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Abstract

This paper investigates the dynamics of multilateral real effective exchange rates (RER) using a panel data set of 19 Latin American countries for the last 45 years. Our results do not support the PPP, so real shocks tend to have permanent effects. Using non-stationary panel econometrics, our estimations show that: a) terms of trade, productivity, capital flows, government spending and net foreign assets exhibit strong relationships with RER; b) exchange rate regimes are not neutral and c) the subsamples regressions are consistent. Finally, we discuss the causes of the persistent overvaluation that the region has experienced in the 2000's.

Keywords: Equilibrium Real Exchange Rate; Panel Unit Roots, Panel Cointegration, Latin America.

J.E.L. Classification: F31, C23.

Resumen

Se estudia la dinámica de los tipos de cambio reales multilaterales (TCR) en un panel de 19 países de América Latina para 45 años. Nuestros resultados no avalan la PPA por lo cual las perturbaciones reales tienen efectos permanentes. Usando cointegración en paneles encontramos que: a) términos de intercambio, productividad, diferentes formas de flujos de capital, gasto público y activos externos netos tienen una fuerte relación con el TCR; b) los regímenes cambiarios no son neutrales y c) las muestras subregionales son consistentes. Finalmente, se discute sobre las causas de la moderada sobrevaluación que caracteriza a la región en los 2000.

Palabras claves**:** Tipo de cambio real de equilibrio, Determinantes macroeconómicos, Raíces Unitarias en Paneles, Cointegración en paneles, América Latina.

Códigos JEL: F31, C23.

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1. INTRODUCTION

The role of exchange rates in Latin American economies has evolved over the last half century. For several decades, the nominal exchange rate dynamics served as the main anchor of nominal stability but also it has been one of the main sources of economic and financial crises. However, since the start of the new century, the role of exchange rates has been transformed to that of shock absorber for booms in capital flows and terms of trade. This same period has seen an expansion of inflation targeting as the dominant monetary policy regime in the region accompanied by the flexible exchange rate regime as its complement (IMF, 2007, 2015 and Cepal, 2014, 2015).In this evolving context started in the seventies, the RER in Latin American countries have shown a dynamic characterized by large swings, strong volatility and persistent misalignments. Therefore, due to the combination of different real and policy factors, real exchange rates volatility and misalignments continues to be a relevant subject of analysis for theoretical and policy reasons.

Table 1 shows the average level of the RER for the last four decades in nineteen South and Central American countries. The 90's saw an increase in the overall level of the RER (relative appreciation) as well as a decrease in volatility relative to the 70's and 80's for the region as a whole. Most likely this was the result of the widespread implementation of exchange rate stabilization plans in South America in the early 90's and the loosening of the external gap with the arrivals of massive capital inflows. With the adoption of exchange rate-based anchors, a large part of Central American countries also succeeded in taming RER volatility during the 90's. Since the late 90's, most of the countries have adopted inflation targeting regimes that have been associated with more flexibility in exchange rate behavior. This change in the institutional policy set up was accompanied by important improvements in terms of trade (especially in South America) and a new wave of capital flows.

The impact of ERER misalignments[1]and volatility in developing countries has been a subject of important debates. Kappler *et al*. (2013), observe that large exchange rate appreciations and revaluations have an impact on the current account as they lead to marked changes of savings and investment within countries. Appreciation shocks impact external balances mainly by reducing exports and stimulating imports. According to them, overall growth is much more affected by overvalued currencies in non-advanced economies than in advanced ones. In the same vein, Aguirre and Calderón (2005) found a nonlinear relation between economic performance and misalignments in a panel of sixty countries: large overvaluations and undervaluations hurt growth, whereas small undervaluations can boost growth. Besides this, ERER misalignments affect internal production and consumption allocations between traded and non-traded goods. Rodrick (2008) claims that sustained undervaluation could improve investment in tradable goods compensating, in a second best fashion, for institutional weakness and market failures. On the other hand, recurrent and large misalignments have been linked to lower growth rates and current account deficits in the long-run, frequently associated with currency and financial crisis (Kaminsky *et al*., 1997 and Kaminsky and Reinhart, 1998).

[1]RER misalignments are defined as the degree of deviation relative to the equilibrium RER level. See Section 2 below.

Table1: **Real Effective Exchange Rates in Latin America (1970-2014)**

	Volatility				Average level (2000 = 100)			
	70 - 79	80 - 89	90 - 00	2001 - 14	70 - 79	80 - 89	90 - 00	2001 - 14
Argentina	0.57	0.71	0.25	0.31	66.81	42.28	98.15	55.13
Bolivia	0.23	0.33	0.13	0.13	166.10	181.00	107.44	99.71
Brazil	0.18	0.35	0.57	0.22	87.81	40.30	85.53	117.47
Chile	0.72	0.35	0.07	0.07	163.71	145.01	103.10	104.65
Colombia	0.07	0.27	0.18	0.16	136.60	127.50	106.97	116.72
Ecuador	0.11	0.35	0.20	0.06	226.65	193.35	116.24	151.95
Paraguay	0.17	0.38	0.22	0.11	196.30	307.58	119.35	103.48
Peru	0.25	0.32	0.21	0.07	49.36	36.05	95.85	106.42
Uruguay	0.19	0.24	0.15	0.14	74.73	93.03	92.48	89.32
Venezuela	0.04	0.38	0.33	0.35	100.17	89.10	63.18	106.00
South America *(10 countries)*	**0.25**	**0.37**	**0.23**	**0.16**	**126.83**	**125.52**	**98.83**	**105.08**
Costa Rica	0.07	0.30	0.05	0.15	193.10	102.44	98.93	112.44
Dominican Rep.	0.06	0.38	0.12	0.12	146.43	128.05	98.38	98.19
Guatemala	0.04	0.27	0.11	0.14	147.03	133.59	100.88	133.42
Honduras	0.06	0.07	0.13	0.09	127.89	150.03	82.78	110.97
Jamaica	0.14	0.26	0.27	0.07	139.40	101.54	83.35	99.08
Mexico	0.17	0.38	0.19	0.09	111.94	67.43	83.06	91.15
Nicaragua	0.06	1.06	0.27	0.04	46.74	122.19	104.05	92.89
Panama	0.11	0.06	0.06	0.05	145.20	142.05	105.13	94.90
Trinidad & Tobago	0.06	0.16	0.12	0.16	96.60	120.74	104.06	126.30
Central America *(9 countries)*	**0.09**	**0.33**	**0.15**	**0.10**	**128.26**	**118.67**	**95.62**	**106.59**
Latin America *(19 countries)*	**0.17**	**0.35**	**0.19**	**0.13**	**127.50**	**122.28**	**97.31**	**105.80**

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

Considering the importance of the link between the principal macroeconomic variables and ERER misalignments for developing countries, determining which fundamentals drive real exchange rate behavior in the long-run is still a crucial issue, especially in Latin America because of its persistent external vulnerabilities. First, commodity exports continue to represent a dominant proportion of total exports (IFM, 2015 and CEPAL, 2015). Given the fact that commodity prices have shown a larger volatility relative to manufactured goods; this is an important source of volatility in the current account. Second, foreign capital inflows are strong and volatile first order determinants of the business cycle in Latin America in a context where push factors -such as the international interest rate or, more generally, global liquidity conditions in the center economies –have increased their relevance (see Rey, 2015; Aizenman, Chinn and Ito, 2016).Third, external private and public debt remains relatively high in most Latin American countries, at the same time that dollarization of residents' portfolios and financial systems are still important in domestic financial systems. Since these stocks are vulnerable to ERER movements, they could affect the financial stability of the countries.

Given this framework, the paper has two main goals. The first one is to investigate economic variables that influence the long-run real exchange rates dynamics. The second one is to estimate equilibrium paths for the nineteen real exchange rates in order to compute the degree of misalignment between the equilibrium and the observed real exchange rate, with the aim of analyzing the behavior of misalignments in the last 45 years in the region. In order to achieve those goals, the rest of the paper is organized as follows. Section 2 briefly reviews the literature on the fundamentals of the real exchange rate. In Section 3, we present the econometric methodology and the database used to estimate the relationship between the real exchange rate and its long-run fundamentals. In Section 4, we analyze the results from the cointegrating regression and perform some robustness checks on regional subsamples. In Section 5, we compute the exchange rate misalignment and analyze it with special attention given to the period beginning in 2000. Finally, Section 6 draws some conclusions.

2. REAL EXCHANGE RATE DETERMINANTS

The oldest and most common theory to determine the ERER is the purchasing-power parity (PPP) approach.

"Under the skin of any international economist lies a deep-stated belief in some variant of the purchasing power parity theory of the exchange rate" (Dornbusch and Krugman, 1976:540)

The PPP approach indicates that the nominal exchange rate of a country is givenby the ratio between the domestic and foreign price level:

$$E^{PPP} = \frac{P}{P*} \tag{1}$$

The real exchange rate based on the PPP approach is a measure of the long-run equilibrium exchange rate. This means that while in the short-run the nominal exchange rate may deviate from that suggested by PPP, the extent of deviation from PPP might be thought of as an over or undervaluation of the home currency (Égert, Halpern and MacDonald, 2006). The PPP approach suggests that the ERER is defined as the exchange rate which equalizes the purchasing power parity of local currencies between countries. According to this approach, ERER, or q, not only tends to 1 but also does not vary through time, and any deviations from the exchange rate equilibrium are only temporally. This implies that:

$$q = \frac{E}{E^{PPP}}=1 \tag{2}$$

The PPP - ERER is based on the law of one price (LOOP). According to the LOOP, if there is perfect competition in both the home and foreign markets, no trade barriers and unrestricted capital movements, in the absence of tariffs and transportation costs, the domestic price of any good i, $\mathrm{P}(i)$, is defined as:

$$P(i) = P^{*}(i)E \tag{3}$$

Where$P^*(i)$ is foreign price the good i and E the exchange rate in domestic currency. Then, if the LOOP holds for every good,E corresponds to PPP based ERER.

In spite of the attractiveness of PPP as a way to determine ERER because of its simplicity, there is no consensus in the specialized literature about its validity. Bastourre, Carrera and Ibarlucía (2004) maintain that at the end of the nineties there was some agreement about that the theory does not hold. However, Sarno y Taylor (2003) argue that the use of panel data, more powerful statistical tests and longer times series introduced a new set of evidence that does not reject the PPP. In the same way, models allowing nonlinear exchange rate dynamics provide further evidence in favor of the PPP acceptance

Meanwhile, other theoretical approaches emerged which provided a more comprehensive way to analyze ERER behavior, based on evidence that reject the PPP approach. According to this vision, commonly referred the macroeconomic approach of ERER (Calderón,2004), the equilibrium 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 (Edwards, 1988, 1989, 1999, 2011).In this framework, while internal equilibrium is achieved when non-traded goods and labor markets clear, external equilibrium is related to the intertemporal budget constraints (*i.e.* the economy is intertemporally solvent). In fact, ERER is not a constant value and its deviations from the PPP are the rule. Based on that idea, Edwards (1989) argues that:

"*Simplified views based on the purchasing power parity theory have suggested that the equilibrium real exchange rate is a constant that does not vary through time. Speaking rigorously, however, there is no reason why the value of the RER required to attain internal and external equilibrium should be a constant number; it would indeed be an extraordinary* coincidence if it was". Edwards (1989: 15).

Within this approach, two types of models were derived: the fundamental equilibrium real exchange rate models (FEER) and the behavioral equilibrium exchange rate models (BEER). Despite the fact that both are macroeconomic approaches, there is a key difference between them. The aim of the FEER models is to estimate the ERER compatible with *desirable macroeconomic conditions*. As a result, that models estimate a desirable ERER (*i.e.* the ERER dynamics that are compatible with desirable future macroeconomic conditions) (Clark y MacDonald, 1999).In contrast, the aim of BEER models is to determine the effective real exchange rate equilibrium, given the dynamics of the fundamentals that define an external and internal equilibrium. The ERER derived from the BEER approach can be formalized as follow:

$$q_T = \beta' X_t + \varepsilon_t \qquad (4)$$

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 $\overline{q}_t$, Clark and MacDonald (1999) suggest using sustainable values or the permanent component of fundamentals. By estimating (4) with a consistent econometric method, we can obtain the equilibrium path for real exchange rate:

$$\overline{q}_t = \hat{\beta}' X_t^p \qquad (5)$$

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 (e.g. the Hodrick-Prescott filter, Beveridge-Nelson decomposition or Gonzalo-Granger methodology). From

(5), 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 $q_t^d \equiv q_t - \bar{q}_t = q_t - \hat{\beta}' X_t^p$.

Within the BEER approach, two types of methodological techniques used to derive ERER according to macroeconomic equilibrium could be distinguished: the set of models that apply a multi-equation approach and the set of models that opt for a single equilibrium model. In spite of that difference, all of them assume that ERER has both real and nominal fundamentals which vary across models, depending on the specification of system of equations used in order to define macroeconomic equilibrium as simultaneous internal and external balance accounts[2].

The last stage of BEER classification includes general equilibrium models. In fact, since the introduction in the middle of the nineties of the New Open Economy Macroeconomic Models (NOEM) several papers in the theoretical ground have used this framework in order to study RER dynamics. The key aspect of that new Keynesian model is to introduce some frictions in wage and price settings in a two countries dynamic stochastic general equilibrium background. This configuration allows models to have a non-neutral role for monetary policy shocks in the real variables dynamics (Obstfeld and Rogoff, 1995). In some sense NOEM model are a microfunded revival of overshooting hypothesis that differentiated short run results than long run results. The empirical implementation of those models has been wide making use of several econometric techniques. As a recent example of different uses of that approach, Berka, Deveraux and Engel, 2014) has use this type of models to analyze the impact on RER of sectoral productivity shocks and unit labor cost differences in a fix exchange rate regime (ie. a common currency like the Eurozone).

A literature review on equilibrium ERER models derived from the BEER approach suggests that only a limited number of variables seem to influence the real exchange rate in the long-run. The most common include the Balassa-Samuelson effect, government spending, terms of trade, trade openness, capital inflows and net foreign assets. Taking this into account, the aim of our paper is not to validate or to reject any particular real exchange rate model formulated in a previous study, but to study the dynamics of real exchange rates with non-stationary panel data methods. Since we adopt an *atheoretical* approach, we are allowed to include the most important fundamentals in the long-run real exchange rate equation that are relevant in empirical literature, reducing the possibility that an omitted fundamental has been captured by another because of misspecification errors in simultaneous equilibrium models. Before building a parametric model of the equilibrium exchange rate, in following subsections, we first discuss the rationale of the selected fundamentals in our paper.

2.1. Productivity effect

According to Balassa (1964) and Samuelson (1964), the relative price of non-tradable goods, q, is determined by the traded/non-traded productivity differential. The explanation is the following. Consider a two sector economy (traded and 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

[2] Without any pretension to be exhaustive, in Appendix A, we present the key differences in the fundamentals included in the models in the latest theoretical research.

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*, with *i* = *T, N*), the formal expression of the Balassa-Samuelson effect is:

$$\hat{q} \equiv \hat{p}_T = \frac{\beta}{\alpha}\hat{\mathrm{A}}_T - \hat{\mathrm{A}}_N \tag{6}$$

Where a hat above a variable denotes growth rate[3]. Thus, according to equation (6), the real exchange rate depends entirely on productivity differentials. Moreover, the Harrod-Balassa-Samuelson effect can be also interpreted as the effect of economic development on real exchange rate, *i.e.* fast growing countries based on tradable sectors tend to experience a real appreciation of their exchange rate. Recently the HBS effect has been introduced into the so called new open macroeconomic models. In that regard, Berka, Deveraux and Engel (2015) develop a DSGE model assuming a common currency to replicate RER behavior in from of sectoral TFP differentiated growth and shocks in labor costs.

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. 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 the goods markets, free movement of factors between the two sectors of production, internationally mobile capital, law of one price for traded goods and constant returns to scale in the two sectors. For example, by introducing monopolistic competition in the non-traded sector in the Lane and Milesi-Ferretti (2004) model, Aguirre and Calderón (2005) allow for demand factors to influence the real exchange in the long run. In that model, 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 exerts an upward pressure on the relative price of non-traded goods and thus appreciates the real exchange rate. In the same fashion, Moreno Badía and Segura-Ubiergo (2014) find that a permanent fiscal adjustment reduces appreciation pressures, especially when is based on current spending. Finally, Galstyan and Lane (2009) and Chatterjee and Mursagulov (2012) show that the composition of government spending influences the long-run behavior of RER. While consumption spending appreciates the currency, public investments in infrastructure tend to depreciate it because of its positive impacts on the productivity of private capital and labor in both the traded and non-traded sectors.

2.3 Terms of Trade

In the economic history of Latin American countries, terms of trade have played a fundamental role in order to explain short term macroeconomics fluctuations and growth paths. The period of commodity prices booms fueled by demand pressure from Asia and massive injections of global

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[3] The formal derivation of equation (6) can be found in Froot and Rogoff (1994) and De Gregorio et al. (1994).

liquidity that started at the very beginning of this century, has put this variable at the center of RER discussion.

All the 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 an improvement in the terms of trade on the real exchange rate is theoretically undefined because of two contrary effects playing in opposite ways. First, an improvement of terms of trade induces a positive income effect (increase in the domestic purchasing power) and results in an augment in the private demand for non-traded goods and then to a real appreciation of the exchange rate. On the other hand, a substitution effect makes the consumption of imported goods relatively cheaper. As a result, there is a shift of demand in favor of the traded goods, and the reestablishment of the equilibrium in the non-traded market is provided by a decrease of the real exchange rate. In fact, the total effect of a term 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.

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[4]. As a result, the real exchange rate depreciates to restore the equilibrium in the non-traded market. The second theoretical influence channel has been emphasized by Obstfeld and Rogoff (2000) and Hau (2002). According to their model 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 allow the convergence of the domestic price index. 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 indices. 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. Bodart *et al.* (2015) shows that commodity price effects on real exchange rates are smoothed when countries are more open to international trade. Finally, Nouira and Sekkat (2015) find that openness is an important variable to explain RER dynamics in a large sample of developing countries.

2.5 Capital Flows and Net Foreign Assets

The impact of capital flow cycles on the real exchange rate has been an outstanding issue in the last decades in Latin America as in other emerging market economies. Most emerging market economies have moved toward capital account openness. This has meant the volume and variety of flows have increased dramatically, but in a way that is far from stable and homogeneous.

[4] 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.

Capital inflows are associated with real exchange rate appreciation in the long run (Corden 1994). 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. In the non-tradable goods market, this excess demand must be matched by a proportional increase in 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 the exchange rate definition in terms of tradable and non-tradable prices, the change in p_N, following the foreign capital inflows, entails an appreciation of the real exchange rate.

However, Nouira and Sekkat (2015) and Combes *et al.* (2012) showed that the impact of capital flows differs by type of flows. Since foreign direct investment, which tends to be more concentrated in the traded sector, is associated with productive improvements, the real appreciation is lower than in the case of portfolio flows. Econometric results from Combes *et al*. (2012) show that real appreciation driven by portfolio capital flow is barely one-seventh of the real appreciations associated with portfolio flows. Moreover, Athukorala and Rajapatirana (2003) compare the effect of FDI and portfolio flows on real exchange rates in Latin America and Asia and concluded that the degree of real appreciation following portfolio inflows is stronger in Latin American countries than in Asia.

In countries where currency mismatch is high, an increase in capital inflows also could weaken financial stability. For example, currency appreciations that are associated with higher capital flows, increases asset prices improving the balance sheet of currency denominated debt and the perception of country risk, which encourages a new round of inflows that expand credit. If inflows are deposited in the domestic financial system, they could increase maturity and currency mismatches between assets and liabilities of domestic banks. So the volume and composition of capital flows have influence not only on competitiveness, but also on financial stability and in the formation of potential macroeconomic crisis if a sudden stop occurs (Calvo, Izquierdo y Talvi, 2003; Reihart y Calvo, 1999; Hofmann, Shim and Shin, 2016).

Considering the impact of the net foreign assets position (NFA) on equilibrium exchange rate, Lane and Milesi-Feretti (2004) argue that the relationship between capital flows, international payments and the real exchange rate - the transfer problem - is one of the classic questions in international economics since the debate between Keynes (1929) and Ohlin (1929). As such, the relationship between net foreign asset positions and the real exchange rate has been analyzed by several intertemporal macroeconomic models (Obstfeld and Rogoff, 1995, Lane and Milesi-Ferretti, 2004) which predict debtor (creditor) countries will havemore depreciated (appreciated) real exchange rates. Indeed, countries with net foreign liabilities need to run a trade surplus to finance interest 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 increase 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 regimes on the real exchange rate is a controversial issue in both the theoretical and empirical literature. In the traditional literature, the dynamics of potential misalignments depend on the level of price stickiness and financial openness. According to some theories of exchange rate determination, such as traditional equilibrium models, the exchange rate regime is neutral in the determination of the level or volatility of the real exchange rate (Baxter and Stockman, 1989; Flood and Rose, 1995; Obstfeld and Rogoff, 2000). However, recent equilibrium literature based on nominal rigidities or market imperfections has found non-neutrality between these variables, but this finding is highly dependent on the type of imperfection that is highlighted (Sarno and Taylor, 2002).

For Latin American countries, the literature on exchange rate-based stabilization plans highlights the significant effects of the adoption of a peg on real exchange rate levels. The stylized fact is the following: when a country suffered from high inflation, one of the preferred policy alternatives in the last three decades was to peg the nominal exchange rate, in most instances, following a strong devaluation. Because of high inflation inertia, changes in prices and wages tend to persist and so the fixed exchange rate regime is associated with a persistent appreciation of the real exchange rate. In several papers, Calvo and Vegh point to the credibility problem associated with this inertial appreciation following fixed exchange rate regime stabilization plans (Calvo and Vegh, 1993). Normally, non-credible exchange rate regime stabilizations are followed by a consumption boom, so we expect that rigid exchange rate regimes lead to real exchange rate appreciation (Calvo and Reinhart, 2000; Caputo, 2015).

Moreover, according to the IMF (2005), the exchange rate stabilization plans introduced in Latin America in the early 90's encouraged a surge in external capital inflows that appreciated the real exchange rates. By adopting a fixed regime, a country provides, at least in the short term, 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 in the domestic monetary base if the central bank does not sterilize it. These capital flows also allow rapid expansions of the domestic credit supply, which exacerbates consumption booms.

Since the mid-90's, most countries in Latin America have adopted gradually regimes of inflation targeting. Now, interest rates more than monetary aggregates or exchange rate are the main instruments of monetary policy. In this regard, the officially declared exchange rate regime is a flexible one. However, several papers have shown that the use of interest rate in a context of open capital accounts tend to incentivize portfolio capital inflows and consequently exchange rate appreciations (Titelman Kardonsky and Pérez Caldentey, 2015).

In order to correctly evaluate the exchange rate regime and its possible impact on RER behavior, Calvo and Reinhart (2002) recommend accounting for the important distortions between the de jure exchange regime (the official one that is reported to the IMF) and the de facto regime which reflects the true policy pursued by the country. Table 2 shows the percent distribution of fixed de facto exchange rate regime over the total sample period. The table illustrates that in the 70's, 76% of the time, the sample countries on average were under a fixed exchange rate regime. This percentage declines over the following decades, reaching 54% in the present century. However, the behavior is different within subgroups. For example, in the 90's in Central America there was an intense increase in trade and financial ties with the United States, including the formal dollarization of some economies. There, the dominant regime continues to be various types of fixation.

Table2: **Exchange Rate Regime Distribution(1970-2014)**

	Percentage of fix regimes				
	70 - 79	80 - 89	90 - 00	2001 - 14	70 - 14
Argentina	0.40	0.10	0.82	0.50	0.47
Bolivia	0.70	0.50	1.00	0.86	0.78
Brazil	0.60	0.40	0.27	0.00	0.29
Chile	0.20	0.60	0.64	0.00	0.33
Colombia	1.00	1.00	0.45	0.07	0.58
Ecuador	0.80	0.40	0.36	1.00	0.67
Paraguay	1.00	0.60	0.73	0.07	0.56
Peru	0.60	0.30	0.55	0.29	0.42
Uruguay	0.50	0.60	0.91	0.07	0.49
Venezuela	1.00	0.60	0.36	0.50	0.60
South America *(10 countries)*	**0.68**	**0.51**	**0.61**	**0.34**	**0.52**
Costa Rica	0.80	0.60	1.00	0.64	0.76
Dominican Rep.	1.00	0.60	0.55	0.57	0.67
Guatemala	1.00	0.70	0.55	0.79	0.76
Honduras	1.00	1.00	0.55	1.00	0.89
Jamaica	0.50	0.60	0.36	0.86	0.60
Mexico	0.80	0.50	0.45	0.07	0.42
Nicaragua	0.90	0.40	0.73	1.00	0.78
Panama	1.00	1.00	1.00	1.00	1.00
Trinidad & Tobago	0.40	0.80	0.73	1.00	0.76
Central America *(9 countries)*	**0.82**	**0.69**	**0.66**	**0.77**	**0.74**
Latin America *(19 countries)*	**0.75**	**0.59**	**0.63**	**0.54**	**0.62**

Notes: Author's calculation.

3. METHODOLOGICAL AND DATACONSIDERATIONS

In this section, we present the panel unit root and cointegration techniques involved in our analysis. Given our relatively short time series(T = 45), estimating the long-run behavior of real exchange rates through non-stationary panel methods, rather than single equation models, yields substantial benefits.

First, by pooling the data in the cross-section dimension (*N*), panel unit root and cointegration tests gain power and outperform their conventional time series counterparts[5] (*i.e.* Dickey-Fuller or Engle-Granger tests). Moreover, panel data provide efficient estimators for cointegration vectors[6] which are superconsistent and converge at rate $T\sqrt{N}$, while in the 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 into account the heterogeneity across different members of the panel. This allows us to test for the presence of a unit root and cointegrating relationship 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 individual countries.

3.1. Test of Panel Cointegration

Adding the cross-section dimension in testing for cointegration increases the power of test statistics just as with detecting unit roots. Several authors have recently proposed alternative methodologies for testing cointegration in a panel data context. 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)[7]. Kao (1999) and Pedroni (1999, 2004) proposed residuals based test for panel cointegration. The tests proposed by Kao (1999) are ADF type, similar to the classical approach adopted by Engle 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. However, 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[8]. Pedroni considers the following cointegration model with *k* regressors for a panel:

$$y_{i,t} = \gamma\, z_{i,t} + \beta_{1,i}\, x_{1,i,t} + \beta_{2,i,t} + \cdots + \beta_{k,i}\, x_{k,i,t} + e_{i,t} \qquad (7)$$

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:

$$\hat{e}_{i,t} = \rho_i\, \hat{e}_{i,t-1} + v_{i,t} \qquad (8)$$

[5] 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.
[6] 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).
[7] 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.
[8] Endogeneity of the regressors is also allowed by Pedroni's tests, which contrats with Kao's (1999) approach where homogeneity and exogeneity are imposed.

Where $v_{i,t}$ are assumed to be identically, independently distributed (i.i.d.) across i and t with $E(v_{i,t})=0$, E(v2) = σ2 < ∞, and E($v_{i,t}$, $v_{j,t}$) = 0 for all i with ¡= j. Pedroni (1999, 2004) considered seven tests (noted i,tl χ_{NT}) based on the residual from the regression (8). Four are based on pooling data along the within dimension (panel-v, 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 $H_a : \rho_i = \rho < 1$ 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:

$$\frac{\chi_{NT} - \mu\sqrt{N}}{\sqrt{v}} \Rightarrow N(0,1) \tag{9}$$

Where χ_{NT} is the appropriately statistic and μ and v are respectively the mean and the variance of χ_{NT}[9].

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 statistic when testing for cointegration.

3.2. 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 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

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

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 $H_a : \rho_i = \beta_a \neq \beta_0$; 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:

$$y_{i,t} = \alpha_i + x'_{i,t}\beta + u_{i,t}$$

$$x_{i,t} = x_{i,t-1} + \varepsilon_{i,t} \qquad (10)$$

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

$$\Omega_i = \begin{bmatrix} \Omega_{ui} & \Omega_{\mu\varepsilon i} \\ \Omega_{\varepsilon ui} & \Omega_{\varepsilon i} \end{bmatrix} = \Omega_i^0 + \Gamma_i + \Gamma'_i$$

$$= \begin{bmatrix} \Omega_{ui}^0 & \Omega_{uzi}^0 \\ \Omega_{eui}^0 & \Omega_{ei}^0 \end{bmatrix} + \begin{bmatrix} \Gamma_{ui} & \Gamma_{u\varepsilon i} \\ \Gamma_{\varepsilon ui} & \Gamma_{\varepsilon i} \end{bmatrix} + \begin{bmatrix} \Gamma'_{ui} & \Gamma_{u\varepsilon i} \\ \Gamma_{\varepsilon ui} & \Gamma'_{\varepsilon i} \end{bmatrix}$$

Where Ω_0 is the contemporaneous covariance and Γ_i is a weighted sum of auto-covariances. Ω_u refers to the long-run variance of the residual $u_{i,t}$, $\Omega_{\varepsilon i}$ is the (*k*×*k*) long-run covariance among the $\varepsilon_{i,t}$ and Ωεui is a (k × 1) vector that gives the long-run covariance between $u_{i,t}$ and $\varepsilon_{i,t}$. Note that the $\Omega_{\varepsilon ui}$ 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:

$$\hat{\beta}_{FM} = N^{-1}\sum_{i=t}^{N}\left(\sum_{t=1}^{T}\left(x_{i,t} - \bar{x}_i\right)\left(x_{i,t} - \bar{x}_i\right)'\right)^{-1}\left(\sum_{t=1}^{T}\left(x_{i,t} - \bar{x}_i\right)y^*_{i,t} - T\hat{\gamma}_i\right) \qquad (11)$$

Where

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

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 $\bar{x}_i = N^{-1}\sum_{T=1}^{N} x_{i,t}$). Under the assumption of cross sectional independence (*i.e.*E$\left|\zeta_{i,t},\zeta'_{i,t}\right|$=0 for all i ≠ j), Pedroni (2000) showed that the group-mean FMOLS is asymptotically unbiased and its t-statistic is standard normal:

$$T\sqrt{N}\left(\hat{\beta}_{FM} - \beta\right) \Rightarrow N(0,v)$$

$$t_{\hat{\beta}FM} \Rightarrow N(0,1) \qquad (12)$$

Where v depends of $\bar{x}_i, \bar{y}_i$ and of the dimension of $x_{i,t}$,*k*[10]. 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 (12), $\hat{\beta}_{FM}$ 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 i_{th} member of the panel as $\hat{\beta} = N^{-1}\sum_{i=1}^{N}\hat{\beta}_{FM,i}$ 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 $\Delta x_{i,t}$ (*i.e.* $\sum_{j=-q}^{q}\Delta x_{i,t+q}$) as additional regressors in (12). 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.

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

3.3. The Data: Source and Construction

Our panel is based on annual 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, Nicaragua, Panama, Paraguay, Peru, Trinidad and Tobago, Uruguay and Venezuela. The sample covers the period 1970-2014 (T = 45).

The real exchange rate (in logarithm) is usually defined as:

$$\log q = \log e + \log p - \log p^{*} \qquad (13)$$

Where e is the nominal exchange rate and p and p^{*} are respectively the national and foreign total

[10] When k = 1 and $\bar{x}_i = \bar{y}_i = 0, v = 2,$, if and/or $\bar{y}_i \neq 0,\ v = 0$

price indices[11]. Assuming that $\log p$ and $\log p^*$ can be split into traded and non-traded prices as follow:

$$\log p = (1-\alpha)\log p_T + \alpha \log p_N \tag{14}$$

$$\log p^* = (1-\alpha)\log p_T^* + \alpha \log p_N^* \tag{15}$$

Where $\log p_T^*$ and $\log p_N^*$ represent the price of foreign traded and non-traded goods respectively and α being the share of the non-traded sector in GDP at home and abroad. In fact, the real exchange rate log q can be decomposed in two components:

$$\log q = (\log e + \log p_T - \log p_T^*) + \ \alpha[(\log p_N - \log p_T) - (\log p_N^* - \log p_T^*)] \tag{16}$$

Then, assuming that: [i] the law of one price is valid for the traded goods and [ii] the domestic country is too small to have an influence on foreign partners' relative prices, the first term in (16) vanishes and $(\log p_N^* - \log p_T^*)$ is given. Thus, the real exchange rate varies only with the domestic relative price of non-traded goods:

$$\log q = \log p_N - \log p_T \tag{17}$$

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[12], such as the most of Latin America countries.

In empirical terms, we construct the real effective exchange rate of the country *i* at time t is as the geometrical weighted average of the real bilateral exchange rates vis-a-vis its ten main trading partners:

$$q_{i,t} = \prod_{j=1}^{10}\left(\frac{P_{i,t}}{E_{ij,t}\ P_{j,t}}\right)^{wj,t} \tag{18}$$

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) and $w_{j,t}$ 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' exports and imports 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 sectoral productivity, the GDP per capita relative to trading

[11] According to (13) 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.

[12] Here we impose a common value for α in both countries. If $\alpha \neq \alpha^*$ equation (16) 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 $(\log p_N - \log p_T)$.

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 $q_{i,t}$. 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[13]. However 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[14]; 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 anBénassy-Quéré *et al.*, 2004 for recent applications)[15]. However, like for sectoral 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).

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

$$f_{c\,i,t} = \frac{(PIA_{i,t} + OIA_{i,t}) - (PIL_{i,t} + OIL_{i,t})}{GDP_{i,t}} \tag{19}$$

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) and $GDP_{i,t}$ 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:

$$fdi_{i,t} = \frac{DIE_{i,t} - DIA_{i,t}}{GDP_{i,t}} \tag{20}$$

Where $DIE_{i,t}$ is direct investment in the country *i* and $DIA_{i,t}$ is 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, except for Haiti, Honduras and Nicaragua. So, in order to complete the sample, we updated the database using the following construction[16]:

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[13] 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).

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

[15] De Loach (2001) argues that the logarithms of CPI and PPI are composed of traded and non-traded goods such that $CPI = \alpha_{PN} + (1-\alpha)_{PT}$ and $PPI = \beta_{PN} + (1-\beta)_{PT}$, where PN and PT are respectively the price of non traded and traded goods. The relative price of non traded goods can be expressed as: $PN - PT = (\alpha - \beta)^{-1}(CPI - PPI)$. Assuming that α>β, an increase in (CPI − PPI) leads to an increase in the relative price of non traded goods.

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

$$\Delta NFA_{i,t} = CA_{i,t} + KA_{i,t} \tag{21}$$

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).

The terms of trade, $tot_{i,t}$ is defined as the ratio of country's export price index to its import price index (source: CEPAL), and 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).

To identify the *de facto* exchange rate regime ($reg_{i,t}$) for our sample of countries[17] we apply the methodology proposed by Coudert and Dubert (2005). 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).

In order to identify potential collinearity among the explanatory variables included in the analysis we perform a correlation analysis (results are exposed in Appendix D). We focus on the bivariate correlation between: a) real exchange rate and its fundamentals -column 2-; and b) the different long-run determinants of the real exchange rate -columns 3-.

Panel bivariate correlations, noted $\overline{R}(x,y)$, are computed as follows: in a first time we compute individual bivariate correlations, noted r_{xy}, between variables *x* and *y* for country *i*. Then we average the absolute value of $\overline{R}(x,y)$ across the *N* dimension that is $\overline{R}(x,y) = N^{-1}\sum_{i=1}^{N}\left|r_i^{xy}\right|$. Note that this construction of $\overline{R}(x,y)$ 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 Appendix C, 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 $(\overline{R}(prod, fci) = 0.37)$. In different reports, IMF (2005, 2014, 2015) stressed the importance of external financial flows as an important element in fueling Latin America growth, and particularly since the early 90's where a large part of capital inflows was composed of portfolio investments. Terms of trade and the productivity variable are strongly correlated $(\overline{R}(prod, tot) = 0.46)$. 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 investments 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

[17] See appendix C for further details

inflows[30]. According to our correlation analysis, a large part of determinants of the real exchange rate show a moderate correlation each other.

4. EMPIRICAL RESULTS

4.1. Testing PPP

Before determining the long-run fundamentals of real exchange rates in Latin America, we apply panel unit root tests in the original and demeaned series (results of panel unit root tests on original[18], first differences and demeaned series are given in Tables1 and 2 of Appendix B).

Applied to original series, The LM-Hadri test clearly rejects the null hypothesis of stationarity for all variables. The IPS and MW tests indicate that the null hypothesis of non-stationarity cannot 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 (fdii,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 inflows variable. This result implies a common feature in the evolution of financial flows and term of trade to Latin America zone[19]. Then 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 series in first differences, the LM-Hadri test strongly reject the null hypothesis of non-stationarity. In fact, for all variables, the LM statistic is below the right tail 5 % critical value of a standard distribution (1.64) in the panel for all series at the 5 % significance level. According to Hlouskova and Wagner's (2005) it is highly efficient to find unit root. Since the IPS and MW also shows a similar result, we are able to accept the null hypothesis of stationarity of the series in first differences.

As a result, we can conclude that the real exchange rate in Latin America follows random walks, implying that deviations from PPP can be permanent. So, the equilibrium real exchange rate is not an immutable value but it varies through time because of fundamentals' fluctuations. This evidence, that implying that PPP does not hold in this region, is in line with the findings of Edwards and Savastano (1999) for Latin America; Dumrongrittikul and Anderson (2015) -who check the invalidity of PPP (using similar panel UR test) for Asian countries- and Nouira and Sekkat (2015), who found also that ERER in developing countries follows is non stationary.

4.2. Heterogeneous cointegration: total sample results

Once we reject PPP approach for our panel, we are able to profit the cointegration proprieties of the series in order to obtain a conditional mean for the ERER. We first apply Pedroni's (2004) 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

[18] 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 a 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).

[19] 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.

can estimate efficiently the influence of each determinant on the real exchange rate. We consider different specifications of the long-run real exchange rate model, where the key issue is to show alternative treatments for the external financial channel in the form of capital flows and NFA. The models are the following:

Model1: $q_{i,t} = \alpha_i+\beta(g_{i,t},prod_{i,t},tot_{i,t},open_{i,t},reg_{i,t})+\varepsilon_{i,t}$

Model2: $q_{i,t} = \alpha_i+\beta(g_{i,t},prod_{i,t},tot_{i,t},open_{i,t},reg_{i,t},fci_{i,t}) +\varepsilon_{i,t}$

Model3: $q_{i,t} = \alpha_i+\beta(g_{i,t},prod_{i,t},tot_{i,t},open_{i,t},reg_{i,t},fdi_{i,t})+\varepsilon_{i,t}$

Model4: $q_{i,t} = \alpha_i+\beta(g_{i,t},prod_{i,t},tot_{i,t},open_{i,t},reg_{i,t},fci_{i,t},fdi_{i,t})+\varepsilon_{i,t}$

Model5: $q_{i,t} = \alpha_i+\beta(g_{i,t},prod_{i,t},tot_{i,t},open_{i,t},reg_{i,t},nfa_{i,t})+\varepsilon_{i,t}$

Where α_i is the fixed effects, $\beta=(\beta_1,\beta_2,\beta_3,...)$ is the vector of coefficients, $\varepsilon_{i,t}$ the residual and α_i a term that captures countries specificities[20].

We consider the model 1 as the framework. It includes as covariates the ratio of government spending to GDP ($g_{i,t}$), a productivity effect ($prod_{i,t}$), the terms of trade ($tot_{i,t}$), the degree of openness ($open_{i,t}$) and the *de facto* exchange rate regime ($reg_{i,t}$). However, Athukorala and Rajapatirana (2003), Nouira and Sekkat (2015) and Combes *et al* (2012) showed that the composition of capital flows matters in determining their influence on the real exchange rate. Thus, in models 2, 3 and 4 we examine 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 to get her in the regression (model4). Moreover, several studies[21] found that the long-run net foreign assets improvements are associated with real exchange rate appreciations. In fact, in order to check the presence of a *transfer effect*, in model 5 the variable net foreign assets ($nfa_{i,t}$) is added.

According to our real exchange rate's definition (an increase in $q_{i,t}$ implies an appreciation of the domestic currency), we would expect $\beta_1>0$, $\beta_2>0$, $\beta_3\gtreqless 0$, $\beta_4<0$ and $\beta_5>0$. Assuming that government spending falls more on non-traded goods, an increase in public consumption will raise 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

[20]It is needed in regressions because real exchange rates, productivity differentials and terms of trade are expressed as indexes and hence are not comparable in levels across countries.
[21]See Calderón (2002), Alberola (2003), Bénassy-Quéré et al. (2004), Lane and Milesi-Ferretti (2004), Aguirre and Calderón (2005) and Dufrenot and Yehoue (2005).

appreciation (*i.e.*β_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.

Table 3 reports estimates of models 1-5 based on the group-mean FMOLS estimator. The last row of table 3 reports the group parametric-t test -$\overline{Z}^{*}_{t\,N,T}$- 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. Hence, there is strong support for a cointegration relationship between real exchange rate and its determinants. Moreover, regarding the coefficients of the proposed long-run determinants of the ERER, all our coefficient estimates are highly statistically significant at the 5% significance level with the expected signs.

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 is mainly oriented to non-traded goods. Our estimates suggest that the elasticity of real exchange rate to government spending changes fluctuates around the 0.29 - 0.36 range, depending on the model specification (the biggest elasticity corresponds to the framework model, while the smallest with model 4). These estimations are fairly close to recent other studies of real exchange rate behavior in developing countries. For example, 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. Ricci *et al*. (2013) show that an increase in the government consumption to GDP ratio of 1 percentage point is associated with an appreciation of equilibrium RER of 3 percent. However, it is important to remark that our estimates remain relatively high in comparison to Drine and Rault (2003) study ($\hat{\beta}$= 0.10), who considered a group of 17 countries of Latin America over the period 1973-1996.

Concerning the Balassa-Samuelson effect, our estimates show that a permanent increase in productivity in the domestic country tend to appreciate the effective real exchange rate with its partners, with coefficients in a range between 0.18 (model 2)and 0.34 (model 3). In that case, the elasticity magnitude shows the biggest variation across models, with a standard deviation of a 0.06. The highest one corresponds to model 3, and the smallest with model 2.

As was explain in section 2, 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. However, 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 0.44-0.57, which means that the income effect is predominant on substitution effect. That result is in line with Jongwanich and Kohpaiboon (2013), who find a similar result for a sample of nine emerging Asian countries. Regarding the magnitude of the net income effect across specifications of our ERER equation, although the variation is smaller than the Balassa-Samuelson one, the standard deviation in that case (0.05) is bigger than for the rest covariates.

In addition, there is a strong relation between real exchange rates and the degree of openness. The coefficient $\hat{\beta}$ appeared strongly significant and negative in all models (the estimated elasticity for openness is quite stable among the regressions, with a mean coefficient of −0.43, and a range between -0.45 -model 3- and -0.4 -model 2-). This result indicates that more integration to world trade leads to real depreciations, and it is in line with Elbadawi (1994), Dufrenot and Yehou (2005) and Ricchi *et al*. (2013) results.

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 who state that the exchange rate regime is neutral regarding the evolution and the volatility of real macroeconomic variables (see Flood and Rose, 1995, and Obstfeld and Rogoff, 2000). In spite of that, Carrera and Vuletin (2011) also found that exchange regimes are non-neutral on the behavior of ERER. In regards to the magnitude coefficients, estimations show that they are quite stable across models (*i.e.* the standard deviation is only 0.02), registering the minimum elasticity for model 4 and the maximum one for the framework model.

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 model 4 result, the real exchange rate appreciates about 3.4% following a permanent 1 % rise in FDI flows as share of GDP (magnitude that reduced to 2.92 when capital inflows is not consider as a fundamental of ERER), whereas an increase of 1% in portfolio investments inflows (as share of GDP) leads to a real exchange appreciation of 1.2%. These results differ from those estimated by Athukorala and Rajapatirana (2003), who found elasticities 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, and (or) because they focused only on a restrict group of six countries in Latin America[22] over the period 1985-2000. Also, our estimations are quite different -in magnitude- from Jongwanich and Kohpaiboon (2013) findings. In spite they also found positive relations between real exchange rate and FDI and capital financial flows for the Asian economy, their estimations show similar values for both elasticities. They explain the difference between theirs results and Athukorala and Rajapatirana (2003) ones by changing in direct investment patterns from tradable sector to service sector in the last years.

Finally, 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 foreign assets coefficient estimate is 0.194, 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.

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

Table 3: **Long-run determinants of real exchange rates: group-mean FMOLS results**

All countries (N = 19 and T = 45)					
	Model 1	Model 2	Model 3	Model 4	Model 5
$g_{i,t}$	**0.356**	**0.337**	**0.326**	**0.298**	**0.322**
	[7.28062]	*[6.83287]*	*[6.71185]*	*[6.28938]*	*[6.72120]*
$prod_{i,t}$	**0.303**	**0.180**	**0.346**	**0.270**	**0.226**
	[7.60239]	*[6.79583]*	*[8.42054]*	*[7.70650]*	*[7.89807]*
$tot_{i,t}$	**0.529**	**0.568**	**0.446**	**0.457**	**0.503**
	[7.69697]	*[7.87214]*	*[7.70678]*	*[7.59970]*	*[6.63239]*
$open_{i,t}$	**-0.414**	**-0.403**	**-0.450**	**-0.420**	**-0.448**
	[-10.17247]	*[-10.09525]*	*[-9.00420]*	*[-9.07039]*	*[-13.13961]*
$reg_{i,t}$	**0.242**	**0.205**	**0.208**	**0.180**	**0.189**
	[11.11616]	*[10.45263]*	*[10.91275]*	*[10.35565]*	*[10.36736]*
$fci_{i,t}$		**1.261**		**1.208**	
		[3.62317]		*[3.30827]*	
$fdi_{i,t}$			**2.918**	**3.398**	
			[2.19458]	*[2.87419]*	
$nfa_{i,t}$					**0.194**
					[9.46115]
$Z_{tN,T}$	**-4.61466**	**-2.32774**	**-5.69112**	**-2.39853**	**-8.49691**

Notes: Heteroskedasticity and autocorrelation consistent t-statistics are reported in parentheses.

4.3. Heterogeneous cointegration: regional samples results

Due to specific characteristics in trade, productive structure and external financial linkages (especially with the dollar area), it is possible that South America (SA) and Caribbean and Central America (CCA) have got different behaviors on their ERER fundamentals. In order to check it, in this section we run the FMOLS regressions for each 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).

Results of UR tests for sub samples series are exposed in Table 3 of Appendix C. It shows that for all the different specification of variables for any group split, the results of UR tests are similar to the ones for the whole sample tests. Furthermore, the Pedroni's cointegration test indicates that there is evidence for cointegration hypothesis in the five models for South American countries, since only models 2, 3 and 5 support the hypothesis of cointegration between the real exchange rate and fundamentals in the case of Caribbean and Central America group.

Table 4 reports group-mean FMOLS estimation results for SA and CCA. In the SA case, most of the coefficients are significant at 5% in the five models. The only exception is Government expenditure in models 1 to 4 that are significant at 10%. Particularly, in model 5 all coefficients are significant at 1%. In the CCA group, all the variables in models 1 to 5 are significant at 1% except foreign direct investment. Finally, it is important to remark that all elasticities, in both groups and for all models, maintain the same signs from the total sample estimation.

By analyzing results comparing with the total sample estimation, it could be note that an increase in government spending has got a bigger impact on ERER (in all models) for CCA group that SA one. The same result is found for the Balassa-Samuelson effect and net foreign assets elasticity. In contrast, term of trade elasticity is always bigger for South American countries sample (except in model 3), as well as *de facto* exchange rate regime, indicating that fixed regimes tend to appreciate more the real exchange rate in South America than in Caribbean and Central America zone. Finally, results of model 5 confirm the existence of the transfer effect in both regions, although an improving in net external positions are associated with a higher appreciating real exchange rates in Central and Caribbean countries.

Table 4: **Long-run determinants of real exchange rates: group-mean FMOLS results**

	South America (N = 10 and T = 45)					**Caribbean and Central America (N = 9 and T = 45)**				
	Model 1	Model 2	Model 3	Model 4	Model 5	Model 1	Model 2	Model 3	Model 4	Model 5
$g_{i,t}$	**0.271**	**0.317**	**0.184**	**0.211**	**0.313**	**0.450**	**0.358**	**0.484**	**0.395**	**0.332**
	[2.14005]	*[2.22544]*	*[1.80305]*	*[1.84765]*	*[3.20256]*	*[8.32269]*	*[7.58212]*	*[7.85151]*	*[7.19067]*	*[6.38988]*
$prod_{i,t}$	**0.283**	**0.148**	**0.329**	**0.214**	**0.072**	**0.325**	**0.217**	**0.364**	**0.332**	**0.397**
	[4.86969]	*[4.40919]*	*[4.92081]*	*[4.51997]*	*[3.69662]*	*[5.91291]*	*[5.22642]*	*[7.04777]*	*[6.43282]*	*[7.57906]*
$tot_{i,t}$	**0.596**	**0.683**	**0.419**	**0.505**	**0.567**	0.455	0.440	0.475	0.403	0.432
	[5.42729]	*[6.12369]*	*[5.04324]*	*[5.62754]*	*[4.53967]*	*[5.46257]*	*[4.98302]*	*[5.88164]*	*[5.11016]*	*[4.85141]*
$open_{i,t}$	**-0.505**	**-0.475**	**-0.527**	**-0.488**	**-0.578**	**-0.313**	**-0.323**	**-0.365**	**-0.345**	**-0.303**
	[-6.60357]	*[-6.70565]*	*[-6.91634]*	*[-6.89590]*	*[-10.46111]*	*[-7.81948]*	*[-7.59968]*	*[-5.79234]*	*[-5.91004]*	*[-8.06443]*
$reg_{i,t}$	**0.298**	**0.260**	**0.257**	**0.222**	**0.254**	**0.178**	**0.144**	**0.154**	**0.133**	**0.116**
	[8.43628]	*[7.9297]*	*[7.97212]*	*[7.48091]*	*[8.40418]*	*[7.25879]*	*[6.82869]*	*[7.4525]*	*[7.16084]*	*[6.20464]*
$fci_{i,t}$		**1.710**		**1.743**			0.762		0.613	
		[3.12865]		*[3.42289]*			*[1.96645]*		*[1.19876]*	
$fdi_{i,t}$			**4.468**	**5.107**				**1.195**	**1.499**	
			[2.47765]	*[3.21785]*				*[0.57698]*	*[0.78419]*	
$nfa_{i,t}$					**0.078**					**0.323**
					[5.19926]					*[8.26622]*
$Z_{tN,T}$	**-3.371**	**-2.489**	**-4.814**	**-3.235**	**-5.641**	**-4.77286**	**-0.75865**	**-5.1056**	**-0.07529**	**-6.39958**

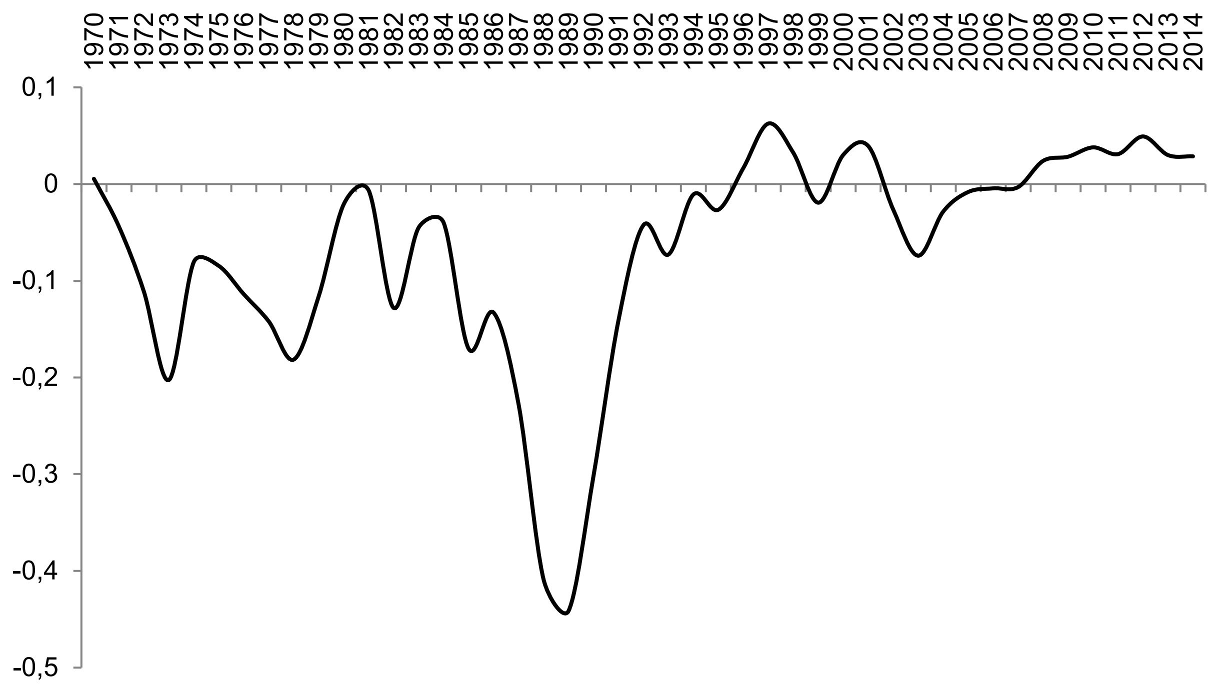
Notes: Heteroskedasticity and autocorrelation consistent t-statistics are reported in parentheses.

5. UNDERSTANDING THEGREAT MODERATION OF MISALIGNMENTS

RER misalignments are defined as the deviation of real exchange rates from an equilibrium level over an extended period of time. As such, RER misalignments can be distinguished from RER volatility, which is defined as frequent but not persistent fluctuations from equilibrium real exchange rates. Appendix E shows the RER misalignments computed according to Coudert and Dubert (2005).

Figure 1 shows the long-run average misalignment of Latin American over the last four and a half decades. The figure could be divided into two different periods. The first period spanning 1970-1995 is characterized by a sustained undervaluation of the RER and a high volatility. The volatility that marked the70´s was replaced by relatively stable levels of misalignment for the first half of the 80´s followed by massive devaluations due to the occurrence of simultaneous currency crises in a number of countries. The second half of the sample displays a more equilibrated pattern, with a persistent and increasing overvaluation of RER (only interrupted by some devaluations in the Southern Cone at the end of the nineties) and much less volatility.

Figure1: **Latin America and Caribbean average misalignment. 1970-2014**



Notes: Author's calculation.

The mechanisms by which Latin American countries brought about this successful outcome of low levels of misalignment and low volatility of RER are of great interest to policymakers. A plausible hypothesis is that the move toward flexible exchange rate regimes resulting from the adoption of inflation targeting since the mid-90s is the key variable in explaining lower volatility and the reduction of average misalignments. In fact, although some small and open economies -mostly in Central America and the Caribbean -applied pure fixed regimes, most of the region opted for *de jure* flexible exchange regimes. In many ways, there was a clear division with larger economies opting for flexible exchange rate regimes and smaller one selecting fixed regimes.

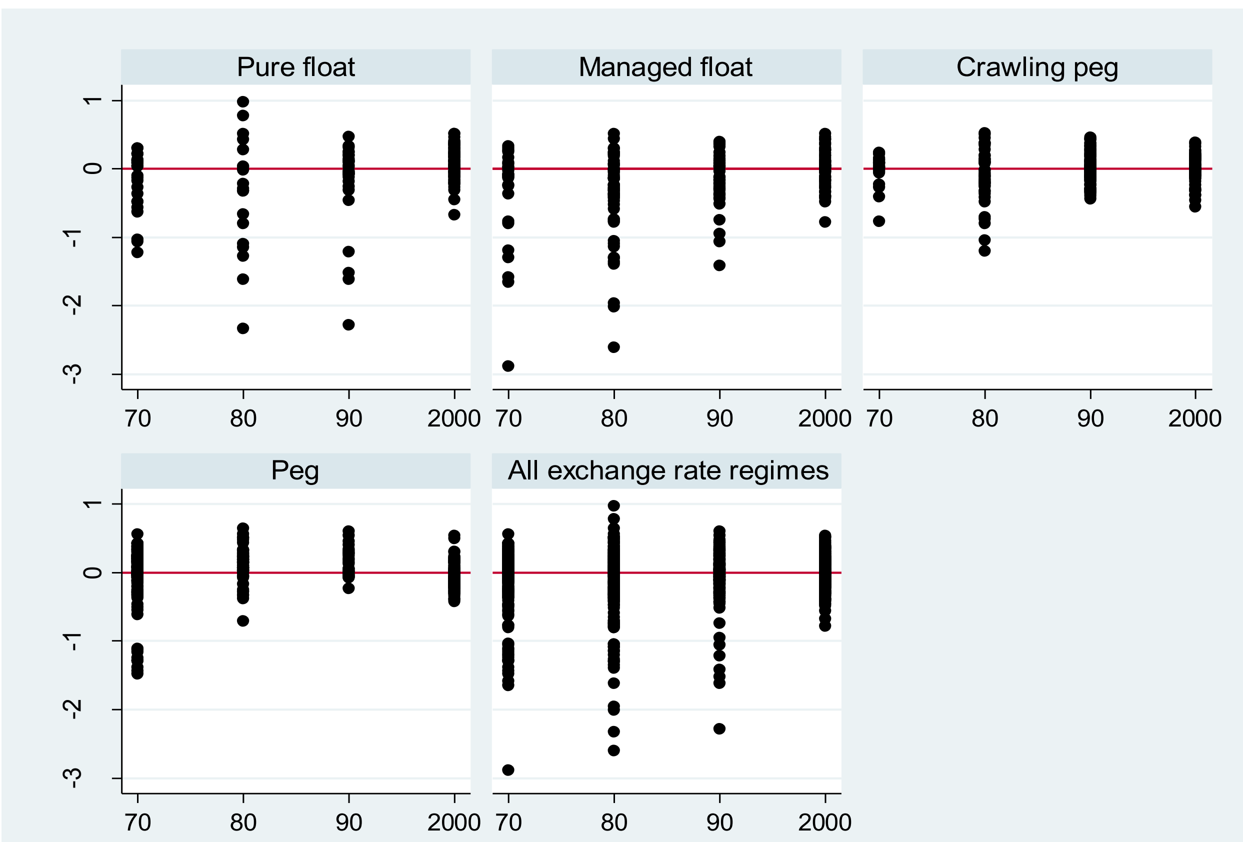
The evidence in the literature is mixed regarding the relationship between exchange rate regime and misalignment. Combes *et al*. (2012) argue that flexible exchange rate regimes are one of the main macroeconomic tools for countries to manage capital flow impacts on RER. Specifically, by introducing short-term volatility, a flexible regime discourages the very short term capital flows associated with important RER misalignments and short term volatility. On the other hand, some studies have found a lower pass-through of exchange rate movements to prices; creating a higher correspondence between nominal and real exchange rate variations which facilitate macroeconomic adjustment of external shocks (Khundrakpam, 2007 and McCarthy, 1999).

Coudert *et al*. (2013) confirm the possibility of the existence of large ERER misalignments, irrespective of the nominal real exchange rate. However, just as our estimations concluded, Holtemoller and Mallick (2013) find that a fixed regime induces more misalignment. Finally, Nouira and Sekkat (2015) find that intermediate regimes cause more volatility.

Part of the explanation of this improved adjustment capacity of flexible exchange rate regimes could be based on a new stylized fact that is the reduced pass-through from exchange rate movements to prices. Some studies have found a reduced pass-through Deveraux and Yetman, (2010) among other find that for low depreciations like 10% inflation rises around 1.5% in the first year and close to 2% in the long run. This low response creates a higher correspondence between nominal and real exchange rate variations which facilitate macroeconomic adjustment of external shocks (Khundrakpam, 2007 and McCarthy, 1999).However, Forbes, (2015) find that pass-through may be larger for global shocks than for country specific ones, he also finds that inflationary pressures of exchange rate movements induced by demand shocks may be smaller than those generated by supply shocks. In fact, it is also possible the presence of non-linearity's in the form that small appreciation or devaluations causes a smaller price movement than the higher ones.

In order to understand the (potential) role of an exchange rate regime on RER misalignment, it is useful to classify misalignment data by each exchange rate regime and by different sub-sample periods. Figure 2 shows that in the Latin American economies, a significant overall reduction in misalignments was achieved during the last sub-period, the 2000`s.This is valid not only on average over the whole sample, but also for each one of the exchange rate regimes considered. This suggests a kind of *great moderation* of exchange rate misalignment and volatility since the start of the new century. Also noteworthy is the trend of overvaluation for all the regimes. One stylized fact visible in the data is that pure float and managed float regimes are more prone to undervaluation, while pegged regimes tend to be more appreciated throughout the sample.

Figure2: **Distribution of Latin American and Caribbean countries misalignment by exchange rate regime and decade. 1970-2014**



Notes: Author's calculation.

Similarly, Table 5 shows that for each exchange rate regime, the average misalignment tends to decrease in the new century. In the same way, volatility (approximated by the coefficient of variation) significantly decreases in the new century, independent of the exchange rate regime considered.

Table 5: **Misalignment by exchange rate regime and decade, 1970-2014**

Average and coefficient of variation

ER Regime	Statistics	Full period	Decades			
			70	80	90	2000
Pure float	Mean	-0.0762054	-0.2138260	-0.3733157	-0.1955	0.09121622
	CV	0.50863407	0.44814837	0.85067208	0.63297605	0.22922384
	Freq dist	17%	12%	10%	16%	26%
Managed float	Mean	-0.2565337	-0.40548	-0.4796551	-0.1699285	-0.0107358
	CV	0.63020855	0.77688884	0.82941313	0.37718258	0.27063296
	Freq dist	21%	13%	31%	22%	19%
Crawling peg	Mean	-0.0196327	-0.02090909	-0.1007954	-0.0051123	0.01285916
	CV	0.25372264	0.22589755	0.38879178	0.21021447	0.19422441
	Freq dist	26%	12%	23%	47%	25%
Peg	Mean	0.01121639	-0.0402333	0.12378261	0.13406897	-0.0480459
	CV	0.31914695	0.41934488	0.23407986	0.2054999	0.19271909
	Freq dist	36%	63%	36%	15%	31%
All exchange rate regimes	Mean	-0.0676081	-0.1070684	-0.1621421	-0.0503631	0.01022456
	CV	0.43559185	0.48298714	0.62979056	0.36314515	0.22426394
	Freq dist	100%	100%	100%	100%	100%

Notes: Author's calculation.

This data serves as preliminary evidence that the adoption of flexible exchange rate regimes in Latin American economies appears not to be the primary determinant of the reductions in average misalignment and volatility. So, what explains this phenomenon in the most recent sub-period of the sample?

Edwards (1989) holds that misalignments are determined by both: a) the observed real exchange rate, which depends on the exchange rate regime and economic policy adopted by the government; and b) the equilibrium real exchange rate, which depends on economic fundamentals. Taking this into account, is it possible that the RER behavior observed in Latin America was the result of favorable external conditions rather than the adoption of flexible exchange regimes? Was the flexible exchange regime a key instrument for moderating real exchange rate dynamics because of inherent characteristics of the regime itself or, indeed, is a particular combination of fundamentals the cause of the RER appreciation, facilitating the reduction in misalignment?

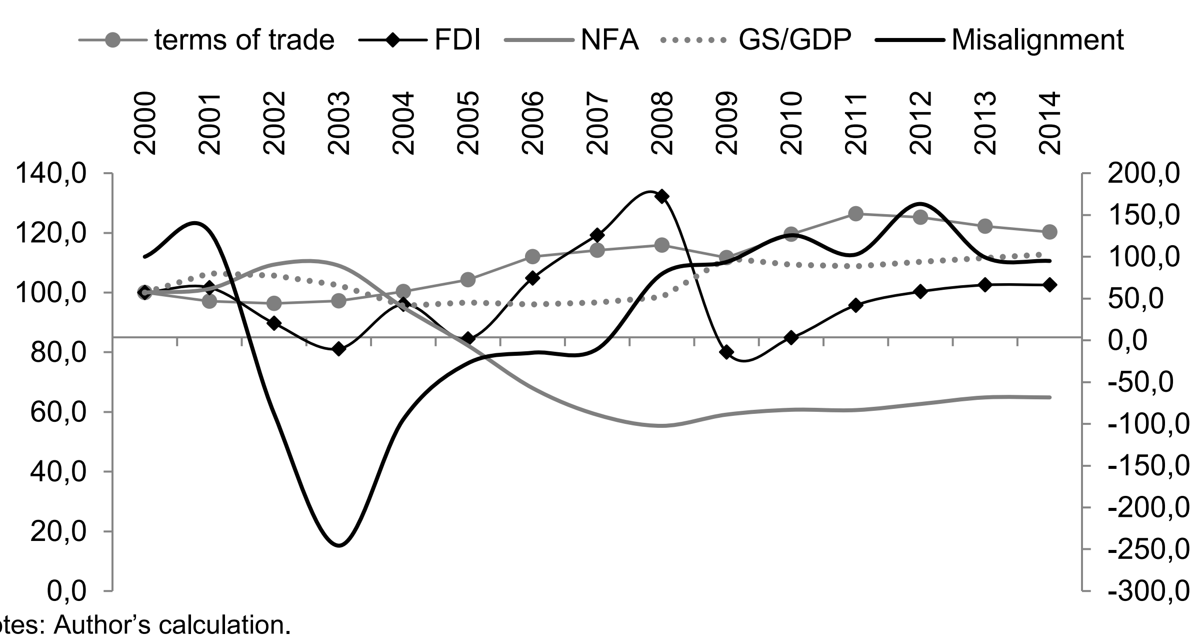
A possible explanation is that external conditions transmitted by economic fundamentals were responsible for more stable RER behavior. In fact, during the period beginning in the new century we can observed a positive development in the set of equilibrium real exchange rate fundamentals associated with external conditions which showed very favorable dynamics for Latin America. This was especially the case for terms of trade, capital flows and public expenditure. While further analysis is beyond the scope of this paper, this is a good starting point for future research.

As demonstrated in the previous subsection, terms of trade is one of the most robust covariates explaining equilibrium RER dynamics and, therefore, misalignments. Even though, there was an important heterogeneity among countries, the average terms of trade of the region significantly increased since the beginning of the new century, especially from 2002 to 2013), reducing regional exposure to volatility and devaluation pressures (see figure 3). The countries of the Pacific Basin plus Venezuela have seen the highest increases. Brazil, Argentina and Uruguay observed more moderate increases, while Mexico, Central America and the Caribbean have seen small increases or even reductions, because of oil and food dependency. Simultaneously, the region has received strong capital inflows in different forms. A substantial part was FDI, mainly oriented at the production of commodities and related service production given the high prices and the strong demand from China and the rest of Asia (CEPAL, 2015). As theory predicts, it is highly probable that this has improved the productivity of the tradable sector in the region, becoming, *via* Balassa Samuelson effect, another source of pressure for a real appreciation of the ERER. In this sense, Bastourre, Carrera and Ibarlucía (2012) demonstrate that for net export countries, there isa positive feedback from commodity prices to capital flows, amplifying current account volatility.

Financial capital inflows to the private and public sector were also very substantial in almost all countries given the increased appetite for risk after the implementation of non-conventional monetary policies in the advanced countries in the aftermath of the international financial crisis of 2007-9. Concomitantly, most of the countries in the region implemented a continuous process of financial liberalization of the capital account during the period that facilitated both capital inflows and outflows. For this reason, both gross and net flows must be considered. The period shows an increase in gross flows (*i.e.* inflows plus outflows) in a context of capital account deregulation that encouraged residents to internationally diversify their financial assets. Consequently, the NFA shows that the region as a whole has moved from a net creditor position at the beginning of the period to a net debtor position.

Additionally, given the tax structure of several countries and the simultaneous easing of access to cheaper external financing, governments were allowed to expand public expenditure on non-tradable goods generating another source of appreciation pressures on ERER. In 2014, government spending was 18% higher than in the local minimum point of 2004.

Figure 3: **Evolution of average misalignment (*right axis*) and external RER fundamentals (left axis). 2000-2014, index 2010=100.**



Notes: Author's calculation.

We can offer a precursory answer to our question about the role of the flexible exchange rate regime in the stylized fact of lower and less volatile misalignments since the turn of the century. More likely, the appreciation pressure was driven by these underlying fundamentals, which created an environment more compatible with the successful adoption of flexible exchange rate regimes. Therefore, the great moderation of RER misalignment may not be the logical result of the generalized adoption of flexible exchange rates regimes but the result of a combination of massive benign appreciating pressures from fundamentals.

In other word, it is possible to hypothesize the existence of a potential asymmetric behavior of exchange rate regimes that makes easier the successes of any exchange rate regime in front of appreciating pressures than in the opposite case of depreciating pressures coming from fundamentals. Certainly, this hypothesis deserves a formal evaluation in future research.

6. CONCLUSION

Since the 70's, Latin American countries have increased their interaction with the rest of the world and among themselves. This region has been subject to several shocks from domestic and external sources and has also introduced important changes in their policy regimes and exchange rate systems. Almost all the countries have gradually opened their economies to trade and capital flows, and most of them have adopted, also gradually, inflation targeting regimes and more flexible exchange rates systems. Additionally, since the period that followed the international crisis started in 2007-8, the region has increased its participation in financial and goods and services markets. That is why the understanding of real exchange rate dynamics is a key issue for the region.

In this context, the main objective of this paper is to show which factors determine ERER in nineteen Latin American countries over the 1970-2014 period. Using panel cointegration techniques, we estimate a model of real exchange rate based on the most accepted fundamentals used as explanatory variables in the ERER literature: the Balassa-Samuelson effect, government spending, terms of trade, the country openness to international trade, foreign capital inflows (differentiating financial flows and direct investments) and the net foreign assets position. In addition, we also include the de facto nominal exchange regime as a relevant factor influencing the evolution of real exchanges rates in Latin America. To identify the de facto regime, we follow the methodology proposed by Coudert and Dubert (2005). Moreover, forthe main set of results based on the whole sample, we divide our countries sample in two subgroups, South America (ten countries) and Caribbean and Central America (nine countries). We run the same long-run regressions for each subgroup which allows us to have a robustness check. 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 strong evidence that, over the long run, Latin American real exchange rates are non-stationary, implying that PPP does not hold in this region. This result is in line with Edwards and Savastano (1999) who claims that different studies focusing on real exchange rates in Latin America do not support the PPP hypothesis and Nouira and Sekkar (2015) who find the same result for a sample of developing countries. Thus, real shocks seem to have a permanent effect on real exchange rate paths.

Second, we identified six real factors that 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 fundamentals, which are statistically and economically significant. Particularly, we find that a higher government spending to GDP ratio, an increase in productivity differential, a positive terms of trade shock, a surge in foreign capital flows (aggregated or separated in FDI and financial flows) and a higher net foreign assets position appreciate the real exchange rate in Latin America. In contrast, an increase in trade openness leads to depreciation in the real exchange rate. In fact, we confirm that the most common explanatory variables in the literature continue to be highly relevant for explaining ERER, even in the recent post-crisis period.

Third, we find that the de facto exchange rate regime has a strong influence on exchange rates in Latin America: rigid regimes (peg or crawling peg) exercise an appreciating pressure on the real exchange rate. This stylized fact has been also documented by the IMF (2005) in examining the exchange rate stabilization plans applied in Latin American at the beginning of the 90's. This finding shows the non-neutrality of exchange rate regime regarding its effects on real exchange rates whatever the credibility level of the fixed regime. The tendency of fixed regimes to appreciate the real exchange rate can be seen as an adverse legacy of the recurrent reliance on exchange rate based stabilization in Latin America.

Fourth, there is evidence that real exchange rate behaviors are quite similar between South America and Central America and the Caribbean. Regarding the magnitude of FMOLS coefficients, there are some differences between the two regions. Particularly, fixed regimes and trade restriction policies are associated with greater appreciations in South America than in the Caribbean and Central America. This seems to be the expected result because small countries have a long tradition of pegging their currency to the dollar and an open trade policy.

Fifth, we present the evolution of misalignments over time for the nineteen countries. Periods of appreciation tend to be more persistent than depreciation periods. For the period in the sample corresponding to the 21st century, it is possible to see a lower degree of a positive overvaluation with a lower volatility. We show some evidence that this behavior is mainly explained by a combination of appreciation pressures from fundamentals such as term of trade, capital inflows, productivity in the tradable sector and government spending; other than to the generalization of flexible exchange rate regime as the complement of inflation targeting schemes. In other words, RER management is more fluid when countries face appreciating pressures rather than when face depreciating ones.

Finally, the appreciating pressures of the fundamentals on the ERER, however, should be managed with care since a large number of currency crises experienced by Latin American countries were preceded by huge and persistent overvaluations (Kaminsky *et al*., 1997 and Kaminsky and Reinhart, 1998). This finding has particular importance and strong policy implications for countries with a *de facto* fixed regime. In fact, our nominal regime classification shows that in 2014, eleven countries maintained a *de facto* fixed regime (peg or crawling peg)[39].

For these countries but also for the ones with a flexible exchange rate regime, the key question is how they will deal with possible pressures for massive devaluation resulting from a simultaneous change in fundamentals (*i.e.* deterioration in terms of trade, capital flows, productivity and government spending) which could be induced by a change in international monetary and financial conditions. Up to now, flexible regimes have softly accompanied the appreciation pressures during the 2000´s. The question is how symmetric the responses are to appreciating vis-à-vis depreciating pressures in the current institutional framework of inflation targeting regimes.

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8. APPENDICES

A. Fundamentals included in BEER macroeconomic approach. Theoretical models.

Year	Title and authors	ERER Fundamentals included
2007	Ravn, M. O., Schmitt-Grohé, S., & Uribe, M. (2007). *Explaining the effects of government spending shocks on consumption and the real exchange rate* (No. w13328). National Bureau of Economic Research.	Anticipated government spending
2010	Choudhri, E. U., &Schembri, L. L. (2010). Productivity, the terms of trade, and the real exchange rate: Balassa–Samuelson hypothesis revisited. Review of International Economics, 18(5), 924-936	B-S effect adjusted by product differentiation (within the tradable and nontradable sectors) with the number of firms and term of trade
2012	Christopoulos, D. K., Gente, K., & León-Ledesma, M. A. (2012). Net foreign assets, productivity and real exchange rates in constrained economies. European Economic Review, 56(3), 295-316.	B-S effect, net foreign assets and capital market access
2012	Chatterjee, M. S., &Mursagulov, M. A. (2012). *Fiscal policy and the real exchange rate* (No. 12-52). International Monetary Fund.	B-S effect adjusted by Government spending on public infrastructure
2013	Bleaney, M., & Tian, M. Net Foreign Assets, Real Exchange Rates and Net Exports Revisited (No. 13/04).	B-S effect, net assets denominated in domestic currency, net foreign assets and total trade
2014	Boero, G., Mavromatis, K., & Taylor, M. P. (2015). Real exchange rates and transition economies. Journal of International Money and Finance, 56, 23-35.	B-S effect, capital account as a proportion of GDP and relative domestic and foreign interest rate
2015	Dumrongrittikul, T., & Anderson, H. M. (2016). How do shocks to domestic factors affect real exchange rates of Asian developing countries? Journal of Development Economics, 119, 67-85.	B-S effect, real GDP, term of trade, public spending, openness coefficient, short-run inflation rate and short-run nominal domestic interest rate
2015	Bodart, V., Candelon, B., &Carpantier, J. F. (2015). Real exchanges rates, commodity prices and structural factors in developing countries. Journal of International Money and Finance, 51, 264-284.	Exchange rate regime, commercial openness coefficient, financial openness coefficient, export concentration indicator and commodities prices
2015	Berka, M. Devereux, M. and Engel, C., (2015) Real Exchange Rates and Sectoral Productivity in the Eurozone. http://www.ssc.wisc.edu/~cengel/WorkingPapers/BDE-15-3-11.pdf	B-S effect, sectoral TFP productivity, unit labor cost, exchange rate regimes

B. The de facto exchange rate regime classification:

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

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:

$$\ln e_t = \alpha + \gamma \, tiempo + \varepsilon_t$$

Where In 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, $\left|\hat{\beta}_j\right|$, is compared to an arbitrarily threshold τ. Following Coudert and Dubert (2005), τ is set to 2 % annually.

2. Comparing exchange rate variances: The second step consists in comparing the annual variance of changes in Δe_t or $\Delta \tilde{e}_t$ 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. s_i^2 denotes the empirical annual variance of Δe_t for the Latin America country *i* and s_B^2 is the average of annual variance of the *benchmark*. Assuming that annual variances follow normal distributions with theoretical variance σ_i^2 for the benchmark, then the ratio $((s_B^2/\sigma_B^2)/(s_i^2/\sigma_i^2)$ follows a Fisher F(n_B, n_i) where n_B and n_i designate degrees of freedom. Since s_B^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 ≤ (172.54)s_B^2, the exchange rate variance of the country *i* is considered as low. If s_i^2 ≥ (172.54)s_B^2, the variance of the country *i* is considered as high.

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 (ΔR). 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 $\tilde{s}_i^2$ ≤ 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	Type of regime
$\hat{\beta} > 0$	high	low	Pure float
$\hat{\beta} > 0$	high	high	Managed float
$\hat{\beta} > 0\, y\, \hat{\beta} < \tau$	low	-	Peg
$\hat{\beta} > 0\, y\, \hat{\beta} > \tau$	low	-	Crawling peg
$\hat{\beta} > 0\, y \lvert\beta\rvert > \tau$	-	low	Pure float
$\hat{\beta} > 0\, y \lvert\beta\rvert > \tau$	-	high	Managed float
$\hat{\beta} > 0\, y \lvert\beta\rvert < \tau$	low	-	Peg
$\hat{\beta} > 0\, y \lvert\beta\rvert < \tau$	high	low	Pure float
$\hat{\beta} > 0\, y \lvert\beta\rvert < \tau$	high	high	Managed float

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

C. Panel Unit Root Test

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 Wtbar 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).

Im, Pesaran and Shin (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, . . . ,*N* of the panel:

$$\Delta y_{i,t} = \gamma z_{i,t} + \rho_i y_{i,t-1} + \sum_{j=1}^{p_i} \theta_{i,j} \Delta y_{i,t-j} + \varepsilon_{i,t} \qquad \text{(A)}$$

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 E($\varepsilon_{i,t}$) = 0, E($\varepsilon_{i,t}^2$) = $\sigma_i^2 < \infty$ y ($\varepsilon_{i,t}, \varepsilon_{j,t}$)for all i *i ≠ 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:

$H_0: \quad p_i = 0, \quad \forall_i = 1, \ldots, N$

$H_a: \quad p_i < 0$, for at least one *i*.

Thus, under the alternative hypothesis IPS allow for p_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 (A) for each *i*, that is $\bar{t}_{NT} = N^{-1} \sum_{i=1}^{N} t_{i,T}$. Im *et al*. (2003) proposed the standardized statistic $\bar{t}_{NT}$:

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

Where E $\left[\left| t_{i,t} \right| \rho_i = 0 \right]$ and Var $\left[\left| t_{i,T} \right| \rho_i = 0 \right]$ are respectively the mean and the variance of $t_{i,T}$ [23].
Under the null hypothesis, the Wtbar 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:

$$P = -2 \sum_{i=1}^{N} \ln \left(p_i \right) \qquad \text{(C)}$$

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 p_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.

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[23] Simulated values of $E \left[\int_0^1 V(r)^2 dr \right]$ and Var $\left[\left| t_{i,T} \right| \rho_i = 0 \right]$ are provided by Im et al. (2003), table 3, page 66.

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:

$$y_{i,t} = z'_{i,t}\gamma + r_{i,t} + \varepsilon_{i,t} \qquad \text{(D)}$$

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 $E(\varepsilon_{i,t}) = 0,\ E(\varepsilon_{i,t}^2) = \sigma^2,\ E(u_{i,t}) = \sigma_u^2$. Using backward substitution, equation (D) can be written as:

$$y_{i,t} = z'_{i,t} + e_{i,t} \qquad \text{(E)}$$

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

$$Z_u = \frac{\sqrt{N}\left(LM - E\left[\int_0^1 V(r)^2\,dr\right]\right)}{\sqrt{Var\left[\int_0^1 V(r)^2\,dr\right]}} \qquad \text{(F)}$$

is exactly standard normal distributed, where V (r) is a standard Brownian motion[24].

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[25].

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[24] The moments $E\left[\int_0^1 V(r)^2\,dr\right]$ and $Var\left[\int_1^0 V(r)^2\,dr\right]$ are derived exactly, whereas for IPS a simulation is needed.

[25] 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.

Table 1: **Results of panel unit root tests on original and demeaned series**

	Original series			Demeaned series		
	IPS	MW	LM	IPS	MW	LM
q	-1.099	43.978	5.934	-0.766	40.571	6.267
g	-1.473	45.085	6.879	-1.477	47.212	7.283
prod	-0.180	36.224	7.745	0.340	36.188	8.796
fci	-1.446	46.688	3.334	-1.517	46.584	4.290
fdi	-3.135	68.261	6.599	-1.356	47.183	8.241
nfa	-0.140	39.120	7.566	-1.122	56.922	6.197
open	-1.193	55.388	6.737	-0.942	45.625	6.863
tot	-4.539	95.013	6.144	-4.978	98.744	6.779

Table 2: **Results of panel unit root tests on first difference and demeaned series**

	Original series			Demeaned series		
	IPS	MW	LM	IPS	MW	LM
Δ q	-7.876	143.342	-2.054	-8.378	153.419	-1.936
Δ g	-6.646	120.773	-1.071	-7.379	133.420	-1.072
Δ prod	-5.150	95.225	0.460	-5.195	98.461	1.195
Δ fci	-8.249	147.865	-2.238	-9.616	174.271	-2.313
Δ fdi	-8.445	151.968	-1.049	-7.963	143.091	-0.749
Δ nfa	-5.174	97.541	-0.650	-6.112	114.445	-1.887
Δ open	-9.085	164.533	-1.736	-8.402	149.387	-1.563
Δ tot	-9.754	179.637	-1.110	-10.427	195.054	-0.767

Table 3: **Results of panel unit root tests on original, first difference and demeaned series. South and Central American and Caribbean countries.**

	Original series			Demeaned series		
	IPS	MW	LM	IPS	MW	LM
			South America (N = 10)			
q	-1.460	26.805	4.167	-0.976	24.756	5.277
g	-0.936	21.729	6.879	-2.840	40.360	4.714
prod	0.396	16.070	4.904	1.059	15.566	6.145
fci	-1.566	25.726	1.751	-1.629	28.020	2.126
fdi	-1.023	22.441	6.630	-2.708	37.261	3.208
nfa	-0.485	22.982	4.134	0.368	17.212	4.956
open	-1.143	30.409	4.438	-0.899	23.833	5.067
tot	-3.145	47.874	5.735	-3.612	57.087	5.799
Δ q	-5.738	143.342	-2.054	-8.378	153.419	-1.936
Δ g	-6.646	120.733	-1.071	-7.379	133.420	-1.072
Δ prod	-5.150	95.225	0.460	-5.195	99.461	1.195
Δ fci	-8.249	147.865	-2.238	-9.616	174.271	-2.313
Δ fdi	-8.445	151.968	-1.049	-7.960	143.091	-2.749
Δ nfa	-5.174	97.541	-0.650	-6.112	114.445	-1.887
Δ open	-6.211	77.440	-0.977	-6.098	75.001	-0.530
Δ tot	-9.754	179.637	-1.110	-10.427	195.054	-0.767

	Caribbean and Central America (N = 10)					
q	-0.127	17.173	4.226	-0.103	16.363	4.135
g	-1.413	23.356	5.080	-1.191	24.401	4.136
prod	-0.626	20.154	6.026	-0.200	20.496	6.772
fci	-0.505	20.962	3.223	0.558	16.582	4.369
fdi	-1.431	28.626	6.496	-1.160	25.620	6.513
nfa	0.269	16.147	6.513	-1.278	28.508	3.845
open	-0.052	18.227	5.544	0.118	16.705	4.723
tot	-3.273	47.139	3.021	-2.629	43.083	4.188
Δ q	-5.412	70.142	-1.668	-6.698	88.179	-1.785
Δ g	-3.724	48.757	-0.656	-4.849	61.818	-1.187
Δ prod	-3.709	48.460	-0.139	-4.111	53.932	0.147
Δ fci	-5.734	73.974	-1.210	-6.093	79.611	-0.948
Δ fdi	-6.310	82.632	-0.444	-5.769	74.028	-0.102
Δ nfa	-3.765	49.320	-0.361	-4.256	55.121	-1.251
Δ open	-6.192	80.851	-1.341	-5.865	75.078	0.815
Δ tot	-6.906	91.905	-1.312	-7.502	101.770	-1.000

D. Panel bivariate correlations. 1970-2014

	q	*g*	*prod*	*tot*	*open*	*reg*	*fci*	*fdi*	*nfa*
q	1								
g	0.37	1							
prod	0.32	0.29	1						
tot	0.37	0.27	0.37	1					
open	0.38	0.37	0.40	0.33	1				
reg	0.34	0.33	0.31	0.28	0.27	1			
fci	0.31	0.18	0.25	0.22	0.26	0.24	1		
fdi	0.30	0.39	0.32	0.29	0.39	0.20	0.19	1	
nfa	0.31	0.25	0.45	0.35	0.32	0.26	0.27	0.31	1

E. Latin American countries misalignments

	Arg	Bol	Bra	Chi	Col	CR	Dom	Ecu	Gua	Hon	Jam	Mex	Nic	Pan	Par	Per	Tri	Uru	Ven
1970	-0,57	0,42	0,12	0,56	0,24	0,34	-0,05	0,32	-0,07	0,08	0,07	0,20	-1,30	0,23	-0,31	-0,33	0,26	-0,55	0,42
1971	-0,49	0,37	0,09	-0,10	0,15	0,32	-0,05	0,30	-0,12	0,03	0,10	0,17	-1,39	0,20	-0,38	-0,33	0,22	-0,29	0,37
1972	0,02	-0,14	0,07	-0,08	0,12	0,32	0,01	0,25	-0,14	0,05	0,03	0,16	-1,44	0,19	-0,47	-0,29	0,13	-1,22	0,31
1973	0,39	-0,03	0,06	-2,88	0,09	0,33	0,07	0,24	-0,11	0,00	0,02	0,16	-1,13	0,14	-0,61	-0,26	0,08	-0,63	0,24
1974	0,42	0,21	0,01	-0,80	0,01	0,25	0,08	0,23	-0,08	0,01	0,12	0,23	-1,11	0,14	-0,51	-0,20	0,08	-0,77	0,15
1975	-1,59	0,42	0,09	-0,25	-0,02	0,30	0,14	0,24	-0,05	0,00	0,18	0,26	-1,16	0,09	0,06	-0,12	-0,09	-0,24	0,11
1976	-1,30	0,37	0,04	0,16	-0,04	0,29	0,17	0,23	-0,02	-0,01	0,23	-0,37	-1,25	0,08	-0,04	-0,36	-0,14	-0,27	0,08
1977	-1,04	0,29	0,00	0,22	0,07	0,24	0,20	0,23	-0,03	-0,03	0,24	-0,18	-1,29	0,03	-0,17	-1,06	-0,21	-0,23	0,02
1978	-0,41	0,20	-0,06	0,21	0,03	0,20	0,15	0,19	-0,08	-0,08	-0,10	-0,11	-1,48	-0,02	-0,26	-1,20	-0,30	-0,27	-0,06
1979	0,13	0,06	-0,48	0,30	0,09	0,23	0,15	0,20	-0,04	-0,04	0,02	-0,03	-1,65	0,01	-0,02	-0,77	-0,35	0,03	-0,05
1980	0,45	0,24	-0,37	0,43	0,08	0,27	0,18	0,22	-0,06	0,02	0,15	0,10	-1,27	0,06	0,02	-0,74	-0,38	0,18	0,03
1981	0,06	0,50	-0,22	0,55	0,13	-1,36	0,19	0,33	0,06	0,10	0,23	-0,10	-0,71	0,14	0,29	-0,48	-0,33	0,37	0,14
1982	-1,30	-0,24	-0,16	0,44	0,14	-0,36	0,28	0,24	0,03	0,14	0,30	-1,30	-0,29	0,18	0,50	-0,71	-0,26	-0,27	0,22
1983	-1,14	0,28	-0,75	0,50	0,13	-0,15	0,31	0,21	0,02	0,19	-0,15	-0,53	-0,06	0,14	0,65	-0,77	-0,08	0,04	0,31
1984	-1,09	0,18	-0,81	0,51	0,09	-0,11	0,48	0,28	0,06	0,24	-0,24	-0,22	0,22	0,16	0,51	-1,05	0,06	-0,01	-0,02
1985	-0,30	-0,74	-1,05	0,11	-0,21	-0,17	-0,06	0,21	0,18	0,21	-0,14	-0,59	0,28	0,14	0,42	-1,09	-0,37	-0,15	0,06
1986	-0,33	0,34	-0,33	0,03	-0,36	-0,12	0,06	-0,04	-0,40	0,26	0,03	-1,21	0,77	0,18	0,03	-0,32	-0,28	-0,20	-0,66
1987	-0,78	0,38	-1,15	-0,02	-0,42	-0,27	-0,42	-0,22	-0,34	0,22	-0,02	-1,39	0,97	0,15	0,33	-0,80	-0,17	-0,06	-0,30
1988	-0,31	0,38	-1,96	0,08	-0,33	-0,18	-0,32	-0,39	-0,30	0,24	0,03	-0,14	-2,33	0,09	0,51	-2,60	-0,24	-0,03	0,06
1989	-4,60	0,52	-2,01	0,09	-0,33	-0,08	0,12	-0,25	-0,43	0,32	-0,01	-0,13	-1,62	0,08	0,30	-0,03	-0,05	0,20	-0,48
1990	0,09	0,06	-0,95	-0,10	-0,44	-0,18	-0,13	-0,23	-0,44	-0,52	-0,12	-0,06	-1,52	0,04	0,08	-0,74	0,05	-0,34	-0,29
1991	0,39	-0,01	-1,42	-0,08	-0,37	-0,31	0,10	-0,32	-0,14	-0,17	-1,06	0,07	0,33	0,01	0,15	0,31	0,12	-0,12	-0,17
1992	0,53	-0,05	-1,62	0,04	-0,19	-0,13	0,11	-0,26	-0,10	-0,16	-0,19	0,18	0,47	0,01	0,11	0,31	0,20	0,05	-0,12
1993	0,60	-0,08	-2,28	0,01	-0,10	-0,17	0,13	0,03	-0,11	-0,32	-0,46	0,24	0,42	-0,03	0,19	0,36	0,03	0,24	-0,08
1994	0,40	-0,34	0,06	-0,04	0,12	-0,15	0,19	0,04	0,03	-0,43	-0,16	-0,27	0,37	-0,05	-0,11	0,35	0,04	-0,04	-0,21
1995	0,29	-0,40	0,37	-0,05	0,11	-0,11	0,27	-0,06	0,03	-0,24	-0,23	-0,38	0,33	-0,06	-0,23	0,32	0,05	-0,12	-0,39
1996	0,27	-0,34	0,43	-0,03	0,23	-0,10	0,19	-0,09	0,09	-0,29	0,09	-0,06	0,27	-0,08	-0,24	0,29	0,02	-0,12	-0,22
1997	0,31	-0,31	0,45	0,06	0,14	-0,11	0,14	0,00	0,10	-0,11	0,11	0,09	0,22	-0,03	-0,29	0,33	0,00	-0,03	0,11
1998	0,34	-0,30	0,43	-0,02	0,11	-0,13	0,01	-0,16	0,04	-0,06	0,13	0,03	0,19	-0,08	-0,37	0,26	-0,04	0,02	0,22
1999	0,45	-0,29	0,24	-0,07	0,04	-0,14	-0,01	-1,21	-0,09	-0,03	0,04	0,19	0,17	-0,06	-0,31	0,25	0,02	0,17	0,26
2000	0,49	-0,30	0,27	-0,08	-0,06	-0,11	-0,03	-0,41	-0,05	-0,01	0,01	0,24	0,18	-0,09	-0,24	0,29	0,02	0,19	0,27
2001	0,54	-0,31	0,22	-0,16	-0,06	-0,08	-0,01	-0,05	-0,05	-0,01	0,01	0,30	0,15	-0,14	-0,41	0,35	0,04	0,15	0,26
2002	-0,07	-0,12	0,00	-0,09	-0,23	-0,10	-0,22	0,05	0,04	-0,04	-0,02	0,23	0,15	-0,11	-0,25	0,36	0,06	0,03	-0,18
2003	0,05	-0,33	0,20	-0,04	-0,22	-0,14	-0,78	0,01	0,03	-0,07	-0,18	0,17	0,12	-0,14	-0,19	0,31	0,01	-0,04	-0,16
2004	0,02	-0,39	0,29	-0,05	-0,05	-0,14	-0,04	-0,06	0,09	-0,10	-0,10	0,17	0,12	-0,19	-0,28	0,34	-0,02	0,11	-0,28
2005	0,04	-0,40	0,43	0,01	0,04	-0,11	-0,16	-0,08	0,15	-0,08	-0,01	0,23	0,14	-0,19	-0,31	0,32	0,00	0,19	-0,35
2006	0,04	-0,42	0,46	-0,12	0,04	-0,08	-0,12	-0,14	0,16	-0,09	-0,01	0,22	0,14	-0,22	-0,12	0,35	0,01	0,16	-0,35
2007	-0,01	-0,45	0,51	-0,14	0,11	0,00	-0,17	-0,23	0,16	-0,08	-0,03	0,21	0,15	-0,24	-0,13	0,35	0,01	0,22	-0,31
2008	0,14	-0,12	0,37	-0,40	0,02	-0,04	-0,19	-0,17	0,20	-0,02	0,00	0,04	0,23	-0,18	0,01	0,36	0,09	0,19	-0,07
2009	0,01	-0,32	0,51	-0,21	0,06	0,00	-0,27	-0,20	0,13	0,02	-0,07	0,11	0,18	-0,18	-0,07	0,37	0,10	0,30	0,08
2010	0,08	-0,40	0,51	-0,21	0,08	0,10	-0,31	-0,22	0,16	0,02	0,06	0,18	0,16	-0,21	-0,12	0,37	0,14	0,26	0,08
2011	0,15	-0,31	0,46	-0,43	0,11	0,11	-0,27	-0,22	0,21	0,06	0,14	0,07	0,22	-0,19	-0,05	0,39	0,15	0,27	-0,27
2012	0,18	-0,29	0,40	-0,33	0,18	0,12	-0,39	-0,23	0,20	0,04	0,09	0,16	0,20	-0,18	0,02	0,42	0,18	0,30	-0,13
2013	0,11	-0,18	0,35	-0,48	0,11	0,15	-0,46	-0,21	0,22	0,04	0,06	0,18	0,22	-0,15	0,04	0,38	0,20	0,27	-0,27
2014	0,17	-0,11	0,32	-0,68	-0,10	0,11	-0,56	-0,15	0,27	0,06	-0,02	0,11	0,18	-0,11	0,03	0,37	0,23	0,23	0,19