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Exchange Rate Movements, Stock Prices and Volatility in the Caribbean and Latin America

Andre Yone Haughton[1], Emma M. Iglesias[2]*†

[1]Department of Economics, University of the West Indies Mona, Kingston 7, Jamaica, [2]Department of Applied Economics, II. Facultad de Economía y Empresa, University of A Coruña, Campus de Elviña, A Coruña, 15071, Spain. *Email: emma.iglesias@udc.es

ABSTRACT

We analyze the interrelationship between stock prices and exchange rates in the only two Caribbean countries with stock market and floating exchange rates: Jamaica and Trinidad and Tobago. We also study the same four Latin American countries as in Diamandis and Drakos (2011). Using their model, our results show a very mild relationship between both variables in Jamaica, Trinidad and Tobago, Argentina and Brazil, but we cannot find any relationship in the other countries as in Diamandis and Drakos (2011). However, when we extend their model including a generalized autocorrelation conditional heteroskedasticity (component to examine the impact of volatility, our results changed drastically: Stock prices significantly impacted the exchange rate in the tranquil sub-period and the full period for Jamaica, over all three periods for Trinidad and Tobago and in the tranquil period for Argentina, Mexico and Chile. This shows the importance of incorporating volatility explicitly in the model. Our results have the policy implications that governments in the previous countries should try to prevent a currency crisis by stimulating economic growth and the expansion of the stock market to attract capital inflow as in Lin (2012).

Keywords: Exchange Rates, Stock Prices, Volatility
JEL Classifications: F31, G01

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1. INTRODUCTION

Research on the interaction between exchange rate and stock prices has received more attention since the recent global financial crisis. Researchers in the Unites States and Asia mainly, have been trying to estimate the direction of causality between both variables. According to Lin (2012), there are two main theories surrounding their interaction: (1) The one proposed by Dornbusch and Fisher (1980) with the "flow oriented" models of exchange rates, which looks specifically at the balance of trade between countries. Theoretically, exchange rate fluctuations influence the output and hence competiveness of firms. If firms are more competitive this has a direct positive effect on its stock prices, since stock prices represent future cash flow streaming for a company. (2) The second approach, the "stock oriented" model of exchange rate proposed by Frankel (1983) and Branson (1993), states that advances in the stock market affect exchange rate through the liquidity and the wealth effects. A decrease in stock prices reduces the wealth of local investors, which lowers their demand for money. Then banks react by lowering interest rates which dampens capital inflows, reducing the demand for local currency and therefore depreciates the local currency. Since domestic and foreign assets are not perfect substitutes in the portfolio balancing effect, as investors adjust their portfolio ratio of domestic to foreign assets in response to changes in economic conditions, the exchange rate responds accordingly. Evidence of either theory is not uniform across countries as various studies employing a range of different techniques revealed varying results. Most of these studies have been focused on North America, Europe, Asia, and Latin America to a lesser extent but none have investigated this issue in the Caribbean, from the best of our knowledge. Recently, Hassanain (2017) has analyzed the case of the Gulf Cooperation Council.

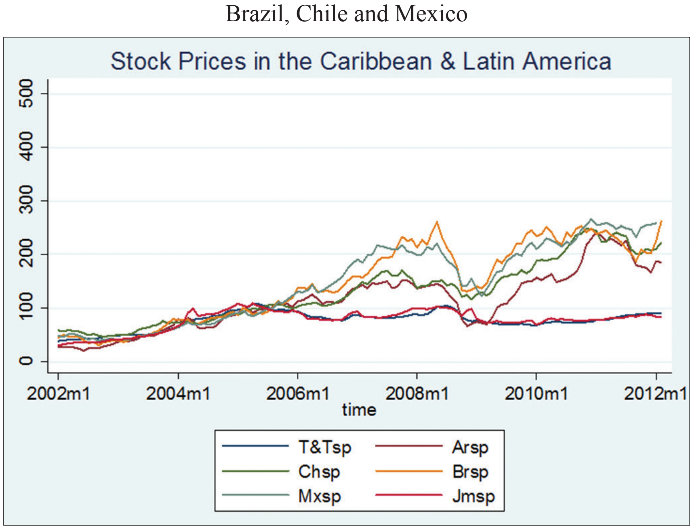
Our objective is to analyze the interrelationship between the stock market and the exchange rates in the two Caribbean countries that have floating exchange rates and stock markets: Jamaica and Trinidad and Tobago. We also study the same four Latin American countries as in Diamandis and Drakos (2011): Argentina, Brazil, Chile and Mexico. We follow Lin (2012) by analyzing the relationships between the exchange rate market and the equity market during the tranquil (2002-2008) and during crisis (2008-2012) periods and we also use the autoregressive distributed lag (ARDL) model bounds test approach proposed by Pesaran et al. (2001). Diamandis and Drakos (2011) claim that the type of exchange rate regime being operated in the particular country will influence the long run relationship between both variables. All the countries in this study operate a float or managed float exchange rate regime.

The relationship between exchange rate and stock prices have a tendency to be greater during crisis periods as returns in asset markets are lower and volatility are higher (Lin, 2012; Guo et al., 2011). Therefore, we extend Diamandis and Drakos (2011) and Lin (2012) studies to incorporate a generalized autocorrelation conditional heteroskedasticity (GARCH)(1,1) component as in Bollerslev (1986) in the ARDL framework to take account of the impact of risk in the model. Note that the importance of volatility is made clear in many contexts nowadays also in the Caribbean countries (see for example Mapp and Wiston (2015) who analyze the impact of the informal economy on the volatility in the Caribbean countries). Our objective in this paper is to show that volatility must be modeled explicitly also when analyzing the relationship between stock prices and exchange rates. Gordon and Pettiford (2016) have also demonstrated how accounting for ARCH effects can be a crucial factor to establish relationships among macroeconomic variables.

Like in Asia, North America and Europe, the Global Financial Crisis in 2008 resulted in an immediate decline in the stock prices (sp in Figure 1) in the Caribbean and Latin American Countries (Figure 1). Notice the behavior is similar for the countries in each group, and also stock prices in Latin America tend to be higher (corresponding to larger economies) than the Caribbean.

Exchange rates in our Caribbean and Latin America countries depreciated as a result of the financial crisis (Figure 2). The shift

Figure 1: Stock prices in Jamaica, Trinidad and Tobago, Argentina, Brazil, Chile and Mexico



for Jamaica in particular is more obvious given that exchange rates are higher and appear to be less stable than the other countries (Figure 3). We therefore expect a stronger relationship between the exchange rate market and the stock market in Jamaica. Even though Caribbean and Latin American countries have floating exchange rates, their central banks can intervene the market to curtail rapid depreciations. In the event of a sudden decline in the value of the domestic currency, the central bank might increase interest rates to attract foreign investors thereby increasing the supply of foreign currency or sell reserves to maintain currency stability. If interest rates are already high (such is the case in Jamaica), further increases will reduce economic activity thereby reducing productivity of firms and therefore the value of their stock. Jamaica level of reserves has not been enough to keep the currency stable. To account for this in our analysis, we follow Lin (2012) by including net international reserves (NIR) and interest rate variables in our model. This will improve our results and correct for omitted variable biases. The following section provides a brief overview of the current literature.

2. LITERATURE REVIEW

Research on the interrelationship between exchange rates and stock prices has been carried out for a variety of countries using various techniques which have produced varying results. In Latin America, Diamandis and Drakos (2011) analyze the long run relationship and short run dynamics between exchange rates and stock prices as well as the impact of exogenous shocks on four countries: Argentina, Brazil, Chile, and Mexico using cointegration techniques and Granger causality tests. They found no significant long run relationship between stock prices and exchange rates for each country. However, after incorporating the US stock market, their results show that stock prices and exchange rates are positively related, with the US stock market facilitating the transmission between the two in these countries. The interaction is independent of the sample choice but Hansen and Johansen (1993) instability tests show that some of their cointegrating coefficients are stable overtime.

Empirically, the research is largely concentrated on more developed countries. Early research by Neih and Lee (2001) examine the dynamic relationship between stock prices and exchange rates for the G7 countries using basic cointegration tests and vector error correction models (VECM) from 1993 to 1996. This research did not account for dual causality between the variables and their findings suggest that there is no long run relationship between stock prices and exchange rate in the G7 counties. Muller and Verschoor (2006) examine how multinational firms in the US are affected by exchange rate fluctuations. They believe that currency movements are a major source of macroeconomic instability which affects a firm's value; a situation they refer to as exchange rate exposure. Theoretically, they outline several reasons why the exchange rate/stock price interaction might be asymmetric. These include the asymmetric impact of hedging on cash flow, firms pricing to market strategies, asymmetry due to hysteric behavior, investors over reaction and mispricing errors and nonlinear currency risk exposure.[1]

1 For more in-depth analysis see Muller and Verschool (2006).

Figure 2: Exchange rates Trinidad and Tobago, Argentina, Brazil and Mexico

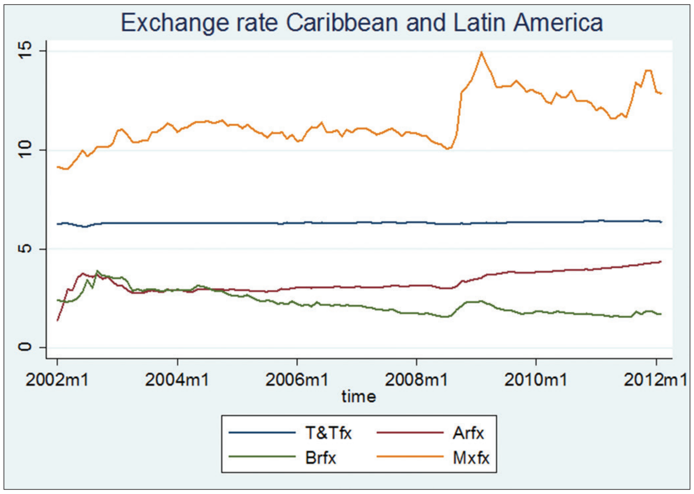
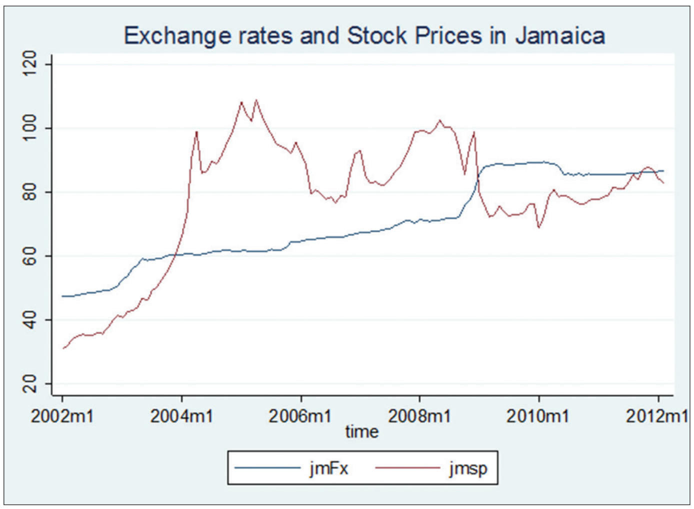


Figure 3: Exchange rates and stock prices in Jamaica



Also on the US, Vygodina (2006) uses a Granger causality test to investigate the relationship between stock prices and exchange rates controlling for the size of the firm from 1987 to 2005. The results found causality from large stock prices to US exchange rate but no causality from small stock prices. The results from the subsamples show that there might be evidence to support the claim that causality between the two variables is changing overtime.

Some have also explored the issue in Asia. Pan et al (2007) used Granger causality tests and vector autoregressions to examine the dynamic linkages between exchange rate and stock prices in seven Asian countries: Hong Kong, Japan, Korea, Malaysia, Singapore, Taiwan and Thailand from 1988 to 1998. Their results showed significant causal relationships for Hong Kong, Japan, Malaysia and Thailand before the financial crisis. They also found evidence of causal relationships between the equity market to the foreign exchange market for Hong, Korea and Singapore. They also found causality from exchange rate to the stock market for all countries except Malaysia, while there is no causality from stock prices to exchange rates. They claim their results are robust to a variety a testing procedures including causality tests and variance decomposition.

Yau and Nieh (2006) empirically investigate the new Taiwan dollar exchange rate against the Japanese yen on stock prices in Japan from 1991 to 2005. Granger causality test showed that no short run causal relationships existed between the variables for both countries. Also, the findings suggest there is no relationship between the exchange rate and the respective stock prices in the long run. Yau and Neih (2009) also examined the relationship between Japanese exchange rate and Taiwan stock market using a threshold error correction model proposed by Enders and Siklos (2001). Their findings suggest that there is a long run equilibrium relationship between the Taiwan dollar, the Japanese Yen and the stock prices of Japan and Taiwan, but asymmetry only exist for Taiwan as the effects of the Japanese exchange rate is symmetric.

Zhao (2010) used monthly data from 1991 to 2009 to examine the dynamic effects between exchange rate and stock prices in China by employing a vector autoregressive approach (VAR) and a multivariate GARCH. Their results show that there is no definite long run relationship between the Chinese Renminbi real effective exchange rate and stock prices in China. They also found no spill over effects between the two variables. The paper goes a step further to examine the cross volatility effects between stock market and the exchange rate using likelihood ratio tests. The results show that there is volatility spill over effects from stock prices to exchange rate and from the exchange rate to stock prices.

More recently, Tsai (2012) uses quantile regression to investigate the relationship between stock price index and exchange rate in six Asian countries: Singapore Thailand, Malaysia, Philippines, South Korea and Taiwan. The results supported a priori information that the two variables are negatively related. More specifically, the negative relationship observed is more pronounced when exchange rates are extremely low or extremely high. This result is supported by the portfolio balancing effect in these two markets which outlines that an increase (decrease) in the returns on stock price index will result in an appreciation (depreciation) of the domestic currency via a decrease (increase) in the exchange rate. Tsai (2012) findings also suggest that the relationship is not homogeneous across countries and across market situations and the coefficients may vary since the portfolio balancing effect is not present all the time in every market. They explain that a significant impact of stock prices on exchange rates exist in time where large sums of capital enter or exit the market.

Our research in this paper follows the method of Lin (2012) who examines the relationship between the exchange rate and stock prices in Asia's emerging markets from 1986 to 2010. Using monthly data, the ARDL model proposed by Pesaran et al. (2001) was employed. This method is designed to account for structural breaks and data that are integrated of different orders. The results from the cointegration tests as well as the short run causality tests indicate that the co-movement between exchange rate and stock prices increases during times of economic crisis and it reduces when the economies are stable. These results correspond with the general literature on exchange rate spill over effects on the stock market. The results also show that most spill overs are in the channel from stock price shocks to exchange rates. In theory, economic slowdown reduces the value of companies stocks causing investors to withdraw their capital which reduces the demand for the domestic currency and out downward pressure on the exchange rate. Apart from the findings for aggregated data, Lin (2012) also examined the issue using industry level data, and the results indicated that the co movement is weak for export oriented industries such as IT for example. All in all, the findings suggest that the interrelationship between exchange rate and stock market is driven by changes in the capital account rather than changes to trade balance in these Asian countries.

Our research estimates the interrelationship between stock prices and exchange rates using the same approach as in Diamandis and Drakos (2011) and Lin (2012) by employing the bounds cointegration tests in the ARDL model, since it provides meaningful long run results even if all the variables are not integrated of the same order. During crisis period, market returns are lower and volatility is higher as the correlation between assets prices tends to be greater see (see e.g., Climent and Meneu (2003) and Guo et al., 2011). This justifies our extension incorporating a GARCH(1,1) component in our ARDL framework as volatility (risk) is also a determining factor in the relationship between prices and exchange rates.

The remainder of the paper is organized as follows: Section 3 outlines the data, data sources, the methodology employed and the results. Section 4 concludes. Appendix A contains some of the Tables A1-A6 and Figures A1 and A2.

3. DATA AND METHODOLOGY

3.1. Data

We examine the interrelationship between the stock index and the exchange rate in the Caribbean and Latin America using monthly data from 2002 to 2012. Data on the share price index (sp), the exchange rate relative to the US dollar (fx), the money market rate (mm) and foreign reserves minus gold (r) are collected from the international monetary fund (IMF) international financial statistics. All data are transformed into logarithms.

The data begins in January 2002 which is approximately the same time the asset bubble began to develop in the international asset markets. In this way we also minimize any effects of the Jamaican and the Mexican financial crisis of the 90's.

On the other hand, the reason to finish our sample size in 2012 is in order to remove the economic effects of the big increase in debt of the economy in Jamaica that happened in that year. In early 2010, the Jamaican Government asked the Jamaica debt exchange to retire high-priced domestic bonds and reduce annual debt servicing. However, debt continued to be a serious concern, forcing the government to negotiate and sign a new IMF agreement in May 2013 to gain access to approximately $1 billion additional funds. As a precursor, the government instigated a second National debt exchange in 2012.

First we analyze the data across the full time period and later we split the data into two parts: (1) The first sub-sample from 2002:01 to 2008:08 (the so called tranquil period) where the asset bubble was developing. (2) The second sub-sample is taken from 2008:09 to 2012:02 (the crisis period). This will provide useful comparisons of the interrelationship between the variables before and after the announcement of the recent global financial crisis.

The summary statistics are provided in Table 1 for the full sample 2002:01-2012:02, as well as each sub-sample periods. Figures A1 and A2 in Appendix A show the evolution of exchange rates in Latin America and the Caribbean.

3.2. Methodology and Results

3.2.1. Unit root tests

To test for a long run relationship between the stock prices and exchange, the order of integration of each variable must first be examined. Through careful examination of the movements of the data for the countries of the Caribbean and Latin America overtime, it is visible that structural breaks may exists at different points in time (Figures 1-3 in Section 1). If such structural breaks

Table 1: Summary statistics

Variables	Mean±SD		
	Tranquil period (2002/01 to 2008/08)	Crisis period (2008/09 to 2012:02)	Full period (2002/01 to 2012/02)
Stock prices			
Jamaica	0.151±0.048	−0.004±0.049	0.008±0.049
Trinidad and Tobago	0.012±0.031	−00.1±0.031	0.007±0.032
Argentina	0.019±0.071	0.013±0.093	0.016±0.079
Brazil	0.019±0.072	0.007±0.074	0.136±0.073
Chile	−0.126±0.171	−0.134±0.123	−0.134±0.123
Mexico	0.017±0.049	0.010±0.063	0.014±0.054
Exchange rate			
Jamaica	0.005±0.009	0.004±0.015	0.005±0.012
Trinidad and Tobago	−0.000±0.003	0.000±0.003	0.000±0.003
Argentina	0.009±0.066	0.008±0.014	0.009±0.054
Brazil	−0.005±0.055	−0.003±0.047	−0.002±0.542
Chile	−0.063±0.081	−0.069±0.185	−0.069±0.185
Mexico	0.001±0.020	0.004±0.042	0.003±0.030
Interest rate			
Jamaica	−0.007±0.242	−0.056±2.370	−0.007±0.222
Trinidad and Tobago	0.012±0.171	−0.169±0.409	−0.049±0.287
Argentina	−0.384±8.614	0.006±0.921	−0.244±6.928
Brazil	−0.078±0.678	−0.073±0.375	−0.071±0.590
Chile	−0.989±0.489	−2.590±0.579	−2.590±0.680
Mexico	0.007±0.549	−0.094±0.234	−0.026±0.465
Foreign reserves			
Jamaica	0.003±0.059	−0.003±0.075	0.000±0.064
Trinidad and Tobago	0.019±0.047	0.004±0.025	0.014±0.041
Argentina	0.015±0.062	0.001±0.019	0.009±0.054
Brazil	0.022±0.059	0.013±0.019	0.189±0.049
Chile	−0.166±0.131	−0.069±0.101	−0.165±0.131
Mexico	0.009±0.022	0.010±0.037	0.009±0.027

All data is analyzed in logarithms. SD: Standard deviation are not accounted for, it increases the likelihood of failing to reject the null of a unit root (e.g., Lin, 2012). Therefore, unit root tests are sensitive to the alternative to a trend break hypothesis. Consequently, along with the usual Augmented Dickey Fuller (1979) (ADF) unit root test, we also employ the Zivot and Andrews (1992) and the Clemente et al. (1998) unit root tests which both account for the presence of structural breaks in the data.

The Zivot and Andrews (1992) allows for a structural break in the data set which may occur in the intercept or trend or both. The break is determined endogenously as the test supports various criteria for detection. The Zivot and Andrews (1992) can be specified in one of three general forms: Model A accounts for break in intercept only; Model B accounts for break in trend only and Model C accounts for break in both intercept and trend. Consider the following: Suppose the structural shift occurs at a period $1<T_B<T$ in the data set, and given the following AR(1) process;

$$y_t=\varphi+y_{t-1}+\varepsilon_t \tag{1}$$

Then the models are outlined as follows:

Model A

$$y_t = \varphi + \beta t + \tau DU_t(\lambda) + \alpha y_{t-1} + \sum_{i=1}^{k} c_t \Delta y_{t-i} + \varepsilon_t \tag{2}$$

Model B

$$y_t = \varphi + \beta t + \vartheta DT_t(\lambda) + \alpha y_{t-1} + \sum_{i=1}^{k} c_t \Delta y_{t-i} + \varepsilon_t \tag{3}$$

Model C

$$y_t = \varphi + \beta t + \tau DU_t(\lambda) + \vartheta DT_t(\lambda) + \alpha y_{t-1} + \sum_{i=1}^{k} c_i \Delta y_{t-i} + \varepsilon_t \tag{4}$$

Where; φ and t represent the deterministic intercept and trend respectively, $DU_t(\lambda)=1$ if $t>T(\lambda)$ and 0 otherwise; allows for break in intercept and, $DT_t=t-t\lambda$ if $t>t\lambda$, and 0 otherwise; allows for break in trend. Δ is the difference operator and the error term $\varepsilon_t \sim IID(0,\sigma^2)$ is assumed to be identically and independently distributed with mean zero and constant variance.

The null hypothesis of the Zivot and Andrews (1992) test is that y_t has unit root I(1) with no exogenous structural break against the alternative that series is a stationary I(0) with structural break at some unknown point in time. The minimum t-value for the Zivot and Andrews test is selected for the endogenously determined break point $1<T_B<T$. Here the test statistics are larger than that of the ADF since it allows for endogenous break points. The lag length is selected using the Schwarz Bayesian Criterion.

To increase robustness of our results, we also employ the Clemente et al. (1998) test which extends Zivot and Andrews (1992) test to include two structural breaks, allowing for double changes in the mean, modeling both additive outliers (AO) schemes and innovative outliers (IO) schemes. The AO scheme is derived from the following:

$$y_t = \varphi + d_1 DU_{1t} + d_2 DU_{2t} + \tilde{y} \tag{5}$$

Where,

$$\tilde{y} = \sum_{j=0}^{k} \tau_{1t} DT_{B1t-i} + \sum_{i=0}^{k} \tau_{2t} DT_{B2t-i} + \alpha \tilde{y}_{t-1} + \sum_{i=1}^{k} \theta \Delta \tilde{y}_{t-i} + \varepsilon_t \tag{6}$$

The test procedure searches for the minimum t-statistics where $\alpha = 1$, in Equation (6) above. If the breaks are as a result of the IO case then the model is estimated as Equation (7) below:

$$y_t = \varphi + d_1 DU_{1t} + d_2 DU_{2t} + \delta_1 DT_{B1t} + \delta_2 DT_{B2t} + \sum_{i=1}^{k} c_t \Delta y_{t-i} + \varepsilon_t \tag{7}$$

Where DU_{1t} and DU_{2t} are defined as above, and DT_{Bi}; (i =1, 2), are impulse variables equal to 1 if $t=T_{Bi}+1$ and 0 otherwiseT_{Bi} represents the time of the shifts in the data. Once more, the error term $\varepsilon_t \sim IID(0,\sigma^2)$ is assumed to be independent and identically distributed with mean zero and constant variance.

The results are provided in Table 2. The results from the unit root tests with and without structural break(s) confirm that our variables of interest, mm, sp, fx and r are integrated of different orders over time and across countries. Table 2 shows the results of the three unit root tests employed in the analysis; the Augmented Dickey Fuller (1979) tests with no structural break(s), the Zivot and Andrew test (1992) with 1 structural break and the Clemente et al. (1998), with two structural breaks. The results of the Augmented Dickey Fuller (1979) tests show that all variables for all six countries are integrated I(1) series except Brazil's exchange rate and interest rates and Mexico's Exchange rate. The Zivot and Andrews (1992) tests shows that most variables from all countries have at least 1 significant structural break in the data. The structural break for stock prices for Jamaica and Trinidad and Tobago occurs late 2003, while the structural break for the Latin American countries occurred mid 2007 or 2008, approximately the time when the asset bubble started peaking. The structural break in the exchange rate for Jamaica, Argentina, Brazil, Chile and Mexico occurs at late 2008; while for Trinidad and Tobago it occurred late 2009. The structural break for interest rate data is a little less aligned: For Jamaica and Brazil is at early 2010; for Trinidad and Tobago it is early 2009; for Mexico early 2004; Argentina 2005 and Chile 2009. The international foreign reserves variables shows the largest disparity, since the structural breaks for all countries are occurring at different years: Jamaica 2004, Trinidad and Tobago 2005, Brazil 2007, Argentina 2009, Chile 2009 and Mexico 2010. Overall, by taking account of one structural break in the data, the results show that all variables are integrated I(1) series except interest rates for Trinidad and Tobago, Chile, Argentina and exchange rates for Chile and Brazil.

To extend the analysis and increase robustness of the results, we further employ the Clemente et al. (1998) unit root test, with two structural breaks. The results show that all variables except interest rates for Trinidad and Tobago, Mexico and Argentina

Table 2: Unit root tests: ADF, Zivot and Andrews (1992) and Clemente et al. (1998) tests

Variables	Jamaica	Trinidad and Tobago	Argentina	Brazil	Chile	Mexico
Unit root test with no structural breaks						
Stock prices	−3.282	−2.447	−2.602	−2.167	−2.634	−1.634
Exchange rates	−2.640	−3.781	−5.110*	−4.452*	−2.686	−3.637*
Interest rates	−3.076	−0.994	−4.386*	−4.484*	−3.315	−2.776
Foreign reserves	−2.813	−0.606	−1.092	−1.819	—0.899	−2.357
Unit root test with one structural break						
Stock prices	−3.800 [03/11]	−3.369 [03/09]	−4.043 [08/06]	−3.586 [08/06]	−4.217 [07/11]	−4.431 [08/06]
Exchange rates	−3.487 [08/09]	−4.711 [09/07]	−6.994 [08/09]	−7.339* [08/09]	−5.330* [08/08]	−6.025 [08/09]
Interest rates	−4.792 [10/03]	−7.707* [09/01]	−11.766* [05/11]	−5.491 [10/01]	−5.845* [09/01]	−4.003 [04/08]
Foreign reserves	−3.820 [04/02]	−2.919 [05/07]	−2.671 [09/12]	−4.680 [07/02]	−2.554 [09/06]	−3.077 [10/07]
Unit root test with two structural breaks						
Stock prices	−2.456 [03/10, 08/11]	−4.373 [03/08,08/06]	−4.035 [08/09, 09/02]	−3.616 [03/01,05/06]	−3.894 [03/02,09/02]	−3.719 [03/02, 09/01]
Exchange rates	−4.445 [05/08, 08/08]	−4.445 [09/05,10/09]	−3.886 [02/10, 08/07]	−2.905 [04/06,04/01]	−3.895 [03/07,08/09]	−6.756 [08/07,09/11]
Interest rates	−4.649 [04/01, 11/01]	−11.778* [06/01, 08/12]	−4.991 [02/05, 05/09]	−4.337 [03/05, 06/07]	−5.319 [05/06,09/11]	−6.613* [03/08,09/02]
Foreign reserves	−3.770 [03/12, 09/06]	−3.691 [04/04, 05/05]	−5.670* [02/12, 05/12]	−4.043 [02/10 ,06/03]	−4.456 [08/03,11/1]	−2.799 [05/04,09/10]

Critical values for 5% significance level for unit root tests with one and two structural breaks are (−4.800) and (−5.490) respectively. The time of the structural break is given in parenthesis [] following the order [year/month]. *Indicates significance at the 5% level, rejecting the null of a unit root. ADF: Augmented Dickey Fuller

foreign reserves are integrated I(1) with a possible structural break in the trend, intercept or both. The structural breaks for stock prices for Jamaica, Trinidad and Tobago are in 2003 and 2008; Argentina 2008 and 2009 and Chile 2003 2009, while the structural break for Brazil occurred in 2003 and 2005. Clearly, it appears that the first structural break for stock prices in 5 of the 6 countries occurred with the beginning of the asset bubble in 2003, while the second structural break for all countries except Brazil occurred either in 2008 or 2009, approximately the time when the asset bubble peaked. The results for exchange rates using the test with two structural breaks are different across all countries: For Jamaica the break occurs in 2005 and 2008; Trinidad and Tobago 2009 and 2010; both breaks for Brazil occurred in 2004, Mexico 2008 and 2009, Argentina 2002 and 2008 and Chile 2003 and 2008.

The structural breaks for interest rate data using this test shows for Jamaica the breaks occurred in 2004 and 2011; for Trinidad and Tobago the breaks occurred 2006 and 2008; Brazil 2003 and 2006; Mexico 2003 and 2009; Argentina 2002 and 2005 while Chile 2003 and 2006. The international foreign reserves variable also show disparity across countries. Using this test the structural breaks are registered in different years: Jamaica 2003 and 2009; Trinidad and Tobago 2004 and 2005; Brazil 2002 and 2006, Mexico 2005 and 2009; Argentina 2002 2005 and Chile 2008 and 2011. Overall, the first structural break in this second test occurs at a similar point in time to the structural break in the Zivot and Andrew (1992) test in most instances. Overall all three tests indicate that unit roots are present in some of the variables. There was no variable where all three tests results agree on no unit root; and therefore there is no clear evidence that any of variables are I(0).

3.2.2. Bounds tests for cointegration

The results from the unit root tests with and without structural break(s) confirm that our variables of interest, mm, sp, fx and r are integrated of different orders over time and across countries. Therefore, the regular Engle and Granger (1987) and VAR based tests of Johansen (1992) and Johansen and Juselius (1998) are mis-specified given these conditions. To correct this, we follow Lin (2012) by employing the ARDL bounds test approach proposed by Pesaran et al. (2001). This approach solves the problem, as it provides valid test results even if the variables are integrated of different orders. Using this bounds test procedure the order of integration of the variables does not have to be the same; i.e., it accounts for the inclusion of both I(0) and I(1) in the same equation.

Here similar to Lin (2002) we employ the bounds test corresponding to Case v, from Pesaran et al. (2001) which accounts for an unrestricted intercept and an unrestricted trend in the model.

The model is specified in VAR terms as follows:

$$z_t = \alpha + \omega t + \sum\nolimits_{i=1}^{p} \gamma_i z_{t-i} + \varepsilon_t,\ t = 1, 2\ldots,\ T \qquad (8)$$

Where α represents a $(k + 1)$ vector of intercepts/drifts and ω is a vector of $(k + 1)$ trend coefficients. Given this, Pesaran et al.

(2001) derived the following VECM corresponding to Equation (8) above, with being the difference operator:

$$\Delta z_t = \alpha + \omega t + \Pi\, z_{t-1} + \sum\nolimits_{i=1}^{p} \theta_i \Delta z_{t-i} + \varepsilon_t,\ t = 1, 2\ldots, T \qquad (9)$$

Here the $(k+1)\times(k+1)$ matrices $\Pi = I_{k+1} + \sum\nolimits_{i=1}^{p} \mu_j$ and $\theta_i = -\sum\nolimits_{i=1}^{p} \mu_j\ i = 1, 2\ldots, p-1$, as outlined by Lin (2012), contain the long run multipliers and the short run dynamic coefficients of the VECM and z_t is the vector of dependent variables y_t and regressors x_t. After regressing the ARDL model, the bounds test procedure requires conducting an F-test on the joint significance of lagged levels of the variables in the model. The null hypothesis is that the lagged levels of the variables are insignificant i.e., $\Pi = 0$ against the alternative that they are jointly significant i.e., $\Pi = 0$. Under the ARDL model, the F-statistic can no longer be compared to the critical values of the F-tables. Instead, the bounds test provides two asymptotic critical values for which to compare the calculated F-statistic. It provides a lower bound critical value assuming the variables are I(0) and an upper bound critical value assuming the variables are I(1). If the F-statistic is greater than the critical value for the upper bound, the null of no cointegrating relationship can be rejected. If the F-statistic falls between the upper and lower bound then the test is inconclusive and if the F-statistic is lower than the critical value for the lower bound then the null of no cointegrating cannot be rejected.

The results for any long run relationship between the variables using the bounds test for cointegration are provided in Table 3. Here the null hypothesis is the no-long run level relationship. In what follows we restrict our analysis to the conclusive cases.

The results of the bounds test show that stock prices movements significantly impact exchange rate movements in the tranquil period as well as over the entire period of the sample for Argentina and Brazil. Stock prices also significantly affect the exchange rate during the crisis period in Jamaica. These results are evidence of the stock oriented model (Frankel, 1983; Branson, 1993), where a decrease in stock prices reduces the wealth of local investors, which lowers their demand for money; banks react by lowering interest rates which dampens capital inflows, reducing the demand for local currency and therefore depreciates the local currency. According to Lin (2012), this suggests that governments should try to prevent a currency crisis by stimulating economic growth and the expansion of the stock market to attract capital inflow.

On the flip side, exchange rate movements significantly impact stock price movements in Jamaica in the non crisis period and in Trinidad and Tobago in the full sample, evidence of the "flow oriented" models of exchange rates. A fall in the exchange rate increases the cost of production to firms and therefore their output and competiveness. If firms are less competitive this has a negative effect on its stock prices. This appears to be the case in Jamaica, as sharp exchange rate depreciations are relayed to stock prices in crisis period. There is no evidence to suggest any significant impact of exchange rate on stock prices for Jamaica over full sample. The results from the bounds test found no evidence to suggest that either exchange rate movements affect stock price movements or vice versa in Trinidad and Tobago, Chile and Mexico. This coincides with the general literature which states there is no long run relationship between the two (Zhao, 2010; Neih and Lee, 2001; Diamandis and Drakos, 2011).

3.2.3. ARDL GARCH(1,1)

We go one step further to include a GARCH(1,1) component in the ARDL framework to incorporate the impact of volatility in the model, similar to Chen et al. (2013), in modeling the effect of oil prices on global fertilizer prices. We specifically examine the impact of stock price shocks on exchange rates since this is the more prevalent channel in our analysis thus far. The new specification of the model is the same as the one given in (9), but when writing the conditional error correction model where the dependent variable is the first difference of the exchange rates, the disturbance is specified to follow a GARCH(1,1) model. The results are provided in Tables A1-A6 in Appendix A.

By including this volatility factor, the results change drastically and now we find strong relationships: For Jamaica there is evidence of the stock oriented model in the crisis period and over the full period of study. The ARCH effects are statistically significant across all three periods, showing the presence of a significant volatility. Trinidad and Tobago also shows evidence of the stock oriented model in all three samples: The tranquil period, the crisis period and the full period in the study. Once more ARCH effects are statistically significant across all three samples and the GARCH effect is present for the full period only showing that there is a less persistent volatility across all three samples and a more persistent volatility across the full period only. The money market rate is significant in the crisis period and the reserves are only statistically significant over the full period of study.

Table 3: Bounds test for cointegration analysis

Bounds test	Jamaica	Trinidad and Tobago	Argentina	Brazil	Chile	Mexico
Full sample from 2002m1 to 2012m06						
F[fx|sp, mm, R]	4.160	4.160	9.20*	6.410*	3.350	0.610
F[sp|fx, mm, R]	5.440	7.510*	2.920	4.600	2.650	3.190
Tranquil period 2002m1 to 2008m08						
F[fx|sp, mm, R]	2.300	3.900	35.06*	6.620*	4.450	4.020
F[sp|fx, mm, R]	6.300*	2.370	2.17	4.700	1.870	2.130
Crisis period 2008m09 to 2012m05						
F [fx|sp, mm, R]	19.005*	2.690	2.65	3.630	5.06	2.340
F [sp|fx, mm, R]	4.406	5.310	4.23	3.240	2.08	3.230

Critical values are from Pesaran et al. (2001), Table CI (v), Case (v), unrestricted intercept and unrestricted trend, lower bound I (0)=4.87 and upper bound I (1)=5.85 at the 5% level of significance. *Indicates that cointegration exists at the 5% level of significance. The results for Table 3 were done for different lag structures, and the results were robust to that

In Argentina, changes in stock prices have a significant effect in all periods; showing evidence of the stock oriented model, similar to the results from Diamandis and Drakos (2011), who found the same relationship using the US stock market as a transmission variable. The ARCH volatility component is also statistically significant in all three periods, showing evidence of a significant less persistent volatility. The money market rate and the reserves are statistically significant in the sub-samples but not over the full period of study. It may be explained by the existence of an unaccounted structural break that may have disrupted the results in the full sample.

In Brazil, changes in stock prices are significant. Here as well, ARCH effects are present in the tranquil period as well as the full period. The money market rate and the reserves are statistically significant in the full period of study. For Mexico, there is evidence of the stock oriented model in the tranquil period; and the stock index is statistically significant. The money market rate is statistically significant in the full period of study. The reserves are statistically significant in the tranquil and full period, but insignificant during crisis period, and the ARCH element is statistically significant only in the crisis period. There is evidence of the stock oriented model in Chile in the tranquil period as well as in the full period but not in the period of crisis. The ARCH effect is present over the full period while the GARCH effect is present in the tranquil period only. Interest rates are statistically significant in in the tranquil period as well as the full period.

4. CONCLUSIONS

We also study the same four Latin American countries as in Diamandis and Drakos (2011): Argentina, Brazil, Chile and Mexico. Following Lin (2012), who examined the same issue in six Asian emerging markets and also employed the ARDL model bounds test approach proposed by Pesaran et al. (2001), we also include interest rates and NIR variables in our analysis to avoid any omitted variable bias. We extend Diamandis and Drakos (2011) and Lin (2012) by expanding the ARDL model including a GARCH component to examine the impact of volatility. First, we use the structural break unit root tests of Zivot and Andrews (1992) and Clemente et al. (1998) to show a significant structural break in the exchange rate, stock prices and our other control variables around the time of the 2008 crisis in all analysed countries, leading us to check our results in three periods: The full sample and in two subsamples before and after 2008. Our results from the bounds test showed a very mild relationship between both variables in Jamaica, Trinidad and Tobago, Argentina and Brazil, but we cannot find any relationship in the other countries as in Diamandis and Drakos (2011). However, when we include the GARCH component in the ARDL framework our results changed drastically: Stock prices significantly impacted the exchange rate in the tranquil sub-period and the full period for Jamaica, over all three periods for Trinidad and Tobago and in the tranquil period for Argentina, Mexico and Chile. This shows the importance of incorporating volatility explicitly in the model. Our results have the policy implications that governments in the previous countries should try to prevent a currency crisis by stimulating economic growth and the expansion of the stock market to attract capital inflow as in Lin (2012).

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APPENDIX A

Table A1: Results from the ARDL-GARCH for Jamaica

Jamaica	Tranquil period (2002/01 to 2008/08)		Crisis period (2008/09 to 2012:02)		Full period (2002/01 to 2012/02)	
	Coefficient	P value	Coefficient	P value	Coefficient	P value
Δsp_{t-1}	−0.028 (0.012)	0.018*	0.000 (0.001)	0.715	−0.028 (0.005)	0.000*
Δmm_{t-1}	−0.003 (0.002)	0.113	−0.000 (0.000)	0.929	−0.005 (0.001)	0.002*
ΔR_{t-1}	0.016 (0.007)	0.016*	−0.001 (0.001)	0.001*	0.002 (0.005)	0.664
sp_{t-1}	0.003 0.002	0.135	0.076 (0.007)	0.000*	−0.004 (0.002)	0.000*
mm_{t-1}	0.001 (0.002)	0.484	0.014 (0.002)	0.000*	0.001 (0.001)	0.342
R_{t-1}	0.001 (0.001)	0.305	0.001 (0.002)	0.657	0.002 (0.000)	0.001*
fx_{t-1}	−0.002 (0.007)	0.781	−0.084 (0.011)	0.000*	−0.005 (0.002)	0.038*
ARCH	1.538 (0.427)	0.000*	2.247 (0.897)	0.012*	2.696 (0.320)	0.000*
GARCH	0.108 (0.063)	0.088	0.012 (0.051)	0.829	0.016 (0.019)	0.393

*Indicates significance at the 5% level. ADRL: Autoregressive distributed lag, GARCH: Generalized autocorrelation conditional heteroskedasticity

Table A2: Results from the ARDL-GARCH for Trinidad and Tobago.

Trinidad and Tobago	Tranquil period (2002/01 to 2008/08)		Crisis period (2008/09 to 2012:02)		Full period (2002/01 to 2012/02)	
	Coefficient	P value	Coefficient	P value	Coefficient	P value
Δsp_{t-1}	0.000 (0.001)	0.715	−0.016 (0.012)	0.178	−0.005 (0.001)	0.000*
Δmm_{t-1}	0.000 (0.000)	0.929	0.000 (0.000)	0.548	0.000 (0.000)	0.666
ΔR_{t-1}	−0.002 (0.001)	0.001	0.000 (0.004)	0.869	0.000 (0.002)	0.793
sp_{t-1}	−0.003 (0.000)	0.000*	0.076 (0.007)	0.000*	−0.001 (0.000)	0.034*
mm_{t-1}	−0.006 (0.001)	0.000*	0.014 (0.002)	0.000*	0.000 (0.000)	0.596
R_{t-1}	0.004 (0.000)	0.000*	0.001 (0.002)	0.657	0.001 (0.000)	0.027*
fx_{t-1}	−0.033 (0.003)	0.000*	−0.084 (0.114)	0.000*	−0.006 (0.003)	0.056
ARCH	4.147 (0.963)	0.000*	2.247 (0.897)	0.012*	1.325 (0.329)	0.000*
GARCH	0.012 (0.021)	0.560	0.010 (0.050)	0.829	0.314 (0.048)	0.000*

*Indicates significance at the 5% level. ADRL: Autoregressive distributed lag, GARCH: Generalized autocorrelation conditional heteroskedasticity

Table A3: Results from the ARDL-GARCH for Argentina.

Argentina	Tranquil period (2002/01 to 2008/08)		Crisis period (2008/09 to 2012:02)		Full period (2002/01 to 2012/02)	
	Coefficient	P value	Coefficient	P value	Coefficient	P value
Δsp_{t-1}	0.025 (0.020)	0.220	0.001 (0.010)	0.865	0.012 (0.007)	0.104
Δmm_{t-1}	−0.001 (0.000)	0.000*	0.000 (0.001)	0.725	−0.000 (0.001)	0.808
ΔR_{t-1}	0.000 (0.017)	0.998	−0.003 0.048)	0.995	−0.046 (0.021)	0.030*
sp_{t-1}	−0.025 (0.004)	0.000*	0.001 (0.005)	0.851	0.003 (0.254)	0.229
mm_{t-1}	0.019 (0.002)	0.000*	0.035 (0.008)	0.000*	−0.001 (0.001)	0.543
R_{t-1}	0.029 (0.002)	0.000*	0.003 (0.001)	0.007*	−0.000 (0.000)	0.445
fx_{t-1}	−0.547 (0.060)	0.000*	0.005 (0.025)	0.839	−0.002 (0.006)	0.714
ARCH	1.954 (0.541)	0.000*	1.424 (0.529)	0.007	2.709 (0.292)	0.000*
GARCH	−0.014 (0.035)	0.692	-	-	0.085 (0.048)	0.077

*Indicates significance at the 5% level. ADRL: Autoregressive distributed lag, GARCH: Generalized autocorrelation conditional heteroskedasticity

Table A4: Results from the ARDL-GARCH for Brazil

Brazil	Tranquil period (2002/01 to 2008/08)		Crisis period (2008/09 to 2012:02)		Full period (2002/01 to 2012/02)	
	Coefficient	P value	Coefficient	P value	Coefficient	P value
Δsp_{t-1}	−0.131 (0.049)	0.007*	−0.053 (0.146)	0.715	−0.029 (0.051)	0.558
Δmm_{t-1}	−0.007 (0.004)	0.084	0.011 (0.034)	0.740	−0.004 (0.005)	0.381
ΔR_{t-1}	−0.102 (0.052)	0.049*	−0.891 (0.483)	0.065	−0.081 (0.064)	0.198
sp_{t-1}	−0.005 (0.018)	0.755	0.172 (0.108)	0.114	0.008 (0.020)	0.067
mm_{t-1}	0.012 (0.017)	0.497	0.078 (0.088)	0.375	0.029 (0.015)	0.047*
R_{t-1}	0.001 (0.005)	0.912	0.049 (0.030)	0.099	0.005 (0.005)	0.038*
fx_{t-1}	−0.026 (0.034)	0.437	0.313 (0.085)	0.000*	−0.015 (0.034)	0.668
ARCH	1.392 (0.488)	0.004*	0.302 (0.315)	0.314	1.345 (0.375)	0.000*
GARCH	−0.0968 (0.074)	0.193	0.536 (0.581)	0.356	-	-

*Indicates significance at the 5% level. ADRL: Autoregressive distributed lag, GARCH: Generalized autocorrelation conditional heteroskedasticity

Table A5: Results from the ARDL-GARCH for Mexico

Mexico	Tranquil period (2002/01 to 2008/08)		Crisis period (2008/09 to 2012:02)		Full period (2002/01 to 2012/02)	
	Coefficient	P value	Coefficient	P value	Coefficient	P value
Δsp_{t-1}	0.011 (0.051)	0.822	0.005 (0.131)	0.967	0.016 (0.054)	0.774
Δmm_{t-1}	−0.001 (0.004)	0.730	0.034 (0.052)	0.511	−0.002 (0.005)	0.782
ΔR_{t-1}	−0.073 (0.089)	0.413	0.133 (0.255)	0.602	−0.018 (0.099)	0.857
sp_{t-1}	−0.010 (0.004)	0.020*	−0.223 (0.164)	0.175	−0.008 (0.005)	0.117
mm_{t-1}	−0.025 (0.015)	0.093	−0.069 (0.094)	0.467	−0.029 (0.012)	0.013*
R_{t-1}	0.017 0.005	0.002*	0.084 0.055	0.125	0.015 (0.004)	0.002*
fx_{t-1}	−0.139 (0.055)	0.011*	−0.330 (0.163)	0.043*	−0.117 (0.039	0.003*
ARCH	−0.089 (0.142)	0.531	0.440 (0.372)	0.236	0.560 (0.117)	0.000*
GARCH	0.241 (1.675)	0.886	0.329 (0.365)	0.367	0.316 (0.182)	0.083

*Indicates significance at the 5% level. ADRL: Autoregressive distributed lag, GARCH: Generalized autocorrelation conditional heteroskedasticity

Table A6: Results from the ARDL-GARCH for Chile.

Chile	Tranquil period (2002/01 to 2008/08)		Crisis period (2008/09 to 2012:02)		Full period (2002/01 to 2012/02)	
	Coefficient	P value	Coefficient	P value	Coefficient	P value
Δsp_{t-1}	0.096 (0.079)	0.226	−0.269 (0.158)	0.088	−0.017 (0.007)	0.826
Δmm_{t-1}	−0.026 (0.124)	0.037	−0.007 (0.287)	0.795	−0.002 (0.101)	0.860
ΔR_{t-1}	−0.057 (0.074)	0.445	−0.224 (0.164)	0.172	−0.007 (0.101)	0.948
sp_{t-1}	−0.062 (0.023)	0.007*	−0.037 (0.069)	0.590	−0.071 (0.254)	0.005*
mm_{t-1}	0.001 (0.010)	0.080	−0.007 (0.008)	0.375	−0.003 (0.005)	0.532
R_{t-1}	0.062 (0.020)	0.002*	0.056 (0.391)	0.151	0.062 (0.021)	0.003*
fx_{t-1}	−0.187 (0.060)	0.002*	−0.184 (0.117)	0.116	−0.178 (0.061)	0.003*
ARCH	0.201 (0.196)	0.304	0.250 (0.398)	0.630	0.371 (0.116)	0.001*
GARCH	0.722 (0.314)	0.020	0.614 (0.639)	0.336	-	-

*Indicates significance at the 5% level. ADRL: Autoregressive distributed lag, GARCH: Generalized autocorrelation conditional heteroskedasticity

Figure A1: Exchange rate in the Caribbean and Latin America

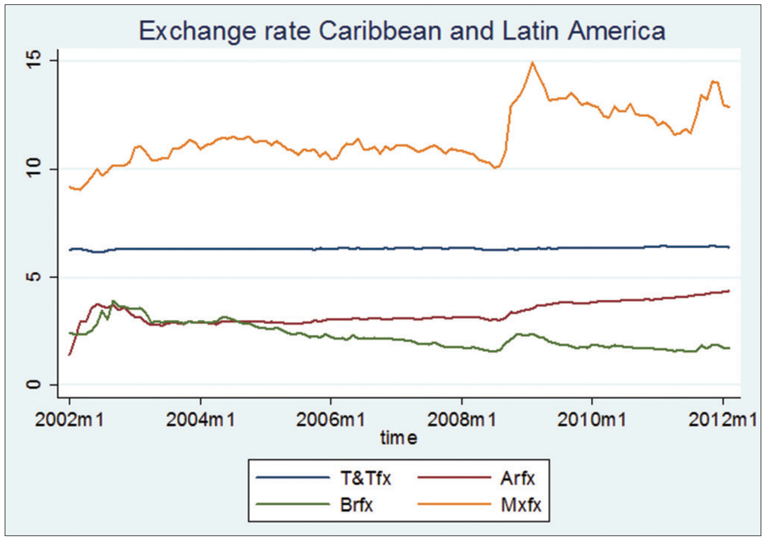

Figure A2: (a and b) Exchange rate in Jamaica and Chile