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# Efficiency Market Hypothesis in an Emerging Market: Does It Really Hold for Malaysia?

## *(Kecekapan Pasaran Hipotesis dalam Pasaran Baru Muncul: Adakah Ia Benar-benar Sesuai untuk Malaysia?)*

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### *ABSTRACT*

*This study revisits the efficient market hypothesis (EMH) with regard to the Kuala Lumpur Stock Exchange (KLSE) at the sectoral level. Based on Liu and Narayan's (2011) GARCH-based unit-root with structural breaks test, the unit-root null is rejected for all except one sector. By contrast, models based on commonly used unit-root tests that ignore heteroskedastic and/or breaks tend to favour the EMH. We find that the half-life estimates based on the local-persistent model are short, with the majority of them taking less than six months to absorb half a shock. All in all, the indices examined are largely inconsistent with weak-form efficiency, which implies that the returns on equity portfolios are indeed predictable.*

*Keywords: Stock prices; unit-root; half-life; structural breaks*

### *ABSTRAK*

*Kajian ini mengkaji semula hipotesis pasaran cekap (EMH) berkaitan dengan Bursa Saham Kuala Lumpur (BSKL) di peringkat sektor. Berdasarkan ujian kepegunan Liu dan Narayan's (2011) berasaskan GARCH dengan selaan struktur, nol kepegunan ditolak untuk semua kecuali satu sektor. Sebaliknya, model berdasarkan ujian kepegunan yang biasa digunakan yang mengabaikan heteroskedastik dan/atau selaan struktur cenderung untuk memihak kepada EMH. Kita dapati bahawa anggaran separuh hayat berdasarkan model "local-persistent" adalah pendek, dengan majoriti daripada mereka yang mengambil kurang daripada enam bulan untuk menyerap separuh kejutan. Kesimpulannya, indeks yang dikaji sebahagian besarnya tidak konsisten dengan kecekapan lemah, yang dapat membayangkan bahawa pulangan ke atas portfolio ekuiti memang boleh diramalkan.*

*Kata kunci: Harga saham; kepegunan; separuh hayat; selaan struktur*

## INTRODUCTION

The efficient market hypothesis (EMH) – the idea that competitive financial markets exploit all available information when setting security prices, has been extensively investigated. The theory maintains that financial information is disseminated efficiently and, consequently, asset prices are unpredictable. One group of researchers maintains that movements in the stock market are driven mainly by a set of micro- and macroeconomic variables (money supply, exchange rates, inflation, etc.); see Campbell and Shiller (2001), Cochrane (2005), Chen (2009), Md Nor et al. (2010) and Gupta and Modise (2013), among others. Despite the enormous efforts made in the past to develop modeling techniques (including nonlinear models), there is a growing consensus among researchers that stock prices are best characterised by a random walk (or unit-root) process, implying that changes in stock prices are basically unpredictable based on past movements in stock prices (see, inter alia, Worthington & Higgs 2004; Lim & Brooks 2011; Narayan & Smyth 2006). The findings from these studies cast serious doubt on various methods used to predict stock prices. By contrast, randomness in asset price is consistent with the weak-form EMH – abnormal returns are unattainable in competitive markets due to instantaneous market responses to the arrival of new information.[1]

Among the emerging equity markets, the Kuala Lumpur Stock Exchange (KLSE) has gone through a series of financial liberalisations (since the 1980s), the Asian financial crisis (AFC 1997-98), the imposition of capital controls (1998-2005), and a global financial crisis (GFC 2007-2009). Investor's confidence is severely affected during the financial turmoil and consequently disrupts the performance of the stock market. A number of studies have already been undertaken to investigate the EMH for this emerging market which reported conflicting results. Lim (2008), for example, found that all sectors, but one (tin and mining) sector exhibits inefficiency during the 1997 AFC period, but the crisis-stricken sectors improved significantly as the economy recovered from the financial turmoil. Lim's work highlights an important point often ignored in earlier studies: statistical tests which impose too restrictive assumptions on the behavior of the financial series may be problematic. Under those assumptions, the standard unit-root tests have very low power, and hence are most likely to produce biased results. Munir et al. (2012), who apply the two-regime threshold autoregressive model, find that stock prices in Malaysia (but not Singapore) are characterised by a non-stationary process, which is supportive of EMH. Their results point to stronger evidence of the fulfillment of EMH, once when nonlinearities are taken into account.

A nonlinear data generating process (DGP) in stock prices may be explained by factors such as market frictions, institutional constraints and transaction costs. An important finding that emerged from both studies is that the speed of adjustment to the fundamental value may not be constant over time. Lim (2008) using a rolling Hinich bicorrelation test statistic shows that the highest inefficiency occurs during the AFC period for all but one economic sector. In our view, this is to be expected as the speed at which stocks revert to fundamental value is faster during periods of economic uncertainty. Additionally, during the last two decades, there have been several global financial event shocks that cause structural breaks in the behavior of stock returns. The speed of mean reversion (as measured by half-life) is significantly higher during large fall in the market which would imply the market is predictable.[2]

The cumulative literature has found evidence to show that emerging equity markets are less predictable than developed ones (Al-Khazali, Ding & Pyun 2007) and that the predictability of stock market returns has declined somewhat in recent years (Shamsuddin & Kim 2010). The changing characteristic (structure) of the KLSE is also highlighted in Cheong, Mohd Nor and Isa (2007). Lee et al. (2010), who applied recently developed panel unit-root tests for a large set of (32 developed and 26 developing) equity markets, buck the trend. Unlike those articles mentioned earlier, they found that real stock prices (including the developed markets) can be characterised by a stationary process, which is inconsistent with EMH. Mean-reversion is found after allowing for structural breaks in the monthly (high frequency) price data. The impact of shocks is not long lasting and profitable arbitrage opportunities among stock markets exits, even in the developed markets. Two points are worth highlighting from their findings. First, the failure of traditional time-series tests (e.g., Augmented Dickey-Fuller (ADF)) to detect reverting behaviour could be due to the lack of statistical power to reject the unit-root null. The problem is often magnified with limited availability of long-span data; see also Poterba and Summers (1988). Second, ignoring structural breaks associated with extreme events (notably AFC, internet bubble in 2000; GFC in 2007) are known to be biased towards rejecting the unit-root null for univariate statistics (see Perron 1989). In the context of East Asian countries, the issue has been addressed in Lean and Smyth (2007), Narayan and Smyth (2006) and Baharumshah, Soon and Hamzah (2013) to account for the large fluctuations that shift the mean of the individual series.

The primary aim of this study is to conduct an empirical analysis on the EMH for the KLSE using data that cover the recent dotcom crisis, GFC and the debt crisis in Europe. We examine the EMH at the sectoral level to address the concerns of bias in the persistence estimates due to aggregation across sectors of the economy (mentioned in Imbs et al. 2005). In doing so, it allows us to identify which sector(s) drives the market to be inefficient and determine whether stock prices revert to their fundamental values after a shock. We extend the data sets beyond the 1997 AFC in part to demonstrate the comparability of our results with those from previous studies (e.g., Lim 2008; Lean & Smyth 2007). The motivation for extending the work lies in a timely question on the problems associated with recent global economic uncertainties being faced by Malaysia. Unlike the AFC, the GFC in 2008 happened in G3 (US, EU and Japan). Many scholars and policymakers view that the impact of the recent GFC would have long lasting consequence of slowdown in external sectors. From another perspective, an assessment at the disaggregated level, as done in the present paper, allows traders to formulate better trading strategies based on more specific information.

We consider Malaysia, an emerging market economy, because economic and financial crisis in this country tends to be more frequent and its impact tends to be larger than that in the major industrialised countries. Therefore, a question that we explore in this paper is whether the nature of the crisis matters for testing market efficiency. Kim and Shamsuddin (2008), Lean and Smyth (2007) and Lim (2008), among others, have investigated the impact of the AFC on selected Asian stock markets. Although the impact of the currency crisis on stock returns is evident, support for the EMH is at best mixed and efficiency appears to vary by both country and time period. By contrast, Hoque, Kim and Pyun (2007) fail to show any significant impact of the AFC on equity markets in the region. According to the authors, all the Asian markets (Hong Kong, Indonesia, Malaysia, the Philippines, Singapore, and Thailand) but South Korea remain inefficient over the sampling period examined. Other studies have found the Malaysian equity market exhibits weak-form efficiency include Ahmad and Hussain (2001), Barnes (1986) and Laurence (1986).

The present paper complements the existing literature in terms of the estimation strategy. To test for the random walk hypothesis, unit-root tests are widely applied in the earlier literature (Lean & Smyth 2007; Munir et al. 2012; Narayan & Smyth 2006; and others).[3] Often, researchers neglected the detail that financial time series are usually characterised by some stylised facts, heteroskedatic and structural breaks volatility clustering (Panagiotidis 2010; Rahman & Saadi 2008), such that standard unit-root tests may not be well equipped to validate the hypothesis. In this study, our focus is on a new methodology that overcomes the size distortion bias arising from heteroskedasticity, which has been largely ignored in the past. The standard Dickey-Fuller test tends to be oversized in the presence of strong generalised autoregressive conditional heteroskedasticity (GARCH) effect; see also Kim and Schmidt (1993), Hamori and Tokihisa (1997) and more recent paper by Su (2011). As mentioned in Hamori and Tokihisa (1997), a simple shift in innovation variance partway through a sample invalidate the usual unit-root asymptotic. For a full account of the volatility in KLSE using a GARCH-family model; see Cheong et al. (2007). Thus, if heteroskedasticity is not properly accounted for, spurious inference about the presence of unit-root may result since the standard unit-root tests have low statistical power in rejecting a false null hypothesis. Besides, the half-life estimates, generate from such models is biased and may lead to non-optimal investment strategies.

While the past literature focus on whether stock prices is an *I*(1) or *I*(1) variable to determine the EMH, we for the first time allow for in-between (locally persistence) process to sidestep that shortfall of earlier approaches of measuring the degree of persistence. Local persistence (introduced by Phillips, Moon & Xiao 2001) implies that the series is more persistent than the stationary case, but less persistent than the unit-root case. It also implies that the stock price is non-stationary but a mean reversion occurs in the long-run and the reject the fulfillment of EMH. Through this "new lens" we are able to reach the same conclusion as in Lim (2008) and others, but our results are more informative than theirs as we provided the speed of convergence to fundamental value which is important for trading strategies (portfolio managers) and regulators. In fact, our finding supports the idea that completely efficient markets were not a realistic assumption even theoretically. Nobel Laureate Paul Krugman (2009) in his article on "How did economists get it so wrong?' has this to say about market efficiency: "the belief in efficient financial markets blinded many if not most economists to the emergence of the biggest financial bubble in history [. . .]. In the same article, Krugman went on to say "Economics, as a field, got in trouble because economists were seduced by a vision of a perfect, frictionless market economy." It should be mentioned here that Krugman made no distinction between the developed and developing markets.

Our paper is distinguished from earlier studies by examining the degree of mean-reversion hypothesis for stock price as we will describe later. Unit-root tests are designed to ascertain whether a series is *I*(0) or *I*(1), and the *I*(0)/*I*(1) distinction implicitly restricts the type of dynamic process allowed in the DGP (Caporale & Gil-Alana 2002).[4] It rules out an intermediate or in-between process and hence raised considerable doubt regarding earlier findings. A stationary process might be consistent with the EMH if shocks to the series are slow to die out (i.e., lengthy half-lives). The bipolar characterisation does not provide the speed of adjustment to an equilibrium value. This means that the application of more sophisticated method seems necessary. In this study, we present for the first time the half-life estimates along with their computed confidence intervals (CIs) based on Phillips et al.'s (2001) local-persistent model. As in the case of unit-root tests, the computation of half-life is likely to be affected by structural breaks (occasionally permanent shocks). As elaborated in Baharumshah et al. (2013) and Basher and Carrion-i-Silvestre (2013), ignoring structural breaks lead over-estimated prediction of persistence (longer half-life). The findings in this paper show that when breaks are not considered, it is almost impossible to reject the existence of unit-root (EMH) in the series. To isolate these shocks, we rely on new structural breaks approaches proposed by Narayan and Popp (2010) and Liu and Narayan (2011) to deal with a high degree of persistence in stock prices, and measuring the speed of adjustment.[5] The latter is to address the problem of innovations that exhibit unstable volatility, especially during periods of economic uncertainty. If unstable volatility is not properly taken into account, spurious inference about the presence or absence of a unit-root may result.

The remainder of this paper is organised as follows: Section 2 provides a brief overview of the estimation strategy. Section 3 describes the data and then presents the empirical results. Section 4 discusses the empirical findings, including the calculated speed of mean-reversion to fundamental value. In section 5, we draw conclusions.

## GARCH (1,1)-UNIT-ROOT TEST WITH TWO ENDOGENOUS STRUCTURAL BREAKS

Traditionally, EMH has been tested by examining whether there is a unit-root in stock prices. We follow the literature to assess whether Malaysian stock prices can be characterised as following a random walk or mean-reverting process. If stock prices follow random walks, weak-form EMH is supported. Otherwise, if stock prices are mean-reverting, the price level will return to its trend over time, making it possible to forecast future movements in stock prices based on past behavior. Thus, there is a potential for investors to gain an advantage through utilising predictability in stock returns. A convenient and common test of stock price properties is via the unit-root test, namely, the ADF and Phillips and Perron (PP) tests. These tests suffer from both low power and often severe size distortions, especially in small samples.

The existing literature on EMH based on the conventional unit-root tests has been criticised for not accounting for structural breaks. Accounting for structural breaks can substantially reduce the persistence within the regimes defined by those breaks (Perron 1989) or even produce spurious evidence of fractional integration. This is of special interest in our study, due to the fact that our sampling period includes the AFC, the GFC and the euro debt crisis that follows after that. Lee and Strazicich (2003) and Narayan and Popp (2010) who recognised ignoring breaks and inaccurate estimation of break dates as important sources of spurious rejections and hence developed testing procedures to account for two breaks. Narayan and Popp (2010) demonstrate that their test is more powerful and precise than the Lee and Strazicich

(2003) one in selecting the endogenous break dates. For this reason, we apply the Narayan and Popp (2010) test of endogenous breaks, noting that the test may be inefficient because it assumes no GARCH effect. Nonetheless, financial time series are potentially characterised by GARCH effect. With this purpose in mind, we further our analysis on the unit-root tests by performing newly improved GARCH-based unit-root test with two structural breaks, which was proposed by Liu and Narayan (2011) to test the EMH for the KLSE.

The key features of the Liu and Narayan's (2011) model are first, they account for GARCH errors and two structural breaks simultaneously. This means that the break test is based on non-identical and independently distributed (iid) errors. According to Ling and Li (2003) and Ling, Li and McAleer (2003), if the error terms in the model followed a GARCH process, the estimation and testing for unit-root involve intrinsic problems. The estimators are no longer efficient if heteroskedasticity in errors is not tackled properly. Second, the critical values (CVs) do not vary much with the GARCH parameters across different break fractions in the case of unknown break dates; however, the CVs in the case of known break dates converge to the traditional Dickey-Fuller distribution as sample size increases. For more details, see Ling et al. (2003) and others.

The GARCH-based unit-root with two structural breaks model has the following form:

$$P_t = \rho P_{t-1} + d_1 B_{1,t} + d_2 B_{2,t} + v_t, \tag{1}$$

where $P_t$ refers to stock price for $t = 1,\ldots,\mathrm{T}$, $B_{i,t} = 1$ for $t \geq T_{B_i}$ otherwise $B_{i,t} = 0$, where $T_{B_i}$ are the structural break points with $i = 1,2$. Here, we assume that $v_t$ follows the first-order GARCH model and is presented as $v_t = \sqrt{h_t \mu_t}$ where, $h_t = \vartheta + \delta v_{t-1}^2 + \varphi h_{t-1}$, $\mu_t \rightarrow N(0,1)$, $\vartheta > 0$, $\delta \geq 0$ and $\varphi \geq 0$.

Liu and Narayan (2011) adopt a joint maximum likelihood estimation approach proposed by Seo (1999) for unit-root equations with a GARCH error process. Accordingly, the maximum likelihood $t$-statistics for $\rho$ are used to test the unit-root $\mathrm{H}_0$: $\rho = 1$ against the alternative $\mathrm{H}_1$: $\rho < 1$ . A heteroskedasticity-consistent covariance matrix as stated by the White (1980) is utilized for the unit-root hypothesis test estimation.

Since the true break dates are unknown, Equation (1) has to be substituted with their estimates $\hat{T}_{B,i}$ $i = 1, 2$, in order to conduct the unit-root test. The authors claim that the sequential procedure is preferable to a simultaneous procedure for selecting the break date, as the former procedure is far less computationally demanding. Following the sequential procedure, a single break is selected according to the maximum absolute $t$-value of the break dummy coefficient $(d_1)$, $\hat{T}_{B,1} = \arg\max_{T_{B,1}} \left| t_{\hat{d}_1}(T_{B,1}) \right|$. Accordingly, when the first break is selected, it is then imposed in the test regression in order to estimate the second break date, $\hat{T}_{B,2}$. $\hat{T}_{B,2}$ is estimated with the maximum absolute $t$-value of the break dummy coefficient $(d_2)$, $\hat{T}_{B,2} = \arg\max_{T_{B,2}} \left| t_{\hat{d}_2}(T_{B,2}) \right|$ . The 5% test CVs for different break fractions and sample sizes ranging from 150 to 500 are tabulated in Liu and Narayan (2011: Table 4).

## EMPIRICAL RESULTS

The monthly frequency equity price indices from 1980 to 2011 are extracted from the FTSE Bursa Malaysia.[6] The consumer price index (CPI, 2005 = 100) source is the International Monetary Fund, *International Financial Statistics*. The real equity price index is constructed by deflating the nominal equity price with the CPI. The start of the sampling period was dictated primarily by the availability of data. We extend Lim's (2008) work by including indices from two additional sectors (technology and plantation sectors) and by considering data beyond 2006 – to account for the impact of the recent GFC and the euro debt crisis in the equity market. The updated data allows us to investigate whether the extreme events in the late 2000s have affected the pricing behavior in the equity market. The ten major sectors studied include finance, industries, plantations, properties, tin and mining, construction, industrial products, consumer products, trade and services, and technology. The extreme events that originate from the US (S&P 500) may have different effects on the various sectors of the KLSE. As shown in Figure 1, the equity prices were severely affected by the AFC and GFC. The extent of these shocks on the individual series varies according to the economic sectors listed in Figure 1. Table A1 provides the descriptive statistics for two sub-sample periods (1980: M1 to 1996: M12 and 1999: M1 to 2011: M7). A quick glance at the statistics reveals that the variance of several of the sectors (e.g., plantations, construction, industrial products, consumer products, and trade and services) has increased significantly after the AFC.

As mentioned earlier, the confirmation of EMH is thus far conflicting and it depends on the methodology (or model specification), country characteristics, and predictive power of the stock indices.[7] The predictability of stock returns can be determined from previous changes in prices by using unit-root tests (Narayan 2005). Rahman and Saadi (2008: 210), however, argue that testing for the difference- (trend-) stationarity in stock indices is not sufficient for the random walk hypothesis, but whether it is sufficient or not also depends on the predictability of the indices. A growing numbers of studies concentrate on the DGP of the series, especially for emerging markets that experience instability in their economic environment.

For comparison, we first apply the conventional Ng and Perron (2001) and Said and Dickey (1984, ADF) unit-root tests to all the indices examined (see Table A2). In all but two sectors (industrial and tin and mining sectors), the ADF and Ng and Perron (2001) tests cannot be rejected, i.e they confirm the unit-root null at the usual significance levels. They seem to indicate that most stock price indices are non-stationary. It is widely acknowledged that these

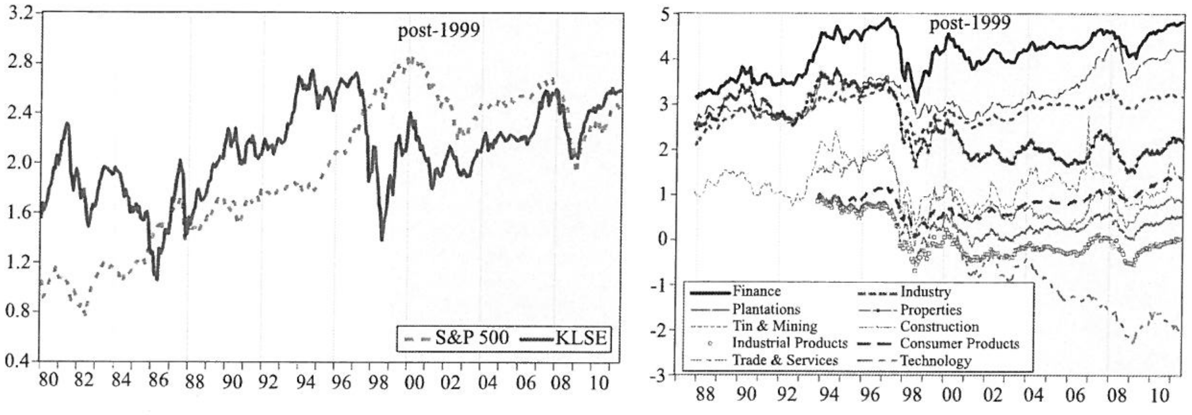

FIGURE 1. Real equity price index

unit-root tests have low power when applied to finite samples. Further, our results might be restrictive since we have ignored the possibility of structural breaks in the DGP. If structural breaks are present in DGP but ignored, the results tend to favor the unit-root null; see the seminal paper by Perron (1989). To address this concern, a unit-root test proposed by Narayan and Popp (2010) is applied to all the series. As shown in Table 1, four sectors (finance, properties/real estate, tin and mining, and construction) and the FTSE KLCI are stationary at 10% significance levels or better (see Model 1). Two other sectors (trade and services and technology) are added to the list when a trend is added to the model (Model 2). When two structural breaks in the DGP are taken into account, the results show that only four sectors are consistent with the EMH. Therefore, limited evidence of the EMH is reported with the Narayan and Popp (2010) two-break test. A potential problem with the approaches applied so far is that they are based on models that assume iid innovations, but financial data are likely to violate this assumption. More to the point, inferences drawn from the above unit-root tests may lead to misleading conclusions when heteroskedastic errors are present in the series (Ling et al. 2003; Liu & Narayan 2011).

TABLE 1. Narayan and Popp (2010) unit-root with breaks test

| | Model 1 | | | | Model 2 | | | |
|---|---|---|---|---|---|---|---|---|
| | TB1 | TB2 | $\hat{\rho}\,[t_{\hat{\rho}}]$ | k | TB1 | TB2 | $\hat{\rho}\,[t_{\hat{\rho}}]$ | k |
| FTSE KLCI | 87:9$^{a}$ | 97:10$^{a}$ | -0.0808$^{a}$<br>[-4.790] | 7 | 97:10$^{a}$ | 98:08$^{a}$ | -0.0782$^{b}$<br>[-4.898] | 1 |
| 10 sectors | | | | | | | | |
| Finance | 97:10$^{a}$ | 98:01$^{a}$ | -0.0763$^{c}$<br>[-3.904] | 11 | 97:10$^{a}$ | 98:02$^{a}$ | -0.1496$^{a}$<br>[-5.476] | 9 |
| Industry | 97:10$^{a}$ | 98:07$^{a}$ | -0.1359<br>[-3.631] | 15 | 97:10$^{a}$ | 98:08$^{a}$ | -0.1714<br>[-3.986] | 15 |
| Plantations | 93:11$^{a}$ | 98:07$^{a}$ | -0.0393<br>[-2.218] | 1 | 93:11$^{a}$ | 98:07$^{a}$ | -0.0558<br>[-2.452] | 1 |
| Properties | 97:10$^{a}$ | 98:01$^{a}$ | -0.0912$^{c}$<br>[-4.076] | 13 | 97:10$^{a}$ | 98:08$^{a}$ | -0.1132<br>[-3.619] | 13 |
| Tin & Mining | 97:10$^{a}$ | 98:05$^{a}$ | -0.1160$^{c}$<br>[-4.002] | 0 | 97:10$^{a}$ | 98:10$^{b}$ | -0.1264<br>[-4.121] | 0 |
| Construction | 97:10$^{a}$ | 98:01$^{a}$ | -0.1819$^{b}$<br>[-4.541] | 13 | 97:10$^{a}$ | 98:01$^{a}$ | -0.1692<br>[-4.315] | 13 |
| Industrial Products | 97:10$^{a}$ | 98:07$^{a}$ | -0.1183<br>[-2.429] | 13 | 97:10$^{a}$ | 98:08$^{a}$ | -0.1781<br>[-4.565] | 9 |
| Consumer Products | 97:10$^{a}$ | 98:07$^{a}$ | -0.1699<br>[-3.919] | 7 | 98:01$^{a}$ | 98:07$^{a}$ | -0.1140<br>[-1.977] | 13 |
| Trade & Services | 97:10$^{a}$ | 98:07$^{a}$ | -0.1247<br>[-3.352] | 7 | 97:10$^{a}$ | 99:03$^{a}$ | -0.2327$^{c}$<br>[-4.927] | 12 |
| Technology | 03:5$^{b}$ | 08:05$^{a}$ | -0.1369<br>[-3.905] | 5 | 03:05$^{b}$ | 08:05$^{b}$ | -0.2401$^{b}$<br>[-4.951] | 5 |

*Notes*: ($^{a}$), ($^{b}$) and ($^{c}$) indicate significance at 1%, 5%, and 10% levels, respectively. TB1 and TB2 are the dates of structural breaks detected. $t_{\hat{\rho}}$ is the test statistic of $\hat{\rho}$ and length lag (k) selected based on the general-to-specific procedure proposed by Hall (1994). Critical values are tabulated in Narayan and Popp (2010).

To take the analysis further, we apply a newly developed GARCH-type unit-root model proposed by Liu and Narayan (2011) that simultaneously accounts for structural breaks and heteroskedastic innovations. Results based on the new test that outperforms its rival statistics are reported in Table 2. The GARCH (1, 1)-two-break unit-root test rejects the unit-root null in all cases except one – the technology sector. We note that the technology sector is relatively new in Malaysia, the sampling period used to examine the half-life property of the local persistence process is much shorter than other sectors. In essence, our findings do not favour the EMH for all 10 sectors (including the FTSE KLCI index). Liu and Narayan (2011), who used a model related to the one explored in this paper, find that 22% of US firm stock indices temporarily overreact by moving away from their fundamental values in response to the information (dubbed as the stock market overreaction hypothesis) based on this new test.

TABLE 2. Liu and Narayan (2011) GARCH (1, 1)-unit-root with breaks test

| | Periods | TB1 | TB2 | $t_{\hat{\rho}}$ |
|---|---|---|---|---|
| FTSE KLCI | 1980M1-2011M7 | 82:06 | 98:08 | -7.039$^a$ |
| 10 sectors | | | | |
| Finance | 1987M11-2011M7 | 98:03 | 03:03 | -4.759$^a$ |
| Industry | 1987M11-2011M7 | 91:08 | 09:08 | -4.555$^a$ |
| Plantations | 1987M11-2011M7 | 01:10 | 05:08 | -6.326$^a$ |
| Properties | 1987M11-2011M7 | 98:09 | 02:04 | -6.573$^a$ |
| Tin & Mining | 1987M11-2011M7 | 91:05 | 91:10 | -5.522$^a$ |
| Construction | 1993M11-2011M7 | 96:08 | 99:10 | -6.810$^a$ |
| Industrial Products | 1993M11-2011M7 | 96:04 | 08:09 | -5.971$^a$ |
| Consumer Products | 1993M11-2011M7 | 96:12 | 09:03 | -8.062$^a$ |
| Trade & Services | 1993M11-2011M7 | 01:12 | 10:04 | -5.930$^a$ |
| Technology | 2000M5-2011M7 | 01:12 | 02:03 | -3.145 |

*Notes*: ($^a$) indicates significance at the 5% level. TB1 and TB2 are the dates of structural breaks detected. $t_{\hat{\rho}}$ is the test statistic of $\hat{\rho}$. Critical values are tabulated in Table 4 of Liu and Narayan (2011)

We now turn to the location of the break dates that were endogenously determined by the Liu and Narayan (2011) test. Evidently, the devaluation of the Thai baht in the mid-1990s triggered a massive reversal of capital flows in East Asia, including Malaysia, and this has affected the stock prices in KLSE. This is closely related to our study, as the KLSE has been affected by a slowdown in the region and it was followed by the imposition of capital controls in 1998 (the KLSE index fell from 1,300 points to as low as 400 points) and the pegging of the Malaysian ringgit to the US dollar. The majority of the dates detected by Narayan and Popp (2010) sequential test coincide with the 1997-98 AFC; except for the technology sector (see Table 1). Note that the sample period for this sector is available only after 2000. The location of the breaks in the technology sector is around mid-2003 and 2008–closely linked to the recession in early 2000s and subprime crisis of the late 2000s, respectively. According to the Liu and Narayan (2011) test, the recent global financial shocks affect sectors such as industry, consumer products and industrial products (see Table 2). In the plantation sector, the estimated break dates are consistent with the economic recession at the beginning of the 2000s. Results reported in Table 2 indicate that break dates occur in the early 1990s for the industry sector. The break corresponds to developments that are specific to that particular sector following major structural reforms in that industry. It also coincides with the period of which inflation, along with the other macroeconomic variables become more stable. Finally, it should be mentioned that not all the sectors are affected by the GFC. The broader KLSE is adversely affected by the high uncertainty associated with the AFC but not the GFC. Again, this observation provides a rationale to investigate the stochastic properties stock prices at the sectoral level.

Sekioua (2008) noted that rejecting the unit-root null does not mean that the alternative automatically holds. Since unit-root tests do not provide the degree of persistence in stock prices, some studies have looked at the speed of convergence or the degree of mean-reversion in stock price series to examine the EMH (e.g., Chaudhuri & Wu 2003; Niarchos & Alexakis 1998; Spierdijk, Bikker & Van den Heok 2012). To complement the results presented earlier and cast the empirical net wider, we proceed with the speed of adjustment (half-life) estimation based on the local-persistent model developed by Phillips et al. (2001) to capture a much wider class of mean-reverting behavior; see Kim and Lima (2010) for application of the model of real exchange rates. Recently, Baharumshah et al. (2013) used this model to confirm real interest rate parity holds in Asian countries.[8] Briefly, the model proposed by Phillips et al. (2001) has two advantages over the traditional approach based on ADF regression. First, it allows for local persistence process. Second, it accounts for the short-run memory dynamics in the residuals of the ADF regression; see Kim and Lima (2010) for the proof. Table 3 reports the degree of local persistence of the stochastic process, the corresponding half-life estimates persistence, and the associated 95% CIs for stock market indices. Most studies rely on point estimates of the half-life, neglecting CIs for proper statistical inference. As mentioned in Rossi (2005), wide CIs provide little information regarding the speed of convergence.

As shown in Table 3, the point estimate of half-lives from the persistence model (Model 1) ranges from 3.81 months (construction sector) to as high as 17.64 months (plantations sector). Likewise, differing degrees of persistence are observed in the Model 2: The point estimate of half-lives ranges from 2.89 months (technology sector) to 12.42 months (plantations sector). Focusing on the Model 2, it takes approximately 2.89-12.42 months for these prices to revert to their fundamental values, following a one-time shock to stock price. At the broader index level (KLCI), the degree of persistence based on models 1 and 2 is about 8.58 and 8.86 months, respectively.

In sum, the short half-life and the tight CIs (around 4-5 months) for local persistence and account for structural breaks encourage us to conclude that shocks to the stock market display mean-reverting behavior and do not lead to a permanent deviation. Looking at individual half-lives and comparing the half-life of the individual stock price series to the broader KLCI index, we can conclude that some (but not all, e.g., plantation sector) sectors are well integrated into a single Malaysian stock market. This finding highlights the aggregation bias that may arise using the broader index. Given the high speed of mean reversion and the tight bounds for the half-life estimates (upper bounds mostly close to point estimates and never exceeds 2.69 years), we may conclude there is no single sector in the equity market that drives the market to be inefficient: Stock prices revert to their fundamental values at high speed of convergence after a shock. The finite half-life upper bounds, with the plantations sector highest upper limit, are not inconsistent with long-run mean reversion in stock prices. The last finding can be used as supportive evidence of market inefficiency.

TABLE 3. The degree of local persistence, half-lives (in months) and confidence intervals

| | Model 1 | | | | Model 2 | | | |
|---|---|---|---|---|---|---|---|---|
| | $\hat{d}$ | HL(M) | se($\hat{\rho}$) | 95%CI | $\hat{d}$ | HL(M) | se($\hat{\rho}$) | 95%CI |
| FTSE KLCI | 0.42 | 8.58 | 0.021 | [4.28,12.88] | 0.43 | 8.86 | 0.020 | [4.35,13.38] |
| 10 sectors | | | | | | | | |
| Finance | 0.46 | 9.08 | 0.023 | [3.68,14.49] | 0.34 | 4.63 | 0.032 | [2.66,6.60] |
| Industry | 0.35 | 5.10 | 0.031 | [2.82,7.38] | 0.31 | 4.04 | 0.035 | [2.44,5.65] |
| Plantations | 0.57 | 17.64 | 0.017 | [3.03,32.25] | 0.51 | 12.42 | 0.020 | [3.79,21.06] |
| Properties | 0.42 | 7.60 | 0.025 | [3.46,11.74] | 0.39 | 6.12 | 0.028 | [3.13,9.11] |
| Tin & Mining | 0.38 | 5.98 | 0.029 | [3.09,8.86] | 0.37 | 5.48 | 0.030 | [2.95,8.02] |
| Construction | 0.32 | 3.81 | 0.041 | [2.11,5.51] | 0.33 | 4.10 | 0.040 | [2.20,5.99] |
| Industrial Products | 0.40 | 5.86 | 0.033 | [2.62,9.10] | 0.32 | 3.89 | 0.041 | [2.13,5.65] |
| Consumer Products | 0.33 | 4.08 | 0.040 | [2.19,5.96] | 0.41 | 6.08 | 0.033 | [2.66,9.50] |
| Trade & Services | 0.39 | 5.56 | 0.034 | [2.56,8.55] | 0.27 | 2.98 | 0.047 | [1.80,4.16] |
| Technology | 0.41 | 5.06 | 0.045 | [1.79,8.33] | 0.29 | 2.89 | 0.060 | [1.47,4.30] |
| Mean | 0.41 | 7.12 | 0.03 | [2.88,11.37] | 0.36 | 5.59 | 0.04 | [2.69,8.49] |
| Median | 0.40 | 5.86 | 0.03 | [2.82,8.86] | 0.34 | 4.63 | 0.03 | [2.66,6.60] |

*Notes*: The persistence parameter denoted as $\hat{d} = 1 - \hat{\rho} = n^{-d}$, where $0 < d < 1$, and $\hat{\rho}$ are drawn from the Narayan and Popp's (2010) model. HL(*M*) is the half-life for local-persistent model measured in months by $\ln(0.5b(\hat{1}) / (-1/n^{\hat{d}})$. The two-sided 95% confidence intervals (CIs) measured in monthly are constructed according to $\hat{h}_{0.50} \pm 1.96\text{se}(\hat{\rho})([-1\text{n}0.5/\hat{\rho}][1\text{n}(\hat{\rho})]^{-2})$ where $\text{se}(\hat{\rho}) = \sqrt{2}/(n^{\frac{1}{2}+\frac{\hat{d}}{2}})$ (see Kim & Lima 2010).

## DISCUSSIONS

We applied a battery of unit-root tests to the stock price series and the result shows that the Liu and Narayan (2011) yield the most rejection of the unit-root null hypothesis. This finding confirmed our prior believe that conflicting results on the Fama and French (1988) and Poterba and Summers (1988) mean-reversion results can arise depending on whether structural breaks and heteroskedasticity are accommodated in the analysis. Clearly, our results show the correct model specification is important not only for determining the order of integration in the series, but also in measuring the degree of persistence in stock prices. It also reveals the risk of failing to reject the null due to model misspecification. The adjustments speed are much faster than those found in previous work based on long-span data from the major industrialised countries; see Balvers, Wu and Gilliland (2000, 3 to 3.5 years), Gropp (2004, 4.5 to 8 years), Kim, Stern and Stern (2010, 2.6 to 17.3) and Spierdijk et al. (2012, 2.0 to 22.6 years). Of course, one needs to be careful in comparing these results as they were based on different sample of countries, time period and model specifications.

Based on a different approach and for a large set of countries, Balvers et al. (2000) not only established mean-reversion, but indicate that it takes approximately 3.5 years for stock prices to absorb half the shock with 90% CI for the half-life equal to [2.4, 5.9] years. The empirical analysis used is based on annual data (1970-1996) and employed a panel-data framework. In a study covering a group of emerging market economies, Chaudhuri and Wu's (2003) show a speed of adjustment to about 2.5 years. Nonetheless, our estimates are in sharp contrast to their findings. Obviously, Chaudhuri and Wu's (2003) results are not directly comparable to those obtained in this study. First, we consider breaks in the calculation of the half-lives as accounting for the change in the persistence due to major historical episodes of economic crisis in recent decades. Second, unlike Chaudhuri and Wu (2003), we used the in-between process to show that the shocks are less long-lasting. More importantly, our CIs are much narrower (compact) when compared to previous works, meaning that price deviations temporary are rapidly arbitrage away. It should be mentioned that those studies on the major industrialized were based on different methods. Except for Kim et al. (2010), the other studies were based on annual frequency data. It should be noted that temporal aggregation from monthly to the quarterly to the annual frequency could induce persistence in the financial series (Paya, Duarte & Holden 2007).

Spierdijk et al. (2012) has highlighted the fact that, during a crisis, half-lives are likely to be shorter. Based on data that extend over more than a century (1900-2009), they find that the half-life ranges from 2.0 years to 22.6 years. Specifically, the speed with which stock prices adjust to shocks and revert to their fundamental value is greater during periods of high economic uncertainty. Similarly, in this study, we show that ignoring structural breaks due to extreme events may lead to spurious result.

Several studies have shown that excess returns can be earned by exploiting the mean-reversion of stock prices (Balvers et al. 2000; Campbell & Shiller 2001; Gropp 2004; Spierdijk et al. 2012, just to name four). If the stock price is mean-reverting, it follows that low returns are followed by higher expected future returns. This means that investors (e.g., portfolio managers) could invest in the equity market after a fall in the stock market to earn excess returns by exploiting the mean-reversion of stock prices (Spierdijk et al. 2012). Our results also highlight that the speed at which stock prices revert to their fundamental value may vary across economic sectors (17.6 to 3.8 months for Model 1; 12.4 to 2.9 for Model 2). It should be noted that a high speed of reversion as well as a high uncertainty of estimated half-lives limit the possibility of exploiting mean-reversion in a trading strategy. Finally, it should be noted that our result favours mean-reversion with breaks. This means that the historical events as well as the policies that follow those events (e.g., imposing capital controls) have only a temporary effect on stock prices (or returns).

For the emerging economies, the literature has yet to show strong evidence of mean-reverting behavior of stock prices. This is due to the difficulties associated with measuring the speed of adjustment largely due to short data span available for the analysis. Investors are likely to overestimate risk exposure if they underestimate the degree of mean-reversion. Mean-reversion in stock prices implies mean-reversion in stock prices, but the reverse is not true. It should be noted that mean-reversion less risky for investors with long investment horizons and so a larger share of wealth may be allocated stocks; see the two papers by Gropp (2004) and Balvers et al. (2000) for more details of such strategy. Mean-reverting low prices are followed by relatively high expected future returns, which could encourage long-term investors (pension fund managers and other institutional managers) to invest in equity market after a stock market downturn.

There are several theories that provide plausible explanations for the mean-reversion in stock prices that we have reported earlier; see Balvers et al. (2000). These theories are based on individual stock prices but can be easily extended to national market level. For example, in Chan (1988), the author presents the argument that after a substantial loss the firms in a country index are more highly leverages (if no adjustments to capital structure are made). Thus, the betas of their equities rise and return are expected to be higher. Finally, Ibrahim and Law (2013) find that the effect of the AFC on Malaysia's economic growth can be substantial (2.7% of the post-crisis gross domestic product) largely due to the stock market (wealth) impact on private consumption. Our findings suggest that the authorities need not pay too much attention to sharp movements in stock prices because they are only temporary, that is, they may be persistent but ultimately mean-reverting a characteristic well described by the local-persistent model.

## CONCLUSION

This study investigates the EMH uses monthly frequency data from 1980 to 2011. This paper makes two main contributions. First, we show that when the correct specification is used, the data reject the unit-root null hypothesis. In particular, when a GARCH-based structural breaks test that controls for heteroskedasticity and structural breaks is used, the unit-root null is easily rejected. Periods of high volatility may not be representative of stock price behavior and ignoring high volatility years may lead to slower speed of mean reversion (presence of unit root). In all, our results suggest that EMH does not hold in the 10 sector indices of the KLSE. Second, we contribute to the on-going debate on whether the stock price is mean-reverting through the lens of local-persistent model. A unique feature of the locally persistence process lies between an exact unit-root and a stationary process. We confirm that stock prices have a low degree of persistence after breaks are introduced in the model. We also find that the CIs constructed from the local-persistent model provide useful information regarding the half-life of stock prices.

From a policy perspective, the rapid speed of mean-reversion as well as the narrow CIs of our results provides further support for the view that excess returns can be earned by exploiting the mean-reversion of stock prices. It is possible for investors (including pension fund managers) to exploit mean-reversion as part of a trading strategy. Stock prices are less risky in the long-run since low returns are followed by higher expected future returns, as seen in the recent economic and financial crises.

## ENDNOTES

[1] As in Lee, Lee and Lee (2010) and Narayan (2005), the terms "“random walk” and “unit-root”" are used interchangeably in this paper.

[2] Lim (2008), for instance, argued that aggregate analysis using the broader market index data could easily mask the impact of crises on the various sectors of the equity market. We depart from Lim's work by providing the speed of mean reversion through a model that considers the persistence of the stock price series. Specifically, we construct confidence intervals for the half-lives in order to provide a more complete picture of the speed of convergence towards fundamental value.

[3] The EMH has been tested in previous studies by showing that successive changes in stock prices are independent of one another and therefore cannot contain information for predicting future prices (Abeysekera 2001). Researchers have also used various variance ratio statistics (for recent survey, see Lim & Brooks 2011) and long-run regression coefficients estimated to examine the random walk model and the EMH but this is beyond the scope of our study.

[4] Fractionally integrated series is denoted by $I(d)$, $0 \leq d \leq 1$. When the series is $I(0)$ (i.e., $d = 0$), the shocks die out at a geometric rate; when the series is $I(1)$ (i.e., $d = 1$) shocks have permanent effect. In the intermediate case (i.e., $0 < d < 1$), the series is mean reverting but the shock to die out slower hyperbolic (rather than geometric) rate. In short, they are more persistent than an $I(0)$ variable.

[5] Thus, failure to account for structural breaks can lead to an upward bias in the autoregressive coefficient of the unit-root regression model (and the corresponding upward bias in the persistence measure), leading to the conclusion that shocks are more persistent than they actually are (Basher & Carrion-i-Silvestre 2013).

[6] The price index is calculated as $PI_t = PI_{t-1} \left[\sum_1^n P_t N_t / \sum_1^n (P_{t-1} N_t f)\right]$, where $PI_t$ is the index value at day $t$, $PI_{t-1}$ is the index value on the previous working day (of $t$), $P_t$ is the unadjusted share price on day $t$, $P_{t-1}$ is the unadjusted share price on the previous working day (of $t$), $N_t$ is the number of shares issued on day $t$, $f$ is an adjustment factor for a capital action accruing on day $t$, and $n$ is the number of constituents in the index.

[7] According to Kim and Shamsuddin (2008) the “Pricing efficiency of one market depends on the level of equity market development as well as the regulatory framework conducive to transparent corporate governance.” Baharumshah, Sarmidi and Tan (2003) suggest that the Asian markets are all closely linked with one another and with the world capital markets in the post-liberalization era.

[8] The potential bias presents in the least squares estimation of half-life has long been recognized in the literature. Several new methods have been proposed to correct the bias. However, these methods have produced mixed and often uninformative results with infinite half-lives; see Kim and Lima (2010), among others.

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## APPENDIX

TABLE A1. Descriptive statistics

| | Period I: 1980:M1-1996:M12 | | | | | Period II: 1999:M1-2011:M7 | | | | |
|---|---|---|---|---|---|---|---|---|---|---|
| | Mean | Median | Max | Min | Std. Dev. | Mean | Median | Max | Min | Std. Dev. |
| FTSE KLCI | 1.999 | 1.963 | 2.741 | 1.055 | 0.390 | 2.214 | 2.190 | 2.602 | 1.742 | 0.215 |
| Finance | 3.888 | 3.647 | 4.797 | 3.121 | 0.542 | 4.297 | 4.276 | 4.827 | 3.631 | 0.267 |
| Industry | 2.847 | 2.815 | 3.329 | 2.087 | 0.286 | 2.915 | 2.946 | 3.326 | 2.368 | 0.224 |
| Plantations | 3.050 | 2.883 | 3.809 | 2.507 | 0.380 | 3.405 | 3.208 | 4.357 | 2.621 | 0.516 |
| Properties | 3.103 | 3.141 | 3.709 | 2.438 | 0.350 | 1.992 | 1.958 | 2.655 | 1.501 | 0.251 |
| Tin & Mining | 1.438 | 1.383 | 2.398 | 0.733 | 0.416 | 1.134 | 1.117 | 2.751 | 0.567 | 0.305 |
| Construction | 1.644 | 1.645 | 1.882 | 1.380 | 0.123 | 0.667 | 0.662 | 1.193 | 0.238 | 0.230 |
| Industrial Products | 0.730 | 0.724 | 0.966 | 0.459 | 0.106 | -0.198 | -0.200 | 0.224 | -0.547 | 0.159 |
| Consumer Products | 0.942 | 0.932 | 1.142 | 0.775 | 0.108 | 0.842 | 0.832 | 1.393 | 0.261 | 0.277 |
| Trade & Services | 0.805 | 0.802 | 0.985 | 0.612 | 0.093 | 0.260 | 0.254 | 0.585 | -0.146 | 0.173 |
| Technology | n.a. | n.a. | n.a. | n.a. | n.a. | -1.153 | -1.248 | 0.595 | -2.280 | 0.617 |

*Source*: Indices are extracted from the FTSE Bursa Malaysia and the statistics computed by authors.

TABLE A2. Conventional unit-root tests

| | ADF | k | MZa | MZt | MSB | MPT | k |
|---|---|---|---|---|---|---|---|
| Panel A: Constant | | | | | | | |
| FTSE KLCI | -2.43 | [3] | -2.91 | -1.00 | 0.34 | 7.95 | [3] |
| 10 sectors | | | | | | | |
| Finance | -2.56 | [13] | -0.93 | -0.43 | 0.47 | 14.82 | [13] |
| Industry | $-2.71^{c}$ | [15] | -0.60 | -0.34 | 0.57 | 20.23 | [15] |
| Plantations | -1.12 | [1] | 0.19 | 0.10 | 0.52 | 21.13 | [1] |
| Properties | -1.52 | [1] | -4.68 | -1.50 | 0.32 | 5.29 | [1] |
| Tin & Mining | $-2.93^{b}$ | [1] | $-13.93^{a}$ | $-2.63^{a}$ | $0.19^{b}$ | $1.81^{b}$ | [1] |
| Construction | -1.86 | [14] | -2.36 | -1.03 | 0.43 | 10.01 | [13] |
| Industrial Products | -2.51 | [1] | -0.88 | -0.56 | 0.64 | 22.09 | [13] |
| Consumer Products | -1.26 | [1] | -4.48 | -1.29 | 0.29 | 5.85 | [1] |
| Trade & Services | -2.29 | [1] | -2.48 | -1.09 | 0.44 | 9.73 | [1] |
| Technology | -1.77 | [3] | 0.41 | 0.33 | 0.79 | 41.10 | [3] |
| Panel B: Constant and Trend | | | | | | | |
| FTSE KLCI | -2.97 | [3] | -10.60 | -2.30 | 0.22 | 8.61 | [0] |
| 10 sectors | | | | | | | |
| Finance | -2.66 | [1] | -12.58 | -2.51 | 0.20 | 7.24 | [6] |
| Industry | -2.80 | [14] | -8.37 | -2.04 | 0.24 | 10.91 | [15] |
| Plantations | -1.87 | [1] | -7.46 | -1.90 | 0.26 | 12.27 | [1] |
| Properties | -2.61 | [1] | -6.96 | -1.86 | 0.27 | 13.11 | [1] |
| Tin & Mining | -2.99 | [1] | $-15.08^{c}$ | $-2.75^{c}$ | $0.18^{c}$ | $6.04^{c}$ | [1] |
| Construction | -1.91 | [14] | -10.61 | -2.25 | 0.21 | 8.86 | [14] |
| Industrial Products | -2.00 | [1] | -4.51 | -1.36 | 0.30 | 19.19 | [1] |
| Consumer Products | -2.02 | [1] | -5.54 | -1.55 | 0.28 | 16.18 | [1] |
| Trade & Services | -1.99 | [0] | -6.59 | -1.71 | 0.26 | 13.89 | [1] |
| Technology | $-3.44^{c}$ | [0] | -10.71 | -2.28 | 0.21 | 8.69 | [3] |

*Notes*: ($^{b}$) and ($^{c}$) denote statistical significance at 5 and 10% level, respectively. The values in [ ] denote the lag length (*k*) based on $K_{max} = 15$. The spectral GLS-detrended AR is based on modified AIC (MAIC) for the Ng-Perron test while ADF automatic is based on the MAIC criterion. Columns three to six refer to four modified test statistics constructed by Ng and Perron (2001). The MZa, MZt, MSB and MPT are modified forms of Phillips ($Z\alpha$, 1987), Phillips and Perron (Zt, 1988), Bhargava ($R_1$, 1986), and Point Optimal by Elliot, Rothenberg and Stock (1996) test statistics, respectively.