# Stock Market Liberalizations and Efficiency: The Case of Latin America

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**Abstract**

This investigation is among the first to examine the impact of stock market liberalization on the efficiency of Latin American stock markets. It is also among the first to apply the martingale hypothesis test and a stochastic dominance approach to study the issue of efficient markets. Daily stock indices from Latin American countries, including Brazil, Mexico, Chile, Peru, Jamaica, and Trinidad and Tobago, are used in our analysis. To examine the impact of stock market liberalization on efficiency, we employ several approaches, including the runs test, Chow-Denning multiple variation ratio test, Wright variance ratio test, the martingale hypothesis test and the SD test, the stock market indices of the countries above. We find that stock market liberalization does not significantly improve stock market efficiency in Latin America.

*Key words:* Market Liberalization; Market Efficiency; Stochastic Dominance, Latin America.

JEL Classifications: G14, G15

## 1. Introduction

The move toward financial liberalization in Latin America has not been a continuous process but one with a bump during the 1980s as a result of financial crisis. Financial liberalization in Latin America has not been the result of isolated policy initiatives but has always been implemented as the components of wide-ranging structural reforms and stabilization programs. These kinds of programs were set up in the region in the second half of the 1970s, and their implementation was mostly completed  in the 1990s. During that period, Latin American countries started the process of stock market liberalization.  These countries undertook massive reforms and consequently reduced the stock market liberalization gap between them and industrial countries.

The literature notes that stock market liberalization could lead to an increase in equity prices in emerging markets (Henry 2000b; Bekaert and Harvey 2000),  liquidity (Han Kim and Singal 2000), and investment (Henry 2000a), economic growth (Bekaert and Harvey 1997; Bekaert et al. 2001), the repricing of risk (Chari & Henry 2001) and a decrease in the cost of capital (Stulz 1999), a decrease in dividends (Bekaert and Harvey, 2000), and a reduced equity premium (Ahimud and Mendelson 1986; Ahimud et al., 1997).

According to the efficient markets hypothesis (EMH), the increase the availability of information in efficient markets should lead to stock prices that are more efficiently priced. In other words, theory suggests that as information becomes more readily available as a result of liberalization and there is increased competition, predictability should decline. Examples of research related to this topic include Groenewold and Ariff (1998), Kawakatsu and Morey (1999), Han Kim and Singal (2000a, 2000b) and Laopodis (2004) and Füss (2005).

The literature about the impact of liberalization on market efficiency is mixed.  For example, Groenewold and Ariff (1998) find that the predictability of Asia-Pacific stock prices remains the same after stock market liberalization.  Füss (2005) finds that liberalization could improve market efficiency in Asian stock markets, and Han Kim and Singal (2000a, 2000b) analyzed 14 emerging countries' stock markets and found find that  these markets became more efficient after market liberalization. But authors like Kawakatsu and Morey (1999), testing the efficiency of nine emerging market countries, did not find evidence that these markets become more efficient after liberalization. Laopodis (2004) finds that Greece's financial market didn't become efficient because the market was already efficient before liberalization. Hence, the impact of liberalization on stock market efficiency remains puzzle in the existing literature, especially for Latin American stock markets.

This investigation is among the first to examine the impact of stock market liberalization on the efficiency of stock markets and also among the first to examine the impact of stock market liberalization on the efficiency of Latin American stock markets. To examine the impact on efficiency and to be rigorous in our investigation, we employ several approaches, including the runs test, Chow-Denning multiple variation ratio test, Wright variance ratio test, the martingale hypothesis test, and the SD test, to studying Latin American stock market indices. Therefore, this study could be an important reference for those countries that want to improve their stock markets' efficiency through liberalization, especially those countries that have conditions similar to those in Latin American countries. This paper is also among the first to apply the martingale hypothesis test and a stochastic dominance approach to the issue of efficient markets.

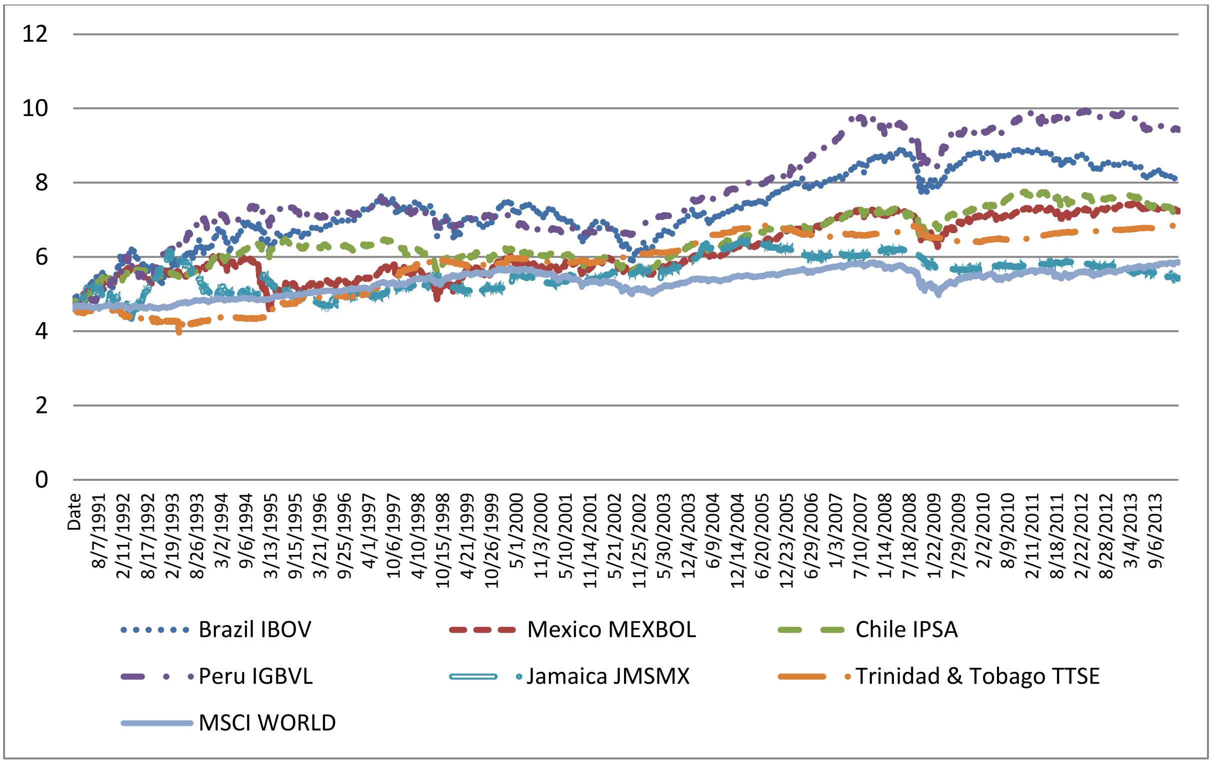
The paper is organized as follows. Section 2 describes data sources and outlines the empirical methodology used, and section 3 discusses the results. Section 4 concludes.

## 2. Data and Methodology

### 2.1 Data

Our data consist of daily closing values of stock market indices for the Latin American countries (Brazil, Mexico, Chile, Peru, Jamaica and Trinidad and Tobago), including the IBOV, MEXBOL, IPSA, IGBVL, JMSMX and TTSE. These countries were chosen because they are important economies in the region. The other countries were excluded from the Latin American list because their stock market data are available only after the date on which their stock markets were liberalized. In addition, we use the MSCI World Index to represent the regional market index. The MSCI World Index captures large- and mid-cap firms across 23 developed market (DM) countries. With 1,611 constituents, the index covers approximately 85% of the free float-adjusted market capitalization in each country. These data are extracted from Datastream. The stock market indices of the Latin American countries from 1999 to 2013 are plotted in figure 1 for reference.

**Figure 1**: Stock market indices of the Latin American countries 1999 to 2013




Note: For an easier comparison, we set all the stock indices at the same basis of 100 on January 2, 1999 and take the natural log of them.

We compare the stock returns of one year before, one year after, and ten years after liberalization for a particular stock exchange. Table 1 lists the date on which each of the 6 countries first liberalized its stock market. We compare stock returns before and after liberalization, since the returns could be affected by the market return. Thus, we substract the reture of the MSCI World Index from the return of each return series to obtain the excess returns in order to eliminate the impact of the regional economy. In other words, we compare the excess returns between the pre-liberalization period and the post-liberalization period for the six stock indices. In this paper, we adopt several tests, including the MV criterion, CAPM statistics, the runs test, the multiple variation ratio test, the martingale hypothesis test and the SD test, to investigate whether stock exchange liberalization improved the performance and efficiency of the markets we are analyzing. We illustrate these tests in the following subsections.

Table 1: First stock market liberalization

| Country | Date of first stock market liberalization |
|---|---|
| Brazil | May 1991 |
| Mexico | May 1989 |
| Chile | January 1992 |
| Peru | January 1992 |
| Jamaica | September 1991 |
| Trinidad &Tobago | April 1997 |

## 2.2 Methodology

As Fama (1970) mentions, the definition of efficient market prices—that they "fully reflect" the available information—is very general. It is so general that it has no direct empirically testable implications. To make the model testable, we must define more exactly what is meant by the term "fully reflect." Various tests are developed to indirectly test the EMH. In this paper we will first use the mean-variance criterion and CAPM statistics to evaluate the performance of stock market indices in the Latin American countries. To ensure a robust result, we will then employ a number of these tests, including the runs test, the multiple variation ratio test, the martingale hypothesis test, and a recently developed stochastic dominance test, to examine the EMH in both the pre- and the post-stock market liberalization period.

### 2.2.1 Performance

#### 2.2.1.1. The Mean-Variance Criterion

For any two investments with returns $X$ and $Y$ with means $\mu_X$ and $\mu_Y$ and standard deviations $\sigma_X$ and $\sigma_Y$, respectively, $Y$ is said to dominate $X$ by the MV criterion for risk averters if $\mu_Y \geq \mu_X$ and $\sigma_Y \leq \sigma_X$ with at least one inequality holds (Markowitz, 1952). Thus, the MV rule for risk averters is to check whether $\mu_Y \geq \mu_X$ and $\sigma_Y \leq \sigma_X$. If both are not rejected with at least one strict inequality relationship, then we conclude that $Y$ dominates $X$ 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 $Y$ dominates $X$ by the MV criterion for risk averters, then risk averters will attain higher expected utility by holding $Y$ than $X$. The theory can be extended to non-differentiable utilities (see Wong and Ma (2008), for details).

#### 2.2.1.2. The CAPM Statistics

We next apply the CAPM [1] analysis, including a beta component, the Sharpe ratio, [2] Treynor's index and Jensen's index (alpha), to measure the performance of the stock indices. The beta of the portfolio measures the marginal contribution of a portfolio to the total market portfolio and the sensitivity of its return to the movement of market portfolio returns. The estimation requires fitting the following CAPM equation for the return $R_{i,t}$ of index *i* at time *t*:

$$R_{i,t} - R_{f,t} = \alpha_i + \beta_i \left( R_{m,t} - R_{f,t} \right) + \varepsilon_{i,t} \tag{1}$$

where $\varepsilon_{i,t}$ is the residual assumed to be i.i.d., $R_{m,t}$ is the return of the market portfolio, and $R_{f,t}$ is the return of the risk-free asset at time *t*. In our paper, we use the return of the MSCI World Index to represent the $R_{m,t}$ and the return of the 3-month Treasury bill as the $R_{f,t}$. From equation (1), three performance indices — the Sharpe ratio ($S_i$), Treynor's index ($T_i$), and Jensen's index ($J_i$) — are then computed using the following formula:

$$S_i = \frac{\bar{R}_i - \bar{R}_f}{\hat{\sigma}_i},\ T_i = \frac{\bar{R}_i - \bar{R}_f}{\hat{\beta}_i} \text{ and } J_i = \hat{\alpha}_i = (\bar{R}_i - \bar{R}_f) - \hat{\beta}_i(\bar{R}_m - \bar{R}_f)\,. \tag{2}$$

where $\hat{\sigma}_i$ is the estimated standard deviation, and $\bar{R}_i$, $\bar{R}_m$ and $\bar{R}_f$ are the expected return of index *i*, the market portfolio and the risk-free asset, respectively.

### 2.2.2. Degree of Efficiency

#### 2.2.2.1 Runs Test

The runs test (Bradley, 1968) is a nonparametric test to determine whether successive price changes are independent. According to Campbell (1997), it could be used to examine the number of sequences of consecutive positive and negative returns tabulated and compared against its sampling distribution under the random walk hypothesis.

If $y_1, \ldots y_N$ is a time series of *N* returns and $y_m$ is their median, the series of signs of residuals, sign $u_1$,..., sign $u_N$ are considered where $u_i = y_i - y_m$ and $i = 1, \ldots, N$. That is, a

[1] Ostermark (1991) uses the capital asset pricing model to analyze two Scandinavian stock markets and finds that the standard CAPM is unable to exhaustively represent the economic forces of capital asset pricing, especially in Sweden.

[2] Ferruz Agudo and Sarto Marzal (2004) apply the Sharpe ratio to analyze the performance of Spanish investment funds.

positive change *"+"* is assigned to each return $y_i$ that is greater than the median, a negative change *"–"* is assigned when the return is less than the median, and the return is omitted when it equals the median. A run is the number of sequences of same signs. For example, the series of signs $+ + + - - + - -$ gives 4 runs.

To perform this test, we let $n_+$ and $n_-$ be the number of runs of *"+"* and *"–"*, respectively and let *U* be the observed number of runs. Too many or too few runs in the sequence are the results of negative and positive autocorrelation, respectively. Under the null hypothesis of randomness or independence, by comparing the observed number of runs (*U*) with the expected number of runs ($\mu_U$), the test of the randomness hypothesis can be constructed. It has been shown that, for large sample sizes where both $n_+$ and $n_-$ are greater than twenty, the standardized test statistic is

$$Z = \frac{U - \mu_U}{\sigma_U} \ ,$$

where $\mu_U = \frac{2n_+n_-}{n} + 1$, $\sigma_U = \sqrt{\frac{2n_+n_-\left(2n_+n_- - n\right)}{n^2\left(n-1\right)}}$ and $n = n_+ + n_-$.

We note that *Z* is approximately normally distributed under the null hypothesis of randomness or independence. If $Z < -Z_{1-\alpha/2}$ ($Z > Z_{1-\alpha/2}$), we reject the null hypothesis and conclude that $Y_i$ is not random and not independent, and thus, we can conclude that $Y_i$ is negatively (positively) autocorrelated.

#### 2.2.2.2 Variation Ratio Test

Variance ratio tests have been widely used and are particularly useful for examining the behavior of stock prices or indices in which returns are frequently not normally distributed. Suppose we have the time series $\{X_t\} = \left(X_0, X_1, X_2, ..., X_T\right)$ satisfying

$$\Delta X_t = \mu + \varepsilon_t , \qquad (3)$$

where $X_t$ is the stock index and $\mu$ is an arbitrary drift parameter. The residual $\varepsilon_t$ satisfies

$E(\varepsilon_t)=0$ and $E\left(\varepsilon_t\varepsilon_{t-j}\right)=0$ when $j\neq 0$ for all $t$. Lo and MacKinlay (1988) provide tests of the null hypothesis of randomness. Variance ratio tests focus on the property that under a random walk with uncorrelated increments in $X_t$, the variance of these increments increases linearly in the observation intervals such that $Var(X_t - X_{t-q}) = qVar(X_t - X_{t-1})$ for any positive integer $q$. The variance ratio is then given by

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

Under the null hypothesis that $\{X_t\}$ follows the random walk model stated in (3), we have $\mathrm{VR}(q)=1$. Lo and MacKinlay (1998) generate the asymptotic distribution of the estimated variance ratios and provide two test statistics, $Z_1(q)$ and $Z_2(q)$,[3] both of which have asymptotic standard normal distributions under the null hypothesis. $Z_1(q)$ is derived under the assumption that the disturbances of equation (3) are homoscedastic, while $Z_2(q)$ treats them as heteroscedastic. The latter test statistic is not only sensitive to the changes in stock prices, but it is also robust to many general forms of heteroscedasticity and non-normality.

The random walk hypothesis implies that $\mathrm{VR}(q)=1$ for any integer $q$. To improve on the work of Lo and MacKinlay (1998), Chow and Denning (1993) show how controlling test size facilitates the multiple variance ratio tests. For a single variance ratio test, under the null hypothesis that $\mathrm{M}_r(q)=\mathrm{VR}(q)-1=0$, we follow Chow and Denning (1993) to consider a set of $m$ tests $\{\mathrm{M}_r(q_i)|i=1,2,\cdots,m\}$ associated with the set of aggregation intervals $\{q_i|i=1,2,\cdots,m\}$. Under the null hypothesis of a random walk, there are multiple sub-hypotheses

[3] Readers may refer to Lo and MacKinlay (1999) for the formula.

$$
\begin{aligned}
&H_{0i}\text{: } \mathrm{M}_r\left(q_i\right)=0 \ \text{ for all } i=1,2,\cdots,m; \\
&H_{1i}\text{: } \mathrm{M}_r\left(q_i\right)\neq 0 \ \text{ there exists any } \ i=1,2,\cdots,m.
\end{aligned} \tag{5}
$$

Rejection of at least one $H_{0i}$ for $i=1,2,\cdots,m$ implies rejection of the random walk. For the homoscedastic situation, we use the test statistics $\left\{Z_1\left(q_i\right)\middle| i=1,2,\cdots,m\right\}$, whereas, for the heteroscedastic situation, we adopt the test statistics $\left\{Z_2\left(q_i\right)\middle| i=1,2,\cdots,m\right\}$. Since the random walk hypothesis is rejected if any of the $\hat{\mathrm{VR}}\left(q_i\right)$ is significantly different from one, we only consider the $Z_1^*\left(q\right)$ and $Z_2^*\left(q\right)$, where

$$
Z_i^*\left(q\right)=\max\left(\left|Z_i\left(q_1\right)\right|,\cdots,\left|Z_i\left(q_m\right)\right|\right), \quad i=1,2\,. \tag{6}
$$

If $Z_1^*\left(q\right)$ ( $Z_2^*\left(q\right)$ ) is greater than the SMM (*α*, *m*, *N*), then the random walk hypothesis is rejected under the homoscedastic (heteroscedastic) assumption, where SMM is the upper *α* point of the studentized maximum modulus distribution (Richmond, 1982) with parameter *m* and *N* (sample size) degrees of freedom. In addition, Wright (2000) indicates two potential advantages of rank- and sign-based tests.

*2.2.2.2.1 Rank-Based Variance Ratio Tests*

Suppose that $Y_t$ is a time series of the asset returns with sample size *T*. Let $r\left(Y_t\right)$ be the rank of $Y_t$ among $Y_1,Y_2,\ldots,Y_T$. Define

$$
r_{1t}=\frac{\left(r\left(Y_t\right)-\dfrac{T+1}{2}\right)}{\sqrt{\dfrac{\left(T-1\right)\left(T+2\right)}{12}}} \qquad \text{and} \qquad r_{2t}=\Phi^{-1}\left(\frac{r\left(Y_t\right)}{T+1}\right), \tag{7}
$$

where $\Phi^{-1}$ is the inverse of the standard normal cumulative distribution function.

The series $r_{1t}$ is a simple linear transformation of the ranks and is standardized to have zero mean and unit variance. The series $r_{2t}$, known as the inverse normal or van der Waerden scores, also has zero mean and unit variance. Wright (2000) substitutes $r_{1t}$ and $r_{2t}$ in place of the return $(X_t - X_{t-q})$ in Lo-MacKinlay's definition of the variance ratio test statistic (assuming homoscedasticity). The rank-based variance ratio test statistics $R_1$ and $R_2$ are defined as

$$R_1 = \left( \frac{\frac{1}{Tk}\sum_{t=k}^{T}\left(r_{1t} + r_{1t-1} + \cdots + r_{1t-k+1}\right)^2}{\frac{1}{T}\sum_{t=1}^{T} r_{1t}^2} - 1 \right) \times \left( \frac{2(2k-1)(k-1)}{3kT} \right)^{-1/2}, \tag{8}$$

$$R_2 = \left( \frac{\frac{1}{Tk}\sum_{t=k}^{T}\left(r_{2t} + r_{2t-1} + \cdots + r_{2t-k+1}\right)^2}{\frac{1}{T}\sum_{t=1}^{T} r_{2t}^2} - 1 \right) \times \left( \frac{2(2k-1)(k-1)}{3kT} \right)^{-1/2}. \tag{9}$$

*2.2.2.2.2 Sign-Based Variance Ratio Tests*

For any series of return $Y_t$, let $u(Y_t, q) = 1(Y_t > q) - 0.5$. So, $u(Y_t, 0)$ is 1/2 if $Y_t$ is positive and −1/2 otherwise. Let $s_t = 2u(Y_t, 0)$. Clearly, $s_t$ is an independent and identically distributed (i.i.d.) series with zero mean and unit variance. Each $s_t$ is equal to 1 with probability 1/2 and is equal to −1 otherwise. The sign-based variance ratio test statistic $S_1$ is defined as

$$S_1 = \left( \frac{\frac{1}{Tk}\sum_{t=k}^{T}\left(s_t + s_{t-1} + \cdots + s_{t-k+1}\right)^2}{\frac{1}{T}\sum_{t=1}^{T} s_t^2} - 1 \right) \times \left( \frac{2(2k-1)(k-1)}{3kT} \right)^{-1/2}. \tag{10}$$

The critical values of $R_1$, $R_2$, and $S_1$ can be obtained by simulating their exact distributions.

### 2.2.2.3. Martingale Hypothesis Test

The distinction between a martingale difference sequence (MDS) and uncorrelatedness is crucial when nonlinear dependence is present, as commonly happens with financial data. For processes with bounded second moments, an MDS is an uncorrelated sequence, but an uncorrelated sequence is not necessarily an MDS. So we should point out that a random walk is a strong form of the efficient market hypothesis and further one may be interested in testing the martingale hypothesis that $X_t$ is a martingale with respect to some filtration $(\mathrm{F}_n)$; or equally to test whether the return sequence $r_t = X_t - X_{t-1}$ forms a martingale difference ($E(r_t \mid \mathrm{F}_{t-1}) = \mu$). To do so, Dominguez and Lobato (2003) derive consistent tests for the null hypothesis that the time series process $r_t$ has constant conditional expectation $\mu$ given the information set $\mathrm{F}_{t-1}$ composed of the current value of some exogenous variables and by a finite number of past values of both the own process and some exogenous variables. Since the asymptotic distribution of the considered test statistic depends on the specific data-generating process, standard asymptotic inference procedures are not feasible. They show that a modified wild bootstrap procedure properly estimates the asymptotic null distribution of the test statistic. That is, they use the wild bootstrap Cramer-von Mises test statistic (denoted as Cp) and wild bootstrap Kolmogorov-Smirnov test statistic (denoted as $\mathrm{K_p}$) (Dominguez and Lobato, 2003) to test whether the return $\{r_t\}$ is a martingale difference sequence. Since there are no other exogenous variables, information set is provided by $\tilde{z}_{t,p} = \{r_{t-1}, \cdots, r_{t-p}\}$. Thus, the considered null hypothesis is: $H_0 : \mathrm{E}\left(r_t \middle| \tilde{z}_{t,p}\right) = \mu \;\; a.s.$. To test $H_0$ here we need some assumptions. Let $\{r_t, \tilde{z}_{t,p}\}$ be an ergodic and strictly stationary process that satisfies $\mathrm{E}|r_t|^{4+\delta} < \infty$ for some $\delta > 0$, and that $r_t$ given $\tilde{z}_{t,p}$ has a bounded conditional density function that is continuous on any conditioning argument.

The proposed test is based on the following equivalence $H_0 \Leftrightarrow R(\tilde{\tau}) = 0$ for almost all $\tilde{\tau} \in R^p$, where $R(\tilde{\tau}) = \mathrm{E}\left[(r_t - \mu) I\left(\tilde{z}_{t,p} \le \tilde{\tau}\right)\right] = \int (s-\mu) I(\tilde{u} \le \tilde{\tau}) dF(s,\tilde{u})$ in which $F(s,\tilde{u})$ is the joint distribution function of the vector $\{r_t, \tilde{z}_{t,p}\}$. Let $F_n$ denote the empirical distribution function of $\{r_t, \tilde{z}_{t,p}\}$ and $\bar{r}$ the sample mean $\bar{r} = n^{-1}\sum_{1}^{n} r_t$.

The estimate of the function $R(\tilde{\tau})$ given by its sample analog $R_n(\tilde{\tau}) = \int (s-\bar{r}) I(\tilde{u} \le \tilde{\tau}) dF_n(s,\tilde{u}) = \frac{1}{n}\sum_{t=1}^{n}(r_t - \bar{r}) I(\tilde{z}_{t,p} \le \tilde{\tau})$ and the two particular test statistics considered here the Cramer–von Mises test statistic is

$$C_{p,n} = \int \left[\sqrt{n} R_n(\tilde{\tau})\right]^2 dF_n(\infty,\tilde{\tau}) = \frac{1}{n^2}\sum_{j=1}^{n}\left[\sum_{t=1}^{n}(r_t - \bar{r}) I(\tilde{z}_{t,p} \le \tilde{z}_{j,p})\right]^2 \tag{11}$$

where $F_n(\infty,\tilde{\tau}) = \lim_{s\to\infty} F_n(s,\tilde{\tau})$, and the Kolmogorov–Smirnov statistic is

$$K_{p,n} = \max_{i=1,\cdots,n}\left|\sqrt{n} R_n(\tilde{z}_{i,p})\right| = \max_{i=1,\cdots,n}\left|\frac{1}{\sqrt{n}}\sum_{j=1}^{n}(r_j - \bar{r}) I(\tilde{z}_{j,p} \le \tilde{z}_{i,p})\right|. \tag{12}$$

Dominguez and Lobato (2003) provide the 10%, 5% and 1% critical values through a bootstrap procedure. In practice, we just need to calculate the value of $C_{p,n}$ or/and $K_{p,n}$, and compare them to the corresponding critical value. If the value of the test statistic is larger than the corresponding critical value, we will reject the null hypothesis. We denote the two martingale difference tests $C_p$ and $K_p$ as MTD and display the tests results in Table 5. Readers may refer to Dominguez and Lobato (2003)) for more details and discussions on testing the martingale hypothesis and the conjecture of the 'martingale property' that generalizes the 'random walk' conjecture in the concept of efficient market.

### 2.2.3 Stochastic Dominance Test

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

$$H_0 = h,\ H_j^A(x) = \int_a^x H_{j-1}^A(t)dt \text{ and } H_j^D(x) = \int_x^b H_{j-1}^D(t)dt \tag{13}$$

for $h = f, g$; $H = F, G$; and $j = 1, 2, 3$.

We call the integral $H_j^A$ the *j*-order ascending cumulative distribution function (ACDF), and the integral $H_j^D$ the *j*-order descending cumulative distribution function (DCDF), for *j* = 1, 2 and 3 and for $H = F$ and $G$.

We define the SD rules as follows (see Quirk and Saposnik, 1962; Fishburn, 1964; Hanoch and Levy, 1969):

*$X$ dominates $Y$ by FASD (SASD, TASD), denoted by $X \succ_1 Y$ ( $X \succ_2 Y$, $X \succ_3 Y$ ) if and only if $F_1^A(x) \le G_1^A(x)$ ( $F_2^A(x) \le G_2^A(x)$, $F_3^A(x) \le G_3^A(x)$ ) for all possible returns $x$, and the strict inequality holds for at least one value of $x$.*

The SD theory for risk seekers has also been well established in the literature. Whereas SD for risk averters works with the ACDF, which counts from the worst return ascending to the best return, SD for risk seekers works with the DCDF, which counts from the best return descending to the worst return (Wong and Li, 1999). Hence, SD for risk seekers is called descending SD (DSD). We have the following definition for DSD (see Hammond, 1974; Wong and Li, 1999):

*$X$ dominates $Y$ by FDSD (SDSD, TDSD)) denoted by $X \succ^1 Y$ ( $X \succ^2 Y$, $X \succ^3 Y$ ) if and only if $F_1^D(x) \ge G_1^D(x)$ ( $F_2^D(x) \ge G_2^D(x)$, $F_3^D(x) \ge G_3^D(x)$ ) for all possible returns $x$, the strict inequality holds for at least one value of $x$, where FDSD (SDSD, TDSD) denotes first-order (second-order, third-order) descending SD.*

We briefly describe the DD test in the following:

Let $\{(f_i, g_i)\}$ $(i = 1, ..., n)$ be pairs of observations drawn from the random variables $X$ and $Y$, with distribution functions *F* and *G*, respectively, and with their integrals $F_j^A(x)$ and $G_j^A(x)$ defined in (13) for $j = 1, 2, 3$. For a grid of pre-selected points $x_1$, $x_2$, ..., $x_k$, Bai et al. (2011) modify the statistic developed by Davidson and Duclos (2000) to obtain the following *j*-order DD test statistic for risk averters, $T_j^A$:

$$T_j^A(x) = \frac{\hat{F}_j^A(x) - \hat{G}_j^A(x)}{\sqrt{\hat{V}_j^A(x)}} \qquad (14)$$

where

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

$$\hat{H}_j^A(x) = \frac{1}{N(j-1)!}\sum_{i=1}^{N}(x - z_i)_+^{j-1},$$

$$\hat{V}_{H_j}^A(x) = \frac{1}{N}\left[\frac{1}{N((j-1)!)^2}\sum_{i=1}^{N}(x-z_i)_+^{2(j-1)} - \hat{H}_j^A(x)^2\right], H = F, G; z = f, g;$$

$$\hat{V}_{FG_j}^A(x) = \frac{1}{N}\left[\frac{1}{N((j-1)!)^2}\sum_{i=1}^{N}(x-f_i)_+^{j-1}\left(x-s_i\right)_+^{j-1} - \hat{F}_j^A(x)\hat{G}_j^A(x)\right].$$

It is not possible to test empirically the null hypothesis for the full support of the distributions. Thus, Bishop et al. (1992) propose to test the null hypothesis for a pre-designed finite number of values *x*. Specifically, for all $i = 1, 2, ..., k;$ the following hypotheses are tested:

$$\begin{aligned} &H_0 : F_j\,(x_i) = G_j(x_i)\ , \text{ for all } x_i; \\ &H_A : F_j\,(x_i) \neq G_j(x_i) \text{ for some } x_i; \\ &H_{A1} : F_j\,\left(x_i\right) \leq G_j\left(x_i\right) \text{ for all } x_i, F_j\,\left(x_i\right) < G_j\left(x_i\right) \text{ for some } x_i; \\ &H_{A2} : F_j\,\left(x_i\right) \geq G_j\left(x_i\right) \text{ for all } x_i, F_j\,\left(x_i\right) > G_j\left(x_i\right) \text{ for some } x_i. \end{aligned} \qquad (15)$$

In order to test SD for risk seekers, the DD statistics for risk averters are modified to be the descending DD test statistic, $T_j^D$, such that:

$$T_j^D(x) = \frac{\hat{F}_j^D(x) - \hat{G}_j^D(x)}{\sqrt{\hat{V}_j^D(x)}} \qquad (16)$$

Where

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

$$\hat{H}_j^D(x) = \frac{1}{N(j-1)!}\sum_{i=1}^{N}(z_i - x)_+^{j-1},$$

$$\hat{V}_{H_j}^D(x) = \frac{1}{N}\left[\frac{1}{N((j-1)!)^2}\sum_{i=1}^{N}(z_i - x)_+^{2(j-1)} - \hat{H}_j^D(x)^2\right], H = F, G; z = f, s;$$

$$\hat{V}_{FG_j}^D(x) = \frac{1}{N}\left[\frac{1}{N((j-1)!)^2}\sum_{i=1}^{N}(f_i - x)_+^{j-1}\left(s_i - x\right)_+^{j-1} - \hat{F}_j^D(x)\hat{G}_j^D(x)\right];$$

where the integrals $F_j^D\left(x\right)$ and $G_j^D\left(x\right)$ are defined in (13) for $j = 1, 2, 3$. For $i = 1, 2, ..., k,$ the following hypotheses are tested for risk seekers:

$H_0 : F_j^D(x_i) = G_j^D(x_i)$ , for all $x_i$;
$H_D : F_j^D(x_i) \neq G_j^D(x_i)$ for some $x_i$ ;
$H_{D1} : F_j^D\left(x_i\right) \geq G_j^D\left(x_i\right)$ for all $x_i, F_j^D\left(x_i\right) > G_j^D\left(x_i\right)$ for some $x_i$;
$H_{D2} : F_j^D\left(x_i\right) \leq G_j^D\left(x_i\right)$ for all $x_i, F_j^D\left(x_i\right) < G_j^D\left(x_i\right)$ for some $x_i$.

Not rejecting $H_0$ or $H_A$ or $H_D$ implies the non-existence of any SD relationship between *X* and *Y*, the non-existence of any arbitrage opportunity between these two markets, and that neither of these markets is preferred to the other. If $H_{A1}$ ($H_{A2}$) of order one is accepted, *X* (*Y*) stochastically dominates *Y* (*X*) at the first order, while if $H_{D1}$ ($H_{D2}$) of order one is accepted, asset *X* (*Y*) stochastically dominates *Y* (*X*) at the first order. In this situation, and under certain regularity conditions,[4] an arbitrage opportunity exists and any non-satiated investors will be better off if they switch from the dominated to the dominant asset. On the other hand, if $H_{A1}$ ($H_{A2}$) [ $H_{D1}$ ( $H_{D2}$ )] is accepted at order two (three), a particular market stochastically dominates the other at the second (third) order. In this situation, an arbitrage opportunity does not exist, and switching from one asset to another will only increase the risk averters' [seekers'] expected utility, though not their expected wealth (Jarrow, 1986; Wong et al. 2008). These results could be used to infer that market efficiency and market rationality could still hold in these markets.[5]

In the above analysis, in order to minimize Type II errors and to accommodate the effect of almost SD,[6] we follow Gasbarro et al. (2007) and use a conservative 5% cut-off point in checking the proportion of test statistics for statistical inference. Using a 5% cut-off point, we conclude that one prospect dominates another prospect only if we find that at least 5% of the statistics are significant.

## 3. Empirical Results and Discussion

[4] Refer to Jarrow (1986) for the conditions.
[5] See Chan et al. (2012) and the references contained therein for further information.
[6] Almost SD allows a small area violation computed from the compared distributions to reveal a preference for "most" decision makers but not for "all" of them. Readers may refer to Leshno and Levy, 2002for more information.

This section discusses empirical results with different tests and approaches in various sub-sections in detail.

### 3.1 Performance

To evaluate the performance of the stock indices, we use the MV approach and the CAPM statistics to compare the performance of Latin American stocks before and after stock market liberalization.

#### 3.1.1 Mean-Variance Criterion and CAPM Statistics

Table 2 provides the descriptive statistics of four daily stock excess returns for the pre- and post-liberalization periods. Academics and practitioners are interested in testing whether stock market liberalization could improve the performance of stock markets. One could provide an answer by checking whether the mean return after merging is higher and the volatility is smaller. From table 2, we find that risk averters prefer the post-liberalization period's stock market in the case of Brazil, Jamaica and Trinidad and Tobago by the MV criterion, since they all have a bigger mean and smaller standard deviation. However, using the MV criterion for risk seekers, we conclude that risk seekers would prefer to invest in the post-liberalization period in the case of Mexico, since it has both bigger mean and standard deviation. However, there is no dominance between the pre- and post- liberalization periods by the MV criterion in the case of Chile and Peru.

On the other hand, from table 2, the *t* statistics are not significant for all stock market indices. Thus, although one may suggest that market liberalization could, in general, result in a higher return for the index after liberalization, no index significantly improves its performance after liberalization. Among them, the F-statistic of the returns between the pre- and post-liberalization periods is significantly smaller than unity only for Brazil and Trinidad and Tobago. This result implies that although Brazil, Mexico, Jamaica and Trinidad and Tobago are more volatile after market liberalization, only Brazil and Trinidad and Tobago are significantly more volatile after market liberalization.

Table 2: Descriptive Statistics of the returns

| Variable | | Mean | StdDev | Skewness | Kurtosis | J-B test | ADF | t-test | F-test |
|---|---|---|---|---|---|---|---|---|---|
| Brazil | 1 year Pre | -0.0009 | 0.0503 | $0.4453^{***}$ | $4.6028^{***}$ | $239.02^{***}$ | $-10.44^{***}$ | | |
| | 1 year Post | 0.0033 | 0.0422 | -0.0699 | -0.325 | 1.36 | $-16.43^{***}$ | -1.0367 | $1.4149^{***}$ |
| | 10 years Post | 0.0005 | 0.0327 | -0.0427 | $3.7829^{***}$ | $1556.44^{***}$ | $-23.37^{***}$ | -0.6495 | $2.3625^{***}$ |
| Mexico | Pre | 0.0014 | 0.0157 | -0.0942 | $1.4164^{***}$ | $22.12^{***}$ | $-7.71^{***}$ | | |
| | 1 year Post | 0.0022 | 0.0165 | $-0.7736^{***}$ | $3.9074^{***}$ | $192.07^{***}$ | $-14.84^{***}$ | -0.5942 | 0.9043 |
| | 10 years Post | 0.0004 | 0.0203 | $-0.8202^{***}$ | $14.7025^{***}$ | $23782.08^{***}$ | $-8.46^{***}$ | 0.7061 | $0.5970^{***}$ |
| Chile | Pre | 0.0027 | 0.0211 | -0.113 | 0.4019 | 2.29 | $-14.05^{***}$ | | |
| | 1 year Post | 0.0012 | 0.0154 | 0.1257 | $1.2089^{***}$ | $16.52^{***}$ | $-5.87^{***}$ | 0.9727 | $1.8717^{***}$ |
| | 10 years Post | 0.0001 | 0.0129 | $0.3097^{***}$ | $4.8742^{***}$ | $2624.38^{***}$ | $-11.28^{***}$ | 2.9022 | $2.6859^{***}$ |
| Peru | Pre | 0.0023 | 0.0234 | -0.2617 | $2.3309^{***}$ | $61.59^{***}$ | $-6.85^{***}$ | | |
| | 1 year Post | 0.0022 | 0.0302 | 0.0039 | $1.2366^{***}$ | $16.57^{***}$ | $-11.60^{***}$ | 0.0482 | $0.5986^{***}$ |
| | 10 years Post | 0.0004 | 0.0165 | 0.0875 | $6.0065^{***}$ | $3925.32^{***}$ | $-8.50^{***}$ | 1.7102 | $1.9999^{***}$ |
| Jamaica | Pre | 0.0014 | 0.0225 | $-0.3262^{**}$ | $4.5925^{***}$ | $233.1^{***}$ | $-4.13^{***}$ | | |
| | 1 year Post | 0.0021 | 0.0327 | $-0.4268^{***}$ | $12.2510^{***}$ | $1646.41^{***}$ | $-3.97^{***}$ | -0.2829 | $0.4735^{***}$ |
| | 10 years Post | 0.0001 | 0.0195 | $0.2706^{***}$ | $16.3804^{***}$ | $29200.04^{***}$ | $-7.88^{***}$ | 0.9511 | $1.3308^{***}$ |
| Trinidad & Tobago | Pre | 0.0004 | 0.0195 | $1.2244^{***}$ | $10.2076^{***}$ | $1198.33^{***}$ | $-4.49^{***}$ | | |
| | 1 year Post | 0.0015 | 0.0183 | $3.2391^{***}$ | $17.8748^{***}$ | $3931.05^{***}$ | $-12.95^{***}$ | -0.6288 | 1.1358 |
| | 10 years Post | 0.0005 | 0.0098 | $4.1536^{***}$ | $48.3305^{***}$ | $261326.98^{***}$ | $-7.17^{***}$ | -0.0350 | $3.9065^{***}$ |

Notes: Lag orders of the ADF test are determined based on AIC. *, **, and *** denote significance at the 10%, 5% and 1% level, respectively.

Table 3 provides the CAPM statistics of four daily stock excess returns for the pre- and post-liberalization periods.

For the CAPM statistics, all Sharpe ratios and the Jensen index are negative. We note that the Sharpe ratios in the post-liberalization period are bigger than in the pre-liberalization period except for the case of Trinidad and Tobago. The Sharpe (1966) ratio is the most conventional formula used in stock evaluation. The Sharpe ratio measures the excess return per unit of risk, where the risk is determined by standard deviation. The higher the Sharpe ratio value, the better the portfolio's returns relative to its risk or the larger the excess return per unit of risk in a portfolio. We find that the Jensen index ratios in the post-liberalization period are bigger than in the pre-liberalization period except for the case of Mexico. A higher Jensen index suggests a higher level of return given the level of risk (systematic or market) on the investment. A low Jensen index, such as a negative number, indicates inferior performance given the level of risk.

On the other hand, the Treynor index has mixed results. The Treynor index for Brazil, Peru and Jamaica in the post-liberalization period is better than in the pre-liberalization period, while Mexico, Chile and Trinidad and Tobago show the reverse relationship. The Treynor (1965) ratio takes into account the systematic risk or market volatility as its measure of risk instead of the standard deviation, as in the Sharpe ratio (1966). Treynor (1966) noted that the relationship of the excess fund return to the beta lies along the security market line. In short, our results using the mean-variance approach and the CAPM statistics confirm that market liberalization could result in the index showing marginally higher returns and more volatility after liberalization, while our results using the CAPM statistics confirm that the performance of the six stock indices improves after liberalization.

Table 3: CAPM Statistics

| Variable | | Beta | Sharpe | Treynor | Jensen |
|---|---|---|---|---|---|
| Brazil | Pre | 0.3438 | -0.0174 | -0.0025 | -0.0009 |
| | 1 Year Post | 0.3786 | 0.0747 | 0.0083 | 0.0032 |
| | 10 Years Post | 3.1930 | 0.0162 | 0.0002 | 0.0005 |
| Mexico | Pre | -0.1387 | 0.1116 | -0.0115 | 0.0016 |
| | 1 Year Post | 0.2731 | 0.1373 | 0.0079 | 0.0022 |
| | 10 Years Post | 2.2475 | 0.0210 | 0.0002 | 0.0004 |
| Chile | Pre | 0.4387 | 0.1510 | 0.0072 | 0.0030 |
| | 1 Year Post | 0.1586 | 0.0594 | 0.0054 | 0.0009 |
| | 10 Years Post | 1.4008 | 0.0072 | 0.0001 | 0.0001 |
| Peru | Pre | 0.3696 | 0.1174 | 0.0073 | 0.0025 |
| | 1 Year Post | 0.6562 | 0.0613 | 0.0028 | 0.0021 |
| | Post | 1.1228 | 0.0196 | 0.0003 | 0.0004 |
| Jamaica | Pre | -0.0018 | 0.0707 | -0.8149 | 0.0014 |
| | 1 Year Post | 0.0775 | 0.0573 | 0.0237 | 0.0019 |
| | 10 Years Post | 0.0406 | 0.0057 | 0.0027 | 0.0001 |
| Trinidad &Tobago | Pre | -0.1838 | 0.0309 | -0.0031 | 0.0006 |
| | 1 Year Post | 0.0802 | 0.1355 | 0.0281 | 0.0022 |
| | 10 Years Post | 0.0033 | 0.0510 | 0.1413 | 0.0005 |

Notes: *, **, and *** denote significance at the 10%, 5% and 1% level, respectively.

### 3.2 Degree of Efficiency

#### 3.2.1 Runs Test

The results of the runs test for daily excess returns, which do not depend on the normality of returns, are presented in table 4. From table 4, we see that all the values of the *Z*-statistic pre-liberalization are statistically significant at the 5 % level in both the pre- and post-liberalization periods except for Brazil, Jamaica and Trinidad and Tobago. The *Z*-statistics of Brazil and Jamaica are statistically insignificant at the 10 % level in both the pre- and post-liberalization periods. The *Z*-statistics of Trinidad and Tobago on the other hand are only statistically significant at the 5 % level in the pre-liberalization period. The null hypothesis of randomness is rejected in both the pre- and post-liberalization periods for Mexico, Chile, Peru and Trinidad & Tobago. The results also show that for all countries except Peru the absolute value of the Z-statistic is smaller in the post-liberalization period. These results could imply that almost all countries are more efficient after stock market liberalization.

Table 4: Results of Runs Test

| Variable | | 1 Year Period | | | | | | 10 Years Period | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | | $n_-$ | $n_+$ | Total Cases | Number of Runs | Z-statistic | p-value | $n_-$ | $n_+$ | Total Cases | Number of Runs | Z-statistic | p-value |
| Brazil | Pre | 131 | 130 | 261 | 130 | -0.1858 | 0.8526 | 131 | 130 | 261 | 130 | -0.1858 | 0.8526 |
| | Post | 129 | 131 | 260 | 143 | 1.492 | 0.1356 | 1299 | 1310 | 2609 | 1293 | -0.4886 | 0.6251 |
| Mexico | Pre | 132 | 128 | 260 | 105 | -3.228*** | 0.0012 | 132 | 128 | 260 | 105 | -3.228*** | 0.0012 |
| | Post | 124 | 137 | 261 | 114 | -2.136** | 0.0327 | 1302 | 1306 | 2608 | 1181 | -6.9128*** | <0.001 |
| Chile | Pre | 135 | 124 | 259 | 108 | -2.778*** | 0.0055 | 135 | 124 | 259 | 108 | -2.778*** | 0.0055 |
| | Post | 138 | 122 | 260 | 114 | -2.059** | 0.0395 | 1333 | 1276 | 2609 | 1128 | -6.9304*** | <0.001 |
| Peru | Pre | 122 | 137 | 259 | 104 | -3.257*** | 0.0011 | 122 | 137 | 259 | 104 | -3.257*** | 0.0011 |
| | Post | 136 | 124 | 260 | 106 | -3.079*** | 0.0021 | 1364 | 1245 | 2609 | 1105 | -7.7621*** | <0.001 |
| Jamaica | Pre | 132 | 128 | 260 | 123 | -0.991 | 0.3219 | 132 | 128 | 260 | 123 | -0.991 | 0.3219 |
| | Post | 136 | 126 | 262 | 123 | -1.092 | 0.2748 | 1386 | 1223 | 2609 | 1288 | -0.4879 | 0.6257 |
| Trinidad &Tobago | Pre | 138 | 123 | 261 | 151 | 2.480** | 0.0131 | 138 | 123 | 261 | 151 | 2.480** | 0.0131 |
| | Post | 153 | 108 | 261 | 130 | 0.304 | 0.761 | 1459 | 1149 | 2608 | 1257 | -1.1751 | 0.2399 |

Note: This table presents the results of the runs test of the excess returns that are computed by deducting the return of the MSCI World Index from each of the return series. *, **, and *** denote significance at the 10%, 5%, and 1% level, respectively.

### 3.2.2 Variation Ratio Tests

The results of the multiple variance ratio test developed by Chow and Denning (1993) are presented in table 5. In table 5, both $Z_1^*(q)$ and $Z_2^*(q)$ of all Latin American countries are statistically significant at the 1% level for both the pre- and post-liberalization periods. These results reject the random walk hypothesis under both homoscedastic and heteroscedastic situations. However, it should also be noted that the $Z$-statistics $(Z_1^*(q), Z_2^*(q))$ of Mexico and Jamaica become smaller in the post-liberalization period, which shows a possible tendency to random walk.

Table 5: Chow-Denning (1993) Multiple Variation Ratio Test Statistics

| Country Indexes | **1 Year Pre** | | **1 Year Post** | | **10 Years Post** | |
|---|---|---|---|---|---|---|
| | $Z_1^*(q)$ | $Z_2^*(q)$ | $Z_1^*(q)$ | $Z_2^*(q)$ | $Z_1^*(q)$ | $Z_2^*(q)$ |
| Brazil | $50.03^{***}$ | $34.84^{***}$ | $46.64^{***}$ | $38.83^{***}$ | $169.65^{***}$ | $108.98^{***}$ |
| Mexico | $28.99^{***}$ | $15.59^{***}$ | $35.29^{***}$ | $21.22^{***}$ | $169.04^{***}$ | $102.49^{***}$ |
| Chile | $45.97^{***}$ | $33.34^{***}$ | $38.92^{***}$ | $28.71^{***}$ | $169.35^{***}$ | $108.44^{***}$ |
| Peru | $45.59^{***}$ | $32.03^{***}$ | $47.78^{***}$ | $34.49^{***}$ | $169.88^{***}$ | $74.86^{***}$ |
| Jamaica | $50.23^{***}$ | $43.01^{***}$ | $42.95^{***}$ | $27.12^{***}$ | $170.50^{***}$ | $88.06^{***}$ |
| Trinidad & Tobago | $31.23^{***}$ | $22.38^{***}$ | $47.83^{***}$ | $41.28^{***}$ | $170.41^{***}$ | $115.47^{***}$ |

Note: $Z_1^*(q)$ is the variance ratio test statistic assuming homoscedasticity. $Z_2^*(q)$ is the variance ratio test statistic, assuming heteroscedasticity. *, **, and *** denote significance at the 10%, 5% and 1% level, respectively. The 10%, 5% and 1% critical values are 2.226268, 2.490915 and 3.022202, respectively.

Table 6 describes the rank and sign test developed by Wright (2000). In order to draw a better picture of the comparison, we adopt the common practice and select lags 2, 4, 8, and 16 in the testing procedure. From table 6, the rank-based test results show that R1 and R2 are statistically significant at the 1% level in both the pre- and post-liberalization periods. The sign-based test results are similar to the rank-based test results. Overall, for every Lag = 2,4,8,16, the null hypothesis that Latin American stock market indices are following a random walk is rejected in both the pre- and post-liberalization periods.

Table 6: Wright (2000) Variance Ratio Test Statistics – Ranks and Signs

| Variables | | 1 Year Pre-Liberalization | | | | 1 Year Post Liberalization | | | | 10 Years Post Liberalization | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | | Number of Lag (q) | | | | | | | | | | | |
| | | 2 | 4 | 8 | 16 | 2 | 4 | 8 | 16 | 2 | 4 | 8 | 16 |
| Brazil | R1 | 15.72$^{***}$ | 24.74$^{***}$ | 35.14$^{***}$ | 47.75$^{***}$ | 15.60$^{***}$ | 24.40$^{***}$ | 34.26$^{***}$ | 44.03$^{***}$ | 50.92$^{***}$ | 81.45$^{***}$ | 119.63$^{***}$ | 170.74$^{***}$ |
| | R2 | 15.68$^{***}$ | 24.66$^{***}$ | 35.09$^{***}$ | 47.83$^{***}$ | 15.42$^{***}$ | 24.00$^{***}$ | 33.07$^{***}$ | 41.29$^{***}$ | 50.76$^{***}$ | 80.99$^{***}$ | 118.26$^{***}$ | 167.33$^{***}$ |
| | S1 | 16.03$^{***}$ | 25.51$^{***}$ | 37.06$^{***}$ | 51.68$^{***}$ | 16.00$^{***}$ | 25.46$^{***}$ | 36.98$^{***}$ | 51.57$^{***}$ | 51.04$^{***}$ | 81.78$^{***}$ | 120.50$^{***}$ | 173.00$^{***}$ |
| Mexico | R1 | 15.47$^{***}$ | 23.86$^{***}$ | 32.56$^{***}$ | 39.70$^{***}$ | 15.59$^{***}$ | 24.21$^{***}$ | 33.53$^{***}$ | 42.60$^{***}$ | 50.90$^{***}$ | 81.40$^{***}$ | 119.45$^{***}$ | 169.90$^{***}$ |
| | R2 | 15.14$^{***}$ | 22.82$^{***}$ | 29.66$^{***}$ | 33.71$^{***}$ | 15.15$^{***}$ | 23.01$^{***}$ | 30.32$^{***}$ | 35.98$^{***}$ | 50.72$^{***}$ | 80.92$^{***}$ | 118.15$^{***}$ | 166.78$^{***}$ |
| | S1 | 16.00$^{***}$ | 25.46$^{***}$ | 36.98$^{***}$ | 51.57$^{***}$ | 16.03$^{***}$ | 25.51$^{***}$ | 37.06$^{***}$ | 51.68$^{***}$ | 51.03$^{***}$ | 81.77$^{***}$ | 120.48$^{***}$ | 172.96$^{***}$ |
| Chile | R1 | 15.88$^{***}$ | 25.06$^{***}$ | 35.81$^{***}$ | 48.12$^{***}$ | 15.56$^{***}$ | 24.18$^{***}$ | 33.77$^{***}$ | 43.77$^{***}$ | 50.91$^{***}$ | 81.36$^{***}$ | 119.31$^{***}$ | 169.73$^{***}$ |
| | R2 | 15.79$^{***}$ | 24.81$^{***}$ | 35.11$^{***}$ | 45.42$^{***}$ | 15.32$^{***}$ | 23.36$^{***}$ | 31.44$^{***}$ | 38.89$^{***}$ | 50.72$^{***}$ | 80.77$^{***}$ | 117.68$^{***}$ | 166.02$^{***}$ |
| | S1 | 15.29$^{***}$ | 24.05$^{***}$ | 34.28$^{***}$ | 45.96$^{***}$ | 15.88$^{***}$ | 25.13$^{***}$ | 36.10$^{***}$ | 50.26$^{***}$ | 50.84$^{***}$ | 81.43$^{***}$ | 119.79$^{***}$ | 171.88$^{***}$ |
| Peru | R1 | 15.67$^{***}$ | 24.53$^{***}$ | 34.65$^{***}$ | 45.94$^{***}$ | 15.85$^{***}$ | 24.99$^{***}$ | 35.70$^{***}$ | 48.06$^{***}$ | 50.93$^{***}$ | 81.43$^{***}$ | 119.51$^{***}$ | 170.22$^{***}$ |
| | R2 | 15.29$^{***}$ | 23.47$^{***}$ | 32.16$^{***}$ | 40.91$^{***}$ | 15.74$^{***}$ | 24.74$^{***}$ | 35.11$^{***}$ | 46.52$^{***}$ | 50.89$^{***}$ | 81.34$^{***}$ | 119.35$^{***}$ | 170.01$^{***}$ |
| | S1 | 16.00$^{***}$ | 25.46$^{***}$ | 36.98$^{***}$ | 51.57$^{***}$ | 16.00$^{***}$ | 25.46$^{***}$ | 36.98$^{***}$ | 51.57$^{***}$ | 51.04$^{***}$ | 81.78$^{***}$ | 120.50$^{***}$ | 173.00$^{***}$ |
| Jamaica | R1 | 15.64$^{***}$ | 24.70$^{***}$ | 35.39$^{***}$ | 47.75$^{***}$ | 15.78$^{***}$ | 24.79$^{***}$ | 34.95$^{***}$ | 45.77$^{***}$ | 50.90$^{***}$ | 81.39$^{***}$ | 119.38$^{***}$ | 169.66$^{***}$ |
| | R2 | 15.51$^{***}$ | 24.38$^{***}$ | 34.64$^{***}$ | 45.84$^{***}$ | 15.44$^{***}$ | 23.98$^{***}$ | 33.26$^{***}$ | 42.77$^{***}$ | 50.87$^{***}$ | 81.35$^{***}$ | 119.33$^{***}$ | 169.62$^{***}$ |
| | S1 | 16.00$^{***}$ | 25.46$^{***}$ | 36.98$^{***}$ | 51.57$^{***}$ | 16.06$^{***}$ | 25.56$^{***}$ | 37.13$^{***}$ | 51.79$^{***}$ | 51.04$^{***}$ | 81.78$^{***}$ | 120.50$^{***}$ | 173.00$^{***}$ |
| Trinidad & Tobago | R1 | 12.37$^{***}$ | 19.12$^{***}$ | 26.94$^{***}$ | 33.45$^{***}$ | 15.66$^{***}$ | 24.64$^{***}$ | 34.97$^{***}$ | 46.17$^{***}$ | 50.94$^{***}$ | 81.53$^{***}$ | 119.85$^{***}$ | 171.26$^{***}$ |
| | R2 | 11.11$^{***}$ | 17.25$^{***}$ | 24.56$^{***}$ | 30.07$^{***}$ | 15.03$^{***}$ | 23.59$^{***}$ | 33.27$^{***}$ | 42.81$^{***}$ | 50.72$^{***}$ | 81.07$^{***}$ | 118.89$^{***}$ | 168.93$^{***}$ |
| | S1 | 16.03$^{***}$ | 25.51$^{***}$ | 37.06$^{***}$ | 51.68$^{***}$ | 16.03$^{***}$ | 25.51$^{***}$ | 37.06$^{***}$ | 51.68$^{***}$ | 51.03$^{***}$ | 81.77$^{***}$ | 120.48$^{***}$ | 172.96$^{***}$ |

Note: R1 and R2 are rank-based variance ratio test statistics, which are defined in equations (6) and (7), S1 is the sign-based variance ratio test statistic, which is defined in equation (8). The *, **, and *** denote significance at 10%, 5% and 1% level, respectively.

### 3.2.3 Martingale Hypothesis Test

The wild bootstrap Cramer-von Mises test statistic (Cp) and wild bootstrap Kolmogorov-Smirnov test statistic (Kp) by Dominguez and Lobato (2003) are reported in table 7. For the pre-liberalization, except for Jamaica, we reject the null hypothesis that the return is a martingale difference sequence. These results show that Jamaica has an efficient stock market but the others do not. For the post-liberalization, except for Brazil, we reject the null hypothesis that the return is a martingale difference sequence. This shows that Brazil is an efficient market and the others are not efficient markets. Thus, only Brazil's results show that liberalization has improved market efficiency when we compare the pre-liberalization 1-year period to the post-liberalization 1-year period. However, this result doesn't hold when we compare the pre-liberalization 1-year period to the post-liberalization 10-year period. Overall, our results do provide evidence that liberalization has improved market efficiency in Latin American stock markets.

**Table 7:** Dominguez and Lobato (2003) test statistics – $C_p$ and $K_p$

| Variable | 1 Year Pre | | 1 Year Post | | 10 Years Post | |
|---|---|---|---|---|---|---|
| | $C_p$ | $K_p$ | $C_p$ | $K_p$ | $C_p$ | $K_p$ |
| Brazil | 0.13 | 0.735*** | 0.072 | 0.605 | 0.8963** | 1.7786*** |
| Mexico | 1.389*** | 1.818*** | 0.256*** | 0.903*** | 6.6550*** | 3.6004*** |
| Chile | 0.607*** | 1.335*** | 0.386*** | 1.378*** | 8.2128*** | 4.1207*** |
| Peru | 1.244*** | 1.901*** | 2.559*** | 2.351*** | 16.0295*** | 5.2880*** |
| Jamaica | 0.064 | 0.633 | 0.182 | 0.832*** | 0.2003 | 0.9767 |
| Trinidad & Tobago | 1.785*** | 2.518*** | 0.101 | 0.750*** | 0.1509 | 0.9866 |

Note: In the wild bootstrap Cramer-von Mises test ($C_p$) and wild bootstrap Kolmogorov-Smirnov test ($K_p$), the number of bootstrap replications is 500, and the lag value p is 1. *, **, and *** denote significance at the 10%, 5% and 1% level, respectively.

### 3.2.4 Stochastic Dominance Approach

We will employ the DD test to examine investors' preference toward liberalization, check whether there is an arbitrage opportunity due to liberalization and examine market efficiency.

Tables 8 and 9 report the percentage of significant modified DD statistics over the negative domain (losses), positive domain (gains) and the entire return distribution (both) for risk averters and risk seekers, respectively. We find that the post-liberalization excess returns do not stochastically dominate the pre-liberalization excess returns at the first three orders for all six indices. We explain the results in more detail below.

As can be seen from table 8, no $T_1$ (ASD and DSD) is significantly positive and negative, indicating that post-liberalization excess returns and e pre-liberalization excess returns do not dominate each other in the sense of FSD. Moreover, the pre- and post-liberalization excess returns also do not dominate each other in both SSD and TSD because both $T_2$ and $T_3$ $T_1$ (ASD and DSD) are not significantly positive and negative. The results show that excess returns in the pre- and post-liberalization period do not dominate each other at the first three orders, implying that there is no "arbitrage opportunity" in the pre- and post-liberalization periods and investors cannot get higher expected wealth by shifting their investment from the pre-liberalization period to the post-liberalization period and vice versa. This result, in turn, shows that all six markets are efficient. To have a better understanding of the pre- and post-liberalization relationship in the sense of FSD, we plot the CDFs for Mexico in figure 2. From figure 2, we notice that the CDFs for both the pre- and post-liberalization periods coincide with each other. This leads us to suspect that there is no difference between pre- and post-liberalization.

Our SD result implies that risk averters are indifferent between the pre- and post-liberalization periods for all indices studied in this paper; there is no arbitrage opportunity owing to liberalization; and the market is efficient and investors are rational.

At last, we note that we have applied the SD test for risk seekers to analyze the data and draw the same conclusion that risk seekers are indifferent between the pre- and post-liberalization periods for all e indices studied in this paper; there is no arbitrage opportunity owing to liberalization; and the market is efficient and investors are rational. Thus, the SD test offers no evidence that stock market liberalization has improved market efficiency.

**Figure 2**: Plot of the CDF of the pre- and post-liberalization periods for Mexico

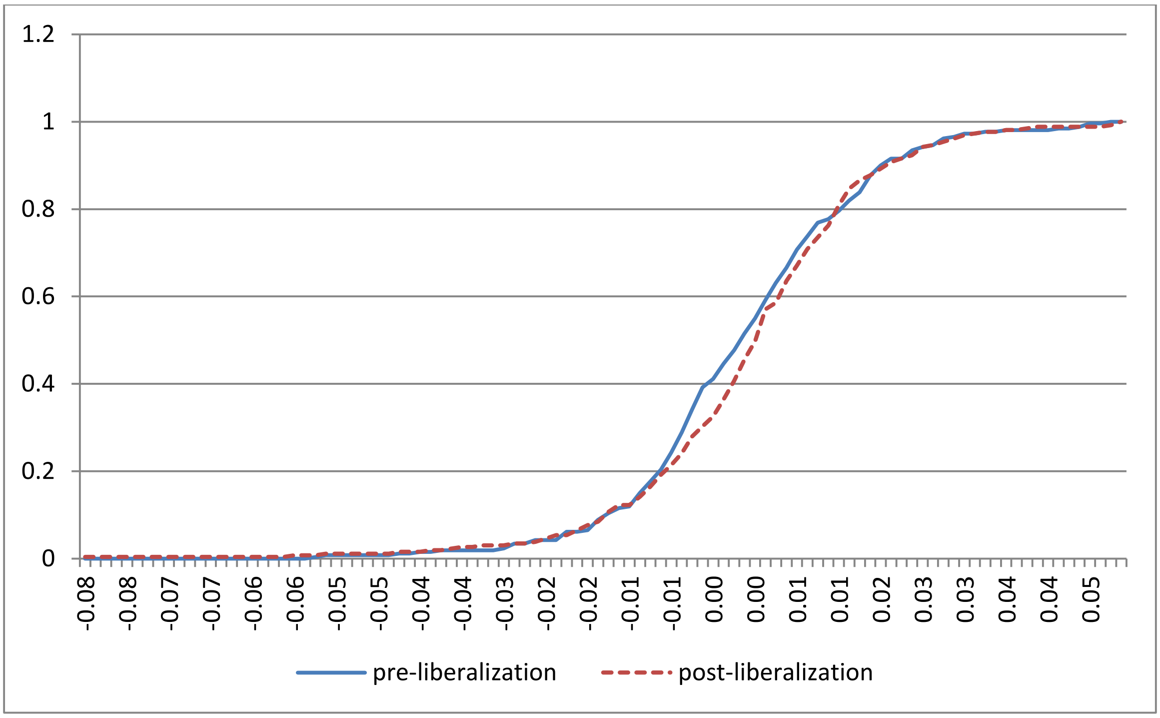


Table 8: Percentages of Significance of Modified Davidson-Duclos Statistics

| | | 1 Year Period | | | | | | 10 Years Period | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| Variable | | FASD | | SASD | | TASD | | FASD | | SASD | | TASD | |
| F=pre, G=post | | $\%T_1^A>0$ | $\%T_1^A<0$ | $\%T_2^A>0$ | $\%T_2^A<0$ | $\%T_3^A>0$ | $\%T_3^A<0$ | $\%T_1^A>0$ | $\%T_1^A<0$ | $\%T_2^A>0$ | $\%T_2^A<0$ | $\%T_3^A>0$ | $\%T_3^A<0$ |
| Brazil | + | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 4 | 0 | 0 | 0 |
| | - | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 |
| Mexico | + | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 |
| | - | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 |
| Chile | + | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 |
| | - | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 |
| Peru | + | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 |
| | - | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 |
| Jamaica | + | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 |
| | - | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 |
| Trinidad & Tobago | + | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 |
| | - | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 |

Note: The numbers in the columns of FSD, SSD and TSD indicate the percentages of the modified DD statistics significantly in the positive domain (+) and negative domain (-) at the 5% level. $T_j$ is defined in (14). $F$ is the return series for the pre-liberalization period, while $G$ is the return series for the post- liberalization period.

Table 9: Percentages of Significance of Modified Davidson-Duclos Statistics

| | | 1 Year Period | | | | | | 10 Years Period | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| Variable | | FDSD | | SDSD | | TDSD | | FDSD | | SDSD | | TDSD | |
| F=pre, G=post | | $\%T_1^D > 0$ | $\%T_1^D < 0$ | $\%T_2^D > 0$ | $\%T_2^D < 0$ | $\%T_3^D > 0$ | $\%T_3^D < 0$ | $\%T_1^D > 0$ | $\%T_1^D < 0$ | $\%T_2^D > 0$ | $\%T_2^D < 0$ | $\%T_3^D > 0$ | $\%T_3^D < 0$ |
| Brazil | + | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 |
| | - | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 |
| Mexico | + | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 |
| | - | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 |
| Chile | + | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 |
| | - | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 |
| Peru | + | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 |
| | - | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 |
| Jamaica | + | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 |
| | - | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 |
| Trinidad & Tobago | + | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 |
| | - | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 | 0 |

Note: The numbers in the columns of FSD, SSD and TSD indicate the percentages of the modified DD statistics significantly in the positive domain (+) and negative domain (-) at the 5% level. $T_j$ is defined in (16). $F$ is the return series for the pre-liberalization period, while $G$ is the return series for the post-liberalization period.

## 4. Conclusion

The impact of liberalization on stock market efficiency still remains a puzzle in the literature, especially for Latin American stock markets. This investigation is among the first to examine the impact of stock market liberalization on the efficiency of stock markets and also among the first to examine the impact of stock market liberalization on the efficiency of Latin American stock markets.

Daily stock indices from Brazil, Mexico, Chile, Peru, Jamaica and Trinidad and Tobago are used for the analysis. We compare stock returns one year before, one year after, and ten years after liberalization for a particular stock exchange. We first employ the mean-variance approach and CAPM statistics to evaluate stock indices' performance. To test the randomness, in order to have a more reliable result, we employ several approaches, including the runs test, the Chow-Denning multiple variation ratio test, the Wright variance ratio test, the martingale hypothesis test and the SD test, to Latin American stock market indices. The runs test implies that these markets' efficiencies are more or less the same after liberalization. The results of the Chow-Denning multiple variation ratio test statistics and the Wright variance ratio test statistics – ranks and signs — show that the indices studied in this paper do not improve their randomness after the introduction of liberalization. The martingale hypothesis test shows that only Brazil's efficiency is improved after liberalization, but this result doesn't hold when we compare the pre-liberalization period with the period 10 years after liberalization. The results from the SD test imply that risk averters and seekers are indifferent between the pre- and post-liberalization periods for all indices studied in this paper; there is no arbitrage opportunity owing to liberalization; and the market is efficient and investors are rational. All of our tests offer no evidence that stock market liberalization has improved stock market efficiency in Latin America. We believe our results are quite reliable, since we use several different approaches and they all give the same result. Therefore, one may need to consider that liberalization may not necessarily significantly improve stock market efficiency when it comes to countries that are similar to Latin America.

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