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Random walk versus breaking trend in stock prices: Evidence from emerging markets

Kausik Chaudhuri [a], Yangru Wu [b,c,*]

[a] *Indira Gandhi Institute of Development Research, Mumbai 400 065, India*
[b] *Department of Finance and Economics, Faculty of Management, Rutgers University, Newark, NJ 07102-1820, USA*
[c] *Hong Kong Institute for Monetary Research, Hong Kong, China*

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Abstract

This paper investigates whether stock-price indexes of seventeen emerging markets can be characterized as random walk (unit root) or mean reversion processes. We implement a test that can account for structural breaks in the underlying series and is more powerful than standard tests. We find that for fourteen countries, stock prices exhibit structural breaks. Furthermore, for ten countries, the null hypothesis of a random walk can be rejected at the one or 5% significance level. Our results indicate that ignoring structural breaks that arise from the liberalization of emerging markets can lead to incorrect inference that these indices are characterized by random walks, and are consistent with the points made by Bekaert et al. [J. Int. Money Finan. 21 (2002) 295]. Our findings hold true regardless of whether stock indexes are denominated in US dollar terms, in local currencies terms, or in real terms.

JEL classification: G15; G14; C22
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1. Introduction

Economists have shown considerable interest in the long-run time-series properties of stock prices, with a particular attention being paid to investigate whether stock

* Corresponding author. Tel.: +1-973-353-1146; fax: +1-973-353-1233.
E-mail address: yangruwu@andromeda.rutgers.edu (Y. Wu).

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prices can be characterized as unit root (random walk) or mean reverting (trend stationary) processes. [1] If stock price follows a mean reverting process, then there exists a tendency for the price level to return to its trend path over time and investors may be able to forecast future returns by using information on past returns. On the other hand, a random walk process says that any shock to stock price is permanent and there is no tendency for the price level to return to a trend path over time. This suggests that future returns are unpredictable based on historical observations. The random walk property also implies that the volatility of stock price can grow without bound in the long run. These time-series properties are not only of interest by themselves but also have important implications for asset pricing.

The existence of mean reversion in stock prices is subject to much controversy. Fama and French (1988) and Poterba and Summers (1988) first document evidence that mean reversion exists in US stock prices. Others are skeptical of these results. For example, using variance-ratio tests, Lo and MacKinlay (1988) report some evidence against mean reversion in weekly US data. Kim et al. (1991) argue that the mean reversion results are only detectable in pre-war data, while Richardson and Stock (1989) and Richardson (1993) find that after correcting for small-sample biases, the Fama–French and Poterba–Summers results may be reversed. McQueen (1992) demonstrates that the results of mean reversion in US stock prices are not robust to outliers or alternative distributional assumptions. A number of researchers have also examined the mean reversion property using international data. For example, Richards (1997) finds evidence of long-term winner–loser reversals for equity index prices for sixteen countries. Balvers et al. (2000) report significant evidence of mean reversion in annual equity indexes for a sample of eighteen developed countries and demonstrate that one can exploit the property of mean reversion to predict annual equity returns using a parametric contrarian investment strategy.

In this paper, we re-examine the random walk hypothesis in stock prices of seventeen emerging markets using monthly data from 1985 through 1997. [2] Compared to developed markets, emerging markets are relatively isolated from capital markets of other countries and have relatively low correlation with developed markets, especially the United States. Therefore, our study provides particularly interesting information from this independent sample, and will complement the existing studies on developed markets. To summarize our results at the outset, we find that among the seventeen stock markets indexes investigated, fourteen of them are subject to structural breaks at the 5% significance level. These breaks either appear in the constant intercept, in the scope of the trend function or in both. Indeed, when the break points are properly accounted for, we show that ten of these series can be reasonably

[1] Here we use the term "random walk" in a loose sense to simply mean that price index is a nonstationary process with a unit root. Therefore, we will use the terms "random walk" and "unit root" interchangeably throughout the paper.

[2] Researchers have recently paid a particular attention to the study of stock price behavior in emerging markets. See, for example, Harvey (1995), Claessens et al. (1995), and Cheung and Lai (1995), among others.

characterized as mean reverting processes rather than as random walk processes at the one or 5% significance level.

Our econometric tests are motivated by the observations that equity markets in these countries might have been affected by shocks because of financial market liberalization and/or other structural changes in the underlying economies. According to the recent study by the International Finance Corporation (1997, IFC), many of these countries are liberalizing their financial markets to various degrees during the sampling period (see Table 1). The existence of possible structural breaks in these markets can also be visualized from the time-series plots of these market indexes. As displayed in Figs. 1–4 for four representative countries in our sample, these indexes in general experience large fluctuations over time and there seem to be some apparent break points. For example, for Greece the equity index level jumped in the late 1980s and the slope of the trend line declined in the 1990s. For Malaysia, a break point seems to occur in both the level and the slope of the trend line around 1993. The market liberalization of the Philippines in 1989 caused a break around the same year. Similarly for Taiwan it appears that a break occurred near its market liberalization date in early 1991. For these countries, if a break point is taken into account, stock prices may be reasonably modeled as a trend-stationary process, rather

Table 1
Market liberalization dates of emerging markets

Country	Market liberalization date	Liberalization status
Argentina	October 1991	Fully open
Brazil	May 1991	49% of voting common stock and 100% of non-voting participating preferred stock
Chile	December 1988	25% of a listed a company's shares[a]
Colombia	February 1991	Fully open
Greece	December 1988	Fully open
India	November 1992	24% of a company's issued share capital
Jordan	December 1988	49% of listed companies' capital
Korea	January 1992	20% on October 1, 1996
Malaysia	December 1988	100% available to foreign investors[b]
Mexico	May 1989	30% of total capital[c]
Nigeria	July 1995	Fully open
Pakistan	February 1991	Fully open
Philippines	October 1989	Investable up to 40%
Taiwan	January 1991	Investable up to 30% on November 1, 1996
Thailand	December 1988	25% for bank stocks and 49% for others
Venezuela	January 1990	100% investable other than bank stocks[d]
Zimbabwe	June 1993	Investable up to 40%

Notes: The table reports market liberalization dates for the seventeen emerging stock markets under investigation in this study. This information is obtained from IFC (1997).

[a] Since January 1996, it is fully open.

[b] Bank Negara, the central bank, restricts foreign ownership of banks to 30% but the statute is not currently enforced.

[c] Certain classes may be freely available to foreign investors.

[d] From January 1994, bank stocks are also fully available to foreign institutional investors.

than as a random walk process. Therefore, in this paper we argue that because of the nature of the particular data, the possible structural breaks should be incorporated when tests for the random walk hypothesis are conducted. To this end, we employ the sequential test procedure developed by Zivot and Andrews (1992) to test for the random walk hypothesis which allows for a one-time change in the constant and/or in the slope of the trend function of the underlying series. Our results are in sharp contrast to those obtained with traditional tests where the possible breaks are not considered. As Perron (1989) and others have demonstrated, failure to take into account the breaking points may significantly reduce the power of traditional tests and mistakenly produce evidence in support of the random walk hypothesis. We find that this may well be the case for emerging equity market data.

Several authors have directly studied the issues of structural breaks and market liberalization in emerging equity markets. See Bekaert et al. (2000, 2002), Kawakatsu and Morey (1999), Henry (2000), Bekaert and Harvey (1998, 2000), Stulz and Wasserfallen (1995), and Errunza et al. (1992). In particular, Bekaert et al. (2000)

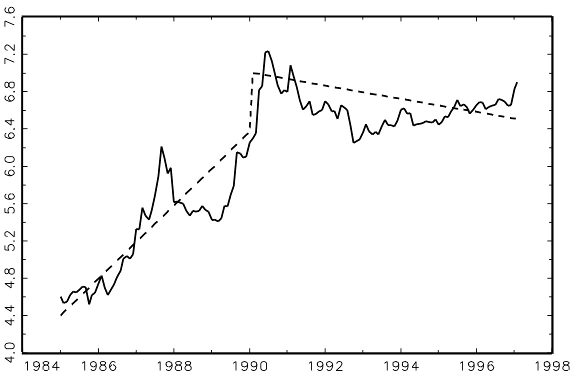

Fig. 1. Time-series plot of stock index: Greece.

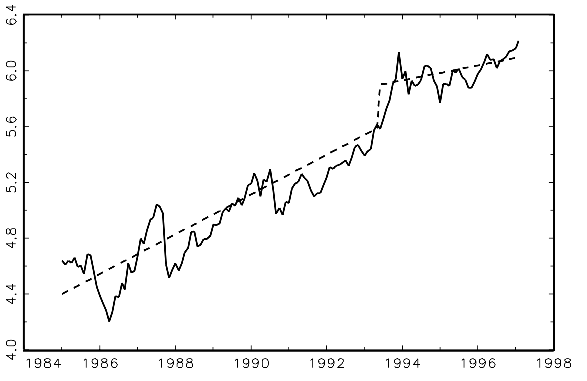

Fig. 2. Time-series plot of stock index: Malaysia.

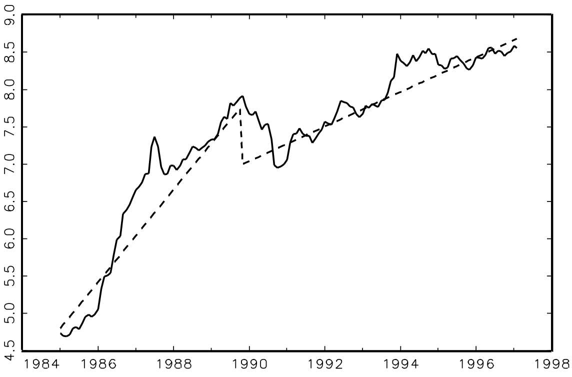

Fig. 3. Time-series plot of stock index: Philippines.

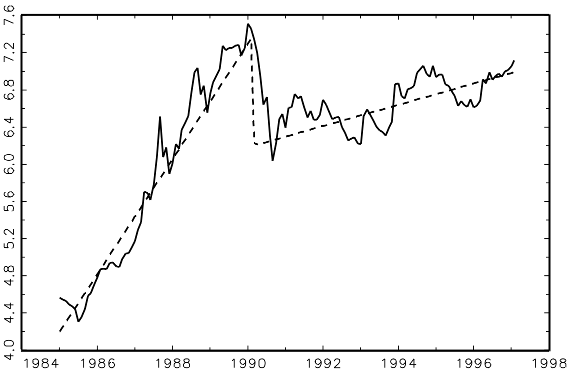

Fig. 4. Time-series plot of stock index: Taiwan.

have provided a comprehensive characterization of the structural changes in emerging markets using prior instrumental variables. They specify a reduced-form model for a number of financial time series (e.g. equity returns and dividend yields) and search for a common break in the process generating the data. Furthermore, Bekaert et al. (2002) employ a vector autoregression to examine multiple break points in capital flows and other variables. Henry (2000) employs an event study methodology to investigate abnormal return and finds on average a 3.3% abnormal return per month during an eight-month window leading up to the implementation of initial stock market liberalization in emerging markets. None of these authors, however, tests for the random walk hypothesis against the alternative hypothesis of mean reversion in emerging equity market. Our focus in this paper is, however, different from these studies. We are primarily interested in testing for mean reversion of equity prices in the presence of structural break, not on directly testing for structural break per se. Nevertheless, the findings on structural breaks in our paper complement those in the aforementioned studies and provide an interesting comparison.

The remainder of the paper is organized as follows. Section 2 presents the empirical methodology. Section 3 describes the data. The main empirical results are reported in Section 4. Section 5 checks for robustness of the results using indexes denominated in local currencies and in real terms, and reports results from the variance-ratio test. The final section concludes the paper.

2. Empirical methodology

Our primary interest in this study is to test whether stock prices in emerging markets follow random walk or mean reverting processes. Let p_t be the natural logarithm of a stock-price index at time t and let the sample size be T. The most popular tests for this hypothesis are the augmented Dickey and Fuller (1979, 1981, ADF) tests and the Phillips–Perron (1988, PP) tests.

Consider the null hypothesis that $\{p_t\}$ is a random walk (unit root) process:

$$p_t = \mu + p_{t-1} + \varepsilon_t \tag{1}$$

where μ is a constant parameter and ε_t is a stationary process which is allowed to be serially correlated. For the augmented Dickey–Fuller (ADF) tests, one estimates the following regressions:

$$\Delta p_t = \mu + \alpha p_{t-1} + \sum_{j=1}^{k} \phi_j \Delta p_{t-j} + \varepsilon_t, \tag{2}$$

$$\Delta p_t = \mu + \beta t + \alpha p_{t-1} + \sum_{j=1}^{k} \phi_j \Delta p_{t-j} + \varepsilon_t. \tag{3}$$

Eq. (2) tests for the null hypothesis of a random walk against a mean stationary alternative, while Eq. (3) tests for the same null against a trend stationary alternative. In both cases, the k extra regressors, Δp_{t-j}, are added to eliminate possible nuisance-parameter dependencies in the asymptotic distributions of the test statistics caused by serial correlation in the error terms. For a given sample, if the estimate of α is not significantly different from zero, then the null hypothesis of a random walk cannot be rejected. On the other hand, if one finds that $\alpha < 0$, then the alternative hypothesis of mean reversion holds. The PP tests work in a similar way except that the extra regressors, Δp_{t-j}, are not included in the regressions, but the serial correlation of the residuals is corrected via a non-parametric approach.

One main shortcoming of the popular ADF and PP tests is that they have quite low power against slow mean reversion alternatives in small samples. Therefore, failure to reject the null hypothesis may not be interpreted as decisive evidence against mean reversion. Perron (1989) points out that these tests perform especially poorly when there is a break in the deterministic trend function and he derives the asymptotic distribution of the test statistic which incorporates the presence of a broken trend. However, Perron's method has been criticized on the grounds that his break point is chosen based on pre-test examination of the data and is hence subject to the problem of "data snooping." One important consequence of using prior information to set the break point is that his procedure will in general overstate the likelihood of the trend-break alternative hypothesis. Zivot and Andrews (1992), among others, have developed methods to endogenously search for a break point and test for the presence of a unit root when the process has a broken trend. [3] They demonstrate that their test is robust and more powerful than the ADF and PP tests, and it is the test that we shall employ in this study. [4]

Let T_B be a potential breaking point in $\{p_t\}$, the Zivot–Andrews sequential test procedure starts by estimating the following three equations:

$$\text{Model (A)}: \Delta p_t = \mu^A + \theta^A DU_t + \beta^A t + \alpha^A p_{t-1} + \sum_{j=1}^{k} \phi_j^A \Delta p_{t-j} + \varepsilon_t, \tag{4}$$

$$\text{Model (B)}: \Delta p_t = \mu^B + \beta^B t + \gamma^B DT_t + \alpha^B p_{t-1} + \sum_{j=1}^{k} \phi_j^B \Delta p_{t-j} + \varepsilon_t, \tag{5}$$

$$\text{Model (C)}: \Delta p_t = \mu^C + \theta^C DU_t + \beta^C t + \gamma^C DT_t + \alpha^C p_{t-1} + \sum_{j=1}^{k} \phi_j^C \Delta p_{t-j} + \varepsilon_t, \tag{6}$$

where the two dummy variables are defined as follows:

$$DU_t = \begin{cases} 1 & \text{if } t > T_B, \\ 0 & \text{otherwise,} \end{cases}$$

$$DT_t = \begin{cases} t & \text{if } t > T_B, \\ 0 & \text{otherwise.} \end{cases}$$

In the above specifications, Model (A) allows for a one-time shift in the intercept; Model (B) allows for a break in the slope of the trend function; while Model (C) includes the hybrid of the two. The selection of the possible break point, T_B, is viewed as the outcome of an estimation procedure designed to fit $\{p_t\}$ by a trend-stationary process with a one-time break in the trend function, which occurs at an unknown point in time. The procedure searches for the break point that gives the most weight to the trend-stationary alternative. Like in the ADF tests, the k extra regressors, Δp_{t-j}, are added to purge serial correlation in the error term. In a sample with T observations, for each specification, to determine the break and to compute the test

[3] Several other authors have also proposed tests for a random walk with an unknown breaking point. See, for example, Banerjee et al. (1992), Christiano (1992), and Perron and Vogelsang (1992). These papers use similar methods to search for break points. We do not intend to exhaust all these tests. For more recent work on testing for structural changes and non-linearity in stock returns and other data, see Kim and Kon (1999) and Bidarkota (2000).

[4] Through extensive Monte-Carlo simulation, Vogelsang and Perron (1998) show that the power of test for a unit root with break is in the range of 60–90% under plausible alternatives, which is higher than traditional unit-root tests without considering break.

statistic for a random walk, an ordinary least squares regression is run with a potential break point at T_B, where T_B ranges from 2 to $T-2$. For each value of T_B, the number of extra regressors, k, is determined using the procedure suggested by Campbell and Perron (1991). That is, we start with a large k_{max} and then estimate the model with k_{max} lags. If the coefficient of the last included lag is significant at the 10% level, select $k = k_{max}$. Otherwise, reduce the lag order by one until the coefficient of the last included lag is significant. It is worth mentioning that the choice of lag length, k, can affect the test results and other procedures to select the lag length have also been proposed in the literature. Ng and Perron (1995) demonstrate that a too parsimonious model can have large size distortions, while an over-parameterized model may result in reduction of test power. But the size problem is in general more severe than the power loss. They show that methods based on sequential tests have an advantage over information-based rules such as the Akaike information criterion and the Schwartz Bayesian information criterion because the former have less size distortions and have comparable power. The procedure adopted in this paper falls into this category of the general-to-specific sequential procedures. Once the optimal lag length is chosen, the t-statistic for testing whether the first-lag coefficient is zero, i.e., $\alpha = 0$, is computed. The test statistic for a random walk is the minimum t-statistic over all $T-2$ regressions and the break point is the time corresponding to such a statistic.

3. The data

The data used in this paper are obtained from International Finance Corporation's *Emerging Market Database* (IFC-EMDB). The sample is monthly from January 1985 to February 1997 with 146 observations and contains US dollar as well as local currency denominated stock indexes for the following seventeen countries: Argentina, Brazil, Chile, Colombia, Greece, India, Jordan, Korea, Malaysia, Mexico, Nigeria, Pakistan, Philippines, Taiwan, Thailand, Venezuela, and Zimbabwe. Our analysis will be primarily focused on the indexes denominated in the US dollar. The indexes include dividends and capital gains and are end-of-month quotes. [5] We choose to use the IFC indexes rather than other local stock price indexes for several reasons. First, these indexes are constructed on a consistent basis by the IFC, making cross-country comparison more meaningful. Second, these indexes include the most active traded stocks in the respective local markets and cover at least 60% of market capitalization. Third, the IFC-EMDB has been used in numerous recent studies.

4. Empirical results

All results reported in this section are based on indexes denominated in the US dollar. For the purpose of comparison, we first apply the standard ADF and PP tests

[5] We have also done the same analysis using stock price indexes without dividends. The results are qualitatively the same. They are not reported but are available upon request.

without breaks to each series and report the results in Table 2. The model is estimated both with and without a time trend. For the ADF tests, the lag length, k, is optimally chosen using the sequential procedure suggested by Campbell and Perron (1991), with the maximum lag length, $h_{\max}$, set to 12, while for the PP tests, the fixed truncation lag is set to 12. For both types of tests, since the distribution of the test statistics is non-standard, we compute critical values for the exact sample size ($T = 146$) using Monte-Carlo simulation with 10,000 replications under the null hypothesis of a random walk with iid normal innovations. It is observed that for most series, the null hypothesis of a random walk cannot be rejected at conventional significance levels. From the ADF tests without a time trend, the null hypothesis of a random walk can be rejected at the 10% level for Korea, the Philippines and Taiwan. With a time trend, the null can be rejected at the 10% level for Argentina and Indian and at the 5% for Malaysia. Results from the PP tests provide even weaker evidence against the random walk hypothesis. The null hypothesis can be rejected at the 10% only for Korea without a time trend and for Argentina with a time trend. Overall, these results tend to suggest that there is no significant evidence of mean reversion in emerging-market stock prices. One plausible reason for the non-rejection of the random walk hypothesis is the mis-specification of the deterministic components

Table 2
ADF and PP tests for random walk in emerging stock market prices: indexes denominated in the US dollar

Country	ADF tests		PP tests	
	No trend	With trend	No trend	With trend
Argentina	−1.930	−3.295*	−1.225	−3.363*
Brazil	−1.058	−3.075	−1.361	−2.581
Chile	−2.531	−1.215	−2.420	−0.961
Colombia	−1.048	−2.264	−1.045	−1.859
Greece	−1.828	−2.166	−1.686	−1.987
India	−1.299	−3.244*	−2.526	−3.150
Jordan	−1.250	−2.300	−1.584	−2.457
Korea	−2.818*	−1.227	−2.682*	−1.331
Malaysia	−0.311	−3.947**	−0.087	−3.391
Mexico	−2.398	−1.489	−1.928	−1.578
Nigeria	−1.379	−2.312	−1.302	−2.169
Pakistan	−1.277	−1.983	−1.318	−1.667
Philippines	−2.666*	−2.389	−2.528	−2.154
Taiwan	−2.647*	−1.701	−1.953	−2.752
Thailand	−1.903	−0.007	−2.022	−0.186
Venezuela	−1.418	−2.291	−1.474	−2.130
Zimbabwe	−1.375	−2.332	−1.861	−2.343

Notes:

1. The table reports ADF and PP tests for the random walk hypothesis for emerging market stock prices. The optimal lag length for the ADF tests is selected as suggested by Campbell and Perron (1991), with the maximum lag set to 12. For the PP tests, the truncation lag is set to 12.
2. The 1%, 5% and 10% critical values are −3.597, −2.943 and −2.522, respectively for the model without a time trend; and −4.097, −3.492 and −3.188 for the model with a time trend. They are computed using Monte-Carlo simulation with 10,000 replications.
3. * and ** denote statistical significance at the 10% and 5% levels, respectively.

included as regressors. It is likely that the series under investigation is characterized by a fundamental structural change. Failure to account for such a change may bias the test in favor of the null hypothesis of a random walk.

We now consider the case in which the stock index is assumed to contain a structural break with the break point determined endogenously. For each series, we estimate all three models (A)–(C) and compute the *t*-statistics for testing for $\alpha^i = 0$ ($i =$ A, B or C). We report the estimation results of the one specification, which shows the strongest evidence against the random walk hypothesis (i.e., the specification that gives the most significant test statistic on α^i). In determining the optimal lag, we follow the procedure suggested by Campbell and Perron (1991), by starting with the maximum lag length, $k_{\max}$, equal to 12. Table 3 contains the test results, with *t*-ratios exhibited in parentheses. We draw inference based on two sets of critical values for the *t*-statistic for $\alpha^i = 0$ ($i =$ A, B or C). The first set is the asymptotic critical values taken from Zivot and Andrews (1992), who obtain them through 5,000 Monte-Carlo replications. We acknowledge, however, that our sample size

Table 3
Test for random walk with a structural break in emerging stock market prices: indexes denominated in US dollar terms

Country	Model	T_B	K	θ	γ	α
Argentina	C	1991.5	6	0.909 (4.329)	−0.006 (−3.065)	−0.429** (−5.443)
Brazil	B	1992.11	12		0.003 (3.576)	−0.300** (−4.856)
Chile	C	1991.2	1	0.326 (3.719)	−0.004 (−3.673)	−0.157 (−4.229)
Colombia	C	1991.7	5	0.273 (3.895)	−0.002 (−2.725)	−0.109 (−4.507)
Greece	C	1990.1	11	0.584 (4.480)	−0.007 (−4.200)	−0.279** (−5.296)
India	C	1991.5	11	0.289 (3.563)	−0.002 (−2.362)	−0.327** (−5.080)
Jordan	C	1992.6	6	0.069 (1.449)	0.001 (0.233)	−0.195 (−3.721)
Korea	C	1987.9	4	0.150 (1.917)	−0.003 (−1.096)	−0.100 (−2.996)
Malaysia	C	1993.4	11	0.257 (2.352)	−0.001 (−1.337)	−0.494*** (−5.757)
Mexico	B	1994.9	7		−0.002 (−4.383)	−0.195** (−4.587)
Nigeria	A	1986.6	0	−0.319 (−4.276)		−0.229** (−4.862)
Pakistan	C	1991.7	5	0.271 (3.649)	−0.002 (−3.012)	−0.150 (−3.921)
Philippines	C	1989.9	12	0.204 (2.081)	−0.006 (−3.041)	−0.224** (−5.254)
Taiwan	B	1990.1	8		−0.005 (−4.690)	−0.136** (−4.848)
Thailand	C	1995.3	4	1.073 (2.243)	−0.009 (−2.399)	−0.086 (−2.109)
Venezuela	B	1993.6	10		−0.001 (−3.012)	−0.140* (−4.118)
Zimbabwe	B	1991.5	8		−0.002 (−4.246)	−0.120*** (−5.164)

Notes:
1. The table reports Zivot–Andrews test for the random walk hypothesis with structural breaks for seventeen emerging market stock prices. For each choice of breaking point, the optimum lag length, *k*, is selected as suggested by Campbell and Perron (1991).
2. Numbers inside parenthesis are *t*-ratios.
3. The 10%, 5% and 1% asymptotic critical values, obtained from Zivot and Andrews (1992), are respectively, as follows. Model A: −4.58, −4.80 and −5.34; Model B: −4.11, −4.42, and −4.93; Model C: −4.82, −5.08 and −5.57. The respective finite-sample critical values obtained from Monte-Carlo simulation with 10,000 replications are as follows. Model A: −4.448, −4.744 and 5.300; Model B: −4.455, −4.751 and 5.311; Model C: −4.808, −5.059 and −5.595.
4. *, ** and *** denote statistical significance at the 10%, 5% and 1% levels, respectively, based on the asymptotic critical values.

($T = 146$) may not be large enough for the asymptotic distribution to provide an adequate approximation to the exact finite-sample distribution. Therefore, we generate the second set of critical values using Monte-Carlo simulation with 10,000 replications under the null hypothesis of a random walk. Both sets of critical values are reported in Table 3 (note 3). Several observations can be drawn from Table 3.

Firstly, a broad examination of the results indicates that the coefficients on the break dummies, θ^i and γ^i ($i = \text{A, B or C}$), are in general significantly different from zero at the 5% level (based on critical values from the standard normal distribution). The exceptions are Jordan and Korea where neither dummy variable is significant, and Malaysia where the time trend dummy is not significant. Therefore, there is overwhelming evidence that these stock indexes may be subject to permanent shocks in sample due to structural or policy changes in the underlying economies.

Secondly, by incorporating the structural break into the trend function, we find that the null hypothesis can be rejected for a great number of countries. Using the asymptotic critical values, we can reject the null at the 1% significance level for Malaysia and Zimbabwe; at the 5% for Argentina, Brazil, Greece, India, Mexico, Nigeria, the Philippines and Taiwan; and at the 10% for Venezuela. When the finite-sample critical values from Monte-Carlo simulation are used, the results are somewhat weaker for three countries. Specifically, the null hypothesis can be rejected only at the 10% for Mexico and at the 5% for Zimbabwe, but cannot be rejected at the 10 level for Venezuela. For the remaining countries, the qualitative results remain the same. These results are in sharp contrast to those reported in Table 2, where no structural changes are taken into account in the tests. They suggest that for most of the emerging markets, stock indexes may be well characterized as mean reverting processes.

Thirdly, the break point, as identified, varies from country to country in general. Recall that these breaks are searched endogenously from the data and our procedure does not rely on pre-test information to determine them, thereby avoiding the possible problem of "data mining". While our main focus in this paper is on studying mean reversion of equity prices, it is still interesting to investigate whether those break points that are found to be significant at the 5% level or better (14 out of 17 cases) can be primarily accounted for by major policy changes or economic events in the associated countries.

We can clearly see that for a number of countries, the break points as identified in Table 3 are close to their corresponding market liberalization dates as documented in Table 1. This occurs to Argentina, Brazil, Colombia, Greece, Pakistan, Philippines and Taiwan. For these countries, the opening/liberalizing of financial markets fundamentally changed the market structures and could have caused some permanent shocks to equity prices. For several other countries, we find that major economic events or policy changes may provide plausible explanations for the breaks. [6] For

[6] Some of the major economic events for emerging markets are documented in Price (1994) and IFC's *Emerging Stock Market Factbook* (various issues).

example, the 1994–1995 peso crisis may be responsible for the structural changes in the Mexican stock market. In Thailand, the economy faced rapidly rising inflation in 1994. While the Thai government attempted to curb liquidity by tightening guidelines for domestic banks and imposing restrictions on foreign exchange holdings, these attempts were not very effective due to the large influx of foreign capital through offshore banking facilities. As a result, the Thai Ministry of Finance relaxed controls on foreign exchange transfers and allowed foreign banks to borrow and lend in the Thai currency. The rising interest rates and increased competition among banks in 1994 led to the big drop in its stock price in 1995. In Venezuela, the conflicts between the central bank and the government over the Roosen Plan, which aimed at stimulating economic growth by easing monetary policy, caused its currency to depreciate by 25% in 1993 and an additional 38% in 1994. This currency crisis and its consequences may be a major factor for the structural break in its stock market in 1993. A similar situation happened to Zimbabwe when its currency was devaluated by nearly a half in 1991 and this event may be mainly responsible for its stock market shock in the same year. In Chile, after the government changed from a military system to a civilian system in 1991, market liberalization took the proper shape, making the Chilean market freely accessible to foreigners. Furthermore, several major changes in its foreign exchange market were made so as to facilitate international capital flows. These might help explain the 1991 break in the Chilean stock market. As for India, the country faced a severe international payment crisis and the Indian Rupee was devalued by 30% in 1991. This, coupled with the new government's liberalization package known as the "New Economic Policy" had a major impact on its financial markets and could be responsible for the structural break in 1991.

The explanations provided above are of course very preliminary and only suggestive and should be interpreted with caution. Other explanations are also possible. For example, three of the six Latin American countries had their breaks in the year 1991 (Argentina, Chile, Colombia). This could have been caused by a global factor and may be anticipated. Admittedly, our test only identifies the possible break point but cannot distinguish a break that is anticipated from the one that is unanticipated. To check for this possibility of a global factor as an explanation, we run the Zivot–Andrews test to the Latin American index and find the possible break point to be in December 1990, although it is not significant at the 10% level. This break point is close to those for the three individual countries. This explanation does not go through for countries in other regions, however. For example, we apply the same test to the Asian index, and find the break point to be in February 1990 at the 5% significance level. However, all seven Asian countries in our sample seem to have quite different break points from each other and only Taiwan has the break point close to that in the Asian index.

In sum, we have identified significant structural changes in 14 emerging equity markets. These breaks are searched endogenously without using prior information. When they are incorporated in the tests for a random walk, we find that the null hypothesis of random walk in equity indexes can be rejected in favor of mean reversion in 11 markets. Furthermore, major economic events seem to provide reasonable explanations for these changes.

5. Robustness of results

The results on significant structural breaks and mean reversion in equity indexes reported in the preceding section are based on indexes denominated in the US dollar. In this section, we check for robustness of the results.

First, we consider the importance of exchange rate fluctuations in affecting these results. Abuaf and Jorion (1990), Engel and Hamilton (1990), Wu (1996) and others have shown that at low frequencies, exchange rates may be mean reverting. Engel and Hamilton (1990), Perron and Vogelsang (1992), and Wu (1997) have documented that exchange rates are subject to structural changes. It is possible that the results obtained in Section 4 are merely picking up the mean reversion and/or structural changes in exchange rates. To check for robustness of our findings, we carry out the Zivot–Andrews test for the indexes denominated in local currencies. The difference in comparing local-currency returns and dollar returns is of course due to exchange rate fluctuations. The test results are presented in Table 4, from which several

Table 4
Test for random walk in emerging stock market prices: indexes denominated in local currency terms

Country	Model	T_B	K	θ	γ	α
Argentina	C	1989.1	9	1.416 (6.468)	−0.014 (−4.143)	−0.114*** (−6.140)
Brazil	B	1993.11	0		−0.058 (−12.623)	−0.553*** (−12.56)
Chile	C	1990.11	3	0.264 (3.639)	−0.003 (−3.085)	−0.104 (−3.398)
Colombia	C	1991.8	1	0.401 (4.989)	−0.003 (−4.378)	−0.120** (−5.100)
Greece	C	1990.2	11	0.602 (4.757)	−0.007 (−4.346)	−0.321** (−5.517)
India	C	1991.5	11	0.324 (3.776)	−0.002 (−2.074)	−0.258* (−4.830)
Jordan	C	1988.8	3	−0.031 (−1.292)	0.002 (3.111)	−0.161 (−4.170)
Korea	C	1989.7	0	0.107 (1.645)	−0.003 (−2.348)	−0.106 (−2.793)
Malaysia	C	1993.8	11	0.379 (2.809)	−0.002 (−2.292)	−0.487** (−5.416)
Mexico	B	1986.3	2		0.013 (2.877)	−0.055 (−3.322)
Nigeria	A	1995.1	1	0.068 (4.104)		−0.066 (−4.058)
Pakistan	C	1991.4	5	0.155 (2.835)	−0.001 (−0.990)	−0.131 (−4.152)
Philippines	C	1989.10	12	0.228 (2.167)	−0.007 (−2.941)	−0.246** (−5.187)
Taiwan	B	1990.2	8		−0.005 (−4.910)	−0.163*** (−4.999)
Thailand	C	1995.4	0	1.120 (2.363)	−0.009 (−2.524)	−0.102 (−2.682)
Venezuela	B	1989.12	10		0.002 (3.231)	−0.111* (−4.408)
Zimbabwe	B	1991.8	8		−0.001 (−3.982)	−0.134*** (−5.234)

Notes:
1. The table reports Zivot–Andrews test for the random walk hypothesis with structural breaks for seventeen emerging market stock prices. For each choice of breaking point, the optimum lag length, k, is selected as suggested by Campbell and Perron (1991).
2. Numbers inside parenthesis are t-ratios.
3. The 10%, 5% and 1% asymptotic critical values, obtained from Zivot and Andrews (1992), are respectively, as follows. Model A: −4.58, −4.80 and −5.34; Model B: −4.11, −4.42, and −4.93; Model C: −4.82, −5.08 and −5.57. The respective finite-sample critical values obtained from Monte-Carlo simulation with 10,000 replications are as follows. Model A: −4.448, −4.744 and 5.300; Model B: −4.455, −4.751 and 5.311; Model C: −4.808, -5.059 and 5.595.
4. *, ** and *** denote statistical significance at the 10%, 5% and 1% levels, respectively, based on the asymptotic critical values.

observations can be drawn. Firstly, consistent with the results in Table 3, the coefficients on the break dummies, θ and γ, are in general significantly different from zero at the 5% level (based on critical values from the standard normal distribution). The only exceptions are Jordan and Korea where the intercept dummy is insignificant and Pakistan where the slope dummy is insignificant. Secondly, we find that the break points identified for local-currency indexes are identical or very close to those with dollar indexes for the majority of the countries, namely, Chile, Colombia, Greece, India, Malaysia, Pakistan, Philippines, Taiwan, Thailand and Zimbabwe. Thirdly, by incorporating the structural breaks into the tests, the null hypothesis of a random walk can be rejected for ten countries out of seventeen. Using the asymptotic critical values, we find that the null hypothesis can be rejected at the 1% significance level for Argentina, Brazil, Taiwan and Zimbabwe, at the 5% level for Colombia, Greece, Malaysia and Philippines, and at the 10% level for India and Venezuela. When the finite-sample critical values are employed, the results are only slightly weaker in that, for Taiwan and Zimbabwe, the null hypothesis is rejected at the 5% level but not at the 1% level. Compared to the results in Table 3, we find that Colombia becomes significant at the 5% level, while Mexico and Nigeria become insignificant.

Second, one may suspect that mean reversion in local currency price of an asset might reflect mean reversion in inflation rates, which tend to vary greatly in emerging markets in our sample period. Therefore, to further check for robustness of our mean reversion results, we carry out the tests using stock indexes expressed in real terms. To this end, for each country, we use its consumer price index to deflate its local currency stock price and test whether the resulting index (in real terms) are mean reverting. The consumer price indexes for all countries are obtained from International Monetary Fund's *International Financial Statistics*. Test results are reported in Table 5. We find that the break dummies are significant and the break points identified are in general consistent with those from Table 4. Furthermore, when the break points are incorporated, the null hypothesis of a random walk can be rejected in favor of mean reversion for 11 countries out of 17. Using asymptotic critical values, the null can be rejected at the 1% significance level for Argentina, Brazil, Taiwan and Zimbabwe; at the 5% level for Greece, Jordan, Malaysia, Nigeria, Philippines and Venezuela; and at the 10% level for Colombia. When finite-sample critical values are used, the results are only slightly weaker in that Nigeria becomes significant at the 10% level, and Taiwan at the 5% level.

In sum, the results obtained using indexes in local currencies and in real terms are qualitatively the same as those from dollar indexes. We therefore conclude that the findings in the preceding section are robust and are unlikely to be driven by structural changes and/or mean reversion in exchange rates or inflation rates.

Finally, following Poterba and Summers (1988) and Lo and MacKinlay (1988), we conduct a variance-ratio test for the random walk hypothesis. Table 6 reports variance-ratio statistics at the 24-, 36-, 48-, and 60-month horizons for the seventeen country index prices (in dollar terms), where the numbers inside parentheses are p-values, which are obtained from Monte-Carlo simulation with 5,000 replica-

Table 5
Test for random walk in emerging stock market prices: indexes denominated in real terms

Country	Model	T_B	K	θ	γ	α
Argentina	C	1989.1	11	0.409 (2.990)	−0.002 (−0.619)	-0.612^{***} (−6.261)
Brazil	B	1993.11	12		−0.092 (−19.618)	-0.958^{***} (−19.750)
Chile	C	1991.3	2	0.224 (2.963)	−0.002 (−2.835)	−0.120 (−3.324)
Colombia	C	1991.8	1	0.345 (4.563)	−0.003 (−3.842)	-0.113^{*} (−4.836)
Greece	C	1990.2	11	0.491 (4.176)	−0.006 (−3.761)	-0.299^{**} (−5.114)
India	C	1991.5	11	0.320 (3.704)	−0.002 (−2.216)	−0.252 (−4.694)
Jordan	A	1992.7	0	0.078 (4.364)		-0.214^{**} (−4.975)
Korea	C	1989.7	0	0.094 (1.475)	−0.003 (2.260)	−0.102 (−2.800)
Malaysia	C	1993.8	11	0.358 (2.634)	−0.002 (−2.317)	-0.430^{**} (−5.108)
Mexico	B	1994.10	1		−0.001 (−2.711)	−0.114 (−3.825)
Nigeria	B	1995.2	1		0.001 (3.997)	-0.108^{**} (−4.458)
Pakistan	C	1991.4	5	0.182 (3.201)	−0.001 (−1.645)	−0.130 (−4.140)
Philippines	C	1989.10	12	0.193 (1.904)	−0.006 (−2.883)	-0.235^{**} (−5.471)
Taiwan	B	1990.2	8		−0.005 (−4.910)	-0.163^{***} (−4.999)
Thailand	C	1995.4	0	1.070 (2.258)	−0.009 (−2.419)	−0.097 (−2.603)
Venezuela	B	1994.1	10		−0.001 (−3.028)	-0.126^{**} (−4.786)
Zimbabwe	A	1991.9	8	−0.197 (−5.118)		-0.142^{***} (−5.991)

Notes:
1. The table reports Zivot–Andrews test for the random walk hypothesis with structural breaks for seventeen emerging market stock prices. For each choice of breaking point, the optimum lag length, k, is selected as suggested by Campbell and Perron (1991).
2. Numbers inside parenthesis are t-ratios.
3. The 10%, 5% and 1% asymptotic critical values, obtained from Zivot and Andrews (1992), are respectively, as follows. Model A: −4.58, −4.80 and −5.34; Model B: −4.11, −4.42, and −4.93; Model C: −4.82, −5.08 and −5.57. The respective finite-sample critical values obtained from Monte-Carlo simulation with 10,000 replications are as follows. Model A: −4.448, −4.744 and 5.300; Model B: −4.455, −4.751 and 5.311; Model C: −4.808, −5.059 and 5.595.
4. *, ** and *** denote statistical significance at the 10%, 5% and 1% levels, respectively, based on the asymptotic critical values.

tions under the null hypothesis of serially independent returns. The variance ratio is equal to one under the null hypothesis of random walk and is less than one under the alternative of mean reversion. We find that while most point estimates of the variance ratios are below one, they are not statistically significant. In particular, at the 10% significance level, the null hypothesis of random walk can be rejected only for Argentina (at the 24-month horizon), Brazil (at the 24- and 36-month horizons), India (at the 24-, 36-, and 60-month horizons), and Malaysia (at all horizons). Overall, the evidence against the random walk hypothesis from the variance-ratio test is weaker than that from the Zivot–Andrews test. We note, however, that the variance-ratio test, which is based on a weighted average of serial correlations, focuses more on short to medium run behavior of stock returns, while the Zivot–Andrews test focuses more on long-run mean reversion. Furthermore, the variance-ratio test does not take into account a structural break and might have lower power. These may explain the different results obtained from the two tests.

Table 6
Variance-ratio test for random walk in emerging stock market prices: indexes denominated in US dollar terms

	Return measurement interval			
	24 Months	36 Months	48 Months	60 Months
Argentina	0.604 (0.065)	0.636 (0.223)	0.632 (0.301)	0.562 (0.287)
Brazil	0.647 (0.093)	0.381 (0.034)	0.546 (0.226)	0.553 (0.295)
Chile	1.009 (0.578)	0.737 (0.332)	0.905 (0.545)	0.941 (0.580)
Colombia	0.905 (0.414)	0.830 (0.418)	0.683 (0.359)	0.458 (0.196)
Greece	0.936 (0.448)	0.707 (0.290)	0.923 (0.538)	1.123 (0.663)
India	0.519 (0.025)	0.432 (0.061)	0.446 (0.137)	0.317 (0.075)
Jordan	1.061 (0.659)	1.221 (0.763)	1.130 (0.693)	1.035 (0.637)
Korea	1.252 (0.856)	1.486 (0.884)	1.510 (0.852)	1.124 (0.666)
Malaysia	0.496 (0.020)	0.442 (0.068)	0.326 (0.044)	0.331 (0.086)
Mexico	0.910 (0.419)	0.804 (0.403)	0.815 (0.478)	0.845 (0.509)
Nigeria	0.811 (0.246)	0.756 (0.326)	0.633 (0.289)	0.433 (0.166)
Pakistan	0.918 (0.441)	0.910 (0.515)	0.930 (0.565)	0.817 (0.494)
Philippines	1.102 (0.685)	1.001 (0.575)	1.069 (0.641)	0.800 (0.474)
Taiwan	1.374 (0.937)	1.651 (0.931)	1.638 (0.877)	1.334 (0.763)
Thailand	1.118 (0.723)	1.120 (0.683)	1.018 (0.614)	0.955 (0.583)
Venezuela	1.078 (0.669)	0.881 (0.476)	0.840 (0.494)	0.696 (0.408)
Zimbabwe	1.210 (0.824)	1.165 (0.723)	1.294 (0.767)	1.451 (0.808)

Notes:
1. The table reports results from the variance-ratio test for seventeen emerging market stock prices. The variance ratio is equal to one under the null hypothesis of random walk and is below one under the alternative hypothesis of mean reversion.
2. Numbers inside parentheses are *p*-values based on 5,000 replications under the null hypothesis of serially independent returns.

6. Conclusion

We believe that our paper contributes to the literature by applying a methodological innovation as well as through our findings of mean reversion in the context of emerging stock markets.

It is well known that standard tests, such as Dickey and Fuller (1979, 1981) and Phillips and Perron (1988), for the random walk hypothesis in stock prices have low power against the alternative hypothesis of mean reversion in small samples. The problem is especially serious when there exist structural changes in the underlying series. Failure to account for the breaks can produce misleading tests and result in incorrect inference. In this paper, we implement the Zivot–Andrews sequential test for a random walk that explicitly takes into account the effects of structural changes in stock prices. This test considerably improves the power over the ADF and PP tests in a given sample size.

Applying the Zivot–Andrews test to stock prices of seventeen emerging markets, we reach several interesting conclusions. The gain in test power allows us to reject the random walk hypothesis at the 1% or 5% significance level in ten markets. This is a useful result as it adds to the controversial evidence of mean reversion first provided for US stock prices by Fama and French (1988) and Poterba and

Summers (1988). The uncovering of a strong relation in an entirely different data set decreases the likelihood of earlier mean reversion findings as attributable to ‘‘data mining’’. Furthermore, we find that for fourteen countries, the stock markets exhibit significant structural breaks, which appear either in the intercept, in the time trend, or in both. These structural changes, as identified by the test, are in general consistent with their corresponding market liberalization dates and/or can be explained by other major economic events in the underlying economies.

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