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Abstract

This paper analyses stability in real multilateral exchange rates in six leading Latin-American economies during the XXth century using a new data set. A univariate approach is complemented by an error-correction model including key fundamentals. Unit-root testing shows a very slow process of mean reversion – if any – in the series in levels; however, mean reversion is found after allowing for trends and structural breaks with half-life values ranges from 0.8 to 2.5 years. We also found reversion to a conditional mean defined by the co-integrating relationship, and that the equilibrium path is largely explained by fundamentals - especially terms of trade and trade openness. Exchange rate policy proved to have only a transitory effect in generating real depreciation.

JEL keywords: Real Exchange Rates, Purchasing Power Parity, Economic Development, Latin America
JEL codes: F41, N16, O11

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1. Introduction

Long-run quantitative research on countries of the periphery remains an exotic subject: both interesting and elusive. Interesting by virtue of these countries' distinct features, varied experiences and, somehow, from a combination of failure and success in their development efforts. Elusive as a consequence of data limitations. For such countries "periphery" denotes a double condition. On the one hand, it describes their subordinate position in the world economic system and, on the other, refers to a relatively marginal attention paid to them by economic studies, particularly those focusing on the long term.

The revival of research on the behaviour of real exchange rates (RER) over a long time horizon conforms to this situation. One key concern of such research is assessing the validity of the Purchasing Power Parity (PPP) as an equilibrium condition for the nominal exchange rate.[1] The PPP doctrine requires that real exchange rates exhibit mean reversion. That is, that the impact of shocks should be temporary, and that, in the absence of further disturbances, the RER should move back towards its mean value. Recent studies covering a time span of over a century or more have supported the PPP hypothesis by finding mean reversion in real exchange rate series. Mark Taylor (2003) provides a recent survey on this empirical work. Meanwhile, Alan Taylor (2002) extended the long-run analysis to a set of twenty countries and also found support for PPP – although after allowing for deterministic trends in a number of cases. This outcome is at odds with most of the evidence coming out of the analysis during the post Bretton Woods period dominated by the floating of the major currencies (Adler & Lehmann, 1983; Enders, 1988).

A commonly-used measure of the speed of mean reversion is the half-life of the process.[2] The prevailing consensus in the long-span and panel unit-root studies focusing on industrialised economies is that the half-life process of real exchange rates (in levels) ranges between 3 and 5 years (Rogoff, 1996; and Frankel & Rose, 1996).[3] However, this apparent slow speed of reversion can be partly caused by the presence of one or more structural breaks, implying the absence of a constant mean for the entire period. For instance, Lothian and Taylor (2004), after allowing for shifts in the equilibrium dollar-sterling real rate over two centuries, suggest that the half-life deviations from PPP may be as low as 2.5 years (compared to 6 years in the original series). Also, Hegwood and Papell (1998) found in their long-span study of six RER series (all from OECD economies) that reversion to the changing mean is much faster than reversion to a fixed mean. After accounting for structural breaks, they estimated half-life values of between 0.5 and 2.5 years.

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[1] In its absolute version, PPP states that the equilibrium nominal exchange rate between two countries will equal the ratio of the countries' price levels. In its relative version it states that the nominal exchange rate equilibrium will change according to the relative change in the countries' price levels.

[2] The number of years that it takes for deviations from equilibrium to subside permanently below 0.5 in response to a unit shock in the level of the series.

[3] This apparently slow speed of adjustment is the origin of the Rogoff's puzzle. Rogoff (1996) argues that the estimated speed of adjustment of real exchange rates is difficult to justify in terms of wage or price stickiness, or shocks related to real factors such as technology or tastes.

However, as Froot and Rogoff (1995) have pointed out, there is the possibility of a "survivors" bias in these recent findings. That is, that the propensity of the RER series to revert to the mean - and its speed - might be exaggerated by the inclusion of wealthy countries, or economies that already displayed high living standards at the start of the last century. Indeed, there are various ways in which the dynamics associated with economic development can affect the RER stability around a constant mean in the long term. The most prominent is via productivity differentials as captured by the Balassa-Samuelson model (hereinafter B-S).[4] Other fundamentals shaping the formation of relative prices are the terms of trade, openness to international trade, and capital flows (Neary, 1988; Montiel, 1999). Furthermore, the fact that most multi-country studies on developing countries relating to real exchange rates focus on the second half of the XXth century (e.g., Edwards, 1989; Wood, 1991; Joyce & Kamas, 2003),[5] compounded by limited availability of comparable data, make it difficult to assess the extent of mean reversion in countries on the periphery.

This paper assesses RER behaviour in the long term in six Latin American economies (LA6): Argentina, Brazil, Chile, Colombia, Mexico, and Venezuela during the period 1900-2000. First, it tests the case of mean reversion using a univariate (time-series) approach - allowing for structural breaks and trend behaviour, making it possible to compare the results with other studies that examine the PPP hypothesis with long-span data. Secondly, it makes a systematic attempt to incorporate fundamental variables - including relative productivities, terms of trade, trade openness, and government spending – with the potential to shape the equilibrium real exchange rates in LA6. To this aim, following Edwards (1989), Baffes et al. (1999), and Richaud et al. (2003), we estimate equilibrium RERs with an error-correction model.[6] The multivariate approach offers an alternative measure of RER reversion, this time to a conditional mean defined by a co-integrating equation.

Our analysis relies on two of the most commonly-used measures of the real exchange rate in empirical work: the purchasing power parity and the price ratio of the world tradable goods to domestic consumption goods - a limited proxy for the ratio of domestic tradable to non-tradable prices. The first measure focuses on intra-country utility comparisons and living standards; the second on macroeconomic equilibrium. There is no theoretical reason for the RER under both definitions to coincide or converge (Edwards, 1989), so whether in fact they tell a similar story or not, needs to be confirmed empirically.

The starting point of our task is the creation of a consistent dataset of multilateral or effective real exchange rates (REER) over the whole century. Although there are

[4] The main prediction is that if the labour productivity of a given country in producing tradable goods relative to their productivity in producing non-tradable goods grows faster than abroad, then the country's currency will appreciate in real terms. The seminal contributions are Balassa (1964) and Samuelson (1964).

[5] An exception is Taylor (2002) who includes Argentina, Brazil, and Mexico in his panel data of twenty countries covering more than 100 years, but he uses bilateral rates and the sample is dominated by developed countries. Also the lack of testing for structural breaks – and, if necessary, correcting for them – has the potential to undermine his assessment of stationarity.

[6] This approach is akin to the econometric analysis of the behaviour of RER (BEER) favoured by Clack and MacDonald (1998).

available long-term real exchange rate series for a number of Latin American countries, they are largely bilateral rates with the US and computed with US price indices to reflect world prices. The use of bilateral rates introduces a bias, particularly in the early and late decades of the last century characterised by a more geographically diversified trade structure. Meanwhile, relying on world prices that are not directly related to the country's trade flows misses out important terms of trade effects. In addition, we benefit from recently constructed comparable series across LA6 for key RER fundamentals available from OxLAD.[7]

The remainder of the paper is structured as follows. **Section 2** presents the two RER empirical measures and their corresponding multilateral indices. **Section 3** tests for non-stationarity and calculates the half-time values of mean reversion in the REER series. **Section 4** assesses the role of fundamental forces as well as the effectiveness of exchange rate policy. Finally, there is a section of conclusions. An **Appendix** includes a description of data sources and additional tables.

[7] The Oxford Latin American Economic History Database. Available at: http://oxlad.qeh.ox.ac.uk/

2. Measures and indices

We focus on two widely-used RER definitions in the empirical literature. First, one adopting the PPP proposition:

(1) $RER1 = E\dfrac{P_C^*}{P_C}$

Under this concept the exchange rate is a measure for relative purchasing power between two countries. As a similar price index is used in the numerator and the denominator, this is also referred to as the symmetric definition. E stands for the nominal exchange rate (domestic currency per unit of foreign currency). P_C denotes the consumer price index (CPI) at home, and P_C^* the CPI in the comparator country.

Our second definition is the price of world tradable goods relative to domestic consumption goods, both denominated in local currency.

(2) $RER2 = E\dfrac{P_T^*}{P_C}$; where P_T^* is the world price index of tradable goods.[8]

Here, the RER is the key equilibrating variable of a country's external accounts (Harberger, 2004). There are several options for P_T^* in (2). For example, to use the producer price index or the wholesale index of the main trading partner, as both indices tend to exclude retail sale services in their derivation. But one main drawback in using these proxies is that the same foreign price index is applied to all countries, without taking into account possible variations in the composition of their consumption baskets (Chinn, 2006). We prefer to use the border import price index facing a particular country. The use of this index results in a <u>real import exchange rate</u>, which can be interpreted as the number of basket of imports that can be bought with one consumption basket in a given country. In this way we also avoid some of the ambiguities that the use of a composite index of tradable goods (including both imports and exports) can create. For instance, a country benefiting from a commodity windfall should experience a real appreciation caused by the increased domestic spending (this assumes a fixed exchange rate regime). However, a RER measure that uses an index comprising all tradable goods can end up indicating a real depreciation if the commodity has sufficient weight to make the tradable index rise.

Based on the above definitions we construct two multilateral real exchange rate indices for each of the LA6 countries.[9] They are:

- REER1: calculated using CPIs for the main trading partners, as well as for the home country (or the GDP deflators).
- REER2: calculated using import unit value indices as proxies for the border import prices and CPIs to reflect general prices at home.

[8] Note that expressions (1) and (2) do not directly include the effect of protection in the home country in the numerator, which impact is only felt when it feeds through to the general price index.

[9] These series, together with methodological notes, are available at OxLAD.

An increase in the value of the real exchange rate indices indicates real depreciation, whereas a fall shows a real appreciation of the domestic currency. Both indices are calculated as geometric weighted averages of bilateral real exchange rates as in Chinn (2006) using at least six main trade partners (US, UK, Germany, France, Japan, and one or two Latin American economies).[10] Bilateral real rates with countries other than the US are derived as cross rates from the corresponding US dollar series.

The bilateral nominal exchange rate used to calculate our indices refers to the rate applicable to imports – or to averages, when multiple rates where in place. This is consistent with (2), which is based on the border price of imports, as well as being the more appropriate rate for utility comparison under the PPP concept. At different times, the LA6 countries adopted various exchange rate regimes, ranging from the gold and gold-exchange standard, fixed, multiple rates (with many variations), crawling pegs, and, more recently, floating rates. But, overall, exchange regimes with a fixed or a controlled rate for most transactions in the trade account were the norm.

The LA6 economies were under fixed, multiple or dual regimes during more than 70% of the time in the last century. Mexico and Venezuela lived under some sort of fixed type regime for about 70% of the time. Floating arrangements – with convertible currencies - were rarely implemented, featuring mostly in the 1990s with the adoption of inflation targeting.[11] The prevalence of fixed exchange rates means that domestic prices were the predominant adjustment channel to shocks. In this case, nominal rigidities in the labour or goods markets have the potential to exacerbate and prolong RER deviations from equilibrium and to slow the speed of mean reversion.

In those instances in which there was a unified rate, or a dual regime with an official (usually fixed) rate applied to most current account transactions and a “free” or market determined rate to convert capital transactions, the selection of the appropriate exchange rate has no complications. In the case of multiple exchange rates, when possible, we are working with an average of those rates applied to imports. When the data are available (e.g., Colombia), the average is weighted by the different trade flows associated with each rate. But in the face of data limitations, we are taking simple averages (e.g., Venezuela during 1961-1964 and 1983-1989). Multiple rates applied to exports are mostly ignored, as we are primarily concerned with imports or importable goods. However, they were a common feature in some of the LA6 countries during the period.

Despite the reputation of being an inflation-prone region, nearly 80% of the time inflation was under control (at low or moderate levels) or the countries underwent periods of deflation. This was the typical adjustment mechanism under the gold standard, and was common in the years prior to the Great Depression. According to our data on internal price, deflation was a feature in Argentina and Mexico during the

[10] Our sample of trading partners covers about 70% of imports up to the 1960s. Thereafter the average share falls to about 60%, reflecting a more diversified trade structure. However, despite the narrowing of the coverage, our multilateral series during the period 1980-2000 behave in line with those constructed by CEI, which encompass 23 trading partners covering at least 80% of trade flows.

[11] See Astorga (2007, Annex A) for an outline of exchange rate regimes and inflation in each country.

1920s and the early 1930s. Brazil and Colombia also experienced deflation during this period and earlier in the century. However, our sample includes countries such as Argentina and Brazil that endured recurrent episodes of hyperinflation and currency changes, vis-à-vis Venezuela and post-revolutionary Mexico, both characterised by a higher degree of price and currency stability.

3. Mean reversion under a univariate approach

The empirical literature dealing with the testing for long-run PPP usually relies on the augmented Dickey-Fuller test (ADF) for unit root in the process driving the real exchange rate. The rejection of the null hypothesis that a time series follows a random walk – the archetypal non-mean reverting process – is taken as evidence of mean reversion (Taylor, 2003). The application of the unit root tests - ADF and Phillips-Perron (PP) - to our REER series in levels (logs) cannot reject the null hypothesis in any of the cases, with the exception of Venezuela's REER2 at the 10% level (see Table A1). This indicates that the series in levels are non-mean reverting; a result that is at odds with the message from recent studies with long-span data. The failure to reject the unit-root hypothesis can be owing to the presence of trends or structural breaks in the series. Once a simple linear trend is included, the unit root test performed over the de-trended series (e.g., the residuals of the regression with a time trend and a constant) still fails to reject the null hypothesis in most cases at the 5% level. But we still need to allow for the possibility of structural breaks in the series, before concluding that they are non-stationary (Perron, 1989 & 1990).

3.1. Structural breaks

We adopt the methodology proposed by Zivot and Andrews (1992) - hereinafter Z&A - to determine endogenous structural breaks. A test statistic is calculated in each period, allowing for the possibility of breaks in the intercept, the slope, or both. If the minimum test value is below (i.e., higher in absolute value) a given critical level, it implies that the non-stationarity of the series is due to the presence of a structural break.[12] In order to identify potential multiple break points, we follow a sequential search on the lines suggested by Bai and Perron (1998). However, the power of the test declines sharply once the sample is subdivided. Table 1 presents the outcome of testing for structural breaks for REER1 and REER2. It informs about the years where a minimum is identified, the test values, and the type of break. When applying the Z&A test to the REER1 series the null hypothesis of unit root is rejected at the 10% level or lower in eleven cases, involving Argentina, Brazil, Colombia, Mexico, and Venezuela. The test values for Chile fail to reject the null at the 10% level of significance. However, the probability of not rejecting the unit root null when it is false is high, even with a long span of data (Lothian & Taylor, 1996).[13] For Chile's REER1, we decide to include a breakpoint in the mean in 1945, where the test is close to the critical value at 10% of significance. Regarding the REER2 series, the null is rejected at the 10% level or lower in twelve cases involving all countries.

[12] We use a program written by Trujillo (2006) in EViews to implement the Z&A test. To improve the power of the test we extend the sample period until 2005.

[13] Facing this level of type II error, the researcher needs to take a view on whether to make allowances for a break. This decision involves giving more weight to one of the two competing conceptions of the nature of the macroeconomic series. Under a view that favours non-stationarity, lasting shocks to the series can be interpreted as low probability realisations of a given data-generating process. On the other hand, under the belief that the world is more akin to stationary processes, a sudden and lasting move is seen as a structural break rather than a low-probability event.

Table 1: Indentification of break points

Zivot & Andrews testing procedure

period 1900-2000	REER1			REER2			TofT		Open		Rprodm	
	year	*t-stat* [1]	*type*	*year*	*t-stat*	*type*	*year*	*type*	*year*	*type*	*year*	*type*
Argentina	1952	5.30***	*m&t*	1952	5.39**	*m&t*			1940	*m*		
	1955/85	3.3/4.62**	*t*	1976/85	4.91/4.9**	*t*	1976	*t*	1972	*t*	1970	*m&t*
	1985	4.51***	*m*	1985	4.90**	*m*						
Brazil	1944	4.73**	*t*	1980	5.15**	*m&t*	1914	*m&t*	1942	*m*		
	1944	4.71	*m&t*	1992	5.05*	*t*	1950	*m*	1990	*t*	1958	*m&t*
Chile	1945	4.27	*m*	1973	4.96**	*m*	1930	*m*	1931	*m*		
	1947	4.49	*m&t*	1973	4.73**	*t*	1975	*m*	1973	*m&t*	1980	*t*
Colombia	1923	4.73*	*m*	1955	4.30	*m*	1915	*t*				
	1958	4.60**	*t*	1955	4.68**	*t*	1958	*m*	1956	*m*	1983	*t*
	1957	4.85*	*m&t*	1956	4.83*	*m&t*						
Mexico	1917	4.58*	*m*	1917 [2]	4.87**	*t*			1914 [3]	*m&t*		
	1917	4.25*	*t*	1977	4.22	*m&t*	1977	*m*	1988	*t*	1974	*t*
	1932	5.2**	*m&t*	1989	4.13*	*t*						
Venezuela	1961	4.24*	*t*	1961	4.30*	*t*	1973	*m*	1925	*m*	1982	*t*
	1961	4.33	*m&t*	1972	4.41	*m&t*			1961	*t*		

(1) all test values are negative ; (2) sample 1900-1988 ; sample 1900-1980.
m : break in the mean; *t* = break in the trend; m*&t* = break in the mean and trend.
*, **, & *** indicate that the critical value is rejected at the 10%, 5%, and 1% levels, respectively.
Critical values for *t* are from Perron (1990); and for *m* and *m&t* from Perron & Vogelsang (1993).

Table 1 also presents information about the presence of major discontinuities in key fundamentals, namely: terms of trade (*TofT*), the country's productivity in manufacturing relative to the US (*Rprodm*), and the gross trade ratio (*Open*) as a proxy for trade policy. Although the link between REER and the fundamentals will be explored with more rigour in Section 3, a simple comparison on the date breaks shows that a number breaks in REERs can be associated with an unusual close event – usually a precedent - in either the terms of trade, openness, or relative productivities. This is particularly the case in Brazil, Colombia, Mexico, and Venezuela. However, it is not the case that all breaks in fundamentals have a corresponding REER one (e.g., terms of trade circa 1930 in Chile), nor do all REER breaks have a matching discontinuity in any of the three fundamentals (e.g., Argentina in 1952). The former suggests that – notwithstanding statistical limitations – joint effects (including offsetting moves, for instance, in the form of capital flows to cushion an external shock) are important in triggering or avoiding a major break in the real exchange rate; whereas the latter points to the need to include additional variables or relevant events.

3.2. Testing for non-stationarity

We follow the two-stage procedure to test for non-stationarity in series with breaks proposed by Perron (2006). A description of this procedure follows.

Consider a trending series generated by $y_t = \mu + \beta t + u_t$, where:

(3) $\Delta u_t = C(L)e_t$

with e_t ~ i.i.d. (0, σ_e^2) and C(L) = $\sum_{j=0}^{\infty} c_j L^j$ such that $\sum_{j=0}^{\infty} j\,|\,c_j\,| < \infty$ and $c_0 = 1$.

Making allowances for a one-time change in the trend function, results in two versions of four different structures: 1) a change in level for a non-trending series; and for trending series, 2) a change in level, 3) a change in slope, and 4) a change in both level and slope. For each of the four cases, two versions allow for different transition effects. The first is labelled the "additive outlier model" (AOM) and specifies that the change to the new trend function occurs instantaneously. The second is the "innovational outlier model" where the change to the new trend function is gradual. We use AOMs because they allow for a joint change in the trend without a break, and in general offer a good description of our series. The AOM specifications for changes at a break date T_1 are as follows:

Model (AO-0) $y_t = \mu_1 + (\mu_2 - \mu_1)\ DU_t + u_t$

Model (AO-A) $y_t = \mu_1 + \beta t + (\mu_2 - \mu_1)\ DU_t + u_t$

Model (AO-B) $y_t = \mu_1 + \beta_1 t + (\beta_2 - \beta_1)\ DT_t^* + u_t$

Model (AO-C) $y_t = \mu_1 + \beta_1 t + (\mu_2 - \mu_1)\ DU_t + (\beta_2 - \beta_1)\ DT_t^* + u_t$

where $DU_t = 1$, $DT_t^* = t - T_1$ if $t > T_1$ and 0 otherwise, and u_t is specified by (3). Under the null hypothesis $C(1) \neq 0$, while under the alternative hypothesis, $C(1) = 0$.

The test procedure consists of a two-step approach. In the first step, the trend function of the series is estimated and removed from the original series via the following regressions for Model (AO-0) to (AO-C), respectively:

(AO-0) $y_t = \mu + \gamma DU_t + \tilde{y}_t$

(AO-A) $y_t = \mu + \beta t + \gamma DU_t + \tilde{y}_t$

(AO-B) $y_t = \mu + \beta t + \gamma DT_t^* + \tilde{y}_t$

(AO-C) $y_t = \mu + \beta t + \theta DU_t + \gamma DT_t^* + \tilde{y}_t$

where $\tilde{y}_t$ is accordingly defined as the de-trended series.

The next step differs according to whether or not the first involves DU_t, the dummy associated with a change in intercept. For Models (AO-0), (AO-A) and (AO-C), the test is based on the value of the t-statistic for testing that $\alpha = 1$ in the following autoregression:

(4) $\tilde{y}_t = \alpha\ \tilde{y}_{t-1} + \sum_{j=0}^{k} d_j D(T_1)_{t-j} + \sum_{i=1}^{k} a_i \Delta\tilde{y}_{t-i} + e_t$

The outcome of the implementation of Perron's procedure to the REER1 series is summarised in Table 2 (see Table A2 for results for REER2). Columns 1 and 2 give the model to be estimated and the breakpoint dates. When more than one break is included, we list them in chronological order. For instance, Brazil's model AO-B&C, 1914 & 1944 indicates the inclusion of a trend break in 1914 and a simultaneous break in the mean and trend in 1944. Columns 3 to 6 present key estimated parameters of the regressions: $\hat{\beta}$ is the estimate of the initial (pre-break) slope of the trend function; $\hat{\theta}_1$ and $\hat{\theta}_2$ are estimates of the change in the intercept of the trend function; and $\hat{\gamma}_1$ and $\hat{\gamma}_2$ are estimates of the change in the slope of the trend function (post first break). The trend coefficients indicate annual rates of growth. In the presence of a break in the trend, annual growth in the post-break period results from adding up the initial trend coefficient $\hat{\beta}$ and the $\hat{\gamma}$ s estimates. For example, in the case of Argentina REER1's AO-B (1952) model, the annual trend growth up to 1952 is estimated at -0.6%. But after 1952, it changes to 0.9% (-0.6% +1.8%).

The trend patterns during the first half of the century or so are mixed, with Argentina, Brazil and Venezuela (REER1) displaying a flat or downward trend (i.e., REER appreciation). However, the second half of the century is dominated by an upward trend (depreciation). The behaviour of the final decades confirms the findings of studies covering a much wider sample of developing countries after 1960 (e.g., Wood 1991). A depreciating trend in the RER is consistent with a lack of economic convergence, a phenomenon that would call for faster productivity growth in those countries attempting to catch up and real exchange rate appreciation over time (as has been the case in economies such as Japan).

Table 2: Non-stationarity test for REER1

	period 1900-2005		First stage						Second stage	
	(1) *model*	(2) *Tb*	(3) $\hat{\beta}$	(4a) $\hat{\theta}_1$	(4b) $\hat{\theta}_2$	(5a) $\hat{\gamma}_1$	(5b) $\hat{\gamma}_2$	(6) *SER*	(7) $\hat{\alpha}$	(8) $t_{\hat{\alpha}}$(ADF)
ARG	AO-B	1952	**-0.006**			**0.018**		*0.187*		
	AO-C	1952	0.000	**-0.358**		**0.017**		*0.164*	**0.62**	-3.25**
	AO-B	1955	**-0.005**			**0.019**		*0.182*		
	AO-C&B	1952 &1985	0.000	**-0.306**		**0.014**	0.011	*0.162*		
BRA	AO-B	1944	**0.012**			-0.002		*0.235*		
	AO-C	1944	**0.020**	**-0.409**		**-0.006**		*0.213*	**0.64**	-3.80***
	AO-B&C	1914&1944	**-0.020**		**-0.587**	**0.053**	**-0.020**	*0.193*	**0.59**	-3.79***
	AO-C	1953	**0.009**	0.120		0.000		*0.235*		
CHI	AO-A	1945	**0.020**	**-0.853**				*0.2*	**0.56**	-4.08***
	AO-C	1947	**0.012**	**-0.778**		**0.012**		*0.182*		
COL	AO-B	1957	**0.006**			**0.015**		*0.211*		
	AO-C	1957	0.002	**0.263**		**0.014**		*0.200*	**0.72**	-4.38***
MEX	AO-B	1917	**0.044**	**-0.037**				*0.238*		
	AO-C	1932	**0.024**	**0.327**	**-0.023**			*0.166*		
	AO-0&0	1917&1932	0.001	**0.502**	**0.442**			*0.149*	**0.59**	-4.0***
	AO-B&C	1917&1932	-0.001	**0.529**	**0.454**	0.002		*0.115*		
VEN	AO-B	1961	-0.002			**0.017**		*0.152*		
	AO-C	1961	**-0.005**	**0.322**		**0.014**		*0.130*	**0.70**	-3.18**

Columns 3-5: estimates in **bold** are significant at least at the 5% level. *SER* : standard error of the regression.
(*), (**), and (***) indicate that the critical value is rejected at the 10%, 5%, and 1% levels, respectively.

In column (6) we report the standard error of the regression (SER). Finally, columns 7 and 8 inform on the second stage of the procedure: $\hat{\alpha}$ is the estimate of the autoregressive coefficient of the regression equation with the residuals (see eq. 4), and $t_{\hat{\alpha}}$ is the associated t-statistic – we use the ADF test - for testing that α =1.[14] In deciding on which particular model to choose for each country (shaded values), we

[14] We used a lag parameter of four for the sum of the autoregressive first differences, and a lag of two for the trend dummies (only necessary in models A and C).

also assess the SER and favour those specifications that minimises it. But, if the difference is small, we opt for the specification with the minimal number of breaks.

3.3. Half-life results

All series in the second stage of the procedure are stationary around a trend after allowing for at least one-time structural change. This is evidence of mean reversion in the adjusted series. Table 3 presents information about the speed of mean reversion. First estimates for the autocorrelation coefficients of the de-trended series are presented with the corresponding half-time values for each of the REER series. The lower part includes information about the series after making allowances for changes in the slope and/or the intercept.[15]

Table 3: Speed of mean reversion

period 1900-2000		**Argentina**	**Brazil**	**Chile**	**Colombia**	**Mexico**	**Venezuela**
De-trended series							
REER1	alpha	0.86	**0.76**	0.90	**0.82**	0.91	0.88
	half-life	*4.6*	*2.5*	*6.5*	*3.5*	*7.0*	*5.4*
REER2	alpha	**0.64**	**0.80**	**0.80**	0.83	0.88	0.84
	half-life	*1.6*	*3.0*	*3.1*	*3.8*	*5.4*	*4.1*
Series de-trended and corrected for structural breaks							
REER1	alpha	**0.62**	**0.59**	**0.56**	**0.72**	**0.59**	**0.70**
	half-life	*1.4*	*1.3*	*1.2*	*2.1*	*1.3*	*1.9*
REER2	alpha	**0.43**	**0.54**	**0.72**	**0.68**	**0.46**	**0.75**
	half-life	*0.8*	*1.1*	*2.1*	*1.8*	*0.9*	*2.4*

Alpha: autocorrelation coeficient.The half life of the process is measured in years.
Values in ***bold*** *are series where the unit root is rejected at the 10% level or lower.*

The half-life of the de-trended REER1 series ranges from 2.5 years in Brazil to 7 years in Mexico in the case of REER1; and from 1.6 years in Argentina to 5.4 years in Mexico for the REER2 series. Note that these half-life ranges are similar to those found in more advanced economies, but with the crucial difference that the latter tend to be estimated without de-trending the original series. For instance, in their assessment of real exchange rate series of the dollar-sterling and franc-sterling spanning two centuries, Lothian & Taylor (1996) found half-life values of about 6 years and a little under 3 years respectively. Thus, the need to de-trend our series in order to remove non-stationarity can be interpreted as a required correction due to the development process (in the spirit of B-S) or, more generally, a sampling bias associated with countries that have failed to join the wealthy club at the end of the last century. After adjustments for breaks, the half-life values drop significantly, now with the REER1 values ranging from 1.2 years in Chile to 2.1 years in Colombia; and the REER2's from less than one year in Argentina to 2.4 years in Venezuela. This second

[15] To calculate the half-time of the process of mean reversion in the second case we use the autocorrelation estimates $\hat{\alpha}$ reported in Tables 2 and A2.

set of estimates is supposed to be "free" of the impact of major discontinuities (e.g., related to lasting changes in fundamentals).

The main limitation of the univariate analysis conducted so far is that it says little about the forces causing permanent shifts or trend behaviour in the real exchange rate. During the period under analysis, our economies went through a drastic process of structural transformation and industrialisation, and secular changes in their terms of trade (in most cases implying the replacement of the dominant export commodity). This is likely to have undermined the case for stability around a constant equilibrium position. Then, in order to account for deviations from the strict PPP prediction, there is the need to bring into the analysis the fundamental forces behind the formation of the RER. This comes next.

4. The role of fundamentals

As is customary in the modelling work on RER determination, our fundamentals combine supply and demand factors. On the demand side, we include the terms of trade (*TofT*) and two limited proxies for public spending on non-tradable goods: total government spending as a share of GDP (*Govexp*) and non-equipment investment as a share of GDP (*Inont*). The former proxy is too gross, including both spending on tradable and non-tradable goods; whereas the latter, although targeting non-tradable spending, is too narrow as it only includes the investment component. On the supply side we include the gross trade ratio (*Open*) as a proxy for trade policy and the country's productivity in manufacturing relative to the US (*Rprodm*) to account for the B-S effect.

Regarding trade openness, the most plausible outcome of higher barriers to international trade corresponds to a real appreciation in the long term - though the direction of the effect during the adjustment process is ambiguous (Edwards, 1989). On the other hand, trade liberalisation should lead to the opposite result via lower prices for import-competing goods.[16] However, note that *Open* can reflect other influences beyond trade policy with different implications for the expected sign of the link. For instance, a natural resource discovery is likely to result in a higher trade share – due to both the development of the new export activity and the increased import capacity – and a RER appreciation (as demonstrated in the Dutch-disease model).

Meanwhile, it is to be expected that relative productivity gains in the tradable sector result in real appreciation (and a negative coefficient for *Rprodm*).[17] This assumes that the tradable sector is the main source of productivity gains and that its prices are tied down by the world price level and the exchange rate.[18] However, regarding the second condition, during the ISI period the tradable sector in the LA6 tended to be sheltered from foreign competition – with the intensity of this effect varying among the countries.[19] As a result the "non-tradability" of manufacturing is likely to weaken the link between its productivity performance and the real exchange rate.

[16] But opening up a country to trade also increases demand for exports. And to the extent that the home country is a large producer of particular export goods, the export goods price can rise and will contribute to the rise in the consumer price level at home.

[17] A change in the intra-country relative productivity in non-tradable goods has an opposite effect, therefore undermining the effect emanating from the tradable sector. This may have played a role during the last three decades when the information technology revolution is likely to have boosted the productivity in US services relative to those in the LA6.

[18] Overall, the evidence on the role of the B-S effect is mixed (Sarno & Taylor, 2002).

[19] There were three main stages in terms of development strategies. First, an export-led growth episode came to an end with the collapse of commodity prices and capital flows in the 1930s. This led to a wave of devaluations and protectionism which shifted relative prices in favour of domestic industries and agricultural production. This transition gave way to a more conscious industrialisation strategy, the so-called import substitution industrialisation model. Then, gradually, policymakers started to give greater emphasis to export promotion. Balance of payments and fiscal difficulties in the 1970s made the opening of trade more pressing in many countries. The debt crisis and subsequent economic reforms pushed the economies further into a new period of export-led growth.

In theory, the impact of changes in the terms of trade on relative prices is ambiguous. But there is a strong presumption that a sustained terms of trade increase will result in a real appreciation (Edwards & Wijnbergen, 1987). This is the case if the income effect - which leads to higher demand for non-tradable goods - dominates the substitution effect associated with a decline in the relative cost of the imported intermediate goods used in the non-tradable sector. Finally, the effect of *Govexp* depends on whether it took place in the tradable or non-tradable goods sector. To the extent that spending is directed towards non-tradable goods, there should be a real appreciation. Meanwhile, *Inont* should mostly reflect spending on non-tradable goods and thus is expected to be negatively correlated with REER.

4.1. Error correction model

Our specification to estimate the equilibrium RER in terms of its fundamentals is based in a dynamic equation derived by Edwards (1989) capturing important features of developing countries. This is expressed in terms of an error correction specification (ECM) that includes both the middle to long-term explanatory power of the fundamentals together with the short term impact of macroeconomic policy, in particular, changes in the nominal exchange rate. The ECM is specified as follows:

(5) $\text{REER}^* = k\, X_t$

(6) $\Delta\text{REER}_t = b_0 + \alpha\,(\text{REER} - \text{REER}^*)_{t-1} + b_i\,\Delta X_t + c_i\,\Delta Z_t + e_t$

where REER_t is the real effective exchange rate (in logs), X_t a set of fundamentals (in logs), k is a vector of long-run multipliers of REER with respect to X_t. Meanwhile, Z_t stands for a set of macroeconomic policy variables, including the nominal exchange rate (ΔNER_t), the budget deficit as a share of GDP (*GB%*), and the exchange rate premium in first differences (ΔNERgap_t) - calculated as the spread between the parallel market rate and the official or controlled rate - to capture the impact of capital controls. Δ stands for the first-difference operator, an asterisk indicates equilibrium values and e_t is the error term.

Equation (5) describes the equilibrium or co-integration relationship between the REER and its fundamentals. Equation (6) includes the feedback or error-correction coefficient (α). It should be negative to indicate that REER deviations from its long-run determinants result in a move back towards equilibrium, and its size ($-1<\alpha<0$) measures the speed of this adjustment. The size of the feedback coefficient depends on the structural features of the economy and on institutional factors such as wage indexation rules and the exchange rate regime. This offers an alternative measure of mean reversion (or, in this case, reversion to the long-run equilibrium path);[20] although note that α only captures the adjustment associated with fundamentals forces, whereas our half-life parameters in the previous section also include the effect of changes in the nominal exchange rate. So, comparisons between both mean-reversion measures need to be interpreted with caution. Finally, the short-run impact coefficients (b_i) measure the effect of contemporaneous or lagged changes in the

[20] The expression to calculate the corresponding half-life values is $t = \log(0.5)/\log(1-\alpha)$; where t is the number of years and α is the feedback coefficient.

fundamentals on $\Delta REER_t$. Meanwhile, c_i measures the impact of changes in the macro-policy variables.

4.2. Estimation issues

The validity of inferences with the above specification depends on meeting two conditions: first, that the series in (5) are co-integrated; and, secondly, that the fundamentals are weakly exogenous for the parameters of interest k and α (Engle et al., 1983). Lack of co-integration means that the model can be estimated in first differences without the need of imposing a long-term relationship; whereas a failure of weak exogeneity indicates that the use of a single-equation framework is not the most appropriate and that variables are better treated as endogenous in a system setting.[21]

We perform the Johansen co-integration test (including structural break dummies as described in footnote 24). In all six countries the null hypothesis of no co-integration is rejected at least at the 5% significance level. In two cases (Brazil and Colombia) the test detects at most two co-integrating equations and only one in the remaining four. We restrict attention to the case of a single co-integration relationship, with the causality running from the fundamentals to the REER – as specified by (5). This has the advantage that the long-run parameters are clearly related to the economic theory that underpins the model (Baffes et al., 1999).

Regarding causality, given that we are dealing with small, open economies, we can assume that the terms of trade are exogenous, entering equation (5) contemporaneously.[22] The proxies for public spending on non-tradables are treated as policy variables entering with a lag, making them pre-determined. To address a potential reverse-causation problem between REER and *Open* we tested for weak exogeneity.[23] The outcome of this test (not shown) indicates that, for the purpose of estimating the ECM, *Open* can be treated as an exogenous variable. Because we use interpolated series the economically active population in the construction of *Rprodm* it is not possible to conduct weak exogeneity test in this case - not to include $\Delta Rprodm_t$ in (6). For the purpose of this exercise relative productivities are instrumented with the used of one-period lagged values (the lag also captures a likely delayed impact on relative prices).

We estimate the ECM using a two-stage procedure (Engle & Granger, 1987). First, we run the long-term regressions (5); then, the resulting residuals lagged one period

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[21] This would imply, for instance, using a structural vector error correction model (e.g., Gauthier & Tessier, 2002; Jang & Ogaki, 2004).

[22] Although some countries were able to have some influence in their terms of trade. For instance, Brazil with the coffee valorisation policies in the period 1880-1930 and Chile with saltpetre during a similar period (Abreu & Fernandes, 2005) and, more recently, Venezuela after 1973 via OPEC quota policies.

[23] The test estimates the feedback coefficient η in the following equation:

(6a) $\Delta \mathrm{Open}_t = b_0 + \eta\,(\mathrm{REER} - \mathrm{REER}^*)_{t-1} + b_i \Delta X_t + c_i \Delta Z_t + e_t$

If α is significant in (6) and η lacks significance in (6a), it can be concluded that *REER* does not contribute to the explanation of the parameters of the equation for $\Delta \mathrm{Open}_t$.

are incorporated into the dynamic model (6). We test for structural breaks in the co-integration errors by applying the Z&A method described earlier. In those cases where a break is detected we add dummy variables to the respective regression and test again for stationarity in the residuals – a requirement for the series to be co-integrated. Note that we are not automatically adding structural break dummies already identified in the REER series, as some of them are accounted for by the fundamentals.[24]

4.3. Regression outcome

Regressions are run for each country over the whole sample in order to determine the explanatory power of the fundamentals, estimate the error correction term, and assess the effectiveness of nominal depreciations. The results of the co-integrating country regressions for each of the REER measures are summarised in the upper part of Table 4. Since all variables are in natural logs, the coefficients can be interpreted as elasticities of the REER with respect to its fundamentals.

Table 4: Error Correction Model regressions

	period 1900-2000	**Argentina**		**Brazil**		**Chile**		**Colombia**		**Mexico**[1]		**Venezuela**	
	Dependent var.	REER1	REER2	REER1	REER2	REER1[2]	REER2[2]	REER1	REER2	REER1[2]	REER2	REER1	REER2
eq. (5) ; vars. in logs	**Rprodm** (*lagged*)	**-0,14**	-0,05	**-0,36**	**0,16**	**-0,21**	-0,06	**0,29**	**0,35**	**0,25**	**0,17**	**0,15**	**0,32**
	Open (*lagged*)	0,04	**-0,40**	**-0,74**	**-0,53**	**0,17**	**0,14**	**-0,89**	**-0,82**	**0,06**	**-0,19**	**-0,31**	**-0,33**
	TofT$_t$	**-0,35**	**-0,52**	**-0,24**	**-0,59**	**-0,14**	**-0,66**	**-0,42**	**-0,30**	-0,11	**-0,14**	**-0,12**	**-0,20**
	Govexp/Inont$_{t-1}$	0,00	-0,07	**0,14**	**0,33**	**-0,23**	**-0,17**	**-0,14**	**-0,27**	**-0,13**	**0,25**	**0,12**	0,04
	Adjusted R2	*0,73*	*0,84*	*0,79*	*0,93*	*0,80*	*0,84*	*0,90*	*0,83*	*0,76*	*0,80*	*0,87*	*0,79*
	SEE	*0,111*	*0,123*	*0,181*	*0,137*	*0,165*	*0,163*	*0,135*	*0,145*	*0,093*	*0,097*	*0,093*	*0,101*
	Error term (*α*)	**-0,39**	**-0,58**	**-0,17**	**-0,29**	**-0,13**	**-0,14**	**-0,20**	**-0,23**	**-0,27**	**-0,50**	**-0,25**	**-0,33**
	Half-life equivalent[3]	*2,1*	*1,5*	*4,4*	*2,7*	*5,9*	*5,4*	*3,7*	*3,4*	*2,9*	*1,7*	*3,1*	*2,4*
eq. (6) ; vars. in growth rates	**ΔOpen**$_t$			-0,16		**-0,13**	**-0,33**			**-0,33**	-0,16		
	ΔTofT$_t$	**-0,15**	**-0,46**	**-0,27**	**-0,52**	0,09	-0,09	**-0,13**	**-0,14**	0,01	0,06	**-0,06**	**-0,08**
	ΔGovexp$_t$ **/Δinont**$_t$		**-0,16**	**-0,22**	**-0,19**	**-0,18**	**-0,19**	0,07	0,08		**0,30**		
	ΔREER$_{t-1}$	0,10	**0,15**	**0,24**	**0,13**	**0,31**	**0,23**	**0,58**	**0,40**	**0,27**	**0,27**	**0,56**	**0,47**
	ΔNER$_t$	**0,51**	0,29	**0,17**	**0,12**	**0,47**	**0,38**	**0,73**	**0,67**	**0,58**	**0,39**	**0,50**	**0,36**
	ΔNER$_{t-1}$	**-0,41**	**-0,44**	**-0,15**	**-0,13**	**-0,31**	**-0,23**	**-0,52**	**-0,49**	**-0,35**	**-0,31**	**-0,45**	**-0,35**
	ΔNERgap$_t$	**-0,22**	**-0,23**			**-0,13**	**-0,10**			**-0,56**	**-0,77**	**-0,15**	**-0,18**
	GB%$_t$ (levels)			**0,01**		**-0,01**						**0,005**	**0,004**
	Adjusted R2	*0,62*	*0,57*	*0,43*	*0,57*	*0,67*	*0,67*	*0,62*	*0,52*	*0,73*	*0,65*	*0,76*	*0,59*
	SER	*0,077*	*0,097*	*0,119*	*0,108*	*0,082*	*0,094*	*0,073*	*0,083*	*0,067*	*0,083*	*0,049*	*0,064*

(1)1925-2000 ; (2) uses *Inont* ; coefficients in **bold** are significant at the 5% level, and in **bold** at 10% level. *SER* : standard error of regression.

(3) Calculated as log(0.5)/log(1- α); where t is the number of years that takes to dissipate 50% of the deviation from equilibrium.
Constant terms and structural dummy are included in the regressions but not shown.

Broadly speaking, the evidence across the six economies is in line with the predictions for the action of the fundamentals. *TofT* coefficients have all the expected

[24] The structural dummies added are: Argentina in 1953 of the mean type (m) to the REER1 regression (R1) and of the mean-and-trend type (m&t) to the REER2 one (R2); Brazil in 1944 (m) to both regressions and in 1906 (m) to R2; Chile in 1947 (t) to both regressions; Colombia in 1983 (t) to both; Mexico in 1977 (m) to both, and in 1932 (m) to R1; and Venezuela in 1960 (t) to R1 and in 1973 (m) to R2.

signs and are statistically significant except in the case of Mexico (REER1).[25] The elasticities are higher when the REER2 is used than in the regressions with the REER1, which is not surprising as the former uses the price of imports in the numerator (see eq. 2). As expected, *Open* is negatively correlated to the REERs with the exception of Chile, and there are little differences between the coefficients of the regressions for each country.

As to intra-country variations, the "commodity lottery" [26] provides a primary source of differences and similarities in the LA6: oil in Venezuela and Mexico, coffee in Brazil and Colombia, copper in Chile and grains and meat in Argentina. This lottery is set to be an influential factor in the RER behaviour of the LA6. First, via their impact on the terms of trade and the external accounts,[27] and indirectly, through their implications for fiscal policy, particularly in mining and oil economies where the commodity revenues are a major contributor to the budget. Thus, the smaller coefficients for *TofT* in the oil economies of Mexico and Venezuela could reflect the fact that oil prices were relatively stable up to the mid 1970s (as until then they were under the control of the main oil multinationals). Whereas, those economies relying on agricultural products tend to show higher exposure to terms of trade fluctuations.

The role of *Rprodm* is less clear. The B-S hypothesis is supported in the case of Argentina, Brazil and Chile, but the remaining countries show positive and significant coefficients. Although this could be attributed to departures from the model's assumptions, there is no obvious explanation for the differences across LA6. Moreover, the lack of price data discriminating between tradable and non-tradable goods makes it difficult to know whether the problem lies in the link between productivity and the relative price of non-tradables or in the workings of the law of one price - or in both.[28] Also mixed is the evidence of government spending, but in this case it is likely to reflect the limitations of the proxies at hand. *Govexp* is negatively correlated with the REERs in Colombia, and *Inont* proved significant and with a negative coefficient in Chile and Mexico (REER2). Otherwise, both variables were either positively correlated or failed to be significant.

The lower section of Table 4 presents the outcome of estimating (6). All the feedback coefficients proved to be statistically significant at the 5% level or lower and with the right sign. The speed of the error correction is significantly higher for the REER2 regressions in Argentina, Brazil, and Mexico. We also include the half-life values implied by the αs, ranging from 1.5 years in Argentina (REER2) to 5.9 years in Chile (REER1). As expected, they are higher than the half-life parameters

[25] Joyce & Kamas (2003) found that terms of trade changes were the most important determinant of RER movements in Argentina, Colombia and Mexico during the period 1976-1995.

[26] The term refers to the joint effect of the magnitude, timing, stability, and product composition of exports (Díaz-Alejandro, 1984).

[27] Country regressions where the terms of trade are explained by the dominant commodity over the whole century and the US whole price index (as proxy for each country's import price index) account for more than half the variations of the dependent variable (results available on request). And the fit improves considerable by changing the main commodity in different periods (e.g., using saltpetre in Chile during the pre-WWII period, and then copper afterwards) and/or by adding a second commodity.

[28] For instance, Engel (1999) in his study of the US RER relative to other developed countries found little evidence for the role of non-tradable relative prices in explaining RER movements.

estimated in Section 2, particularly in countries such as Chile and Brazil. The country ranking of the "reactiveness" to disequilibrium is similar under both REERs, with Argentina the most reactive and Chile the least. Differences in the feedback parameters can be attributed to variations in institutional factors such as the rigidity of wages and prices and the cost of labour mobility. But their specific role in each of the LA6 is beyond the scope of this work.[29]

The ECM regressions can also inform about the effectiveness of the exchange rate policy in speeding up, or slowing down, the automatic adjustment brought about by the fundamentals. Table 4 includes the estimated coefficients linking $\Delta REER_t$ with ΔNER_t and ΔNER_{t-1}. For instance, the outcome of the REER1 regression in Chile implies that a 21% nominal depreciation is required for a 10% real depreciation (i.e., 10%/0.47) in the same year. However, the negative sign of ΔNER_{t-1} indicates that the effectiveness of exchange rate action is eroded by a real appreciation in the following period – for instance, owing to the inflationary impact of the nominal depreciation. Thus, again in the case of Chile, the second effect represents a 6.5% REER appreciation (0.31*21%), resulting in a final real depreciation of 3.5% (10% - 6.5%) at the end of the second year.[30] The corresponding final real depreciation values for the rest of the countries (REER1 regressions) are: Argentina 1.9% (from ΔNER_t = 19.6%), Brazil 0.8% (60.1%), Colombia 2.9% (13.7%), Mexico 4% (17.4%), and Venezuela 1% (19.9%).[31]

The effectiveness of nominal depreciation in the LA6 is broadly consistent with the countries' inflationary record. The more inflation-prone economies of Argentina, Brazil, and Colombia show a relatively less effective use of this policy instrument. Venezuela is an odd case, combining a low inflationary history with a poor effectiveness of nominal depreciation. On other short-run effects, terms of trade improvements are correlated with a move toward real appreciation in all the economies with the exception of Chile and Mexico. The coefficient of $\Delta NERgap_t$ proved significant and negative in Argentina, Chile, Mexico, and Venezuela, suggesting that distortions on the exchange rate market led to accelerated REER appreciation. Finally *GB%* lacked significance in most cases.

[29] In particular, it is puzzling that Argentina appears as the economy with the fastest speed of adjustment (a similar result was found with the univariate analysis) because this country was one of the most affected by price instability and wage indexation (delaying the adjustment to a new equilibrium position) compared to Mexico (after the revolution) and Venezuela with a better record of price and currency stability. A higher speed of adjustment also suggests that overall nominal rigidities were less prevalent in that country than in the rest of LA6, though there is no compelling reason to believe that this was the case. This is a topic for further investigation.

[30] We draw from Richaud et al. (2003) for this analysis.

[31] The corresponding values for REER2 are: Argentina -5% (34%), Brazil -0.9% (82.5%), Chile 3.8% (26.6%), Colombia 2.6% (15%), Mexico 1.9% (25.8%), and Venezuela 0.3% (27.5%).

5. Conclusions

Our univariate analysis of real multilateral exchange rates in the six leading Latin American economies over the last century found a very slow process of mean reversion – if any – in the original series in levels. This contrasts with recent results of long-span studies focusing on developed economies that found evidence of mean reversion in the unadjusted RER series. Moreover, we showed that the initial non-stationarity can be removed by making allowances for trends and structural breaks. The required adjustment can give an insight into the extent of the presence of a "survivors" bias in long-run studies that concentrate on more advanced economies. In our sample of Latin American countries, the half life of the process of the adjusted series ranges from 0.8 to 2.5 years - compared to a 1.6 to 7 years for the series that have only been de-trended.

The failure to reject the unit root hypothesis in the series in levels has important practical implications for the construction of PPP benchmarks for international income comparisons, as well as for the use of PPP exchange rate estimates to determine the degree of misalignment of the nominal exchange rate and the appropriate policy response (Sarno and Taylor, 2002). The presence of non-stationarity, for instance, undermines the use of a constant PPP benchmark in long-span studies involving these countries (Astorga et al., 2005), or a wider sample of economies including the LA6 (Prados de la Escosura, 2004; Maddison, 1989). When comparing GDP per-capita at PPP values, the weakening of the REER in the closing decades of the century suggests that the LA6 welfare record might be worse than that shown by estimates using a constant PPP pre-1980 benchmark.

The outcome with the error-correction specification also shows a reversion but in this case to a conditional mean defined by the co-integrating relationship - though there are important variations in the speed of these adjustments among countries and RER indicators. It also broadly confirms the prediction that the equilibrium REER is determined by its fundamentals. Terms of trade coefficients and, to some extent, those of trade openness, proved significant and with the expected signs; whereas the evidence is mixed in the case of relative productivities and public spending. In the short term, REER movements have responded to both changes in the nominal exchange rate and in fundamentals. Our estimates indicate that nominal devaluation initially had a moderate success in generating a real depreciation, but that this effect was mostly eroded by the end of the second year or so.

Overall, the main results (i.e., on mean reversion and on the significance and signs of the coefficients associated with the fundamentals) are broadly robust to the use of the REER measure - whether the PPP definition or the real import version. However, there are significant country differences, for instance, on the estimated half-life parameters and the explanatory power of the fundamentals. In general, results using the real import series tend to show faster mean reversion, which can be explained by the fact that, by construction, they exclude the prices of non-tradable goods in the trading partners, and with it, potential rigidities linked to the formation of those prices.

Our findings by no means exhaust the enquiry into real exchange rates over the long run in the region. Many interesting aspects remain to be explored, and now additional data make them less elusive. For instance, the apparent difficulties of the LA6 countries in adjusting their real exchange rates, as well as the intra-country differences in the estimated mean-reversion parameters, open further questions concerning to what extent this reflects structural rigidities in reallocating resources after an external shock; or delays in policy response owing to, for example, constraints on real wages. There are other factors that can shed additional light on the behaviour of the RERs that deserve a careful examination, such as the type of the exchange rate regime in place and the use of capital flows to cushion adjustment (though the study of the latter is made difficult by data limitations during the first half of the last century).

As to the wider implications for development and economic convergence, our time-series analysis of the RERs in the LA6 indicates a depreciating trend over the last century (although some countries experienced appreciation during the 1900-1950 period or so). The behaviour of the final decades is in line with evidence found in studies covering a much wider sample of developing countries. This pattern is consistent with a lack of economic convergence, a phenomenon that would call for faster productivity growth in those countries attempting to catch up and real exchange rate appreciation over time. However, we did not find conclusive evidence for the role of relative productivities or support for the Balassa-Samuelson effect, a result that requires further research.

Appendix

Table A1: Testing for unit roots					
Period 1900-2000 (all series in logs)		**Levels & intercept**		**Level, trend & intercept**	
		ADF	*P-P*	*ADF*	*P-P*
Arg	REER1	-0.82	-1.8	-1.49	-2.41
	REER2	-0.71	-1.69	-2.92	-4.19***
Bra	REER1	-1.69	-1.86	-3.25*	-3.81**
	REER2	-1.4	-1.38	-2.75	-3.74**
Chi	REER1	-1.33	-1.56	-1.99	-2.42
	REER2	-1.79	-1.92	-3.13	-3.51**
Col	REER1	-1.72	-1.37	-3.49**	-3.40*
	REER2	-2.27	-1.9	-2.91	-3.11
Mex	REER1	-1.93	-1.65	-2.14	-2.35
	REER2	-2.1	-2.1	-1.87	-2.65
Ven	REER1	-1.06	-1.63	-2.07	-2.52
	REER2	-2.49	-2.63*	-2.82	-2.89

ADF: Augmented Dickey-Fuller test; P-P: Phillips-Perron test.
Critical values for series in levels: -3.50 (at 1% of sig.); -2.89 (5%); -2.58 (10%);
and with trend & intercept: -4.05 (1%); -3.45 (5%); -3.13 (10%).
(*), (**), & (***) indicate that the unit-root hypothesis is rejected at the 10%, 5%, and 1% level of significance, respectively. Truncation lag = 4 in all cases.

Table A2: Non-stationarity test for REER2										
period 1900-2000			**First stage**						**Second stage**	
	(1) *model*	(2) *Tb*	(3) $\hat{\beta}$	(4a) $\hat{\theta}_1$	(4b) $\hat{\theta}_2$	(5a) $\hat{\gamma}_1$	(5b) $\hat{\gamma}_2$	(6) *SER*	(7) $\hat{\alpha}$	(8) $t_{\hat{\alpha}}$ (ADF)
ARG	AO-B	1952	**0.003**			**0.012**		*0.200*		
	AO-C	1952	**0.007**	**-0.287**		**0.012**		*0.188*	**0.43**	-4.26***
	AO-A	1985	**0.006**	**0.321**				*0.199*		
	AO-C	1976	**0.004**	**0.198**		**0.011**		*0.195*		
BRA	AO-C	1980	**0.016**	**0.493**		**-0.042**		*0.208*	**0.54**	-4.72***
	AO-B	1992	**0.017**			**-0.050**		*0.231*		
	AO-C&C	1914&1980	**-0.046**	**0.424**	**0.466**	**0.063**	**-0.042**	*0.187*	**0.54**	-4.22***
CHI	AO-A	1974	**0.004**	**0.511**				*0.224*	**0.72**	-3.95***
	AO-B	1974	**0.008**			**0.009**		*0.261*		
	AO-C	1974	**0.004**	**0.584**		-0.006		0.223		
COL	AO-A	1956	0.002	**0.460**				*0.191*	**0.68**	-3.78***
	AO-B	1955	**0.008**			0.000		*0.224*		
	AO-C	1956	0.003	**0.463**		-0.001		*0.192*		
MEX	AO-C	1917	-0.017	**0.704**		0.020		*0.208*		
	AO-B	1989	**0.011**			**-0.060**		*0.190*		
	AO-0&0	1917&1989	**0.008**	**0.363**	**-0.467**			*0.171*		
	AO-C&B	1917&1989	**-0.017**	**0.582**		**0.025**	**-0.050**	*0.154*	**0.46**	-4.2***
VEN	AO-A	1961	-0.001	**0.358**				*0.164*	**0.75**	-3.82***
	AO-B	1961	**0.003**			0.002		*0.188*		
	AO-C	1972	0.002	**0.427**		**-0.014**		*0.157*	**0.68**	-4.08***

Columns 3-5: estimates in **bold** are significant at least at the 5% level. SER: standard error of the regression.
(*), (**), and (***) indicate that the critical value is rejected at the 10%, 5%, and 1% levels, respectively.

Data Sources

Unless otherwise indicated data comes from OxLAD.

Real bilateral exchange rates. *Argentina*. 1900-92: uses the series “real exchange rate imports” from Veganzones & Winograd (1997). We multiply this series by the ratio CPIus/WPIus to obtain our RER$1 series. 1993-2005: uses growth rates of a RER index obtained from nominal exchange rate (NER) and price data. *Chile*. Own calculations using data from Díaz et al. (2003). Our real bilateral rate uses average inflation rather than year-end values, and our multilateral rates comprise five trading partners in addition to the US and the UK. *Colombia*. 1900-05 & 2000-2005: own estimates using NER and price data (see below). Otherwise, uses GRECO’s RER series without tariff. For *Brazil*, *Mexico* and *Venezuela* we construct series from NER and price data.

Nominal exchange rate (local currency per US$). *Argentina*. OxLAD, official rate. *Brazil*. 1900-33 & 1938 from Abreu (1990). During these years there was a unified rate. In 1934 we use the controlled rate; and in 1935-37 & 1939-46, the “taxas de câmbio livre” - which was applied to imports. For 1947-2000 we employ IPEA’s commercial exchange rates. *Chile*. Díaz et al. (2003), except in 1932-33 from Lüders (1968). In 1900-30 we use the official rate, and in 1960-2000 annual averages of the rate applied to banking transactions. *Colombia*. 1900-04: the exchange rate peso/US$ is calculated based on the peso/sterling parity (López Mejía, 1990). 1905-30: uses GRECO (2002). 1931-74: uses a weighted average of the rate applied to imports (Romero, 2005). The series is completed forward by applying growth rates of GRECO’s series - which mostly reflects rates applied to trade-account flows. *Mexico*. OxLAD for 1900-80, except in 1915-17 when estimates from Cardenas & Manns (1987) are used. After 1981: uses Banco de México - in ITAM (2004). *Venezuela*. Izard (1970) in 1900-37 & BCV (2000) thereafter. During the episodes with differentiated rates for imports (1960-64; 1983-89; & 1994-95) we use simple averages of the rates applied to imports.

With the exception of Argentina, the parallel or market-determined exchange rate between 1950 to 1999 comes from Reinhart & Rogoff (2004). The sources for 1900-50 are: Brazil, “taxa de câmbio livre” from Abreu (1990); Chile, market rate from Díaz et al. (2003); Colombia, rate applicable to short-term capital flows from Romero (2005); and Mexico, market rate from ITAM (2004).

Domestic price indices and inflation. *Argentina*.1900-80: uses OxLAD. 1980 onwards: CEI. *Brazil*. For the sake of inter-temporal consistency we chose the implicit GDP deflator estimated by IBGE (2003) as our measure for the internal price index. *Chile*. 1900-27: Mamalakis (1983). 1928-2000: Instituto Nacional de Estadísticas (INE). During 1970-77 the INE series is corrected using Cortazar & Marshall (1980). 1971-72: inflation in 1971 is calculated by applying the same adjustment used by Díaz et al. (2003) in their Dec.-Dec. series. For 1972 we assume a correction factor that reflects the lower acceleration in annual inflation relative to the year-end values. *Colombia*. 1900-05: inflation estimates from López Mejía (1990). GRECO thereafter. *Mexico*. 1900-13: WPI in Mexico City from ITAM (2004). During 1915-17 inflation grows in line with currency depreciation (from Cardenas & Manns, 1987). 1918 onwards: OxLAD. *Venezuela*. 1900-44: general price index from Baptista (1997). After 1945: CPI from BCV.

Import unit values (IUV). *Argentina*. 1900-86: OxLAD. The series in completed to 2005 applying growth rates of the IUV index published by the Economy Ministry. *Brazil*. 1901-2000: IBGE (2003). 1900: estimated using the rate of growth of the US producer price index (PPI). The series is completed to 2005 using rate of growth of IUV from ECLAC/SYLA. *Chile*. Díaz et al. (2003). Completed to 2005 using rate of growth from ECLAC. *Colombia*. uses OxLAD, completed to 2005 using ECLAC/SYLA. *Mexico*. 1900-27: grows in line with the US PPI. 1928-72: CEPAL (1976). ECLAC/SYLA thereafter. *Venezuela*. 1900-83: OxLAD. Post 1984: own estimates using BCV data of unit import volumes and import dollar values.

Import weights (calculated as the ratio of annual import values to the total). *Argentina*. Mitchell (1993) for 1900-51; IMF Historical Trade Statistics (HTS) for 1952-80; and CEI for 1990-2000. 1980-90: figures are interpolated. Japan data in 1929, 1932, 1936 and 1937 come from US Tariff Commission (1942). Countries included: France, Germany, Japan, UK, US (G5 hereinafter), and Brazil. *Brazil*. IBGE (2003), except 1938-47 from Mitchell (1993). Countries included: G5 and Argentina. *Chile*. Mitchell (1993) for 1900-51 and IMF/HTS for 1952-80; with the exception of US and UK' weights in 1900-60 from Díaz et al. (2003). After 1990 data comes from ECLAC and Servicio Nacional de Aduanas. 1980-90: interpolated. Import shares from Brazil prior to 1950 are assumed to be zero. Countries included: G5, Argentina, and Brazil. *Colombia*. Mitchell (1993) for 1900-51 & 1988; IMF/HTS for 1952-80; and DANE for 1994-2000. 1980-87 & 1989-93: interpolated. Import shares from Venezuela prior to 1926 are assumed to be 0.6% (1926-28 average) and from Brazil prior to 1941 to be 0.1%. Countries included: G5, Venezuela, and Brazil.

Mexico. Mitchell (1993) for 1900-47; IMF/HTS for 1948-80; and INEGI thereafter. 1914-19: interpolated. Import values from the US and the rest of Latin America during 1900-11 & 1932-79 from ITAM (2004). Countries included: G5, and rest of Latin America. *Venezuela*. Mitchell (1993) for 1900-47 & 1988; IMF/HTS for 1948-80. Weights for US, UK, and Germany in 1920-29 are from Machado & Padrón, (1987, 82). After 1993 uses ECLAC and CEI. 1980-87 & 1989-93: interpolated. Import shares from Colombia and Brazil prior to 1948 are assumed to be 0.4% (1948 value) and Brazil prior to 1948 to be 0.1%, respectively. Countries included: G5, Colombia, and Brazil. Note: with the exception of Brazil and Argentina, import data between the LA6 and Japan start circa 1950. In such cases we assume that import flows from Japan during the first half of the century were insignificant.

RER-related data on third countries. *France*. The CPI in 1900-89 is from Maddison (1991), and IMF-IFS thereafter. We use Officer (2002) for the NER to the US$. *Germany*. The CPI in 1900-59 comes from Mitchell (1993), and IMF/IFS thereafter. We use Officer (2002) for the NER to the US$. *Japan*. The CPI in 1900-60 comes from Maddison (1991), and IMF/IFS thereafter. We use Officer (2002) for the NER to the US$ after 1916. *UK*. Index of producer prices 1900-70 is from Mitchell (1993), and IMF thereafter. CPI (retail price index) is from McCuster (2001). We use Officer (2001) for the rate US$ per sterling. *US*. PPI in 1900-12 is from USDC (1975) and in 1913-2000 from BLS (2002). CPI for 1900-70 from Mitchell (1993), and IMF thereafter. We use Officer (2001) for the US$/sterling rate.

Net barter terms of trade. We use OxLAD for Argentina, Colombia, Mexico, and Venezuela. Brazil is sourced from IBGE (2003) and Chile from Díaz et al. (2005).

Labour productivity in manufacturing relative to the US. Own calculations from OxLAD. In the following years the series grow in line with relative overall productivity: Brazil 1900-19; Chile 1900-07; Colombia 1900-24; Venezuela 1900-35. There is no data for Mexico between 1911 and 1920.

Gross trade share as % of GDP (Openness). The procedure to construct the trade ratios is as follows: i) applies growth rates of export and import quantum indices to exports and imports dollar values in 1970; ii) applies growth rates of GDP series at constant prices to the GDP dollar value in 1970; iii) adds up the resulting export and import series and divide them by the GDP series. **Export quantum.** Brazil: IBGE (2003); Chile: Díaz et al. (2005); Mexico: the revolution gap is filled in with values of the series of export of goods at constant dollars (export values deflated by the unit value of exports). **Import quantum.** Argentina: before 1910 figures are calculated from import series at constant dollars. Brazil: IBGE (2003); Chile: Díaz et al. (2005); Colombia: GRECO (2002) for 1905 onwards. Before 1905 uses imports at constant dollars. Venezuela: from 1920 onwards uses an index based on imports from national accounts at 1968 prices (Baptista, 1997). Prior to 1920 uses import series at constant dollars.

Non-machinery investment spending. 1900-94: calculated form data in Hofman (2000), and from official sources thereafter. **Government spending as % of GDP**. Calculated from data in OxLAD.

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