4. REVISITING MEAN REVERSION IN THE STOCK PRICES OF NINE TRANSITION COUNTRIES: THRESHOLD UNIT ROOT TEST

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Abstract

In this study, we use the threshold unit root test proposed by Caner and Hansen (2001) to re-investigate the time-series properties of stock prices for the nine transition countries during the 2000.10 to 2010.11 period. The empirical results from our threshold unit test indicate that the null hypothesis of I(1) unit root in stock prices can not be rejected for any of these transition countries, with the exception of Estonia and Latvia two countries. Our results highlight the efficient market hypothesis does hold in these transition stock markets, with the exception of the Estonia and Latvia two stock markets.

Keywords: Mean Reversion; Stock Prices; Transition Countries; Threshold Unit Test
JEL classification: C14; C22; G12

1. Introduction

Ever since the seminal work of Nelson and Plosser (1982), a plethora of studies have been devoted to investigating the potential non-stationarity of important macroeconomic and/or financial variables. Whether or not stock prices are characterized by a unit root has important implications for the efficient market hypothesis, which asserts that returns of a stock market are unpredictable from previous price changes. If stock prices are an I(0) stationary process, then any shock

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effect is temporary. Thus, shifting the stock price from one level to another will eventually return it to its equilibrium level. From an investment point of view, this ensures that one can forecast future movements in stock prices based on past behavior, and trading strategies can be developed so as to earn abnormal returns.[4] However, if it is found that stock prices are non-stationary (or I(1) process) then shocks will have a permanent effect, implying that stock prices will attain a new equilibrium and future returns cannot be predicted based on historical movements in stock prices. This proposition was termed the efficient market hypothesis (hereafter, EMH). A non-stationary stock price also implies that volatility in stock markets will increase in the long run without bound (Narayan, 2008). Nelson and Plosser (1982) pointed out that whether stock prices are modeled as a trend stationary or as a difference stationary process has important implications *vis-à-vis* modeling, testing, and forecasting. Studies on this issue are of considerable concern to researchers conducting empirical studies and investors alike.

Though numerous studies have found support of a unit root in stock prices, but critics have staunchly contended that drawing such a conclusion may be attributed to the lower power of the conventional unit root tests employed when compared with near-unit-root but stationary alternatives.[5] Enders and Granger (1998) also show that the standard tests for unit root all have lower power in the presence of misspecified dynamics. Sarno (2000) and Taylor and Peel (2000) demonstrate that the adoption of linear stationarity tests is inappropriate for the detection of mean reversion if the true process of the data generation is in fact a stationary non-linear process. The presence of nonlinear mean-reverting adjustment for stock prices has been advanced by recent theoretical developments that emphasize the role of transaction costs. Taylor and Peel (2000), Taylor and Taylor (2004), Juvenal and Taylor (2008) and Lothian and Taylor (2008) have argued that different speeds of adjustment at the disaggregated goods level average up to smooth nonlinearity at the aggregate level. An alternative view is that nonlinearity at the aggregate level is caused by other influences, such as the effects of official government intervention (Menkhof & Taylor, 2007; Reitz & Taylor, 2008) or heterogeneous agents (Kilian & Taylor, 2003). Additionally, the existence of structure changes in stock prices might imply broken deterministic time trends and the result is a nonlinear pattern (Bierens, 1997). It should, therefore, not be unexpected that these shortcomings have seriously called into question many of the earlier findings based on a unit root in stock prices.

The central aim of this study contributes significantly to this field of research because, first of all, we examine evidence for mean reversion in these nine transition countries, using the threshold autoregressive model (hereafter, TAR) and the test statistics suggested by Caner and Hansen (2001). The main advantage of this procedure is that it allows one to simultaneously test for nonlinearities and nonstationarity.

[4] *In other words, if stock prices are mean reverting, then short-selling assets that have performed well and buying assets with relatively poor performance in the past (i.e. contrarian strategies) should provide higher returns (see, Debondt and Thaler (1985); Chan (1998); Richards (1997); Balvers and Wu (2006)).*

[5] *For example, see Chaudhuri and Wu (2003), Grieb and Reyes (1999), Poterba and Summeer(1988), Narayan (2005, 2006, 2008), and Narayan and Smyth (2004).*

Secondly, to the best of our knowledge, this study is the first of its kind to utilize the threshold unit root test for mean reversion in the stock prices of these nine transition countries. We found the transition countries to be an interesting sample to investigate stock market behaviour since they had moved from centrally planned economies towards market driven economies recently. This empirical study contributes to the field of empirical research by determining whether or not the unit root process is characteristic of the stock prices in these nine transition countries.

The remainder of this study is organized as follows. Section 2 presents the data used. Section 3 first describes the methodology employed and then discusses the empirical findings and policy implications. Section 4 summarizes up of the conclusions we draw.

2. Data

The data set consists of weekly stock market indices for nine transition countries: Bulgaria, the Czech Republic, Estonia, Hungary, Latvia, Lithuania, Poland, Romania, and the Russia.[6] The stock market indices for these nine transition countries are the Sofia Stock Exchange Index for Bulgaria, Prague Stock Exchange Index for the Czech Republic, the OMX Tallin Stock Exchange Index for Estonia, the Hungary-DS Market - Price Index for Hungary, the OMX Riga Stock Exchange Index for Latvia, OMX Vilnius Stock Exchange Index for Lithuania, Poland-DS Market – Price Index for Poland, Romania Bet (L) – Price Index for Romania, and Russia RTS Index – Price Index for Russia. Sample periods cover from October 2000 to November 2010. Table 1 reports the summary statistics of the data studied. We find that Poland had the highest average stock market returns of -0.07% and both the Czech Republic and Hungary have the lowest average stock market returns of -0.12% over this sample period. The measures for skewness and excess kurtosis show that the stock market return series are highly leptokurtic and negatively skewed with respect to the normal distribution, indicating that all stock market returns are non-normal. This result is consistent with that of the current literature.

Figure 1 presents the graphs of stock price indices for the selected transition countries. Visual inspection of the stock price index series for these nine markets reveals significant upward and downward trends (due to 2008 global financial crisis) in the data series during this sample period.

3. Methodology and Empirical Results

3.1. Caner and Hansen's (2001) Threshold Unit Root Test

Following the work of Caner and Hansen (2001), we adopt a two regime threshold autoregressive (TAR(k)) model with an autoregressive unit root as follow:

$$\Delta P_t = \theta_1' x_{t-1} I_{\{Z_t<\lambda\}} + \theta_2' x_{t-1} I_{\{Z_t \geq \lambda\}} + e_t , \quad t = 1, \dots , T \qquad [1]$$

[6] In our sample countries, Estonia, Latvia, Lithuania (the three Baltic countries), Poland, the Czech Republic, Hungary (the three Central European countries), Romania, and Bulgaria (Southeastern Europe), and the Russia.

Where P_t is the stock price indices for $t = 1,2....,T$, $x_{t-1} = (P_{t-1}, v_t', \Delta P_{t-1}, \ldots, \Delta P_{t-k})'$, $I_{\{\bullet\}}$ is the indicator function, e_t is an i.i.d. disturbance, $Z_{t-1} = P_{t-1} - P_{t-m}$ is the threshold variable, *m* represents the delay parameter and $1 \le m \le k$, v_t is a vector of exogenous variables including an intercept and possibly a linear time trend. The threshold value λ is unknown and takes the values in the compact interval $\lambda \in \Lambda = [\lambda_1, \lambda_2]$, where λ_1 and λ_2 are selected according to $P(Z_t \le \lambda_1) = 0.15$ and $P(Z_t \le \lambda_2) = 0.85$.[7] The components of θ_1 and θ_2 can be partitioned as follows:

$$\theta_1 = \begin{pmatrix} \rho_1 \\ \beta_1 \\ \alpha_1 \end{pmatrix} \quad \theta_2 = \begin{pmatrix} \rho_2 \\ \beta_2 \\ \alpha_2 \end{pmatrix} \qquad [2]$$

where ρ_1 and ρ_2 are scalar terms. β_1 and β_2 have the same dimensions as v_t, and α_1 and α_2 are *k*-vectors. Thus (ρ_1, ρ_2) are the slope coefficients on P_{t-1}, (β_1, β_2) are the slopes on the deterministic components, and (α_1, α_2) are the slope coefficients on $(\Delta P_{t-1}, \ldots, \Delta P_{t-k})$ in the two regimes.

The threshold effect in Equation [1] has the null hypothesis of $H_0 : \theta_1 = \theta_2$, which is tested using the familiar Wald statistic: $W_T = W_T(\hat{\lambda}) = \sup_{\lambda \in \Lambda} W_T(\lambda)$.[8] The stationarity of the process P_t can be established in two ways. The first is when there is a unit root in both regimes (a complete unit root). Here, the null hypothesis $H_0 : \rho_1 = \rho_2 = 0$ is tested against the unrestricted alternative $H_2 : \rho_1 \ne 0$ or $\rho_2 \ne 0$ using the Wald statistic. This statistic is:

$$R_{2T} = t_1^2 + t_2^2 \qquad [3]$$

Here, t_1 and t_2 are the t ratios for $\hat{\rho}_1$ and $\hat{\rho}_2$ from the least squares estimation. The parameters of ρ_1 and ρ_2 from the Equation [1] will control the regime-dependent unit root process of the stock price. If $\rho_1 = \rho_2 = 0$ holds, then we say that the stock price is I(1) and can be described as having a "unit root." Second, when there is a unit root

[7] According to Andrews (1993), this division provides the optimal trade-off between various relevant factors, which include the power of the test and the ability of the test to detect the presence of a threshold effect.

[8] $W_T = W_T(\hat{\lambda}) = \sup_{\lambda \in \Lambda} W_T(\lambda) = T\left(\frac{\hat{\sigma}_0^2}{\hat{\sigma}^2(\lambda)} - 1\right)$, *where* $\hat{\sigma}_0^2$ *and* $\hat{\sigma}^2$ *are residual variances from least squares estimation of the null linear and TAR models, respectively.*

in only one of the regimes, a case of partial unit root, the alternative hypothesis is of the form, $H_1 : \rho_1 < 0$ and $\rho_2 = 0$, or $\rho_1 = 0$ and $\rho_2 < 0$. However, Caner and Hansen (2001) claim that the two-sided Wald statistic may have less power than a one-sided version of the test. As a result, they propose the following one-sided Wald statistic:

$$R_{1T} = t_1^2 I_{\{\hat{\rho}_1 < 0\}} + t_2^2 I_{\{\hat{\rho}_2 < 0\}} \quad [4]$$

To distinguish between the stationary case given as H_1 and the partial unit root case given as H_2, Caner and Hansen (2001) suggest using individual t statistics t_1 and t_2. If only one of $-t_1$ and $-t_2$ is statistically significant, this will be consistent with the partial unit root case H_2. This means stock price behaves like a "nonstationary process" in one regime; but exhibits a "stationary process" in the other regime, vice versa. Caner and Hansen (2001) show that both tests R_{1T} and R_{2T} will have power against both alternatives.[9] To obtain maximum power from these tests, critical values are generated using bootstrap simulations with 10,000 replications, as suggested by Caner and Hansen (2001).

3.2. Empirical Results

Table 2 presents the country-by-country results for the unit root and stationary tests (i.e., the ADF, P-P and the KPSS). At first sight, the individual unit test statistics seem to show that stock prices are non-stationarity for all transition countries under study. As stated earlier, there is a growing consensus that the stock prices exhibit nonlinearities, and consequently, conventional unit root tests such as the ADF test, have low power in detecting the mean reversion of the stock prices. Therefore, we proceed to test the stock prices by using Caner and Hansen's (2001) TAR unit root test.

First, we use the Wald test W_T to examine whether or not we can reject the linear autoregressive model in favor of a threshold model. The results of the Wald test along with the bootstrap critical values generated at conventional levels of significance are reported in Table 3. The bootstrap *p*-value for threshold variables of the form $Z_{t-1} = P_{t-1} - P_{t-m}$ for delay parameters *m* varies from 1 to 12. Since the parameters *m* is generally unknown, there is no reason to assume the optimal delay parameter will be the same across countries. To circumvent this, Caner and Hansen (2001) suggest making *m* endogenous by selecting the least squares estimate of *m* that minimizes the residual variance. This amounts to selecting *m* at the value that maximizes the W_T statistic. We find that the W_T statistic is maximized for Estonia, Latvia, and Lithuania when $m = 10$, for the Russia when $m = 8$, for Bulgaria and Poland when $m = 5$, for

[9] *As stated by Caner and Hansen (2001) that R_{1T} has more power than that of R_{2T}, here we only report the results of R_{1T} in our study.*

Hungary when $m = 4$, for the Czech Republic when $m = 3$, and for Romania when $m = 2$. Taken together, these results imply strong statistical evidence against the null hypothesis of linearity at least at the 10% significance level in all markets and indicate that simple linear models are inappropriate and the TAR model is our preference.

Next, we explore the threshold unit root properties of stock prices based on the R_{1T} statistic for each delay parameter *m*, ranging from 1 to 12, paying particular attention to the results obtained for our preferred model. The R_{1T} test results, together with the bootstrap critical value at the conventional levels of significance and the bootstrap *p*-value, are reported in Table 4. We are able to reject the unit root null hypothesis only for the markets of Estonia and Latvia at the 10% significance level. Taken together our results provide strong support for the EMH of these transition countries, with the exception of the markets of Estonia and Latvia for which the stock prices are characterized as non-linear stationary.

The one-sided test statistic of the R_{1T}, however, is not able to distinguish the complete and partial unit root in stock prices, we examine further evidence on the unit root hypothesis (partial unit root) by examining the individual t statistics, t_1 and t_2. The results are reported in Table 5. Also, with the exception of the stock markets of Estonia and Latvia, the statistics for both t_1 and t_2 are smaller than the critical value at the 5% level of significance, and this leads us to the conclusion that stock prices in these transition countries are nonlinear processes that are characterized by a unit root process, consistent with the EMH.

Several important policy implications emerge from our study. First, overwhelming evidence in favor of the I(1) unit root hypothesis is found, implying that most of the stock markets in these nine transition countries are characterized by the EMH, with the exception of Estonia and Latvia. Second, it is noteworthy that different conclusions are made for the markets of Estonia and Latvia in which we find both the R_{1T} and individual t statistics (either t_1 or t_2) strongly reject the unit root hypothesis indicating that the stock prices of Estonia and Latvia both exhibit nonlinear stationary behaviour. This result shows that the presence of profitable arbitrage opportunities only exist in the Estonia and Latvia stock prices but not the stock prices of the other 7 transition countries. On the basis of a consensus view that investors tend to have informational advantages in their home markets, say the Estonia (and/or Latvia) stock market, we argue that when favorable news becomes available in the home market, foreign investors raise their valuation by more than domestic investors do because domestic investors naturally have precise information and might have received the market news earlier (see, Brennan & Cao, 1997). Third, our findings suggest that shocks to stock price are not temporary for most of the transition countries' stock markets. This result implies that following a major structural change in these stock markets, stock prices will not return to their original equilibrium over a period of time. The fact that stock prices show I(1) unit root indicates that it should not be possible for the series to forecast future movement in stock prices based on past behavior.

Equally important, the results here are consistent with those of Narayan (2005, 2006) and Munir and Mansur (2009), for these three studies also used the TAR unit root test of Caner and Hansen (2001) and determined that the stock markets of the U.S., Australia and New Zealand, and Malaysia exhibit nonlinear behaviors with a unit root process, respectively. Our results contrast with that of Narayan (2008) which support the notion of stationarity of real stock prices for G-7 when the breaking-trend specifications are introduced in the analysis.

4. Conclusion

In this empirical study, we employ the threshold unit test to reassess the non-stationary properties of stock prices for the nine transition countries over the 2000.10 to 2010.11 period. Our major contribution is that, for the first time in the literature, we test for threshold unit root in the stock prices for nine transition countries. Caner and Hansen's (2001) TAR unit root test indicates that a unit root in stock prices is not rejected for most of the transition countries, with the exception of Estonia and Latvia two countries. Our results indicate that the EMH does hold in these transition countries' stock markets, with the exception of the Estonia and Latvia two markets. These results might cast some doubts about the active investment strategies of international mutual funds among these transition countries.

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References

Andrews, D., 1993.) Tests for parameter instability and structural change with unknown change point. *Econometrica*, *61*, pp.821–856.

Balvers, R. and Y. Wu., 2006. Momentum and mean reversion across national equity markets, *Journal of Empirical Finance. 13*, pp.24-48.

Brennan, M. J. and H. H. Cao., 1997. International portfolio investment flows, *Journal of Finance, 52,* pp.1851-1880.

Bierens, H. J., 1997. Testing the unit root with drift hypothesis against nonlinear trend stationarity, with an application to the U.S. price level and interest rate, *Journal of Econometrics*, 81, pp.29-64

Chan, K. C., 1998. On the contrarian investment strategy. *Journal of Business, 61,* pp.147-163.

Caner, M. and B. E. Hansen., 2001. Threshold autoregression with a unit root. Econometrica, 69, pp.1555–96.

Chaudhuri, K and Y. Wu., 2003. Random walk versus breaking trend in stock prices: Evidence from emerging markets. *Journal of Banking and Finance, 27, pp.575-592.*

Debondt, W. and R. Thaler., 1985. Does the stock market overreact? *Journal of Finance, 40,* pp.793-805.

Enders, W. & Granger, C.W. F., 1998. Unit-Root Tests and Asymmetric Adjustment with an Example using the Term Structure of Interest Rates. *Journal of Business Economics & Statistics*, 16, pp.304-311

Grieb, T.A. and M. G. Reyes., 1999. Random walk tests for Latin American equity indexes and individual firms. *Journal of Financial Research, 22, pp.371-383.*

Juvenal, L. and M. P. Taylor., 2008 Threshold adjustment of deviations from the law of one price, *Studies in Nonlinear Dynamics and Econometrics*: 12(3), Article 8.

Kilian, L. and M. P. Taylor., 2003. Why is it so difficult to beat the random walk forecast of exchange rates? *Journal of International Economics*, 60(1), pp.85-107.

Lothian, J. R. and M. P. Taylor., 2008. Real exchange rates over the past two centuries: How important is the Harrod-Balassa-Samuelson effect? *Economic Journal*, 118 (532), pp.1742-63.

Menkhoff, L. and M. P. Taylor., 2007. The obstinate passion of foreign exchange professionals: Technical analysis, *Journal of Economic Literature*, 45(4), pp.936-972.

Munir, Q. and K. Mansur, 2009. Is Malaysian Stock Market Efficient? Evidence from Threshold Unit Root Tests, *Economic Bulletin,* 29(2), pp.1359-1370.

Narayan, P. K., 2005 Are the Australian and New Zealand Stock Prices nonlinear with a unit root? *Applied Economics. 37,* pp.2161-2166.

Narayan, P. K., 2006. The behaviour of US stock prices: Evidence from a threshold autoregressive model. *Mathematics and Computers in Simulation, 71,* pp.103–108.

Narayan, P.K., 2008..Do shocks to G7 stock prices have a permanent effect? Evidence from panel unit root tests with structural break. *Mathematics and Computers in Simulations, 77,* pp.369–373.

Narayan, P.K. and R. Smyth., 2004 Is South Korea's stock market efficient? *Applied Economics Letters, 11, pp.707-710.*

Nelson, C. and C. Plosser., 1982 Trends and random walks in macroeconomic time Series. *Journal of Monetary Economics, 10,* pp.139-162.

Newey, W. and K. West., 1987 A simple positive semi-definite, heteroskedasticity and autocorrelation consistent covariance matrix. *Econometrica, 55,* pp.703-708.

Potebra, J. M. and L.H. Summers., 1988. Mean reversion in stock prices: evidence and implications. *Journal of Financial Economics, 22, pp.27-59.*

Reitz, S. and M.P. Taylor., 2008. The coordination channel of foreign exchange intervention: A nonlinear microstructure analysis. *European Economic Review*, **52**(1), 55-76.

Richards, A., 1997. Winner-loser reversals in national stock market indices: Can they be explained? *Journal of Finance, 52,* pp.2129-2144.

Sarno, L., 2000. Real exchange rate behavior in the Middle East: a re-examination. *Economics Letters*, 66**,** pp.127-36.

Taylor, M.P. and D.A. Peel., 2000. Non-linear adjustment, long-run equilibrium and exchange rate fundamentals. *Journal of International Money and Finance***,** 19**,** pp.33-53.

Taylor, A. M. and M. P. Taylor., 2004. The purchasing power parity debate. *Journal of Economic Perspectives*, 18(4), pp.135-58.

Table 1

Summary Statistics of the Data: $\Delta \ln P$ (2000.10-2010.11)

Statistic	Bulgaria	Czech	Estonia	Hungary	Latvia	Lithuania	Poland	Romania	Russia
Mean	-0.009	-0.012	-0.009	-0.012	-0.010	-0.009	-0.007	-0.012	-0.010
Median	0.002	0.005	0.003	0.003	0.002	0.003	0.004	0.004	0.007
Maximum	0.294	0.156	0.160	0.149	0.242	0.248	0.229	0.157	0.342
Minimum	-5.832	-7.062	-6.530	-7.003	-5.966	-5.986	-4.809	-8.571	-7.379
Std. Dev.	0.258	0.310	0.287	0.308	0.263	0.263	0.215	0.377	0.327
Skewness	-21.938	-22.430	-22.430	-22.372	-22.188	-22.384	-21.256	-22.411	-21.928
Kurtosis	495.528	510.281	510.352	508.558	503.023	508.979	474.987	509.732	495.198
J-B	5348.629	5673.204	5674.782	5634.804	5512.332	5644.170	4912.665	5660.955	5341.476
Probability	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000
Sum Sq. Dev.	34.904	50.409	43.122	49.652	36.242	36.292	24.245	74.348	55.917

*Notes: 1. *** and ** indicate significance at the 0.01 and 0.05 levels, respectively.*

2. $\Delta \ln P = \ln P_t - \ln P_{t-1}$.

Table 2

Univariate Unit Root Tests: (2000.10-2010.11)

Country	*Levels*			*First Differences*		
	ADF	PP	KPSS	ADF	PP	KPSS
Bulgaria	-1.5265 (3)	-1.6352 (12)	1.6462*** (17)	-9.6558*** (2)	-22.0634*** (11)	0.6161** (12)
Czech	-1.2878 (4)	-1.1408 (8)	2.0730*** (17)	-10.2693*** (3)	-21.3509*** (7)	0.1993 (8)
Estonia	-1.2829 (2)	-1.3493 (9)	1.8647*** (17)	-12.5998*** (1)	-19.3451*** (8)	0.2218 (9)
Hungary	-1.0645 (0)	-1.1374 (8)	1.7895*** (17)	-21.7624*** (0)	-21.7956*** (8)	0.1212 (8)
Latvia	-1.7155 (8)	-1.7405 (7)	1.4042*** (17)	-7.2832*** (7)	-21.1564*** (9)	0.3388 (7)
Lithuania	-1.2427 (3)	-1.1806 (13)	1.8963*** (17)	-8.1930*** (2)	-20.5466*** (12)	0.1779 (13)
Poland	-1.2261 (0)	-1.2779 (8)	2.1310*** (17)	-24.7027*** (0)	-24.6538*** (8)	0.0988 (7)
Romania	-2.3431 (0)	-2.1423 (9)	2.0554*** (17)	-10.5771*** (2)	-22.9353*** (10)	0.4949** (10)
Russia	-1.4416 (0)	-1.4800 (9)	2.3317*** (17)	-21.6767*** (0)	-21.8701*** (8)	0.1308 (9)

*Note: *** and ** indicate significance at the 0.01 and 0.05 levels, respectively. The number in parenthesis indicates the lag order selected based on the recursive t-statistic, as suggested by Perron (1989). The number in the brackets indicates the truncation for the Bartlett Kernel, as suggested by the Newey-West test (1987).*

Table 3

Threshold test

Countries	Wald Statistic	Bootstrap p-value	Optimal delay parameter m	Threshold parameter $\hat{\lambda}$	Number of observations in Regime 1 and its percentage
Bulgaroa	46.927	0.0258	5	0.0451	327(63.87%)
Czech	79.775	0.0010	3	-0.0481	76(14.84%)
Estonia	66.759	0.0000	10	-0.0911	76(14.84%)
Hungary	40.794	0.1011	4	-0.0575	76(14.84%)
Latvia	117.930	0.000	10	0.1224	434(84.76%)
Lithuania	45.444	0.032	10	0.1623	434(84.76%)
Poland	84.986	0.000	5	-0.085	76(14.84%)
Romania	56.001	0.005	2	-0.041	76(14.84%)
Russia	73.961	0.000	8	-0.079	86(16.79%)

Note: Following much of the existing empirical literature on monthly stock prices, we set a maximum lag of 12 and base all our bootstrap tests on 10,000 replications. All of the statistics are significant, which supports the presence of threshold effects

Table 4

One sided unit root tests

Countries	Optimal delay parameter m	R_{1T} Statistic	Bootstrap critical values			Bootstrap p-value
			10%	5%	1%	
Bulgaria	5	3.983	9.575	11.959	17.125	0.484
Czech	3	4.725	9.212	11.294	17.006	0.382
Eestonia	10	10.758	9.396	11.555	16.934	0.064
Hungary	4	1;366	9.331	11/587	17.028	0.808
Latvia	10	17.775	9.960	12.428	18.719	0.011
Lithuania	10	3.569	9.955	12.419	18.256	0.541
Poland	5	0.959	9.309	11.542	16.521	0.866
Romania	2	6.572	9.286	11.581	16.301	0.229
Russia	8	8.327	9.559	11.845	17.245	0.144

Table 5

Partial unit root results

Countries	Optimal delay parameter m	t_1^2 statistic	Bootstrap p-value	t_2^2 Statistic	Bootstrap p-value
Bulgaria	5	1.131	0.487	1.644	0.314
Czech	3	2.168	0.156	0.147	0.776
Estonia	10	3.257	0.024	0.386	0.718
Hungary	4	0.471	0.687	1.069	0.501
Latvia	10	2.222	0.164	3.583	0.015
Lithuania	10	0.651	0.648	1.774	0.276
Poland	5	0.522	0.682	0.828	0.582
Romania	2	2.338	0.129	1.051	0.501
Russia	8	2.774	0.116	0.795	0.600

Figure 1

Plots of Stock Price Indices for the Nine Transition Countries