat	γ	c
G-20	-0.031	-0.54	-0.245	-4.14	17.461	-0.143
G-7	Linear					
Non G-7	0.024	0.404	-0.255	-2.75	18.013	-0.092
Asia	0.037	0.63	-0.330	-4.06	41.846	-0.018
Crisis	0.097	1.01	-0.240	-2.06	16.955	-0.112

Notes: Model chosen according to BIC and the lowest p-value in the linear tests.

[22] As mentioned in the previous section, the coefficient of the lagged endogenous term was not significant. That is why we have dropped $\Delta q_{i,t-1}$ from our final specification, which allows us to apply the PSTR methodology since our model does not contain any dynamic component.

[23] We have discriminated between logistic and logistic quadratic panel smooth transition functions according to two criteria: we selected first those with the lowest p-value in the linear test and then selected the one that exhibited the lowest Schwarz information criterion (BIC).

Table 3: **PSTR model with $\Delta q_{i,t-1}$ as the threshold variable**

	Regime 1		Regime 2		Transition	
	θ	T-stat	$\theta + \theta^*$	T-stat	γ	c
G-20	Linear					
G-7	1.029	3.77	-0.251	-4.12	27.555	-0.145
Non G-7	Linear					
Asia	Linear					
Crisis	Linear					

Notes: Model chosen according to BIC and the lowest p-value in the linear tests.

The main parameters of interest here are the error correction coefficients in the two extreme regimes θ and $\theta + \theta^*$, the threshold parameter c and the speed of transition γ.[24] Regarding the results for the G-20, the threshold estimate is -0.143 (corresponding to an over-valuation of 14%) which is the lower band below which deviations from the real exchange rate equilibrium level (i.e. when $g\left(q_{i,t};\gamma,c\right)=0$) are not corrected. Note that θ is not significant in the first regime, which means that there is no convergence process towards the BEER value for the real exchange rate in t when the over-valuation exceeds 14 pp. However, once the misalignment crosses this threshold, there is a strong tendency of the real exchange rate to go back to its equilibrium value ($\theta + \theta^*$ is significant and strongly negative in the second regime).

This result can be understood as a confirmation of the asymmetric property of the real exchange rate adjustment towards equilibrium. Indeed, as the distribution of the threshold variable confirms (see Figures 1 and 2 in Appendix B), even if the threshold c is not fixed at 0,[25] most of the points that are above the threshold are positive figures (i.e. there are more points above 0 than between the threshold and 0). This implies that the adjustment process is more effective in case of an undervaluation than when an overvaluation occurs. This result is particularly true for emerging economies and developing

[24] In most of the cases, the logistic transition function shows better properties than the logistic quadratic one. This implies that the predominant type of asymmetry is that which distinguishes between positive or negative deviations from equilibrium. In other words, the short-term adjustment that occurs, being nonlinear, corrects deviations from the equilibrium position by giving more weight to the sign of the deviations - whether it is an over-valuation or an undervaluation of the currency i - than to their magnitude. Our results are then based on these models.

[25] Recall that a negative (resp. positive) value for $z_{i,t-1}$ corresponds to an over (resp. under) valuation of the real exchange rate.

Asia sub-samples, with threshold variables estimated at −0.092 and −0.018 respectively. This is consistent with the fact that emerging countries' currencies, and especially those of China, Indonesia and India, appear rather undervalued in average (see Bénassy-Quéré et al. (2008)).

It is important to notice that the convergence process in the nonlinear model is more pronounced than that in the linear specification, with 24% of the adjustment taking place within a year corresponding to a half-life of 3.2 years versus 4.8 in the linear estimation for the G-20. In the rest of the groups, the adjustment is even higher with respect both to the nonlinear G-20 specification and to the figures obtained in the linear models. The correction is particularly crucial below an appreciation of 2% in emerging Asia (reaching 33% within a year, which corresponds to a half-life of 2.4 years versus 5.7 years in the linear estimation).

Figures 3, 4, 5, 6 and 7 report the values of both the threshold variable and the transition function against time for each country belonging to our different sub-panels. For all the considered groups, the movements of the disequilibrium error above (below) zero are associated with undervalued (overvalued) real exchange rate. As it can be noticed, undervaluations are corrected faster than over-valuations, confirming our former conclusions. Besides, the transition function changes from the lower to the higher regime quite often. As a result, the transition function is, indeed, smooth and observations are distributed from each side of the threshold, with a relatively higher presence of observations above the threshold.

However, the case of the advanced economies alone is completely different from the rest of the panel. Indeed, the first interesting feature in this group is that linearity is not rejected when the previous misalignment is used as threshold variable (Table 2). Therefore, in industrialized countries, reversion to equilibrium is a characteristic that happens regardless of the size of the deviation from equilibrium, confirming previous studies in time series (see López Villavicencio (2008) among others). Second, when the selected transition variable is the real exchange rate variation ($\Delta q_{i,t-1}$), linearity cannot be rejected in any of the other panels but the G-7 (Table 3). For those groups of countries it is more past misalignments that matter than the magnitude of exchange rate variations.

The estimated parameters of the nonlinear model for the G-7 can be found in Table 3. As observed, reversion is much higher in the nonlinear model above a depreciation of 14% than in the linear specification, with associated half-lives of 3.1 and 5.6 years respectively. Yet, as observed on Figure 4, this has only been the case in Japan between 1987-88 and in the euro zone in 1987. Therefore, the consistency of our results with respect to the nonlinear behavior in the short-run adjustment model seems to depend critically on the presence of just a few observations.[26] As expected, this is reflected in the transition function showing most of the observations to the right of the location parameter, where reversion to equilibrium is higher.

We also checked the linearity of the adjustment process with net foreign asset and current account gaps as threshold variables (see Section 2 for the construction of data). Indeed, it could have been reasonable to think that, as the BEER corresponds to an exchange rate level consistent with the net foreign asset position being at an equilibrium value (characterized by $\overline{nfa}_{i,t}$), the adjustment speed would be fastened if the gap between the current and the equilibrium values had gone beyond a certain threshold. The same explanation holds for the current account gap, the stabilization of the stock implying that of the flow. However, when NFA or current account gaps are used as the threshold variable ($s_{i,t}$) in our PSTR model (see Equation (5)), the null hypothesis of linearity cannot be rejected.[27] This means that according to those two specifications, the real exchange rate will converge linearly towards its long-run BEER value for all our sub-groups of countries, whatever the size of the disequilibria in the NFA position or in the current account balance. In standard models, the real exchange rate is the only channel through which the current account may adjust. This explanation appears insufficient regarding our empirical result, since it implies that the real exchange rate adjusts totally independently from the level of global imbalances. The real exchange rate cannot be then the only adjustment variable, which confirms previous results on the pre-eminence or coordinated/combined effects of other factors, such as wealth effects or valuation effects.[28] Ultimately, it appears that the real exchange rate may not solve by its own global imbalances.

[26] In order to check this, we eliminate Japan and the euro zone from this group and proceed to linearity tests. The results confirm our intuitions since the null of linearity is not rejected.

[27] All results are available upon request to the authors.

[28] See Gourinchas (2007) for more details on valuation effects. In particular, our findings corroborate those of Bénassy-Quéré et al. (2008) showing that the real exchange rate may probably not be the key of global imbalances' unwinding.

4.4 Robustness checks

As previously mentioned, since many approaches have been developed to estimate long-run exchange rate values, there is not a unique "measure" of equilibrium exchange rate and misalignment. Indeed, it is well known that the concepts of equilibrium and misalignment depend on the underlying model, i.e. on the considered exchange rate fundamentals. Another crucial point when estimating equilibrium exchange rates concerns the choice of the productivity measure.[29]

In the previous sections, our real exchange rate misalignments were based on a reduced-form equation that includes the net foreign asset to GDP ratio and the relative productivity differential proxied by the relative CPI-to-PPI ratio (Equation (12)) as the two main fundamentals. By doing so, we showed that the speed of adjustment towards equilibrium, rather than being continuous and constant, depends on the size and sign of the misalignment.

Yet, the exclusion of some important fundamentals, such as terms of trade, as well as the choice of the relative productivity measure can be important when computing exchange rate misalignments and, therefore, may have an impact on our previous results. To test the robustness of our main conclusion—the evidence of a different adjustment of the exchange rate towards its equilibrium value depending on the size and sign of the misalignment—we conduct two additional sets of estimations, investigating (i) the impact of accounting for terms of trade in the BEER equation, and (ii) the influence of the productivity measure.

4.4.1 Inclusion of terms of trade in the BEER equation

In addition to the net foreign asset position of a country and the Balassa-Samuelson effect, the real exchange rate can also be affected by commodity price shocks through their impact on the terms of trade and the international competitiveness of a country. Overall, a lasting deterioration of the terms of trade of a country should result in a depreciation of its real exchange rate.
Given the non-stationarity of the terms of trade and because there also exists a cointegrating relationship when this variable is included as a determinant of the real exchange rate, we proceed to the estimation of the PSTR model by using a new misalignment series based on the following long-term relationship:

[29] See, among others, De Gregorio, Giovanni and Krueger (1994), Canzoneri, Cumby and Diba (1999), MacDonald (1998a), Schnatz and Osbat (2003), and Bénassy-Quéré et al. (2008).

$$\widehat{q}_{i,t} = \widehat{\mu}_i - 0.283nfa_{i,t} - 0.878rpi_{i,t} - 0.419tot_{i,t} \quad (14)$$

As for Equation (12), all the determinants have the expected sign, with a depreciation of the exchange rate if terms of trade decline. In addition, the inclusion of the terms of trade yields very similar estimates to our previous measure of real exchange rate misalignment.

Based on this alternative series, we estimate a linear and a nonlinear specifications (Model 2 in Tables 4 and 5 respectively). By looking at the sign and size of the error correction term in both specifications, we can see that the inclusion of the terms of trade insures that our results are not biased by the omission of this variable. Indeed, not only the speed of convergence is higher in the nonlinear specification versus the linear one, but also it increases as the size of the undervaluation gets higher. As previously, θ is not significant in the first regime, meaning that there is no convergence process towards the equilibrium value when the currency over-valuation exceeds 19 pp. Our main finding relating to the existence of a different adjustment process of the currency when facing over or under-valuations is thus robust to the inclusion of terms of trade as a determinant of the real exchange rate.

Table 4: **Linear estimates of the error correction coefficient with additional fundamentals and alternative measures of productivity (complete G-20 panel)**

	Model 2 nfa, rpi & tot	Model 3 nfa & prod1	Model 4 nfa & prod2
θ	-0.136	-0.094	-0.120
T-stat	-5.80	-3.65	-4.30

Notes: nfa: net foreign asset to GDP; rpi: relative CPI-to-PPI ratio; tot: terms of trade; prod1: relative GDP per capita; prod2: relative GDP per person employed.

4.4.2 The choice of the productivity measure

As mentioned before, the previous estimations were based on the CPI-to-PPI ratio, which serves as a proxy for the relative productivity in the tradable sector. As a robustness check and to provide a complete analysis, we re-estimate effective misalignments and PSTR models using other productivity proxies. Following Bénassy-Quéré et al. (2008) among others, we consider two alternative measures: (i) the relative GDP per capita

($prod1$), and (ii) a measure of relative labor productivity based of the number of persons employed ($prod2$).[30]

Table 5: **Robustness checks. PSTR model with additional fundamentals and alternative measures of productivity (complete G-20 panel)**

Model	Regime 1		Regime 2		Transition	
	θ	T-stat	$\theta + \theta^*$	T-stat	γ	c
Model 2: nfa, rpi & tot	-0.0843	-1.37	-0.1801	-3.34	12.7398	-0.1899
Model 3: nfa & prod1	0.0150	0.21	-0.1239	-2.53	14.2058	-0.2684
Model 4: nfa & prod2	0.0382	0.46	-0.1614	-3.15	12.6578	-0.2533

Notes: nfa: net foreign asset to GDP; rpi: relative CPI-to-PPI ratio; tot: terms of trade; prod1: relative GDP per capita; prod2: relative GDP per person employed.

Once again, the two additional PSTR models in Table 5 (Models 3 and 4) are very similar to our baseline estimation. The error correction term is significant only in the second regime, when the misalignment exceeds a threshold of around 25 pp. On the whole, these findings offer solid evidence that, whatever the considered measure of productivity, there is a different adjustment of the exchange rate towards its equilibrium value, depending on the size and sign of the misalignment.

5 Conclusion

In this paper, we have studied the nonlinear convergence process of the real exchange rate towards its equilibrium BEER value using a Panel Smooth Transition Regression model framework. We have shown that the real exchange rate dynamics in the long run is proved to be nonlinear for emerging economies, whereas industrialized countries exhibit a linear pattern, confirming previous studies in time series (see López Villavicencio (2008) among others). More especially, there exists an asymmetric behavior of the real exchange rate when facing an over- or an undervaluation of the domestic currency. The adjustment speed appears drastically accelerated in case of an undervaluation, which is consistent with the fact that developing economies and especially emerging Asian countries are more inclined to exhibit undervalued currencies. Another conclusion of our findings is that the convergence process towards the long-run equilibrium is independent from the magnitude of the current account or the net foreign asset imbalances, which confirms

[30] These data are obtained from WDI and the Groningen database.

that the real exchange rate may be unable to solve alone global imbalances as suggested by Bénassy-Quéré et al. (2008).

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A Tables

Table 6: **Country samples**

G-20	Argentina (ARG), Australia (AUS), Brazil (BRA), Canada (CAN), China (CHN), United Kingdom (GBR), Indonesia (IDN), India (IND), Japan (JPN), Korea (KOR), Mexico (MEX), Turkey (TUR), United States (USA), South Africa (ZAF), and Euro area (ZZM)
G-7	Australia (AUS), Canada (CAN), United Kingdom (GBR), Japan (JPN), United States (USA), and Euro area (ZZM)
Non G-7	Argentina (ARG), Brazil (BRA), China (CHN), Indonesia (IDN), India (IND), Korea (KOR), Mexico (MEX), Turkey (TUR), and South Africa (ZAF)
Asia	China (CHN), Indonesia (IDN), India (IND), and Korea (KOR)
Crisis	Argentina (ARG), Brazil (BRA), Indonesia (IDN), Korea (KOR), Mexico (MEX), and Turkey (TUR)

B Graphics

Figure 1: **Kernel density estimate of $z_{i,t-1}$**

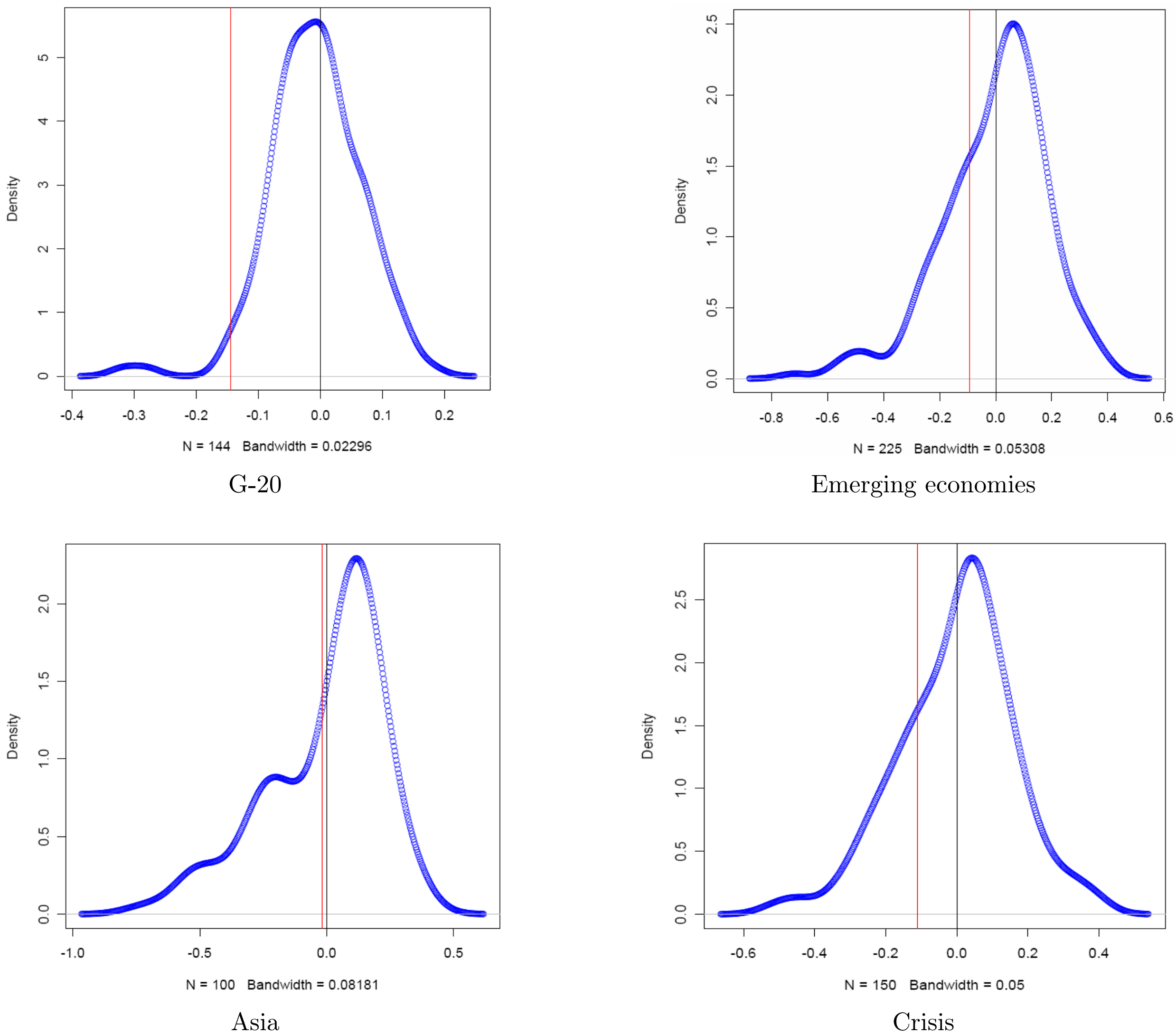


Figure 2: **Kernel density estimate of** $\Delta q_{i,t-1}$

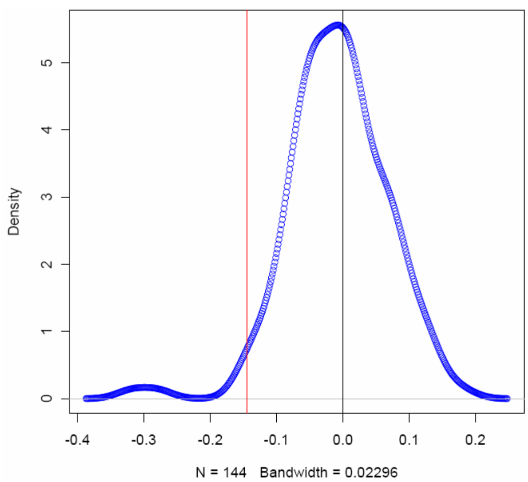
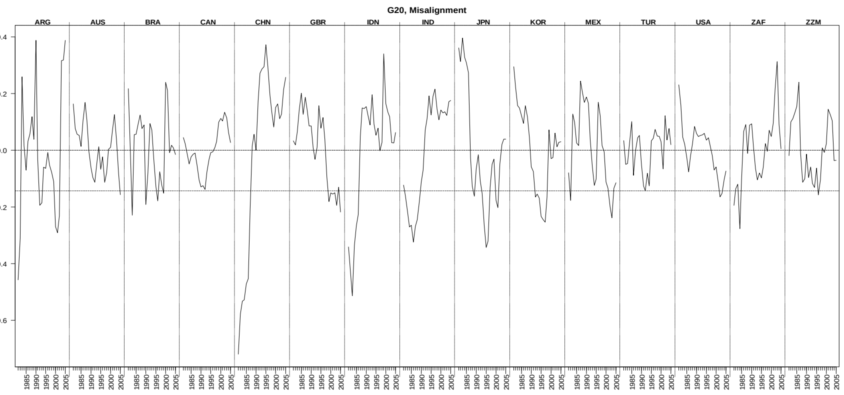


Figure 3: **G-20**


G20, Misalignment
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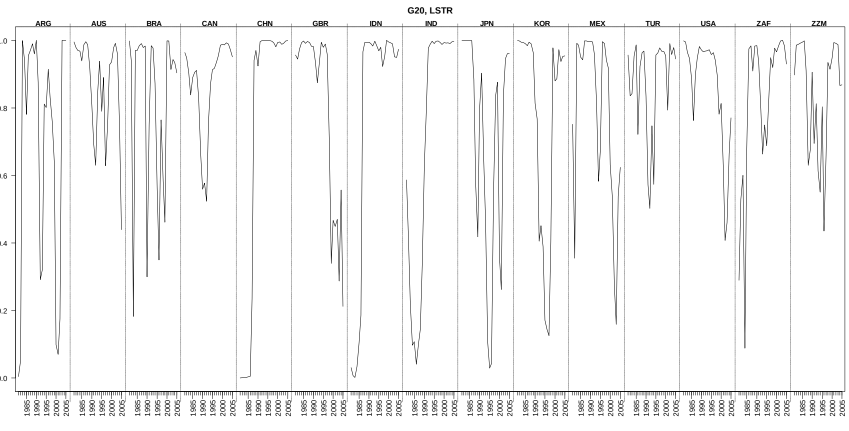
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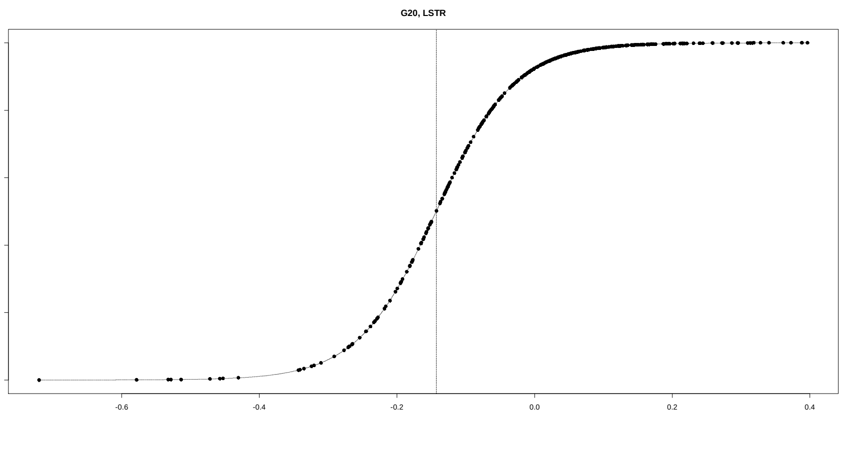
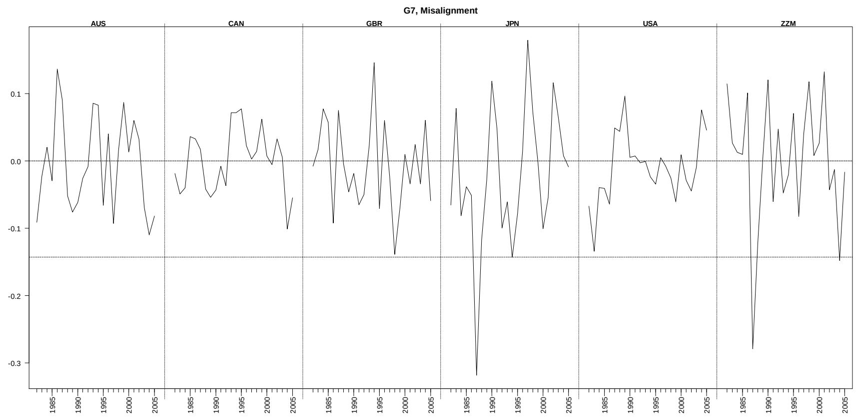

G20, LSTR

Figure 4: **G-7**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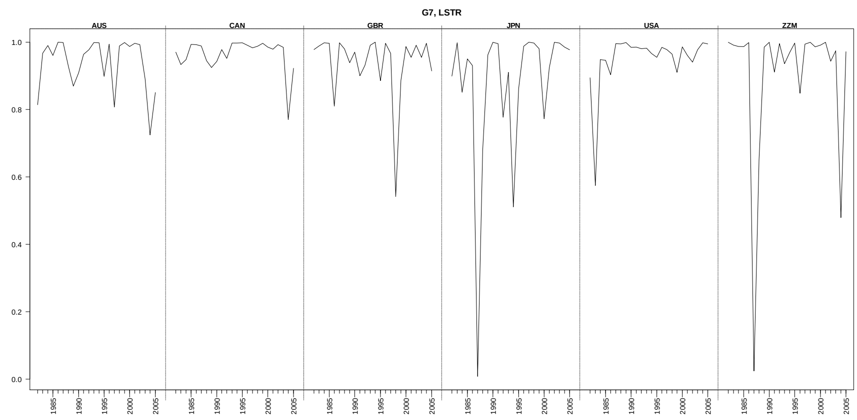
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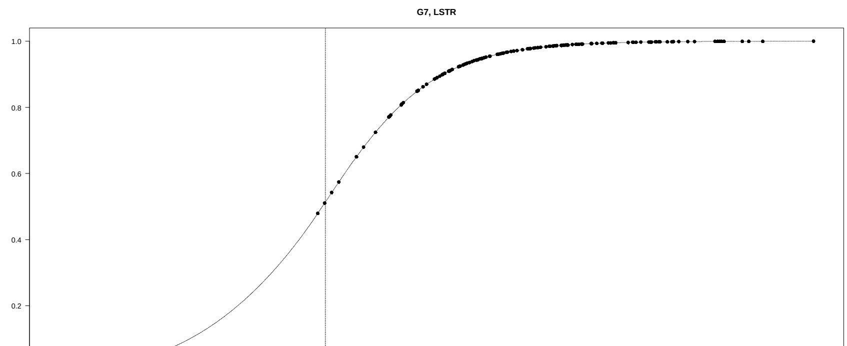
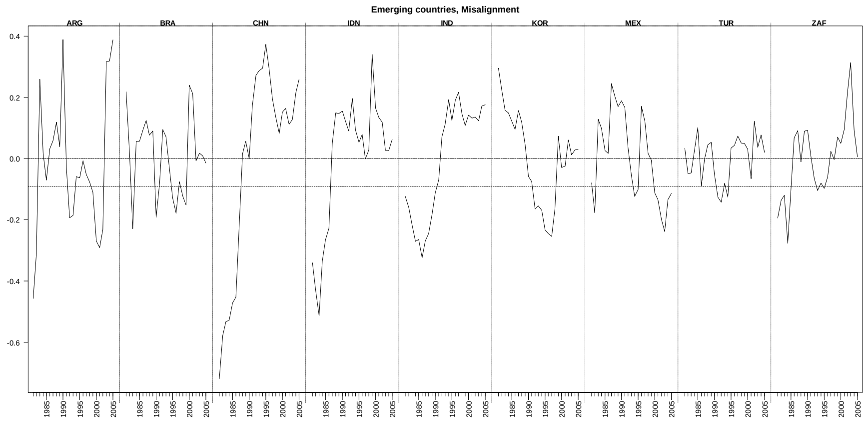
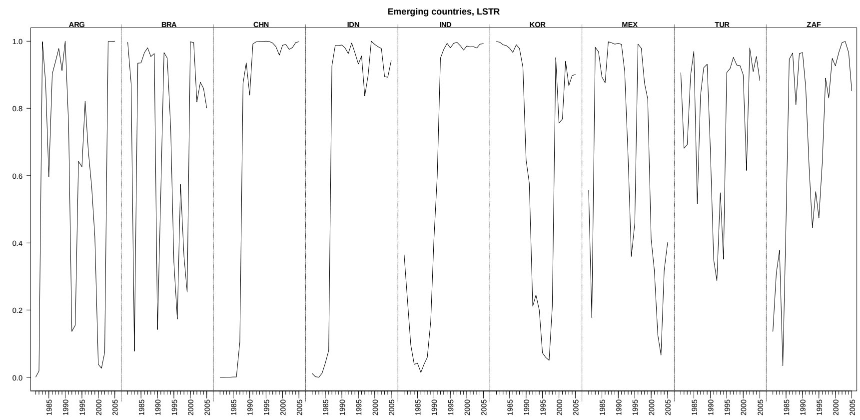

G7, LSTR

Figure 5: **Emerging economies (Non G-7)**





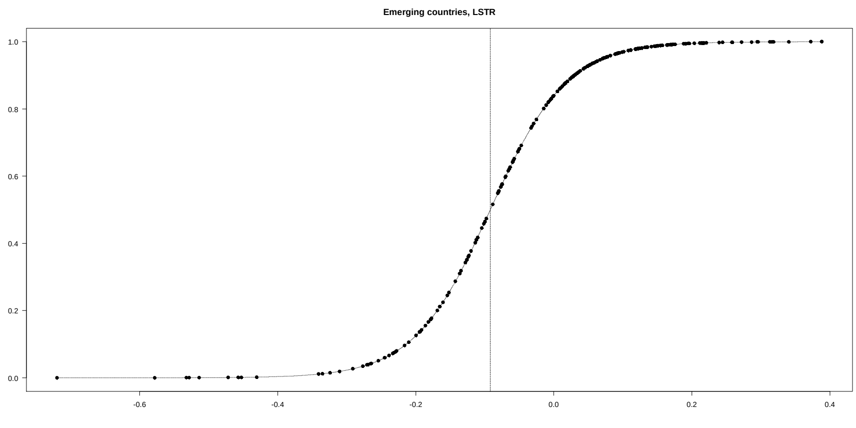
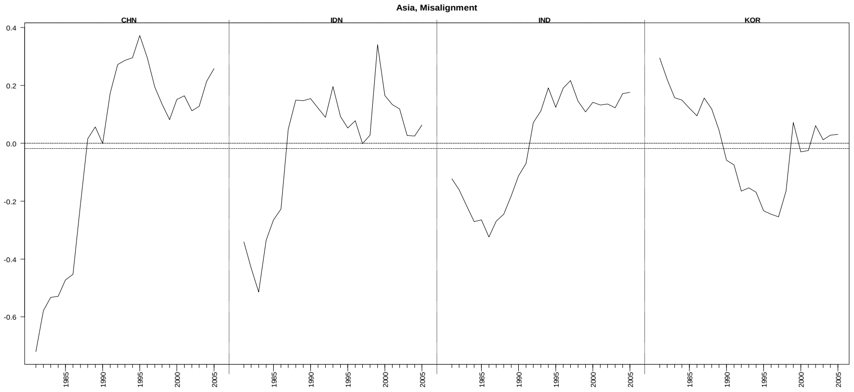
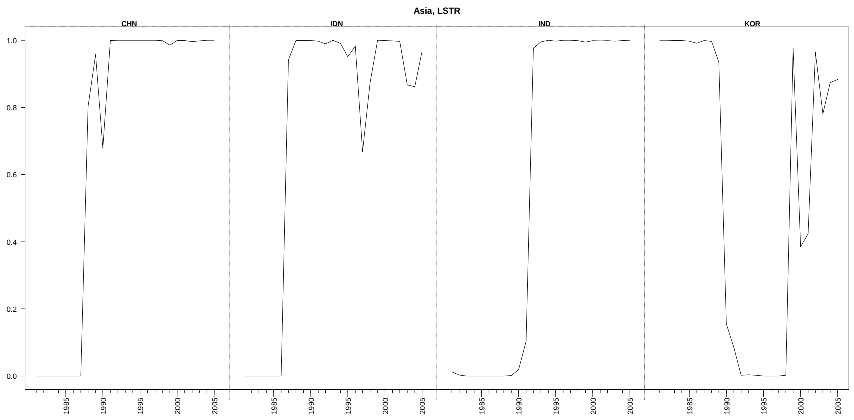


Figure 6: **Developing Asia**





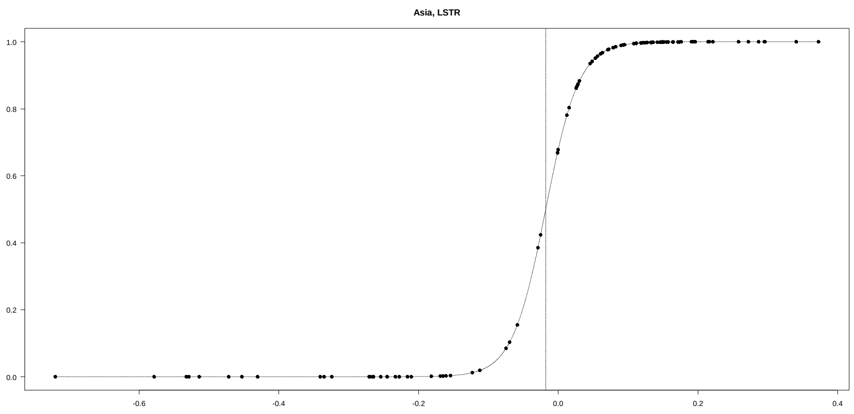


Figure 7: **Countries having overcome a recent financial crisis**