Nonlinear Mean Reversion across National Stock Markets: Evidence from Emerging Asian Markets

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Abstract

This paper seeks empirical evidence of nonlinear mean-reversion in *relative* national stock price indices for Emerging Asian countries. It is well known that conventional linear unit root tests suffer from low power against the stationary nonlinear alternative. Implementing the nonlinear unit root tests proposed by Kapetanios, et al. (2003) and Cerrato, et al. (2009) for the relative stock prices of Emerging Asian markets, we find strong evidence of nonlinear mean reversion, whereas linear tests fail to reject the unit root null for most cases. We also report some evidence that stock markets in China and Taiwan are highly localized.

Key Words: Linear Unit Root Test, Nonlinear Unit Root Test, Nonlinear Panel Unit Root Test, International Relative Stock Prices

JEL Classification: C22, G10, G15

Running head: Nonlinear Mean Reversion across Emerging Asian Stock Markets

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I. Introduction

In the field of finance, mean reversion properties of asset prices have been widely investigated to examine the validity of the contrarian investment strategy.[1] Despite extensive studies, empirical evidence on mean reversion of stock prices is still mixed at best.[2] A growing amount of literature has also started investigating mean reversion among international stock price indices. Among others, Kasa (1992) reported cointegrating relations for the national stock indices of five developed countries, while Richards (1995) found no such relations when he used proper critical values.

More recently, Balvers, et al. (2000) employed a seemingly unrelated regression (SUR) technique for stock prices in eighteen developed countries relative to a reference index, such as the U.S. stock index. They reported strong evidence of mean reversion. Similar evidence has been reported by Chaudhuri and Wu (2004) for seventeen emerging equity markets.

It should be noted, however, that their technique is subject to the following problems. First, their SUR estimation imposes a homogeneity assumption that assumes all countries share *identical* speeds of mean reversion. This is a very strong assumption and contradicting to the common wisdom. Second, their panel unit root test may have a serious size distortion problem in presence of cross-section dependence, which was pointed out by Phillips and Sul (2003).

We note that SUR/panel unit root tests are not the only way of improving the power of unit root tests, and take a different approach by implementing a *nonlinear* unit root test proposed by Kapetanios, et al. (2003) for each of nine Emerging Asian countries to improve

[1] If asset prices are mean-reverting, short-selling assets with relatively better performance and buying assets with poor performance in the past may create excess returns. See DeBondt and Thaler (1985).

[2] For example, Fama and French (1988) and Poterba and Summers (1988) found evidence that favors mean-reversion in U.S. stock prices. Yet, many others questioned the validity of mean-reversion on the robustness issue with regards to the choice of sample period (Kim, et al., 1991), the distributional assumptions (Kim, et al., 1991; McQueene, 1992), and small sample bias (Richardson and Stock, 1989; Richardson, 1993).

the power of unit root tests. In addition, we employ a nonlinear panel unit root test recently proposed by Cerrato, et al. (2009). These tests allow different mean reversion rates across countries, thus do not require the homogeneity assumption.

Unlike the conventional *linear* unit root test, their tests allow smooth transition between the stationary regime and the nonstationary regime around the long-run equilibrium value, which can be justified by nonlinear adjustments of financial market variables in presence of fixed transaction costs.[3] Using the Morgan Stanley Capital International (MSCI) stock index data for these countries, we find very strong evidence of nonlinear mean-reversion *across* these countries. We also find that only some, but not all Emerging Asian countries possess nonlinear cointegrating relations with the U.S. stock index as well as the World stock index, which provides less support for the homogeneity assumption of the SUR unit root test.

Similar work has been done by Lim and Liew (2007) and Hasanov (2007), who test the nonlinear mean reversion for *individual* Asian equity prices. It should be noted, however, that they test the unit root null hypothesis for the *nominal* equity *prices* without taking any economic fundamentals (e.g., price-earning ratio or dividend yield) into consideration.[4] Our work is different from theirs, since we test the nonlinear mean reversion for the *relative* prices or stock price deviations from a reference index. When a country shares a fundamental value, possibly a unit root process, with a reference country or index, the stock price deviation from the reference index should be mean-reverting.

To deal with the second issue, we implement an array of panel unit root tests including a newly proposed nonlinear panel unit root test by Cerrato, et al. (2009). We do

[3] For example, in the presence of market frictions or transaction costs, arbitrages occur only when the deviations from the fundamental values are big (see, among others; Dumas, 1992; Michael, et al., 1997). In other words, when the deviations are relatively small, asset prices may exhibit *local nonstationarity* around the long-run equilibrium values in the absence of any arbitrage. When dealing with an aggregate price index, smooth transition model would make more sense, since the transaction costs might be different across the products.

[4] One may test the unit root null for the nominal price deviations from a fundamental variable such as the dividend yield and the price-earning ratio.

find that controlling cross-section dependence substantially lowers rejection probabilities of the panel unit root tests.[5] However we find much stronger evidence of nonlinear mean reversion relative to its linear counterpart irrespective of the treatment of cross-section dependence.

The rest of the paper is organized as follows. Section II describes our baseline linear model of stock indices in two countries. We extend this model to a nonlinear adjustment model in Section III and to a nonlinear panel model in Section IV. Section V reports our main findings. Section VI concludes.

II. The Linear Cointegration Model

We first consider a linear model for the stock markets in two countries, *A* and *B*. Let p_t^i and f_t^i be the log of the stock index and the log of its fundamental value for country *i*, respectively. If p_t^i is mean-reverting around f_t^i, its stochastic process can be represented as the following error correction model,

$$\Delta(p_{t+1}^i - f_{t+1}^i) = a^i + \lambda(p_t^i - f_t^i) + u_{t+1}^i,\ \ i = A, B, \tag{1}$$

where $-1 < \lambda < 0$ is a common convergence rate parameter for *A* and *B*, and u_t^i is an idiosyncratic mean-zero i.i.d. process. The time-varying fundamental term f_t^i, a possible unit root process, is not directly observable but is assumed to obey the following stochastic process,

$$f_t^i = b^i + f_t^c + v_t^i,\ \ i = A, B, \tag{2}$$

where f_t^c is the common component for p_t^A and p_t^B, b^i is a country-specific constant, and v_t^i is an idiosyncratic zero-mean, possibly *serially correlated* stationary process.

Combining (1) and (2), we obtain

[5] Kim (2009) finds similar results for stock price indices of 18 countries with well-developed capital markets.

$$\Delta(p_{t+1}^{A} - p_{t+1}^{B}) = \alpha + \lambda(p_t^A - p_t^B) + \varepsilon_{t+1}, \tag{3}$$

where

$$\alpha = (a^A - a^B) - \lambda(b^A - b^B)$$
$$\varepsilon_{t+1} = (u_{t+1}^A - u_{t+1}^B) + (v_{t+1}^A - v_{t+1}^B) - (1+\lambda)(v_t^A - v_t^B)$$

For notational simplicity, let r_t denote the stock price deviations (or the relative stock price), $p_t^A - p_t^B$. Lagging time subscript by one, we get

$$\Delta r_t = \alpha + \lambda r_{t-1} + \varepsilon_t .$$

Or equivalently,

$$r_t = \alpha + \rho r_{t-1} + \varepsilon_t, \tag{4}$$

where $\rho = 1 + \lambda$ is the persistence parameter of the deviation.

Note that the error term ε_t is serially correlated even when v_t^i is an i.i.d. process. In order to control this serial correlation, we augment the equation (4) as follows:

$$r_t = \alpha + \rho r_{t-1} + \sum_{j=1}^{k} \beta_j \Delta r_{t-j} + e_t, \tag{5}$$

where e_t is a martingale difference sequence that generates ε_t.

Note that the regression equation (5) is a conventional augmented Dickey-Fuller (ADF) regression equation with a known cointegrating vector [1 -1] for the integrated processes p_t^A and p_t^B. When p_t^A and p_t^B share a common unit root process f_t^c in (2), the stock price deviation r_t should be stationary ($0 < \rho < 1$), and the conventional ADF test applies to test such a linear cointegration relation across the stock markets in *A* and *B*.

III. The Nonlinear Cointegration Model

We extend the regression model (5) to a nonlinear cointegration model that allows nonlinear adjustments of the stock price deviation r_t. Stock prices may adjust to its long-run equilibrium only when the deviation is big enough in the presence of fixed transaction cost. Then, r_t may follow a unit root process locally around the long-run equilibrium value, when the transaction cost is prohibitively high. Such a stochastic process can be represented by the following exponential smooth transition autoregressive process. Abstracting from a constant for notational simplicity,

$$r_t = r_{t-1} + \lambda r_{t-1}\left\{1 - \exp\left(-\kappa r_{t-d}^2\right)\right\} + \varepsilon_t, \tag{6}$$

where κ is a strictly positive scale parameter so that $0 < \exp(-\kappa r_{t-d}^2) < 1$, and d is a delay parameter.

Note that when r_{t-d} is very big, put differently stock price indices significantly deviate from each other, $\exp(-\kappa r_{t-d}^2)$ becomes about zero, and the equation (6) reduces to a stationary $AR(1)$ process, where $1 + \lambda = \rho < 1$. On the other hand, if r_{t-d} is close to zero, $\exp(-\kappa r_{t-d}^2)$ is about unity, which leads to a unit root process.

Since λ is not identified under the unit root null hypothesis[6], Kapetanios, et al. (2003) transformed the equation (6) to

$$\Delta r_t = \lambda r_{t-1}\left\{1 - \exp\left(-\kappa r_{t-d}^2\right)\right\} + \varepsilon_t . \tag{7}$$

By the Taylor approximation of (7), they obtained the following equation

$$\Delta r_t = \delta r_{t-d}^3 + \varepsilon_t . \tag{8}$$

They show that, under the unit root null, the least squares t-statistic for δ $(= \hat{\delta} / s.e.(\hat{\delta}))$ has the following asymptotic distribution

$$\frac{\frac{1}{4}W(1)^2 - \frac{3}{2}\int_0^1 W(z)^2\, dz}{\sqrt{\int_0^1 W(z)^6\, dz}}, \tag{9}$$

where *W(z)* is the standard Brownian motion defined on $z \in [0,1]$.

[6] This is the so-called Davies' Problem.

When error terms (ε_t) are serially correlated, the equation (8) can be augmented as follows

$$\Delta r_t = \delta r_{t-d}^3 + \sum_{j=1}^{k} \beta_j \Delta r_{t-j} + e_t . \tag{10}$$

IV. The Nonlinear Panel Cointegration Model

Lastly, we consider a nonlinear panel cointegration test proposed by Cerrato, et al. (2009), which is an extension of Kapetanios, et al. (2003) and Pesaran (2007). Their nonlinear test is more powerful than conventional linear panel unit root tests such as the IPS test by Im, et al. (2003) and can allow for cross-section dependence.

For this purpose, rewrite (7) as follows.

$$\Delta r_{i,t} = \lambda_i r_{i,t-1}\left\{1 - \exp\left(-\kappa r_{i,t-d}^2\right)\right\} + \varepsilon_{i,t} \text{ and } \varepsilon_{i,t} = \delta_i f_t + u_{i,t} , \tag{11}$$

where δ_i is a country specific factor loading, f_t is a common factor, and $u_{i,t}$ is a possibly serially correlated idiosyncratic shock.[7] Cerrato, et al. (2009) suggest the following nonlinear cross-section augmented IPS-type statistics.

$$\bar{t}_{N,T} = N^{-1} \sum_{i=1}^{N} t_i(N,T) , \tag{12}$$

where $t_i(N,T)$ is the *t*-statistic for $\beta_{i,0}$ from the following least squares regression,

$$\Delta r_{i,t} = \alpha_i + \beta_{i,0} r_{i,t-1}^3 + \gamma_{i,0} \bar{r}_{t-1}^3 + \gamma_{i,1} \Delta \bar{r}_t + e_{i,t} \tag{13}$$

$$\Delta r_{i,t} = \alpha_i + \beta_{i,0} r_{i,t-1}^3 + \gamma_{i,0} \bar{r}_{t-1}^3 + \sum_{j=1}^{p} \left(\beta_{i,j} \Delta r_{i,t-j}^3 + \gamma_{i,j} \Delta \bar{r}_{t-j}^3\right) + e_{i,t} , \tag{14}$$

[7] Recall, in Section II, that we construct $r_{i,t}$ as a deviation of individual stock price from its fundamental value. Therefore, f_t can be interpreted as any remaining common shock components that originate from the emerging Asian countries.

for serially uncorrelated error case and for serially correlated error case, respectively, and $\bar{r}_t$ is the cross-section average at time t, which proxies the common factor component for $i = 1, \cdots, N$

Note that, in absence of cross-section dependence, $\gamma_{i,j} = 0$ for all i and j and the test statistic is reduced to nonlinear IPS-type statistic.

V. Empirical Results

We use the monthly data obtained from the Morgan Stanley Capital International (MSCI) for stock market indices of nine Emerging Market (EM) Asian countries, the U.S. stock index, and the World stock index as well as two local reference indices, the EM-Asia and the EM-Far East indices. The data covers the period from December 1987 through December 2007 with the exceptions of China, India, and Pakistan. [8] The observations are end-of-period value-weighted stock prices of many companies in each market. The indices include reinvested gross dividends and are transformed to the U.S. dollar terms using end-of-period foreign exchange rates.

Table 1 presents descriptive statistics for the logarithm of the stock price indices for Emerging Asian countries and reference indices.

>>> Insert Table 1 Here <<<

Following Balvers, et al. (2000), we begin our analysis by implementing the ADF test for the stock price deviations of EM-Asia indices relative to the U.S. stock index and the World stock index. We choose the number of lags (k) by the General-to-Specific Rule (Hall, 1994) as recommended by Ng and Perron (2001) and implement the tests when an intercept is

[8] For these countries, the observations span from December 1992 ending December 2007.

included and when an intercept and time trend are included.[9] As shown in Table 2, the ADF test rejects the unit root null for virtually no country. The only exception was the Taiwan index deviation relative to the World index when trend term is included.

In contrast to the results from the ADF test, our nonlinear unit root test rejects the null of unit root for four countries, Indonesia, Korea, Malaysia, and Pakistan, at the 5% significance level irrespective of the choice of the reference index. When we relax the significance level to 10%, the unit root null is rejected for two more countries, Taiwan and Thailand. Such findings imply that the stock price indices in many EM Asian markets exhibit so-called "coupling" relations with these reference indices in the long-run. Our findings also suggest that there exist nonnegligible sources of market frictions in EM-Asia markets. It is interesting to see that we find strong evidence of mean-reversion for a subset of these countries. This finding implies that the homogeneity assumption by Balvers, et al. (2000) may be problematic.

>>> Insert Table 2 Here <<<

Next, we turn our attention to pairwise unit root tests across EM Asian countries. Again, the linear test hardly rejects the unit root null. The only exception is Korea, where the test rejects the null for a maximum of four out of eight local partners. Surprisingly, the nonlinear test with an intercept rejects the unit root null for eighteen out of thirty-six pairs favoring nonlinear mean reversion. By allowing trend stationarity, we obtain seven additional rejections totaling twenty-five rejections out of thirty-six pairs.

[9] Note that mean-reversion property is more closely related to the ADF test with an intercept only, since rejecting the unit root null from the ADF test with both intercept and time trend implies that the series is trend stationary. Therefore, the ADF test with both deterministic terms should be understood as a supplementary test when the test with an intercept only does not reject the unit root null.

We also consider the cases when a *local* aggregate stock index such as the EM-Asia index or the EM-Far East index serves as a reference index. Again, we obtain very strong evidence of nonlinear mean reversion for the deviations of China, Indonesia, Korea, and Taiwan relative to these local reference indices. It is interesting to see that the stock indices of China and Taiwan exhibit very strong tendencies toward these local indices, whereas they have relatively weak long-run relations with the U.S. stock index and the World stock index. We interpret this as the evidence of *localized* stock markets for those countries.

>>> Insert Table 3 Here <<<

>>> Insert Table 4 Here <<<

Lastly, we implement an array of panel unit root tests with four different reference indices, the U.S. index, the World index, the EM-Asia index, and the EM-Far East index. Results are reported in Table 5.

We first test the null of a unit root with the linear stationarity alternative hypothesis using the IPS panel unit root test. The IPS test fails to reject the null hypothesis for all cases even at the 10% significance level. However, nonlinear IPS-type panel unit root test, based on (13) or (14) with a restriction of $\gamma_{i,j} = 0$, rejects the null of a unit root at the 5% significance level for all cases.

Next, we implement the cross-section augmented IPS-type (CIPS) test by Pesaran (2007) as well as the nonlinear CIPS (NCIPS) test by Cerrato, et al. (2009). It should be noted that allowing for cross-section dependence reduces the probability of rejection of the null hypothesis substantially, which may be consistent with the findings of Phillips and Sul (2003).[10] It should be also noted, however, that the *p*-values of NCIPS statistics are uniformly

[10] This casts doubt on the results of Balvers, et al. (2000) and Chaudhuri and Wu (2004).

less than their linear counterparts. Therefore, we find much stronger evidence in favor of nonlinear mean reversion no matter how we treat cross-section dependence.

>>> Insert Table 5 Here <<<

VI. Concluding Remarks

This paper investigates nonlinear mean reversion across international stock markets using Morgan Stanley Capital International monthly stock index data for nine Emerging Asian countries along with both the global and the local reference indices. As a preliminary analysis, we implement conventional linear unit root tests for the stock price deviations relative to reference indices. The linear test fails to reject the unit root null for most countries. Pairwise tests yield similar results.

As Taylor, et al. (2001) noted, such results may result from a low power problem of the ADF test when the true data generating process is nonlinear transition autoregressive model, which can be theoretically justified by transaction cost arguments. By implementing univariate and panel nonlinear unit root tests, we find strong evidence of mean reversion favoring nonlinear adjustments of stock prices toward the fundamental values. Hence, our results imply that nonnegligible sources of market frictions exist such as strictly positive transaction costs. We also find some evidence of highly localized stock markets for China and Taiwan.

Table 1. Descriptive Statistics for the Stock Price Indices

Country	Mean	Standard Error	Minimum	Maximum	Jarque Bera
China	3.813	0.569	2.818	4.967	9.212 *
India	5.048	0.546	4.374	6.753	13.99 *
Indonesia	5.855	0.729	4.160	7.077	15.22 *
Korea	5.129	0.512	3.698	6.518	8.071 *
Malaysia	5.540	0.451	4.237	6.454	1.035
Pakistan	4.781	0.667	3.581	6.168	4.670
Philippines	5.556	0.591	4.541	6.652	11.20*
Taiwan	5.594	0.301	4.605	6.275	0.350
Thailand	5.511	0.634	4.105	6.672	7.738*
EM-Asia	5.590	0.426	4.605	6.746	2.300
EM-FE	5.550	0.425	4.605	6.607	6.333*
World	7.629	0.500	6.721	8.550	16.34*
USA	7.576	0.694	6.216	8.520	19.15*

Notes: i) Observations span from 1987:12-2007:12 (241 observations) with the exceptions of China, India, and Pakistan (1992:12-2007:12, 181 observations). ii) The superscript * indicates the null hypothesis of normality is rejected at the 5% significance level.

Table 2. Unit Root Test for the Log Stock Price Deviations Relative to Reference Indices

	US Index				*World Index*			
Country	ADF_c	$ADF_{c,t}$	$NLADF_c$	$NLADF_{c,t}$	ADF_c	$ADF_{c,t}$	$NLADF_c$	$NLADF_{c,t}$
China	-1.517	0.328	-1.292	-0.565	-1.512	0.260	-1.410	-0.453
India	-0.306	0.080	-0.514	-0.398	0.188	0.114	0.529	-0.207
Indonesia	-1.525	-0.901	-2.961^{b}	-3.118	-2.001	-2.265	-3.734^{c}	-4.216^{c}
Korea	-2.369	-1.888	-5.058^{c}	-4.799^{c}	-1.918	-1.655	-3.806^{c}	-3.663^{b}
Malaysia	-2.003	-2.407	-4.097^{c}	-3.949^{c}	-2.240	-2.668	-4.086^{c}	-3.966^{c}
Pakistan	-1.464	-1.343	-3.044^{b}	-3.119	-1.349	-1.340	-3.038^{b}	-3.028
Philippines	-1.263	-1.550	-1.977	-2.301	-1.239	-1.894	-1.888	-2.192
Taiwan	-2.199	-2.440	-2.865^{a}	-3.591^{b}	-1.999	-3.405^{a}	-2.787^{a}	-3.661^{b}
Thailand	-1.569	-1.877	-2.757^{a}	-3.592^{b}	-1.632	-2.149	-2.825^{a}	-3.652^{b}

Notes: i) Observations span from 1987:12-2007:12 (241 observations) with the exceptions of China, India, and Pakistan (1992:12-2007:12, 181 observations). ii) The number of lags (k) was chosen by the general-to-specific rule (Hall, 1994). iii) ADF_c and $ADF_{c,t}$ refer the ADF-*t* statistics when an intercept is included and when an intercept and time trend are included, respectively. iv) The NLADF tests were implemented by the Taylor-approximated ESTAR process by Kapetanios, et al. (2003). v) For the NLADF tests, each time series was either demeaned or demeaned and detrended depending on the specifications about deterministic terms. vi) $NLADF_c$ and $NLADF_{c,t}$ refer the nonlinear ADF-*t* statistics when an intercept is included and when an intercept and time trend are included, respectively. vii) The superscripts a, b, and c refer the cases when the unit root null is rejected at the 10%, 5%, and 1% significance levels, respectively. viii) The asymptotic critical values were obtained from Harris (1992) for ADF-*t* statistics and Kapetanios, et al. (2003) for NLADF-*t* statistics.

Table 3. Linear Unit Root Test for the Log Stock Price Deviations across EM-Asia Countries

County		Chi	Ind	Ids	Kor	Mal	Pak	Phi	Tai	Tha
China	ADF_{c}	-	-1.626	-2.225	-2.573^{a}	-2.122	-1.364	-2.565	-2.084	-3.052^{b}
	$ADF_{c,t}$	-	-1.552	-2.466	-3.604^{b}	-2.124	-1.522	-4.149^{c}	-0.042	-3.042
India	ADF_{c}	-1.626	-	-2.062	-2.746^{a}	-1.169	-1.913	-0.506	1.626	-1.281
	$ADF_{c,t}$	-1.552	-	-2.447	-2.832	-2.433	-1.899	-2.567	-0.507	-2.233
Indonesia	ADF_{c}	-2.225	-2.062	-	-2.145	-1.688	-2.361	-1.813	-1.716	-3.008^{b}
	$ADF_{c,t}$	-2.466	-2.447	-	-3.539^{b}	-1.886	-3.122	-1.778	-1.673	-3.005
Korea	ADF_{c}	-2.573^{a}	-2.746^{a}	-2.145	-	-2.558	-3.303^{b}	-1.192	-1.574	-1.199
	$ADF_{c,t}$	-3.604^{b}	-2.832	-3.539^{b}	-	-2.380	-3.291^{a}	-1.731	-1.968	-1.967
Malaysia	ADF_{c}	-2.122	-1.169	-1.688	-2.558	-	-1.427	-1.661	-2.633	-1.526
	$ADF_{c,t}$	-2.124	-2.433	-1.886	-2.380	-	-2.267	-2.175	-2.562	-2.245
Pakistan	ADF_{c}	-1.364	-1.913	-2.361	-3.303^{b}	-1.427	-	-0.936	-0.582	-1.573
	$ADF_{c,t}$	-1.522	-1.899	-3.122	-3.291^{a}	-2.267	-	-2.910	-1.670	-4.144^{c}
Philippines	ADF_{c}	-2.565	-0.506	-1.813	-1.192	-1.661	-0.936	-	-1.756	-1.754
	$ADF_{c,t}$	-4.149^{c}	-2.567	-1.778	-1.731	-2.175	-2.910	-	-1.953	-1.747
Taiwan	ADF_{c}	-2.084	1.626	-1.716	-1.574	-2.633^{a}	-0.582	-1.756	-	-2.107
	$ADF_{c,t}$	-0.042	-0.507	-1.673	-1.968	-2.562	-1.670	-1.953	-	-2.301
Thailand	ADF_{c}	-3.052^{b}	-1.281	-3.008^{b}	-1.199	-1.526	-1.573	-1.754	-2.107	-
	$ADF_{c,t}$	-3.042	-2.233	-3.005	-1.967	-2.245	-4.144^{c}	-1.747	-2.301	-
Local Aggregate Indices										
EM-Asia	ADF_{c}	-2.193	-0.668	-1.961	-2.397	-3.106^{b}	-1.478	-1.190	-2.450	-1.225
	$ADF_{c,t}$	-2.238	-2.915	-2.482	-2.714	-3.135^{a}	-2.275	-2.032	-2.411	-2.285
EM-FE	ADF_{c}	-2.519	-0.758	-1.888	-2.048	-2.943^{b}	-1.429	-1.386	-2.598^{a}	-1.415
	$ADF_{c,t}$	-2.653	-2.988	-2.215	-2.631	-3.005	-2.464	-1.951	-2.530	-2.107

Notes: i) Observations span from 1987:12-2007:12 (241 observations) with the exceptions of China, India, and Pakistan (1992.12-2007:12, 181 observations). ii) The number of lags (k) was chosen by the general-to-specific rule (Hall, 1994). iii) ADF_{c} and $ADF_{c,t}$ refer the ADF-*t* statistics when an intercept is included, and when an intercept and time trend are included, respectively. iv) The superscripts a, b, and c refer the cases when the unit root null is rejected at the 10%, 5%, and 1% significance levels, respectively. v) The asymptotic critical values were obtained from Harris (1992).

Table 4. Nonlinear Unit Root Test for the Log Stock Price Deviations across EM-Asia Countries

		Chi	Ind	Ids	Kor	Mal	Pak	Phi	Tai	Tha
China	$NLADF_c$	-	-2.102	-4.699[c]	-5.547[c]	-3.005[b]	-1.730	-2.786[a]	-3.077[b]	-6.746[c]
	$NLADF_{c,t}$	-	-2.040	-4.806[c]	-6.761[c]	-3.368[a]	-3.474[b]	-4.184[c]	-0.038	-6.739[c]
India	$NLADF_c$	-2.102	-	-2.436	-5.487[c]	-1.664	-2.735[a]	-1.082	2.322	-1.402
	$NLADF_{c,t}$	-2.040	-	-4.827[c]	-5.325[c]	-3.563[b]	-2.777	-2.329	-1.474	-3.654[b]
Indonesia	$NLADF_c$	-4.699[c]	-2.436	-	-3.230[b]	-2.024	-2.179	-2.336	-2.784[a]	-2.962[b]
	$NLADF_{c,t}$	-4.806[c]	-4.827[c]	-	-4.348[c]	-2.084	-4.787[c]	-2.434	-2.961	-2.972
Korea	$NLADF_c$	-5.547[c]	-5.487[c]	-3.230[b]	-	-3.256[b]	-5.615[c]	-2.582	-3.053[b]	-1.882
	$NLADF_{c,t}$	-6.761[c]	-5.325[c]	-4.348[c]	-	-2.919	-5.612[c]	-2.760	-3.341[a]	-1.899
Malaysia	$NLADF_c$	-3.005[b]	-1.664	-2.024	-3.256[b]	-	-1.596	-2.761[a]	-3.630[c]	-2.017
	$NLADF_{c,t}$	-3.368[a]	-3.563[b]	-2.084	-2.919	-	-4.059[c]	-2.627	-3.136[a]	-2.650
Pakistan	$NLADF_c$	-1.730	-2.735[a]	-2.179	-5.615[c]	-1.596	-	-1.427	-1.606	-1.241
	$NLADF_{c,t}$	-3.474[b]	-2.777	-4.787[c]	-5.612[c]	-4.059[c]	-	-3.632[b]	-2.552	-4.219[c]
Philippines	$NLADF_c$	-2.786[a]	-1.082	-2.336	-2.582	-2.761[a]	-1.427	-	-2.038	-2.804[a]
	$NLADF_{c,t}$	-4.184[c]	-2.329	-2.434	-2.760	-2.627	-3.632[b]	-	-2.514	-2.766
Taiwan	$NLADF_c$	-3.077[b]	2.322	-2.784[a]	-3.053[b]	-3.630[c]	-1.606	-2.038	-	-2.999[b]
	$NLADF_{c,t}$	-0.038	-1.474	-2.961	-3.341[a]	-3.136[a]	-2.552	-2.514	-	-3.441[b]
Thailand	$NLADF_c$	-6.746[c]	-1.402	-2.962[b]	-1.882	-2.017	-1.241	-2.804[a]	-2.999[b]	-
	$NLADF_{c,t}$	-6.739[c]	-3.654[b]	-2.972	-1.899	-2.650	-4.219[c]	-2.766	-3.441[b]	-
Local Aggregate Indices										
EM-Asia	$NLADF_c$	-3.225[b]	-0.886	-2.769[a]	-4.700[c]	-3.079[b]	-2.126	-2.318	-3.965[c]	-2.415
	$NLADF_{c,t}$	-3.342[a]	-2.783	-3.921[b]	-4.570[c]	-3.073	-3.117	-2.786	-4.184[c]	-3.475[b]
EM-FE	$NLADF_c$	-4.009[c]	-0.851	-2.562	-4.176[c]	-2.625	-1.893	-2.159	-4.329[c]	-2.493
	$NLADF_{c,t}$	-4.474[c]	-3.017	-3.332[a]	-4.092[c]	-2.577	-3.380[a]	-2.627	-4.229[c]	-3.370[a]

Notes: i) Observations span from 1987:12-2007:12 (241 observations) with the exceptions of China, India, and Pakistan (1992.12-2007:12, 181 observations). ii) The number of lags (k) was chosen by the general-to-specific rule (Hall, 1994) from linear models. iii) The estimations were implemented by the Taylor-approximated ESTAR process by Kapetanios, et al. (2003). iv) Each time series was either demeaned or demeaned and detrended depending on the specifications about deterministic terms. v) $NLADF_c$ and $NLADF_{c,t}$ refer the nonlinear ADF-*t* statistics when an intercept is included and when an intercept and time trend are included, respectively. vi) The superscripts a, b, and c refer the cases when the unit root null is rejected at the 10%, 5%, and 1% significance levels, respectively. vii) The asymptotic critical values were obtained from Kapetanios, et al. (2003).

Table 5. Panel Unit Root Tests for EM-Asia Countries

	Cross-Section Independence	
	IPS	NIPS
US	-1.723 (0.257)	-2.238 (0.016)
World	-1.691 (0.296)	-2.246 (0.015)
EM-Asia	-1.496 (0.565)	-2.250 (0.015)
EM-FE	-1.642 (0.360)	-2.346 (0.005)
	Cross-Section Dependence	
	CIPS	NCIPS
US	-1.946 (0.342)	-2.031 (0.203)
World	-1.896 (0.398)	-1.984 (0.252)
EM-Asia	-1.576 (0.746)	-2.080 (0.161)
EM-FE	-1.551 (0.768)	-1.934 (0.309)

Note: i) Observations span from 1992:12-2007:12 (181 observations for each of N countries). ii) IPS refers the panel unit root test statistics by Im, et al. (2003). iii) NIPS is an average t statistics based on individual $NLADF_c$. iv) CIPS refers to Pesaran's (2007) cross-section augmented IPS test statistic. v) Nonlinear CIPS is Cerrato, et al.'s (2009) nonlinear cross-section augmented IPS test statistic. vi) p-values are in parentheses and obtained from 100,000 Monte Carlo summations with $N = 9$ and $T = 181$.

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