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Long Run Determinants of Real Exchange Rates in Latin America

Jorge Carrera[1] and Romain Restout[2]

Abstract

This paper investigates the long run behavior of real exchange rates in nineteen countries of Latin America over the period 1970 - 2006. Our data does not support the Purchasing Power Parity (PPP) hypothesis, implying that real shocks tend to have permanent effects on Latin America's real exchange rates. By exploiting the advantage of non stationary panel econometrics, we are able to determinate factors that drive real exchanges rate in the long run : the Balassa-Samuelson effect, government spending, the terms of trade, the openness degree, foreign capital flows and the *de facto* nominal exchange regime. The latter effect has policy implications since we find that a fixed regime tends to appreciate the real exchange rate. This finding shows the non neutrality of exchange rate regime regarding its effects on real exchange rates. We also run estimations for country subgroups (South America *versus* Caribbean and Central America). Regional results highlight that several real exchange rates determinants are specific to one geographic zone. Finally, we compute equilibrium real exchange rate estimations. Two main results are derived from the investigation of misalignments, [*i*] eight real exchange rates are quite close to their equilibrium level in 2006, and [*ii*] our model shows that a part of currencies crises that arose in Latin America was preceded by a real exchange rate overvaluation.

Keywords: Equilibrium Real Exchange Rate; Panel Unit Roots, Panel Cointegration.
J.E.L. Classification: F31, C23.

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The views expressed in this paper are those of the authors, and do not necessarily reflect those of the Central Bank of Argentina.

1 Introduction

Since 1970 until 2006, the Real Effective Exchange Rate (REER, hereafter)in Latin American countries shows dynamics characterized by important swings and a strong volatility. Table 1 illustrates the average level for the three last decades of nineteen countries form South and Central America. The REER level of the 90´ s for the whole region was more depreciated than in the previous decades with a volatility lower than in the 80´s. Exchange rate stabilization plans that were implemented in South America in the early 90's and the loosening of external gap with the arrivals of massive capital inflows possibly contributed to reduce REER's volatility. Except in the cases of Argentina and Peru, volatility reduction was associated with a REER depreciation. With the adoption of exchange rate-based anchors, a large part of Central American countries succeeded also to break down REER volatility during the 90's. In Costa Rica, Guatemala and Panama, REER's volatility is extremely low compared to others countries.

Table 1: ***Real Effective Exchange Rates in Latin America (1970- 2006)***

	Volatility				Average level (2000 = 100)			
	70 - 80	80 - 90	90 - 00	01 - 06	70 - 80	80 - 90	90 - 00	01 - 06
Argentina	0.57	0.71	0.25	0.46	66.81	42.28	98.15	53.69
Bolivia	0.23	0.33	0.13	0.11	166.10	181.00	107.44	94.43
Brazil	0.18	0.35	0.57	0.20	87.81	40.30	85.53	95.32
Chile	0.72	0.35	0.07	0.08	163.71	145.01	103.10	105.32
Colombia	0.07	0.27	0.18	0.12	136.60	127.50	106.97	100.85
Ecuador	0.11	0.35	0.20	0.04	226.65	193.35	116.25	148.32
Paraguay	0.17	0.38	0.21	0.08	196.30	305.52	118.99	93.77
Peru	0.25	0.32	0.21	0.04	49.36	36.05	95.85	101.18
Uruguay	0.19	0.24	0.15	0.10	74.73	93.03	92.48	78.95
Venezuela	0.04	0.38	0.33	0.17	100.17	89.10	63.18	76.95
South America	**0.25**	**0.37**	**0.23**	**0.14**	**126.83**	**125.31**	**98.79**	**94.75**
Costa Rica	0.07	0.30	0.05	0.03	193.10	102.44	98.93	97.39
Dominican Rep.	0.06	0.38	0.12	0.19	146.43	128.05	98.38	92.96
El Salvador	0.05	0.21	0.15	0.00	44.49	69.12	89.27	101.41
Guatemala	0.04	0.27	0.10	0.10	147.03	133.68	100.59	116.68
Honduras	0.06	0.07	0.13	0.02	127.89	150.03	82.78	101.66
Jamaica	0.14	0.26	0.27	0.07	139.40	101.54	83.35	95.51
Mexico	0.17	0.38	0.19	0.07	111.94	67.43	83.06	97.93
Panama	0.11	0.06	0.06	0.02	145.20	142.05	105.13	95.88
Trinidad & Tobago	0.06	0.16	0.12	0.03	96.60	120.74	104.06	108.71
Central America	**0.08**	**0.23**	**0.13**	**0.06**	**128.01**	**112.79**	**93.95**	**100.90**
Latin America	**0.17**	**0.30**	**0.18**	**0.10**	**127.39**	**119.38**	**96.50**	**96.67**

Notes : Author's calculation based on IMF's data.

Recently, it has been stressed that REER's volatility and misalignments (compared to an equilibrium level, see section 2 below) affect economic growth in developing countries. More exactly, recurrent and large misalignments are linked to lower growth rates and current account deficits in the long run and very frequently with currency and financial crisis. REER's movements affect also internal production and consumption allocations between non traded and traded goods. Aguirre and Calderón (2005) found a non linear relation between economic performance and misalignments in a panel of sixty countries : large overvaluations and undervaluations hurt growth, whereas small undervaluations can boost growth. Considering the importance of the link between economic growth and REER misalignments for developing countries, the question of what are the determinants that drive the real exchange rate in the long run is still a crucial issue. In the case of Latin America area, this question is reinforced by its inherent external vulnerabilities. First, primary commodities trade still represents an important part of exports (67 % [1]). Since commodities prices are invoiced in U.S. dollars and mostly exogenous for Latin American countries, real exchange rates fluctuations can lead to serious contractions of exports and trade balance. Second, foreign capital inflows are an important but volatile element of Latin America growth[2]. In this context, real exchange rate stability contributes to provide a stable macroeconomic environment that in turn encourages capital inflows. And, finally external debt remains higher in many Latin American

[1]Data cover exports of Argentina, Brazil, Chile, Colombia, Costa Rica, Ecuador, El Salvador, Guatemala, Honduras, Jamaica, Mexico, Paraguay, Peru, Trinidad and Tobago, Uruguay and Venezuela (source : CEPAL, period 1990-2002).

[2]See IMF (2005) for a discussion of external financing flows impact on growth.

countries, a large portion of these debts are indexed to exchange rate and/or denominated in foreign currencies, leaving stocks of debt vulnerable to REER's movements.

This paper have two main aims, first to investigate economic variables that influence the long run real exchange rates' paths. We consider a panel of nineteen countries of Latin and Central America. In order to determine the relevant real exchange rate's fundamentals, we use the more recent panel unit roots and cointegration techniques. These fundamentals include the Balassa-Samuelson effect, government spending, terms of trade, the degree of openness and foreign capital flows. This paper adds as a determinant, the *de facto* nominal exchange rate regime which can have a potentially impact on the real exchange rate. We follow the classification developed by Coudert and Dubert (2004) to identify the *de facto* exchange regime. Second, based on cointegration results, we estimate equilibrium path for the nineteen real exchange rates and then compute the degree of misalignment between the equilibrium and the observed real exchange rate. Next, these misalignments are used to construct an early warning indicator of currency crisis.

The rest of the paper is organized as follows. Section 2 briefly reviews the literature on fundamentals of real exchange rate. In section 3, we present the econometric methodology used to estimate the relationship between the real exchange rate and its long run fundamentals. Section 4 deals with empirical estimations and analysis of the results. It describes the data used in the regressions, presents cointegration regression results and computes real exchange rate misalignments. Finally, section 5 concludes.

2 The Determinants of the Real Exchange Rate

In this work, we define the real exchange rate (q) as the relative price of the non traded to the traded goods, *i.e.*:

$$q = \frac{p_N}{p_T}, \tag{1}$$

where p_N and p_T are respectively the price of non traded and traded goods[3]. This definition is called the internal real exchange rate and is appropriate for developing countries whose exports are predominantly primary products subject to the law of one price, *i.e.* $p_T^* = e\, p_T$, where e is the nominal exchange rate and p_T^* is the foreign price of traded goods. As noted by Edwards (1989), this definition provides a consistent index of the country's tradable sector competitiveness and also guides the resource allocation since an increase in q causes a shifting of resources away from the traded to the non traded sector.

According to definition (1), the real exchange rate's path is only driven by the dynamic of the internal relative price of non traded goods. This result depends upon two hypotheses [*i*] the law of one price is valid for the traded goods and [*ii*] the country is too small to have an influence on foreign partners' relative prices. Formally, introducing the common definition of the real exchange rate (in logarithm) as $\log q = \log e + \log p - \log p^*$, where p and p^* are respectively the national and foreign total price indices, and assuming that $\log p$ and $\log p^*$ can be split into traded and non traded prices as $\log p = (1-\alpha)\log p_T + \alpha \log p_N$, and $\log p^* = (1-\alpha)\log p_T^* + \alpha \log p_N^*$, with $\log p_N^*$ the price of foreign non traded goods and α being the share of the non traded sector in GDP at home and abroad, the real exchange rate $\log q$ can be decomposed in two components :

$$\log q = (\log e + \log p_T - \log p_T^*) + \alpha\left[(\log p_N - \log p_T) - (\log p_N^* - \log p_T^*)\right] \tag{2}$$

Under the hypotheses [*i*] and [*ii*], the first term in (2) vanishes and $(\log p_N^* - \log p_T^*)$ is given[4]. Thus, the real exchange rate varies only with the domestic relative price of non traded goods, $(\log p_N - \log p_T)$.

Several theoretical models of real exchange rate determination were implemented recently for developing countries, including Edwards (1989, 1994), Elbadawi (1994), Obstfeld and Rogoff (1996), Montiel (1999) and Lane and Milesi-Ferretti (2004). Except Edwards (1994) and Elbadawi (1994) ones, all these models are based on strong microfundations with a representative agent which maximizes his intertemporal utility in a two sectors framework (a traded and a non traded ones). However, these models differ from underlying hypotheses on which they are based. Whereas the model of Edwards (1989) assumes perfect competition and constant returns to scale in both sectors, Obstfeld and Rogoff (1996) consider

[3] According to (1) a real appreciation (depreciation) is reflected by an increase (decrease) in q. See Edwards (1989), Williamson (1994) and Edwards and Savastano (1999) for discussions about theoretical foundations of this real exchange rate concept.

[4] Here we impose a common value for α in both countries. If $\alpha \neq \alpha^*$, equation (2) has the general form :

$$\log q = (\log e + \log p_T - \log p_T^*) + \alpha\,(\log p_N - \log p_T) - \alpha^*\,(\log p_N^* - \log p_T^*)\,.$$

Under assumptions [*i*] and [*ii*], the real exchange rate still reduces to $(\log p_N - \log p_T)$.

that the non traded goods are produced under a monopolistic competition structure and with a production function characterized by diminishing returns to scale[5]. Except Montiel's model, the others ones assume a perfect world capital market, the domestic interest rate reaches the exogenously given world interest rate. However, world markets capital are used to claim risk premium for developing countries. Montiel incorporates this idea in his model by assuming that the economy faces an upward-sloping schedule relating the domestic interest rate to a country-specific risk premium that increases with the national stock of debt.

Despite these theoretical differences, previous models are similar in several points. First, they are based on the *single equation approach* that allows to derive a reduced form for the long run equilibrium real exchange rate. According to Edwards' (1989) definition, the equilibrium real exchange rate is defined as the relative price of non traded to traded goods that is compatible with the simultaneous attainment of the internal and external equilibrium. Internal equilibrium is achieved when non traded goods and labor markets clear. External equilibrium is related to the intertemporal budget constraint, *i.e.* the economy is intertemporally solvent.

Second, the equilibrium real exchange rate is long run driven by a set of foreign and domestic real variables, called *fundamentals* by Edwards and Savastano (1999). Usually, theoretical models link the equilibrium real exchange rate with government spending, sectoral productivity differentials (the Balassa-Samuelson effect), terms of trade, country's openness to international trade, foreign capital inflows and net foreign assets among other variables. In the short run, both real and nominal variables affect the equilibrium real exchange rate (see Edwards, 1989), but it responds only to fundamentals variations in the long run. Thus, all these theoretical approaches reject models based on Purchasing Power Parity (PPP), the equilibrium real exchange rate is not an immutable value but it varies through time. Real exchange rate movements do not necessarily reflect disequilibrium situations, fundamentals variations generate endogenously real exchange rate fluctuations. Moreover, it assumes that the actual real exchange rate is mean reverting, *i.e.* it returns rapidly to its equilibrium value in the long run. Here, misalignments that can arise due to an inadequate macroeconomic policies for example, are only temporary.

Finally, the equilibrium real exchange rate can be calculated by an appropriate use of econometric techniques. In empirical terms, a baseline equation for the long run equilibrium real exchange rate can be estimated, that is :

$$q_t = \beta' X_t + \varepsilon_t, \tag{3}$$

where X_t are the *fundamentals*, β the vector of long run parameters and ε_t an error term. To construct the equilibrium real exchange rate path, noted $\bar{q}_t$, Clark and MacDonald (1999) suggest to use sustainable values or the permanent component of fundamentals. By estimating (3) with a consistent econometric method, we can obtain the equilibrium path for real exchange rate :

$$\bar{q}_t = \hat{\beta}' X_t^p, \tag{4}$$

where the vector $\hat{\beta}$ contains efficient estimators of β and X_t^p is the permanent component of the *fundamentals* which can be computed from time series decomposition techniques (Hodrick-Prescott filter, Beveridge-Nelson decomposition or Gonzalo-Granger methodology). From (4), the real exchange rate misalignment, q_t^d, is computed as the deviation of the observed real exchange rate, q_t, from its equilibrium level, that is $q_t^d \equiv q_t - \bar{q}_t = q_t - \hat{\beta}' X_t^p$.

In this paper, we follow the *single equation approach* in order to determine *fundamentals* that drive evolutions of long run real exchange rates in the Latin America region. Since the literature on theoretical models of real exchange rate determination is abundant, we mainly focus here on the empirical application of these models, *i.e.* we estimate a long run relationship between the real exchange rate and a limited number of *fundamentals* which have a theoretical influence on long run real exchange rate according to models presented below. The existent empirical literature on equilibrium real exchange rates in Latin America suggests that only a restricted number of real variables seems to influence the real exchange rate in the long run[6]. These variables include a Balassa-Samuelson effect, government spending, terms of trade, degree of openness and capital inflows[7]. By adopting this empirical approach, our goal is not to valid or to reject any particular real exchange rate model quoted previously but is to study dynamics of real exchange rates with the help of non stationary panel data econometrics methods. Moreover, as shown by Edwards and Savastano (1999), the choice and the number of *fundamentals* included in the real exchange rate equation is model dependent. For example, Lane and Milesi-Ferretti's (2004) model do not include

[5] Montiel (1999) and Lane and Milesi-Ferretti (2004) also assume diminishing marginal returns technologies but keep the assumption of perfect competition in both sectors.

[6] See Edwards (1989), Gay and Pellegrini (2003), Alberola (2003), Escudé and Garegnagni (2005) for recent empirical investigations.

[7] The choice of *fundamentals* is obviously constraints by data availability.

government spending as a determinant for the real exchange rate, whereas the Balassa-Samuelson effect is not explicitly derived from models of Edwards (1989,1994) and Elbadawi (1994). Thus, by adopting an empirical approach, we do not restrict our analysis to be consistent with the reduced form of a particular model, and we are allow to include the most important fundamentals in the long run real exchange rate equation.

2.1 Productivity Effect

The productivity effect refers to the Balassa-Samuelson model (Balassa, 1964 and Samuelson, 1964). According to this hypothesis, the relative price of non traded goods, q, is determined by the traded-non traded productivity differential. The explanation is the following. Consider a two sectors economy (traded-non traded) where wages are the same in both sectors and are linked to productivity in the open sector. Assume that the law of one price holds in the traded sector and that the interest rate is entirely exogenous. When productivity improves faster in the traded sector than in the non traded sector, wages are expected to rise in the entire economy. In the non traded sector where the wage increase is unmatched by an equivalent productivity improvement, the price p_N is expected to rise. This in turn leads to an increase in the relative price of non traded goods q, i.e. an appreciation of the home country's real exchange rate. Assuming a Cobb Douglas production function with constant returns to scale in both sectors ($Y_T = A_T K_T^{1-\alpha} L_T^{\alpha}$ and $Y_N = A_N K_N^{1-\beta} L_N^{\beta}$, where A_i, L_i and K_i represent respectively total factor productivity, labor and capital in sector i, $i = T, N$), the formal expression of the Balassa-Samuelson effect is :

$$\hat{q} \equiv \hat{p}_N - \hat{p}_T = \frac{\beta}{\alpha}\hat{A}_T - \hat{A}_N, \tag{5}$$

where a *hat* above a variable denotes growth rate[8]. Thus, according to equation (5), the real exchange rate depends entirely on productivity differentials. Moreover, the Balassa-Samuelson effect can be also interpreted as the effect of the economic development on real exchange rate, *i.e.* fast growing countries tend to experience a real appreciation of their exchange rate.

2.2 Government Spending

A crucial feature of the Balassa-Samuelson model is that the real exchange rate is fully determined by the supply side of the economy, demand factors do not matter. As noted by Froot and Rogoff (1995) and De Gregorio et *al.* (1994), this result depends on assumptions of the Balassa-Samuelson model. Demand factors can have an effect on the relative price of non traded goods if one of the following assumptions is relaxed : perfect competition in goods markets, factors move freely between the two sectors of production, capital is mobile internationally, law of one price for traded goods and constant returns to scale in the two sectors. By introducing monopolistic competition in the non traded sector in Lane and Milesi-Ferretti's (2004) model, Aguirre and Calderón (2005) allow for demand factors to influence the real exchange in the long run. The impact of public demand on real exchange rate is traditionally linked to the hypothesis that government spending generally falls disproportionately on non traded goods. An increase in government spending exercises an upward pressure on the relative price of non traded goods and thus appreciates the real exchange rate.

2.3 Terms of Trade

All theoretical models quoted previously stress the importance of terms of trade disturbances as a potential source of real exchange rate fluctuations. However, the impact of a terms of trade worsening on the real exchange rate is theoretically undefined because two contrary effects play in opposite way. First, a deterioration of terms of trade induces a negative income effect (decline in the domestic purchasing power) and results in a reduction in the private demand for non traded goods and then to a real depreciation of the exchange rate. On the other hand, a substitution effect makes the consumption of imported goods relatively more expensive. As a result there is a shift of demand in favor of the non traded goods, the reestablishment of the equilibrium in the non traded market is provided by an increase of the real exchange rate. The total effect of a terms of trade deterioration on real exchange rate depends on the strength of the income and substitution effects. However, recent empirical studies found that the income effect is predominant, hence, terms of trade improvements are associated with real appreciation in the long run.

[8] The formal derivation of equation (5) can be found in Froot and Rogoff (1994) and De Gregorio et *al.* (1994).

2.4 Openness

The degree of openness influences the real exchange rate through two main channels. The first one stands that trade-liberalizing reforms tend to depreciate the long run real exchange rate. An increase in the openness variable, such as a reduction in tariff, leads to a decline in the domestic price of imported goods. This in turn entails an excess demand for imported goods and reduces domestic demand for the non traded good[9]. As a result, the real exchange rate depreciates to restore the equilibrium in the non traded market. The second theoretical influence channel between real exchange rate and degree of openness has been emphasized by Obstfeld and Rogoff (2000) and Hau (2002). According to their models' predictions, real exchange rate volatility is negatively related to economic openness. Since the non traded sector is the locus of the monopoly, non traded goods increase the degree of aggregate price rigidity, whereas traded goods permits the convergence of the domestic price indice. Following a real shock, larger real exchange rate changes are needed for a more closed economy to restore equilibrium on domestic markets. As noted by Hau (2002), more open countries behave more like flexible prices economies with smaller real exchange rate fluctuations since more imported goods provide a channel for quick adjustment of the national price indice. Using a panel of forty eight countries (including eight countries from Latin America), Hau (2002) provided evidence of the negative relationship between real exchange rate volatility and trade openness.

2.5 Capital Flows and Net Foreign Assets

According to Corden (1994), the impact of foreign capital flows on the real exchange rate refers to the *real exchange rate problem*, that is capital inflows are associated with real exchange rate appreciation in the long run. The intuition for this effect is straightforward. A foreign capital surge affects the economy by raising the domestic absorption which leads to an increase in consumption demand for both traded and non traded goods. On non traded goods market, this excess demand has to matched to a proportional increase of the non traded supply in order to ensure market equilibrium. This in turn leads to a rise of the price of non traded goods, p_N. The traded consumption increase will cause the trade balance to deteriorate without any effects on p_T since it is entirely determined by the law of one price. According to definition (1), the change in p_N, following the foreign capital inflows, entails an appreciation of the real exchange rate. Athukorala and Rajapatirana (2003) showed that the magnitude of real exchange rate appreciation depends on the composition of foreign inflows. Foreign direct investment (FDI, hereafter) tends to be more concentraded in the traded sector, and thus the appreciation resulting from a FDI entry is lower compared to others types of capital inflows. In contrast, portfolio inflows bring a greater real appreciation. Econometric results from Athukorala and Rajapatirana (2003) suggest that the composition of capital flows matters in determining their impact on the real exchange rate. Moreover, comparing the effect of FDI and portfolio on real exchange rates in Latin America and Asia, they can conclude that the degree of real appreciation following a portfolio inflows is stronger in Latin America countries than in Asia.

The relationship between net foreign asset (NFA, hereafter) positions and the real exchange rate has been analyzed by several theoretical models (Obstfeld and Rogoff, 1995, and Lane and Milesi-Ferretti, 2004) which predict that debtor (creditor) countries having more depreciated (appreciated) real exchange rates (*the transfer problem*). Indeed, countries with net foreign liabilities need to run trade surplus to finance interests and dividends payments. Similarly, countries with positive NFA must have trade deficits. Obstfeld and Rogoff (1995) claim that the *transfer problem* can also operate through the impact of wealth effects on labor supply. A deterioration of the NFA position reduces national wealth. To prevent a large drop in consumption, households rise their labor supply, thus increasing the non traded goods supply. Since the non traded goods market is in equilibrium each period, the price p_N has to fall, *i.e.* the real exchange rate depreciates.

2.6 Nominal Exchange Rate Regime

The effect of exchange rate regime on real exchange rate is a controversial issue both in theoretical and empirical terms. In traditional literature the dynamics of potential misalignments depends on the level of price stickiness and financial openness. From the point of view of some theories like the equilibrium ones the regime is neutral regarding the level or volatility of real exchange rate for Baxter and Stockman (1989), Flood and Rose (1995), and, Obstfeld and Rogoff (2000). Recent equilibrium literature based on nominal rigidities or market imperfections tends to find non neutrality between these variables but is highly dependent of the type of imperfection that is highlighted.

[9]Unlike in the case of terms of trade shocks, the income effect is absent following a fall in the tariff rate. Indeed, the tariff reduction has to be financed by an increase in taxes (we assume here that the government budget is balanced in every period), which in turn offsets the initial wealth effect induced by the reduction of the imported goods price.

From a different perspective the literature on exchange rate-based stabilizations applied to Latin America countries normally remarks the strong effects of fixation on real exchange rate levels. The stylized fact is the following: when a country suffered from high inflation, one of the most popular alternatives in the last three decades was to implement a fixation of (nominal) exchange rate, sometimes it happens after a strong devaluation. Because of the high inertial dynamic of inflation the changes in prices and wages tends to persist and so the fixed exchange rate regime is associated with a persistent appreciation of real exchange rate. In several papers Calvo has remarked the credibility problem associated with this appreciation following to an exchange regime stabilization plan (Calvo and Vegh, 1993). Normally non credible exchange regime stabilizations are followed by a boom in the consumption of non traded goods, so we expect that rigid exchange rate regimes lead to real exchange rate appreciation.

Moreover, according to IMF (2005) exchange rates stabilization plans introduced in Latin America countries in the early 90's encouraged an external capital inflows surge that appreciated the real exchange rates. By adopting a fixed regime, a country provides, at least in the short run, a perceived stable environment which in turn attracts foreign capital inflows. To maintain the desired parity of the nominal exchange rate the central bank is forced to purchase these excessive flows, leading to an increase of the domestic money base if the central bank does not run sterilization operations. This expansion of the domestic credit supply has lead in the Latin America zone to consumption of non traded goods boom which in turn has appreciated the real exchange rate.

Calvo and Reinhart (2002) show that there are important distortions between *de jure* exchange regime (the official one that is reported to IMF) and the *de facto* regime which reflects the policy pursuing by the country. Table 2 shows the distribution percentage of fixed *de facto* exchange rate regime in the total sample period. It is possible to perceive that in the seventies 76 % of the time was on average under a fixed exchange rate regime. This percentage decreases in the following decades reaching 58 % in the present century. For example in Central America there was in the 90´s an intense increase in the trade and financial relationships with United States and the dollarization of the economies but the regime is more flexible than in the previous decades.

Table 2: ***Exchange Rate Regime Distribution (1970 -2006)***

	Percentage of fix regimes				
	70 - 80	80 - 90	90 - 00	01 - 06	70 - 06
Argentina	0.40	0.10	0.82	0.33	0.43
Bolivia	0.70	0.50	1.00	1.00	0.78
Brazil	0.60	0.40	0.27	0.00	0.35
Chile	0.20	0.60	0.64	0.17	0.43
Colombia	1.00	1.00	0.45	0.33	0.73
Ecuador	0.80	0.40	0.36	1.00	0.59
Paraguay	1.00	0.60	0.73	0.00	0.65
Peru	0.60	0.30	0.55	0.50	0.49
Uruguay	0.50	0.60	0.91	0.17	0.59
Venezuela	1.00	0.70	0.36	0.33	0.62
South America	**0.68**	**0.54**	**0.61**	**0.38**	**0.57**
Costa Rica	0.90	0.60	1.00	1.00	0.86
Dominican Rep.	1.00	0.60	0.55	0.33	0.65
El Salvador	1.00	0.90	0.82	1.00	0.92
Guatemala	1.00	0.70	0.55	0.67	0.73
Honduras	1.00	1.00	0.55	1.00	0.86
Jamaica	0.50	0.60	0.36	0.83	0.54
Mexico	0.80	0.50	0.45	0.33	0.54
Panama	1.00	1.00	1.00	1.00	1.00
Trinidad & Tobago	0.40	0.80	0.73	1.00	0.70
Central America	**0.84**	**0.74**	**0.67**	**0.80**	**0.76**
Latin America	**0.76**	**0.63**	**0.64**	**0.58**	**0.66**

3 A Review of Panel Econometric Methods

In this section, we present panel unit root and cointegration techniques involved in our analysis. Given our relatively short time span ($T = 37$), examining the long run behavior of real exchange rates by using non stationary panel econometrics, instead of individual time series, yields substantial benefits.

First, by allowing data to be pooled in the cross section dimension (N), panel unit root and cointegration tests gain power and outperform their conventional time series counterparts[10] (i.e. Dickey-Fuller or Engle-Granger tests). Moreover, panel data provide efficient estimators for cointegration vectors[11] which are superconsistent and converge at rate $T\sqrt{N}$ while in time series dimension, the convergence is slower at a rate T. So, even in the case of relatively small time and cross section dimensions, these estimators are extremely precise.

Second, recent panel unit root and cointegration tests (see among others Im, Pesaran and Shin, 2003, and Pedroni, 1999) take account of the heterogeneity across different members of the panel. This allows us to test the presence of unit root and cointegration relation in the panel while permitting the short run dynamics, error variances and fixed effects to be heterogeneous among individual members. As noted by Alberola *et al.* (2003), this flexibility is useful for studies that only focus on the long run behavior of data since the short run dynamics and the long run equilibrium are likely to be different across individuals.

Finally, by adding the cross section dimension, non stationary panel permits to limit the time span of data, and so is able to reduce the presence of structural breaks since long time series include, potentially, serious structural shifts. As shown first by Perron (1989), conventional time series unit root and cointegration tests tend to be biased in the presence of structural breaks.

3.1 Panel Unit Root Tests

Because we mainly focus on the long run determinants of real exchange rates, we first test for unit root in panel data. Here, we apply three unit root tests, the W_{tbar} test proposed by Im, Pesaran and Shin (2003, hereafter, IPS), the Fisher type test suggested by Maddala and Wu (1999, hereafter, MW) and the Hadri (1999) LM test. Whereas the latter takes as the null hypothesis of the stationarity against the alternative of a unit root in panel data, IPS and MW considered the non stationarity (i.e. presence of a unit root) as the null hypothesis. All these three tests are designed for cross sectionally independent panels, i.e. there is no cointegration between pairs or groups of individuals in the across section dimension. This assumption of independence across individuals is quite strong but essential in order to apply the Lindberg-Levy central limit theorem that permits to derive limiting distributions of tests (Baltagi and Kao, 2000)[12].

Im *et al.* (2003) proposed a test that allows for residual correlation, and heterogeneity of the autoregressive root and error variances across individual members of the panel. IPS is based on the use of Augmented Dickey-Fuller (ADF) test to each individual series. If we consider a sample of N cross sections observed over T time periods, the following ADF regression is estimated for each individual $i = 1, \ldots, N$ of the panel :

$$\triangle y_{i,t} = \gamma z_{i,t} + \rho_i y_{i,t-1} + \sum_{j=1}^{p_i} \theta_{i,j} \triangle y_{i,t-j} + \varepsilon_{i,t}, \tag{6}$$

where $z_{i,t}$ is the deterministic component (fixed effects α_i and/or individual time effect $\beta_i\, t$), $\varepsilon_{i,t}$ are assumed to be identically, independently distributed ($i.i.d.$) across i and t with $\mathrm{E}(\varepsilon_{i,t}) = 0$, $\mathrm{E}(\varepsilon_{i,t}^2) = \sigma_i^2 < \infty$, and $\mathrm{E}(\varepsilon_{i,t}, \varepsilon_{j,t}) = 0$ for all $i \neq j$. Notice that the lag length p_i is permitted to vary across individual members of the panel.

The null and alternative hypotheses are defined as :

$$H_0 : \quad \rho_i = 0, \quad \forall i = 1, \ldots, N,$$

against

$$H_a : \quad \rho_i < 0, \quad \text{for at least one } i.$$

Thus, under the alternative hypothesis IPS allow for ρ_i to be individual specific, and in this sense, is more general than the homogeneous alternative (i.e. $\rho_i = \rho < 0$ for all i) developed by Quah (1994) and Levin and Lin (1992). The IPS test averages the individual ADF t-statistics ($t_{i,T}$) that are obtained from estimating (6) for each i, that is $\bar{t}_{NT} = N^{-1} \sum_{i=1}^{N} t_{i,T}$. Im *et al.* (2003) proposed the standardized $\bar{t}_{NT}$ statistic :

$$W_{tbar} = \frac{\sqrt{N}\left(\bar{t}_{NT} - N^{-1} \sum_{i=1}^{N} \mathrm{E}\left[t_{i,T} | \rho_i = 0\right]\right)}{\sqrt{N^{-1} \sum_{i=1}^{N} \mathrm{Var}\left[t_{i,T} | \rho_i = 0\right]}}, \tag{7}$$

[10]For example, with $T = 50$, the power of the standard Dickey-Fuller test is only 0.151. With $N = 10$, power of Levin and Lin (1993) and Im, Pesaran and Shin (2003) tests reaches 0.555 and 0.752, respectively.

[11]The various estimators include within and between Fully Modified OLS (FMOLS) and Dynamic OLS (DOLS) proposed by Pedroni (2000, 2001) and Kao and Chiang (2000).

[12]O'Connell (1998) showed that PPP tests which ignore to control for cross sectional dependence suffer from significant distortions.

where $\mathrm{E}[t_{i,T}|\rho_i = 0]$ and $\mathrm{Var}[t_{i,T}|\rho_i = 0]$ are respectively the mean and the variance of $t_{i,T}$[13]. Under the null hypothesis, the W_{tbar} statistic is asymptotically standard normal distributed.

Maddala and Wu (1999) proposed a Fisher type test based on the p-values from individual unit root statistics, that is :

$$P = -2\sum_{i=1}^{N} \ln(p_i), \tag{8}$$

where p_i denotes the p-value of the individual unit roots test (ADF or Phillips-Perron (1988) for example) applied to cross section i. Like IPS, MW permits heterogeneity of the autoregressive root ρ_i under the alternative H_a. Under the assumption of cross sectional independence, P is distributed as a chi-squared with $2N$ degrees of freedom.

Hadri (1999) proposed a residual base Lagrange Multiplier (LM) test of the null that the time series for each i is stationary around a deterministic trend against the alternative of a unit root in the panel data. Consider the following model :

$$y_{i,t} = z_{i,t}'\gamma + r_{i,t} + \varepsilon_{i,t}, \tag{9}$$

where $z_{i,t}$ is the deterministic component and $r_{i,t}$ a random walk process ($r_{i,t} = r_{i,t-1} + u_{i,t}$). The $\varepsilon_{i,t}$ and $u_{i,t}$ are independent and *i.i.d.* across i and t with $\mathrm{E}(\varepsilon_{i,t}) = 0$, $\mathrm{E}(\varepsilon_{i,t}^2) = \sigma_\varepsilon^2$, $\mathrm{E}(u_{i,t}) = 0$, $\mathrm{E}(u_{i,t}^2) = \sigma_u^2$. Using backward substitution, equation (9) can be written as :

$$y_{i,t} = z_{i,t}'\gamma + e_{i,t}, \tag{10}$$

where $e_{i,t} = \sum_{j=1}^{t} u_{i,j} + \varepsilon_{i,t}$. Let $\hat{e}_{i,t}$ be the residuals from the regression (10) and $\hat{\sigma}_e^2$ be the consistent estimator of the true variance σ_e^2 under H_0. Then, the LM statistic is :

$$LM = \frac{1}{\hat{\sigma}_e^2}\frac{1}{NT^2}\left(\sum_{i=1}^{N}\sum_{t=1}^{T} S_{i,t}^2\right),$$

where $S_{i,t}$ is the residual partial sum ($S_{i,t} = \sum_{j=1}^{t} \hat{e}_{i,j}$). Under the null hypothesis of stationarity, the statistic test :

$$Z_\mu = \frac{\sqrt{N}\left(LM - \mathrm{E}\left[\int_0^1 V(r)^2 dr\right]\right)}{\sqrt{\mathrm{Var}\left[\int_0^1 V(r)^2 dr\right]}}, \tag{11}$$

is exactly standard normal distributed, where $V(r)$ is a standard Brownian motion[14].

Maddala and Wu (1999) investigated the finite sample performance of panel unit root tests. Since IPS and Fisher tests have the same alternative hypothesis, they are directly comparable. The major conclusion of their study is that the Fisher test seems superior to the IPS (the Fisher test has smaller size distortions and comparable power). In an extensive study, Hlouskova and Wagner (2005) studied performance of seven panel unit root tests including IPS, Fisher and Hadri-LM ones. They find that the stationary LM test of Hadri (2000) performs very poorly and often leads to a rejection of the null hypothesis[15].

3.2 Tests of Panel Cointegration

Adding the cross section dimension in testing for cointegration should also offer the same advantages in terms of power that are present when detecting for unit root. Several authors have recently proposed alternative methodologies for testing cointegration in a panel data context [16]. Using a multi-equation framework, Larsson *et al.* (2001) presented a likelihood-based (LR) test for cointegration rank in heterogeneous panels based on the average of the individual rank trace statistics developed by Johansen (1995)[17]. Kao (1999) and Pedroni (1999, 2004) proposed residuals based test for panel cointegration. The tests proposed by Kao (1999) are ADF type tests similar to the classical approach adopted by Engle

[13] Simulated values of $\mathrm{E}[t_{i,T}|\rho_i = 0]$ and $\mathrm{Var}[t_{i,T}|\rho_i = 0]$ are provided by Im *et al.* (2003), table 3, page 66.

[14] The moments $\mathrm{E}\left[\int_0^1 V(r)^2 dr\right]$ and $\mathrm{Var}\left[\int_0^1 V(r)^2 dr\right]$ are derived exactly, whereas for IPS a simulation is needed.

[15] Hlouskova and Wagner restricted their investigation to the case of homogenous panels, so IPS and MW, that are specially adapted from heterogeneous panels, appear to be disadvantaged by the imposition of homogeneity.

[16] Like when testing for unit root, tests of panel cointegration assume that the individual processes are independent and *i.i.d.* cross sectionally.

[17] If the LR test allows for more than one cointegrating relations, the size of the test is severely distorted even if the panel has large cross sectional and time dimensions.

and Granger (1987). He developed five tests under the null hypothesis of no cointegration. All Kao's tests constrained the cointegration vector and short run dynamics to be homogeneous across the individual members of the panel. This assumption of homogeneity has been relaxed by Pedroni (1999, 2004) who developed tests that allow for considerable heterogeneity across individuals. Like the IPS unit root test, Pedroni's tests allow individual short run dynamics, individual fixed effects and also allow the cointegration vector to differ across members under the alternative hypothesis[18]. Pedroni considers the following cointegration model with k regressors for a panel :

$$y_{i,t} = \gamma z_{i,t} + \beta_{1,i}x_{1,i,t} + \beta_{2,i}x_{2,i,t} + \cdots + \beta_{k,i}x_{k,i,t} + e_{i,t}, \tag{12}$$

where $z_{i,t}$ is the deterministic component (fixed effects α_i and/or individual time effect $\delta_i\, t$) and $x_{i,t}$ are the k regressors which are assumed to be $I(1)$ (i.e. $x_{i,t} = x_{i,t-1} + u_{i,t}$) and not cointegrated with each other. Pedroni's approach focuses on testing for unit roots in panel estimates of $e_{i,t}$, that is :

$$\hat{e}_{i,t} = \rho_i \hat{e}_{i,t-1} + v_{i,t}, \tag{13}$$

where $v_{i,t}$ are assumed to be identically, independently distributed ($i.i.d.$) across i and t with $\mathrm{E}(v_{i,t}) = 0$, $\mathrm{E}(v_{i,t}^2) = \sigma_i^2 < \infty$, and $\mathrm{E}(v_{i,t}, v_{j,t}) = 0$ for all $i \neq j$. Pedroni (1999, 2004) considered seven tests (noted χ_{NT}) based on the residual from the regression (13). Four are based on pooling data along the within dimension (panel-ν, panel-rho, panel non parametric-t and panel parametric-t) and three are calculated pooling data along the between dimension of the panel (group-rho, group non parametric-t and group parametric-t). Using the within approach, the test of the null of no cointegration is $H_0 : \rho_i = 1$ for all i against the alternative $H_a : \rho_i = \rho < 1$ for all i. Thus, all within statistics presume a common value $\rho_i = \rho$, whereas the between estimators are less restrictive in that they allow for considerable heterogeneity since the alternative hypothesis is $H_a : \rho_i < 1$ for all i. The between statistics provide an additional source of heterogeneity since the autoregressive coefficients, ρ_i, are allowed to vary across individual members of the panel. Pedroni (1999, 2004) found that each of the seven within and between statistics are distributed under the standard normal distribution as :

$$\frac{\chi_{NT} - \mu\sqrt{N}}{\sqrt{\nu}} \quad \Rightarrow \; \mathrm{N}(0,1), \tag{14}$$

where χ_{NT} is the appropriately statistic and μ and ν are respectively the mean and the variance of χ_{NT}[19].

Pedroni (2004) explored finite sample performances of the seven statistics. He showed that in terms of power all the proposed statistics do fairly well for $T > 90$ and $N = 20$. In addition to be less restrictive, Pedoni's simulations showed that between statistics have lower small sample size distortions than within ones. Moreover, for small time span ($T < 20$), the between group parametric-t statistic is the most powerful. Given our relatively short time span ($T = 37$) and size adjusted power results found by Pedroni (2004), we will only consider the group parametric-t statistic when testing for cointegration.

3.3 Estimation of Panel Cointegration Models

In a cointegrated system, only under restrictive conditions, *i.e.* exogeneity of the regressors and homogeneity of the dynamics across members of the panel, the OLS estimator for the cointegrating vector is asymptotically consistent and has a standardized distribution. Otherwise, the OLS estimator is biased and its asymptotic distribution will be dependent on nuisance parameters associated with the dynamics of the underlying system (Pedroni, 2000). Like testing for unit roots and for cointegration, alternatives procedures are proposed to provide efficient cointegrating vector estimators and thus to infer in the cointegrated panel model. The recent various approaches include within and between estimators of the Fully Modified OLS (FMOLS) and Dynamic OLS (DOLS). FMOLS is a non-parametric approach to adjusting for the effects of endogenous regressors and serial correlation while DOLS estimator adds leads and lags of first differences regressors in the cointegrating equation to correct these issues.

Pedroni (2001) argues that between (or group-mean) estimators allow for greater flexibility in estimating cointegrating vectors, in the sense that group-mean estimators can be interpreted as the mean value of the individual cointegrating vectors. Pesaran and Smith (1995) found that, when the cointegrating vectors are heterogeneous across individuals, group-mean estimators provide consistent estimates of the sample mean of the heterogeneous cointegrating vectors, while within dimension estimators do not. Furthermore, group-mean estimators allow for heterogeneity when inferencing in the cointegrating vector. Within estimators test the null hypothesis $H_0 : \beta_i = \beta_0$ for all i against the alternative $H_a : \beta_i = \beta_a \neq \beta_0$

[18]Endogeneity of the regressors is also allowed by Pedroni's tests, which contrats with Kao's (1999) approach where homogeneity and exogeneity are imposed.

[19]See table 2, page 666 of Pedroni (1999) for values of μ and ν.

where β is the cointegrating vector and β_a is the same value for all i. Group-mean estimators are designed to test the null hypothesis $H_0 : \beta_i = \beta_0$ for all i against the alternative $H_a : \beta_i \neq \beta_0$, so that heterogeneity is allowed and all the individual β_i are not constrained to have a common β_a value. Finally, Pedroni (2000) investigated the finite sample of the two within FMOLS (residual-FMOLS and adjusted-FMOLS) and of the group-mean FMOLS. He found that the group-mean FMOLS suffers from much lower small sample size distortions than the within estimators.

The group-mean FMOLS estimator is based on the estimation of the following cointegrated system for a panel :

$$\begin{aligned} y_{i,t} &= \alpha_i + x'_{i,t}\beta + u_{i,t}, \\ x_{i,t} &= x_{i,t-1} + \varepsilon_{i,t}, \end{aligned} \tag{15}$$

where α_i are the fixed effects, β is a $k \times 1$ vector of the slope parameters, $x_{i,t}$ is a $k \times 1$ vector of integrated regressors, and the vector error process $\xi_{i,t} = (u_{i,t}, \varepsilon_{i,t}')'$ is a stationary process with an asymptotic covariance matrix Ω_i, which can be decomposed as :

$$\begin{aligned} \Omega_i \equiv \begin{bmatrix} \Omega_{u_i} & \Omega_{u\varepsilon_i} \\ \Omega_{\varepsilon u_i} & \Omega_{\varepsilon_i} \end{bmatrix} &= \Omega_i^0 + \Gamma_i + \Gamma'_i, \\ &= \begin{bmatrix} \Omega^0_{u_i} & \Omega^0_{u\varepsilon_i} \\ \Omega^0_{\varepsilon u_i} & \Omega^0_{\varepsilon_i} \end{bmatrix} + \begin{bmatrix} \Gamma_{u_i} & \Gamma_{u\varepsilon_i} \\ \Gamma_{\varepsilon u_i} & \Gamma_{\varepsilon_i} \end{bmatrix} + \begin{bmatrix} \Gamma'_{u_i} & \Gamma_{u\varepsilon_i} \\ \Gamma_{\varepsilon u_i} & \Gamma'_{\varepsilon_i} \end{bmatrix}, \end{aligned}$$

where Ω_i^0 is the contemporaneous covariance and Γ_i is a weighted sum of auto-covariances. Ω_{u_i} refers to the long run variance of the residual $u_{i,t}$, Ω_{ε_i} is the $(k \times k)$ long run covariance among the $\varepsilon_{i,t}$ and $\Omega_{\varepsilon u_i}$ is a $(k \times 1)$ vector that gives the long run covariance between $u_{i,t}$ and $\varepsilon_{i,t}$. Note that the $\Omega_{\varepsilon u_i}$ captures the endogenous feedback effect between $y_{i,t}$ and $x_{i,t}$. Thus, by considering this feedback effect, the group-mean FMOLS estimator eliminates the bias due to the endogeneity of the regressors, that is :

$$\hat{\beta}_{FM} = N^{-1} \sum_{i=1}^{N} \left(\sum_{t=1}^{T} (x_{i,t} - \bar{x}_i)(x_{i,t} - \bar{x}_i)' \right)^{-1} \left(\sum_{t=1}^{T} (x_{i,t} - \bar{x}_i) y^*_{i,t} - T\hat{\gamma}_i \right), \tag{16}$$

where

$$y^*_{i,t} = (y_{i,t} - \bar{y}_i) - \frac{\hat{\Omega}_{\varepsilon u_i}}{\hat{\Omega}_{\varepsilon_i}} \triangle x_{i,t}, \qquad \hat{\gamma}_i \equiv \hat{\Gamma}_{\varepsilon u_i} + \hat{\Omega}^0_{\varepsilon u_i} - \frac{\hat{\Omega}_{\varepsilon u_i}}{\hat{\Omega}_{\varepsilon_i}} \left(\hat{\Gamma}_{\varepsilon_i} + \hat{\Omega}^0_{\varepsilon_i} \right),$$

and $\bar{y}_i$ $(\bar{x}_i)$ is the simple average of $y_{i,t}$ $(x_{i,t})$ over the cross section dimension (*i.e.* $\bar{y}_i = N^{-1} \sum_{i=1}^{N} y_{i,t}$ and $\bar{x}_i = N^{-1} \sum_{i=1}^{N} x_{i,t}$). Under the assumption of cross sectional independence (*i.e.* $\mathrm{E}\left[\xi_{i,t}, \xi'_{j,t}\right] = 0$ for all $i \neq j$), Pedroni (2000) showed that the group-mean FMOLS is asymptotically unbiased and its t-statistic is standard normal :

$$\begin{aligned} T\sqrt{N}\left(\hat{\beta}_{FM} - \beta\right) &\Rightarrow \mathrm{N}(0,\ \nu), \\ t_{\hat{\beta}_{FM}} &\Rightarrow \mathrm{N}(0, 1), \end{aligned} \tag{17}$$

where ν depends of $\bar{x}_i$, $\bar{y}_i$ and of the dimension of $x_{i,t}$, k[20]. The group-mean FMOLS estimator is consistent and converges at rate $T\sqrt{N}$ to β, so even when T and N are relatively small, $\hat{\beta}_{FM}$ is relatively precise. Finally, in the expression (16), $\hat{\beta}_{FM}$ follows a summation over the cross sectional dimension, it can also be constructed as the average of the conventional time series FMOLS estimator applied to the ith member of the panel as $\hat{\beta}_{FM} = N^{-1} \sum_{i=1}^{N} \hat{\beta}_{FM,i}$, where $\hat{\beta}_{FM,i}$ is the individual time series FMOLS estimator. Likewise, the group mean t-statistic can be computed as $t_{\hat{\beta}_{FM}} = N^{-1/2} \sum_{i=1}^{N} t_{\hat{\beta}_{FM,i}}$, where $t_{\hat{\beta}_{FM,i}}$ is the t-statistic of the individual FMOLS estimator.

The group-mean DOLS estimator, proposed by Pedroni (2001) adds leads and lags of $\triangle x_{i,t}$ (i.e. $\sum_{j=-q}^{q} \triangle x_{i,t+q}$) as additional regressors in (15). This correction allows to take care of a possible endogeneity of the regressors and to correct for correlation between $u_{i,t}$ and $\varepsilon_{i,t}$. Kao and Chiang (2000) showed the superiority of the within DOLS over the within FMOLS. To our knowledge, a comparison of group-mean FMOLS and group-mean DOLS finite sample properties has not yet been investigated by empirical studies[21]. According to us, the DOLS estimator suffers from two drawbacks. First, DOLS estimators are very sensitive to the number of leads and lags included in the regression, small sample properties of these estimators are improved when adding leads and lags (see Kao and Chiang, 2000).

[20] When $k = 1$ and $\bar{x}_i = \bar{y}_i = 0$, $\nu = 2$, if $\bar{x}_i$ and/or $\bar{y}_i \neq 0$, $\nu = 6$.

[21] However, Pedroni's (2000) Monte Carlo simulations reveal that the group-mean DOLS has relatively small size distortion relative to the within DOLS estimator.

Moreover, there is no statistical methodology to choose the optimal number of leads and lags for the DOLS estimators (within or group-mean). Second, given our relatively limited time span ($T = 37$), even for a DOLS estimator with only one lead and one lag the number of degrees of freedom is quite short. For example, with $k = 5$, the total number or regressors is the cointegrating relation is twenty for the DOLS estimator with one lead and one lag, leaving only seventeen degrees of freedom when T is set to thirty seven.

4 Empirical Results

Our sample is based on data availability and included 19 countries of Latin America : Argentina, Bolivia, Brazil, Chile, Colombia, Costa Rica, Dominican Republic, Ecuador, El Salvador, Guatemala, Honduras, Jamaica, Mexico, Panama, Paraguay, Peru, Trinidad and Tobago, Uruguay and Venezuela. The sample covers the period 1970-2006 ($T = 37$) and data are annual. The variables used in this section are the real effective exchange rate ($q_{i,t}$), government spending ($g_{i,t}$), the productivity differential ($prod_{i,t}$), the financial capital inflows ($fci_{i,t}$), the foreign direct investment ($fdi_{i,t}$), the net foreign assets ($nfa_{i,t}$), the terms of trade ($tot_{i,t}$), the openness ($open_{i,t}$) and the *de facto* exchange rate regime ($reg_{i,t}$) for $i = 1, \ldots, 19$ and $t = 1, \ldots, 37$.

4.1 The Data : Sources and Construction

Data is drawn from IMF (International Financial Statistics, *IFS*, and Direction of Trade Statistics, *DTS*), World Bank (World Development Indicators, *WDI*) and CEPAL (Statistical Yearbook) databases.

The real effective exchange rate of the country i at time t is constructed as the geometrical weighted average of the real bilateral exchange rates *vis-a-vis* its ten main trading partners :

$$q_{i,t} = \prod_{j=1}^{10} \left(\frac{P_{i,t}}{E_{ij,t} \; P_{j,t}} \right)^{\omega_{j,t}},$$

where $P_{i,t}$ and $P_{j,t}$ are domestic and foreign CPI price indexes respectively, $E_{ij,t}$ is the nominal exchange rate (in units of domestic currency). $\omega_{j,t}$ is the 3-years moving average trade weight of partner j in total trade of the home country i. Prices and nominal exchange rates series are taken from IMF-*IFS* database. Trade weights were constructed using countries' exportations and importations data from IMF-*DTS* database. All $q_{i,t}$ are expressed as an index (100 = 2000) and converted in logarithms. According to our definition, an increase in $q_{i,t}$ represents an appreciation of the real exchange rate.

For government spending, $g_{i,t}$, we take the share of government consumption in the GDP extracted from World Bank-*WDI* database.

Due to the lack of data availability on sectorial productivity, the GDP per capita relative to trading partners was used as a proxy for the Balassa-Samuelson effect ($prod_{i,t}$). Partner countries' weights are the same as those used in the construction of $q_{i,t}$. GDP per capita data were taken from World Bank-*WDI* database. The recent empirical literature on the Balassa-Samuelson hypothesis has often focused on total factor productivity (*i.e.* Solow residual) or labor productivity differentials between non traded and traded sectors to explain real exchange rate movements[22]. However, these data are unavailable for developing countries. As discussed by Canzoneri *et al.* (1999), the Solow residual specification as a proxy for the Balassa-Samuelson is subject to a variety of limitations : first, it tends to be correlated with variations in aggregate demand[23], second, Solow residual involves data on sectoral labor and capital stock and estimates of labor's share in production that are mostly unavailable for developing countries, and third it is generally associated with a Cobb-Douglas production function which is a restrictive assumption. The ratio of the consumer price index (CPI) to the producer price index (PPI) is also often used as a proxy for the relative productivity effect (see DeLoach, 2001, Alberola, 2003, Bénassy-Quéré *et al.*, 2004 for recent applications)[24]. Like for sectorial labor and production data, PPI indexes are unavailable for a large part of Latin America countries (only eleven countries provide a PPI index with a sufficient time span).

[22] See, among others, De Gregorio, Giovannini and Wolf (1994), Asea and Mendoza (1994), Chinn and Johnston (1997), Canzoneri, Cumby and Diba (1999) and Lee and Tang (2003).

[23] See Hall (1988). Coto-Martinez (2000) and, Coto-Martinez and Reboredo (2003) studied the effect of the fiscal policy on Solow residual.

[24] DeLoach (2001) argues that the logarithms of CPI and PPI are composed of traded and non traded goods such that $CPI \equiv \alpha p_N + (1-\alpha)p_T$ and $PPI \equiv \beta p_N + (1-\beta)p_T$, where p_N and p_T are respectively the price of non traded and traded goods. The relative price of non traded goods can be expressed as : $p_N - p_T = (\alpha-\beta)^{-1}(CPI - PPI)$. Assuming that $\alpha > \beta$, an increase in $(CPI - PPI)$ leads to an increase in the relative price of non traded goods.

Following Athukorala and Rajapatirana (2003), we construct the financial capital inflows as :

$$fci_{i,t} = \frac{(PIA_{i,t} + OIA_{i,t}) - (PIL_{i,t} + OIL_{i,t})}{GDP_{i,t}},$$

where PI and OI are respectively portfolio investments and others investments, with the letter A indicating assets and the letter L liabilities. Others investments cover both private flows (bank loans) and public flows (monetary authorities and general government). PI and OI are both expressed in U.S. dollars (source : CEPAL). $GDP_{i,t}$ is the nominal GDP of the country i, also expressed in U.S. dollars (source : World Bank-*WDI* database).

We define the foreign direct investment, $fdi_{i,t}$, as :

$$fdi_{i,t} = \frac{DIE_{i,t} - DIA_{i,t}}{GDP_{i,t}},$$

where $DIE_{i,t}$ are direct investment in the country i, and $DIA_{i,t}$ are direct investment abroad (source : CEPAL). $DIE_{i,t}$, $DIA_{i,t}$ and $GDP_{i,t}$ are all expressed in U.S. dollars.

The variable $nfa_{i,t}$ is the ratio of net foreign assets to GDP, both expressed in U.S. dollars. Lane and Milesi-Ferretti (2001) provide net foreign assets data for the period 1970-2003. We updated the database to 2006 using the following construction[25] :

$$\triangle NFA_{i,t} = CA_{i,t} + KA_{i,t}, \tag{18}$$

where $NFA_{i,t}$ is the net foreign assets in U.S dollars, $CA_{i,t}$ is the current account and $KA_{i,t}$ is the capital account balance (source : IMF-*IFS* database). Data for Haiti, Honduras and Nicaragua are not provided in Lane and Milesi-Ferretti's database. Thus, we estimated $NFA_{i,t}$ for the period 1971-2006 using equation (18). As initial value for 1970, we took the total external debt of each country (source : World Bank-*WDI* database).

The terms of trade, tot_{it} is defined as the ratio of country's export price index to its import price index (source : CEPAL). Terms of trade are expressed as an index (100 = 2000).

The openness ($open_{i,t}$) is the ratio of imports plus exports to GDP. All variables are in U.S. dollars (source : World Bank-*WDI* database).

We apply the methodology proposed by Coudert and Dubert (2005) to identify the *de facto* exchange rate regime ($reg_{i,t}$) for our sample of countries[26]. This classification is based on three statistical criteria. The first one consists on the estimations of annual trends in the (monthly) nominal exchange rate level in order to distinguish crawling peg from peg regimes. The second criterion allows to separate fixed regimes (pegs and crawling peg) to flexible ones (pure and managed float) by building a comparison test of nominal exchange rate volatility between the Latin America country and a benchmark group of floating currencies. Then, for regimes classified as pure float or managed float in former steps, a third test is applied to distinguish between these two types of regime. More precisely, it is a comparison test of percentage change of official reserves variance with the benchmark group. Thus, we construct our *de facto* regime dummy variable which stands 1 for a fixed regime (peg or crawling peg) and 0 for a flexible regime (pure or managed float).

4.2 Unit Root Tests

Before determining the long run determinants of real exchange rates in Latin America, we first apply panel unit root tests presented in section 3.1 to our series. All these tests are designed under the assumption of cross section independence. O'Connell (1998) emphasized the importance of controlling for cross sectional dependence when testing for unit roots in panels of real exchange rates. Without taking care of correlation between individuals of the panel, he rejected the null hypothesis of unit root in a panel of 64 real exchange rates. By contrast, when controlling for cross sectional correlation, no evidence against the unit root null hypothesis can be found in all the sample and in four geographic sub-samples. Moreover, O'Connell showed that cross sectional dependence adversely affects size and power of panel unit root tests. For controlling for such dependence, one can demean the data over the cross section dimension by subtracting average such as $x^*_{i,t} = x_{i,t} - N^{-1}\sum_{i=1}^{N} x_{i,t}$. Results of panel unit root tests on original[27] and demeaned series are given in table 3.

[25]This construction is equivalent to equation (5) in Milesi-Ferretti (2001).

[26]See appendix A for further details.

[27]The variables $q_{i,t}$, $g_{i,t}$, $prod_{i,t}$, $open_{i,t}$ and $tot_{i,t}$ are directly converted in logarithms. $nfa_{i,t}$ is directly expressed as ratio of GDP. The variables $fci_{i,t}$ and $fdi_{i,t}$, which are also expressed as ratio of GDP are converted into logarithms as $\ln(1+X)$.

Table 3: ***Panel unit root tests results (series in levels)***

	Original series			Demeaned series		
	IPS	MW	LM	IPS	MW	LM
$q_{i,t}$	-0.826	39.791	6.762*	-1.152	43.345	6.928*
$g_{i,t}$	-0.869	36.326	5.470*	-0.877	40.606	5.732*
$prod_{i,t}$	0.873	26.892	7.908*	-0.308	37.135	8.668*
$fci_{i,t}$	-2.484*	55.645*	2.696*	-3.123*	61.015*	2.892*
$fdi_{i,t}$	-2.041*	55.102*	5.917*	-1.337	45.819	6.832*
$nfa_{i,t}$	0.607	36.249	8.255*	0.101	34.458	7.631*
$open_{i,t}$	-0.998	47.681	7.409*	-1.177	43.569	7.649*
$tot_{i,t}$	-4.290*	88.742*	5.864*	-5.004*	100.316*	5.891*

Notes : Critical value at the 5 % significance level for a χ^2 is 53.38 with 2N = 38, is 1.64 for a N(0,1). * : rejection of the null hypothesis at the 5 % significance level.

Applied to original series, the IPS and MW tests indicate that the null hypothesis of non stationarity can not be rejected in favor of the alternative hypothesis of stationarity at the 5 % significance level for all variables except for the financial capital inflows ($fci_{i,t}$,) the foreign direct investments ($fdi_{i,t}$) and the terms of trade ($tot_{i,t}$). However, when controlling for cross sectional dependence, IPS and MW do not reject the null hypothesis of unit root for the foreign direct investments variable. This result implies a common feature in the evolution of FDI flows to Latin America zone[28]. For all other variables, tests results are consistent whatever the specification of series and thus support the hypothesis of a weak correlation between individuals among the variables of the panel. For the terms of trade, even when using the demeaned series, IPS and MW tests do not reject the null hypothesis of non stationarity at the 5 % significance level. The presence of a common structural break for terms of trade during the 70's can being a possible explanation for this puzzle.

Finally, the LM-Hadri test clearly rejects the null hypothesis of stationarity for all variables. However, Hlouskova and Wagner (2005) found that this test has poor size and often leads to over-rejection of the null hypothesis, so they suggest to use the Hadri test in order to find unit root since rejection of the null does not imply acceptance of the alternative of unit root. Thus, when rejecting the null hypothesis of stationary, the LM test results have to be taken with caution.

In order to determine the order of integration of our series, we apply the panel unit roots to series in first differences. The results are reported in table 4.

Table 4: ***Panel unit root tests results (series in first differences)***

	Original series			Demeaned series		
	IPS	MW	LM	IPS	MW	LM
$\triangle q_{i,t}$	-8.415*	147.278*	-1.817	-8.648*	152.227*	-1.817
$\triangle g_{i,t}$	-6.256*	113.843*	-0.873	-7.065*	130.960*	-0.949
$\triangle prod_{i,t}$	-5.648*	97.559*	0.298	-6.593*	113.597*	1.031
$\triangle fci_{i,t}$	-10.071*	178.443*	-2.553	-10.892*	193.545*	-2.538
$\triangle fdi_{i,t}$	-8.539*	150.116*	-0.093	-8.969*	157.970*	-0.253
$\triangle nfa_{i,t}$	-5.310*	95.506*	0.358	-6.095*	108.783*	-0.134
$\triangle open_{i,t}$	-9.117*	158.674*	-1.896	-8.922*	152.996*	-1.848
$\triangle tot_{i,t}$	-10.369*	186.413*	-1.132	-10.729*	195.525*	-0.887

Notes : See notes table 3.

The IPS and MW tests strongly reject the null hypothesis of non stationarity in the panel for all series at the 5 % significance level. In this case, we can follow the LM-Hadri conclusions since Hlouskova and Wagner's (2005) simulations showed that it is more designed to find unit root. For all variables, the LM statistic is below the right tail 5 % critical value of a standard distribution (1.64) so we are able to accept the null hypothesis of stationarity.

In summary, we can conclude that the real exchange rate and all its potential long run determinants, except the terms of trade are integrated of order one (order zero for financial capital inflows and the terms of trade for the period 1970 - 2006). In other words, real exchange rates in Latin America follow

[28]One possible explanation of this common evolution of FDI across Latin America countries can be the incapability of financial markets to discriminate between individual creditworthiness. For example, after the Argentina 2001 crisis, FDI flows dropped in 2002 for 29.06 % in Argentina, 42.9 % in Brazil, 38.4 % in Chile and 38.6 % in Mexico.

random walks, implying that deviations from PPP can be permanent. These deviations can be explained by fundamentals' fluctuations.

4.3 Heterogenous Cointegration : all Sample Results

In this section, we first apply Pedroni's (2003) cointegration tests to find evidence of heterogeneous long run relationships amongst the real exchange rate and its determinants, and second by using the group-mean FMOLS estimator we can estimate efficiently the influence of each determinants on the real exchange rate. Due to the variety of possible explanatory variables, we consider different specifications of the long run real exchange rate model :

$$
\begin{aligned}
&\underline{\text{model 1 :}} \qquad & q_{i,t} &= \alpha_i + \beta\left(g_{i,t}, prod_{i,t}, tot_{i,t}, open_{i,t}, reg_{i,t}\right) + \varepsilon_{i,t}, \\
&\underline{\text{model 2 :}} \qquad & q_{i,t} &= \alpha_i + \beta\left(g_{i,t}, prod_{i,t}, tot_{i,t}, open_{i,t}, reg_{i,t}, fci_{i,t}\right) + \varepsilon_{i,t}, \\
&\underline{\text{model 3 :}} \qquad & q_{i,t} &= \alpha_i + \beta\left(g_{i,t}, prod_{i,t}, tot_{i,t}, open_{i,t}, reg_{i,t}, fdi_{i,t}\right) + \varepsilon_{i,t}, \\
&\underline{\text{model 4 :}} \qquad & q_{i,t} &= \alpha_i + \beta\left(g_{i,t}, prod_{i,t}, tot_{i,t}, open_{i,t}, reg_{i,t}, fci_{i,t}, fdi_{i,t}\right) + \varepsilon_{i,t}, \\
&\underline{\text{model 5 :}} \qquad & q_{i,t} &= \alpha_i + \beta\left(g_{i,t}, prod_{i,t}, tot_{i,t}, open_{i,t}, reg_{i,t}, nfa_{i,t}\right) + \varepsilon_{i,t},
\end{aligned}
$$

where α_i is the fixed effects, $\beta = (\beta_1, \beta_2, \beta_3, \ldots)'$ is the vector of coefficients and $\varepsilon_{i,t}$ the residual. The term α_i captures the country specificity and is needed in regressions because both real exchange rates, productivity differentials and terms of trade are expressed as indexes and hence are not comparable in levels across countries. We consider the model 1 as the framework. This model includes as regressors : the ratio of government spending to GDP, a productivity effect, the terms of trade, the degree of openness and the *de facto* exchange rate regime. Athukorala and Rajapatirana (2003) showed that the composition of capital flows matters in determining their influence on the real exchange rate. Thus, we examine here the impact of two types of capital flows : net foreign direct investment ($fdi_{i,t}$) and foreign capital inflows ($fci_{i,t}$). In a first time, we test the magnitude of each category of capital inflows separately (model 2 and 3). Then, net foreign direct investment and foreign capital inflows are included together in the regression (model 4). Finally, the variable net foreign assets is added to the regression (model 5). Recently, several studies[29] found a *transfer effect*, *i.e.* in the long run net foreign assets improvements are associated with real exchange rate appreciations, thus we expect β_6 to be positive in the model 5.

According to our real exchange rate's definition (an increase in $q_{i,t}$ implies an appreciation of the domestic currency), we would expect $\beta_1 > 0$, $\beta_2 > 0$, $\beta_3 \gtrless 0$, $\beta_4 < 0$, $\beta_5 > 0$. Assuming that government spending fall more on non traded goods, an increase in public consumption will rise total demand for non traded goods and thus rising its relative price and the real exchange rate. The coefficient β_2 measures the impact of the Balassa-Samuelson effect which claims that an increase of traded sector productivity relative to non traded sector should appreciate the real exchange rate. Theoretically, the influence on real exchange rate of the terms of trade is ambiguous since a terms of trade improvement generates two contrary effects (income *versus* substitution). Consequently, the impact of terms of trade on real exchange rate depends whether the income or substitution effect dominates. An increase in the openness degree leads to a convergence of international prices, limiting pressure on the real exchange rate. Hence, a greater openness to trade, through trade-liberalizing reforms for example, is expected to lead to a depreciation of the real exchange rate ($\beta_4 < 0$). The exchange rate regime should affect the real exchange rate mainly through a boom of the consumption of non traded goods which entails an increase of the non traded goods' price, so we expect that rigid exchange rate regimes lead to real exchange rate appreciation (i.e. $\beta_5 > 0$). Finally, coefficients of capital flows in models 2, 3 and 4 (β_6 and β_7 in model 4) are expected to be positively signed.

Before estimating the long-run exchange rate models, we perform a correlation analysis. We focus on the bivariate correlation between, first real exchange rate and its fundamentals (column 2 of table 5 below), and second between the different long run determinants of the real exchange rate (columns 3-8). Bivariate correlations between the determinants of the real exchange rates allow us to identify potential collinearity among the explanatory variables included in models 1-5. Multicollinearity problems can

[29] See Calderon (2002), Alberola *et al.* (2003), Bénassy-Quéré *et al.* (2004), Lane and Milesi-Ferretti (2004), Aguirre and Calderón (2005) and Dufrenot and Yehoue (2005).

adversely affect the estimating cointegrating vector $\hat{\beta}$ and t-statistics. Panel bivariate correlations, noted $\bar{R}(x, y)$, are computed as follows : in a first time we compute individual bivariate correlations, noted r_i^{xy}, between variables x and y for country i, then we average the absolute value of r_i^{xy} across the N dimension that is, $\bar{R}(x, y) = N^{-1}\sum_{i=1}^{N}|r_i^{xy}|$. Note that, this construction of $\bar{R}(x, y)$ provides an indicator of the magnitude of the correlation between variables x and y, and do not permit to determine the sign of the correlation since it is calculated on absolute values of individual correlations. As shown in table 5, the real exchange rate is strongly correlated with the government spending, the productivity effect, terms of trade, the degree of openness and the *de facto* exchange rate regime. We can also find a strong correlation between financial capital inflows and the productivity effect ($\bar{R}(prod, fci) = 0.34$). In a recent report, IMF (2005) stressed the importance of external financial flows as an important element in fuelling Latin America growth, and particulary since the early 90's where a large part of capital inflows were composed of portfolio investments. Terms of trade and the productivity variable are strongly correlated ($\bar{R}(prod, tot)$ = 0.36). This huge link was also found by Lane and Milesi-Ferretti (2004) for their panel of 42 developing countries (see Barro and Sala-i-Martin, 1995 and Mendoza, 1997 for related literature on the link between growth and terms of trade). Foreign direct investment are strongly linked with government spending and the openness. Finally, the variable net foreign assets is correlated with a majority of real exchange rate's determinants : the Balassa-Samuelson effect, terms of trade, the degree of openness, the *de facto* regime and financial capital inflows[30].

Table 5: ***Cross-sectional correlations*** $\bar{R}(x, y)$ ***(1970 - 2006)***

	$q_{i,t}$	$g_{i,t}$	$prod_{i,t}$	$tot_{i,t}$	$open_{i,t}$	$reg_{i,t}$	$fci_{i,t}$	$fdi_{i,t}$	$nfa_{i,t}$
$q_{i,t}$	1.00								
$g_{i,t}$	0.41	1.00							
$prod_{i,t}$	0.45	0.30	1.00						
$tot_{i,t}$	0.36	0.27	0.36	1.00					
$open_{i,t}$	0.40	0.36	0.33	0.35	1.00				
$reg_{i,t}$	0.38	0.29	0.27	0.25	0.21	1.00			
$fci_{i,t}$	0.32	0.22	0.34	0.23	0.33	0.27	1.00		
$fdi_{i,t}$	0.32	0.43	0.29	0.25	0.42	0.16	0.22	1.00	
$nfa_{i,t}$	0.35	0.28	0.51	0.44	0.46	0.36	0.48	0.26	1.00

According to our correlation analysis, a large part of determinants of the real exchange rate are correlated each other. Thus, in order to obtain consistent estimates of the cointegrating vector $\hat{\beta}$ and to reduce possible collinearity among explanatory variables, the variables $tot_{i,t}$, $fci_{i,t}$ and $fdi_{i,t}$ enter into models 1-5 with one lag. This strategy allows us to obtain robust results since it increases the exogeneity of regressors included into cointegration regressions. Table 6 reports estimates of models 1-5 based on the group-mean FMOLS estimator.

The last row of table 6 reports the group parametric-*t* test, $\tilde{Z}^*_{tN,T}$, of cointegration proposed by Pedroni (1999, 2004). In all five regressions, the statistic test is significant and clearly indicates a rejection of the null hypothesis of no cointegration[31]. Hence, there is strong support for a cointegration relationship between real exchange rate and its determinants. Moreover, all our coefficient estimates are highly statistically significant at the 5 % significance level with the expected signs.

For all variables, coefficient estimates have the sign predicted by the theoretical literature. First, a permanent increase in government expenditure as a share of GDP tends to appreciate the real exchange rate in the long run. This result supports the theoretical prediction that government spending are mainly composed of non traded goods. Our estimates suggest that the elasticity of real exchange rate to government spending changes fluctuates around the 0.29 - 0.42 range. These estimations are fairly close to recent studies of real exchange rate behavior in developing countries. Thus, Dufrenot and Yehou (2005) found a coefficient $\hat{\beta}_1$ fluctuating about 0.15 - 0.19, whereas Aguirre and Calderón (2005) found an elasticity of 0.22 for their panel of 38 developing countries. However, our estimates remain relatively high in comparison to Drine and Rault (2003) study ($\hat{\beta}_1 = 0.10$) who considered a group of 17 countries of Latin America over the period 1973 - 1996.

According to theoretical models of equilibrium real exchange rates a permanent change in terms of trade has an ambiguous impact on the real exchange since it generates two contrary effects. In our five regressions, the coefficient of terms of trade is positive and statistically significant at the 5 % level. Terms of trade improvements entail real exchange rate appreciations in Latin America with an elasticity of 0.16 in average, which means that the income effect is predominant. In addition, there is a strong

[30] On the link between terms of trade and net foreign assets, see, for example, Lane and Milesi-Ferretti (2004).

[31] Note that we were able to reject the null hypothesis of no cointegration when in the five models the variables $tot_{i,t}$, $fci_{i,t}$ and $fdi_{i,t}$ were contemporaneous to the others.

relation between real exchange rates and the degree of openness. The coefficient $\hat{\beta}_3$ appeared strongly significant and negative in all models, this result indicates that liberalization of commercial policy lead to real depreciations. The estimated elasticity for openness is quite stable among the regressions with a coefficient close to -0.35. This value is comparable with that of Elbadawi (1994) and Dufrenot and Yehou (2005).

Table 6: ***Long-run determinants of real exchange rates : group-mean FMOLS results***

All countries (N = 19 and T = 37)

	Model 1	Model 2	Model 3	Model 4	Model 5
$g_{i,t}$	0.369***	0.321***	0.327***	0.286**	0.420***
	(5.41)	(5.16)	(5.32)	(2.10)	(5.77)
$prod_{i,t}$	0.548***	0.534***	0.631***	0.585***	0.449***
	(6.87)	(5.56)	(7.03)	(5.78)	(4.96)
$tot_{i,t-1}$	0.162***	0.164***	0.141***	0.150***	0.164**
	(2.92)	(2.69)	(3.29)	(3.36)	(2.04)
$open_{i,t}$	-0.308***	-0.349***	-0.370***	-0.404***	-0.364***
	(-7.21)	(-7.88)	(-6.91)	(-7.87)	(-8.73)
$reg_{i,t}$	0.267***	0.260***	0.252***	0.245***	0.229***
	(12.29)	(12.19)	(11.96)	(12.02)	(10.22)
$fci_{i,t-1}$		0.622*		0.660*	
		(1.71)		(1.65)	
$fdi_{i,t-1}$			2.972**	2.649*	
			(2.05)	(1.92)	
$nfa_{i,t}$					0.121***
					(2.68)
$\tilde{Z}^*_{tN,T}$	-4.362***	-3.530***	-4.224***	-2.167**	-3.798***

Heteroskedasticity and autocorrelation consistent t-statistics are reported in parentheses. * (respectively **, ***) : rejection of the null hypothesis at 10 % significance level (respectively 5 % and 1 %).

Moreover, our empirical results confirm that rigid *de facto* exchange rate regimes tend to appreciate the real exchange rate. Hence, the exchange rate regime is not neutral regarding its effects on real exchange rate. This *non neutrality* contrasts with the view of many authors (see Flood and Rose, 1995, and Obstfeld and Rogoff, 2000) which state that the exchange rate regime is neutral regarding the evolution and the volatility of real macroeconomic variables.

The estimated coefficients of both types of capital inflows are positive and statistically significant at the 5 % level, implying that a surge in foreign capital flows are accompanied by real exchange rate appreciations in Latin America. The positive coefficients $\hat{\beta}_6$ and $\hat{\beta}_7$ in model 4 suggest that an increase in net capital inflows rises domestic absorption and induces a reallocation of output factors towards non traded sector. This shift in the composition of output exercises an upward pressure on the price of non traded goods and thus appreciates the real exchange rate. However, the impact on the real exchange rate of the two categories of capital inflows exhibit different elasticity magnitudes. According to models 3 to 4 results, the real exchange rate appreciates about 2.8 % following a permanent 1 % rise in FDI flows as share of GDP, whereas an increase of 1 % in portfolio investments inflows (as share of GDP) leads to a real exchange appreciation of 0.6 %. Our estimates are quite different from those estimated by Athukorala and Rajapatirana (2003) who found an elasticity about 1.70 for financial capital flows and -0.06 for foreign direct investment. These differences can be partially explained by methodologies employed in their study comparing to this one. Indeed, they focused only on a restrict group of six countries in Latin America[32] over the period 1985-2000.

The results concerning net foreign assets confirm a significant transfer effect, that is, permanent improvements in net foreign assets tend to appreciate the real exchange rate in the long run. Our net

[32] Argentina, Brazil, Chile, Colombia, Mexico and Peru.

foreign assets coefficient estimate is 0.121, which is close to those obtained by Calderón (2002), and Lane and Milesi-Ferretti (2004), respectively 0.15 - 0.22 and 0.19 - 0.29 ranges.

4.4 Heterogenous Cointegration : Sample Splits Results

In this section, we run the FMOLS regressions for country subgroups. The entire sample is split according to geographical criteria. We divide our sample in two areas : *South America* (Argentina, Bolivia, Brazil, Chile, Colombia, Ecuador, Paraguay, Peru, Uruguay and Venezuela) and *Caribbean and Central America* (Costa Rica, Dominican Republic, El Salvador, Guatemala, Honduras, Jamaica, Mexico, Panama and Trinidad and Tobago). This methodology is applied in order to take account of a possible heterogeneity between the two zones. It is possible to think that the two groups have specific characteristics in trade, productive structure and external financial linkages, specially with the dollar area.

As a previous analysis, panel unit roots are implemented for sub samples series. Results are displayed in appendix B, table 13. As in the whole sample case, we are able to reject the null hypothesis of non stationarity for the variable terms of trade for any group split and for any specification of this variable. According to IPS and MW results, this implies that the terms of trade are integrated of order zero. For Caribbean and Central America sub-group, all other variables are stationary at the 5 % significance level. In the South America sample, the IPS and MW tests applied to original series indicate that the null hypothesis of non stationarity can not be rejected for all variables, except for terms of trade, portfolio inflows and foreign direct investment. Moreover, when controlling for cross sectional dependence, portfolio inflows and foreign direct investment variables reject the null of non stationarity, implying for a possible common trend that driven the evolution of these two variables across the individual dimension[33].

Table 7: ***Long-run determinants of real exchange rates : group-mean FMOLS results***

South America (N = 10 and T = 37)

	Model 1	Model 2	Model 3	Model 4	Model 5
$g_{i,t}$	0.499***	0.410***	0.393***	0.327***	0.531***
	(3.77)	(3.40)	(3.35)	(3.04)	(4.95)
$prod_{i,t}$	0.682***	0.699***	0.837***	0.806***	0.580***
	(4.92)	(4.63)	(5.61)	(5.30)	(5.60)
$tot_{i,t-1}$	0.315***	0.297***	0.232***	0.213**	0.180***
	(2.82)	(2.39)	(2.53)	(2.11)	(2.68)
$open_{i,t}$	-0.378***	-0.394***	-0.458***	-0.460***	-0.524***
	(-4.30)	(-3.79)	(-4.88)	(-4.46)	(-6.48)
$reg_{i,t}$	0.286***	0.284***	0.267***	0.266***	0.274***
	(6.54)	(6.41)	(6.44)	(6.32)	(6.34)
$fci_{i,t-1}$		0.891		0.894	
		(0.96)		(0.89)	
$fdi_{i,t-1}$			2.599	2.264	
			(1.43)	(1.23)	
$nfa_{i,t}$					-0.566***
					(-2.85)
$\tilde{Z}^*_{tN,T}$	-2.325***	-1.297*	-1.949**	-0.969	-2.275**

Notes : See notes table 6.

The estimation results for the South America group are reported in table 7. In model 1, all coefficient estimates are significant at the 1 % level and have the expected sign. Furthermore, the Pedroni's cointegration test statistics is significant, hence there is evidence for cointegration hypothesis.

In models 1, 2, 3 and 5, we are also able to reject the null of no cointegration at least at the 10 % significance level. However, in models 3 and 4, financial capital inflows and foreign direct investments

[33]However, as we will see latter, these two variables are not significant in cointegration relations and thus the issue of mixed regressions (i.e. regressions with both I(1) and I(0) variables) is not matter in this case.

are not significant. Finally, in model 5 we see that all regressors are significant, and regarding the net foreign assets variable, the coefficient is now negative.

Table 8 reports estimates of models 1-5 for Caribbean and Central America group. In model 1, one variable is not statistically significant : the terms of trade. For many countries in Central America, the terms of trade fluctuations are instead an important source of exogenous fluctuations in the GDP. When the terms of trade variable is removed from equation (model 1b), the others variables remains similar.It is possible that in the Caribbean and Central America case it could be interesting to disentangle the separate effects of exports prices and imports prices, given the particular characteristics of these economies (most of them very small and mono-producers). Model 2 supports the hypothesis of cointegration between the real exchange rate and fundamentals like government spending, the Balassa Samuelson effect, the degree of openness, the exchange rate regime and portfolio inflows. All coefficients are statistically significant with the expected sign, with the exception of the portfolio inflows variable .An increase of public spending, productivity, trade liberalization and a surge in foreign direct investments tend to appreciate the real exchange rates in the Caribbean and Central America group. Like in the South America case, rigid *de facto* exchange rate regimes lead to an upward pressure on the real exchange rate. However, the coefficient is less stronger here than previously, the implication is that fixed regimes tend to appreciate more the real exchange rate in South America group than in Caribbean and Central America zone. In models 3 and 4, we can reject the null hypothesis of no cointegration, foreign direct investments have an impact on the real exchange rate in the long run. Results of model 5 confirm the existence of the *transfer effect* in Caribbean and Central America countries : improving net external positions are associated with appreciating real exchange rates. This evidence contrasts with the opposite relation found it in South America.

Table 8: ***Long-run determinants of real exchange rates : group-mean FMOLS results***

Caribbean and Central America (N = 9 and T = 37)

	Model 1	Model 1b	Model 2	Model 3	Model 4	Model 5
$g_{i,t}$	0.226***	0.178***	0.190***	0.140***	0.152***	0.249**
	(3.88)	(3.24)	(3.82)	(2.89)	(3.07)	(3.51)
$prod_{i,t}$	0.399***	0.265***	0.225*	0.289***	0.252*	0.216**
	(4.80)	(3.19)	(1.95)	(3.17)	(1.78)	(2.26)
$tot_{i,t-1}$	-0.007					
	(-1.26)					
$open_{i,t}$	-0.231***	-0.311***	-0.346***	-0.259***	-0.310***	-0.332***
	(-5.94)	(-6.58)	(-7.37)	(-5.90)	(-6.93)	(-7.97)
$reg_{i,t}$	0.247***	0.254***	0.239***	0.234***	0.225***	0.171***
	(10.96)	(10.66)	(10.45)	(10.33)	(10.18)	(7.56)
$fci_{i,t-1}$			0.350		0.192	
			(1.13)		(0.60)	
$fdi_{i,t-1}$				-2.294***	-1.992*	
				(-2.49)	(-1.93)	
$nfa_{i,t}$						0.362***
						(4.34)
$\tilde{Z}^*_{tN,T}$	-1.908**	-3.127***	-3.497***	-2.801***	-3.106***	-3.334***

Notes : See notes table 6.

In summary, our results support the hypothesis of cointegration between the real exchange rate and a set of real variables that includes a Balassa-Samuelson effect, government spending, the degree of openness, foreign capital inflow (particulary foreign direct investments), net foreign assets and the nominal exchange rate regime. However, sub-sample regressions showed that exchange rate responses to real shocks are different in Latin America and in Central America. In both zones, the long run real exchange rate responds to the government spending, the Balassa-Samuelson effect, changes in the openness degree, net foreign assets and the nominal exchange regime. Nevertheless, fixed regimes and trade restrictions in South America generate greater real appreciations than in Caribbean and Central

America.

Others real factors seem to be region specific. An improvement of the terms of trade entail a real appreciation in South American countries. On the other hand, net foreign assets has an opposite effect in South America than in Caribbean and Central America.

4.5 Equilibrium Exchange Rates and Misalignments

From our estimates, we compute the equilibrium real exchange rate, $\bar{q}_t$ using sustainable values of the fundamentals, that is, $\bar{q}_t = \hat{\beta}' X_t^p$ where X_t^p is the permanent part of fundamentals and $\hat{\beta}'$ the FMOLS vector of coefficients. Here we extract the permanent component of fundamentals using the Hodrick and Prescott filter. Misalignments or deviations between the current real exchange and its equilibrium level are therefore measured by :

$$q_t^d = q_t - \bar{q}_t$$

According to our real exchange rate construction, if $q_t^d > 0$ it would suggest that the actual real exchange rate is overvalued. Similarly, the real exchange rate is undervalued when $q_t^d < 0$.

In order to compute consistent estimations for equilibrium exchange rates in our panel, we retain coefficients estimates from subgroups regressions. Indeed, since these estimations are more country specific than estimations for the whole panel, we expect to obtain more precise misalignment measures. In both cases, we retain a version of model 5, that includes as fundamentals the government spending, the Balassa-Samuelson effect, the terms of trade, the trade openness, and the net foreign assets for South America countries. In the case of Caribbean and Central America countries, a version of model 5 containing as fundamentals government spending, the Balassa-Samuelson effect, the trade openness and the net foreign assets. In fact, the model 5 is very similar in both geographic areas with four regressors in common and permits comparisons. For the South America zone, models 2 to 4 can not be used in order to provide consistent estimations for equilibrium exchange rates because there is no cointegration relation in model 4, and in models 2 and 3 the variables $fci_{i,t}$ *and* $fdi_{i,t}$ are not significant. The same arguments hold for Caribbean and Central America estimations (no cointegration in models 3 and 4)[34].

Table 9 presents real exchange rate deviations from equilibrium for 2006[35]. Eight currencies over nineteen are close to their equilibrium level in 2006, indeed the misalignment is less than ±10% which can be interpreted as an equilibrium situation under cyclical circumstances. For countries like Ecuador Honduras and Jamaica, the deviation is even less than 4 %. This finding contrasts with undervalued currencies for which the undervaluation is in average about 43.2 %. The table shows a very large undervaluation in Bolivia and Paraguay. On the other hand, big countries as Brazil and Mexico exhibit a huge overvaluation in 2006 with a deviation of 45.1 % and 27.9 % respectively. El Salvador, Guatemala Venezuela and Peru's real exchange rates are also strongly above their equilibrium level.

Table 9: ***Real Exchange Rate Misalignments in 2006***

Undervalued currencies		Equilibrated currencies		Overvalued currencies	
Bolivia	-78.6 %	Argentina	-6.2 %	Brazil	45.1 %
Dominican Rep	-13.5 %	Colombia	-5.5 %	El Salvador	42.0 %
Panama	-14.6 %	Costa Rica	-6.9 %	Guatemala	25.2 %
Paraguay	-66.0 %	Chile	-10.0 %	Mexico	27.9 %
		Ecuador	3.2 %	Peru	51.6 %
		Honduras	3.1 %	Uruguay	15.3 %
		Jamaica	2.2 %	Venezuela	31.4 %
		Trinidad & T.	-5.2 %		

Notes : A currency is designed as equilibrated if the misalignment is less than ±10%.

Real exchange rate misalignments are often used as a leading indicator of currency crises (Kaminsky *et al.*, 1997 and Kaminsky and Reinhart, 1998). It has been showed that real exchange rate overvaluations provide an early warning indicator of possible currency crashes. Real appreciations of the domestic currency can signal potential problems in the current account position. Moreover, since the real exchange rate guides the internal resource allocation, a persistent overvaluation induces a non optimal allocation between sectors of production. In order to test the efficiency of our model specification as a leading indicator of crisis, we compare the misalignment level before a currency crisis with the misalignment observed after the crash. Currency crises are detected using the analysis of Frankel and Rose (1996)[36].

[34] Misalignments were also computed using model 1 in both zones and model 5 in Caribbean and Central America group. Results, not reported here to conserve space, are very close to those of model 2.

[35] Misalignments for each country over the period 1970 - 2006 are reported in appendix, table 14, page 28.

[36] They define a currency crash as a large change of the nominal exchange rate that is also a substantial increase in the rate of change of nominal depreciation.

The table 10 presents the evolution of misalignments before and after several currency crises that occurred in Latin America and captures a number of interesting features. First, our results show that real exchange rates were strongly overvalued the years that precede the crisis, the misalignment reaches his peak the year just before the crisis (except for Brazil, 1999 and Uruguay, 2002). In this sense, our equilibrium real exchange rate estimates confirm the hypothesis that real exchange rate misalignments can be used to prevent currency crises.

Table 10: ***Real Exchange Rate Misalignments and Currency Crises***

Country	Crisis Year (= t)	t-2	t-1	t	t+1	t+2
Argentina	1975	92.47	100.00	-260.07	-207.03	-160.31
Argentina	1981	45.66	100.00	28.98	-225.22	-203.10
Argentina	2002	91.95	100.00	-60.64	34.06	38.64
Bolivia	1982	42.60	100.00	-91.84	36.83	7.47
Brazil	1999	108.41	100.00	52.42	56.03	41.74
Chile	1982	71.86	100.00	64.44	77.28	80.06
Costa Rica	1981	88.35	100.00	-367.35	-82.46	-22.51
Dominican Rep.	1985	60.43	100.00	-41.11	-15.15	-148.47
El Salvador	1986	5.65	100.00	-156.21	-41.38	78.17
El Salvador	1990	40.62	100.00	10.99	23.10	2.24
Guatemala	1986	59.32	100.00	-161.28	-149.67	-144.16
Honduras	1990	71.56	100.00	-157.26	-48.55	-38.80
Jamaica	1978	97.24	100.00	-5.74	30.07	65.70
Mexico	1994	73.74	100.00	-63.84	-94.09	11.56
Paraguay	1984	82.65	100.00	74.45	56.58	-14.03
Uruguay	1990	21.91	100.00	-69.86	4.44	55.36
Uruguay	2002	153.95	100.00	-30.40	-114.43	46.63
Venezuela	1984	63.27	100.00	-26.77	3.10	-278.86
Venezuela	2002	94.66	100.00	32.89	50.79	48.20

Notes : The year t refers to the beginning of the crisis. The misalignments are expressed as an index (100 = t-1). In every case, the misalignment is positive in year t- 1. A sign + (-) indicates an overvaluation (undervaluation) of the real exchange rate.

Second, estimates from model 2 indicate that following the crisis, the real exchange rate tends to undershoot for a short period its equilibrium level, *i.e.* after the crisis the real exchange is undervalued. For example after the 1994 crisis in Mexico, the peso was strongly below its equilibrium rate in years 1994 and 1995. After the initial strong undervaluation, the real exchange rate appreciates and converges towards its equilibrium rate (see Bolivia, 1982, Costa Rica, 1981 and Uruguay, 1990 for example). Third, the effect of devaluations on initial misalignment is country specific. Some nominal devaluations permit to well correct the pre-crisis overvaluation (Venezuela, 2002, El Salvador, 1990, and Jamaica, 1978), whereas in Brazil (1999) and Paraguay (1984) the devaluation reduces only partially the gap between the current real exchange rate and its equilibrium level. In these cases, the real exchange rate continues to be overvaluated despite a nominal devaluation.

Finally, regarding the Argentinean convertibility regime, the table 11 shows a high degree of real exchange rate appreciation when the *currency board* was introduced in April 1991. At this time, the peso exhibited an overvaluation of 42.7 %. After 1993, the overvaluation began to decline and persisted around 30 % during the period 1995 - 1998. This is consistent with the appreciation of the Brazilian real over those years following the implementation of the Real Plan (see table 14)[37]. The gap between the peso and its equilibrium level accentuated after the 1999 Brazilian devaluation (the misalignment jumped from 31.4 % in 1998 to 41.2 % in 1999) and reached 46.3 % in 2001 just before the collapse of the regime in December 2001. Note that our pre-crisis misalignment, 46.3 % is in line with those found by Gay and Pellegrini (2004) and Alberola *et al.* (2004), 44 % and 53 % respectively. Williamson's (1995) analysis on the currency board system show its difficulty to correct endogenously an initial real exchange rate misalignment. The initial overvaluation and the incapability of the currency board to correct it led to a persistent peso's overvaluation and to a dramatic loss in competitiveness which in turn weakened the convertibility regime[38].

[37] Note that Brazil is the first trading partner for Argentina and thus the Brazilian real enter with an important weight when constructing the real effective exchange rate of Argentina (about 33 % over the period 1990 - 2006).

[38] See Carrera (2002) for a general discussion of the currency board experience in Argentina.

Table 11: ***Misalignment of the Argentinean peso***

Year	Misalignment	Year	Misalignment
1991	+ 42.7 %	1999	+ 41.2 %
1992	+ 56.5 %	2000	+ 42.6 %
1993	+ 63.1 %	2001	+ 46.3 %
1994	+ 43.6 %	2002	- 28.1 %
1995	+ 32.8 %	2003	- 15.8 %
1996	+ 29.4 %	2004	- 17.9 %
1997	+ 31.5 %	2005	- 11.5 %
1998	+ 31.4 %	2006	- 6.2 %

Notes : + = overvaluation, - = undervaluation.

5 Conclusions

The main goal of this paper was to determinate what factors influence real exchange rates in nineteen Latin American countries over the 1970 - 2006 period. Using panel cointegration techniques, we estimated several models of real exchange rate according to fundamentals that are included in regressions. Referring to the theoretical literature, six traditional fundamentals retain our attention : the Balassa-Samuelson effect, government spending, the terms of trade, the country's openness to international trade, foreign capital inflows and the net foreign assets position. In addition, we also include the *de facto* nominal exchange regime as a factor determining the evolution of real exchanges rates in Latin America. Hence, we follow the methodology proposed by Coudert and Dubert (2005) to identify the *de facto* regime. After, we divide our countries sample in two subgroups, that is South America (ten countries) and Caribbean and Central America (nine countries). We ran long run regressions in both subgroups and found some interesting differences. Finally, we estimate the equilibrium levels real exchange rates and compute the degree of misalignment.

The main empirical results are the following. First, there is a strong evidence that, over the long run, Latin America's real exchanges rates are non stationary implying that PPP does not hold in this region. This result is in line with the findings of Edwards and Savastano (1999) who claim that recent studies focusing on real exchange rates in Latin America do not support the hypothesis of PPP. Thus, real shocks have a permanent effect on the real exchange rates' paths.

Second, we identify six real factors which have a potential effect on real exchange rates. Estimations for the whole sample (nineteen countries) confirm the theoretical links between the real exchange rate and its determinants. That is, a higher government spending to GDP ratio, an increase in productivity differential, a positive terms of trade shock, a surge in foreign capital flows and an higher net foreign assets position affect positively the real exchange rate in Latin America. Whereas an increase in trade openness leads to an depreciation of the real exchange rate. The *de facto* exchange regime has also a strong influence on real exchange rates in Latin America : rigid regimes (peg or crawling peg) exercise an upward pressure on the real exchange rate. This stylized fact has been also highlighted by IMF (2005), exchange rate stabilization programs in Latin America at the beginning 90's introduced inflation stickiness which in turn leads to an increase of the real exchange rate. This finding shows the non neutrality of exchange rate regime regarding its effects on real exchange rates whatever the credibility's level of the fixed regime. The tendency of fixed regime to appreciate the real exchange rate can be seen as an adverse effect of exchange rate stabilization implementations in Latin America.

Third, there is evidence that real exchange rates behaviors are different between South America and Caribbean and Central America. Some fundamentals play a role in only one of the two subgroups, foreign direct investment matter for Caribbean and Central America area solely, whereas the terms of trade is significant only in South America countries. Regarding the magnitude of FMOLS coefficients, there are also differences between the two zones. Fixed regimes and trade restrictions policies are associated with greater appreciations in South America than in Caribbean and Central America.

Fourth, our equilibrium real exchange rates estimations confirm that persistent overvaluation can provide a strongly early warning for currencies crises in Latin America. A large number of currency crises experienced by Latin American countries were preceded by huge and persistent overvaluations. This finding has an important policy implication for countries with a fixed *de facto* regime. Our nominal regime classification shows that in 2006, sixteen countries maintained a *de facto* fixed regime (peg or crawling peg)[39]. El Salvador and Peru exhibit significant and persistent overvaluations since 1998, with an average gap between actual real exchange rate and its equilibrium level about 39.0 % and 44.60 %

[39] Argentina, Bolivia, Chile, Colombia, Costa Rica, Dominican Rep, Ecuador, El Salvador, Guatemala, Honduras, Jamaica, Mexico, Panama, and Trinidad and Tobago, Uruguay and Venezuela.

respectively. For countries like Chile, Colombia, Costa Rica, Ecuador, Honduras, Jamaica, and Trinidad and Tobago, the real exchange rate observed in 2006 was in line with its equilibrium level (see tables 9 and 14).
Moreover, our model specification offers explanations for the recent currency history of Argentina. The last three important devaluations experiences in 1974, 1980 and 2001 were preceded by high real exchange rate overvaluations : 49.1 % , 50.6 % and 46.3 % respectively. As shown by Edwards (1994), expansive and inconsistent policies can generate, under a fixed regime, persistent overvaluation, and when there is a real exchange misalignment, nominal devaluations can be a powerful tool to restore equilibrium.

Finally, further research might attempt to improve our empirical analysis by taking account of others shocks that influence real exchange rates in Latin America. Specially in the Central America case, remittances from United States represent a growing part of income for countries like Dominican Republic, El Salvador and Honduras. Even for a "big"country like Mexico they are an important part of his money inflows. These flows reached, on average 12 % of GDP in El Salvador during the 90's, and 3 % of GDP in Honduras (source : IMF, 2005). Since the beginning of the 90's, remittances exhibit an upward tendency, so we can expect that it will play an increasing influence on real exchange rates in the future.

Appendix

A The *de facto* exchange rate regime classification :

The classification process of Coudert and Dubert (2005) can be summarized as follows :

1. Step 1 : **Assessing annual trend in the exchange rate**

   By using monthly exchange rates (against U.S. dollar), the annual trend is extracted from the following regression :
   $$\ln e_t = \alpha + \gamma \text{ time} + \varepsilon_t$$
   where $\ln e_t$ is the logarithm of the monthly nominal exchange rate against the U.S. dollar, time is a linear trend and ε_t the residual term. The annual trend of the year j , denoted by β_j is constructed from the OLS estimator of γ as $\hat{\beta}_j = (1+\hat{\gamma})^{12} - 1$. If $\hat{\beta}_j$ is found positive, series of monthly exchange rates are detrended ($\tilde{e}_t$ designed the detrended exchange rate). If the annual trend is negative, its absolute value, $|\hat{\beta}_j|$, is compared to an arbitrarily threshold τ. Following Coudert and Dubert (2005), τ is set to 2 % annually.

2. Step 2 : **Comparing exchange rate variances**

   The second step consists in comparing the annual variance of changes in $\triangle e_t$ (or $\triangle \tilde{e}_t$ if β_j is found positive in step one) to the average variance of a benchmark floating currencies. The benchmark sample of floating currencies is made up of the Japanese Yen, the British Pound and the German Deutsche Mark (after 1999, the euro stands in for the deutsche mark). By considering a floating currencies benchmark, we can compute Fisher tests applied to variance of nominal exchange rates.

   Let s_i^2 denote the empirical annual variance of $\triangle e_t$ for the Latin America country i and s_B^2 the average of annual variance of the benchmark. Assuming that annual variances follow normal distributions with theoretical variance σ_i^2 for the Latin America country and σ_B^2 for the benchmark group, then the ratio $(s_B^2/\sigma_B^2)/(s_i^2/\sigma_i^2)$ follows a Fisher distribution $F(n_B, n_i)$ where n_B and n_i designate degrees of freedom. Since s_i^2 and s_i^2 are respectively calculated with 36 and 12 data n_B is equal to 35 and n_i to 11. The null hypothesis is, for a given year, the variance of exchange rate changes in the Latin America country is smaller than the one in the benchmark panel, that is $H_0 : \sigma_i^2 < \sigma_B^2$. Note that the 5 % critical value of an $F(35, 11)$ is 2.54. If $s_i^2 < (1/2.54)\ s_B^2$, the exchange rate variance of the country i is considered as *low.* If $s_i^2 \geq (1/2.54)\ s_B^2$, the variance of the country i is considered as *high.*

3. Step 3 : **Comparing changes in international reserves variances**

   In this stage we compute the same test as in second step to variance of changes in official reserves ($\triangle R$). Thus, changes in foreign reserves empirical variance for the Latin America country i (denoted by $\tilde{s}_i^2$ with a theoretical value $\tilde{\sigma}_i^2$) will be compared to the average variance of changes in foreign reserves in the benchmark sample (noted $\tilde{s}_B^2$ with a theoretical value $\tilde{\sigma}_B^2$). Assuming that monthly rates of change in reserves follow normal distributions, a new Fisher test can be computed. In this case, the null hypothesis is, for a given year, the variance of the reserves change in Latin America country $\tilde{\sigma}_i^2$is greater than the one in the benchmark group $\tilde{\sigma}_B^2$, that is $H_0 : \tilde{\sigma}_i^2 > \tilde{\sigma}_B^2$. Then the ratio $(\tilde{s}_i^2/\tilde{\sigma}_i^2)/(\tilde{s}_B^2/\tilde{\sigma}_B^2)$ follows a Fisher distribution $F(35, 11)$.

   If $\tilde{s}_i^2 > 2.54\ \tilde{s}_B^2$, the variance of international reserves is considered as *high.* Otherwise, if $\tilde{s}_i^2 \leq 2.54\ \tilde{s}_B^2$, the variance of international reserves is considered as *low.*

At final, according to results in the three steps, nine cases can be distinguished :

Annual trend	Exchange rate variance	Official reserves variance	Type of regime
$\hat{\beta} > 0$	*high*	*low*	**pure float**
$\hat{\beta} > 0$	*high*	*high*	**managed float**
$\hat{\beta} > 0$ and $\hat{\beta} < \tau$	*low*	$-$	**peg**
$\hat{\beta} > 0$ and $\hat{\beta} > \tau$	*low*	$-$	**crawling peg**
$\hat{\beta} < 0$ and $\|\beta\| > \tau$	$-$	*low*	**pure float**
$\hat{\beta} < 0$ and $\|\beta\| > \tau$	$-$	*high*	**managed float**
$\hat{\beta} < 0$ and $\|\beta\| < \tau$	*low*	$-$	**peg**
$\hat{\beta} < 0$ and $\|\beta\| < \tau$	*high*	*low*	**pure float**
$\hat{\beta} < 0$ and $\|\beta\| < \tau$	*high*	*high*	**managed float**

All data on monthly nominal exchange rates and official reserves are extracted from the IMF-*IFS* database.

The classification is reported in table 12.

Table 12: ***De Facto Exchange Rate Regime Classification***

	Arg	Bol	Bra	Chi	Col	CR	Dom	Ecu	Sal	Gua	Hon	Jam	Mex	Pan	Par	Per	Tri	Uru	Ven
1970	1	4	1	4	3	4	4	2	4	4	4	1	4	4	4	4	1	4	4
1971	1	4	1	1	3	4	4	1	4	4	4	1	4	4	4	4	1	4	4
1972	2	2	3	2	3	4	4	4	4	4	4	1	4	4	4	4	1	2	4
1973	4	2	4	2	3	4	4	4	4	4	4	4	4	4	4	4	2	2	4
1974	4	4	3	2	3	2	4	4	4	4	4	4	4	4	4	4	4	2	4
1975	2	4	3	2	3	4	4	4	4	4	4	4	4	4	4	2	2	2	4
1976	2	4	1	2	3	4	4	4	4	4	4	4	2	4	4	1	1	2	4
1977	1	4	3	1	3	4	4	4	4	4	4	4	1	4	4	1	4	3	4
1978	3	4	3	3	3	4	4	4	4	4	4	2	4	4	4	2	4	3	4
1979	3	2	1	1	3	4	4	4	4	4	4	2	4	4	4	3	4	3	4
1980	3	2	2	4	3	4	4	4	4	4	4	4	3	4	4	3	4	3	4
1981	2	4	3	4	3	2	4	4	4	4	4	4	3	4	4	3	4	3	4
1982	2	2	3	2	3	2	4	2	4	4	4	4	2	4	4	3	4	2	4
1983	2	2	2	3	3	2	4	2	4	4	4	2	2	4	4	2	4	2	4
1984	2	2	4	2	3	2	4	3	4	4	4	2	3	4	1	2	4	2	1
1985	2	2	4	2	3	3	2	2	4	4	4	2	2	4	1	1	2	2	4
1986	2	3	1	1	3	3	2	2	2	2	4	4	3	4	1	1	4	3	1
1987	2	3	1	3	3	3	2	1	4	4	4	4	2	4	4	1	4	3	4
1988	2	3	2	3	3	3	2	2	4	2	4	4	2	4	4	2	2	3	4
1989	2	3	2	3	3	3	4	3	4	2	4	2	3	4	2	2	4	3	2
1990	2	3	2	3	3	3	2	3	2	2	4	2	2	3	1	4	2	2	4
1991	2	3	2	3	3	3	3	2	2	2	2	2	3	4	4	2	4	3	3
1992	4	3	1	3	3	3	1	1	3	3	3	1	4	4	3	1	4	3	1
1993	4	3	1	3	3	3	3	3	4	3	1	1	4	4	3	3	2	3	1
1994	4	3	2	1	1	3	2	2	4	2	2	2	2	4	3	2	2	2	2
1995	4	3	2	1	3	3	4	3	4	3	3	2	2	4	4	3	4	3	2
1996	4	3	3	3	1	3	1	1	4	1	2	1	1	4	3	3	3	3	2
1997	4	3	3	3	1	3	2	3	4	3	3	3	3	4	3	3	4	3	2
1998	4	3	3	3	1	3	3	2	4	3	3	3	1	4	2	3	3	3	3
1999	4	3	2	1	1	3	4	1	4	1	3	3	1	4	2	1	2	3	3
2000	4	3	2	1	1	3	3	2	4	4	3	3	1	4	3	4	4	3	3
2001	4	3	2	1	4	3	4	4	4	3	3	3	1	4	2	1	4	2	3
2002	2	3	2	1	1	3	2	4	4	1	3	3	1	4	2	3	4	2	2
2003	1	3	1	1	1	3	2	4	4	3	3	2	1	4	1	4	4	2	2
2004	1	3	1	1	1	3	2	4	4	1	3	3	3	4	1	1	4	2	2
2005	2	4	2	2	2	3	2	4	4	4	4	3	1	4	1	3	4	2	2
2006	4	4	2	3	4	3	4	4	4	4	4	3	3	4	1	1	4	4	4

Notes: 4 = peg, 3 = crawling peg, 2 = managed float, 1 = pure float.
Arg = Argentina, Bol = Bolivia, Bra = Brazil, Chi= Chile, Col = Colombia, CR = Costa Rica, Dom = Dominican Republic, Ecu = Ecuador, Sal = El Salvador, Gua = Guatemala, Hon = Honduras, Jam = Jamaica, Mex =Mexico, Pan = Panama, Par = Paraguay, Per = Peru, Tri = Trinidad and Tobago, Uru = Uruguay and Ven = Venezuela.

B Panel unit root tests in sample splits

Table 13: ***Panel unit root tests results : sample splits***

	Original series			Demeaned series		
	IPS	MW	LM	IPS	MW	LM
			South America (N = 10)			
$q_{i,t}$	-1.522	27.819	4.363*	-1.472	28.708	4.865*
$g_{i,t}$	-0.765	20.635	3.588*	-1.209	26.620	4.307*
$prod_{i,t}$	0.638	16.320	5.774*	-0.254	21.353	6.847*
$fci_{i,t}$	-2.608*	35.987*	1.083	-2.600*	33.917*	2.019*
$fdi_{i,t}$	-1.743*	31.554*	2.178*	-0.835	25.654	5.192*
$nfa_{i,t}$	-1.399	22.146	4.542*	0.405	21.507	5.519*
$open_{i,t}$	-1.336	32.536*	4.563*	-1.398	27.296	4.839*
$tot_{i,t}$	-3.784*	59.899*	5.305*	-4.674*	67.433*	5.400*
$\triangle q_{i,t}$	-6.747*	86.888*	-1.328	-7.128*	91.583*	-1.256
$\triangle g_{i,t}$	-5.491*	74.958*	-1.076	-6.856*	92.793*	-0.996
$\triangle prod_{i,t}$	-4.359*	55.266*	0.690	-4.536*	57.479*	1.453
$\triangle fci_{i,t}$	-7.253*	93.365*	-2.073	-8.124*	105.632*	-2.022
$\triangle fdi_{i,t}$	-5.535*	71.753*	-0.842	-6.207*	80.565*	-0.166
$\triangle nfa_{i,t}$	-4.554*	58.422*	-0.675	-5.102*	65.774*	-0.460
$\triangle open_{i,t}$	-6.629*	84.403*	-1.461	-6.733*	84.113*	-1.479
$\triangle tot_{i,t}$	-7.607*	101.390*	0.024	-7.579*	101.269*	-0.048
			Caribbean and Central America (N = 9)			
$q_{i,t}$	0.403	11.972	5.226*	-0.123	14.637	4.938*
$g_{i,t}$	-0.456	15.691	4.166*	0.001	13.986	4.073*
$prod_{i,t}$	0.596	10.572	5.403*	-0.179	15.782	5.378*
$fci_{i,t}$	-0.860	19.657	2.776*	1.797*	27.097	2.073*
$fdi_{i,t}$	-1.127	23.548	6.303*	-1.062	20.166	4.454*
$nfa_{i,t}$	2.358	7.103	7.206*	-0.574	12.951	5.269*
$open_{i,t}$	-0.027	15.145	5.956*	-0.237	16.274	6.012*
$tot_{i,t}$	-2.244*	28.943*	2.929*	-2.344*	32.883*	2.867*
$\triangle q_{i,t}$	-5.115*	60.389*	-1.241	-5.053*	60.644*	-1.317
$\triangle g_{i,t}$	-3.301*	38.885*	-0.134	-3.037*	38.167*	-0.330
$\triangle prod_{i,t}$	-3.612*	42.294*	-0.294	-4.797*	58.118*	-0.033
$\triangle fci_{i,t}$	-6.987*	85.078*	-1.497	-7.263*	87.913*	-1.557
$\triangle fdi_{i,t}$	-6.572*	78.363*	0.751	-6.849*	77.405*	-0.194
$\triangle nfa_{i,t}$	-2.915*	37.084*	1.233	-3.477*	43.001*	0.289
$\triangle open_{i,t}$	-6.259*	75.271*	-1.216	-5.866*	68.883*	1.126
$\triangle tot_{i,t}$	-7.046*	85.023*	-1.671	-7.598*	94.256*	-1.340

Notes : See notes table 3. Critical values at 5 % significance level for a χ^2 are 31.41 with 2N = 20 and 28.87 with 2N = 18. * : rejection of H_0 at 5 % significance level.

C Real Exchange Rate Misalignments

Table 14: ***Real Exchange Rate Misalignments***

	Arg	Bol	Bra	Chi	Col	CR	Dom	Ecu	Sal	Gua	Hon	Jam	Mex	Pan	Par	Per	Tri	Uru	Ven
1970	-0.467	0.448	0.439	0.631	0.293	0.294	-0.043	0.381	-0.740	0.029	-0.084	0.093	0.091	0.213	0.200	-0.193	0.060	0.230	0.004
1971	-0.368	0.406	0.414	0.024	0.233	0.286	-0.048	0.372	-0.799	-0.022	-0.122	0.138	0.074	0.182	0.156	-0.202	0.050	-0.062	0.017
1972	0.111	-0.059	0.403	0.005	0.221	0.294	0.005	0.340	-0.810	-0.039	-0.089	0.085	0.083	0.173	0.092	-0.164	-0.012	-0.873	-0.003
1973	0.454	0.043	0.392	-2.687	0.215	0.317	0.057	0.333	-0.924	-0.013	-0.125	0.091	0.104	0.120	-0.005	-0.151	-0.020	-0.414	-0.047
1974	0.491	0.273	0.350	-0.766	0.167	0.242	0.070	0.322	-0.821	0.018	-0.102	0.191	0.189	0.116	0.047	-0.100	0.023	-0.565	-0.098
1975	-1.276	0.464	0.386	-0.263	0.160	0.314	0.126	0.333	-0.654	0.049	-0.091	0.256	0.240	0.069	0.396	-0.037	-0.108	-0.116	-0.090
1976	-1.016	0.411	0.336	0.125	0.160	0.316	0.142	0.323	-0.650	0.089	-0.093	0.297	-0.382	0.054	0.308	-0.269	-0.113	-0.150	-0.090
1977	-0.786	0.319	0.288	0.165	0.256	0.281	0.177	0.309	-0.669	0.089	-0.104	0.305	-0.162	0.006	0.200	-0.918	-0.137	-0.117	-0.121
1978	-0.244	0.203	0.222	0.135	0.226	0.255	0.124	0.262	-0.671	0.056	-0.148	-0.018	-0.082	-0.035	0.095	-1.038	-0.172	-0.163	-0.169
1979	0.231	0.037	-0.133	0.208	0.277	0.284	0.119	0.250	-0.587	0.094	-0.099	0.092	0.013	0.010	0.227	-0.635	-0.176	0.107	-0.137
1980	0.506	0.189	-0.091	0.333	0.269	0.322	0.145	0.248	-0.555	0.070	-0.035	0.201	0.141	0.066	0.213	-0.604	-0.166	0.236	-0.031
1981	0.147	0.445	-0.009	0.463	0.300	-1.182	0.147	0.323	-0.361	0.173	0.049	0.272	-0.044	0.151	0.389	-0.378	-0.097	0.399	0.092
1982	-1.140	-0.408	0.012	0.299	0.306	-0.265	0.233	0.198	-0.243	0.136	0.096	0.330	-1.196	0.191	0.541	-0.615	-0.023	-0.224	0.175
1983	-1.028	0.164	-0.543	0.358	0.283	-0.072	0.260	0.129	-0.129	0.113	0.148	-0.098	0.470	0.149	0.655	-0.717	0.127	0.070	0.276
1984	-1.012	0.033	-0.641	0.371	0.247	-0.037	0.430	0.174	0.007	0.133	0.204	-0.193	-0.194	0.164	0.488	-1.035	0.241	0.015	-0.074
1985	-0.262	-1.048	-0.914	-0.132	-0.021	-0.096	-0.177	0.073	0.133	0.223	0.188	-0.106	-0.577	0.141	0.371	-1.127	-0.122	-0.108	0.009
1986	-0.297	0.237	-0.271	-0.202	-0.166	-0.056	-0.065	-0.241	-0.207	-0.360	0.243	0.052	-1.218	0.167	-0.092	-0.371	-0.062	-0.137	-0.770
1987	-0.741	0.290	-1.105	-0.206	-0.234	-0.202	-0.638	-0.475	-0.055	-0.334	0.204	-0.011	-1.423	0.118	0.223	-0.082	0.006	0.014	-0.384
1988	-0.279	0.308	-1.935	-0.034	-0.187	-0.133	-0.540	-0.681	0.104	-0.322	0.231	0.015	-0.160	0.044	0.430	-2.745	-0.074	0.061	0.002
1989	-4.397	0.479	-2.017	0.032	-0.213	-0.040	-0.037	-0.509	0.255	-0.491	0.322	-0.042	-0.136	0.020	0.195	-0.045	0.062	0.277	-0.565
1990	0.133	0.017	-0.960	-0.122	-0.355	-0.145	-0.325	-0.473	0.028	-0.516	-0.507	-0.177	-0.055	-0.040	-0.043	-0.722	0.130	-0.194	-0.356
1991	0.427	-0.032	-1.426	-0.069	-0.336	-0.286	-0.060	-0.565	0.059	-0.211	-0.157	-1.198	0.090	-0.077	0.047	0.349	0.173	0.012	-0.220
1992	0.565	-0.046	-1.616	0.060	-0.211	-0.111	-0.036	-0.482	0.006	-0.166	-0.125	-0.279	0.209	-0.091	0.023	0.368	0.228	0.154	-0.164
1993	0.631	-0.064	-2.253	0.033	-0.173	-0.158	-0.002	-0.137	0.197	-0.182	-0.261	-0.585	0.283	-0.133	0.114	0.437	0.034	0.321	-0.108
1994	0.436	-0.301	0.077	-0.016	0.019	-0.140	0.082	-0.119	0.252	-0.019	-0.350	-0.266	-0.181	-0.155	-0.222	0.445	0.030	0.063	-0.227
1995	0.328	-0.349	0.381	-0.047	-0.048	-0.108	0.196	-0.228	0.302	-0.014	-0.149	-0.330	-0.266	-0.165	-0.370	0.434	0.035	-0.027	-0.382
1996	0.294	-0.296	0.443	-0.048	0.058	-0.108	0.127	-0.253	0.337	0.061	-0.174	0.025	0.033	-0.171	-0.425	0.416	-0.004	-0.039	-0.187
1997	0.315	-0.282	0.463	0.020	-0.097	-0.117	0.094	-0.140	0.350	0.081	0.001	0.049	0.178	-0.115	0.536	0.454	-0.016	0.017	0.162
1998	0.314	-0.303	0.427	-0.093	-0.158	-0.141	-0.018	-0.314	0.363	0.030	0.055	0.077	0.125	-0.159	-0.711	0.396	-0.053	0.047	0.288
1999	0.412	-0.331	0.224	-0.169	-0.275	-0.154	-0.016	-1.451	0.368	-0.083	0.091	-0.003	0.271	-0.128	-0.723	0.380	0.011	0.165	0.353
2000	0.426	-0.382	0.239	-0.204	-0.417	-0.130	-0.027	-0.515	0.367	-0.034	0.112	-0.030	0.315	-0.157	-0.723	0.413	0.011	0.163	0.399
2001	0.463	-0.451	0.178	-0.308	0.404	-0.106	-0.012	-0.084	0.380	-0.017	0.111	-0.011	0.374	-0.197	-1.061	0.459	0.032	0.106	0.421
2002	-0.281	-0.288	-0.067	-0.220	-0.595	-0.122	-0.221	0.049	0.389	0.082	0.085	-0.038	0.310	-0.163	-0.964	0.468	0.057	-0.032	0.139
2003	-0.158	-0.577	0.147	-0.145	-0.539	-0.167	-0.804	0.055	0.396	0.086	0.058	-0.185	0.245	0.189	-0.849	0.433	-0.002	-0.121	0.214
2004	-0.179	-0.689	0.249	-0.127	-0.269	-0.163	-0.054	0.026	0.409	0.161	0.025	-0.092	0.248	-0.229	-0.981	0.471	-0.044	0.048	0.203
2005	-0.115	-0.728	0.403	-0.025	-0.109	-0.114	-0.179	0.043	0.417	0.226	0.040	0.018	0.291	-0.166	-0.998	0.474	-0.049	0.151	0.239
2006	-0.062	-0.786	0.451	-0.101	-0.055	-0.069	-0.135	0.032	0.420	0.252	0.032	0.022	0.278	-0.146	-0.660	0.516	-0.052	0.153	0.314

Notes: Figures in the table correspond to the deviations of the actual real exchange rate, q_t, to its equilibrium level, $\bar{q}_t$. Misalignments are expressed as a fraction of q_t, that is, $\frac{q_t - \bar{q}_t}{q_t}$.

+ = overvaluation, - = undervaluation.

Arg = Argentina, Bol = Bolivia, Bra = Brazil, Chi = Chile, Col = Colombia, CR = Costa Rica, Dom = Dominican Republic, Ecu = Ecuador, Sal = El Salvador, Gua = Guatemala, Hon = Honduras, Jam = Jamaica, Mex = Mexico, Pan = Panama, Par = Paraguay, Per = Peru, Tri = Trinidad and Tobago, Uru = Uruguay and Ven = Venezuela.

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