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Nonlinear exchange rate pass-through in Latin America

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Abstract

The study aimed to investigate the exchange rate pass-through (ERPT) to domestic prices in four major Latin American economies between 1999 and 2017. The underlying assumption is that currency appreciations have different degree of pass-through when compared to depreciations. The methodology is based on NARDL model, which allows cointegration analyses after the inclusion of threshold variables. The results showed satisfactory presence of cointegration and consistency towards what the literature predicts. In the short run, the results were in accordance with the literature, with small coefficients, many of them non-significant. In the long run, significant coefficients varying from 9% up to 45% were estimated. For Brazil and Mexico, the currency has a sharper transmission to domestic inflation when it depreciates. In the Chilean case, the opposite outcome prevailed, as appreciations had a larger impact than depreciations. In Colombia, the ERPT coefficients were significant, although linear.

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1. Introduction

Investigating how the exchange rate affects the inflation environment in emerging economies is relevant, since these markets faced several monetary reforms throughout the eighties and nineties. In the run-up to the millennium, more flexible exchange rate regimes were introduced after years of pegged regimes anchored to the U.S. dollar. The change in the nominal anchor brought forth inflation-targeting regimes. In this context, currency fluctuation is likely related to the effectiveness of monetary policy, as monetary authorities in these markets have a constant wariness of currency and financial stabilities. Moreover, the "fear of floating" literature introduced by Calvo and Reinhart (2002) demonstrated that policymakers in developing economies worry that the process of depreciation may lead to an inflation crisis, harming their institutions' credibility.

Following a recent trend in the open macroeconomics literature, the aim of this study is to investigate the exchange rate pass-through (ERPT), in a non-linear fashion, for four major Latin American countries – Brazil, Chile, Colombia and Mexico – with monthly data from 1999 to 2017. The main issue is to understand whether domestic prices are more sensitive to currency appreciations or depreciations (i.e. asymmetric ERPT). According to Bussière (2013), if non-linearities were intense, any inference based on linear models would be misleading. A time series econometric framework based on a NARDL (non-linear autoregressive distributed lag) model, theorised by Shin *et al.* (2014), is thus undertaken. In this model, the exchange rate can be split into two or more variables based on a given threshold. Subsequently, the long-run consistency of the system is assessed through cointegration tests, as per Pesaran *et al.* (2001).

The ERPT is defined as the percentage variation of a price index due to a 1% change in the nominal exchange rate; in other words, it is the exchange rate elasticity of prices. The ERPT is classified as complete when it is equal to one; null, when it is equal to zero; and incomplete when it lies between zero and one. The first case holds if the assumption of producing currency pricing (PCP) is true for all tradable goods. In this case, the imports are invoiced in a foreign currency. Thus, the local price at the dock is the price in foreign currency directly converted by the exchange rate. The prices in question closely follow the exchange rate, and the degree of ERPT is close to one. The ERPT will be close to zero if the local currency pricing (LCP) holds, i.e. if the invoice occurs in the domestic currency. If exporting firms decide to choose the price in the local currency, a change in the exchange rate may not directly affect prices. In such an extreme, firms practise strategic pricing such as *pricing-to-market* (PTM) and the ERPT is thus zero.

Both hypotheses are extreme cases. According to Sarno and Taylor (2002), if the power purchase parity (PPP) and the law of one price (LOOP) hypotheses are validated, the degree of ERPT should be complete. However, Rogoff (1996) pointed out that changes in the exchange rate may not be totally enacted due to several adjustment costs. Likewise, for the validation of the LCP hypothesis, firms may need to have sufficient market power to set prices regardless of the devaluations and overvaluations in the currency. The practice of PTM is associated with highly differentiated goods and low-competitive markets. Indeed, the empirical results supported incomplete ERPT, refuting both "pure" LCP and PCP. However, what causes the occurrence and magnitude of incomplete ERPT to imports prices? Aron *et al.* (2004) surveyed the main causes found in the literature. Accordingly, there are three channels depending on: (i) the degree of mark-up; (ii) the marginal cost and (iii) nominal rigidities. The first two channels wholly depend on the market structure. The variation in the mark-ups allows firms to alter their prices, leading to incomplete ERPT. The presence of variable mark-up depends, in turn, on the functional form attributed to the demand curve, which is composed of the level of competition in the domestic market.

The remaining topics of this article are organised as follows: Section 2 briefly surveys the main theoretical aspects of the non-linear ERPT literature; in the third section, the methodological procedures are formally presented and the data are described. Section 4 analyses the results, and the fifth concludes with the main findings, policy implications and further commentaries.

2. Non-linearity: theoretical assumptions

The following four assumptions describe how exporters set their prices when faced with an exchange rate fluctuation. The first theoretical channel is the market constraints hypothesis (also known as the bottleneck hypothesis).[1] If the importers' native currency appreciates and the foreign exporters' production capacity is already at its maximum short-run level, or if the costs of adjustment are high, it is unlikely that the foreign producers will expand their supply.[2] Thus, the excess of demand for imported goods will lead to a hike in their prices, offsetting the initial effect of the appreciation in the destination market and leading to a higher ERPT for depreciations.

The second hypothesis is the presence of market share objectives.[3] Consider again a depreciation (appreciation) in the importer (exporter) currency; an appreciation in the exporters' currency causes – *ceteris paribus* – a loss of price competitiveness and, therefore, a loss of overseas market share. To offset the appreciation, the exporters can lower the prices in their currency. Thus, from the exporters' point of view, this hypothesis yields a larger ERPT for appreciations (as they are not offset by a movement in prices) than for depreciations.

There may exist, however, a threshold from which the producers do not regard a decrease in their prices as the optimal option. One can thus define the third hypothesis, which is the presence of *downward price rigidities*. If the foreign appreciation is large, demanding an equally large reduction in prices and profit margins, the firm could reach negative mark-up. The exchange rate fluctuation is thus better offset in an episode of devaluation, leading to a lower ERPT. Indeed, increases in prices are said to be more frequent and more likely to occur than decreases. Turning to the local currency standpoint, this hypothesis leads to a large ERPT after a devaluation.

The fourth channel able to explain non-linear ERPT is given by production switching.[4] Suppose there exists a product with international prices that is produced by a foreign price-taker exporting firm; consider also that this firm can alter its production between technologies intensive in both domestic inputs or in imported inputs. In the first case, an exchange rate fluctuation will affect both marginal cost and revenue equally but the production decision will remain unaltered. In the second situation, an appreciation (depreciation) in the foreign (local) currency increases only the marginal revenue of the exporter firm, leading to increasing supply, decreasing prices and ERPT > 0. This hypothesis results in a larger ERPT for appreciations.

The literature concerning ERPT non-linearities and asymmetries is somewhat recent. Bussière (2013) tested both the non-linear and asymmetric assumptions using data from the G7 economies and estimated both cross-country panel and country-specific time series. The extent of

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[1] See Baldwin and Foster (1986), Knetter (1994) and Webber (1999).

[2] The same mechanism may occur in the case of a trade restriction. In this case, the constraint is not a limit in the industrial capacity, but rather a restriction imposed in the destination market.

[3] See Froot and Klemperer (1989), Martson (1990) and Krugman (1987).

[4] See Ware and Winter (1988).

asymmetric pass-through significantly varied among the countries. Caselli and Roitman (2016) estimated the non-linear ERPT based on consumer price indexes in 28 emerging markets. One of the goals was to assess whether the pass-through is larger in episodes of high devaluation and whether it is symmetric concerning appreciations and depreciations. To adequately define episodes of high depreciation, a threshold value was chosen. The evidence suggested that in depreciations of more than 10 and 20 per cent, the ERPT coefficient is twice that of lower fluctuations.

For South American countries, Carneiro *et al.* (2002) focused on Brazil between 1994 and 2001. A backward-looking Phillips Curve with a fixed pass-through coefficient was proposed. Therefore, the inflation is explained in a non-linear fashion by the exchange rate and the seasonally adjusted unemployment rate. In the best specification achieved, the quarterly ERPT is 6.4 per cent after the transition from a fixed to a floating regime. Subsequently, the main implication is that overlooking non-linearities in ERPT estimations can overestimate their degree. Correa and Minella (2010) adopted the Phillips Curve model with thresholds to investigate the non-linear ERPT in the Brazilian economy. It was found that in the short-run, the pass-through is higher when (i) the country is in a high-pace economic growth, (ii) the exchange rate depreciates after a given threshold value and (iii) the exchange rate volatility is lower. These results are related to the presence of PTM, menu costs and uncertainty over the persistency of currency fluctuations.

3. Methodology

We start with the assumption that there is a linear relationship between both prices and the nominal effective exchange rate (e_t) and other control variables (z_t):

$$p_t = \alpha_0 + a_1 e_t + \alpha_2 z_t + \varepsilon_t \tag{01}$$

where p_t is the price index representing domestic inflation; α_1 e α_2 are coefficients associated with the nominal exchange rate and further control variables; α_0 is a constant and ε_t an error.

To achieve our final specification, one should build a standard ECM (error correction model) structure:

$$\Delta p_t = \beta_0 + \sum_{i=1}^{p} \gamma_1 \Delta p_{t-s} + \sum_{i=0}^{q} \pi_1 \Delta e_{t-i} + \sum_{i=0}^{s} \pi_2 \Delta z_{t-i} + \varphi ECM_{t-1} + e_t \tag{02}$$

where the ECM term is the OLS residuals series from the long-run equation:

$$p_t = a_0 + a_1 e_t + a_2 z_t + v_t \tag{03}$$

$$ECM_t = \hat{v}_t = p_t - (\hat{a}_0 + \hat{a}_1 e_t + \hat{a}_2 z_t) \tag{04}$$

Replacing (02) with (04) yields the final specification of an $ARDL\ (q, q_1, q_2)$ model:

$$\Delta p_t = \beta_0 + \alpha_0 p_{t-1} + \sum_{h=1}^{k} \alpha_p w_{h,t-1} + \sum_{s=1}^{q-1} \gamma_1 \Delta p_{t-s} + \sum_{h=1}^{k} \sum_{l_w=0}^{q_w} \pi_{h,l_j} \Delta w_{j,t-l_w} + u_t \tag{05}$$

where $w_{h,t} = \{e_t, z_t\}$ is a k-vector of the explicative variables, $q_w = \{q_1, q_2\}$.

Equation (05) implies a dynamic model in which both the short- and the long-run relationships are estimated. The advantage of the ARDL model compared with traditional econometrics is that, under the null hypothesis, the F_{pss}[5] statistic asymptotic distribution is non-standard regardless of whether the regressors are I(0) or I(1). The bound test has two critical significance values; the lower value is based on the assumption that all of the variables are I(0), and the upper-bound critical value is based on the assumption that all of the variables are I(1). The null hypothesis to be tested in the Pesaran *et al.* (2001) framework is of no cointegration: $\alpha_0 = \alpha_1 = \alpha_2 = 0$. If the F_{pss} statistic obtained in the underlying estimation exceeds the upper-bound value, one can assert the presence of a long-run relationship in the system as the null hypothesis is rejected. On the other hand, if the F_{pss} statistic lies beyond the lower bound, one cannot reject the null hypothesis of no cointegration. Lastly, if the value lies between the bounds, the test is inconclusive. In these cases, the ECM coefficient is tested according to the critical values of Banerjee *et al.* (1998). The presence of cointegration is related to a negative and strongly significant coefficient for the ECM term in the error correction form.

A crucial feature of this modelling tool is the needlessness of testing for a unit root, given that the new critical values account for variables with both orders of integration. According to Nieh and Wang (2005), owing to its advantages in solving the typical problems of integration – an issue that has been of substantial focus for the time series literature and for dealing with the small-sample issue[6] – the ARDL bound test has been widely applied in several studies in recent years.

According to Shin *et al.* (2014), one can set a threshold value that breaks the seemingly non-linear variable into two or more sub-variables. As stressed in the literature review, the interest in the exchange rate non-linearity concerns both its positive and negative variations and large and small fluctuations. Thus, we firstly define a threshold value equal to zero. As a result, two regressors are defined in terms of a partial sum, accounting separately for the actual exchange rate depreciations and appreciations:

$$e_t^{(+)} = \sum_{s-1}^{t} \Delta e_s^{(+)} = \sum_{s=1}^{t} \max(\Delta e_s, 0) \quad (06)$$

$$e_t^{(-)} = \sum_{s-1}^{t} \Delta e_s^{(-)} = \sum_{s=1}^{t} \min(\Delta e_s, 0) \quad (07)$$

The non-linear concept outlined above in equations (06) and (07) is then united with equation (05), as one can replace the original linear variable by the two partial sums.

The next step is choosing a proper set of control variables to include in the vector z_t. The ERPT empirical literature provides some guidance, albeit there is no single way of defining the variables without also affecting prices.

The price dynamics are correlated with the output pace in the short-run, in what is often called demand-pull inflation. When the economy operates in full employment and suffers a demand shock, prices tend to rise.[7] The swinging movements of the income around its long-run trend (output gap) is a typical proxy for short-run deviations of demand. The second set of variables relate to the costs shocks as a result of firms' mark-ups. This approach can be easily derived from a model of imperfect competition. Accordingly, prices are a mark-up over the costs of production factors – namely, labour and capital. For the latter, unit labour costs (ULCs) and minimum wages

[5] PSS is the usual abbreviation used in the literature for the work of Pesaran, Shin and Smith (2001), which we also refer to as Pesaran *et al.* (2001).
[6] See also Narayan (2005).
[7] Such hypothesis have a long history in the macroeconomics literature, in which the most recent and preeminent is the one regarding the New-Keynesian economics. For a complete specification of the New-Keynesian model, see Galí (2008).

are recurrent proxies; for the former, producer price indexes (PPIs), energy and oil prices have been chosen in recent literature.[8] The PPIs are highly correlated with internal prices as foreign PPIs are also frequently used as a more exogenous measure of costs shocks.

To account for these channels, through which diverse shocks may affect inflation, we propose a different set of control variables, as it is not clear which proxy is the most adequate for costs-push inflations. Therefore, we interchange foreign PPI (model 01) and oil prices (model 02). As described in Section 4**,** the results presented robustness between models 01 and 02.

The time span for our estimations must run after the adoption of free-floating regimes; however, this would result in a vastly different number of cross-country observations. To adopt a similar number of observations and avoid the residues of the economic crisis that took place in 1995 in Mexico and in 1997–98 in emerging markets, our observation starts in 1999 and runs until December 2017.[9] All the inflation data came from seasonally adjusted Consumer Price Indexes (CPIs), as previously depicted in graphical analysis[10]. The entire data source encompasses: the central banks of each country, the IMF International Financial System (IFS), the Bank of International Settlements (BIS), the Federal Reserve of St. Louis (FRED), the Chilean INE (Instituto Nacional de Estadísticas) and the OECD data.

4. Results

In the estimation of equation 05,[11] out of 43 reported ERPT coefficients, 30 were significant and 21 showed the expected sign (which is appreciation lowering inflation and depreciation fostering it). The ERPT is larger in the long-run adjustment, as most of the coefficients lay between 10% and 30%. In the short-run, non-significant coefficients were frequent and the ERPT was shown to be small (varying from close to zero to 4.6%, see Table I).

Both specifications for the Brazilian case featured significant ERPTs in the long-run. These coefficients were robust enough throughout different lag specifications; thus, in the long-run, 9% to 15% of currency appreciation is passed through domestic price, while 25% of currency depreciation is transmitted. In the short-run, little evidence was found for either model, as only a lagged coefficient of 1.88% for the depreciation variable and one of -0.78% for the appreciation were noteworthy.

In the estimated equations for the Chilean data, the ERPT of appreciations showed more relevance, especially in the long-run, when they tended to lower inflation more than depreciations tended to hike it. The highly significant coefficients were 37% and 45%. Still in the long-run, an unexpected sign arrived in the second model, as depreciations seemed to lower inflation by 10%.[12]

[8] See Bussière (2013), Correa and Minella (2010) and Delatte and López-Villavicencio (2012).
[9] For Chile and Colombia, it starts in September 1999; for Brazil in March and for Mexico in January.
[10] Appendix 4 depicts how their domestic inflation has behaved alongside nominal exchange rates since 1999.
[11] Beforehand, we tested for the presence of unit roots. The presence of I(2) variables would make the NARDL approach unfeasible; we found no evidence of it. The complete results of the tests are omitted and can be obtained from the authors.
[12] Remember that the variable $e^{(-)}$ accounts for negative changes in the exchange rate. Thus, it either remains in the same previous value or decreases in time. As a result, its estimated coefficients should be interpreted with opposing signs. Therefore, all the expected signs are negative.

To comprehensively check the robustness of this result, we also estimated models with increasing lags and this outcome prevailed. In both short-run models, 4.6% of a depreciation passed through in the first month and around 3% in the following periods. Intriguingly, appreciations seemed to raise prices in lagged coefficients (3.4% in model 01 and 4.5% in model 02).

In the Colombian case, highly significant ERPT coefficients were estimated in the long-run. However, their values were close to one another (lying between 26% and 32%). As a result, the Wald tests failed to reject the null hypothesis. Therefore, in this period, the ERPT in Colombia is linear: 30% of an exchange rate fluctuation passes through prices, regardless of whether the currency appreciates or depreciates. In the short-run, counter-intuitive coefficients occurred again, albeit none of them were strongly significant.

Lastly, the estimations for Mexico showed feeble results regarding the ECM. Depreciations apparently have a lagged impact of around 1.8%. Although appreciations were remarkably non-significant, in the long-run, the pass-through is more intense for a devaluation (20% and 28.5%). In model 01, the Wald test rejected the null hypothesis of linearity.

Table I - Results for the Nominal Effective ERPT

Country	Model Specification				ERPT Coefficients					
					Short-run				Long-run	
	Model	Control Variables	Lag Selection	F-Statistics *PSS (2001)*	$\Delta e_t^{(+)}$	$\Delta e_{t-i}^{(+)}$	$\Delta e_t^{(-)}$	$\Delta e_{t-i}^{(-)}$	$e^{(+)}$	$e^{(-)}$
Brazil	*Model 01*	PPI (USA), PTI_GAP, Linear Trend, Dummy 2008m09	2, 1, 5, 4, 0	4.486	0.0130	-	-0.0066	-0.0188**	-0.1456**	-0.2680***
					(0.011)		(0.008)	(0.009)	(0.057)	(0.033)
	Model 02	Oil Prices, PTI_GAP, Linear Trend, Dummy 2008m09	4, 0, 1, 3, 0	10.594	-0.0078***	-	-0.0045	-	-0.0948**	-0.2554***
					(0.003)		(0.007)		(0.044)	(0.020)
Chile	*Model 01*	PPI (USA), MPI_GAP, Linear Trend, Dummy 2008m07	6, 4, 2, 2, 3	7.340	-0.0451***	0.0343**	-0.0460***	-0.0291**	-0.3762***	0.0136
					(0.016)	(0.016)	(0.013)	(0.014)	(0.061)	(0.034)
	Model 02	Oil Prices, MPI_GAP, Linear Trend, Dummy 2008m07	6, 4, 3, 4, 0	6.196	-0.0258	0.0453**	-0.0460***	-0.0343**	-0.4545***	0.1081**
					(0.018)	(0.018)	(0.014)	(0.015)	(0.105)	(0.053)
Colombia	*Model 01*	PPI (USA), PTI_GAP, Linear Trend	2, 4, 4, 1, 5	7.819	-0.0125	0.0209**	0.0138*	-0.0180**	-0.2608***	-0.3236***
					(0.009)	(0.009)	(0.007)	(0.007)	(0.065)	(0.093)
	Model 02	Oil Prices, PTI_GAP, Linear Trend	2, 5, 4, 1, 0	7.654	-0.0073	0.0174*	0.0107	-0.0152*	-0.3052***	-0.2964**
					(0.01)	(0.01)	(0.008)	(0.008)	(0.082)	(0.131)
Mexico	*Model 01*	PPI (USA), PTI_GAP	3, 0, 5, 0, 0	9.820	0.0051	-	-0.0114	-0.0177**	0.1108*	-0.2007***
					(0.003)		(0.008)	(0.008)	(0.065)	(0.047)
	Model 02	Oil Prices, PTI_GAP	3, 0, 5, 0, 0	8.338	0.0016	-	-0.0135	-0.0184**	0.0458	-0.2852**
					(0.003)		(0.008)	(0.008)	(0.088)	(0.057)

Source: Author's calculations; Eviews 9.0 and Microfit 5.0 were used to obtain these calculations.

Notes:

i) PTI_GAP stands for the gap of the Production of Total Industry (PTI) series; IPM_GAP stands for the gap of the Manufacture Production Index (MPI) series. The latter was used only in the Chilean case, as the PTI time series is not long enough in this country.

ii) The time dummies account for periods of sharp movements in economic activity. For Colombia and Mexico, these dummies were not significant.

iii) In all cases, the dependent variable is the natural log of the seasonally adjusted CPI of each country;

iv) The lags for i vary from 1 to 3. We showed the most recent significant lagged pass-through coefficients;

v) The complete output for the estimations are available from the authors. Diagnostic tests can be found in Appendices 1 and 3.

5. Conclusion

This paper sought to investigate the ERPT in four major Latin American countries in a non-linear fashion. From the methodological framework proposed in Pesaran *et al.* (2001) and Shin *et al.* (2014), we estimated dynamic NARDL models, with threshold variables accounting for exchange rate appreciations and depreciations separately.

The models proposed (two for each country) showed strong evidence of cointegration (i.e. a long-run relationship). In general, the ERPT effect is more intense in the long-run, as these coefficients were larger and more significant than those estimated in the ECM structure. Moreover, out of six applicable Wald tests, four rejected the null hypothesis (see Appendix 2), meaning that in these cases, the non-linear approach fits the data well. For Brazil and Mexico, the currency has a sharper transmission to domestic inflation when it depreciates. In Chile, the opposing result prevailed, as appreciations of the Chilean peso had a larger impact than depreciations. In Colombia, the ERPT effect was significant and pre-eminent, although the tests revealed that the effect is likely to be linear. Lastly, it is noteworthy that some results had opposing signs compared to those expected; namely, lagged currency appreciations in the short-run in Chile and Colombia and in the long-run in Mexico, and currency depreciations in the short-run in Colombia and in the long in Chile. However, these results might be evidence of PTM behaviour from the exporting firms. Nevertheless, the results showed reasonable consistency towards what the literature predicts.

Subsequent progress in this line of research includes estimating the pass-through effect in more countries, as long as adequate data are available. The same study can also be conducted with disaggregated data, as pass-through effects may vary across sectors. However, the availability of price indexes for specific industries might act as an obstacle. Lastly, the non-linear approach also encompasses how different large and small currency movements may affect prices. In this study, we initially focused on the asymmetric behaviour, as our results shed some light on the currency-price dynamics for countries where few authors have covered the topic.

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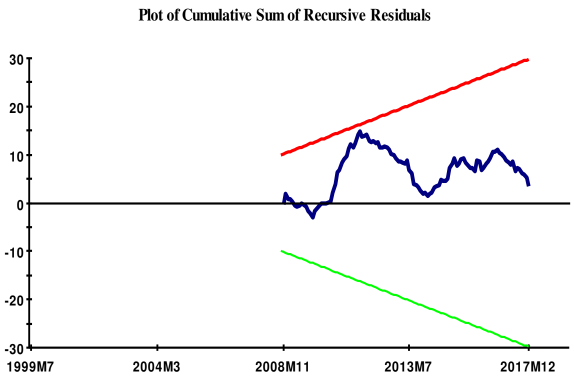
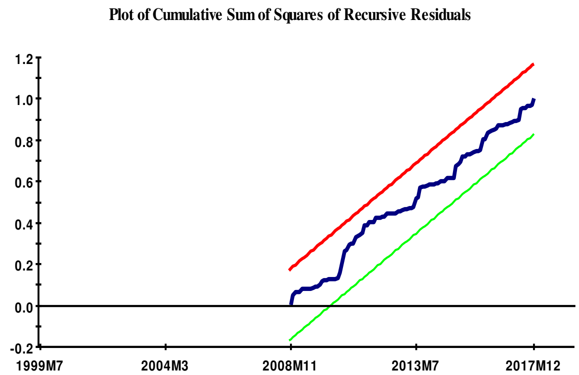
APPENDIX 1 – Stability Diagnostics

CUSUM in the left-hand side; CUSUMQ in the right-hand side.

BRAZIL – Model 01

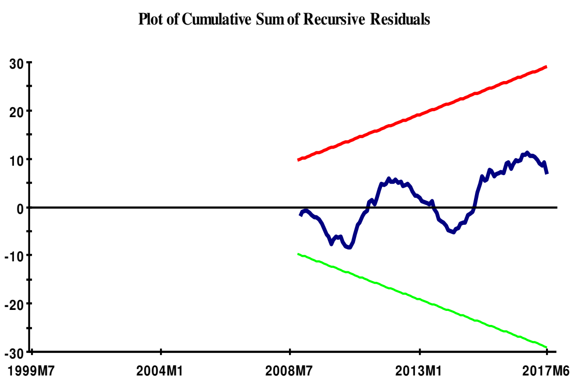


The straight lines represent critical bounds at 5% significance level

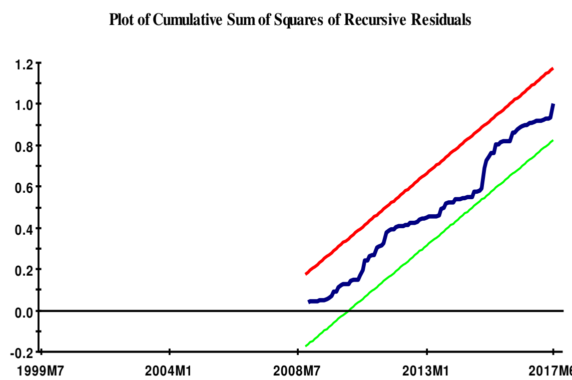


The straight lines represent critical bounds at 5% significance level

BRAZIL – Model 02

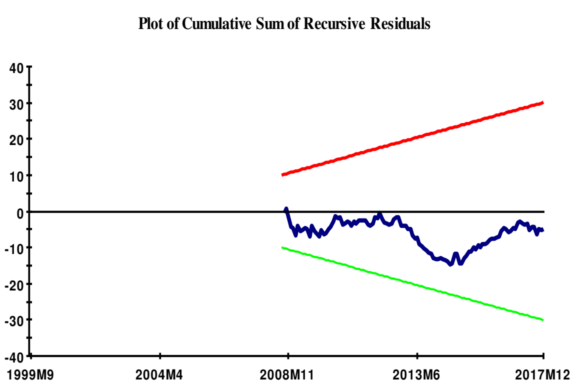


The straight lines represent critical bounds at 5% significance level

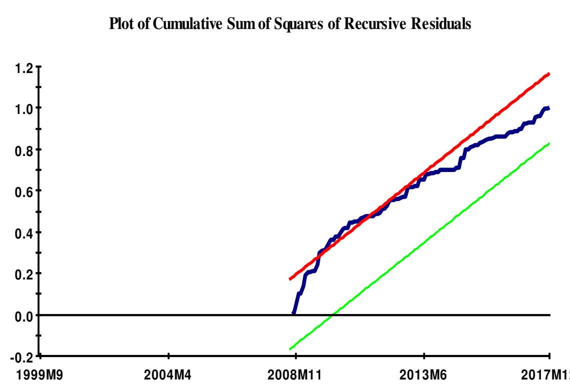


The straight lines represent critical bounds at 5% significance level

CHILE – Model 01

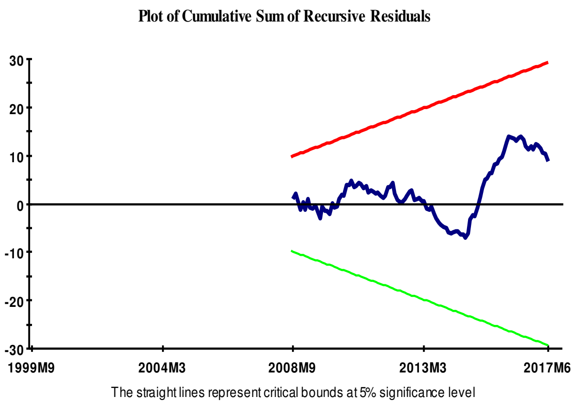


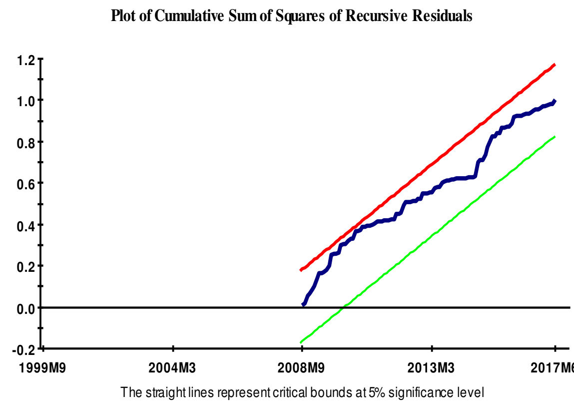
The straight lines represent critical bounds at 5% significance level



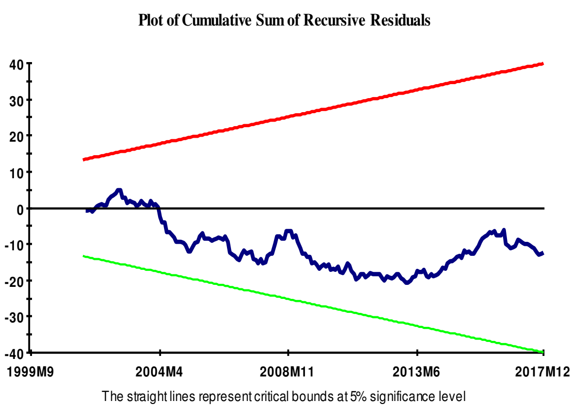
The straight lines represent critical bounds at 5% significance level

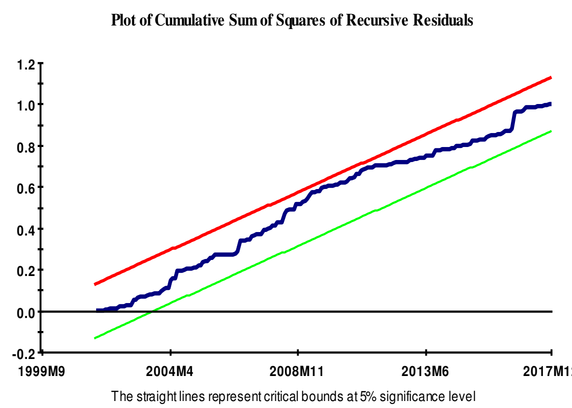
CHILE – Model 02



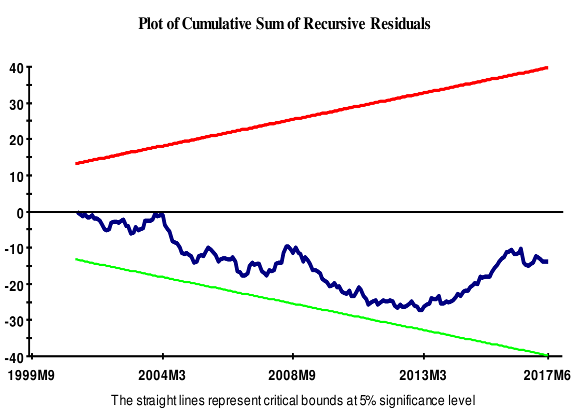


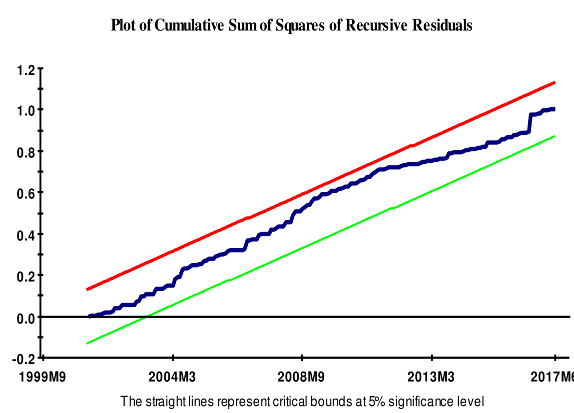
COLOMBIA – Model 01



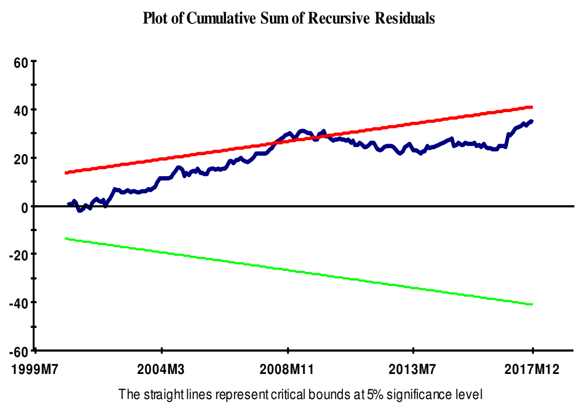


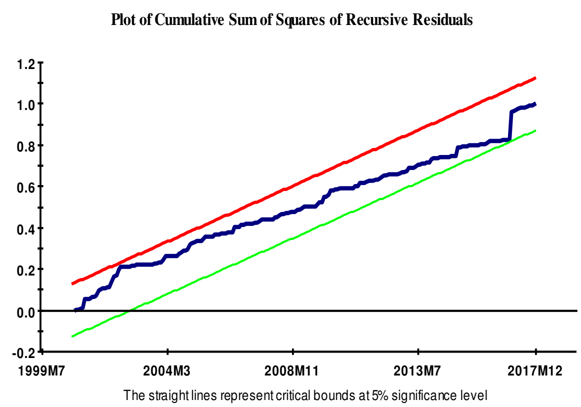
COLOMBIA – Model 02



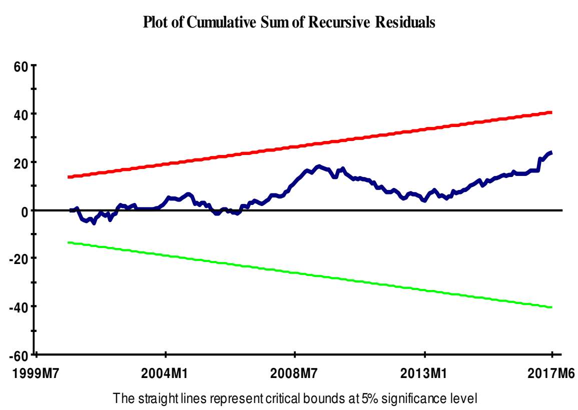


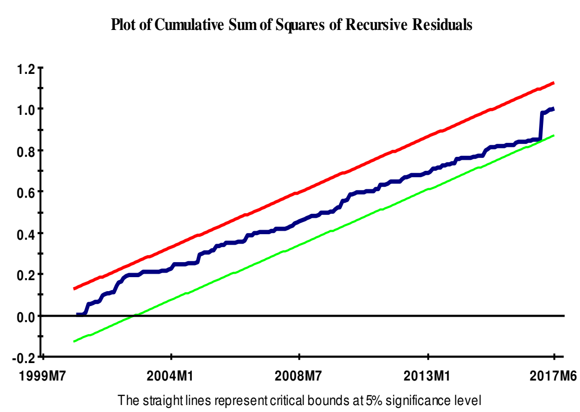
MEXICO – Model 01





MEXICO – Model 02





Appendix 2 - Testing for the presence of nonlinearities

Country	Model	Wald Test Statistic	Wald Test Prob.	NL
Brazil	Model 01	8.145	0.004	Yes
	Model 02	25.073	0.000	Yes
Chile	Model 01	Not applicable		Yes
	Model 02	19.041	0.000	Yes
Colombia	Model 01	0.443	0.506	No
	Model 02	0.005	0.941	No
Mexico	Model 01	147.503	0.000	Yes
	Model 02	Not applicable		Yes

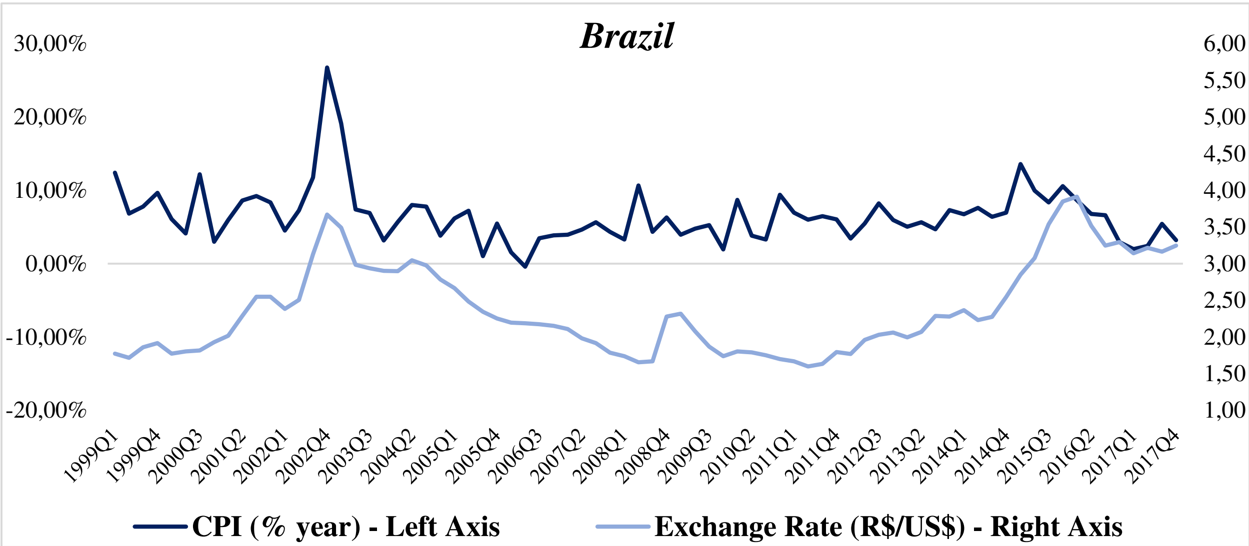
Notes: the test is applicable when both coefficients are significant.

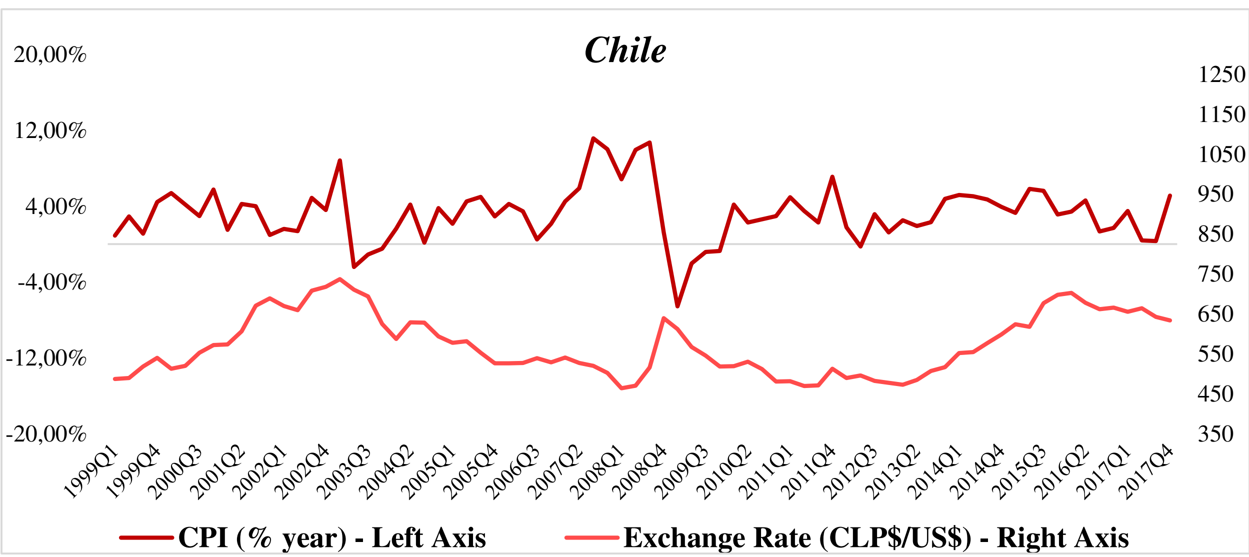
Appendix 3 – Diagnostics

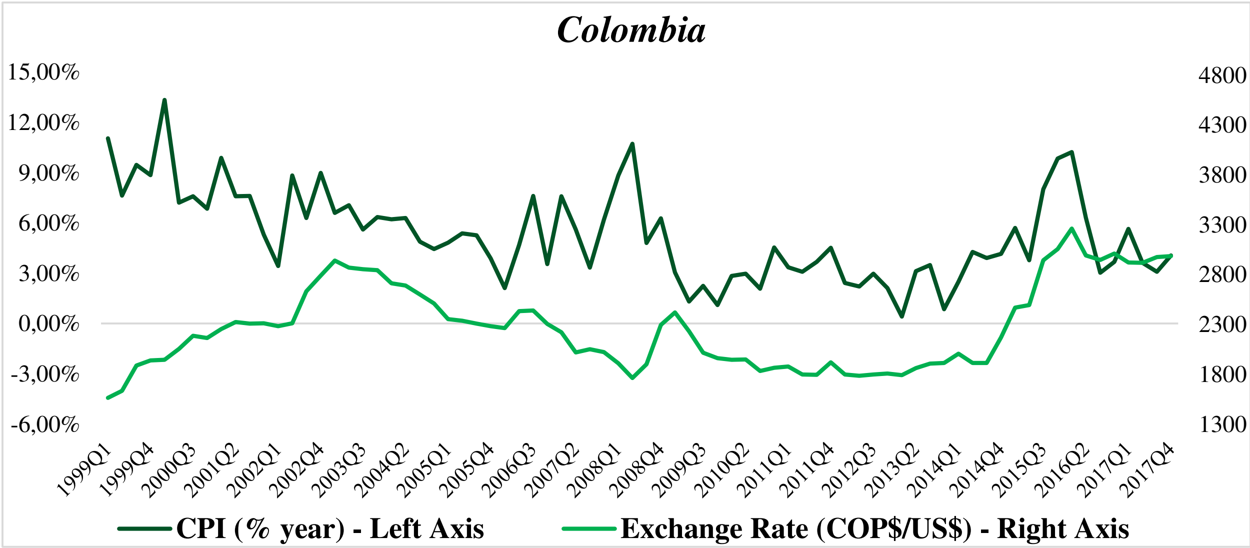
Country	**Model**	**Heteroskedasticity**		**Serial Correlation**
		ARCH	*Harvey*	*LM Test*
Brazil	Model 01	0.242	1.355	0.786
		0.962	0.157	0.582
	Model 02	0.167	1.484	0.228
		0.985	0.120	0.967
Chile	Model 01	2.300	1.188	1.688
		0.036	0.260	0.126
	Model 02	2.123	1.318	1.195
		0.052	0.160	0.311
Colombia	Model 01	0.633	0.846	1.678
		0.704	0.660	0.128
	Model 02	0.664	0.964	1.346
		0.679	0.501	0.239
Mexico	Model 01	0.265	0.260	0.654
		0.953	0.994	0.687
	Model 02	0.288	0.370	0.668
		0.942	0.973	0.675

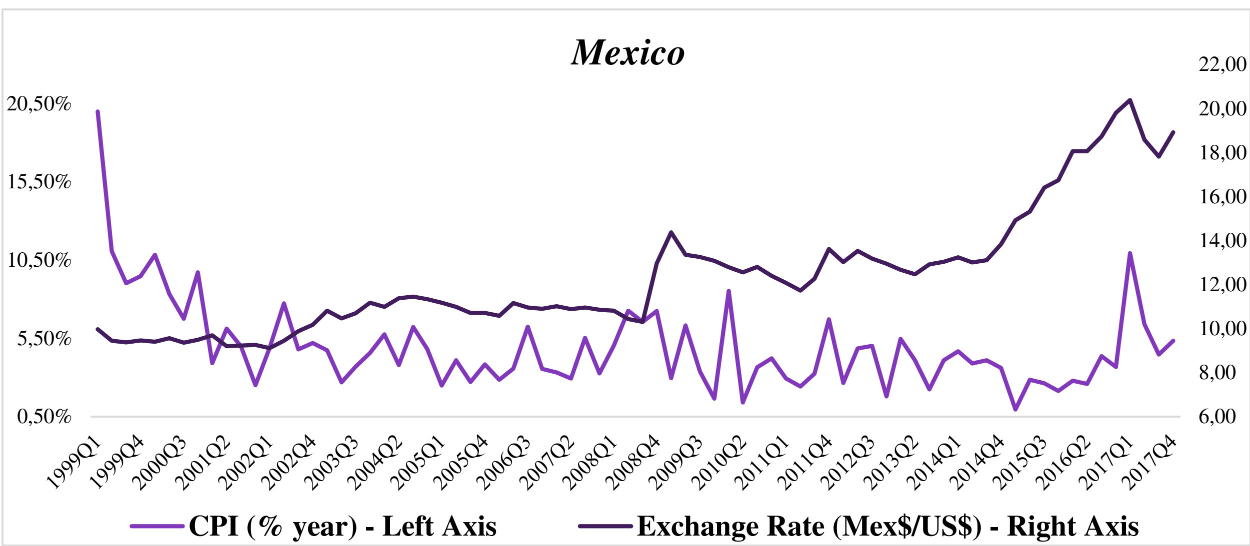
Notes: For each case, the first value is the F-statistic and the second is the p-value associated with an underlying distribution.

Appendix 4 – Overview of Exchange Rate and Domestic Inflation






Colombia
15,00%
12,00%
9,00%
6,00%
3,00%
0,00%
-3,00%
-6,00%
4800
4300
3800
3300
2800
2300
1800
1300
1999Q1
1999Q4
2000Q3
2001Q2
2002Q1
2002Q4
2003Q3
2004Q2
2005Q1
2005Q4
2006Q3
2007Q2
2008Q1
2008Q4
2009Q3
2010Q2
2011Q1
2011Q4
2012Q3
2013Q2
2014Q1
2014Q4
2015Q3
2016Q2
2017Q1
2017Q4
CPI (% year) - Left Axis
Exchange Rate (COP$/US$) - Right Axis


Mexico
20,50%
15,50%
10,50%
5,50%
0,50%
22,00
20,00
18,00
16,00
14,00
12,00
10,00
8,00
6,00
1999Q1
1999Q4
2000Q3
2001Q2
2002Q1
2002Q4
2003Q3
2004Q2
2005Q1
2005Q4
2006Q3
2007Q2
2008Q1
2008Q4
2009Q3
2010Q2
2011Q1
2011Q4
2012Q3
2013Q2
2014Q1
2014Q4
2015Q3
2016Q2
2017Q1
2017Q4
CPI (% year) - Left Axis
Exchange Rate (Mex$/US$) - Right Axis