Testing Weak Form Market Efficiency for Emerging Economies: A Nonlinear Approach

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Abstract

In this paper, we address weak form stock market efficiency of Emerging Economies, by testing whether the price series of these markets contain unit root. Nonlinear behavior of stock prices is well documented in the literature, and thus linear unit root tests may not be appropriate in this case. For this purpose, we employ the nonlinear unit root test procedure recently developed by Kapetanios *et al.* (2003) and nonlinear panel unit root test Ucar and Omay (2009) that has a better power than standard unit root tests when series under consideration are characterized by a slower speed of mean reversion. Large power gains are achieved through combining cross-sectional information and nonlinear estimation techniques in computing unit root tests. The results of ADF and PP indicate that Bulgarian, Greek, Hungarian, Polish, Romanian, Russian, Slovenian and Turkish stock markets are weak form efficient, while the results of nonlinear unit root test implies that Russian, Romanian and Polish stock markets are not weak form efficient. Moreover, the linear panel unit root test suggest that this group as all efficient where as nonlinear panel unit root test suggest as a group they are not efficient.

Keywords: Linear and Nonlinear Unit root and Panel Unit Root, Emerging Markets, Market Efficiency
JEL Code: C1, C12, G14

1. Introduction

The Efficient Market Hypothesis states that security prices fully reflect all available information and that the price fluctuations are unpredictable. Since the market absorbs all relevant information as it becomes available, stock prices should fluctuate as random white noise. The concept of market efficiency is mainly based on the reaction of stock price to new information which means a surprise because if it were to be predictable, then the market should have already compensated for it. Following the argument that the stock prices already incorporate all available information and the stock price changes require a news release which is itself unpredictable by definition, then price changes should be unpredictable and random. The hypothesis indicates that if price formation of a financial asset is random and the return from such a financial asset is unpredictable, then the market is informationally efficient and as Aguirre and Saidi (1998) argue, in such an efficient market it is impossible for an investor to gain excess returns through speculation, because prices do reflect all available information (Azad and Bashar 2010:3). Thus, in an efficient market, price changes can be argued to follow a "random walk". Hence, the Efficient Market Hypothesis carries a close relation with the Random Walk Model. If stock prices follow a random walk which is satisfied by the unpredictability of stock returns, then stock prices are characterized by a unit root.

The liberalization of financial markets and advances in technology coupled with lower costs of investing in international markets has created an increased demand for such transactions in emerging markets. As these markets become more integrated with global equity markets, they increasingly attract international investors hoping to benefit from abnormal high returns as well as portfolio risk diversification. The study of efficient markets hypothesis has some implications for understanding the price formation in capital markets, may prove to be a worthy weapon to develop trading strategies and to build a general idea of the investor's behaviour of a market. Market Efficiency also has important implications for managerial decisions, especially those pertaining to common stock issues, stock repurchases, and tender offers (Brigham and Gapenski 1997: 321). Actually, as Seiler and Rom (1997: 49) discussed, market efficiency is directly or implicitly tested at any time a study is performed to identify stock price reactions to certain events such as dividend announcements (Bajaj and Vijh 1995, 1990), earnings announcements (Bamber 1987), stock splits (Copeland 1979), large block transactions (Holthausen et.al. 1987; Kraus and Stoll 1972), repurchase tender offers (Lakonishok and Vermaelen 1990), and other public announcements (Kim and Verrecchia 1991a; 1991b) while a more encompassing or macro evaluation of market efficiency can be made by testing whether or not the returns in a market follow a random walk process over a longer period of time.

In recent years, although, predictability and efficiency of emerging markets have attracted interest of financial economists (e.g., Emerson *et al.*, 1997; Dockery and Vergari, 1997; Liu *et al.*, 1997; Zalewska-Mitura and Hall, 1999; Rockinger and Urga, 2001; Harrison and Paton, 2004; Cajueiro and Tabak, 2006), no consensus on whether or not efficient market hypothesis holds for these markets is attained yet. A common feature of these studies is that possible nonlinearities in conditional mean of the series have not been taken into account in testing efficiency of these markets. However, it is well known that many economic and financial time series follow nonlinear processes (e.g., Granger and Teräsvirta, 1993; Franses and van Dijk, 2000). Therefore, possible nonlinearities in data generating process should explicitly be taken into account in analysing financial time series in order to avoid spurious results.

The economic theory suggests a number of sources of nonlinearity in the financial data. One of the most frequently citied reasons of nonlinear adjustment is presence of market frictions and transaction costs. Existence of bid-ask spread, short selling and borrowing constraint and other transaction costs render arbitrage unprofitable for small deviations from the fundamental equilibrium. Subsequent reversion to the equilibrium, therefore, takes place only when the deviations from the equilibrium price are large, and thus arbitrage activities are profitable. Consequently, the dynamic behaviour of returns will differ according to the size of the deviation from equilibrium, irrespective of the sign of disequilibrium, giving rise to asymmetric dynamics for returns of differing size (e.g., Dumas, 1992; Shleifer, 2000). In addition to transaction costs and market frictions, interaction of heterogeneous agents (e.g., Hong and Stein, 1999; Shleifer, 2000), diversity in agents' beliefs (e.g., Brock and Hommes, 1998) also may lead to persistent deviations from the fundamental equilibrium.

Recent developments in nonlinear time series analysis allow modelling financial time series more appropriately (e.g., Granger and Teräsvirta, 1993; Franses and van Dijk, 2000). If dynamics of the market differ according to the size of deviations from equilibrium as the economic theory suggests, then such nonlinearities are more suitably modelled by an exponential smooth transition autoregressive (ESTAR) model, a class of smooth transition autoregressive (STAR) models popularized by Granger and Teräsvirta (1993) and Teräsvirta (1994). ESTAR models have extensively been used in empirical literature to test nonlinear mean reversion of financial time series, mainly for testing purchasing power parity (see, inter alia, Michael *et al*., 1997; Taylor and Peel, 2000; Taylor *et. al*, 2001; Gallagher and Taylor, 2001). For example Hasanov and Omay (2008) have shown that the predictability of Greek and Turkish stock markets is increasing when these markets are modeled by a STAR model. This result is a confirmation of weak form inefficiencies for these markets which verifies our results in this study. Recently, Kapetanios *et al*. (2003) have developed a unit root test procedure in an ESTAR framework, which has a better power than conventional Dickey-Fuller test. On the other hand, Ucar and Omay (2009) have developed a panel unit root test procedure in an ESTAR framework, which has a better power than conventional IPS (Im, Pesaran and Shin) test. In this paper we apply Kapetanios *et al.* (2003) and Ucar and Omay (2009) nonlinear unit root and panel unit root tests respectively to eight emerging markets, namely, Bulgarian, Greek, Hungarian, Polish, Romanian, Russian, Slovenian and Turkish stock price indices to test whether the series contain unit root. To provide basis for comparing the results of nonlinear unit root tests, we also apply unit root tests that do not take account of nonlinearity in the series, namely ADF, PP and IPS.

The results of ADF and PP indicate that Bulgarian, Greek, Hungarian, Polish, Romanian, Russian, Slovenian and Turkish stock markets are weak form efficient, while the results of nonlinear unit root test implies that Russian, Romanian and Polish stock markets are not weak form efficient. Moreover, we apply linear and nonlinear panel unit root test to this group of countries. The linear panel unit root test suggest that this group as all weak form efficient where as nonlinear panel unit root test suggest as a group they are weak form inefficient. These results show that the markets in this group are weak form efficient in linear tests, however, the true data generating process is nonlinear and stationary, hence we can conclude that the linear test gives spurious result of market efficiency.

In this study, we contribute to the controversy literature on the validity of weak form market efficiency in the emerging markets by concentrating on the European emerging markets. Since, some of those markets are also among the so called transition markets, it also contributes to the relatively limited literature on the transition economies. Another important contribution of this research lies in the methodology employed. During the analysis, not only conventional ADF and PP unit root tests is used, but also nonlinear unit root test recently proposed by Kapetanios *et al.* (2003) and Ucar and Omay (2009) are applied. By applying the nonlinear and panel version of the unit root tests, we improve the power of the tests as much as possible by both combining cross-sectional information with nonlinearities in the data. Hence, nonlinear panel version of these tests gives us more vigorous result with respect to market efficiency. It is the first time a nonlinear panel unit root test is used in the market efficiency literature. On the other hand, we have taken into account the possible nonlinearities in conditional mean of the series in testing efficiency of these markets which is a deviation from the vast literature. Furthermore, we have explicitly dealt with the cross section dependency problem in panel unit root tests.

The remaining of the study is organized as follows. In part 2 the methodology of test procedure is given. In part 3 data and unit root test results are provided. Finally, section 4 concludes.

2. Methodology

In this section we briefly discuss the nonlinear unit root tests procedures developed by Kapetanios *et al.* (2003) and Ucar and Omay (2009). First of all, we explain the Kapetanios *et al.* (2003). Think about a univariate smooth transition autoregressive (STAR) [1] model of order 1:

$$y_t = \beta y_{t-1} + \gamma y_{t-1} F(\theta; y_{t-d}) + \varepsilon_t \,, \tag{1}$$

where y_t is a mean zero stochastic process for $t = 1,\ldots,T$, $\varepsilon_t \sim iid(0,\sigma^2)$, and β and γ are unknown parameters. The transition function $F(\theta; y_{t-d})$ is assumed to be of the exponential form:

[1] For a discussion of STAR models see and Granger and Teräsvirta, (1993) and Teräsvirta, (1994).

$$F(\theta; y_{t-d}) = 1 - \exp(-\theta y^2_{t-d}) , \tag{2}$$

where it is assumed that $\theta > 0$, and $d \geq 1$ is the delay parameter. The exponential function is bounded between zero and one, and is symmetrically U-shaped around zero. The parameter θ is slope coefficient and determines the speed of transition between two regimes that correspond to extreme values of the transition function. Using (2) in (1) one obtains the following exponential STAR (ESTAR) model:

$$y_t = \beta y_{t-1} + \gamma y_{t-1}\left[1 - \exp(-\theta y^2_{t-d})\right] + \varepsilon_t , \tag{3}$$

which after reparameterising can be written conveniently as

$$\Delta y_t = \phi y_{t-1} + \gamma y_{t-1}\left[1 - \exp(-\theta y^2_{t-d})\right] + \varepsilon_t , \tag{4}$$

where $\phi = \beta - 1$. The ESTAR model has a nice property that it allows modelling different dynamics of series depending on the size of the deviations from the fundamental equilibrium (e.g., Teräsvirta and Anderson, 1992). As briefly argued above, the arbitrageurs shall not engage in reversion strategies if deviations from the equilibrium are small in size and therefore arbitrage is not profitable. If the deviations from equilibrium are large enough, however, arbitrageurs shall engage in profitable reversion trading strategies, and thus bring the prices to their equilibrium levels. In the context of ESTAR model, this would imply that while $\phi \geq 0$ is possible, one must have $\gamma < 0$ and $\phi + \gamma < 0$ for the process to be globally stationary. Under these conditions, the process might display unit root for small values of y^2_{t-d} , but for larger values of y^2_{t-d} it has stable dynamics, and as a result, is geometrically ergodic. As shown by Kapetanios *et al.* (2003), ADF test may not be very powerful when the true process is nonlinear yet globally stationary.

Imposing $\phi = 0$ (which implies that y_t follows a unit root in the middle regime) the ESTAR model can be written as

$$\Delta y_t = \gamma y_{t-1}\left[1 - \exp(-\theta y^2_{t-d})\right] + \varepsilon_t , \tag{5}$$

The global stationarity of the process y_t can be established by testing the null hypothesis $H_0 : \theta = 0$ against the alternative $H_1 : \theta > 0$. However, testing the null hypothesis directly is not feasible since the parameter γ is not identified under the null. To overcome this problem, Kapetanios *et al.* (2003) follow suggestion of Luukkonen *et al.* (1988) to replace the transition function by its suitable Taylor approximation to derive a t-type test statistic. Substituting the transition function with its first order Taylor approximation yields the following auxiliary regression:

$$\Delta y_t = \delta y^3_{t-d} + e_t , \tag{6}$$

where e_t comprises original shocks ε_t as well as the error term resulting from Taylor approximation. The test statistic for $\delta = 0$ against $\delta < 0$ is obtained as follows:

$$t_{NL} = \hat{\delta} / s.e.(\hat{\delta}) , \tag{7}$$

where $\hat{\delta}$ is the OLS estimate and s.e.($\hat{\delta}$) is the standard error of $\hat{\delta}$.

To accommodate stochastic processes with nonzero means and/or linear deterministic trends, one needs following modifications. In the case where the data has nonzero mean, i.e., $x_t = \mu + y_t$, one must replace the raw data with de-meaned data $y_t = x_t - \bar{x}$ where $\bar{x}$ is the sample mean. In the case where the data has a nonzero mean and a nonzero linear trend, i.e., $x_t = \mu + \alpha t + y_t$, one must instead use the de-meaned and de-trended data $y_t = x_t - \hat{\mu} - \hat{\alpha} t$ where $\hat{\mu}$ and $\hat{\alpha}$ are OLS estimators of μ and α.

In the more general case where errors in (5) are serially correlated, one may extend (5) to

$$\Delta y_t = \sum_{j=1}^{p} \rho_j \Delta y_{t-j} + \gamma y_{t-1}\left[1 - \exp(-\theta y_{t-d}^2)\right] + \varepsilon_t \tag{8}$$

The t_{NL} statistic for testing $\theta = 0$ in this case is given by the same expression as in (7), where $\hat{\delta}$ is the OLS estimate and s.e.($\hat{\delta}$) is the standard error of $\hat{\delta}$ obtained from the following auxiliary regression with *p* augmentations:

$$\Delta y_t = \sum_{j=1}^{p} \rho_j \Delta y_{t-j} + \delta y_{t-d}^3 + e_t \tag{9}$$

In practice, the number of augmentations *p* and the delay parameter *d* must be selected prior to the test. Kapetanios *et al*. (2003) propose that standard model selection criteria or significance testing procedure be used for selecting the number of augmentations *p*. They also suggest that the delay parameter *d* be chosen to maximize goodness of fit over $d = \{1,2,\ldots,d_{\max}\}$.

We also applied the non-linear panel unit root test newly proposed by Ucar and Omay (2009), which we called as the UO test. The UO test has a good power when the series under investigation follow a non-linear process. A brief review of the UO test can be given as follows.

Let z_{it} be panel exponential smooth transition autoregressive process of order one (PESTAR(1)) on the time domain t = 1,2,…,T for the cross-section units i = 1,2,…,N. Consider z_{it} generated by the following PESTAR process with fixed effect parameter α_i:

$$\Delta z_{it} = \alpha_i + \phi z_{it-1} + \gamma_i z_{it-1}\left[1 - \exp(-\theta_i z_{it-d}^2)\right] + \varepsilon_{it} \tag{10}$$

where $d \geq 1$ is the delay parameter and $\theta_i \geq 0$ represents the speed of revision for all units; ε_{it} is a serially and cross-sectionally uncorrelated disturbance term with zero mean and variance σ_i^2.

Following previous literature, Ucar and Omay (2009) set $\phi_i = 0$ for all i and d=1 which gives specific PESTAR(1) model:

$$\Delta z_{it} = \alpha_i + \gamma_i z_{it-1}\left[1 - \exp(-\theta_i z_{it-d}^2)\right] + \varepsilon_{it} \tag{11}$$

Non-linear panel data unit root test based on regression (11) with augmented lag variables in empirical application is simply to test the null hypothesis $\theta_i = 0$ for all i against $\theta_i \geq 0$ for some i under the alternative. However, direct testing of the null hypothesis is problematic since γ_i is not identified under the null. This problem can be solved by taking first-order Taylor series expansion to the PESTAR(1) model around $\theta_i = 0$ for all i. Hence the obtained auxiliary regression is given by:

$$\Delta z_{it} = \alpha_i + \delta_i z_{it-1}^3 + \varepsilon_{it} \tag{12}$$

where $\delta_i = \theta_i \gamma_i$. In empirical application equation (12) augmented by lagged variables of dependent variable by using AIC and SIC criteria. Based on equation (12), hypothesis for unit root testing is

$$H_0 : \delta_i = 0, \quad \text{for all i,} \quad \text{(Linear Nonstationary)}$$
$$H_0 : \delta_i < 0, \quad \text{for all i,} \quad \text{(Non-linear Stationary)}$$

The UO test is constructed by standardizing the average of individual KSS statistics across the whole panel. First, the KSS test for the i^{th} individual is the t-statistics for testing $\delta_i = 0$ in equation (12) defined by:

$$t_{i,NL} = \frac{\Delta z_i' M_t z_{i,-1}^3}{\hat{\sigma}_{i,NL}\left(z_{i,-1}' M_t z_{i,-1}\right)^{3/2}}$$

where $\hat{\sigma}_{i,NL}^2$ is the consistent estimator such that $\hat{\sigma}_{i,NL}^2 = \Delta z_i' M_t z_i /(T-1)$, $M_t = I_T - \tau_T\left(\tau_T'\tau_T\right)^{-1}\tau_T'$ with $\Delta z_i = \left(\Delta z_{i-1}, \Delta z_{i-2}, \ldots \Delta z_{i-T}\right)'$ and $\tau_T = (1,1,\ldots,1)$.

Furthermore, when the invariance property and the existence of moments are satisfied, the usual normalization of $\overline{t}_{NL}$ statistic yields as follows:

$$\bar{Z}_{NL} = \frac{\sqrt{N}\left(\bar{t}_{NL} - E(t_{i,NL})\right)}{\sqrt{\operatorname{var}(t_{i,NL})}}$$

where $\bar{t}_{NL} = N^{-1}\sum_{i=1}^{N} t_{NL}$; $E(t_{i,NL})$ and $\operatorname{var}(t_{i,NL})$ can be found in Table 1 of Ucar and Omay (2009).

Up until here, we have not seen anything about cross-section dependency. Most of the panel data models assume that disturbances in panel models are cross-sectionally independent. However, cross-section dependence may arise for several reasons often, due to spatial correlations, spillover effects, economic distance, omitted global variables and common unobserved shocks. In the presence of cross-section dependence, it is well known that neglecting cross-section dependence can lead to biased estimates and produce misleading inference. In large panels, where N is sizeable amount cross-section dependency is not a serious problem to control. But Pesaran (2004) pointed out that cross-section dependency continues to exist in large panel as well as small panels. Therefore, we have to make misspecification tests. Thus, we have made a diagnostic check for cross-section dependency for non-linear panel models following Omay and Kan (2010). Pesaran (2004) showed that his CD test can also be applied to a wide variety of models, including small/large N and T. Additionally, this simple diagnostic test does not require an a priori specification of connection or spatial matrix. CD test is based on simple average of all pair-wise correlation coefficients of the OLS residuals from the individual regressions in the panel:

$$\Delta y_{it} = \mu_i + \beta_i^{'} x_{\text{it}} + u_{it} \tag{13}$$

where, on the time domain t = 1,2,…,T, for the cross-section units i = 1,2,…,N. $x_{\text{i,t}}$ is a kx1 vector of observed time-varying regressors. The individual intercepts, μ_i and slope coefficients β_i are defined on a compact set permitted to vary across i. For each i, $u_{it} \sim iid(0, \sigma_{i,u}^2)$, for all t although they could be cross-sectionally correlated.

The sample estimate of the pair-wise correlation of the residuals is:

$$\hat{\rho}_{ij} = \hat{\rho}_{ji} = \frac{\sum_{t=1}^{T} e_{it} e_{jt}}{\left(\sum_{t=1}^{T} e_{it}^2\right)^{1/2} \left(\sum_{t=1}^{T} e_{jt}^2\right)^{1/2}} \tag{14}$$

And the e_{it} is the OLS estimates of u_{it} defined by

$$e_{it} = \Delta y_{it} - \hat{\mu}_i - \hat{\beta}_i^{'} x_{\mathrm{it}} \tag{15}$$

The proposed CD test by Pesaran (2004) is:

$$CD_{LM} = \sqrt{\frac{2T}{N(N-1)}} \left(\sum_{i=1}^{N-1} \sum_{j=i+1}^{N} \hat{\rho}_{ij} \right) \tag{16}$$

$$CD_{LM2} = \sqrt{\frac{1}{N(N-1)}} \left(\sum_{i=1}^{N-1} \sum_{j=i+1}^{N} (T.\hat{\rho}_{ij}^2 - 1) \right)$$

$$CD_{LM1} = T. \left(\sum_{i=1}^{N-1} \sum_{j=i+1}^{N} \hat{\rho}_{ij} \right)$$

CD test statistic has exactly mean zero for fixed values of T and N, under a broad class of panel data models. The CD test is based on simple average of all pair-wise correlation coefficients of the NLLS residuals from the individual regressions in the smooth transition panel model Omay and Kan (2010):

$$\Delta y_{it} = \mu_i + \beta_0^{'} x_{it} + \beta_1^{'} x_{it} F(s_{it}; \gamma, c) + u_{it} \tag{17}$$

and the e_{it} is the NLLS estimates of u_{it} defined by

$$e_{it} = \Delta y_{it} - \hat{\mu}_i - \hat{\beta}_0^{'} x_{it} - F(\dot{s}_{it}; \hat{\gamma}, \hat{c}) \hat{\hat{\beta}}_1^{'} x_{it} \tag{18}$$

Where $F(\dot{s}_{it}; \hat{\gamma}, \hat{c}) = \dfrac{1}{1 + e^{-\hat{\gamma}(\dot{s}_{it} - \hat{c})}}$

These are the estimated values of the slope (γ) and threshold (c) parameters. The dot on the transition variable means that it is selected from the linearity tests. In non-linear models, the definition of the residual is ambigous and can be defined in a number of different ways. The above representation is the definition of disturbance of the non-linear models analogous to linear case. Thus the CD_{LM} tests are used in the study as proposed by Omay and Kan (2010).

3. Data and unit root test results

In this paper, major European emerging markets are tested for weak form efficiency. The investigated markets are Bulgarian, Greek, Hungarian, Polish, Romanian, Russian, Slovenian and Turkish markets. The data are monthly and sourced from Datastream. To test the weak form of market efficiency in these markets, stock prices in those markets are searched for whether they contain unit root. A finding of unit root would imply that stock prices are random walk processes, and thus, weak form efficient. For this purpose we carried out conventional ADF and PP unit root tests as well as nonlinear unit root test recently proposed by Kapetanios *et al.* (2003).

Table 1. Description of stock price series

Country	Series	Datastream Code	Period covered	Number of observations
Bulgaria	BSE Sofia Lazard	BSLAZ10	2002:01 – 2010:05	101
Greece	Total Market PI	TOTMKGR	2002:01 – 2010:05	101
Hungary	BUX	BUXINDX	2002:01 – 2010:05	101
Poland	Total Market PI	TOTMKPO	2002:01 – 2010:05	101
Romania	Total Market PI	TOTMKRM	2002:01 – 2010:05	101
Russia	AKM Composite	RSAKMCO	2002:01 – 2010:05	101
Slovenia	Total Market PI	TOTMKSL	2002:01 – 2010:05	101
Turkey	ISE 100	TOTMKTK	2002:01 – 2010:05	101

Series names, periods, and Datastream codes for the data are provided in Table 1.

It is well known that stock prices may contain time trend (see, for example, Beechey *et al.,* 2000). If the market is efficient, however, fluctuations in the stock prices away from trend should be unpredictable. Therefore, in conducting the above described nonlinear unit root test we consider de-meaned and de-trended series. The de-meaned and de-trended series were obtained by regressing the natural logarithms of index series on a constant and a linear time trend.

Preliminary tests for nonstationarity of the series and their differences, based on ADF (Dickey and Fuller, 1981) and PP (Phillips and Perron, 1988) tests are provided in Table 2. Both tests suggest that all stock price indices are *I(1)* processes, consistent with the efficient market hypothesis.

Table 2. Linear unit root test results

	ADF		PP	
Country	Log Level[a]	First Difference[b]	Log Level[a]	First Difference[b]
Bulgaria	-0.785	-6.983*	-0.760	-7.314*
Greece	-0.841	-6.890*	-0.874	-6.890*
Hungary	-1.651	-7.773*	-1.739	-7.765*
Poland	-1.186	-8.469*	-1.541	-8.465*
Romania	-1.445	-7.890*	-1.469	-7.899*
Russia	-2.142	-4.560*	-1.770	-7.356*
Slovenia	-1.722	-5.632*	-1.182	-5.686*
Turkey	-1.887	-9.719*	-2.015	-9.721*

Notes:
a) Regressions include an intercept and linear time trend.
b) Regressions include only intercept.
Optimal lag length in ADF test was selected using AIC with maximum lag order of 12. *, ** and *** indicate significance at 1%, 5% and 10% significance levels, respectively.

To carry out the nonlinear unit root tests, we firstly estimated an AR(12) model for each series and excluded insignificant (at 10% significance level) augmentation terms. Then, we estimated regression with selected augmentations to compute the t_{NL} statistics. We selected the delay parameter d that maximized R^2 over $d = \{1,2,...,12\}$. Unlike the case of testing linearity against STAR type nonlinearity, the t_{NL} test does not have an asymptotic standard normal distribution. Therefore, we bootstrapped the t_{NL} test statistic with 10,000 replications.

Table 3. Nonlinear unit root test results

Country	t_{NL}
Bulgaria	-1.324
Greece	-2.821
Hungary	-3.044
Poland	-3.138***
Romania	-3.217***
Russia	-3.203***
Slovenia	-1.754
Turkey	-2.230

Notes: The t_{NL} statistic was computed by bootstrapping with 10,000 replications. Asymptotic critical values of the t_{NL} statistic at 1%, 5% and 10% significance levels are -3.93, -3.40 and -3.13. These values are taken from Table 1, Kapetanios et al. (2003, p. 364). * and ** denote significance at 1% and 5% levels, respectively.

As the Table 3.3 reveals, the null hypothesis of unit root is rejected at 10% significance level for Russian, Romanian and Polish series suggesting that these markets are not efficient. The null of unit root is not rejected at conventional levels for the Bulgarian, Greek, Hungarian, Slovenian and Turkish series, implying that these markets are weak form efficient.

Now it is time to deal this group of countries in panel unit root context.

Table 4. Linear and nonlinear panel unit root test results without cross section dependency

	IPS		UO	
	Log Level[a]	First Difference[b]	Log Level[a]	First Difference[b]
t_{NL}	-1.458	-7.240*	-2.583***	-9.721*
z_{tbar}	2.598	-18.816	3.912***	24.564*

Notes:
a) Regressions include an intercept and linear time trend.
b) Regressions include only intercept.
Optimal lag length in IPS and UO tests were selected using AIC with maximum lag order of 12. *, ** and *** indicate significance at 1%, 5% and 10% significance levels, respectively.

Notes: asymptotic critical values of $\bar{t}_{NT}$ for UO test statistics at 1%, 5% and 10%significance levels are −2.44, −2.21, and −2.08 and for trend-intercepts are −2.94,−2.72, and −2.57. For intercept only, the values are taken from Table 2 of Ucar andOmay (2009, p: 6). Asymptotic critical values of t-bar statistics at 1%, 5% and 10%significance levels are−2.20,−1.95 and−1.85 and for the trend-intercepts are−4.50,−3.35, and −3.02. These values are taken from Table 2 IPS (2003, p 61–62). *, **, and *** denote significance at 1%, 5% and 10% levels, respectively. Besides, optimal lag length in these tests were selected using AIC with maximum lag order of 8.

The test of panel unit root explained in the previous section was based on the assumption of independence over cross-section units. However, we see from the below diagnostic check that this assumption is violated.

Table 5 Cross section dependency test

	Test statistics	P value
CD_{LM1}	44.933	0.000
CD_{LM2}	5.465	0.000
CD_{LM3}	4.492	0.000

Notes: Under the null hypothesis the CD statistics converge to a normal standard distribution. The values in the parentheses are *p* values.

To overcome the cross-section dependency problem, we implemented Sieve bootstrap approach which is very well outlined in Ucar and Omay (2009). The test results for the UO and IPS with Sieve bootstrap is given in the below Table 6:

Table 6. Linear and nonlinear panel unit root test results with cross section dependency

	IPS		UO	
	Log Level[a]	First Difference[b]	Log Level[a]	First Difference[b]
t_{NL}	-1.377	-7.240	-1.857***	-9.721*
	(0.18)	(0.000)	(0.09)	(0.000)
z_{tbar}	3.184	-18.816	4.852***	26.891*
	(0.18)	(0.000)	(0.09)	(0.000)

Notes:
a) Regressions include an intercept and linear time trend.
b) Regressions include only intercept.
Optimal lag length in IPS and UO tests were selected using AIC with maximum lag order of 12. *, ** and *** indicate significance at 1%, 5% and 10% significance levels, respectively.

As can be seen from Table 6, the UO and IPS tests have different results with respect to weak form market efficiency. As regard to the IPS test this group of emerging countries failed to reject the null hypothesis of unit root which means that they are efficient as a group. On the other hand, UO test rejected the null hypothesis that this group does not constitute a group of efficient market. This result may be due to the fact that the IPS test has a low power against non-linear stationary process. Hence, linear unit root and the panel unit root tests suggest that these are individually and as a group efficient market where as nonlinear unit root and panel unit root tests suggest that some of them individually efficient but as a group they are seen to be inefficient in weak form sense.

4. Conclusion

In this paper we have tested whether Bulgarian, Greek, Hungarian, Polish, Romanian, Russian, Slovenian and Turkish stock price series contain unit root, consistent with weak form efficiency. For this purpose we carried out conventional ADF and PP unit root tests as well as nonlinear unit root test recently proposed by Kapetanios *et al.* (2003). The results of ADF and PP indicate that Bulgarian, Greek, Hungarian, Polish, Romanian, Russian, Slovenian and Turkish stock price series contain unit root. Using nonlinear unit root test due to Kapetanios *et al*. (2003), we are able to reject the null hypothesis of unit root for Russian, Romanian and Polish stock price series, implying that these markets are not weak form efficient. Moreover, we apply linear and nonlinear panel unit root tests to this group of countries. The linear panel unit root test suggest that this group as all efficient market where as nonlinear panel unit root test suggest as a group they are inefficient in the weak form sense.

The efficient market hypothesis states that security prices fully reflect all available information and that the price fluctuations are unpredictable. Unpredictability of returns is satisfied if stock prices follow a random walk, that is, stock prices are characterized by a unit root. These results show that the markets in this region seem to be weak form efficient in linear sense, however linear test are not taken into consideration of nonlinearities and this can be seen as model misspecification. By applying nonlinear test, first of all we see that the data generating process is nonlinear. With respect to this information, we obtain the true results about the market efficiencies of these region namely emerging markets of Europe. In this respect we make two important contributions to this literature. First, we have taken into account the possible nonlinearities in conditional mean of the series in testing efficiency of these markets which is a deviation from the vast literature. The second one, we have used Ucar and Omay (2010) nonlinear panel unit root test which increase the power of nonlinear unit root test (One way to obtain a more powerful test is to pool the estimates from a number of separate series and then test the pooled values). Furthermore, this is the first time a nonlinear panel unit root test is used in the market efficiency literature.

References

Aguirre, M. S., and Saidi, R. (1998) Evidence of forward discount determinants and volatility behvior, *Journal of Economic Studies*, **25**, pp. 538-552.

Azad, A.S.M. S. and Bashar, H.M.N. O. (2010) Market efficiency and mean-reversion in developing stock markets: Evidence from South Asia, SSRN Working Paper Series No: 1541803

Bajaj, M. and Vijh, A. (1990) Dividend clienteles and the information content of dividend changes, *Journal of Financial Economics*, **26**, pp. 193-219

Bajaj, M. and Vijh, A. (1995) Trading behaviour and the unbiasedness of the market reaction to divided announcements, *Journal of Finance*, **50**, pp. 255-279

Bamber, L. (1987) Unexpected earnings, firm size, and trading volume around quarterly earnings announcements, *The Accounting Review*, **62**, pp. 510-532

Beechey, M., Gruen, D. Vickery, J. (2000) The efficient market hypothesis: a survey, Reserve Bank of Australia, Research Discussion Paper No: 2000-01

Brigham, E. F. and Gapenski, C. G. (1994) *Financial Management: Theory and Practice.* The Dryden Press, U.S.A.

Brigham, E. F. (1995) *Fundamentals of Financial Management.* The Dryden Press, U.S.A.

Brock, W.A. and Hommes, C.H. (1998) Heterogeneous beliefs and routes to chaos in a simple asset pricing model, *Journal of Economic Dynamics and Control*, **22**, pp. 1235-74.

Cajueiro, D.O. and Tabak, B.M. (2006) Testing for predictability in equity returns for European transition markets, *Economic Systems*, **30**, pp. 56-78

Chaudhuri, K. and Wu, Y. (2003) "Random walk versus breaking trend in stock prices: evidence from emerging markets", *Journal of Banking and Finance*, **27**, pp. 575–92.

Copeland, T. (1979) Liquidity changes following stock splits, *Journal of Finance*, **34**, pp. 115-141.

Courtault, J. M., Bru, B., Crepel, P., Lebon, I. and Le Marchand, A. (2000) Louis Bachelier on the centenary of Theorie de la Speculation, *Mathematical Finance*, **10**, pp. 341-353.

Cowles, A. (1933) Can stock market forecasters forecast?, *Econometrica*, **1**, pp. 309-324.

Cowles, A. and Jones, H. (1937) Some a posteriori probabilities in stock market action, *Econometrica*, **5**, pp. 280-294.

Cowles, A. (1944) Stock market forecasting, *Econometrica*, **12**, pp. 206-214.

Dickey, D. A. and Fuller, W. A. (1981) Likelihood ratio statistics for autoregressive time series with a unit root, *Econometrica*, **49**, pp. 1057–72.

Dockery, E. and Vergari, F. (1997) Testing the random walk hypothesis: evidence for the Budapest stock exchange, ***Applied Economics Letters***, **4**, pp. 627-29.

Dumas, B. (1992) Dynamic equilibrium and the real exchange rate in a spatially separated world, *Review of Financial Studies*, **5**, pp. 153–180.

Emerson, R., Hall, S.G. and Zalewska-Mitura, A. (1997) Evolving market efficiency with an application to some Bulgarian shares, *Economics of Planning*, **30**, pp. 75–90.

Fama, E. F. (1970) Efficient capital markets: a review and empirical work, *Journal of Finance*, **25**, pp. 383-417.

Fama, E. F. (1991) Efficient capital markets: II, *Journal of Finance*, **46**, pp. 1575-617.

Francis, J. C. (1991) *Investments: Analysis and Management.* McGraw-Hill Inc., Singapore.

Franses, P.H. and van Dijk, D. (2000) *Non-Linear Time Series Models in Empirical Finance*, Cambridge University Press.

French K. and Roll, R. (1986) Stock return variances: The arrival of information and the reactions of traders, *Journal of Financial Economics*, **17**, pp. 5-26.

Gallagher, L. and Taylor, M.P. (2001) Risky Arbitrage, Limits of Arbitrage and Nonlinear Adjustment in the Dividend-Price Ratio, *Economic Inquiry*, **39**, pp. 524-36.

Grieb, T. A. and Reyes, M. G. (1999) Random walk tests for Latin American equity indexes and individual firms, *Journal of Financial Research*, **22**, pp. 371–83.

Harrison B. and Paton, D. (2004) Transition, the evolution of stock market efficiency and entry into EU: The case of Romania, *Economics of Planning*, **37**, pp. 203-23.

Hasanov, M. ve T. Omay, (2008) "Non-linearities in Emerging Stock Markets: Evidence from Europe's Two Largest Emerging Markets," *Applied Economics*, **40**(23), 2645-2658.

Holthausen, R., Leftwich, R. and Mayers, D. (1987)The effect of large block transactions on security prices: A cross-sectional analysis, *Journal of Financial Economics*, **19**, pp. 237-267.

Hong, H. and Stein, J.C. (1999) A unified theory of underreaction, momentum trading, and overreaction in asset markets, *Journal of Finance*, **54**, pp. 2143–84.

Kapetanios, G., Shin, Y. and Snell, A. (2003), Testing for a unit root in the nonlinear STAR framework, *Journal of Econometrics*, **112**, pp. 359-79.

Kim, O. and Verrecchia, R. (1991a) Trading volume and price reactions to public announcements, *Journal of Accounting Research*, **29**, pp. 302-321.

Kim, O. and Verrecchia, R. (1991b) Market reaction to anticipated announcements, *Journal of Financial Economics*, **30**, pp. 273-309.

Kraus, A., and Stoll, H. (1972) Price impacts of block trading on the New York Stock Exchange, *Journal of Finance*, **27**, pp. 569-588.

Lakonishok, J. and Vermaelen, T. (1990) Anomalous price behavior around repurchase tender offers, *Journal of Finance*, **45**, pp. 455-478.

Liu, X., Song, H. and Romilly, P. (1997) Are Chinese stock markets efficient? A cointegration and causality analysis, *Applied Economics Letters*, **4**, pp. 511–5.

Lo, A.W. and MacKinlay, A. C. (1988) Stock market prices do not follow random walks: evidence from a simple specification test, *Review of Financial Studies*, **1**, pp. 41- 66.

Luukkonen, R., Saikkonen, P. and Teräsvirta, T. (1988) Testing linearity against smooth transition autoregressive models, *Biometrika*, **75**, pp. 491-99.

Magnusson, M. A., and Wydick, B. (2002) How efficient are Africa's emerging stock markets?, *The Journal of Development Studies*, **38**, pp. 141-156.

Malkiel, B. G., (2003) The efficient market hypothesis and its critics, CEPS Working Paper No: 91.

Michael, P., Nobay, R.A., and Peel, D.A. (1997) Transaction costs and nonlinear adjustment in real exchange rates: an empirical investigation, *Journal of Political Economy*, **105**, pp. 862-79.

Mishra, P. K. (2009) Indian capital market – Revisiting market efficiency, *Indian Journal of Capital Markets*, **2** , pp. 30-34.

Mussavian, M. and Dimson, E. (1998) A brief history of market efficiency, *European Financial Management*, **4**, pp. 91-103.

Omay, T. and Kan E. Ö. (2010) "Re-examining the threshold effects in the inflation–growth nexus with cross-sectionally dependent non-linear panel: Evidence from six industrialized economies." *Economic Modelling,* 27(5).

Phillips P.C.B. and Perron, P. (1988) Testing for unit root in time series regression, *Biometrica*, **75**, pp. 335-46.

Rockinger, M. and Urga, G. (2001) A time varying parameter model to test for predictability and integration in stock markets of transition economies, *Journal of Business & Economic Statistics*, **19**, pp. 73-84.

Ross, S. A., Westerfield, R. W., and Jordan, B. D. (2001) *Essentials of Corporate Finance*. McGraw-Hill Inc., U.S.A.

Sarno, L. (2000) Real exchange rate behaviour in high inflation countries: empirical evidence from Turkey, 1980-1997, *Applied Economics Letters*, **7**, pages 285-91.

Sharpe, W. F., Alexander, G. J., and Bailey, J. V. (1999) *Investments.* Prentice Hall, U.S.A.

Shleifer, A. (2000) *Inefficient Markets. An Introduction to Behavioural Finance*. Clarendon Lectures in Economics, Oxford University Press, Oxford.

Seiler, M. J. and Rom, W. (1997) A historical analysis of market efficiency: Do historical returns follow a random walk, *Journal of Financial And Strategic Decisions*, **10**, pp. 49-57.

Taylor, M.P., Peel, D.A., Sarno, L. (2001) Nonlinear mean-reversion in real exchange rates: towards a solution to the purchasing power parity puzzles, *International Economic Review*, **42**, pp. 1015–42.

Teräsvirta, T. (1994) Specification, estimation, and evaluation of smooth transition autoregressive models, *Journal of the American Statistical Association*, **89**, pp. 208-18.

Uçar N. and T. Omay, (2010) "Testing for Unit Root in Non-linear Heterogeneous Panels," *Economics Letters,*104(1),pp.5-8.

Zalewska-Mitura, A. and Hall, S.G. (1999) Examining the first stages of market performance: a test for evolving market efficiency, *Economics Letters*, **64**, pp. 1-12.