THE EFFECT OF EURO ADOPTION ON NASDAQ OMX BALTIC STOCK EXCHANGE: ANALYSIS BY STRUCTURAL BREAK TESTS

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Abstract. *Although the euro adoption in Estonia in 2011 and changing the trading and clearing currency at the NASDAQ OMX Vilnius for euro on 22 November 2010 were foreseen as a possibility to attract more foreign investors, last year the Baltic stock exchange underwent some extreme fluctuations, both positive and negative. In this paper, shown are statistically significant euro adoption-caused trend breaks underlying the data set of NASDAQ OMX stock exchanges in Tallinn and Vilnius. Also, the possible factors that may have been driving them are discussed. The assessment is carried out using three different structural break tests.*

Key words: *structural break, cointegration, euro adoption, stock exchange*

Introduction

The greater attractiveness of local financial markets is supposed to be one of the most obvious advantages of the euro adoption. It is assumed that trading cost-minimization, ability to compare stock prices in different markets without a struggle, to use financial arbitrage possibilities more extensively and to enhance trust in public finance (remember Maastricht criteria which are necessary to meet in order to adopt the euro) may cause a "cascade effect" by building confidence on the financial markets and attracting new investors one after another. On 1 January 2011 the euro was adopted in Estonia, and from 22 November 2010 it is traded in euro at the NASDAQ OMX Vilnius; moreover, in the middle of 2010 it was stated that trading in euro would be one of the consecutive steps to create a common stock market in the Baltics (NASDAQ OMX Vilnius; Press Center, 26. 07. 2010.). The consequences of these events are possibly long-term but outweighed by the global macroeconomic situation. This encouraged me to challenge the hypothesis that the adoption of euro as the trading currency had a statistically significant impact on the trading volume at the NASDAQ OMX stock exchanges in Tallinn and Vilnius, e.g., on attracting investments into the Baltic financial markets.

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The effect of the euro as a trading currency on foreign investors' attitude towards the Baltic financial markets is assessed by invoking the structural break definition which describes it as a statistically significant alteration in the mean value, a slope of trend or both (Perron, 1989; Enders, 2004). Three structural break tests were chosen: the Chow test (1960), the Perron test (1989) and the Lütkepohl et al. test (2004), which have expanded the possibilities of this research.

The paper contains a literature review which covers some related studies and their inferences used to define the frame of the paper, methodological framework, three other parts dedicated to each of the tests, and conclusions.

1. Literature review

Although Pierre Perron, one of the predecessors of the structural break testing, revealed most of macroeconomic variables to be trend-stationary time-series and confirmed only two structural breaks to have a pervasive impact (the 1929 crash and the 1973 oil price shock) (Perron, 1989), nowadays the structural breaks and parameter stability issues are often discussed concerning the economic data observed in developing countries (Çağli, 2010; Eizaguirre, 2009) and the tendencies underlying the financial markets (Moon, 2010). This is determined by the financial data and the specific features of developing countries' economies: they are characterized by the indicators' volatility allowing more possible or false trend changes; moreover, the impact of structural breaks on conclusions might be very controversial.

Whereas ignoring possible structural breaks may lead to inadequate implications, recently it has been suggested to carry out a time series analysis by using the alternative unit root Lagrange multiplier test (Narayan, 2008)) and cointegration tests (e. g., the Gregory–Hansen test), which allow 1–2 potential trend changes (Çağli, 2010). This proposal gets even more importance when dealing with the relationship of macroeconomic indicators with the data of financial markets. For instance, including a structural break allows rejecting the efficient-market hypothesis for both developed and developing countries, because stock index return statistics was determined to be stationary (Lee, 2010). Continuing research on the role of structural breaks in modeling financial processes, there was identified a greater importance of factors not connected to firms directly (not mergers or dividend news (Coutts, 1997)). According to the cited authors, in most cases structural breaks are caused not by the global factors (e. g., bird flu), but by local ones such as market liberalization, etc. (Eizaguirre, 2009). Wage-regulation rejection, drought or military conflicts were acknowledged as statistically significant contributors to the change of economic course, too (Narayan, 2008). Thus, the structural breaks while considering country specifics (agrarian, developing, politically unstable), might be determined by both economic and social factors. Invoking the aforementioned conclusions, the factor defined in the hypothesis seems to be reasonable. Because of its nationwide effect, which outweighs separate microeconomic processes but is not considered to be global, the adoption of euro as the trading currency is validated not only by the logical reasoning, but also by the results of similar studies, too.

The methods used to carry out the dating of structural breaks vary from CUMSUMQ to the Markov Chain Monte Carlo (MCMC); the Chow test, published in 1960, is considered to be traditional (Coutts, 1997). Although one of the predecessor-tests, CUSUM, was criticized because of the underlying requirement of non-correlation between regressors and residuals, Krämer et al. (1988) proved it to be extended to the dynamic models as well. However, they suggested using alternative test if there is prior information about the occurrence of the structural break rather late in the sample. In most cases, despite the model estimation method, the hypothetical trend change is included as a dummy variable. Although including a dummy variable is approached in the same way also in Perron's test (Perron, 1989), it has superiority over the Chow test because it does not diminish the degrees of freedom by splitting a sample. Analysis of the structural breaks in the vector error-correction model (VECM), when the breakpoint is unknown, may be proceeded by a Lütkepohl et al. test (Lütkepohl et al., 2004). It is designed to investigate the cointegrating relationships which might be rejected in the case of underlying an AR(1) process by a structural break. In the Lütkepohl et al. test, this possibility is examined by constructing an appropriate model and evaluating its statistics, concerning cointegration hypothesis.

One more step closer to reality is multiple structural break treatment, which was extended to use it in the linear multivariate regression models as well (Qu, Perron, 2007). Jushan Bai and P. Perron had discussed the estimation of multiple breaks using OLS much earlier (1998). It covers the forming of confidence levels for the break dates, testing them under certain conditions and estimating also the number of breaks (Bai, Perron, 2003). Although the hypothesis of this study did not require testing multiple breaks, these works might be considered as material for the further research.

2. Methods

In order to confirm or reject the hypothesis, it was decided to use a data set consisting of the main indicator – trading volume – and stock indices as the additional time series. Considering the financial markets' volatility (their extreme reaction to the outstanding news) and relatively high daily trade fluctuations, the selected indicators seem to be appropriate for detecting stock trading changes day by day. Therefore, the hypothesis states that the more active the Baltic financial markets, the higher the trading volume and the stock index value (influenced by the increasing demand). The point of a statistically significant positive change of these indicators' dynamics is considered to be the date of the structural break at the NASDAQ OMX Baltic, which might be caused by the above-mentioned currency issues.

The structural break tests intended for performing the structural break analysis are used to evaluate statistically significant trend changes at certain moments. Considering the fact that financial markets are characterized by a strong reaction to external shocks, it has been decided to ignore only one-week or shorter reaction (change in the dependent variable values) lags and consider them as a measuring error. The scope of the study includes the period between January 2010 and 18 July 2011, so the author seeks to avoid underestimating relatively recent structural breaks. This might be caused by a disproportionally long backward time series. In this paper, data of NASDAQ OMX Baltic stock exchange are used.

The first test used to carry out an assessment of the structural break was the Chow test. It is based on parameter stability evaluation for different observations groups (Chow, 1960). In order to detect the break, two regressions are composed:

$$y_t = \beta_1 z_{1t} + \beta_2 z_{2t} + u_t \tag{1}$$

and

$$y_t^* = \beta_1^* z_{1t}^* + \beta_2^* z_{2t}^* + u_t^*, \tag{2}$$

where y_t, z_{1t}, z_{2t} belong to the first group of observations and y_t^*, y_{1t}^*, y_{2t}^* to the second one. The z time series defines some independent variables that might differ for different models, but not for equations (1) and (2) describing the same model. u_t and u_t^* are white-noise disturbances. Regarding the uneven distribution of the possible structural breaks and consequently the difference in samples' size, the shock variance differs across the equations. It is recommended to modify the data by dividing them by the standard deviation of the corresponding equation disturbances (Maddala, 2002). Then, in order to investigate the null hypothesis for β_1 statistically equal to β_1^* and β_2 to β_2^*, the F value is estimated. This statistics is based on the variance difference across the equations and enables to assess the statistical significance of parameter stability criteria by quantifying them (Wooldridge, 2006):

$$F = \frac{\dfrac{RSS - RSS_{ur}}{k+1}}{\dfrac{RSS_{ur}}{n-2(k+1)}} \sim F\,(k+1,\, n-2(k+1)), \tag{3}$$

where RSS denotes the residual sum of squares for the general regression and RSS_{ur} represents the residual sum of squares from equations (1) and (2), k is the number of independent variables in both regressions (1) and (2), and n defines the size of a full sample.

If no statistically significant trend changes are detected, the parameter stability hypothesis should not be rejected.

Another group of the structural break tests was intended to avoid splitting the sample in order to save the degrees of freedom. So, it was preferable to use a single test based on the full sample (Enders, 2004). One of these tests, which is quite general to approach both the structural shift in the drift and in the slope of the trend, was proposed by Perron

(1989). Since it is basically used for a unit root testing, he found most of macroeconomic variables to be trend-stationary with a constant slope and a change in the level around 1929 (Enders, 2004). The dilemma appears to distinguish between a unit root process, which exhibits a single pulse in a series, and a variable which is stationary within each of the subperiods divided by structural breaks. If a break point happens at time period τ, consider the null hypothesis of a change in the level and the drift of a unit root process:

$$y_t = a_0 + y_{t-1} + \beta_1 D_P + \beta_2 D_L + u_t, \tag{4}$$

where u_t denotes the detrended series, D_p is a dummy variable ($D_L = 1$, when $t = \tau$, and to zero otherwise), and D_L is a dummy variable which represents a possible change in the slope of the trend ($D_L = 1$, when $t > \tau$, and zero otherwise).

And the alternative one:

$$y_t = a_0 + a_1 t + \beta_2 D_L + \beta_3 D_T + u_t, \tag{5}$$

where D_T is a trend dummy equal to $t = \tau$ for $t > \tau$ and zero otherwise.

The procedure requires estimating equation (5) and testing the detrended series for a unit root (the ADF test). Upon estimating the break point and the full sample ratio (τ / T), one has the λ value and can compare the t value with the critical value, which corresponds to a certain λ (Perron, 1989). If the t value is greater in its absolute value than Perron's one, the null hypothesis can be rejected.

The third test, suggested by Lütkepohl et al. (2004), was performed to detect the structural breaks underlying the VECM when the break point is unknown. Since it is designed to assess the possibility of a cointegrating relationship, it examines whether there are certain indications of the AR (1) with a structural break in the model. Therefore, in order to estimate such a model, it is necessary to compose a vector of variables which are at most I (1) and hypothetically are characterized by a long-term equilibrium. Therefore, it is presumed that $\{y_t\}$ (the vector of variables) is regressed by the constant μ_0, the linear trend term t and the trend break which is included as the dummy variable $d_{t\tau}$:

$$y_t = \mu_0 + \mu_1 t + \delta d_{t\tau} + x_t, \tag{6}$$

where $d_{t\tau} = 0$, when $t < \tau$, and $d_{t\tau} = 1$ when $t \geq \tau$. It is presumed that the break point τ is unknown. It is expressed as part of the sample:

$$\tau = [T\varphi], \tag{7}$$

with $0 < \tilde{\varphi} \leq \varphi \leq \overline{\varphi} < 1$. $\tilde{\varphi}$ and $\overline{\varphi}$ define real numbers. Henceforth, the break point is not supposed to be at the very beginning or at the very end of a sample.

The vector $\{x_t\}$ helps to represent the VAR(p) process concerning the $\{y_t\}$ dynamics, e. g.:

$$y_t = \gamma_0 + \gamma_1 t + \delta d_{t\tau} + A_1 y_{t-1} + ... + A_p y_{t-p} + \varepsilon_{t\tau}, \tag{8}$$

where A_i denotes the coefficient matrix and $\varepsilon_{t\tau}$ is a residual vector. So, to estimate the break point, the following expression has to be solved:

$$\hat{\tau} = \arg\min_{\tau \in \check{T}} \det (\Sigma_{t=p+1}^{T} \ \hat{\varepsilon}_{t\tau} \hat{\varepsilon}'_{t\tau}), \tag{9}$$

where $\check{T} = [T\tilde{\varphi}, T\overline{\varphi}]$. After revealing the estimator (9), the data are modified in the following way:

$$\hat{x}_t = y_t - \hat{\mu}_0 - \hat{\mu}_1 t - \hat{\delta} d_{t\hat{\tau}} . \tag{10}$$

This procedure enables assessing the implications of the involvement of a structural break in the model by estimating λ_{trace} and λ_{max} statistics. λ_{trace} formulates the hypothesis that the number of cointegrating vectors is equal to r (alternative hypothesis assumes that there exists the $r + 1$ number of cointegrating vectors). It should be noted that, if the analysis of the initial data allows rejecting the hypothesis of $r = 0$ and the case of modified data is different, the structural break point is obviously significant. To evaluate the possibility of a cointegrating relationship, it is suggested to use critical values from Trenkler (Pfaff, 2006).

The estimation of the structural break point provides the researcher with a possibility to assess whether the break points underlying a certain data set coincide with the hypothetical ones.

3. Estimation results

3.1. The Chow test

In order to perform the Chow test, separate models were composed for the NASDAQ OMX Vilnius and the NASDAQ OMX Tallinn stock exchanges by choosing the trading volume as an endogenous variable. Its lags, stock index value and linear trend are included into the model as independent components.

The linear regression is modified for the Lithuanian case, considering t statistics of estimated parameters, residual autocorrelation coefficients' values and results of the Ljung–Box test. It is approached in this way: whereas lags mostly contribute to eliminating systemic components from the residuals or reflecting the inertia of the dependent variable and have little to do with the analysis of a certain parameter (which is not the main purpose of this research, either), it is focused on disturbance characteristics and the statistical significance of exogenous variables but not on the lags of endogenous ones. Consequently, the equation that fulfils the model correctness criteria needs to be estimated:

$$\nu_t = a_1 + a_2 t + a_3 \Delta \nu\nu_t + \Sigma_{p=1}^{4} g_p \nu_{t-p} + \varepsilon_t, \tag{11}$$

where ν_t denotes the logarithm of the daily stock trading volume, $\Delta\nu\nu_t$ is a daily change of the stock index value (the level value is found to be a random walk with a drift and a trend (see table 1)), and a_1 is a constant term which is included considering the dynamics of variables with a drift; ε_t denotes normally distributed residuals of this model. Since there are no indications of developed biasness towards an increase or decrease (see Fig. 1), the linear trend term t may be not included in the model, yet a look at the ADF unit root test results (see Table 1) finds these elements relevant to keep the stock trading volume series stationary. The inference is confirmed the by φ_2 and φ_3 values (see Table 3) which show the constant and the trend consistency with the model.

TABLE 1. **ADF unit root test**

Variable	*t*	*t* (with a constant)	*t* (with a constant and a linear trend term)
ν_t	–0.2061	–4.8656*	–5.1196*
$\nu\nu_t$	1.873	–2.3066	–2.1758
$\Delta\nu\nu_t$	–12.1202*	–12.291*	–12.378*
τ_t	0.0356	–5.2515*	–5.5876*
$\tau\nu_t$	1.5274	–3.2072**	–2.7628
$\Delta\tau\nu_t$	–12.1188*	–12.235*	–12.4203*

Note: null hypothesis that the time series variable has a unit root can be rejected at 1% (*) or at 5% (**) levels.

Source: NASDAQ OMX Baltic, author's estimation.

TABLE 2. **ADF unit root test**

	τ_t	Critical values		
		1%	5%	10%
τ	–5.59	–3.98	–3.42	–3.13
φ_2	10.432	6.15	4.71	4.05
φ_3	15.63	8.34	6.30	5.36

Source: NASDAQ OMX Baltic, author's estimation.

TABLE 3. **ADF unit root test**

	ν_t	Critical values		
		1%	5%	10%
τ	–5.12	–3.98	–3.42	–3.13
φ_2	8.74	6.15	4.71	4.05
φ_3	13.11	8.34	6.30	5.36

Source: NASDAQ OMX Baltic, author's estimation.

The case of the NASDAQ OMX Tallinn stock exchange requires an analogous procedure to be performed. The ADF unit root test reveals that the logarithm of the daily stock trading volume (τ_t) has a drift term, and the stock index (OMXT) value variable ($\tau\nu_t$) is generated by a I(1) process with a positive trend (table 1). In addition, it is confirmed by the graphical analysis (see Figs. 2 and 3). Whereas the difference of the OMXT value has no trend features any more (Fig. 4), it is supposed to be relevant to detrend the stock trading variable only.

After assessing the error term autocorrelation, the number of the dependent variable lags is limited to 4. Besides, the index value change is included in the model as a statistically significant defining component (Table 5). Finally, this model has to be estimated:

$$\tau_t = b_1 + b_2\Delta\tau\nu_{\mathrm{t}} + b_3 t + \Sigma_{\mathrm{p}=1}^{4} c_{\mathrm{p}}\tau_{t-p} + \omega_t, \tag{12}$$

where a_2 denotes a constant term essential to correct the variables' expectation values (not equal to 0), and ω_t consists of model residuals.

Henceforth, we get the final OLS estimates from the models, which follow an AR(4) process (see Tables 4 and 5). ν_t is an especially inert variable which possesses a statis-

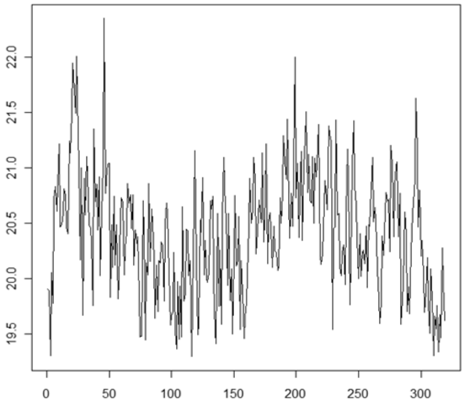

FIG. 1. **The logarithm of daily stock trading volume at the NASDAQ OMX Vilnius**

Source: NASDAQ OMX Baltic, author's calculation.

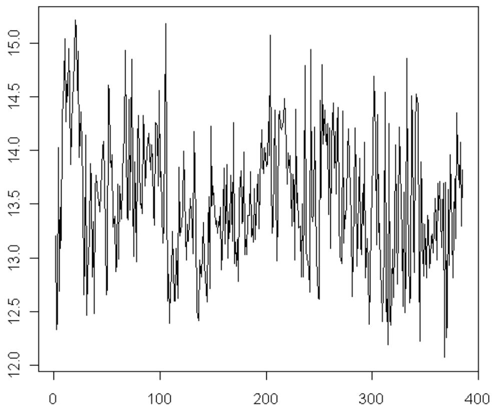

FIG. 3. **The logarithm of OMXT**

Source: NASDAQ OMX Baltic, author's calculation.

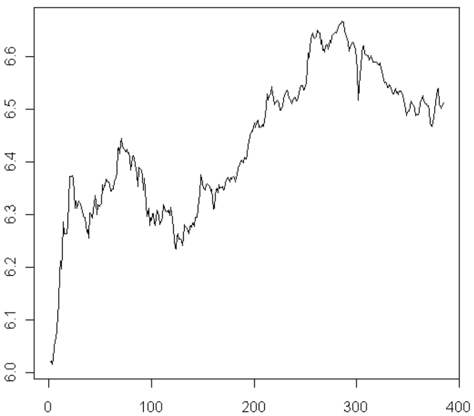

FIG. 2. **The logarithm of daily stock trading volume at the NASDAQ OMX Tallinn**

Source: NASDAQ OMX Baltic, author's calculation.

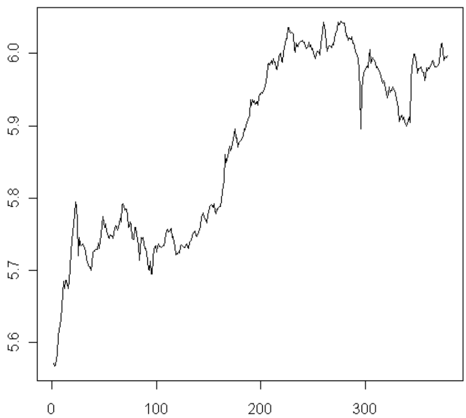

FIG. 4. **The logarithm of OMXV**

Source: NASDAQ OMX Baltic, author's calculation.

tically significant response to almost all of its lags. In this case, the null hypothesis of statistical insignificance cannot be rejected for ν_{t-3} only (see Table 4 again).

TABLE 4. **Model parameters and statistics**

	Estimator	SE	*t* value	Pr(>|t|)
(a_1)	5.3912	0.8705	6.193	1.58e-09
$\Delta\nu\nu_t$	6.2658	2.5708	2.437	0.0153
t	–0.0004	0.0003	–1.637	0.1026
ν_{t-1}	0.3027	0.0511	5.928	7.07e-09
ν_{t-2}	0.1618	0.0533	3.035	0.0026
ν_{t-3}	–0.0356	0.0534	-0.666	0.5059
ν_{t-4}	0.1762	0.0513	3.437	0.0007

Source: NASDAQ OMX Baltic, author's estimation.

TABLE 5. **Model parameters and statistics**

	Estimator	SE	*t* value	Pr(>|t|)
(a_2)	6.1749	0.9551	6.465	3.16e-10
$\Delta\tau\nu_t$	5.0219	2.0672	2.429	0.0156
t	–0.0005	0.0003	–1.652	0.0994
τ_{t-1}	0.2687	0.0509	5.278	2.22e-07
τ_{t-2}	0.0636	0.0524	1.216	0.2248
τ_{t-3}	0.0755	0.0523	1.443	0.1498
τ_{t-4}	0.1420	0.0503	2.824	0.0050

Source: NASDAQ OMX Baltic, author's estimation.

Although this paper is not dedicated to the regressors' analysis, only some inference concerning the stock trading volume variable dynamics could be developed. There is too little evidence to confirm the passivity of NASDAQ OMX stock exchanges in Vilnius and Tallinn, because the inertia found is only a consequence of the lack of other daily variables to include into the model. When continuing the research, foreign stock exchanges' characteristics should be considered as additional factors reflecting the relationship among the financial markets.

The possible trend change in the end of November 2010 is challenged by constructing two samples consisting of 218 and 161 observations. The residual sum of squares (RSS) is calculated for each of the models, based on the samples made. The RSS values differ across the samples marginally (44.7699 and 48.4698, respectively), so the F statistics is estimated for the hypothesis $var(\varepsilon_t^2) \geq var(\varepsilon_t^1)$. On the basis of $F = 1.4834$, where $F \sim F(154, 211)$, there is no evidence to confirm a statistically significant difference between the aforementioned variances, so the variables are not modified by dividing them by the corresponding standard deviation of the model error term (Wooldridge, 2006; Maddala, 2002). The F value (Chow criterion) is equal to 1.2258, where $F{\sim}F(7, 365)$ and the approximate critical value is 2.09.

In the Estonian case, the potential breakpoint separates also two samples (251 and 134 observations). The RSS is calculated for each model, based on the samples made. Whereas RSS values differ across the samples significantly (65.4833 and 44.7183 respectively; for the hypothesis $var(\omega_t^1) \geq var(\omega_t^2)$ $F = 0.7622$, where $F \sim F(244, 127)$), in order to ascertain the presumption of homoscedasticity the variables are modified by dividing them by the corresponding standard deviation of the model error term (Wooldridge, 2006; Maddala, 2002). The f value (Chow criterion) is equal to 1.5944, where $F \sim F(7, 371)$, and the approximate critical value is 2.09.

Henceforth, the null hypothesis cannot be rejected for either of cases; e.g., there is no evidence to determine a statistically significant indication of parameter instability during the period concerned.

3.2. Perron's test for the structural change

Equation (5), which represents the alternative hypothesis of a trend-stationary variable against a unit root process, possesses the same structure for both stock exchange markets, e. g., while investigating whether there is a statistically significant change in both a drift and a slope of the trend. However, some differences appear when applying the ADF test to a detrended series. Consider the stock trading volume as a dependent variable. Then,

$$H_0^{\nu}: \nu_t = a_0 + \nu_{t-1} + \beta_1 D_P + \beta_2 D_L + u_t, \quad \text{and} \tag{13}$$

$$H_0^{\tau}: \tau_t = c_0 + \tau_{t-1} + \gamma_1 D_P + \gamma_2 D_L + z_t, \tag{14}$$

$$H_1^{\nu}: \nu_t = a_0 + \mathrm{a}_1 t + \beta_2 D_L + \beta_3 D_T + u_t\,, \quad \text{and} \tag{15}$$

$$H_1^{\tau}: \tau_t = c_0 + \mathrm{c}_1 t + \gamma_2 D_L + \gamma_3 D_T + z_t\,,$$

where u_t and z_t denote the detrended series, D_p is a dummy variable ($D_L = 1$ when $t = \tau$, and zero otherwise), D_L is a dummy variable which represents a possible change in the slope of the trend ($D_L = 1$ when $t > \tau$, and zero otherwise), and D_T is a trend dummy which is equal to $t = \tau$ for $t > \tau$ and zero otherwise.

The ADF test yields

$$\hat{u}_t = \alpha_1 \hat{u}_t + \Sigma_{k=1}^{13} \alpha_k \Delta \hat{u}_{t-k} + e_{1t} \quad \text{and} \quad \hat{z}_t = \delta_1 \hat{z}_t + \Sigma_{k=1}^{5} \delta_k \Delta \hat{z}_{t-k} + e_{2t}\,,$$

where the differenced series lags are included in order to get the white-noise residuals e_{1t} and e_{2t} (a drift or a trend term is not included in the ADF test regression because the series are already detrended). The t values for α_1 and α_2 are equal to –0.2294 and –0.0026, respectively. Since $\lambda^{\nu} = 0.5752$ and $\lambda^{\tau} = 0.6519$, the critical Peron's values are approximately –4.24 and –4.18 at a 5% level (Perron, 1989) and exceed the t values. The mismatch of a unit root testing (remember the data analysis before applying the Chow test) could be generated of a rather high order of lags of differences, which might have implied some multicollinearity, and a two-step residual estimating procedure (*à la* the Engle–Granger methodology). In conclusion, both daily stock trading volume series approache a unit root process, and regressions (13) and (14) seem to better describe the stock trading dynamics. Then, the OLS estimators (t values are in the parentheses) are as follows:

$$\hat{\nu}_t = \underset{(8.57)}{9.85} + \underset{(3.00)}{0.26}\hat{\nu}_{t-1} - \underset{(-0.37)}{0.23}D_P + \underset{NA}{NA}\cdot D_L , \tag{16}$$

$$\hat{\tau}_t = \underset{(9.20)}{10.57} + \underset{(2.49)}{0.21}\hat{\tau}_{t-1} - \underset{(-0.15)}{0.09}D_P + \underset{NA}{NA}\cdot D_L, \tag{17}$$

where the coefficients of D_L cannot be revealed because of matrix singularities. Whereas the coefficients of D_P are not found to be statistically significant, there is no evidence to confirm statistically significant breakpoints for 22 November 2010 and 1 January 2011.

3.3. Structural break at an unknown point of time test

Whereas statistically significant relationships between the trading volume and the index value series are acknowledged, the Johansen methodology used to assess co integrating relationships can be applied to the sets of variables of the Tallinn and Vilnius stock exchanges. The purpose is, following the Lütkepohl et al. instruction, to examine whether the underlying series behave as an AR(1) process with a structural shift.

The stock index values and trading volumes are generated by the processes of different order, so the daily trading volumes are cumulated in order to avoid the rejection of a potential long-term interrelationship and to construct a typical cointegration model. Keeping in mind that parameter interpretation is of little interest to us this time, such modification is justified. The variables that are derived from it ($c\tau_t$ and $c\nu_t$) are assessed by the ADF unit root test, which unambiguously states that $c\nu_t$ is integrated of order 1, and ct_t is a trend-stationary series (Table 6). This discrepancy between the characteristics

TABLE 6. **ADF unit root test**

Variable	t	t (with a constant)	t (with a constant and a linear trend) term)
$c\nu_t$	4.9013	–2.4474	–1.2862
$\Delta c\nu_t$	–7.2616*	–8.9773*	–9.2146*
$c\tau_t$	4.6476	–5.7206*	–4.034*
$\Delta c\tau_t$	–4.0266 *	–6.5409 *	–8.6173*

Note: the null hypothesis that the time series variable has a unit root can be rejected at 1% (*) or 5% (**) levels.

Source: NASDAQ OMX Baltic, author's estimation.

TABLE 7. **ERS unit root test**

Variable	t (with a constant)	t (with a constant and a linear trend term)
$c\nu_t$	2.754	–2.2218
$c\tau_t$	3.2142	–0.1902

Note: the null hypothesis that the time series variable has a unit root can be rejected at 1% (*) or 5% (**) levels.

Source: NASDAQ OMX Baltic, author's estimation.

of two trading volume series of the same region financial markets leads to a repeated stationarity evaluation by the Elliot, Rothenberg and Stock (ERS) unit root test which is always applied to detrended data. The ERS statistics confirm the suspection of validating the non-stationarity hypothesis (Table 7). Thus, it should be concluded that both the index values and the cumulated trading volumes are generated by the I(1) processes. They also represent some obvious dynamic tendencies which need to be eliminated by including a trend term into the models.

Further, there are composed two models consisting of 10 lags (equation 18) and 19 lags (equation 19), all validated by the information criteria (see Tables 1 and 3 in the Appendix), persistently significant values of the residuals' autocorrelation coefficients and the Ljung–Box test results. Probably this unusually high order of lags is required because of the modification of daily trading volume series.

$$\begin{bmatrix} cv_t \\ vv_t \end{bmatrix} = \begin{bmatrix} \rho_{10} \\ \rho_{20} \end{bmatrix} + \begin{bmatrix} \rho_{11} \\ \rho_{21} \end{bmatrix} \cdot t + \begin{bmatrix} \sigma_1 \\ \sigma_2 \end{bmatrix} \cdot d_{tv} + \sum_{p=1}^{10} A_p \cdot \begin{bmatrix} cv_{t-p} \\ vv_{t-p} \end{bmatrix} + \begin{bmatrix} \in_{t\tau 1} \\ \in t\tau 2 \end{bmatrix} \qquad (18)$$

$$\begin{bmatrix} c\tau_t \\ \tau v_t \end{bmatrix} = \begin{bmatrix} \gamma_{10} \\ \gamma_{20} \end{bmatrix} + \begin{bmatrix} \gamma_{11} \\ \gamma_{21} \end{bmatrix} \cdot t + \begin{bmatrix} \delta_1 \\ \delta_2 \end{bmatrix} \cdot d_{t\tau} + \sum_{p=1}^{10} A_p \cdot \begin{bmatrix} c\tau_{t-p} \\ \tau v_{t-p} \end{bmatrix} + \begin{bmatrix} \varepsilon_{t\tau 1} \\ \varepsilon_{t\tau 2} \end{bmatrix} \qquad (19)$$

The estimated λ_{trace} and λ_{max} values prove the existence of one long-term equilibrium for both the Lithuanian and Estonian data sets (the $r = 1$ hypothesis cannot be rejected for either of the statistics (see Tables 5 and 6 in the Appendix)). When dealing with modified data (involving the structural break term), the λ_{trace} values are lower, but they still support the latter conclusion (Table 7 in the appendix), so statistically significant cointegrating relationships in these models cannot be denied. Thus, the AR(1) process with A structural shift is not proved for these data sets, although the 5th observation (8 January 2010) is assumed to be a potential breakpoint in the NASDAQ OMX Tallinn series, and the 11th observation (15 January 2010) seems to mark a shift in the Vilnius stock exchange. A look at Fig. 5 reveals the variable modification method to be a probable factor driving it: whereas the logarithmic function shows a greater growth rate at the beginning of the period, the logarithm of cumulated variable values is supposed to demonstrate fewer rapid changes in the middle or in the end of the sample.

Otherwise, the index value dynamics (Figs. 3 and 4) underwent a similar jump: in the first two weeks of 2010, the Tallinn stock exchange was a world-leader by the index growth rate (+20.6%). This can be explained by the January effect or especially optimistic expectations inspired by the outstanding Baltics' export indicators and more courageous economic recovery forecasts for this region. According to the finance analysts, the news about prospective euro adoption in Estonia was one of the reasons, too (Čiulada, 22.01.2010). A one-year reaction lag supports the idea of the impact of the expectations which prove to be essential on the financial markets. Although this influence may decrease after the news announcement, it affects the market's moves incrementally. The limited power of the Chow test or Perron's test is to a certain degree explained by the volatility and formulation of expectation specifics observed on the financial markets.

Whereas the index value dynamics (Figs. 3 and 4) prompts at least two more potential trend changes, it was decided to cut the sample by the first 50 observations and, after delivering this partial elimination of the impact of variables' structure on the estimation results, to proceed investigating the local structural breaks. The evaluation is based on the VAR(7) model for the NASDAQ OMX Tallinn and on the VAR(2) for NASDAQ OMX Vilnius data sets, which after the sample cut exhibit a lower variance (without January-located stock exchange galloping) and a consequently decreasing number of lags necessary to eliminate the autocorrelation of residuals (see Tables 2 and 4 in the Appendix). The λ_{trace} and λ_{max} values indicate one independent cointegrating vector either before the break point estimation (see Tables 5 and 6 in the Appendix) or after it (Table 7).

This time, one of the possible breaks is supposed to happen on 19 March 2010 which is infamous for the tense situation in the financial markets, when the rescue plan for Greece suffering from the budget deficit and a full-blown sovereign debt was considered. While investors' concern about Greece and other Southern European countries' default was reflected by the rising CDS price and dropping stock index values, the EU and the IMF representatives agreed on the funding the plan on 25 March 2010.

Another change of the trend, which might have happened on 27 August 2010, also belongs to the fluctuations caused by the European sovereign debt crisis. While Germany was hesitating to provide assistance and rating agencies downgraded credit-ratings for Greece and Portugal repeatedly, the tensions in financial markets were not slackened (Mandaro, 28.04.2010). Therefore, most of the breakpoints mark the Baltic financial markets recovering after the rallies. This inference could be verified by a glance at the variable dynamics, too (see Fig. 3 and 4 again).

Whereas the sample was cut by 50 observations again, some other points appeared to be outstanding (Table 8). All of them could be attributed to the reaction of financial markets to European governments' struggle to cope with financial problems, but they also indicate the markets' recovering after the economic downturn as well (see Fig. 6). For instance, messages about the possible debt restructuring boosted interest yields of governments' debt (Kennedy, 2011.05.17). The break on 10 October 2010 is close to that of 5 October when the prime-minister of Ireland confirmed bigger than expected bail-outs for the banking sector. On the other hand, the Bank of Japan lowered its policy interest rate and said it would take further easing measures (Kennedy, 2010.10.05). This announcement was followed by the stock markets' gains in Western Europe, and consequently it took some further steps upwards in the Baltics. However, it is obvious that changing the trading and clearing currency to euro and the forthcoming euro adoption

TABLE 8. **The estimated breakpoints**

	Break point (BP)	Trade vol. change	Index change	Modified series	Possible causes	Table of λs in the Appendix
Lithuania	2010.01.15	Positive	Positive	Cointegrating relationship	Macro-situation, the January effect, news about euro adoption	5, 6, 7
	2010.08.27	Positive	Positive	Cointegrating relationship	Euro adoption, sovereign debt crisis	6, 7
	2011.05.17	Positive	Positive	Cointegrating relationship	Sovereign debt crisis (the threat of debt restructuring)	6, 7
Estonia	2010.01.09	Positive	Positive	Cointegrating relationship	Macro-situation, the January effect, news about euro adoption	5, 6, 7
	2010.03.19	Negative	Positive	AR(1) with a BP	Sovereign debt crisis (focused on Greece)	6, 7
	2010.06.04	Negative	Positive	Cointegrating relationship	Sovereign debt crisis	6, 7
	2010.10.11	Positive	Positive	AR(1) with a BP	Euro adoption, sovereign debt crisis (focused on Ireland)	6, 7

Source: NASDAQ OMX Baltic, MarketWatch, author's estimation and interpretation.

did not have a far-gone and pervasive impact on either the stock trading volumes or the index values. Besides, even in the presence of a negative change in trading volume, it is not necessarily followed by a slump of indices (see Table 8). Also, the negative external shocks had a greater impact on investors' attitudes and consequently on the financial markets than the positive ones. The fact that in foreign investors' eyes foreign news had outweighed announcements about local events seems reasonable.

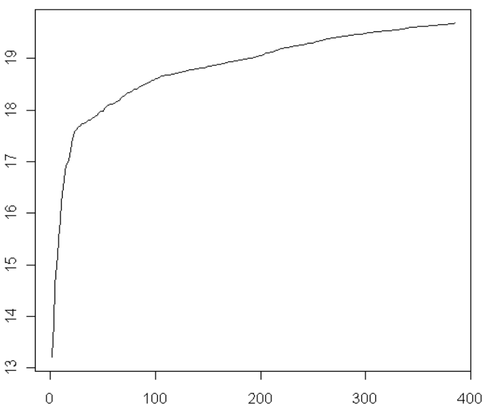

FIG. 5. **The logarithm of cumulated stock trading volume of the NASDAQ OMX Tallinn**

Source: NASDAQ OMX Baltic, author's calculation.

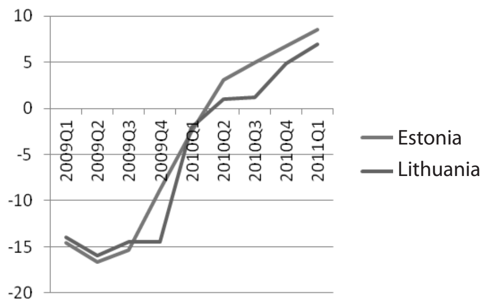

FIG. 6. **The GDP index change, % (quarter over quarter)**

Source: Statictics Lithuania, Statistics Estonia.

Besides, only two points – 19 March and 11 October 2010 are confirmed to be statistically significant. October 11, 2010 is the closest point to the euro adoption, but it is still not a week before 1 January 2011. It is quite possible that investors had been allocating their portfolios earlier, but perhaps it could be concluded to be a mix of the macroeconomic situation and optimistic expectations (including the forthcoming euro adoption) that could have lifted the markets upwards. The rest of the detected potential structural breakpoints should not be underestimated, even if they had not let to reject the cointegrating relationship hypothesis. They could be used for the analysis of markets' movements which might appear to be slighter than the structural breaks, but could also provide some useful insights into the markets' dynamics.

The other point is a difference between the potential structural points' allocation in the Lithuanian and Estonian time series. Although the stock trading volumes and index value dynamics seem parallel (see Figs. 1–4), the trend changes are estimated to occur at different time periods. This could be explained by rather distinct autoregressive processes of the aforementioned series, which could have had a marginal but critical impact on the estimation results.

To sum up, the first *à la* trend change, which had happened a year before the euro adoption in Estonia, is closely related to the macroeconomic situation in the Baltics and in the rest of the world, not to Estonian affairs only. The others mark the dates of the financial market response to negative shocks (in most cases – recovering after them).Thus, the euro adoption in Estonia and changing the trading and clearing currency to euro in Lithuania had no statistically significant influence on the trading volume and index value dynamics on the local stock exchanges, e.g., on attracting more foreign investors to the financial market or encouraging the actual ones to increase their investment in the Baltics.

All estimations are delivered using the *R* code.

Conclusions

The Chow test, which was used to investigate parameter stability in the period between 4 January 2010 and 18 July 2011, delivered no statistically significant evidence for rejecting the null hypothesis either for the Tallinn or for the Vilnius stock exchange data. Therefore, this test prompted to come to the conclusion that the euro adoption and changing the trading currency to euro had no statistically significant impact on attracting new investors to the NASDAQ OMX Baltic stock exchange. This finding might be explained by the limited power of the Chow test when assessing the structural shifts on volatile and expectations-influenced time series (e. g., financial indicators), because, as was confirmed by Lütkepohl et al. test, the potential breakpoints might appear to differ from hypothetical ones.

Perron's test also denied the statistical significance of hypothetical structural breaks and provided some controversial results of unit root testing: the stock trading volume series were confirmed to be I(1), in contrast to the ADF test. This mismatch reveals the complexity of the financial data and the possible critically different reactions to a slightly different structure of the generating processes (lags, a trend term, etc.).

In most cases, performing a test of structural break at an unknown point of time did not reject the hypothesis about cointegrating relationships between the index value and the cumulated stock trading volume; therefore, it led to a statistically significant denial of the hypothesis which declares that the euro adoption boosts investment in stocks. The only breakpoint (11 October 2010), which seems to be more related to the forthcoming euro adoption, is still too far to be excluded from the conjuncture of the other factors that were driving the Baltics' economies upwards in the meantime.

The other potential breakpoints, which have not been confirmed as statistically significant, revealed the extent of the impact of expectations on the financial markets. There was observed a one-year-long impact of prospective events (e. g., the euro adoption), which played a significant role in driving the financial markets as real-time statistical announcements do. Most of the potential breakpoints mark the recoveries after negative external shocks. Thus, the conclusion is that, in the foreign investors' eyes, the euro circulation in the NASDAQ OMX stock exchanges in Vilnius and Tallinn was not as significant as the news about the economic and political instability abroad. It is obvious that the potential possibility to attract more investment in the Baltics' financial market was wasted accidentally. Since the daily trading volume and index value change can experience a statistically significant gain only because of an additional capital flow from institutional foreign investors, it seems to be reasonable to interpret the local news to be faded by the world news which are escalated globally.

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APPENDIX

TABLE 1. **Information criteria (full sample; NASDAQ OMX Vilnius)**

	AIC(n)	HQ(n)	SC(n)	FPE(n)
1	-18.2827	-18.2569	-18.2178	1.15E-08
2	-18.7731	-18.7301	-18.6649	7.03E-09
3	-18.8473	-18.7871	-18.6959	6.53E-09
4	-18.8764	-18.7989	-18.6816	6.34E-09
5	-18.8777	-18.7831	-18.6398	6.33E-09
6	-18.9445	-18.8327	-18.6633	5.92E-09
7	-18.9678	-18.8387	-18.6433	5.79E-09
8	-18.9496	-18.8034	-18.5818	5.89E-09
9	-18.9778	-18.8144	-18.5668	5.73E-09
10	-18.9977	-18.8170	-18.5434	5.62E-09
11	-18.9879	-18.7900	-18.4903	5.67E-09
12	-19.0515	-18.8364	-18.5106	5.32E-09
13	-19.0356	-18.8033	-18.4515	5.41E-09
14	-19.0268	-18.7773	-18.3994	5.46E-09
15	-19.2335	-18.9668	-18.5628	4.44E-09
16	-19.2260	-18.9421	-18.5121	4.47E-09
17	-19.2257	-18.9246	-18.4685	4.48E-09
18	-19.2303	-18.9119	-18.4298	4.46E-09
19	-19.2360	-18.9005	-18.3923	4.43E-09
20	-19.2415	-18.8887	-18.3545	4.41E-09

Source: NASDAQ OMX Baltic, author's calculation.

TABLE 2. **Information criteria (NASDAQ OMX Vilnius)**

	AIC(n)	HQ(n)	SC(n)	FPE(n)
1	-20.904	-20.875	-20.831	8.35E-10
2	-20.918	-20.870	-20.798	8.23E-10
3	-20.899	-20.832	-20.730	8.39E-10
4	-20.885	-20.798	-20.668	8.51E-10
5	-20.874	-20.768	-20.608	8.61E-10
6	-20.852	-20.727	-20.539	8.79E-10
7	-20.835	-20.690	-20.473	8.94E-10
8	-20.818	-20.654	-20.408	9.10E-10
9	-20.801	-20.618	-20.343	9.26E-10
10	-20.788	-20.586	-20.282	9.38E-10
11	-20.777	-20.555	-20.222	9.48E-10
12	-20.756	-20.515	-20.154	9.68E-10
13	-20.739	-20.479	-20.088	9.85E-10
14	-20.721	-20.441	-20.022	1.00E-09
15	-20.698	-20.400	-19.951	1.03E-09
16	-20.698	-20.380	-19.903	1.03E-09
17	-20.686	-20.349	-19.842	1.04E-09
18	-20.723	-20.366	-19.831	1.00E-09
19	-20.701	-20.326	-19.761	1.02E-09
20	-20.682	-20.287	-19.694	1.05E-09

Source: NASDAQ OMX Baltic, author's calculation.

TABLE 3. **Information criteria (full sample; NASDAQ OMX Tallinn)**

	AIC(n)	HQ(n)	SC(n)	FPE(n)
1	-18.7410	-18.7156	-18.6769	7.26E-09
2	-19.2809	-19.2384	-19.1740	4.23E-09
3	-19.3774	-19.3179	-19.2278	3.84E-09
4	-19.3675	-19.2911	-19.1752	3.88E-09
5	-19.3611	-19.2677	-19.1260	3.90E-09
6	-19.3706	-19.2602	-19.0928	3.87E-09
7	-19.3706	-19.2432	-19.0501	3.87E-09
8	-19.3544	-19.2101	-18.9912	3.93E-09
9	-19.3565	-19.1951	-18.9505	3.92E-09
10	-19.3914	-19.2130	-18.9426	3.79E-09
11	-19.4325	-19.2372	-18.9410	3.64E-09
12	-19.5327	-19.3204	-18.9985	3.29E-09
13	-19.5137	-19.2844	-18.9367	3.35E-09
14	-19.5639	-19.3176	-18.9442	3.19E-09
15	-19.6146	-19.3513	-18.9521	3.03E-09
16	-19.5974	-19.3172	-18.8922	3.09E-09
17	-19.5838	-19.2866	-18.8359	3.13E-09
18	-19.5716	-19.2574	-18.7809	3.17E-09
19	-19.6219	-19.2907	-18.7885	3.01E-09
20	-19.6139	-19.2657	-18.7378	3.04E-09

Source: NASDAQ OMX Baltic, author's calculation.

TABLE 4. **Information criteria (NASDAQ OMX Tallinn)**

	AIC(n)	HQ(n)	SC(n)	FPE(n)
1	-20.3358	-20.3073	-20.2645	1.47E-09
2	-20.3249	-20.2774	-20.2060	1.49E-09
3	-20.3096	-20.2431	-20.1432	1.51E-09
4	-20.2860	-20.2005	-20.0720	1.55E-09
5	-20.2800	-20.1755	-20.0185	1.56E-09
6	-20.2657	-20.1423	-19.9567	1.58E-09
7	-20.3204	-20.1780	-19.9639	1.50E-09
8	-20.3107	-20.1493	-19.9066	1.51E-09
9	-20.2926	-20.1122	-19.8410	1.54E-09
10	-20.2900	-20.0906	-19.7908	1.54E-09
11	-20.3035	-20.0851	-19.7567	1.52E-09
12	-20.2823	-20.0449	-19.6880	1.56E-09
13	-20.2670	-20.0106	-19.6252	1.58E-09
14	-20.2499	-19.9745	-19.5606	1.61E-09
15	-20.2362	-19.9418	-19.4993	1.63E-09
16	-20.2341	-19.9207	-19.4496	1.63E-09
17	-20.2345	-19.9021	-19.4025	1.63E-09
18	-20.2404	-19.8890	-19.3608	1.62E-09
19	-20.2200	-19.8496	-19.2929	1.66E-09
20	-20.1989	-19.8096	-19.2243	1.69E-09

Source: NASDAQ OMX Baltic, author's calculation.

TABLE 5. **λ_{trace} statistics**

The break point	2010.01.15	2010.01.09	Critical values		
			10%	5%	1%
r< = 1	2.29	2.62	10.49	12.25	16.26
R = 0	76.55	31.77	22.76	25.32	30.45

Source: NASDAQ OMX Baltic, author's estimation.

TABLE 6. **λ_{max} statistics**

The break point	2010.01.15	2010.01.09	2010.03.19	2010.06.04	2010.10.11	2010.08.27	2011.05.17	Critical values		
								10%	5%	1%
R = 1	2.29	2.62	1.91	9.80	6.12	4.29	4.57	10.49	12.25	16.26
R = 0	74.26	29.16	30.36	14.69	14.18	84.27	43.60	16.85	18.96	23.65

Source: NASDAQ OMX Baltic, author's estimation.

TABLE 7. λ_{trace} **on the modified data**

The break point	2010.01.15	2010.01.09	2010.03.19	2010.06.04	2010.10.11	2010.08.27	2011.05.17	Critical values		
								10%	5%	1%
r< = 1	2.96	3.54	0.55	0.83	2.60	6.20	2.97	5.42	6.79	10.04
R = 0	32.36	19.13	4.65	16.51	13.84	25.94	18.41	13.78	15.83	19.85

Source: NASDAQ OMX Baltic, author's estimation.