5. Asymmetric Interaction between Stock Price Index and Exchange Rates: Empirical Evidence for Romania

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Abstract

This paper examines the interaction between the stock market and the foreign exchange market of Romania. This relationship has important implications from the viewpoint of the two economic theories - the traditional theory and the portfolio balance theory - and especially in the light of the increasing openness of the economy to international trade and investment.

The data set, from March 2000 to March 2014, covers different market phases and stock market crashes, such as the recent global financial crisis and the Euro Area debt crisis.

Due to the nonlinear nature of the relation between the variables, the study employs a threshold error-correction model based on two distinct regimes extended to incorporate asymmetries related to short-term good or bad news from the two markets.

Within this framework, the empirical evidence shows that there is a long-run equilibrium between the two variables during the time period investigated. There are also short-run non-linear relationships sensitive to short-term good or bad news in the regime with fewer observations, called 'extreme regime'.

Keywords: exchange rates, stock prices, causality, non-linearity, asymmetric threshold model

JEL Classification: C32; G15; F21

I. Introduction

The relation between stock and currency markets was the topic of many theoretical and empirical analyses over the past decades. The classical economic theory has two approaches to this problem - the traditional and the portfolio balance. The traditional approach is built on the hypothesis that exchange rates cause movements in stock prices through the international trading effect. It is based on the idea that movements in the exchange rate influence domestic firms through the costs of their imports and

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exports. Domestic currency depreciation leads to increase in competitiveness of exports and the cost of imports and cause positive or negative effect depending on the specific orientation to export or import dependency of local firms and directly influences the stock prices of export-oriented firms.

On the contrary, the portfolio balance theory postulates that movements in stock prices can determine movements in exchange rates via capital account transactions. If considerable quantities of foreign capital enter or leave the stock market, indirect influence on exchange rate could be observed. Changes in stock prices determine the position of investors regarding domestic assets and, thus, causes movements in exchange rate.

The empirical results are mixed in terms of the causal direction between currency and stock prices or the sign of their correlation. For example, cointegration and causality studies by Bahmani-Oskooee and Sohrabian (1992) demonstrate a bi-directional relation only in the short run; Granger *et al*. (2000) found evidence of dual causality between the exchange rate and stock. Ma and Kao (1990) found exchange rates driving stock prices, Yu and Nieh (2009) also argue for the traditional approach in Taiwan in the long run. Diamandis and Drakos (2011) used a VECM model and found that stock indexes and exchange rates are positively related in Brazil, Argentina, Chile, and Mexico for monthly data; Tsai (2012), using a quantile regression model to provide more details of the relationship, found negative relationships between the stock index and exchange rate in Malaysia, Singapore, South Korea, the Philippines, Taiwan, and Thailand, while Zhao (2010) did not find a long-term equilibrium relationship between real effective exchange rate and stock price.

Other authors find a very week correlation between the two variables. Ravazzolo and Phylaktis (2000) show that the financial crisis had a temporary effect on the long-run co-movement between the various markets.

Horobets and Ilie (2007) found unilateral or bi-lateral causality from the stock prices to exchange rates in Romania depending on the sub-period used and unilateral causality for the entire period1999 to 2007.

The explanations for contrasting results in the empirical studies can be found in the idea that not all the time and in all the countries the portfolio effect is present. If the amount of foreign capital that enters or leaves the stock market is significant, the influence on exchange rate causing appreciation or depreciation of the domestic currency should be observed.

Literature on the study of the relationship refers to three methodologies that relate first to a flow-oriented model, and then a portfolio balance approach and, finally, a cointegration and causality approach.

Many of the papers consider a linear relation between exchange rate and stock prices, but there are also studies which investigate the nonlinear relation. Yu and Nieh, (2009) used threshold cointegration and found long-term equilibrium and asymmetric relationship in Taiwan and Japan.

Some various non-parametric methods for nonlinear Granger causality are developed by Baek and Brock (1992) using correlation integral between time series, by Hiemstra and Jones (1994), by Diks and Panchenko (2005) which show limitations of the

Hiemstra and Jones test in large sample due to ignoring variations in conditional distributions.

The aim of this paper is to test the relation between the stock price in Romania and the Romanian Leu against Euro with a nonlinear cointegration model. The nonlinearity was taken into account by estimating a two-regime Threshold Error-Correction model (TAR-ECM), where the threshold variable is the deviation from the long-run equilibrium between the two variables. Also, a Momentum Threshold Error-Correction model (MTAR-ECM) is estimated taking as threshold the previous period change in the deviation from the long-run equilibrium.

The rest of the paper is organized as follows. Section 2 introduces the data, Section 3 explains the structure of the TAR-ECM and MTAR-ECM models employed in the investigation, and presents the empirical findings of the paper. Finally, the paper presents conclusions and policy proposals.

II. Data

This study used monthly data of the stock and foreign exchange markets from March 2000 to March 2014 (169 observations) considering that there is more fluctuation in daily data. I used the Leu/Euro monthly average of the Romanian Stock Exchange Index and *BET* Index prices in Lei. Thus, an increase in exchange rate means that the value of the domestic currency depreciates.

Table 1

Summary Statistics

	BET	Ex
Mean	4397.190	3.697774
Std. Dev.	2590.941	0.677604
Maximum	10207.09	4.562370
Minimum	505.2986	1.845948
Skewness	0.112563	-0.969609
Kurtosis	2.160343	3.280160
Jarque-Bera Test	5.321427	27.03334

This specific period is chosen due to data availability and to the fact that it contains the last financial crisis beginning in 2008, which supports the idea that the relationship should be nonlinear.

The historical time series of stock price index and exchange rate are shown in Figures 1 and 2.

Figure 1

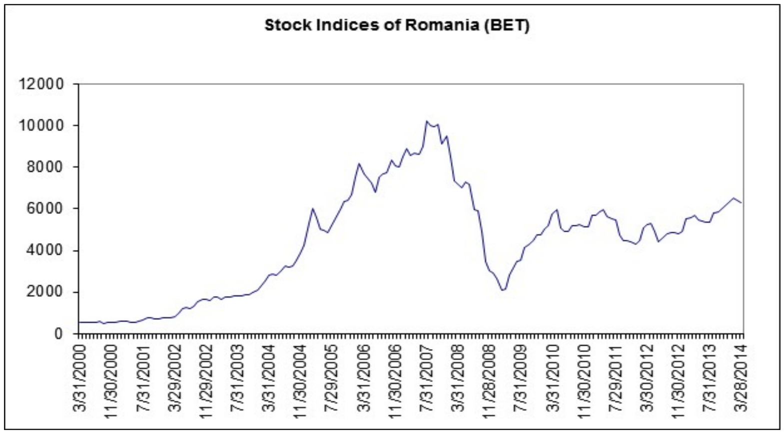
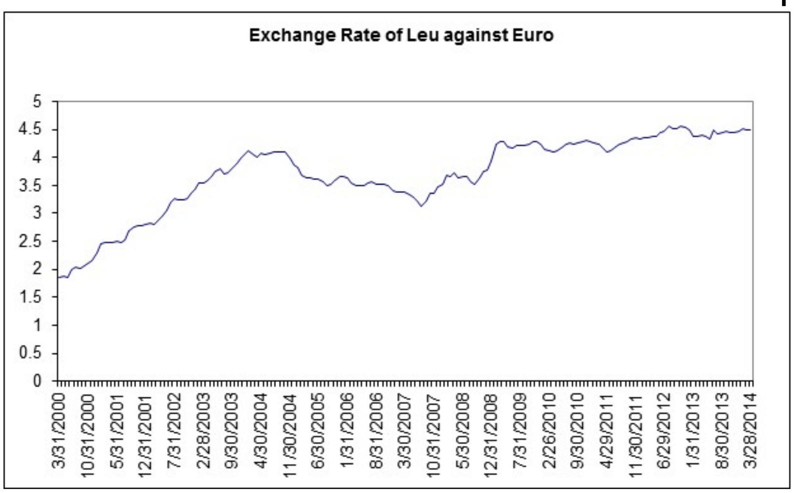


Figure 2



In Figure 1 we notice that stock indices experienced a mild increase until the beginning of 2004, followed by an upward trend until October 2007, and by a sharp decrease until March 2009, when a recovery commenced, followed by an almost stationary price until the end of the period. Correspondingly, the movements in exchange rate (Figure 2) are negatively correlated with changes in the stock price in the period with sharp changes (from the beginning of 2004 until March 2009) and reveal almost stationary levels until the end. Although these two series seem to be negatively related, there are some periods wherein they show positive co-movement. Table 1 presents the summary descriptive statistics of data.

III. Methodologies and Empirical Results

The paper empirically investigates the asymmetric causal relationships between the Leu/Euro exchange rate and the stock price of Romania *BET* using nonlinear models for two-regime threshold cointegration and vector error-correction of Hansen and Seo (2002).

III.1. Unit Root Tests

Table 2 summarizes the results of three linear unit root tests (ADF, NP and KPSS) and the non-linear stationary test proposed by Kapetanios *et al*. (2003) (denoted by KSS) to determine whether the variables are stationary. For the ADF, NP and KSS tests the null hypothesis of non-stationarity cannot be rejected for levels of series, but after apply the difference operator the null is rejected at least at the 5% level of significance. The null for KPSS is I(0), which is rejected for levels of the variables. The conclusion is that all variables are I(1) type series.

Table 2

The Results of Unit Root Tests

		BET	EXU
ADF	Level	-0.067[1]	1.511[1]
	Difference	-9.231[1]***	-8.206[0]***
NP	Level	-0.852[5]	0.702[1]
	Difference	-6.130[0]***	-6.102[0]***
KPSS	Level	0.748[10]***	1.194[10]***
	Difference	0.131[5]	0.338[5]
Non-linear KSS-test	Level	-2.168[1]	-1.910[0]
	Difference	-3.764[2]**	-5.869[0]***

Notes: 1. BET and EXU are the symbols for BET Index and RON/EURO exchange rate

*2. *** and ** denotes significance at 1% and 5% level.*

3. The numbers in brackets are the appropriate lag-length by AIC information criteria for ADF, NP and KSS, whereas for KPSS are the optimal bandwidth decided by Bartlett kernel of Newey-West.

4. The test statistic of NP is MZt.

III.2. Long-run Equilibrium (Cointegration) Tests

Let x_t be a p-dimensional I(1) time series cointegrated with the vector β and $w_{t-1}(\beta) = \beta' x$ denoting the I(0) error-correction term. According to Engle and Granger (1987), two series integrated in the order d, I(d), are cointegrated, if the linear combination of the two series $w_{t-1}(\beta) = \beta' x$ is stationary in less than order d.

A linear error-correction model is as follows:

$$\Delta x_t = \theta' X_{t-1} + e_t \qquad [1]$$

where: $X_{t-1}(\beta) = (1, w_{t-1}(\beta)_t, \Delta x_{t-1}, \ldots, \Delta x_{t-l})'$, *l* represents the optimal lag orders of endogenous variables determined by the selection criterion, and e_t is the vector of innovations.

In the presence of nonlinearities, the linear conventional methodologies determine inefficient estimations and inference tests with low power. One solution to deal with nonlinear data generation process is the threshold cointegration introduced by Balke and Fomby (1997), and further developed by Enders and Granger (1998), Enders and Siklos (2001), Hansen and Seo (2002).

The two-regime threshold cointegration model elaborated by Hansen and Seo is a two-stage estimation procedure that formulates a two-regime threshold autoregressive TAR model for the residuals of the first stage equation:

$$\Delta x_t = \theta_1' X_{t-1} I_{\{Z_t < \lambda\}} + \theta_2' X_{t-1} I_{\{Z_t \geq \lambda\}} + e_t, \ t = 1,2....,T, \qquad [2]$$

where: x_t is a p-dimensional I(1) time series for $t = 1,2....,T$, $X_{t-1}(\beta) = (1, w_{t-1}(\beta)_t, \Delta x_{t-1}, \ldots, \Delta x_{t-l})'$, $I_{\{\bullet\}}$ is the Heaviside indicator function, e_t is an i.i.d. disturbance, $w_{t-1}(\beta) = \beta' x$ denoting the I(0) error-correction term is also the threshold variable. $Z_{t-1} = w_{t-1}(\beta)$ The threshold value λ is unknown and takes on the values in the compact interval $\lambda \in \Lambda = [\lambda_1, \lambda_2]$, where λ_1 and λ_2 are selected according to $P(Z_t \leq \lambda_1) = \pi_0$ and $P(Z_t \leq \lambda_2) = 1 - \pi_0$.

Equation [2] indicates that the two markets follow regime 1 when the threshold variable exceeds or is equal to the threshold value; otherwise they follow regime 2. If there is no significant difference between the estimated parameters, i.e. $\theta_1 = \theta_2$, the threshold model collapses into a linear model as in equation [1]. The model allows us to estimate the regime-dependent parameters and the threshold value (λ) endogenously.

The proposed estimation procedure is maximum likelihood under the assumption that the errors are i.i.d Gaussian. For p = 2, a grid search over the two-dimensional space (β, λ) was chosen by Hansen and Seo as a suitable algorithm for maximization that is not smooth.

A second model MTAR is estimated in this paper replacing the threshold variable $Z_{t-1} = w_{t-1}(\beta)$ from equation [1] with the decay depending on previous period change in the error-correction term $Z_{t-1} = \Delta w_{t-1}(\beta)$. According to Enders and Granger this model can reveal the asymmetric adjustment with more ‘momentum’ in one direction.

The test for the presence of threshold is *SupLM* test (proposed by Davies, 1987) whose asymptotic distribution depends on the covariance of the data (Hansen and Seo, 2002).

The null hypothesis (there is no threshold, the model is linear *H0:* $\theta_1 = \theta_2$) is tested against the hypothesis of two-regime threshold model (1). [2]

$$SupLM = \sup_{\lambda \in \Lambda = [\lambda_1, \lambda_2]} LM(\tilde{\beta}, \lambda) \quad [3]$$

where: $\tilde{\beta}$ is the estimation of beta under the null.

In this paper, the trimming parameter π_0 is set 0.05 and 0.15. [3]

III.3. Validity of Threshold Error-Correction Models

In this paper, p = 2 and $x_t = (BET, EXU)$. To facilitate the calculus, *BET* is chosen as stock index *BET* prices divided by 1000.

Since all the variables are non-stationary at level, I further tested for linear cointegration by means of Johansen cointegration. Table 3 shows these tests, which proved the existence of long-term equilibrium relationship between Leu against Euro (*EXU*) nominal exchange rate and stock prices (*BET*).

Table 3

Johansen Cointegration Test

Unrestricted Cointegration Rank Test (Trace)				
Hypothesized No. of CE(s)	Eigenvalue	Trace Statistic	0.05 Critical Value	Prob.
None *	0.095258	17.13259	12.32090	0.0073
At most 1	0.002482	0.415044	4.129906	0.5829
Trace test indicates 1 cointegrating eqn(s) at the 0.05 level				
Unrestricted Cointegration Rank Test (Maximum Eigenvalue)				
Hypothesized No. of CE(s)	Eigenvalue	Max-Eigen Statistic	0.05 Critical Value	Prob.
None *	0.095258	16.71755	11.22480	0.0050
At most 1	0.002482	0.415044	4.129906	0.5829
Max-eigenvalue test indicates 1 cointegrating eqn(s) at the 0.05 level				

Tables 4, 5 and 6 in the Appendix present the results of the estimation of three types of models: linear, threshold autoregressive (TAR) and threshold autoregressive with

2 Also for maximization, a grid search over the interval for λ is performed in Hansen and Seo (2002).

[3] For π_0 between 0.05 and 0.15 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 (Andrews, 1993; Andrews and Ploberger, 1994).)

momentum (M-TAR) error-correction models. The parameters estimates were calculated over a 300x300 grid on the parameters (β, λ) and for trimming parameter $\pi_0 = 0.05$ and $\pi_0 = 0.15$.

Based on principle of parsimony, the Akaike informaton criterion (AIC) and Bayesian information criterion (BIC), the optimal number of lags is 1 for both TAR and M-TAR models with $\pi_0 = 0.05$.

The estimated threshold for TAR-ECM with lag=1 is $\lambda = 13.1$. The first regime occurs when $BET_t + 1.4923EXU_t \leq 13.1$, for almost 90% of the observations and we can consider this as typical regime. The second regime, with only 10% of the observations, can be considered the extreme regime and occurs when $BET_t + 1.4923EXU_t > 13.1$. The response of the returns to the error correction term is more pronounced in the second regime (Figure 4).

Similar partitions of the observations (90% for $BET_t + 1.439EXU_t \geq 12.54$ and 10% for $BET_t + 1.439EXU_t > 12.54$) are made for the TAR-ECM with lag = 2 and threshold value $\lambda = 12.54$.

The extreme regimes belong to 2006m10-2008m1 and 2006m7-2008m5 for lag 1 TAR and lag 1 MTAR models, respectively, which corresponds to the pre-crisis period and also incorporates the moment of becoming member of the EU.

The *SupLM* test statistics is calculated and the fixed regressor bootstrap and the parametric residual bootstrap algorithms as in Hansen and Seo (2002) are applied to approximate the sampling distribution that depends on the data, and finds strong evidence for a threshold effect significant at 5.14% and 2.26%, respectively, for TAR-ECM models with lag 1 and 2 (bootstrap p-values), respectively.

For the MTAR models, the period that contains the 'extreme regime' is included in 2005m1-2007m3. The estimated threshold for MTAR-ECM with lag = 1 is $\lambda = 0.455$ and the 'typical' regime is present in 90% of observations when $\Delta(BET_t - 1.35238EXU_t) \leq 0.455$. The second regime with only 10% of the observations can be considered the extreme regime and occurs when $\Delta(BET_t - 1.35238EXU_t) > 0.455$.

Similar partitions of the observations (87% when $\Delta(BET_t - 1.04698EXU_t) \leq 0.33$ and 13% in the rest of the cases) are made for the MTAR-ECM with lag = 2 and threshold value $\lambda = 0.33$.

The *SupLM* test with 300 grid points and p_values calculated by parametric residual bootstrap computed with 5000 simulation replications as in Hansen and Seo (2002) is used to assess the evidence for threshold cointegration.

The p_values and the statistics for the selected models (TAR and MTAR with one lag) are presented in Table 5S and Table 6S in the Appendix.

The empirical results give evidence for threshold cointegration in all TAR and MTAR models with 5% significance level based on two methods for approximating asymptotic distribution of the test (Fixed Regressor bootstrap and parametric residual bootstrap), except for the Fixed Regressor (Asymptotic) P-Value of 14% calculated for TAR lag = 1 model. However, the bootstrap method is more appropriate to this test because it has no formal theory, so that the conclusion is that the hypothesis of purely autoregressive model against threshold model is rejected.

The fact that the values of the tests for Equality of Error-Correction (EC) Coefficient are significant with 0.01% (or TAR_ECM with lag = 1 and for M-TAR_ECM with lag = 1), gives evidence of asymmetric adjustment in both TAR and M_TAR error correction models for *BET* and *EXU*. The validity of the models with two regimes is asserted by the p-value of the Test for Equality of Dynamic Coefficients.

III.4. Granger Causality Tests

Given the threshold error correction model as follows:

$$\Delta x_t = (a^1 + b^1 Z_{t-1} + \sum_{i=1}^{k1} \delta_i^{\,1} \Delta x_{1t-i} + \sum_{i=1}^{k2} \gamma_i^{\,1} \Delta x_{1t-i}) I_{\{Z_t<\lambda\}} +$$

$$(a^2 + b^2 w_{t-1} + \sum_{i=1}^{k1} \delta_i^{\,2} \Delta x_{1t-i} + \sum_{i=1}^{k2} \gamma_i^{\,2} \Delta x_{1t-i}) I_{\{Z_t\geq\lambda\}} + e_t$$

where: the upper indices denote the regime, $x_t = (BET, EXU)$, and $Z_{t-1} = w_{t-1}(\beta)$ for TAR-ECM and $Z_{t-1} = Diff(w_{t-1}(\beta))$ for M-TAR-ECM is the error-correction term.

The Granger Causality Tests examine whether all the coefficients of Δx_{1t-i} or Δx_{2t-i} are jointly statistical different from zero.

Thus, we tested the rejection of the null H0: $\delta_i^{\,r} = 0, \quad for \quad all \quad i = 1,..,lag$ to see if *BET* Granger-cause *EXU* for regime r and the null H0: $\gamma_i^{\,r} = 0, \quad for \quad all \quad i = 1,..,lag$ to see if *EXU* Granger-cause *BET* for regime r in the short-run.

Similarly, based on the null H0: $\lambda = \delta_i^{\,r} = 0, \quad for \quad all \quad i = 1,..,lag$, respectively H0: $\lambda = \delta_i^{\,r} = 0, \quad for \quad all \quad i = 1,..,lag$ we may see if *BET* Granger-cause *EXU* for regime r or *EXU* Granger-cause *BET* for regime r in the long run.

The empirical results show evidence of Granger causality in all cases (short run and long run) for both TAR-ECM and MTAR-ECM models with lag = 1 with significance level 5%, except for the case of short run causality from *BET* to *EXU* in MTAR-ECM model regime 1, which rejects the null at only 12%.

To observe the portfolio balance effect, which shows that the stock market leads the foreign exchange market in a negative way, we should see the ECM relationships between the returns of stock price index as the explanatory variable, and that of exchange rate as the dependent variable. The coefficients that stand for this relationship (δ_1) in TAR-ECM and MTAR-ECM are all significantly negative, except for the second regime (10% of observations) in the TAR-ECM model. When presenting the portfolio effect, it means that the increase (decrease) in the returns of stock price index will cause the appreciation (depreciation) of the domestic currency.

The results show that the portfolio effect is present almost all time, except in the period 2006m10-2008m1 (see Figure 3) that is characterized by a more volatile stock market. In times of great volatility and low level of absorption of foreign capital in the stock market, the portfolio effect may not occur.

The effect of exchange rate shocks on stock returns (coefficient γ_1) was proved to be negative in both regimes (Table 5), showing that the international trading effect was not positive. The traditional approach claims that domestic currency depreciation leads to increase in the competitiveness of exports and the cost of imports and causes positive or negative effect depending on the specific orientation to export or import dependency of local firms and directly influences the stock prices of export-oriented firms. The negative relation shows the dependence of local firms on imports and proves that exporting firms did not profit from exchange rate depreciation, probably because the volatility was not significant.

III.5. Threshold Error-Correction Models Sensitive to Short-Term Good or Bad News

We further investigate the sensitivity of the models to short-term good or bad news. In order to disentangle the asymmetric effect, we partition the returns from stock prices and exchange rate into positive (good news) and negative (bad news) parts:

$$\Delta x^{+}{}_{t} = (I_t \Delta BET, I_t \Delta EXU)$$

$$\Delta x^{-}{}_{t} = [(1-I_t)\Delta BET, (1-I_t)\Delta EXU],$$

where: $x_t = (BET, EXU)$, $I_t = 1$ if ΔBET respectively ΔEXU are positive.

The model is a threshold error-correction model extended to incorporate positive and negative movements in prices.

$$\Delta x_t = (a^1 + b^1 Z_{t-1} + \delta_1^1 \Delta x_{1t-1}^{+} + \delta_2^1 \Delta x_{1t-1}^{-} + \sum_{i=2}^{k1} \delta_{i+1}^1 \Delta x_{1t-i} + \gamma_1^1 \Delta x_{2t-1}^{+} + \gamma_2^1 \Delta x_{2t-1}^{-} + \sum_{i=2}^{k2} \gamma_{i+1}{}^1 \Delta x_{2t-i}) I_{\{Z_t < \lambda\}} +$$

$$+ (a^2 + b^2 w_{t-1} + \delta_2^2 \Delta x_{1t-1}^{+} + \delta_2^2 \Delta x_{1t-1}^{-} + \sum_{i=2}^{k1} \delta_{i+1}^2 \Delta x_{1t-i} + \gamma 2 \Delta x_{2t-1}^{+} + \gamma_2^2 \Delta x_{2t-1}^{-} + \sum_{i=2}^{k2} \gamma_{i+1}{}^1 \Delta x_{2t-i}) I_{\{Z_t \geq \lambda\}} + e_t$$

where the upper indices denote the regime, and $Z_{t-1} = w_{t-1}(\beta)$ for TAR-ECM and $Z_{t-1} = Diff(w_{t-1}(\beta))$ for MTAR-ECM is the error-correction term.

The TAR-ECM model with asymmetric effect of good or bad news selected based on AIC and BIC is the model with one lag, which has two regimes (rejecting the hypothesis of equality of Error-Correction coefficients at 3% and the null of equality of Dynamic coefficients at 1% significance level) and the presence of the threshold effect is empirically demonstrated with 6% (*SupLM* test).

The two regimes partition the data into two sets of 94% and 6%, whether $\Delta(BET_t - 0.1264 EXU_t) \leq 8.094$ or $\Delta(BET_t - 0.1264 EXU_t) > 8.094$.

The tests statistics (Table 8S, Appendix) rejects the hypothesis of no Granger causality in the short and long run, except for the short-run causality from BET to exchange rate in the longest regime (p-value=12%).

The asymmetric adjustment to the positive and negative movements in stock prices (good and bad news) can be tested with the null H0: $\delta_1 = \delta_2$ against the alternative H1: $\delta_1 \neq \delta_2$ and the corresponding test for adjustment of the return in stock prices to positive and negative changes in exchange rate.

The estimation of TAR-ECM with good/bad news showed asymmetric adjustment only in second ('extreme') regime.

The MTAR-ECM model with asymmetric effect of good or bad news selected based on AIC and BIC is the model with one lag, which has two regimes (rejecting the hypothesis of equality of Error-Correction coefficients and the null of equality of Dynamic coefficients at 1% significance level and the presence of threshold effect is empirically demonstrated with 5% by *SupLM* test).

The two regimes separate the observations into 90% and 10%, whether $\Delta(BET_t - 1.79418EXU_t) \leq 0.454$ or $\Delta(BET_t - 1.79418EXU_t) > 0.454$.

The test statistics (Table 8S, Appendix) reject the hypothesis of no Granger causality in short and long run, except for the short-run causality from BET to exchange rate in the longest regime.

The estimation of MTAR-ECM with good/bad news showed asymmetric adjustment only in the second ('extreme') regime.

The asymmetric adjustment to the positive and negative movements in stock prices and exchange rate (good and bad news) is found only in the 'extreme' regime that occurs in the 2006-2008 period. In the typical regime, the symmetric adjustment cannot be rejected at 10% significance.

IV. Conclusions

The paper empirically demonstrates that there is a nonlinear two-regime threshold autoregressive relationship between the Romanian stock price and the exchange rate, with asymmetric adjustment. This result proves a co-movement of the two markets and supports the long-run equilibrium relationship between the two prices. Within this framework, the empirical evidence shows that there is a long-run equilibrium between the two variables during the investigated period. There is also a short-run non-linear relationships sensitive to short-term good or bad news in the regime with fewer observations that is called 'extreme regime'.

The selection of the VAR lag based on AIC and BIC criterion consistently choose lag = 1 across all specifications. However, we report results for both lag = 1 and lag = 2 for robustness.

In this paper, we found bi-directional causality between stock price (*BET*) and the Leu/Euro exchange rate in the long-run and in the short-run, with one exception, namely the case of short run causality from *BET* to *EXU* in the MTAR-ECM model regime 1

(87% of the observations). From the theoretical point of view, this validates both the traditional and portfolio theories.

The best model with asymmetries related to short-term good or bad news proved to be the TAR-ECM model with one lag, which shows asymmetric adjustment only in the second ('extreme') regime, which corresponds to the pre-crisis period and also incorporates the moment of Romania becoming member of the EU. The results proved that the portfolio effect is present in almost all times. The effect of domestic currency depreciation on stock market (international trading effect) was not positive, so that we may conclude that the depreciation of exchange rate was not significant enough for the exporting firms to fully benefit.

The analysis of the links between the two markets is particularly important for the implication that the central bank's decision about monetary policy would have a strong impact on both financial markets.

Romania as an emerging economy has capital markets still in a catching-up phase, but quite attractive to the investors who look for diversification. In this respect, the results of this paper could be useful for international portfolio management or for the predictability of the exchange rate movements.

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References

Ajayi, R.A., Mougoue, M., 1996. On the Dynamic Relation between Stock Prices and Exchange Rates. *The Journal of Financial Research*, 19, pp.193-207.

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

Andrews,D.W.K. Ploberger, W., 1994. Optimal tests when nuisance parameters is present only under the alternative. *Econometrica*, 62, pp. 1383-1414.

Baek, E. Brock, W. A., 1992. General test for nonlinear Granger causality: bivariate model. *Working Paper*. Iowa State University and University of Wisconsin.

Balke, N.S. and Fomby, T.B., 1997. Threshold Cointegration, *International Economic Review*, 38, 627-645.

Chan, K.S., 1993. Consistency and limiting distribution of the least squares estimator of a threshold autoregressive model. *The Annals of Statistics*, 21. pp. 520-533.

Davies, R.B., 1987. Hypothesis testing when a nuisance parameter is present only under the alternative. *Biometrica*, 74, pp. 33-43.

Diamandis, P.F. Drakos, A.A., 2011. Financial liberalization, exchange rates and stock prices: Exogenous shocks in four Latin America countries. *Journal of Policy Modeling* 33, pp. 381–394

Diks, C.G.H. Panchenko, V., 2005. A note on the Hiemstra-Jones test for Granger non-causality. *Studies in Nonlinear Dynamics and Econometrics*, 9(2), pp.1-7.

Diks, C.G.H. Panchenko, V., 2006. A new statistic and practical guidelines for nonparametric Granger causality testing. *Journal of Economic Dynamics and Control*, 30, pp.1647-1668.

Granger, C.W. Huang, B. and Yang, C., 2000. A Bivariate Causality between Stock Prices and Exchange Rates: Evidence from Recent Asian Flu. *Quarterly Review of Economics and Finance*, 40, pp. 337–354.

Hansen, B.E. and Seo, B., 2002. Testing for two-regime threshold cointegration in vector error-correction models. *Journal of Econometrics*, 110, pp. 293-318.

Hiemstra, C Jones, J.D., 1994. Testing for linear and nonlinear Granger causality in the stock price-volume relation, *The Journal of Finance*, 49(5), pp.1639-1664.

Horobet, A. and Ilie, L., 2007. On the dynamic link between stock prices and exchange rates: evidence from Romania. *MPRA Paper* No. 6429.

Kapetanios, G. Shin, Yongcheol, and Snell, A. 2003. Testing for a unit root in the nonlinear STAR framework, *Journal of Econometrics*, 112, pp. 359-379.

Kim, K., 2003. Dollar Exchange Rate and Stock Price: Evidence from Multivariate Cointegration and Error Correction Model. *Review of Financial Economics*, 12, pp. 301-313.

Ma, C.K. and Kao, G.W., 1990. On Exchange Rate Changes and Stock Price Reactions. *Journal of Business Finance and Accounting*, 17(2), pp. 441–449.

Madison. Bahmani-Oskooee, M. and Sohrabian, A., 1992. Stock Prices and the Effective Exchange Rate of the Dollar. *Applied Economics*, 24, pp. 459-464.

Ravazzolo, F. and Phylaktis, K., 2000. Stock Prices and Exchange Rate Dynamics. EFMA 2000 Athens; *Cass Business School Research Paper*. Available at SSRN: http://ssrn.com/abstract=251296 or http://dx.doi.org/10.2139/ssrn.251296

Tsai, I.C., 2012. The relationship between stock price index and exchange rate in Asian markets: A quantile regression approach. *Journal of International Financial Markets, Institutions & Money*, 22, pp. 609-621.

Yau, H-Y. and Nieh, C-C., 2009. Testing for cointegration with threshold effect between stock prices and exchange rates in Japan and Taiwan. *Japan and the World Economy*, 21, pp.292-300.

Yu Hsing, 2011. Effects of Macroeconomic Variables on the Stock Market: The Case of the Czech Republic. *Theoretical and Applied Economics*, (7), pp. 53-64.

Zhao, H., 2010. Dynamic relationship between exchange rate and stock price: evidence from China. *Research in International Business and Finance*, 24, pp. 103–12.

Appendix

Table 4

Linear-ECM model with 1 and 2 lags

	Lag=1		Lag=2	
	BET	EXU	BET	EXU
Z_{t-1}	-0.019 (-1.8521)	-0.004 (-0.3879)	-0.0173 (-1.708)	-0.0045 (-0.443)
Constant	0.19562 (2.71384)	0.0431 (0.59814)	0.2131 (2.22937)	0.05698 (0.5961)
BET(-1)	0.25246 (2.76289)	-0.013 (-0.1421)	0.2526 (2.4148)	-0.0098 (-0.0941)
EXU(-1)	-1.1343 (-2.5651)	0.3138 (0.70964)	-1.1096 (-2.6102)	0.31883 (0.75)
BET(-2)			-0.005 (-0.0431)	-0.0193 (-0.1668)
EXU(-2)			-0.0555 (-0.0968)	-0.077 (-0.1344)
AIC	-646.58		-635.82	
BIC	-644.8		-633.18	
Log-Like:	-662.58		-659.82	

Note: Numbers in parenthesis are t-statistic and p-values, respectively.

Table 5

TAR-ECM model

TAR ECM	First regime		Second regime	
Variable's Coefficient	ΔBET	ΔEXU	ΔBET	ΔEXU
	Threshold Estimate: 13.1064			
	Percentage of obs. 0.898		Percentage of obs. 0.102	
Z_{t-1}	-0.007 (0.01028)	-0.005 (0.00161)	-0.1534 (0.20578)	0.0917 (0.01277)
Constant	0.12359 (0.0568)	0.0607 (0.01678)	2.08166 (2.83061)	-1.2865 (0.17526)
ΔBET(-1) (δ_1)	0.36575 (0.0864)	-0.0321 (0.01365)	-0.2456 (0.28917)	0.0305 (0.02460)
ΔEXU(-1) (γ_1)	-0.9292 (0.41005)	0.2831 (0.09202)	-2.7559 (2.70498)	0.1859 (0.23289)
AIC	-647			
BIC	-643.43			
Log-Like:	-679			

Note: 1. Numbers in parenthesis are Eicker-White standard errors and p-values, respectively.

2. Threshold error-correction model (the upper indices denote the regime):

$$\Delta x_t = (a^1 + b^1 Z_{t-1} + \sum_{i=1}^{k1} \delta_i{}^1 x_{1t-i} + \sum_{i=1}^{k2} \gamma_i{}^1 x_{1t-i}) I\{Z_t < \lambda\}$$
$$+ (a^2 + b^2 w_{t-1} + \sum_{i=1}^{k1} \delta_i{}^2 x_{1t-i} + \sum_{i=1}^{k2} \gamma_i{}^2 x_{1t-i}) I\{Z_t \geq \lambda\} + e_t$$

Table 5S

Statistics for TAR-ECM

	First regime		Second regime	
TAR ECM	***Threshold Estimate:***	***13.1064***		
First regime	***Percentage of obs.***	***0.898***	***Percentage of obs.***	***0.102***
Wald Test for Granger Causality $\delta_i = 0, \quad for \quad all \quad i = 1,..,lag$				
		5.5296 (0.0186)		4.994 (0.0254)
Wald Test for long-run Granger Causality H0: $\lambda = \delta_i = 0, \quad for \quad all \quad i = 1,..,lag$				
		6.2732 (0.0000)		182.042 (0.0000)
Wald Test for Granger Causality H0: $\gamma_i = 0, \quad for \quad all \quad i = 1,..,lag$				
	5.1346 (0.0235)		45.1702 (0.000)	
Wald Test for long-run Granger Causality H0: $\lambda = \gamma_i = 0, \quad for \quad all \quad i = 1,..,lag$				
	5.1717 (0.0229)		484.9629 (0.000)	
Model Test Statistics				
Wald Test for Equality of Dynamic Coefficient			24.149 (7.5E-05)	
Wald Test for Equality of EC Coefficient			57.088 (0.0000)	
Lagrange Multiplier Threshold Test			16.945	
Fixed Regressor (Asymptotic) 0.05 Critical Value:			18.989	
Bootstrap 0.05 Critical Value:			16.781	
Fixed Regressor (Asymptotic) / Bootstrap P-Value:			0.127 / 0.0462	

Table 6

MTAR-ECM model

	First regime		Second regime	
	Δ BET	Δ EXU	Δ BET	Δ EXU
MTAR ECM	***Threshold Estimate: 0.455***			
	Percentage of obs. 0.904		***Percentage of obs. 0.096***	
Z_t (λ)	-0.0076 (0.0091)	-0.0053 (0.0017)	-0.1528 (0.0314)	0.0181 (0.0057)
Constant	0.1451 (0.0737)	0.0606 (0.0164)	0.7689 (0.3039)	-0.0985 (0.0473)
Δ BET(-1) (δ_1)	0.3877 (0.1123)	-0.0217 (0.0130)	1.6426 (0.3814)	-0.1677 (0.0943)
Δ EXU(-1) (γ_1)	-1.3053 (0.3999)	0.2933 (0.0919)	3.1479 (0.9286)	0.1859 (0.7980)
AIC	-645.91		-628.644	
BIC	-642.36		-623.361	
Log-Like:	-677.92		-676.64	

Note: Numbers in parenthesis are Eicker-White standard errors and p-values, respectively. The threshold variable is the decay depending on previous period change in the error-correction term.

Table 6S

Statistics for MTAR-ECM model

	First regime		Second regime	
MTAR ECM	***Threshold Estimate: 0.455***			
	Percentage of obs. 0.904		***Percentage of obs. 0.096***	
Wald Test for Granger Causality $\delta_i = 0, \quad for \quad all \quad i = 1,..,lag$				
		2.4269 (0.1192)		24.707 (0.0001)
Wald Test for long-run Granger Causality H0: $\lambda = \delta_i = 0, \quad for \quad all \quad i = 1,..,lag$				
		3.6831 (0.0000)		24.575 (0.0000)
Wald Test for Granger Causality H0: $\gamma_i = 0, \quad for \quad all \quad i = 1,..,lag$				
	10.656 (0.0011)		61.980 (0.0000)	
Wald Test for long-run Granger Causality H0: $\lambda = \gamma_i = 0, \quad for \quad all \quad i = 1,..,lag$				
	10.981 (0.0010)		544.352 (0.0000)	
Model Test Statistics				
Wald Test for Equality of Dynamic Coefficient			24.707 (0.000058)	
Wald Test for Equality of EC Coefficient			24.575 (0.000005)	
Lagrange Multiplier Threshold Test			19.9512	
Fixed Regressor (Asymptotic) 0.05 Critical Value:			18.8668	
Bootstrap 0.05 Critical Value:			19.8083	
Fixed Regressor (Asymptotic) / Bootstrap P-Value:			0.0292 / 0.0478	

Table 7

TAR-ECM model with asymmetries related to short-term good or bad news from the two markets

	First regime		**Second regime**	
TAR ECM (news)	Δ BET	Δ EXU	Δ BET	Δ EXU
	Threshold Estimate: 8.0940			
	Percentage of obs. 0.934		***Percentage of obs. 0.066***	
Zt (λ)	-0.0245 (0.0140)	-0.0068 (0.0022)	-0.3069 (0.1569)	0.0877 (0.0285)
Constant	0.1829 (0.0439)	0.0375 (0.0106)	2.1607 (1.4452)	-0.7313 (0.2585)
Δ BET+(-1) (δ_1)	0.3044 (0.1571)	-0.0362 (0.0279)	-0.9937 (0.4577)	0.1059 (0.0773)
Δ EXU+(-1) (γ_1)	-1.7670 (0.4099)	0.2599 (0.1352)	6.2322 (1.4920)	-0.3537 (0.4371)
Δ BET- (-1) (δ_2)	0.3827 (0.1485)	-0.0195 (0.0197)	-0.1989 (0.2736)	-0.0706 (0.0602)
Δ EXU- (-1) (γ_2)	0.0549 (1.3341)	0.2299 (0.1526)	-22.214 (6.1323)	1.8912 (0.8264)
AIC	-658.16		-650.36	
BIC	-654.60		-645.07	
Log-Like:	-690.16		-698.36	

Note: 1. Numbers in parenthesis are Eicker-White standard errors and p-values, respectively.

2. Threshold error-correction model:

$$\Delta x_t = (a^1 + b^1 Z_{t-1} + \delta_1^1 \Delta x_{1t-1}^+ + \delta_2^1 \Delta x_{1t-1}^- + \sum_{i=2}^{k1} \delta_{i+1}^1 \Delta x_{1t-i} + \gamma_1^1 \Delta x_{2t-1}^+ + \gamma_2^1 \Delta x_{2t-1}^- + \sum_{i=2}^{k2} \gamma_{i+1}{}^1 \Delta x_{2t-i}) I_{\{Z_t<\lambda\}} +$$

$$+ (a^2 + b^2 w_{t-1} + \delta_2^2 \Delta x_{1t-1}^+ + \delta_2^2 \Delta x_{1t-1}^- + \sum_{i=2}^{k1} \delta_{i+1}^2 \Delta x_{1t-i} + \gamma 2 \Delta x_{2t-1}^+ + \gamma_2^2 \Delta x_{2t-1}^- + \sum_{i=2}^{k2} \gamma_{i+1}{}^1 \Delta x_{2t-i}) I_{\{Z_t \geq \lambda\}} + e_t$$

Where the upper indices denote the regime: $x_t = (BET, EXU)$ *,* $\Delta x^+{}_t = (I_t \Delta BET, I_t \Delta EXU)$ $\Delta x^-{}_t = [(1-I_t)\Delta BET, (1-I_t)\Delta EXU]$ *given* $I_t = 1$ *if* ΔBET *and* ΔEXU *are positive respectively negative; and* $Z_{t-1} = w_{t-1}(\beta)$ *is the error-correction term.*

Table 7S

Statistics for TAR-ECM model with asymmetries related to short-term good or bad news from the two markets

	First regime		**Second regime**	
TAR-ECM(news)	***Threshold Estimate: 8.0940***			
First regime	***Percentage of obs. 0.934***		***Percentage of obs. 0.066***	
Wald Test for Granger Causality $\delta_i = 0, \quad for \quad all \quad i = 1,..,lag$				
		4.389 (0.111)		6.247 (0.0440)
Wald Test for long-run Granger Causality H0: $\lambda = \delta_i = 0, \quad for \quad all \quad i = 1,..,lag$				
		4.949 (0.0000)		6.544 (0.0379)
Wald Test for Granger Causality H0: $\gamma_i = 0, \quad for \quad all \quad i = 1,..,lag$				
	21.909 (0.00001)		29.934 (0.0000)	
Wald Test for long-run Granger Causality H0: $\lambda = \gamma_i = 0, \quad for \quad all \quad i = 1,..,lag$				
	22.496 (0.00001)		30.185 (0.0000)	
Wald Test for H0: $\delta_1 = \delta_2$				
		0.2401 (0.624)		3.250 (0.0714)
Wald Test for H0: $\gamma_1 = \gamma_2$				
	1.704 (0.1917)		20.316 (0.0000)	
Model Test Statistics				
Wald Test for Equality of Dynamic Coefficient			127.408 (0.0000)	
Wald Test for Equality of EC Coefficient			11.847 (0.0027)	
Lagrange Multiplier Threshold Test			21.0190	
Fixed Regressor (Asymptotic) / Bootstrap 0.05 Critical Value:			23.967 / 15.256	
Fixed Regressor (Asymptotic) / Bootstrap P-Value:			0.1834 / 0.0054	

Table 8
MTAR-ECM model with asymmetries related to short-term good or bad news from the two markets

	First regime		Second regime	
MTAR-ECM (news)	Δ BET	Δ EXU	Δ BET	Δ EXU
	Threshold Estimate: 0.454			
	Percentage of obs. 0.904		***Percentage of obs. 0.096***	
Zt (λ)	-0.0092 (0.0085)	-0.0052 (0.0016)	-0.2077 (0.0310)	0.0209 (0.0102)
Constant	0.1209 (0.0719)	0.0665 (0.0178)	1.9702 (0.4113)	-0.1950 (0.1608)
Δ BET+(-1) (δ_1)	0.5694 (0.2565)	-0.0223 (0.0407)	1.2531 (0.2432)	-0.1262 (0.1039)
Δ EXU+(-1) (γ_1)	-0.9064 (0.4923)	0.2451 (0.1295)	-1.3154 (2.1937)	0.8069 (0.4812)
Δ BET- (-1) (δ_2)	0.3386 (0.1496)	-0.0221 (0.0186)	79.862 (20.209)	-6.3435 (8.2594)
Δ EXU- (-1) (γ_2)	-2.1799 (1.2813)	0.4131 (0.1795)	4.6386 (1.3639)	-0.1644 (0.2876)
AIC	-649.33		-634.83	
BIC	-645.77		-629.55	
Log-Like:	-681.331		-682.83	

Note: 1. Numbers in parenthesis are Eicker-White standard errors and p-values, respectively.
2. Threshold error-correction model:

$$\Delta x_t = (a^1 + b^1 Z_{t-1} + \delta_1^1 \Delta x_{1t-1}^+ + \delta_2^1 \Delta x_{1t-1}^- + \sum_{i=2}^{k1} \delta_{i+1}^1 \Delta x_{1t-i} + \gamma_1^1 \Delta x_{2t-1}^+ + \gamma_2^1 \Delta x_{2t-1}^- + \sum_{i=2}^{k2} \gamma_{i+1}{}^1 \Delta x_{2t-i}) I_{\{Z_t < \lambda\}} +$$

$$+ (a^2 + b^2 w_{t-1} + \delta_2^2 \Delta x_{1t-1}^+ + \delta_2^2 \Delta x_{1t-1}^- + \sum_{i=2}^{k1} \delta_{i+1}^2 \Delta x_{1t-i} + \gamma_1^2 \Delta x_{2t-1}^+ + \gamma_2^2 \Delta x_{2t-1}^- + \sum_{i=2}^{k2} \gamma_{i+1}{}^2 \Delta x_{2t-i}) I_{\{Z_t \geq \lambda\}} + e_t$$

Where the upper indices denote the regime, $x_t = (BET, EXU)$, $\Delta x^+{}_t = (I_t \Delta BET, I_t \Delta EXU)$ $\Delta x^-{}_t = [(1-I_t)\Delta BET, (1-I_t)\Delta EXU]$ *given* $I_t = 1$ *if* ΔBET *and* ΔEXU *are positive respectively negative, and* $Z_{t-1} = Diff(w_{t-1}(\beta))$ *is the error-correction term.*

Table 8S

Statistics for MTAR-ECM model with asymmetries related to short-term good or bad news from the two markets

M-TAR-ECM(news)	First regime		Second regime	
	Percentage of obs. 0.934		Percentage of obs. 0.066	
Wald Test for Granger Causality $\delta_i = 0, \quad for \quad all \quad i = 1,..,lag$				
		2.7056 (0.2585)		6.0951 (0.0475)
Wald Test for long-run Granger Causality H0: $\lambda = \delta_i = 0, \quad for \quad all \quad i = 1,..,lag$				
		4.1122 (0.0000)		10.6887 (0.0048)
Wald Test for Granger Causality H0: $\gamma_i = 0, \quad for \quad all \quad i = 1,..,lag$				
	9.3896 (0.0091)		12.390 (0.002)	
Wald Test for long-run Granger Causality H0: $\lambda = \gamma_i = 0, \quad for \quad all \quad i = 1,..,lag$				
	9.4589 (0.0088)		51.997 (0.000)	
Wald Test for H0: $\delta_1 = \delta_2$				
		1.8e-005 (0.9966)		0.5665 (0.4516)
Wald Test for H0: $\gamma_1 = \gamma_2$				
	0.8608 (0.3536)		5.313 (0.0212)	
Model Test Statistics				
Wald Test for Equality of Dynamic Coef.			50.157 (0.0000)	
Wald Test for Equality of EC Coef.			40.3017 (0.0000)	
Lagrange Multiplier Threshold Test			21.9501	
FixedRegressor(Asymptotic)/ Bootstrap0.05Critical Value:			24.202 / 20.0140	
Fixed Regressor (Asymptotic) / Bootstrap P-Value:			0.1282 / 0.0188	