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Article

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The impact of the global financial crisis on the efficiency and performance of Latin American stock markets*

El impacto de la crisis financiera en la eficiencia y desempeño de mercados latinoamericanos de acciones

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WING-KEUNG WONG*****

Abstract

We analyze the impact of the most recent global financial crisis (GFC) on the seven most important Latin American stock markets. Our mean-variance analysis shows that the markets are significantly less volatile and, in general, investors prefer to invest in the post-GFC period. Our results from the Hurst exponent and runs and variance-ratio tests show that the randomness and efficiency

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have been improved after the GFC. The stochastic dominance test shows that the markets are efficient, there is no arbitrage opportunity due to the GFC in our studying period, and, in general, investors prefer investing in the post-GFC period. The results confirm that the 2008 global financial crisis does have some positive impacts on Latin American stock markets. Our findings provide important information for investors and market regulators in their decision making in investment and setting regulations.

Key words: *Latin American stock markets, randomness, market efficiency, stochastic dominance.*

JEL Classification: *G14, G15.*

Resumen

Este trabajo analiza el impacto de la crisis financiera global del 2008 en siete mercados accionarios latinoamericanos. Usando diversas técnicas, se demuestra que los mercados son menos volátiles, más eficientes, que no hay oportunidades de arbitraje y que los inversionistas prefieren invertir en el periodo posterior a la crisis. Nuestros hallazgos proveen información importante para inversionistas y reguladores.

Palabras clave: *Mercados accionarios latinoamericanos, aleatoriedad, mercados eficientes, dominancia estocástica.*

Clasificación JEL: *G14, G15.*

1. INTRODUCTION

Financial crises normally have a strong impact not only on developed but also on developing countries, creating high volatility in the prices of financial assets. A recent example was the financial crises that started in 2007¹. During the turbulent period of a crisis, prices do not reflect full information, creating a challenge for the efficient market theory. According to Fama (1970), market efficiency suggests that at any given time, prices fully reflect all available information on a particular stock and/or market, and in this way, it is impossible for investors to consistently earn excess or abnormal returns using information from the market.

¹ The introduction of regulation and increased government intervention based on Fannie Mae and Freddie Mac that subsidizing people to purchase homes that they don't have financial capacity to pay lead to increasing exposures to aggregate risks and to the financial crises that began in late 2007 (Oesterle, 2010).

Since the 1990s, most Latin American countries have implemented liberalization programs to create a competitive market (Arbelaez and Milman, 2000) and the impact of opening the markets to foreign investors made Latin American stock markets become more efficient (Groenewold and Ariff, 1998; Kim and Singal, 2000; Füss, 2005), and part of this success is also related to improvement in market regulation (Antoniou, *et al.*, 1997). Market liberalization is a long process. So, it is important to examine market efficiency at different stages of a country's development. Only in this way it is possible to capture the impact of changes in market regulation across time. If the problem of inefficiency is not rectified by the authorities and policy makers, this could seriously limit the stock market's ability to allocate funds to the most productive sectors of the economy and could potentially hamper long-term growth (Mookerjee and Yu, 1999).

Ball (2009) criticizes regulators' for excessive reliance on the efficiency of markets, which has led to a lax regulatory framework for capital markets. Complement this information, Claessens and Kodres (2014) argue that it is necessary to increase regulation to create a stable and efficient financial system.

Financial crises are normally associated with a negative impact in all areas of the economy, including stock markets. For example, Furceri and Mourougane (2012) note that a financial crisis negatively and permanently affects potential output, and Bassanini and Duval (2009) find that financial crisis can lead to an increase in the structural unemployment rate for economies with rigid labor market institutions.

To investigate whether our conjecture that financial crises could have positive impact to stock markets, we first apply the mean-variance (MV) criterion to examine the performance of the most important Latin American stock markets, including Argentina, Chile, Colombia, Ecuador, Mexico, Peru and Brazil before and after the global financial crisis. Thereafter, we employ the Hurst exponent, the runs test, and the multiple variation ratio test to analyze whether the markets have improved their randomness after the recent global financial crisis (GFC). We also employ the stochastic dominance (SD) test to investigate investors' preferences for the markets before and after the GFC and check whether there is any arbitrage opportunity in the markets and whether the markets are efficient. We state our findings, the inferences from our findings, and our contribution to the literature in our Conclusion Section.

The paper is organized as follows. The next section summarizes the relevant literature review and Section 3 describes the data and presents the methodology of the different statistics. Section 4 discusses our empirical results, Section 5 concludes and draw inference from our findings.

2. BACKGROUND AND RELEVANT LITERATURE

The efficient market theory has been a major topic in the financial literature since 1970. Analyzing market efficiency is very important because this

will have implications for financial theories and investment strategies, and so academicians, speculators, investors, and regulatory authorities are interested in this issue.

The literature on stock market efficiency in Latin American countries is mixed. For example, using data from 1975 to 1991, Urrutia (1995) finds that the stock markets of Argentina, Brazil, Chile, and Mexico are not efficient, but Ojah and Karemera (1999) document that Mexico, Brazil, and Argentina are weak-form efficient. Hinich and Patterson (1985) apply a test can distinguish between white noise and purely random noise to return of 15 common stocks and conclude that these daily stock returns are produced by a nonlinear process. Mexico has also been analyzed in terms of being weak-form efficient by Bonilla, *et al.* (2011) who find that all of the return series are characterized by a few periods with highly significant non-linearity.

Using a new measure for capital market efficiency to estimate the correlation structure of the long-term and short-term memory returns and local herding behavior for the 41 stock indices, Kristoufek and Vosvrda (2013) find that efficiency is dominated by European stock indices and the less efficient markets are mainly in Latin America, Asia, and Oceania. In addition, Kristoufek and Vosvrda (2014) investigate the efficiency of 38 stock market indices across the world and find that the most efficient markets are in the Eurozone (the Netherlands, France, and Germany) while the least efficient ones are in Latin America (Venezuela and Chile).

Another interesting tool is to use the Hurst exponent to analyze the efficiency in the stock indices. Cajueiro and Tabak (2004) use this tool to examine the US, Japan, and 11 emerging markets and find that the Asian equity markets show greater inefficiency than those of Latin America markets (except Chile) while the developed markets are of the most efficiency. In addition, Duarte and Pérez-Iñigo (2014) analyze the 5 principal stock markets in Latin America and find that the markets have changed from inefficient to efficient in 2007 (México), 2008 (Brazil and Colombia), 2011 (Chile) and 2012 (Perú).

So far, only a few studies analyze the impact of financial crises or crashes on market efficiency. Moreover, these studies do not focus on Latin American countries. For example, examining the weak-form efficiency of eight emerging Asian stock markets during the pre-crisis (1990-1997) and post-crisis (1998-2004) periods, Hoque, *et al.* (2007) document that financial crises have no significant effect on the degree of efficiency of stock markets from Hong Kong, Indonesia, Malaysia, Philippines, Singapore, and Thailand because they remain inefficient even after the crisis, while the opposite occurs for Korea. Taiwan is the only market that has recorded improved efficiency from the pre-crisis to post-crisis period. Kim and Shamsuddin (2008) also note that the Asian crisis does not coincide with a significant change in the level of market efficiency, both for the efficient (Hong Kong, Japan, Korea, and Taiwan) and inefficient groups (Indonesia, Malaysia, and Philippines), with Singapore and Thailand being the exceptional cases that attain efficiency after the crisis.

Investigating the impact of the 1997 financial crisis on the efficiency of the eight Asian stock markets, Lim, *et al.* (2008) find that the crisis affects the efficiency of most Asian stock markets, in which Hong Kong is the hardest hit, followed by the Philippines, Malaysia, Singapore, Thailand, and Korea, but most of the markets recover in the post-crisis period in terms of improved market efficiency. According to the authors, during turbulent crisis period efficiency sometimes decreases, but when the crisis disappears, markets become more efficient than before the crisis.

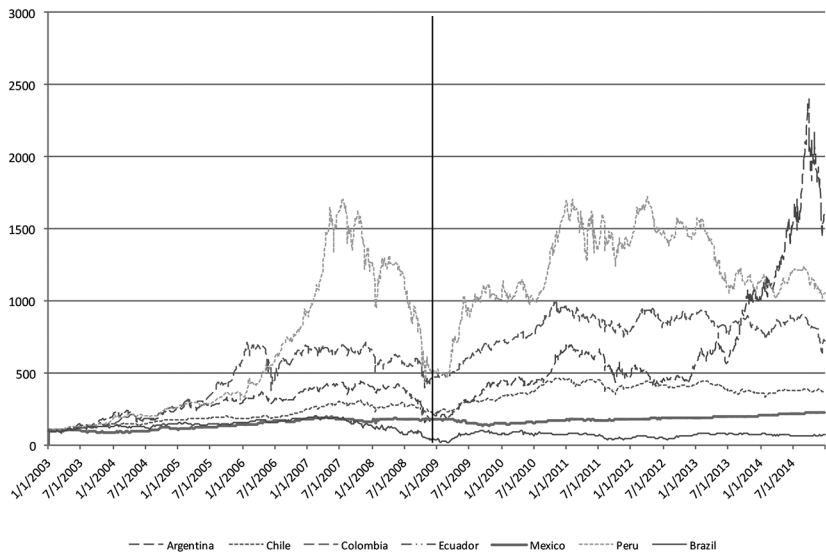
The impact of the financial crisis has also been analyzed from other perspectives besides the efficiency of markets. For example, analyzing diversification opportunities during the most recent financial crisis, Huang, *et al.* (2016) find stronger cross-asset linkages and fewer diversification opportunities. Investigating the worldwide contagion effects using three different econometric models, Morales and Andreosso-O'Callaghan (2014) do not find any significant evidence supporting contagion effects derived from the US stock market, neither worldwide nor regionally. Mayordomo, *et al.* (2014) find asymmetries in commonalities around financial distress episodes such that the effect of market liquidity is stronger when prices in the CDS market increase. In addition, Climent and Meneu (2003) analyse the effects of the crisis on the relationships between the Southeast Asian stock markets and the stock markets of Europe, North America, and Latin America. They find that the integration between Asian markets and other international stock markets increase after the crisis. As far as we know, no study has analyzed the impact of the most recent global financial crisis in terms of stock market efficiency in Latin American stock markets. Thus, our paper bridges the gap to contribute to the literature in this area.

3. DATA AND METHODOLOGY

3.1. Data

The data from January 1, 2003 to December 31, 2014 used in our study comprise the daily closing prices of Latin American stock markets, including Argentina, Chile, Colombia, Ecuador, Mexico, Peru, Brazil. These countries were chosen because they are among the most important economies in the region. These data are collected from Datastream. We plot the daily indices of Latin American stock markets in Figure 1 and use January 1, 2009 as the cut-off point in our study because it is clear from Figure 1 that there is a significant breakpoint on January 1, 2009 for most, if not all, of the indices being studied in our paper due to the recent global financial crisis. In addition, many studies, for example, Karim and Karim (2012), classify the period after January 1, 2009 to be the post-subprime crisis period. In this paper we will analyze the behaviors of log-return $R_t = \log(P_t / P_{t-1})$ for the markets where P_t is the stock index price at time t .

FIGURE 1
TIME SERIES PLOTS OF LATIN AMERICAN STOCK MARKET INDICES FROM
JANUARY 2003 TO DECEMBER 2014



Note: For easier comparison, we fix all values at the same basis of 100 on January 1, 2003.

To test whether January 1, 2009 is a good cut-off point used in this paper, we first conduct the Hansen (2000) “Sampled splitting” test to access whether there is any break in the period from January 1, 2008 to December 31, 2009 and exhibit the results Appendix II. According to Hansen’s (2000) method, we let the MSCI World Index be the threshold variable, and all the seven Latin America countries be the dependent variables. We find that threshold estimation of all Latin America countries occur in July 2008. Thus, we re-set the breakpoint at July 2008 and re-do all our analysis. Thereafter, we apply the heteroscedasticity and autocorrelation consistent (HAC) estimator (Newey and West, 1987; Andrews, 1991) to test whether the breakpoint we used is accurately. We exhibit the results in Appendix III. From the table, we find that the breakpoint of most of the countries occur between 2008 and 2009. Similarity, we re-set the breakpoint for each of the country and re-do all our analysis. We find that the conclusion drawn from the new breakpoints obtained by using HAC estimator is the same as the one (January 1, 2009) we are using. In this connection, we conclude that the results are not sensitive to the exact breakpoint selected if the breakpoint is not too far away from the breakpoint we selected, and thus, we keep the results by using January 1, 2009 to be the cut-off point in this paper.

3.2. Methodology

3.2.1. Mean-Variance Criterion

The mean-variance (MV) criterion for risk averters (Markowitz, 1952) is that for any two returns X and Y with means μ_X and μ_Y and standard deviation σ_X and σ_Y , respectively, X is said to dominate Y by the MV criterion for risk averters, if $\mu_X \geq \mu_Y$ and $\sigma_X \leq \sigma_Y$ in which the inequality holds in at least one of the two. If the above statements are not rejected with at least one strict inequality relationship holding, then one could conclude that X dominates Y significantly by the MV rule. Wong (2007) has proved that if both X and Y belong to the same location-scale family or the same linear combination of location-scale families, and if X dominates Y by the MV criterion, then risk averters will attain higher expected utility by holding X than Y . The theory can be extended to non-differentiable utilities; see Wong and Ma (2008) for details.

3.2.2. Hurst exponent

The Hurst exponent was first developed by Hurst (1951), who studies the statistical properties of the Nile River overflow. Along with the sharp increase in Hurst exponent applications, the methods of exponent calculations have also been improved. In this study, we use the most popular method –rescaled range (R/S) analysis– to estimate the Hurst exponent. It provides a direct estimation of the Hurst exponent that can be used to test the state of randomness for a time series. It can also be used to reveal the existence of long-term dependence that, in turn, can be used to test whether the time series follows a random walk model.

Given a time series with n observations X_1, X_2, \dots, X_n , the R/S statistic is defined as $R/S(n) = \frac{1}{S} \left[\max_k \sum_{i=1}^k (X_i - \bar{X}) - \min_k \sum_{i=1}^k (X_i - \bar{X}) \right]$ for any $1 \leq k \leq n$, in which

\bar{X} is the mean and $S = \sqrt{\frac{1}{n} \sum_{i=1}^n (X_i - \bar{X})^2}$ is the standard deviation from the mean.

With this R/S value, Hurst finds the generalization formula: $E[R/S] = Cn^H$ as $n \rightarrow \infty$, in which H is the Hurst exponent. The value of H can be obtained by running the following simple linear regression over a sample with increasing time horizons:

$$(1) \quad \log(R/S) = \log(C) + H \log(n)$$

In finance, the Hurst exponent is used to measure the efficiency of markets such that a value of the Hurst exponent $H=0.5$ is often required by the efficient market hypothesis. When the time series is persistent, H will be greater than

0.5, and when it is anti-persistent, H will be less than 0.5. If the time series is white noise, $H = 0$, while if the series follows a simple linear trend, then $H = 1$. We note that H must lie between 0 and 1.

3.2.3. *Runs test*

The runs test is a nonparametric test first introduced by Bradley (1968) to determine whether successive price changes are independent. The test is computed based on the signs of deviations from the median observation. The runs test for randomness is used to test the hypothesis that a series of numbers is random based on whether a set of sequential values (called runs) are either above or below the mean. To carry out the test, the total number of runs is computed along with the number of positive and negative values. A positive (negative) run is a sequence of values greater (less) than zero. We can then test whether the number of positive and negative runs is distributed equally in time.

To perform this test, in this paper we let n_+ (n_-) be the number of returns that equals or exceeds (below) the mean. Too many or too few runs in the sequence are the result of negative and positive autocorrelation, respectively. Under the null hypothesis of randomness or independence, the test of the randomness hypothesis can be constructed by comparing the observed number of runs (U) with the expected number of runs (μ_U) and standard deviation (σ_U). It has been shown that, for large sample sizes where both n_+ and n_- are greater than twenty, the standardized test statistic:

$$(2) \quad Z = \frac{U - \mu_U}{\sigma_U}$$

is approximately normally distributed with $\mu_U = \frac{2n_+n_-}{n} + 1$ and

$$\sigma_U = \sqrt{\frac{2n_+n_-(2n_+n_- - n)}{n^2(n-1)}} \quad \text{in which } n = n_+ + n_-.$$

Multiple variance ratio test

Suppose we have the time series $\{X_t\} = (X_0, X_1, X_2, \dots, X_T)$ satisfying:

$$(3) \quad \Delta X_t \equiv X_t - X_{t-1} = \mu + \varepsilon_t$$

Then, the variance of the increments X_t increases linearly in its observation intervals under the random walk hypothesis. That is, $\text{Var}(X_t - X_{t-q})$ is q times of $\text{Var}(X_t - X_{t-1})$. The variance ratio can then defined as:

$$(4) \quad VR(q) = \frac{\frac{1}{q} \text{Var}(X_t - X_{t-q})}{\text{Var}(X_t - X_{t-1})} = \frac{\sigma^2(q)}{\sigma^2(1)}$$

such that under the null hypothesis that $\{X_t\}$ follows the random walk model as stated in Equation (3), we have $VR(q) = 1$. Using this property, Lo and MacKinlay develop the asymptotic distribution of the estimated variance ratios and provide two test statistics, $Z(q)$ and $Z^*(q)$ ² under the null hypothesis of a homoscedastic and heteroskedastic random walk, respectively. If the null hypothesis is correct, both $Z(q)$ and $Z^*(q)$ have asymptotic standard normal distributions. To improve the test developed by Lo and MacKinlay, Chow and Denning (1993) introduce the multiple variance ratio tests by considering a set of m tests $\{M_r(q_i) | i = 1, 2, \dots, m\}$ associated with the set of aggregation intervals $\{q_i | i = 1, 2, \dots, m\}$. Under the random walk null hypothesis, there are multiple sub-hypotheses $H_{0i} : M_r(q_i) = 0$ for all $i = 1, 2, \dots, m$; $H_{1i} : M_r(q_i) \neq 0$ for any $i = 1, 2, \dots, m$. The rejection of at least one $H_{0i} (i = 1, 2, \dots, m)$ implies rejection of the random walk model for the series being tested. Since the random walk hypothesis is rejected if any of the estimated variance ratios are significantly different from one, the largest absolute value of the test statistics:

$$(5) \quad Z_1^*(q) = \text{Max}[Z(q_1), Z(q_2), \dots, Z(q_m)], Z_2^*(q) = \text{Max}[Z^*(q_1), Z^*(q_2), \dots, Z^*(q_m)],$$

will be considered. The rules of the decision then become $P\{Z_j^*(q) \leq \text{SMM}(\alpha; m; N)\} \geq 1 - \alpha$, $j = 1, 2$, where $\text{SMM}(\alpha; m; N)$ is the upper α point of the Studentized Maximum Modulus (SMM) distribution (Richmond, 1982) with parameters m (the number of variance ratios) and N (sample size) degrees of freedom. When N approaches infinity, we have $\lim_{N \rightarrow \infty} \text{SMM}(\alpha; m; \infty) = Z_{\alpha^*/2}$ where $Z_{\alpha^*/2}$ is the critical value under standard normal distribution with $\alpha^* = 1 - (1 - \alpha)^{1/m}$. Thus, if $Z_1^*(q)$ ($Z_2^*(q)$) is greater than the $\text{SMM}(\alpha; m; N)$, then the random walk hypothesis is rejected under the homoscedastic (heteroskedastic) assumption. The critical values of $Z_1^*(q)$ and $Z_2^*(q)$ are the same. When N is large, the SMM critical values at $m = 4$ and α are 2.23, 2.49 and 3.02 at the 10%, 5% and 1% levels of significance, respectively.

² Readers may refer to Lo and MacKinlay (1988) for the equations of $Z(q)$ and $Z^*(q)$.

3.2.4. Stochastic Dominance Test

The weakness of using the MV model is that it uses only mean and variance to characterize a distribution that may ignore important information in the distribution (Wong, 2007). To circumvent the limitations of the MV model, academics, for example, Hanoch and Levy (1969), recommend to use stochastic dominance approach that provides a general set of rules for evaluating the performance of financial assets. The stochastic dominance approach has been demonstrated to be a powerful tool in both theory and applications (Levy 2015, Sriboonchita, *et al.*, 2009). Stochastic dominance is practically useful. There are several consistent SD tests, for example, Davidson and Duclos (DD, 2000), Barrett and Donald (2003), and Linton *et al.* (2005). Bai, *et al.* (2015) extend the DD test by deriving the limiting process of stochastic dominance statistics when the underlying processes are dependent or independent. We apply the tests developed by Davidson and Duclos (2000) and Bai, *et al.* (2015) in our paper because the tests have been demonstrated to be powerful, robust to non-iid data and, yet, not conservative in size (Lean, *et al.*, 2008).

Supposing that X and Y represent two series of returns that have a common support of $\Omega = [a, b]$, with $a < b$, their cumulative distribution functions (CDFs), F and G , and their corresponding probability density functions (PDFs), f and g , respectively, we let

$$(6) \quad H_0 = h, H_j(x) = \int_a^x H_{j-1}(t) dt \quad \text{for } h = f, g; H = F, G; j = 1, 2, 3,$$

for $h = f, g; H = F, G$; and for any integer j where the integral of H_{j-1} is the j^{th} order cumulative distribution function (CDF).

According to Quirk and Saposnik (1962), Fishburn (1964), Hanoch and Levy (1969), Levy (2015) and Guo and Wong (2016), we define the SD rule as follows:

Definition 1: X dominates Y by n SD, denoted $X \succ_n Y$ if and only if $F_n(x) \leq G_n(x)$ (and $F_k(b) \leq G_k(b)$ for any $k = 1, \dots, n-1$ if $n > 2$) for all possible returns x , with the strict inequality holds for at least one interval of x and $\mu_X \geq \mu_Y$ if $n > 2$.

When $n = 1, 2$, and 3 , we call FSD, SSD, and TSD be the first, second, third order stochastic dominance. We discuss more on the stochastic dominance test in Appendix A1. Readers may refer to Bai, *et al.* (2015) for more information on the test.

4. EMPIRICAL RESULTS

We apply mean-variance criterion, Hurst exponent, runs test, multiple variance ratio tests, and stochastic dominance analysis for our empirical study. We first discuss the results of mean-variance criterion.

4.1. Mean-Variance Criterion

Table 1 provides a summary of the descriptive statistics, including sample means, standard deviations, skewness, kurtosis, Jarque-Bera, t, and F statistics of Latin American daily stock returns for both pre- and post-GFC periods are reported in the table. The aim of our study is to test whether the global financial crisis could improve the performance of stock markets. One could provide an answer by checking whether the mean return after the financial crisis is higher and the volatility is smaller.

From Table 1, we find that the mean returns between the pre- and post-GFC periods are not significantly different for all countries being studied in our paper, since all the t statistics are not statistically significant. For the standard deviations of the stock returns between the pre- and post-GFC periods, the F statistic shows that the standard deviations of the stock returns in Argentina are not significantly different between the pre- and post-GFC periods, the standard deviation of the stock returns in the pre-GFC period in Brazil is significantly smaller than that in the post-GFC period, and the returns of all other stock markets are significantly smaller in the post-GFC period.

The result implies that except Argentina and Brazil, all markets become significantly less volatile in the period after the financial crisis. Hence, using the MV criterion (Markowitz, 1952), we conclude that there is no dominance between the pre and post-GFC periods for Argentina, investors prefer to invest in the pre-GFC period for Brazil, and prefer to invest in the post-GFC period for the remaining stock markets and gain higher expected utility under certain conditions (Wong, 2007; Guo, *et al.*, 2018). Thus, in general, our results show that the global financial crisis had a positive impact on Latin American stock markets in the sense that, in general, the stock markets in Latin America are significantly less volatile and risk averters prefer to invest in all of the markets studied in our paper and gain higher expected utility under certain conditions in the post-GFC period.

One may think that since the Jarque-Bera rejects normality for the data, if there exist time dependence or heteroscedasticity, the results of the t and F tests are not valid³. We note that since the log returns are stationary and the t and F test statistics satisfy all the conditions in the central limit theorem for strong mixing stationary sequence (Ibragimov and Maslova, 1971), the t and F test statistics follow asymptotic normal distributions, and thus, the results and inferences made by the t and F tests used in our paper are valid.

³ We would like to show our appreciation to the anonymous referee for his/her helpful comment to improve our paper.

TABLE 1
DESCRIPTIVE STATISTICS FOR LATIN AMERICAN STOCK MARKETS

Country		Mean	Std deviation	Skewness	Kurtosis	JB Statistic	t test/F test
Argentina	Pre	0.1161	0.3012	-0.7285***	5.9598***	2454.53***	-1.2539
	Post	0.3337	0.3104	-0.4581***	3.3170***	772.19***	0.9418
Chile	Pre	0.131	0.1304	-0.0725	14.7039***	14099.73***	0.6839
	Post	0.0822	0.1207	-0.3606***	5.8251***	2246.53***	1.1671***
Colombia	Pre	0.2492	0.2538	-0.3356***	11.3185***	8383.16***	1.5151
	Post	0.0694	0.1518	-0.1462**	2.0021***	266.96***	2.7929***
Ecuador	Pre	0.0937	0.1752	-0.2723***	97.1816***	615864.32***	0.6363
	Post	0.0432	0.0915	-2.5093***	68.2053***	304988.73	3.6648***
Mexico	Pre	0.2086	0.2206	0.1328**	6.7452***	2971.44***	0.9147
	Post	0.1057	0.173	-0.1022*	4.2578***	1184.89***	1.6252***
Peru	Pre	0.2612	0.2592	-0.5559***	11.0565***	8052.12***	1.0408
	Post	0.1194	0.2194	-0.4495***	9.0934***	5444.78***	1.3962***
Brazil	Pre	-0.1439	0.3985	0.3920***	13.9596***	12747.16***	-0.8233
	Post	0.0872	0.5749	1.7025***	42.1606***	116664.78***	0.4806***

Note: All means and standard deviations are in annualized values, estimated by multiplying the daily values by 252 and $\sqrt{252}$, respectively. JB statistics (Jarque-Bera) are tests for the normality of the distribution. The upper (lower) values in the last column represent the t test (F test). The symbols *, **, and *** denote significance at the 10%, 5%, and 1% levels, respectively.

4.2 Hurst Exponent

In this study, we apply the Hurst exponent, H , to test the efficient market hypothesis (Fama, 1970). If $H = 0.5$, the process is completely random; that is, a random process with no (positive or negative) long-range memory. The efficient market hypothesis (EMH), thus, assumes $H = 0.5$. Values ranging from 0.5 to 1 indicate a persistent and trend-reinforcing series with positive long-range dependence. On the other hand, positive values smaller than 0.5 suggest anti-persistence, implying that past trends tend to reverse in the future with negative long-range dependence. We exhibit the estimates of the Hurst exponent in Table 2. According to Weron (2002), if the series are random, the value of the Hurst exponent will fall into the intervals (0.3412, 0.6534), (0.3763, 0.6167), and (0.3954, 0.5976) with confidence levels of 99%, 95%, and 90%, respectively. From Table 2, we draw the following conclusion: 1) all of the Hurst exponents are bigger than 0.5 in both the pre- and post-GFC periods for all the stock markets in Latin America, implying that stock returns in the markets do not possess any negative long-range dependence and past trends do not tend to reverse in the future. 2) All of the Hurst exponents are not rejected to be 0.5 in the post-GFC period for all the stock markets in Latin America, while the Hurst exponents of both Peru are rejected to be 0.5 in the pre-GFC period, implying that all stock indices in the post-crisis period follow random walk models in Latin America but Peru market in the pre-crisis period do not follow any random walk model but possess positive long-range dependence and are persistent and trend-reinforcing. 3) All of the Hurst exponents are smaller in the post-GFC period than in the pre-GFC period except Chile and Ecuador. Thus, our results from the Hurst exponent show that, in general, stock markets in Latin America improved their efficiency and randomness after the global financial crisis.

TABLE 2
HURST EXPONENT FOR LATIN AMERICAN STOCK MARKETS

Country	Pre	Post
Argentina	0.5631	0.5523
Chile	0.5721	0.592
Colombia	0.5958	0.5944
Ecuador	0.5162	0.5438
Mexico	0.5743	0.5504
Peru	0.6343**	0.5901
Brazil	0.5596	0.542

Note: *, **, and *** indicate significance at the 90%, 95%, and 99% level, respectively.

4.3. Runs Test

We now apply the runs test to test for the randomness of the returns and exhibit the results in Table 3. From the table, we observe the following: 1) we do not reject the null hypothesis of randomness for Argentina and Ecuador in both the pre- and post-GFC periods. 2) For Mexico and Brazil, we reject the null hypothesis of randomness in the pre-GFC period but not in the post-GFC periods. 3) For Chile, Colombia, and Peru, we reject the null hypothesis of randomness in both the pre- and post-GFC periods. Nonetheless, the absolute values of the Z-statistics are smaller in the post-GFC period than in the pre-GFC period. Overall, using the runs test we conclude that the GFC have a positive impact on Latin American markets in the sense that their stock returns are more independent in the post-GFC period than in the pre-GFC period, implying that Latin American markets become more efficient and more random after the global financial crisis.

TABLE 3
RUNS TEST FOR LATIN AMERICAN STOCK MARKETS

Country		Cases< mean	Cases> mean	Total Cases	Number of Runs	Z statistic	p-value
Argentina	Pre	774	791	1565	780	-0.1724	0.8632
	Post	827	738	1565	754	-1.3683	0.1712
Chile	Pre	760	805	1565	630	-7.7365***	<.0001
	Post	808	757	1565	686	-4.8940***	<.0001
Colombia	Pre	786	779	1565	616	-8.4702***	<.0001
	Post	811	754	1565	728	-2.7579***	0.0058
Ecuador	Pre	1238	327	1565	520	0.1263	0.8995
	Post	1211	354	1565	556	0.5165	0.6055
Mexico	Pre	756	809	1565	728	-2.7645***	0.0057
	Post	803	762	1565	779	-0.2006	0.841
Peru	Pre	797	768	1565	687	-4.868***	<.0001
	Post	821	744	1565	697	-4.2891***	<.0001
Brazil	Pre	753	812	1565	749	-1.6909*	0.0909
	Post	825	740	1565	755	-1.3285	0.184

Note: *, ** and *** denote significance at the 90%, 95% and 99% level, respectively.

4.4. Multiple Variance Ratio Tests

To further investigate whether Latin American stock markets became more efficient after the global financial crisis, we employ the multiple variance ratio tests to test for the randomness of the returns and display in Table 4 the results

of $Z_1^*(q)$ and $Z_2^*(q)$ by using Equation (4). Under the multiple variance ratio procedure, only the maximum absolute values of the test are examined. The critical values of $Z_1^*(q)$ and $Z_2^*(q)$ at the 10%, 5%, and 1% levels of significance are 2.23, 2.49, and 3.02, respectively. For each set of the multiple variance ratio tests, an asterisk denotes the maximum absolute value of the test statistic that exceeds the critical value, and thereby, indicates that the null hypothesis of a random walk is rejected. In principle, the rejection of the homoscedastic hypothesis could imply that the returns are either from heteroscedasticity or autocorrelation. In the pre-GFC period, all of the stock markets reject the null hypothesis of a homoscedastic random walk. This suggests that all of the markets studied in our paper are inefficient. However, the null hypothesis of a heteroskedastic random walk is not rejected for Argentina, Ecuador and Brazil, implying that rejection of the null hypothesis of a homoskedastic random walk could be due to the heteroskedasticity in the returns.

TABLE 4
MULTIPLE VARIANCE RATIO TESTS FOR LATIN AMERICAN STOCK MARKETS

Country	Pre		Post	
	$Z_1^*(q)$	$Z_2^*(q)$	$Z_1^*(q)$	$Z_2^*(q)$
Argentina	2.4764*	1.4266	1.2556	0.9177
Chile	7.1340***	3.1302***	7.7094***	4.2352***
Colombia	6.5716***	2.8235**	4.1555***	3.1730***
Ecuador	10.1869***	1.7991	1.538	1.4075
Mexico	3.6010***	2.2455*	2.4749*	2.0786
Peru	8.2572***	2.9821**	4.9843***	2.9506**
Brazil	2.5106**	1.2171	3.1046***	1.2527

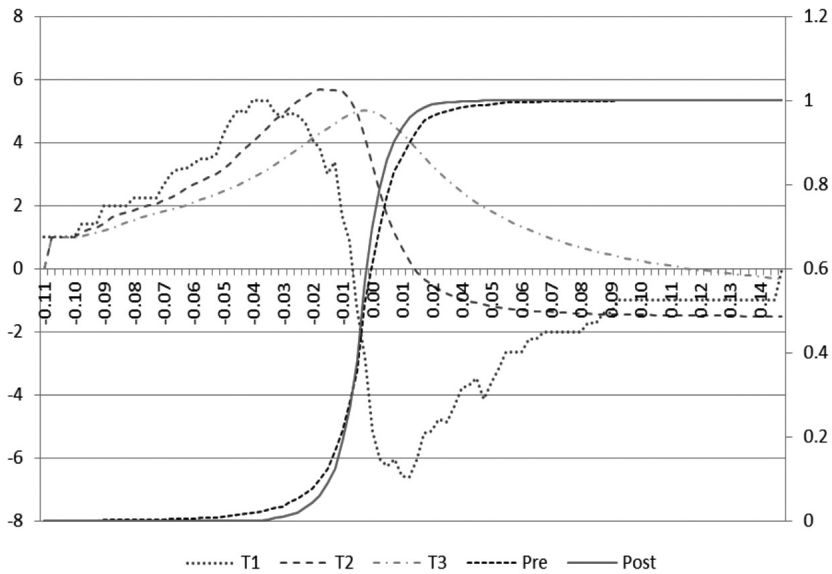
Note: $Z_1^*(q)$ is the test statistic for the null hypothesis of a homoscedastic incremental random walk and $Z_2^*(q)$ is the test statistic for the null hypothesis of a heteroskedastic incremental random walk with the lag-vector (2, 4, 8, 16). *, **, and *** denote significance at the 10%, 5%, and 1% levels, respectively.

We now examine the results in the post-GFC period. Comparing them with the results in the pre-GFC period, from Table 4 we notice the following differences: 1) $Z_1^*(q)$ of Argentina and Ecuador become insignificant in the post-crisis period, 2) $Z_1^*(q)$ of Mexico shows significance at the 10% level in the post-crisis period, a change from 1% significance in the pre-crisis period, and 3) $Z_2^*(q)$ of Mexico becomes insignificant in the post-crisis period. Our results from the MVR test confirm that the stock markets studied in this paper improve their randomness after the recent financial crisis.

4.5. Stochastic dominance analysis

Last, we employ the SD test to examine the preferences of investors in the global financial crisis, examine market efficiency, and check whether there is any arbitrage opportunity due to the GFC. To illustrate the SD relationship, we first plot the cumulative distribution functions (CDFs) of the returns, F and G , for the pre- and post-GFC periods, respectively, and the corresponding SD statistics for Colombia in Figures 2 as an example. For simplicity, we skip reporting the figures of other stock markets. We first apply the SD test to study the investors' preference in the pre- and post-GFC periods for Colombia. Figure 2 shows that there is no first-order SD between the pre- and post-GFC periods for Colombia. From Figure 2, we find that G lies below F in the negative returns, while F is below G in the positive returns, implying that the return in the post-GFC period is preferred in the negative returns, while the return in the pre-GFC period is preferred in the positive returns. In addition, we can see clearly that the first-order SD statistic (T_1) is positive when the returns are negative, and it becomes negative when the returns are positive.

FIGURE 2
CDFs OF RETURNS AND DD STATISTICS



Note: Pre and Post are the CDFs of pre-GFC and post-GFC for Colombia.

TABLE 5
STOCHASTIC DOMINANCE TEST FOR LATIN AMERICAN STOCK MARKETS

Country		FSD $j = 1$	SSD $j = 2$	TSD $j = 3$	Summary
Argentina	$T_j > M_{j, \alpha} \text{ (%)}$	0	0	0	=
	$T_j < -M_{j, \alpha} \text{ (%)}$	0	0	0	
Chile	$T_j > M_{j, \alpha} \text{ (%)}$	0	0	0	=
	$T_j < -M_{j, \alpha} \text{ (%)}$	0	0	0	
Colombia	$T_j > M_{j, \alpha} \text{ (%)}$	22	26	33	< _{2,3}
	$T_j < -M_{j, \alpha} \text{ (%)}$	18	0	0	
Ecuador	$T_j > M_{j, \alpha} \text{ (%)}$	2	2	1	=
	$T_j < -M_{j, \alpha} \text{ (%)}$	7	0	0	
Mexico	$T_j > M_{j, \alpha} \text{ (%)}$	6	24	22	< _{2,3}
	$T_j < -M_{j, \alpha} \text{ (%)}$	7	0	0	
Peru	$T_j > M_{j, \alpha} \text{ (%)}$	1	3	2	=
	$T_j < -M_{j, \alpha} \text{ (%)}$	0	0	0	
Brazil	$T_j > M_{j, \alpha} \text{ (%)}$	0	0	0	=
	$T_j < -M_{j, \alpha} \text{ (%)}$	0	0	0	

Notes: This table reports the stochastic dominance results to test whether the return in the pre-GFC period strictly dominates that in the post-GFC period in the sense of the j - order stochastic dominance for $j = 1, 2, 3$. For example, if we report pre $\succ_{2,3}$ post for Brazil, this means that the return in the pre-GFC period stochastically dominates that in the post-GFC period in the sense of second and third orders. When we report pre = post, this means there is no dominance of the returns between the pre- and post-GFC periods. The numbers in the columns of FSD, SSD, and TSD indicate the percentages of the first-, second-, and third-order modified DD statistics significantly in the positive or negative domain at the 5% level. T_j is defined in Equation (7) in the appendix for $j = 1, 2$ and 3 with F and G denoting the return series for the pre- and post-GFC periods, respectively.

Figure 2 gives us some ideas of the SD relationship of the returns before and after the financial crisis. To test the SD formally, we report in Table 5 the percentages of the first three orders of significantly modified SD statistics for the return distributions in the pre- and post-crisis period. For example, the second- and third-order SD statistics (T_2 and T_3) are significant positive, regardless of whether the returns are positive and negative for Colombia. From Figure 2 and Table 5, we draw the following conclusions: 1) the returns in the pre- and post-GFC periods do not stochastically dominate each other at the first three orders for Argentina, Chile, Ecuador, Peru and Brazil; 2) there is no first-order SD between the returns in the pre- and post-GFC periods for all Latin America stock markets; but 3) the post-GFC period is preferred in the negative returns, while the pre-GFC period is preferred in the positive returns, and 4) returns in the post-GFC period stochastically dominate those in the pre-GFC period at the second and third order SD for Colombia and Mexico.

To summarize, our SD results imply that the markets we studied in this paper are efficient, there is no arbitrage opportunity for the markets due to the GFC in our studying period, and second and third order investors will prefer investing in the post-GFC period to pre-GFC period for Colombia and Mexico⁴.

5. CONCLUSION AND INFERENCE

The previous literature on financial crises focuses on the negative effects of crises on financial markets. This paper investigates whether the most recent global financial crisis has a positive impact in terms of efficiency on Latin American stock indices (Argentina, Chile, Colombia, Ecuador, Mexico, Peru, and Brazil). To conduct the analysis, we use the MV analysis, runs test, multiple variance ratio test, and SD tests.

Using the MV approach, we find that there is no dominance between the pre- and post-GFC periods for Argentina, investors prefer to invest in the pre-GFC period for Brazil, and for the remaining stock markets, investors prefer to invest in the post-GFC period. Thus, in general, our MV results show that the global financial crisis has a positive impact on Latin American stock markets in the sense that, in general, the stock markets in Latin America are significantly less volatile and investors prefer to invest in the markets in the post-GFC period.

The runs test shows that the GFC had a positive impact on Latin American markets in the sense that their stock returns are more independent or closer to independent in the post-GFC period than in the pre-GFC period, implying that Latin American markets becomes more efficient and more random after the global financial crisis. The multiple variance ratio test confirms that the stock markets studied in this paper improve their efficiency and randomness after the recent financial crisis.

Our stochastic dominance result implies that the markets we studied in this paper are efficient, there is no arbitrage opportunity in all of the markets studied in our paper due to the GFC in our studying period, and second and third order risk-aversers will prefer investing in the post-GFC period to pre-GFC period for Colombia and Mexico.

We note that our stochastic dominance result implies that there is no arbitrage opportunity in the markets due to the GFC in our studying period does not mean that we are sure that there is no arbitrage opportunity in the markets. Arbitrage opportunities could arise again until the next crisis shows up, and the try to make a profit arbitraging away the abnormal returns that are going to show up at some point during the crisis (Hinich and Patterson, 1985).

⁴ Readers may refer to Wong, *et al.* (2008), Clark, *et al.* (2015), Guo, *et al.* (2017), and the references therein for more discussions on testing market efficiency and test whether there is any arbitrage opportunity in the markets.

Based on all of the results, we conclude that, in general, Latin American stock markets perform better and becomes more efficient and mature after the most recent global financial crisis and, in general, investors prefer investing in the post-GFC period. The results confirm that the 2008 global financial crisis does have some positive impacts on Latin American stock markets. Our findings provide important information for investors and market regulators for their decision making in their investment and setting regulations.

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APPENDIX

I. Stochastic Dominance Test

We assume $\{f_i\} (i = 1, 2, \dots, n_f)$ and $\{g_i\} (i = 1, 2, \dots, n_g)$ are observations drawn from the returns X and Y , with distribution functions F and G , respectively, and with their integrals $F_j(x)$ and $G_j(x)$ defined in (3) for $j = 1, 2, 3$. For a grid of pre-selected points x_1, x_2, \dots, x_k , the j -order SD test statistic, T_j , proposed by Davidson and Dulcos (2000) and modified by Bai, *et al.* (2015) is:

$$(7) \quad T_j(x) = \frac{\widehat{F}_j(x) - \widehat{G}_j(x)}{\sqrt{\widehat{V}_j(x)}}$$

where

$$\widehat{V}_j(x) = \widehat{V}_{F_j}(x) + \widehat{V}_{G_j}(x) - 2\widehat{V}_{FG_j}(x);$$

$$\widehat{H}_j(x) = \frac{1}{N_h(j-1)!} \sum_{i=1}^{N_h} (x - h_i)_+^{j-1},$$

$$\widehat{V}_{H_j}(x) = \frac{1}{n_h} \left[\frac{1}{n_h((j-1)!)^2} \sum_{i=1}^{N_h} (x - h_i)_+^{2(j-1)} - \widehat{H}_j(x)^2 \right], H = F, G; h = f, g;$$

$$\widehat{V}_{FG_j}(x) = \frac{1}{n_h} \left[\frac{1}{n_h((j-1)!)^2} \sum_{i=1}^{N_h} (x - f_i)_+^{j-1} (x - g_i)_+^{j-1} - \widehat{F}_j(x) \widehat{G}_j(x) \right].$$

For all $i = 1, 2, \dots, k$; we test the following hypotheses:

$$(8) \quad \begin{aligned} H_0 &: F_j(x_i) = G_j(x_i), \text{ for all } x_i; \\ H_A &: F_j(x_i) \neq G_j(x_i), \text{ for some } x_i; \\ H_{A1} &: F_j(x_i) \leq G_j(x_i), \text{ for all } x_i, F_j(x_i) < G_j(x_i) \text{ for some } x_i; \\ H_{A2} &: F_j(x_i) \geq G_j(x_i), \text{ for all } x_i, F_j(x_i) > G_j(x_i) \text{ for some } x_i. \end{aligned}$$

We note that in the above hypotheses, H_A is set to be exclusive of both H_{A1} and H_{A2} . This means that if the test does not reject H_{A1} or H_{A2} , it will not be classified as H_A . Therefore, Bai *et al.* (2015) modify the decision rules to be:

$$\begin{aligned}
 (9) \quad & \max_{1 \leq k \leq K} |T_j(x_k)| < M_{\alpha}^j, \text{ accept } H_0 : X =_j Y \\
 & \max_{1 \leq k \leq K} T_j(x_k) > M_{\alpha}^j \text{ and } \min_{1 \leq k \leq K} T_j(x_k) < -M_{\alpha}^j, \text{ accept } H_A : X \neq_j Y \\
 & \max_{1 \leq k \leq K} T_j(x_k) < M_{\alpha}^j \text{ and } \min_{1 \leq k \leq K} T_j(x_k) < -M_{\alpha}^j, \text{ accept } H_{A1} : X \succeq_j Y \\
 & \max_{1 \leq k \leq K} T_j(x_k) > M_{\alpha}^j \text{ and } \min_{1 \leq k \leq K} T_j(x_k) > -M_{\alpha}^j, \text{ accept } H_{A2} : Y \succeq_j X
 \end{aligned}$$

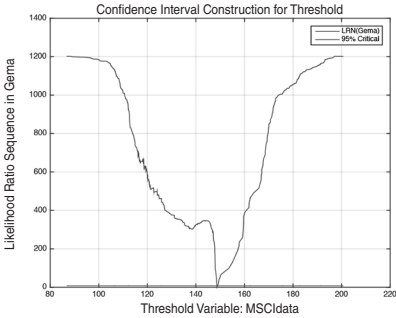
where M_{α}^j is the bootstrapped critical value of the j -order SD statistic. The test statistic is compared with M_{α}^j at each point of the combined sample⁵. We follow Fong *et al.* (2005), Gasbarro *et al.* (2007), and others and choose $K = 100$. We also follow Bai *et al.* (2015) to use $\max_x |T_j(x)|$ in the comparison. In order to minimize Type II errors and to accommodate the effect of almost SD⁶, we follow Gasbarro *et al.* (2007) and others and use a conservative 5% cut-off point in checking the proportion of test statistics for statistical inference. Using a 5% cut-off point implies that one prospect dominates another only if at least 5% of the statistics are significant.

⁵ Readers may refer to Bai *et al.* (2015) for the construction of the bootstrapped critical value M_{α}^j .

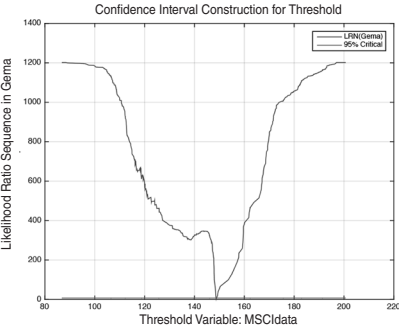
⁶ Readers may refer to Leshno and Levy (2002) and Guo, *et al.* (2013, 2014, 2016) and the references therein for more information. Leshno and Levy (2002) use an example of a 1% violation to state the problem of almost SD.

II. Hansen (2000) “Sampled splitting” test to access if there is a break in the period of 1 January 2008 to 31 December 2009,

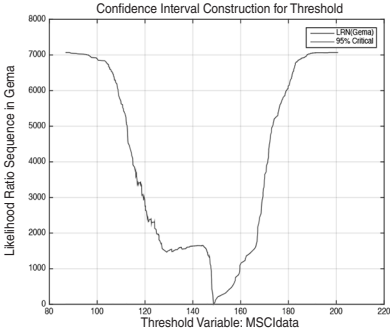
Argentina



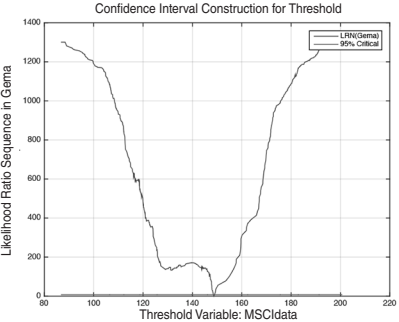
Chile



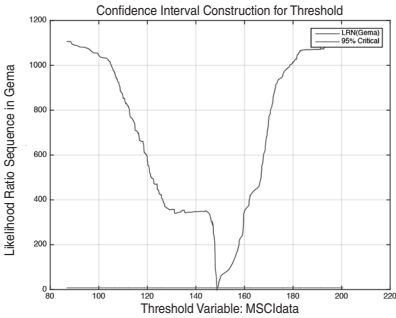
Colombia



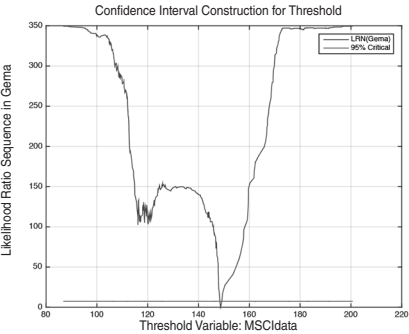
Ecuador



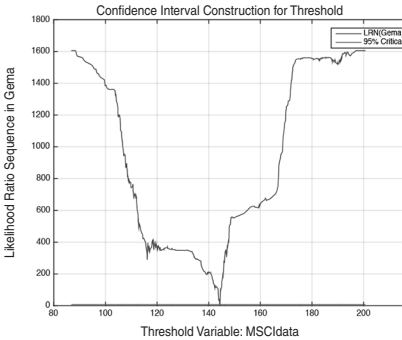
Mexico



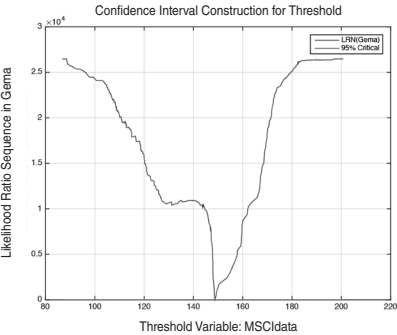
Peru



Brazil



Jamaica



III. HAC Robust Inference

Country	Breakpoint at observation number	Date
Argentina	1463	8/8/2008
Chile	1835	1/12/2010
Colombia	1315	1/15/2008
Ecuador	1630	3/31/2009
Mexico	1789	11/9/2009
Peru	1464	8/11/2008
Brazil	1407	5/21/2008