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Inflation Targeting and the Dynamics between Exchange Rates and Interest Rates: Evidence from Latin America

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Abstract To ensure financial market stabilities, many Latin America countries implemented pure floating and inflation targeting (FIT) policies following the IMF's suggestions. The effectiveness of such policies is under investigation. This paper examines the long-run relationship between the real exchange rates (RERs) and real interest rate(RIR) differentials in major Latin America countries over 1993-2009. It shows there are long-run cointegrations between the RERs and RIR differentials in Argentina, Chile and Columbia, as well as long-run causal relationships in Brazil, Mexico and Venezuela. The results support that the FIT regime has facilitated the regional money market and currencies stabilizations in Latin America. The findings have important implications for policy makers and international investors in emerging markets.

Keywords: *exchange rate, interest rate, cointegration, Granger causality, Latin America*

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1. Introduction

After the Asian financial crises in the late 1990s, there is a prevailing suggestion that developing countries should adopt either a pure floating exchange rate or a hard peg regime to maintain capital flow stability. Following the suggestion[i] of the IMF, most of the major Latin American countries, such as Mexico, Brazil, Colombia and Chile implement pure floating and inflation targeting (FIT) policies. In 1999, all of them officially announce that they switched to an FIT regime[ii], although none of the Latin American FIT countries did really let their currency float the way assumed under a conventional FIT arrangement [10]. Evidence shows that central banks in these countries set their target interest rate in response to nominal exchange rate (NER) movements. Given that the NER is an important transmission mechanism for monetary policy, FIT central banks can use the interest rate to influence inflation through its effects on the NER [33].

However, the central banks in Latin America FIT countries intervene in the foreign exchange (FX) market regularly to affect NER movements through adjusting interest rates. Such behaviors are documented as "fear of floating" [7], albeit there are several arguments about this explanation[iii]. First, all central banks have made explicit statements that they do not pursue exchange rate targets [10]. This implies that interventions are only undertaken when the exchange rate go far beyond the fundamental equilibrium. Second, the RERs in the FIT countries are substantially more volatile than in the countries with a stable and competitive real exchange rate (SCRER)[iv], such as Chile with the crawling bands (1985-1995) or Argentina with the managed floating after the convertibility crisis (2003-2008). In the other Latin America FIT countries, significant appreciation trends were also observed. When the inflation targets were threatened, central banks did not hesitate to induce NER appreciation to meet them. Therefore, investigating the long-run relationship between the real exchange rates (RERs) and real interest rate (RIR) differentials among Latin American FIT countries have important implication for policy makers and international investors.

Previous research on the long-run relationship between the RERs and RIR differentials mainly focuses on the developed economies, few studies examine the market in emerging Latin America countries. The theoretical framework of the law of one price (LOP) and uncovered interest parity (UIP) illustrates that there should be a systematic relationship between the RERs and RIR differentials, although the empirical evidence in support of the classical economic theories is controversial. Two earlier seminal papers by Campbell and Clarida [9] and Meese and Rogoff [32] both conclude that the RER movements cannot be explained by shifts in RIR differentials. Later, Coughlin and Koedijk [13], Blundell-Wignall and Browne [3] and Edison and Pauls [17] argue that the RERs and RIR differentials may be cointegrated given an extended sample period. After around 10 years, by using a longer period of time series data and advanced statistical techniques, Chortareas and Driver [12] reveal that the RERs and RIR differentials cointegration holds better for small open economies rather than G7 countries. More recently, both Nakagawa [35] and Kim [30] are in favor of a long-run linkage between the RERs and RIR differentials with the help of advanced dynamic statistical approaches.

This paper contributes to the previous research in the following ways. First, this is the first paper testing the long-run relationship between the RERs and RIR differentials in Latin America emerging markets, while previous research has overwhelmingly focused on major developed countries. Second, we investigate the cointegration and causality relation country-by-country to draw inferences on individual-specific characteristics; therefore, our methodology is quite different from the extant pooled panel approaches which could miss individual unique properties. Finally, we investigate the RER-RIR relationship in a dynamic way by constructing a Vector Error Correction Model (VECM) allowing for deviations from the long-run equilibrium. Thus, our model provides valuable information on how the equilibrium error in each series could be adjusted by the changes in subsequent periods.

To test for the long-run linkage hypothesis, we retrieve the monthly data over 1993-2009 to perform both Johansen cointegration test and Granger causality test for each nation. The empirical results show that the long-run linear relationship between the RERs and RIR differentials is supported in 3 out of 6 Latin America economies: Argentina, Chile and Columbia. Moreover, the causal relationship between the RERs and RIR differentials is detected for another 3 Latin America emerging countries: Brazil, Mexico and Venezuela, although the pattern of causality is mixed and country-specific. In sum, the causal relationship between the RERs and RIR differentials is found in most of the FIT countries at the 5% level over the past 17 years. The findings of this paper have important implications for policy makers and international investors in emerging markets.

The rest of the paper is organized as follows. Section 2 presents the literature reviews; Section 3 describes the data; Section 4 explains the methodology with a brief review of the theoretical background. In section 5, empirical evidence is reported and interpreted simultaneously. Section 6 discusses the limitations and concludes the paper.

2. Literature Review

The long-run relationship between the RERs and RIR differentials has drawn considerable attention since 1980s. The most well-known papers are those of Campbell and Clarida [9] and Meese and Rogoff [32]. Campbell and Clarida [9] examine whether the RER movements can be explained by shifts in RIR differentials. They find that expected RIR differentials have simply not been persistent enough, and their innovation variances are not large enough to account for much of the fluctuation in the dollar's real exchange rate. Meanwhile, Meese and Rogoff [32] attempt to detect the cointegration relationship, and they find that they cannot reject the null hypothesis of no cointegration between long-term the RERs and RIR differentials. They suggest that a variable that may be omitted from the relationship, possibly the expected value of some future real exchange rate, may have a large variance which, if included, would lead to the finding of cointegration. This conjecture of an important missing variable is also in line with the empirical results of Campbell and Clarida [9].

While in 1990s, the contradicting empirical results of no cointegration between the RERs and RIR differentials become suspect because the time span used in the previous research is not persuasive. Coughlin and Koedijk [13] and Blundell-Wignall and Browne [3] argue that the RERs and RIR differentials may be cointegrated when the sample period is extended. By using more recent data to extend the sample period, Coughlin and Koedijk [13] find cointegration for the Mark/Dollar exchange rate. The ability of Blundell-Wignall and Browne [3] to find cointegration is due to the inclusion of the difference in the share of the cumulated current account relative to GNP in the relevant countries. In addition, Edison and Pauls [17] re-assess the relationship between the RERs and RIRs over the period 1974-1990 and conclude that there might be a long-run relationship between these variables, but it cannot be verified. Although most of the final conclusions in the 1990s showed little empirical evidence in support of a systematic relationship, the researchers also agree that the robustness of the results varies across exchange rates, time periods, and the measures of expected inflation rates.

However, in the 21st century, with the help of longer period of time series data and advanced statistical techniques, more and more empirical evidence on the hypothesized link incline to support the sticky-price exchange rate theories. Chortareas and Driver [12]'s tests are conducted with a panel of 18 OECD economies in the post-Bretton Woods era. When they divide the sample into two subsamples--the G7 and eleven small open economies, they find strong evidence in favor of cointegration for the panel of small open economies. In contrast, there is weak evidence of cointegration in a panel that consists purely of the G7 economies. Their results of supporting a long-run relationship between the RERs and RIR differentials appear to be more positive. More recently, Nakagawa [35] conjectures that the empirical difficulty is due to a failure to recognize nonlinearity in real exchange rate adjustment. When he introduces threshold nonlinearity into a traditional model to take account of a transaction cost-induced "band of inaction" for price adjustment, only outside the band will exhibit mean reversion and bear an association with RIR differentials. Meanwhile, Kim [30] examines the link between the RERs and RIR differentials for traded and non-traded goods with a systematic method in a dynamic seemingly unrelated (SUR) regression for panel data. His empirical result shows that the linkage between the RERs and RIR differential is more favorable for traded goods than for general and non-traded goods.

Recent studies have a special interest in the link between exchange rates and interest rates in emerging market economies, this interest is further spurred by the fact that many of these economies have recently introduced changes in their foreign exchange and monetary policies, moving towards an inflation targeting framework which operate officially under flexible exchange rate regimes [v]. In the case of EMEs, both theoretical and empirical work should take into consideration the specificities of these economies regarding the behavior of interest rates and exchange rates. Authors such as Calvo [5], Calvo and Reinhart [6,7] and Eichengreen have insisted that there are large discrepancies between advanced economies and EMEs.

These differences include credibility problems, the presence of liability dollarization, a high degree of exchange rate pass-through[vi] and non-stationarity in the inflationary process. In this paper, we contribute to the previous research by performing advanced statistical techniques to investigate the cointegration and the causality relation between the RERs and RIR differentials for 6 major Latin American economies to draw inferences on policy strategies and portfolio diversification benefits for investors and policy makers in emerging markets.

3. DATA

The monthly data of the interest rate and exchange rate are collected for 6 Latin American emerging countries and the U.S. over the period 1993:M1 to 2009:M12[vii]. The sample currencies include ARS (Argentine peso), BRL (Brazilian real), CLP (Chilean peso), COP (Colombian peso), MXN (Mexican peso) and VEF (Venezuelan Bolívar). The Consumer Price Index (CPI) for each nation is identified as the percent change over the same period of the previous year. In addition, we set the monthly CPI of the year 2000 as 100. The monthly exchange rate for each country is quoted as foreign currency units per unit of the U.S. dollar (USD). Further, we use annualized average rate on 3-month negotiable certificates of deposit as the benchmark interest rate. The interest and exchange rate data are retrieved from DataStream (Thomson Financial Limited, 2010). To check the data accuracy, we plot each time series and visually inspect the data for errors according to relevant literatures on structural breaks. The CPI data are drawn from the International Financial System (IFS), further details and the methodology of the CPI data specification are available on the International Monetary Foundation (IMF, 2010) official website.

4. Methodology

Dornbusch's [16] seminal paper claims that goods market prices adjust sluggishly in response to anticipated disturbances and to excess demand. Consequently, less than perfectly anticipated monetary innovations can cause temporary deviations in the real exchange rate from its long-run equilibrium value. Further studies on sticky price assumptions are investigated by Frankel [20], Campell and Clarida [9] and Meese and Rogoff [32], *etc*. The interpretation of the empirical tests in this paper depends mainly on three assumptions, expressed in equations (1), (2), and (4) below. First, equation (1) is a feature of a broad class of sticky-price rational-expectations monetary models by Obstfeld and Rogoff [37]. It assumes that any temporary deviation of the real exchange rate from its flexible-price equilibrium value is expected to vanish at a constant rate (in the absence of further shocks). That is, the logarithm of the real exchange rate q_t follows:

$$E_t(q_{t+k} - \bar{q}_{t+k}) = \theta^k (q_t - \bar{q}_t) \qquad (1)$$

where $q_t = s_t + p_t^* - p_t$, s_t is the logarithm of the nominal exchange rate (domestic currency per foreign currency unit), p and p* are the logarithm of the domestic and foreign price level at time t, respectively. $E_t(.)$ is an expectation operator at time t, $\bar{q}_t$ is the real exchange rate that would prevail at time t if all prices were fully flexible, and θ (0< θ <1) is a speed adjustment parameter.

In general, $E_t\bar{q}_{t+k}$ and $\bar{q}_t$ will equal, unless there are no real shocks or all real shocks are random-walk processes. If the ex-ante Purchasing Power Parity (PPP) holds in the long run, it also assumes that

$$E_t\, \bar{q}_{t+k} = \bar{q}_t \qquad (2)$$

Substituting equation (2) into equation (1) obtains

$$q_t = \beta(E_t\, \bar{q}_{t+k} - q_t) + \bar{q}_t$$
$$where \quad \beta = 1/(\theta^k - 1) < -1. \qquad (3)$$

The third important building block is the Uncovered Interest Parity (UIP) relation, which can be written as

$$i_{t,k} - i^*_{t,k} = E_t s_{t+k} - s_t \qquad (4)$$

where $i^*_{t,k}(i_{t,k})$ is the k[th] period foreign (domestic) nominal interest rate at time t. It implies that

$$E_t(q_{t+k} - \bar{q}_{t+k}) = r_{r,k} - r^*_{t,k} \qquad (5)$$

where the k-period real interest rate $r_{t,k}$can be obtained by subtracting the expected inflation rate from the nominal interest rate $i_{t,k}$.

$$r_{t,k} \equiv i_{t,k} - (E_t p_{t+k} - p_t)$$

Substituting equation (5) into equation (3) yields

$$q_t = \beta(r_{t,k} - r^*_{t,k}) + \bar{q}_t \qquad (6)$$

Relaxing equation (4) by allowing for an exogenous risk premium will add a forcing factor to equation (6). Equation (6) relates the real exchange rates to the real interest rate differentials and to the flexible-price RERs. Thus, sticky price theories of exchange rate determination predict that an increase in domestic RIR relative to foreign rates gives rise to an appreciation of the domestic currency. However, in the empirical literatures, a link between the RERs and RIR differentials estimated by Meese and Rogoff [32] can be written as the following version of the equation

$$q_{it} = \alpha_i + \beta_i(r_{it,k} - r^*_{it,k}) + \varepsilon_{it} \qquad (7)$$

where α_i is a constant term for a domestic country i (*i=1,…, N*) and a composite of inflation forecast errors.

The assumption of the classical regression model necessitates that both the $\{q_{it}\}$ and $\{r_{it,k} - r^*_{it,k}\}$sequences be stationary and the errors have a zero mean and finite variance. In the presence of nonstationary variables, there might be what Granger and Newbold [24] call a spurious regression. In other words, the unit root issue arises in the context of the standard regression model.

The results from such spurious regressions are meaningless in that all errors are permanent. In this case, it is often recommended that the regression equation be estimated in first differences as follows

$$\Delta q_{it} = \alpha_i + \beta_i \Delta(r_{it,k} - r^*_{it,k}) + \Delta\varepsilon_{it} \qquad (8)$$

Since $\{q_{it}\}$, $\{r_{it,k}\text{-}r^*_{it,k}\}$and $\{\varepsilon_t\}$ each contains unit roots, the first difference of each is stationary. Hence, the usual asymptotic results apply. In section 5.4, we will check the stationarity property for each RER and RIR differential sequence.

If the nonstatioanry$\{q_{it}\}$ and $\{r_{it,k}\text{-}r^*_{it,k}\}$sequences are integrated of the same order and the residual sequence is stationary, $\{q_{it}\}$ and $\{r_{it,k}\text{-}r^*_{it,k}\}$sequences are cointegrated. This relation can be investigated by conducting Johansen [28] cointegration rank test.

If the cointegration between $\{q_{it}\}$ and $\{r_{it,k}\text{-}r^*_{it,k}\}$sequences does not exist, a bivariate VAR can be constructed to draw inferences about causal relationships. In this circumstance, If both coefficients in the VAR model are statistically significant, it implies bi-directional long-run causality [26,29].

5. Empirical Results

5.1. The Real Exchange Rates

The real exchange rate $\{q_t\}$ sequence is defined as the nominal exchange rate (s_t) that is adjusted by the ratio of the foreign price level (p_t*) to the domestic price level (p_t). In logarithmic form, it can be shown as

$$q_t = s_t + p_t^* - p_t \tag{9}$$

where p_t refers to each nation's Consumer Price Index (CPI) and p_t* refers to the CPI of the U.S. in time t relative to a base year (The CPI of the year 2000 is set as 100).

5.2. The Real Interest Rate Differentials

The RIR differential is defined as the difference between the RIR in each Latin American economy and that of the United States. RIR is the nominal interest rate subtracting the inflation rate. Mathematically, it can be written as

$$\Delta i = (i - \lambda) - (i^* - \lambda^*) \tag{10}$$

where i and λ are the nominal interest rate and inflation rate for Latin American countries; i*and λ* are the nominal interest rate and inflation rate for the United States. Δi is the real interest rate differential.

5.3. Descriptive Statistics

Statistical inspection of data clearly highlights the explosion in inflation experienced during the early 1990s in Argentina, Brazil, and Venezuela. This evidence is also confirmed by the high sum of squared deviations in the last column of the summary statistics that are displayed in Table 1. The data elucidates that there is clear evidence of rejection of normality. Moreover, all series display significant skewness and excess kurtosis. Therefore, the null hypothesis of normal distribution is overwhelmingly rejected for all the series of the RERs and RIR differentials by the Jarque-Bera test [27] at the 5% level. We further eliminate the outliers (most of them are in early1990s) which may affect the robustness of the estimation, yielding 204 observations for each time series.

Table 1. Summary of Descriptive Statistics of the RERs and RIR Differentials of Latin America Countries

Panel A: Real Exchange Rates									
	Mean	Std. Dev.	Median	Max	Min	Skewness	Kurtosis	Jarque-Bera	SSR
ARS	0.332	0.039	1.104	-0.112	0.400	0.230	1.293	26.583**	32.56
BRL	0.874	0.574	4.250	0.132	0.977	2.102	6.315	243.677**	193.81
CLP	6.290	6.274	6.601	6.058	0.130	0.490	2.415	11.074**	3.41
COP	7.568	7.565	7.859	7.209	0.146	-0.163	2.508	12.966**	4.34
MXN	2.359	2.280	3.008	2.129	0.186	1.482	4.628	97.155**	7.00
VEF	-0.077	-0.095	0.834	-0.830	0.373	0.395	2.878	5.436**	28.26
Panel B: Real Interest Rate Differentials									
	Mean	Std. Dev.	Median	Max	Min	Skewness	Kurtosis	Jarque-Bera	SSR
ARGENTINA	0.028	0.028	0.434	-0.213	0.085	1.738	10.431	572.015**	1.48
BRAZIL	2.304	0.096	85.150	-39.233	12.894	3.868	23.054	3927.176**	33748.70
CHILE	0.018	0.010	0.178	-0.051	0.034	1.677	7.906	300.225**	0.24
COLOMBIA	0.036	0.027	0.164	-0.018	0.036	1.253	4.397	69.952**	0.27
MEXICO	-0.021	-0.016	0.254	-0.199	0.059	0.971	7.266	186.734**	0.71
VENEZUELA	-0.139	-0.091	0.273	-0.958	0.228	-1.643	6.077	172.307**	10.60

Note: This table depicts the summary of descriptive statistics of the real exchange rates and real interest rate differentials of Latin America countries. ARS = Argentine peso, BRL = Brazilian real, CLP = Chilean peso, COP = Colombian peso, MXN = Mexican peso, VEF = Venezuelan Bolívar. Panel A of this table shows the summary statistics of real exchange rates while Panel B provides the summary statistics of real interest rate differentials. SSR = sum of squared deviations. Total number of observations for each series is 204. ** denotes significance at the 5% level.

5.4. Unit Root Tests

To formally test for the presence of unit roots in each series, both the Augmented Dickey-Fuller (ADF) [14,15] and Ng and Perron [36] tests are employed to examine the stationary property of the data. Because the ADF test is known to suffer potentially severe finite sample power and size problems, the Ng-Perron [36] unit root test, building on some of their own work [38] deals with both of these problems. To improve finite sample performance, we employ the Ng and Perron [36] approach as a preferred alternative to the traditional ADF test. The testing equation can be written as

$$\Delta y_t = \alpha + \beta t + \lambda y_{t-1} + \delta_1 \Delta y_{t-1} + \ldots + \delta_p \Delta y_{t-p} + \varepsilon_t \quad (11)$$

where α is a constant, β is the coefficient on a time trend and *p* is the lag order of the autoregressive process. To determine the optimal lag length *p*, a 12-month maximum lag length is imposed, and Schwarts Information Criterion (SIC) is followed to select the optimal lag augmentation.

Table 2. Unit Root Tests of the RERs and RIR Differentials of Latin America Countries

	ADF t-Statistic			Ng-Perron MZ_t test		
	level	1st difference		level	1st difference	
Panel A: Real Exchange Rates						
ARS	-1.472	-6.646***		-1.144	-4.834***	
BRL	-2.527	-5.356***		-0.693	-4.155***	
CLP	-1.746	-11.928***		-1.618	-3.861***	
COP	-1.886	-11.994***		-0.848	-7.036***	
MXN	-1.873	-12.894***		-1.409	-7.131***	
VEF	-2.217	-15.847***		-0.353	-7.080***	
Panel B: Real Interest Rate Differentials						
ARGENTINA	-2.677*	-10.908***		-1.442	-3.327***	
BRAZIL	-1.955	-5.245***		-1.709*	-4.082***	
CHILE	-1.476	-10.018***		-1.382	-9.046***	
COLOMBIA	-2.000	-10.668***		-1.606	-7.336***	
MEXICO	-2.504	-10.257***		-1.765*	-8.260***	
VENEZUELA	-2.270	-9.889***		-1.493	-6.981***	
		level	t-Stat.		level	MZ_t
	Critical values:	1%	-3.46	Critical values:	1%	-2.58
		5%	-2.88		5%	-1.98
		10%	-2.57		10%	-1.62

Note: This table reports the ADF and Ng-Perron unit root tests of the real exchange rates and real interest rate differentials of Latin America countries. ARS = Argentine peso, BRL = Brazilian real, CLP = Chilean peso, COP = Colombian peso, MXN = Mexican peso, VEF = Venezuelan Bolívar. Panel A shows the unit root tests of real exchange rates while Panel B provides the unit root tests of real interest rate differentials. The null hypothesis for each test is that the tested sequence has a unit root. The tests contain a constant and a linear trend. Optimal lags are selected automatically based on SIC criterion, and a maximum lag of 12 is imposed. The asterisks ***, * denote significance at the 1% and 10% level, respectively.

Examining the ADF test results in Table 2, there is no evidence to suggest that the RER and RIR differential series are stationary in level at the 5% level except that the RIR differential of Argentina is marginally significant. Furthermore, we test for the stationary properties in the first difference of all the RER and RIR differential series with intercept and trend, we overwhelmingly reject the null of having a unit root for the all series at the 1% level by ADF test. Seemingly, we conduct the Ng-Perron test and get the similar results. We cannot reject the null of non-stationarity in level at the 5% level for all series except that the RIR differentials for Brazil and Mexico are significant at the 10% level. However, in the first difference, we uniformly reject the null of having a unit root for all the RER and the RIR differentials at the 1% level. In sum, all the selected RER and RIR differentials series over the sample period are following I(1) processes.

5.5. Johansen Cointegration Tests

The method of Johansen [28] involves investigation of the *p*-dimensional vector autoregressive (VAR) process of k^{th} order:

$$\Delta Y_t = \mu + \sum_{i=1}^{k-1} \Gamma_i \Delta Y_{t-i} + \Pi Y_{t-k} + \varepsilon_t \quad (12)$$

where Δ is the first-difference lag operator, Y_t is a (p×1) random vector of time series variables with order of integration less than or equal to one, I(1). μ is a (p×1) vector of constants, Γi are (p×p) matrices of parameters, ε_t is a sequence of zero mean p-dimensional white noise processes, and Π is a (p×p) matrix of parameters. The rank (r) of Π contains information about long-run relationships among the variables. The tested hypothesis is that the number of cointegrating vectors is at most r (r=1, . . . , p), we conduct Johansen rank tests by using both the maximum-eigenvalue and trace test statistic on the above unrestricted VAR model.

An inspection of Table 3 tells us that cointegration between the RERs and RIR differentials exist in 3 out of 6 Latin America countries: Argentina Chile and Columbia. For the other 3 Latin American economies, no cointegration is found between the RERs and RIR differentials at least at the 5% significant level.

Table 3. Johansen Cointegration Tests Between the RERs and RIR Differentials of Latin America Countries

	Hypothesized No. of CE(s)	Eigenvalue	Trace Statistic	Hypothesized No. of CE(s)	Eigenvalue	Max-Eigenvalue Statistic
ARGENTINA	None	0.254	61.769**	None	0.254	58.270**
	At most 1	0.017	3.499	At most 1	0.017	3.499
BRAZIL	None	0.184	50.780**	None	0.184	40.353**
	At most 1	0.051	10.428**	At most 1	0.051	10.428**
CHILE	None	0.084	20.886**	None	0.084	17.444**
	At most 1	0.017	3.443	At most 1	0.017	3.443
COLOMBIA	None	0.066	16.812**	None	0.066	16.490**
	At most 1	0.017	3.322	At most 1	0.017	3.322
MEXICO	None	0.211	56.070**	None	0.211	47.177**
	At most 1	0.044	8.893**	At most 1	0.044	8.893*
VENEZUELA	None	0.054	12.775	None	0.054	10.968
	At most 1	0.009	1.807	At most 1	0.009	1.807

Note: This table reports the Johansen cointegration tests between the real exchange rates and real interest differentials of Latin America countries. There are total 199 observations after adjustments over the sample period: January 1993 to December 2009. Optimal lags are selected based on VAR estimations. The critical values are based on the MacKinnon-Haug-Michelis (1999) p-values. Linear deterministic trend is assumed. The asterisk ** denotes significance at the 5% level.

5.6. Vector Error Correction Model

If two variables, Y_t and X_t, are cointegrated, then there must exist an error-correction representation of the form

$$\begin{aligned}\Delta Y_t &= \alpha_0 + \beta_0 Z_{t-1} + \gamma_{0i}\sum_{i=1}^{T}\Delta Y_{t-i} + \eta_{0i}\sum_{i=1}^{T}\Delta X_{t-i} + e_{0t}\\ \Delta X_t &= \alpha_1 + \beta_1 Z^*_{t-1} + \gamma_{1i}\sum_{i=1}^{T}\Delta X_{t-i} + \eta_{1i}\sum_{i=1}^{T}\Delta Y_{t-i} + e_{1t}\end{aligned} \tag{13}$$

where Δ is the first differenced operator (i.e., $\Delta Yt = Y_t - Y_{t-1}$), e_{it} follows iid with zero mean and constant variance, Z_{t-1} and Z^*_{t-1} are the lagged residuals obtained from the following cointegration regressions:

$$\begin{aligned}Y_t &= w_0 + v_0 X_t + Z_t\\ X_t &= w_0 + v_0 Y_t + Z^*_t\end{aligned} \tag{14}$$

Since we find significant cointegrations between the RER and RIR differentials in Argentina, Chile and Columbia, we further employ the above VECM models to investigate the speed of adjustment to equilibrium in these three markets. The results are shown in Table 4.

Table 4. VECM Estimations of the RERs and RIR Differentials of Latin America Countries

	ARGENTINA		CHILE		COLOMBIA	
	D(RIR)	D(RER)	D(RIR)	D(RER)	D(RIR)	D(RER)
Adjustment Coefficient	-0.0036	0.0107	-0.0176	0.0284	-0.0023	0.0296
	[-0.5971]	[1.6566]	[-1.1018]	[1.9575]	[-0.4456]	[1.8545]
D(RIR(-1))	0.3062	-0.7099	-0.2956	0.0084	0.2844	-0.1120
	[4.3289]	[-9.3309]	[-4.0750]	[0.1276]	[3.9819]	[-0.5057]
D(RIR(-2))	0.0891	0.5062	-0.1252	0.0297	-0.0418	-0.0813
	[1.1846]	[6.2596]	[-1.7425]	[0.4533]	[-0.5784]	[-0.3630]
D(RER(-1))	0.1277	0.2961	-0.0318	0.1901	0.0180	0.1730
	[2.2696]	[4.8927]	[-0.3985]	[2.6116]	[0.7795]	[2.4154]
D(RER(-2))	0.1145	0.3353	0.0291	-0.0365	-0.0003	0.0177
	[2.1533]	[5.8628]	[0.3593]	[-0.4945]	[-0.0121]	[0.2457]
Constant	-0.0006	0.0013	0.0004	-0.0009	0.0001	-0.0016
	[-0.2437]	[0.4828]	[0.2097]	[-0.4548]	[0.1417]	[-0.6741]
Adj. R^2	0.1623	0.4047	0.0725	0.0262	0.0569	0.0197
F-stat	8.7514	28.1950	4.1290	2.0742	3.4134	1.8055
SIC	-3.6793	-3.5335	-4.0320	-4.2175	-6.0635	-3.8003
LR test for binding restrictions (rank = 1)						
Chi-squares	55.122***		28.188***		5.141**	
P-values	0.000		0.000		0.023	

Note: This table reports the vector error correction estimations between the real exchange rates and real interest rate differentials of Argentina, Chile and Columbia. There are total 201observations after adjustments over the sample period: January 1993 to December 2009. Cointegration restriction is set β=-1. All t-statistics are reported in brackets. The asterisks ***, ** denote significance at the 1% and 5% level, respectively.

Table 4 confirms that the long-run cointegration holds for Argentina, Chile and Columbia over the sample period from January 1993 to December 2009. Comparing the adjustment coefficients of real interest rate differentials to those of the real exchange rates in Argentina, Chile and Columbia, we can readily conclude that the RIR differentials have an adverse response to the previous period's deviation from long-run equilibrium, while the

RERs have a positive response to the previous period's deviation from long-run equilibrium. However, the RIR differentials are less responsive to their last period's equilibrium error in all these three economies; while the responsiveness of the RERs to their last period's equilibrium error is greater in magnitude. Only 0.36%, 1.76% and 0.23% of the equilibrium error in the RIR differentials could be adjusted by the subsequent period's changes, while 1.07%, 2.84% and 2.96% of the equilibrium error in the RERs could be adjusted by the subsequent period's changes in Argentina, Chile and Columbia, respectively. This is not surprising because the foreign exchange markets tend to react to the market information and policy shocks much quicker than the monetary market. Therefore, the short run dynamic deviations from the long-run relationship are more likely to be influenced by the RERs rather than the RIR differentials in Argentina, Chile and Columbia.

5.7. Granger Causality Tests

If the cointegration does not exist, the following bivariate VAR can be used to draw inferences about causal relationships. In equation (12), If both β_0 and β_1 are statistically significant, this indicates bi-directional long-run causality [26,29]. If $\beta_0 = \beta_1 = 0$, there is no long-run equilibrium relationship between Y and X, then it reduces to the standard Granger (1983) causality test. The following equations are used to explore the subtle relations between Y and X

$$\begin{aligned} \Delta Y_t &= \alpha_0 + \sum_{i=1}^{T} \alpha_{1i}\Delta Y_{t-i} + \sum_{i=1}^{T} \alpha_{2i}\Delta X_{t-i} + e_{1t} \\ \Delta X_t &= \beta_0 + \sum_{i=1}^{T} \beta_{1i}\Delta X_{t-i} + \sum_{i=1}^{T} \beta_{2i}\Delta Y_{t-i} + e_{2t} \end{aligned} \tag{15}$$

where Y_t is the real exchange rate, X_t is the real interest rate differential. Rejecting H_0: $\alpha_{21}=\alpha_{22}=\ldots=\alpha_{2k}=0$ implies that the RIR differentials Granger cause the RERs. Similarly, Rejecting H_0: $\beta_{11}=\beta_{12}=\ldots=\beta_{1k}=0$ implies that the RERs Granger cause the RIR differentials.

We further employ a bivariate VAR model to explore the causal relationships between the RERs and RIR differentials of Brazil, Mexico and Venezuela. Table 5 illustrates that a causal relationship does exist in all the other Latin America countries, albeit the pattern of causality is mixed. Based on the Granger causality test results, we uniformly reject the null of no Granger causality between the RERs and RIR differentials in Brazil, Mexico and Venezuela at the 1 % level. It elucidates that interest rate changes in these countries could cause foreign exchange market fluctuations; on the other hand, the currency markets also have substantial impacts on the domestic money market. In conclusion, the detection of causal relationships between the RERs and RIR differentials in most of these Latin America countries indicates that the changes in RIR differentials and the changes in RERs are mutually affected in the long run, although the pattern of interactions is mixed and country-specific in Brazil, Mexico and Venezuela.

Table 5. The Granger Causality Tests Between the RERs and RIR Differentials of Latin America Countries

Null Hypothesis		RER --/--> RIR		RIR --/--> RER	
	N	F-Statistic	p-value	F-Statistic	p-value
BRAZIL	192	2.556***	0.004	6.424***	0.000
MEXICO	192	9.697***	0.000	3.067***	0.001
VENEZUELA	192	4.188***	0.000	2.558***	0.004

Note: This table reports the Granger causality tests between the real exchange rates and real interest rate differentials of Latin America countries: Brazil, Mexico and Venezuela. There are total 192observations after adjustments over the sample period: January 1993 to December 2009. Optimal lags are selected automatically based on SIC criterion with a maximum lag of 12. --/--> implies 'does not Granger cause'. The asterisk *** denotes significance at the 1% level.

Table 6. The Evolution of Exchange Rate Stabilization Programs in Latin America Countries

Country and program	Starting Date	Ending Date	Description
Argentina,			
FER1967	Mar-67	May-70	Fixed exchange rate (FER)
Tablita	Dec-78	Nov-81	Pre-announced crawling peg
Austral	Jun-85	Mar-86	Fixed exchange rate (FER)
Austral	Mar-86	Sep-86	Pre-announced crawling peg
MBS	1989	1991	Money-based stabilization
Cobnvertibility	Apr-91	present	Currency board 1:1 parity
Brazil,			
FER1964	Mar-64	Aug-68	Fixed exchange rate with periodic devaluation
Cruzado	Feb-86	Nov-87	Fixed exchange rate (FER)
MBS	1990	1991	Money-based stabilization
Chile,			
MBS	1975	1977	Money-based stabilization
Tablita	Feb-78	Jun-79	Pre-announced crawling peg
Tablita	Jun-79	Jun-82	Fixed exchange rate (FER)
Mexico,			
ERB	Dec-87	Dec-94	Exchange rate band
FER1988	Feb-88	Dec-88	Fixed exchange rate (FER)
PACP	Jan-89	Nov-91	Pre-announced crawling peg
Uruguay,			
FER1968	Jun-68	Dec-71	Fixed exchange rate (FER)
Tablita	Oct-78	Nov-82	Pre-announced crawling peg
ERB1990	Dec-90	Present	Exchange rate band

Notes: Data sources are obtained from Calvo and Vegh (1999) and Alfaro (2002).

6. Conclusion

This paper investigates the long-run cointegration relationship between the RERs and RIR differentials in 6 major Latin America emerging markets during the inflation targeting period. The empirical results show the long-run cointegration between the RERs and RIR differentials holds for Argentina, Chile and Columbia over 1993-2009. Moreover, the causal relationship between the RERs and RIR differentials is also supported for the other three Latin America countries: Brazil, Mexico and Venezuela, albeit the pattern of causality is mixed and country-specific. The results are robust and largely significant at the 5% level over the sample period.

We find either cointegrations or causal relations between the RERs and RIR differentials in most of the Latin American economies during the inflation targeting period. The reasons could be explained as follows: First, the Latin American financial markets have switched to a FIT regime since early 1990s, thus, most of the major Latin American economies have undergone currencies stabilization progresses (see Table 6). For example, Chile and Argentina have begun a gradual adjustment process in exchange rate, while other countries have also implemented some adjustment programs to eliminate their relatively high inflation trends. Second, intervention in foreign exchange (FX) has been regularly undertaken by several Latin American countries, such as Mexico and Venezuela. Third, many countries implement inflation targeting program. Since 1999, Brazil, Chile, Columbia and Mexico have fully implemented inflation targeting programs, although weaker forms of inflation targeting measures have been undertaken since early 1990s. For example, the soft inflation targeting policy in Chile started in 1991, and followed by Columbia and Mexico in 1995. [34]. Under such a regime, central banks in these countries set their target interest rate in response to nominal exchange rate (NER) movements. In addition, local currency operations in credit markets are gradually being displaced by dollar denominated operation in Latin America [40], for both deposits and loans. In sum, combining the stylized facts in Latin American financial markets, our empirical findings may have more important implications of portfolio diversification for policy makers as well as international investors in emerging markets.

Future studies can be conducted by employing more advanced statistical techniques, such as nonlinear models, threshold integration or panel cointegration approaches allowing for structural break. Moreover, one may better explain the dynamics between the RERs and RIR differentials for each country by observing the specific omitting variables between the relationships and test the possibilities of applications for other emerging markets. In addition, there are big discrepancies in inflation rates among Latin America countries in early 1990s, which call for some adjustments, such as the exchange rates of Brazil and inflation rates in most of the countries. Finally, we start the data from January 1993. Further studies can be done by extending the sample period back to 1980s if the adjustment on inflation rates is appropriate.

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Notes

[i] The IMF suggests that a pure floating regime is a better institutional setting (Fischer, 2001).

[ii] Despite their public statements about their exchange rate regime choice, both Chile and Colombia have already been utilizing annual targets of inflation since 1990 and 1991 respectively. Peru had been using a managed floating policy jointly with a monetary regime since the early 1990s.

[iii] This means that, despite the recently proclaimed switch to floating exchange rates, the evidence seems to suggest a reversion to some degree of exchange rate management, albeit one which seems to be less tight than before the crisis. In this regard, some analysts have found considerable discrepancies between the de jure exchange rate classifications and the de facto regimes (Reinhart and Rogoff, 2004).

[iv] The RER under the Brazilian FIT has been 259% and 209% more volatile than under the managed floating in Argentina and the crawling bands in Chile, respectively (Frenkel and Rapetti,2010).

[v] Some Asian emerging countries have declared that their currencies have floated in post-Asian-crisis period, accompanied by a switch to inflation targeting. Such moves were taken by South Korea in 1998, Indonesia in 2000, Thailand in 2000, and the Philippines in 2001. In Latin America, inflation targeting has been adopted by Chile in 1990 (together with an exchange rate float only since 1999), Mexico and Colombia in 1999, Brazil in 2000, and Peru in 2002. Among Eastern and Central European countries, EU new member states Czech Republic and Poland have also moved to comparable monetary and exchange rate policy frameworks (in 1998 and 1999, respectively), while South Africa and Israel count among other middle-income inflation targeters.

[vi] Ca'Zorzi et al. (2004) find that not all EMEs display degrees of exchange rate pass-through above those seen in advanced economies. In particular, while pass-through tends to be high in countries in Eastern and Central Europe and Latin America, it is relatively low in many Asian economies.

[vii] The early 1990s marked the turning point in the recent economic development of Latin America with the implementation of debt relief programs and the resumption of capital inflows to the region (Frenkel and Ros, 2006).