Economic Forces, Sentiment and Emerging Eastern European Stock Markets

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Abstract

The aim of the current study is to explore the effect of macroeconomic news on stock returns in Eastern European countries, combining market and macroeconomic data over the period of 2000-2009, during which the markets experienced excessive optimistic and pessimistic episodes. Hypothesising the asymmetry in stock price responses to good and bad news, we seek to test its degree under the specific market conditions. The error correction models for each country are extended with fixed effects panel data specification, for capturing the cross-sectional effects of the state of the market on the return responses to macroeconomic news. The key methodological problem addressed in recent research is how to relate daily stock data and monthly macroeconomic data. The aggregation of stock returns to monthly averages has several advantages over macroeconomic data disaggregation to irregular frequencies or calendar methods. Since the macroeconomic data is large[1], we exploit data mining techniques to reasonably fit the monthly averages of the stock returns and panel data analysis to capture the common patterns typical to Eastern European stock markets. Consistently with prior empirical evidence, we find that the market reacts negatively to news when the market and/or economy is strong, but macroeconomic news is more of a long-term driver, and, therefore, affects trends more than short-run fluctuations.

JEL classification codes: **G14, G12**
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[1] By "large" data set we mean the number of different variables tried, a cross-section of 10 countries and a lag structure. Therefore backward stepwise methods are simply unfeasible, and the all model search, even for variables and 2 time "ticks," will make $2^{1+15\cdot 2}$ variants for the cointegration equation to choose from.

1. Introduction

Emerging Eastern European stock markets have experienced a rapid growth for the last decade and have become an important part of the common European stock market. Although a large body of current research examines the effect of macroeconomic news on stock prices in the developed markets, only a few relate it to the state of the economy or the state of the market. Stock return responses to macroeconomic news under varying market conditions in the emerging Eastern European markets remain largely unexplored. This paper aims to fill in this sizeable gap by providing evidence on the asymmetry in market responses to good and bad macroeconomic news under certain economic and market conditions in the stock markets of new members of European Union. We use market data and macroeconomic data of Eastern Europe countries, including Lithuania, Latvia, Estonia, Romania, Bulgaria, Hungary, Slovenia, Slovakia, the Czech Republic, and Poland over the period of 2000-2009, during which the above-mentioned emerging markets experienced excessive optimistic and excessive pessimistic episodes. We focus on the aggregate market reaction to macroeconomic news across countries and aim to identify the pervasive economic forces that affect all asset prices.

To capture the asymmetry of market responses to news under the different states of the economy and/or the market, we must distinguish between "good" and "bad" times. We define "good" times, when the stock market experiences optimistic episodes and/or the economy is in an expansion; and "bad" times, when it experiences pessimistic episodes and/or the economy is in a recession. Prior research reports the use of the smooth transition regression (STR) model that allows for smooth change from good to bad times, with the threshold estimated endogenously (Laakkonen and Lanne, 2008). They demonstrate that this model can be generalised to more than two regimes. However, further empirical tests suggest the use of a two-regime model; as it shows better performance. We suggest a more natural alternative of how to include smooth transitions from one state to another – the mixed effects model with a transition mechanism dependent on economic sentiment index (ESI) or a similar proxy variable. Mixed effects models may be interpreted as models with random parameters, when key coefficients are conditional on the sentiment index and provide smooth and more natural transitions from one state to another.

Assuming the more pronounced effect of macroeconomic factors than microeconomic ones on stock returns in the long run, we contribute to the strand of the literature that examines the relations of macroeconomic forces and the stock market. We also hypothesise the asymmetry in market responses to macroeconomic news under the prevailing market sentiment and therefore, contribute to the behavioural finance literature. Most importantly, we extend the existing empirical evidence in the field, by providing statistically and economically significant evidence in emerging markets.

We find that macroeconomic news (interest rate, industrial production index, exports relative to imports of goods, industrial turnover gap and market sentiment) have a significant impact on stock returns across the Eastern European stock market. The nature of macro information affects the direction of subsequent changes in index returns. Allowing for variation in the economic and market conditions, we find that market sentiment affects the asymmetry in responses of index returns. The returns are influenced by asymmetric reaction that lowers gap impact on return variation in "good" times and significantly increases it in "bad" times. Assuming commonality and cointegration in the movements of Central and

Eastern Europe, we employ the panel error correction model. The asymmetry effects are found to be significant across markets and at cross sections of the Bulgarian, Czech, Polish and Romanian index returns.

The paper is structured as follows. At first, we discuss the underlying research problem in the prior literature. The next section describes the data and methodology. Then we present the empirical tests and discussions and conclude in the last section.

2. Literature Review

2.1. The Interaction between Economic Forces, Stock Market and Business Conditions

Commencing with the study of Fama and French (1989), scholars examine economic entities and their interaction that have an impact on market responses to macroeconomic news. Fama and French (1989) relate time-varying expected returns on bonds and stocks to business conditions represented by the default spread variable, which is high during recessions and low during expansion of the economy. Business cycle effects were examined in stock markets by McQueen and Roley (1993), Veronesi (1999), Flannery and Protopapadakis (2002), Conrad et al. (2002), Adams et al. (2004), Boyd et al. (2005), and Andersen et al. (2007). McQueen and Roley (1993) find more pronounced relations of stock returns and fundamental macro news after allowing for various phases of the business cycle. They report negative market responses to higher real economic activity when an economy is strong and associate positive price responses with the effect of the weak economy. They attribute the varying responses of stock prices across economic states to expected cash flows. The responses of equity discount rate proxies to new economic information are not significantly different across economic states. In contrast, unanticipated increases in economic activity in a weak economy raise expectations about future economic activity and cash flows. On the other hand, the same information in a strong economy does not lead to higher expected cash flows. The source of the varying response of stock prices across economic states appears to be expected cash flows, not necessarily connected to the macroeconomic issues. Therefore, the responses of equity discount rate proxies to new economic information are not significantly different across economic states.

Veronesi (1999) develops a rational expectations' equilibrium model of asset prices; where, the drift of fundamentals shifts between two unobserved states of the market at random times. The high level of investors' uncertainty about the state of the market affects the volatility of stock returns, and investors require the extra discount. When times are good, bad news makes investors increase the discount over expected future returns in order to bear the risk of higher uncertainty. This results in the greater price reduction, due to bad news, than the reduction in expected future returns. And good news in bad times makes investors increase the expected future returns and also increase the discount in order to hold an asset. Hence, the increase in price is lower than the increase in expected future dividends. Therefore, this conceptual paper introduces the directional asymmetry in market responses to bad and good news under different market conditions and gives a new direction for further research.

Andersen et al. (2007) supplements the above-mentioned theoretical model with empirical evidence. They report state-dependent market responses to macroeconomic news in the U.S., German and U.K. markets. Bad news has a positive impact during expansions and a negative impact during recessions. They explain this by temporal variation in the competing "cash flow" and "discount rate" effects for equity valuation. Another explanation, in our opinion, could be linked to the phenomenon of speculative bubbles, which cause huge distortions to the market fundamentals, expecting that stock market's "newbies" will not interpret the bad news signals correctly in the short run. Following McQueen and Roley (1993), Adams et al. (2004) use state classification regression to define the economic states. They test for coefficient stability across high and low economies and prove the stock-inflation relationship to be state dependent. Flannery and Protopapadakis (2002) employ the GARCH model of daily stock returns for estimating the impact of 17 macroeconomic announcements on stock returns and their conditional volatility. Consistently with Adams et al. (2004), they find strong stock return responses to inflation and money growth announcements. Boyd et al. (2005) find that an announcement of unemployment is "good" news for stocks during an economic expansion and "bad" news during economic contractions. They relate the increase in unemployment to the decline in interest rates and to the decline in future corporate earnings and dividends. A decrease in interest rates is "good" news, whereas a decline in dividends and corporate earnings is "bad" news for equity markets. They conclude that one of these two effects dominates depending on the state of the economy. Information on interest rates prevails during expansion, and information on future corporate earnings prevails during contractions.

Generally, all the above mentioned studies focus on the aggregate market reaction to macro news under certain economic conditions and find that investors are more sensitive to bad news than good news. Conrad et al. (2002) explore the interaction between aggregate market conditions and responses to firm-specific news, and find that stock price responses to negative earnings shocks increase when the market is optimistic. Although the authors focus on the market reaction to firm-specific news that are excluded from our analysis, the state of the market (or market sentiment) introduced in their study motivates our interest and opens for further research in emerging Eastern European stock markets that remain largely unexplored.

2.2. Economic Forces and Emerging Stock Markets

Extensive research explores equity return responses to news about underlying fundamental economic conditions in Eastern Asian emerging markets; whereas, Eastern European emerging markets receive far less scholarly attention. The comprehensive literature overview suggests dividing it into three strands.

The first group of studies covers Eastern European emerging markets. Hanousek and Filler (2000) Babecký et al. (2008) and Samitas and Kenourgios (2007) explore how macro variables affect stock returns in Poland, the Czech Republic, Slovakia and Hungary. Samitas and Kenourgios (2007) examine whether current and future industrial production and interest rate variables explain long and short run stock returns. Samitas and Kenourgios (2007) and Babecký et al. (2008) focus on the integration of these four markets into euro currency union; whereas, Hanousek and Filler (2000) analyse interconnections of stock prices fluctuations and economic variables. Hypothesising equity markets to be semi-strong efficient, they find that economic factors (broad money supply, imports, exports, and foreign capital inflow and industrial production) generate contemporaneous stock price changes in the Czech market. Since lagged values have no impact on stock returns, this supports the conventional concepts of semi-strong market efficiency in the Czech market. On the other hand, the effect of lagged economic factors on stock prices in Hungarian, Polish and Slovakian equity markets proves them to be inefficient. We doubt this statement however, because in our opinion, lagged influence of macroeconomic data may also be linked to the announcement calendar.

The second group of studies focuses on Eastern Asian and African emerging markets. Mookerjee and Yu (1997) test for information inefficiencies in the equity market of Singapore in the short and long run; whereas, Frimpong (2009) performs similar analysis in the equity market of Ghana. They find co-integrated relationships of the lagged macro factors and stock returns that suggest market inefficiency in both Singaporean and Ghanaian markets. Ibrahim and Aziz (2003) and Asmy et al. (2009) also employ co-integration analysis and vector autoregression techniques. However, they explore the linkages between stock prices and macroeconomic variables in the Malaysian equity market. Both studies report positive coefficients for inflation rate and money supply and negative for exchange rate in the regression models, both for pre- and post-crisis periods. Kandir (2008) examines how macro variables and MSCI World Equity Index returns affect portfolio returns in the Turkish stock market. They find that the exchange rate and MSCI affect all portfolio returns and fail to prove the impact of industrial production, money supply and oil prices on portfolio returns. He reports inconclusive results for inflation rate effect on portfolio returns. Kwon et al. (1997) explore the linkage of the South Korean stock market to macro variables and find that stock returns are more sensitive to real economic activity and international trading activities measured in terms of foreign exchange rates, trade balance, the money supply, and the production index, rather than inflation or interest rate variables. This suggests the use of different investment strategies in emerging markets as investors' perceptions of stock differ from those of investors' in the developed markets.

The third group of studies explores the relationships between stock returns and macroeconomic variables in more than one emerging market. Gay (2008) explores the exchange rate and oil price effect on stock indices' returns for Brazil, Russia, India, and China (BRIC) – employing the Box-Jenkins ARIMA model. He fails to find significant relationships and attributes it to the influence of other macroeconomic factors. Furthermore, he finds no significant relationships between past and present market returns, suggesting the weak-form of efficiency for BRIC markets. Muradoglu et al. (2000) explore 19 emerging markets and Bilson et al. (1999) investigate 20 emerging markets. Muradoglu et al. (2000) analyse the effect of exchange rates, interest rates, inflation and industrial production on stock returns. They argue that the macro factors' significance is determined by the size of the respective market and the level of integration into the world market. Bilson et al. (1999) employ money supply, inflation, industrial production and exchange rates as proxies for local risk sources to explore the degree of commonality across 20 emerging stock. They find that good prices and real activity have limited ability to explain the variation in stock returns. Money supply has an explanatory power, but it is significantly lower compared to the exchange rate and world market returns. Furthermore, they find little evidence of commonality when emerging markets are considered collectively. However, they prove considerable commonality at the regional level.

Summarising the previous findings, we should notice that, while there is a growing interest in how the return generating process depends on macroeconomic factors in developed markets and emerging Eastern Asian markets, there has been surprisingly little empirical work on Eastern European emerging markets. Prior research provides mixed evidence on the importance of macroeconomic factors in different emerging markets. Narrowly and broadly defined, money supply is the important factor for the equity markets of the Czech Republic, Ghana, Singapore, South Korea and Malaysia, but insignificant for the Turkish equity market. The foreign exchange rate factor explains variation in stock returns in Turkish, South Korean, Malaysian and Ghanaian markets, as well as in 20 emerging markets explored in Bilson et al. (1999), but not for BRIC markets. Inflation rate is inconclusive for the Turkish stock market, but seems to have an impact on the stock returns in South Korea, Malaysia, Ghana and 20 emerging markets explored in Bilson et al. (1999). Changes in industrial production are also not important for Turkish stock returns, but are significant for the Ghanaian (Frimpong, 2009) and the Malaysian (Kwon et al., 1997) markets and for most of the emerging markets explored in the Muradoglu et al. (2000) and Bilson et al. (1999) studies.

3. Data and Methodology

Prior research suggests a wide variety of relevant macroeconomic factors that influence stock returns, both in the developed and the emerging markets, therefore it is worth to begin with a general discussion of data and econometric instruments used in the paper.

The analysed dataset covers stock market and macroeconomic data of all ten new member-states of the European Union countries in Central and Eastern Europe, including: Lithuania (LT), Latvia (LV), Estonia (EE), Romania (RO), Bulgaria (BG), Hungary (HU), Slovenia (SI), Slovakia (SK), the Czech Republic (CZ), and Poland (PL), over the period of 2000-2009, during which the above-mentioned emerging[2] markets experienced excessive optimistic and excessive pessimistic episodes.

There is difficulty in trying to link high frequency (not to mention functional, intraday) stock returns' data to macro observations. Macro level data, especially related to prices, production, consumption, and trade, is not observed on a higher than monthly basis. Since the number of working days (and also weeks) within a month is irregular, there is no clear algorithm to temporally disaggregate the macroeconomic time series to a daily frequency but only some sort of naïve interpolation. On the other hand, an aggregation of stock market data would lead to a more or less standard (compared to estimation that exploits both high and low frequency information in one regression) econometric analysis. For aggregation we suggest monthly averages, although first, last, weighted averages, minimum, maximum transformations are also possible.

Another alternative option would be to use a data announcement calendar. However, this narrows the focus of the current study on the immediate or short-run impacts of the announcements and usually the magnitude of impacts is left aside. We instead assume that macroeconomic data has long-run impacts, but due to the lags in announcements this is reflected in stock market averages being lagged up to several months. We suggest however

[2] Emerging markets are classified by the Dow Jones classification. Recently Slovenia, as the only country that already adopted the euro, switched to the developed status, but in most of the analysed period all ten countries can be characterised as having emerging markets.

not digging deeper than for 1 month. First, this seems relevant to produce feasible estimates for cointegration relationship applying backward stepwise selection. Secondly, monthly data usually experiences little lags in the announcements compared to quarterly or yearly ones, and thus, markets react to the actual values much sooner.

All variables used in the analysis could be formally divided into 3 groups: market data, sentiment data – the focus of the research, and other variables in which inclusion is uncertain, but motivated by prior research.

Therefore, in the empirical part we analyse the monthly averages of total return indices (*Y*), normalised to 1 at the first period. The other market data variables such as market capitalisation (*MC*), and share price average (*ST*) are explanatory and characterise typical market trends. To measure market sentiment we employ one of the following indicators: economic sentiment indicator (*ESI*), consumer confidence index (*CCI*) or industry confidence index (*ICI*). The last group of variables covers exports (*X*) and imports of goods (*M*), or their log-difference (log(*X*/*M*)); and either nominal effective exchange rate (*NEER*) or real effective exchange rate (*REER*) adjusted for effects of inflation as the proxies for the markets openness and competitiveness; harmonised consumer price index (*HCPI*) or the inflation rate based on it; 1-month interest rate for money market instruments (*R*); broad money supply (*M2*) in million euro, which includes *M1* and foreign currency deposits; unemployment rate (*UR*); industrial production index (*IPI*) or industry turnover index (*ITI*) as the measure of industries fundamental performance, and producer price index (*PPI*). The latter is suggested to be included into the focus group of the regressors as the key variable describing the current state of the economy. We also tried log-difference of industry turnover index from the trend (log(*ITI*/*ITI**)) following the idea presented in McQueen and Roley (1993) as the proxy for the economy state in state classification regression.

All variables, except for total return indices taken from DataStream, are sourced from the Eurostat database. Eurostat contains the macro information for each EU member and insures methodological comparability of the time series. It is also worth mentioning that due to some missing observations in the beginning of the sample period (especially the data for the Slovenian economy and money aggregates) the panel data is unbalanced.

Differing from other studies, we do not apply seasonal corrections for two reasons: first, because this may distort short-run impacts on stock returns, and secondly, any smoothing of unusual behaviour may introduce autoregression in residual terms that in turn requires more sophisticated econometrics (iterative procedures) to be implemented. We use the Hodrick-Prescott filter to extract for the business cycle component. However, we apply an optimal smoothing parameter to minimise the risk of the Slutsky effect being present in the business cycle. The idea of choosing the optimal smoothing parameter is based on obtaining the best Hodrick-Prescott approximation to the ideal filter, minimising compression and leakage effects of the former by choosing a smoothing parameter to minimise the Q statistic (Pedersen, 2001).

Following previous empirical studies, it is known that at the aggregate level the weighted geometrical means (multiplicative models) experience little aggregation bias compared to the weighted averages. Besides, if additive cointegrated models assume that absolute differences behave stationary, empirical facts contradict this assumption since proportions (structures) are more likely to be stationary in the long-run. Therefore, most of the variables, excluding rates and confidence indices, are taken in natural logs, and we assume a log-linear form in the long-run (static) equation:

$$\log y_t = \alpha_0 + \sum_{i=1}^{K} \sum_{j=0}^{1} \alpha_{ij} \log z_{t-j}^{(i)} + \varepsilon_t \,, \qquad (1)$$

where the right-hand side includes all relevant regressors detected by the data mining approach, which are also allowed to be lagged up to one month, due to possible delay in the data announcement calendar and not efficient and fast enough adoption of macroeconomic information.

Since the long-run stationary (non-spurious) relationship in (1) could exist only if the data of the same order is integrated, the first step is to estimate the integration order of the series. To perform this econometric step, we apply Kwiatkowski, Phillips, Schmidt and Shin (KPSS, 1992) and Zivot and Andrews (ZA, 1992) stationarity tests. We choose KPSS over Augmented Dickey-Fuller (ADF) or Phillips–Perron (PP) tests. The reasoning is the following, ADF and PP tests too often suggest unit roots, even if there is no clear evidence for them. So it is good to formulate the test, which is consistent under the null "stationary" hypothesis versus a wide range of non-stationary reasons. Since KPSS is also sensitive to structural breaks, we also performed the ZA test. By performing the ZA test, which is robust against the inclusion of structural breaks, we searched for possible breaks both in trend and an intercept term. The analysis of time series showed that most of the 169 time series (data for Slovenia's M2 is missing) are integrated of the first order ($I(1)$) processes, and only some (13 by ZA and 8 by KPSS, see Appendix) can be viewed as stationary in levels. Note, that stationary time series can still be used together with first order differences in the Error Correction Model (ECM). The latter is chosen as an appropriate model form to catch both the correction mechanism of short-run deviations from the long-run behaviour and the long run cointegration relationship (1) itself.

Seeking to lower the bias in estimates of the parameters, we estimate both cointegration and short-run parameters in one regression:

$$\Delta \log y_t = \gamma\alpha_0 - \gamma \log y_{t-1} + \gamma \sum_{i=1}^{K} \sum_{j=1}^{2} \alpha_{ij} \log z_{t-j}^{(i)} + \sum_{i=1}^{L} \sum_{j=1}^{2} \beta_{ij} \Delta \log z_{t-j+1}^{(i)} + \eta_t \,, \qquad (2)$$

or in more common form:

$$\Delta \log y_t = -\gamma \left(\log y_{t-1} - \alpha_0 - \sum_{i=1}^{K} \sum_{j=1}^{2} \alpha_{ij} \log z_{t-j}^{(i)} \right) + \sum_{i=1}^{L} \sum_{j=1}^{2} \beta_{ij} \Delta \log z_{t-j+1}^{(i)} + \eta_t \,, \qquad (3)$$

where γ represents the error correction parameter, Δ denotes the first order difference and the residual term η is assumed to be the white noise innovation process not correlated with ε from (1). Short run correction regressors may contain other variables not necessarily presented in (1), for instance; first difference of lagged dependent variable, some other $I(0)$ variables. Note, that we extensively test (by the Wald test) for the restriction $\gamma\,\alpha_{ij} = \beta_{ij}$, where possible, to lower the number of estimated parameters, and therefore, increase the degrees of freedom. Thus:

$$\gamma\alpha_{ij} = \beta_{ij} \Rightarrow \gamma\alpha_{ij} \log z_{t-j}^{(i)} + \beta_{ij} \Delta \log z_{t-j+1}^{(i)} = \beta_{ij} \Delta \log z_{t-j+1}^{(i)} \qquad (4)$$

Due to the super-consistency of the ECM, ordinary least squares' (OLS) estimates (2) can be validly estimated by the OLS, when the residual terms behave adequately (see statistics presented in the empirical section). Therefore, it is not much different from the usual regression. Note, that we put the results in the estimated form (2), not (3), but all the parameters are easily calculated from the estimated form, dividing the corresponding coefficients by the error correction parameter γ.

We suggest employing the general-to-specific methodology driven by backward stepwise selection for the data mining. The number of parameters to search in equation (1) is about 25, which is less than the time length, and the backward selection is, in principle, feasible. Finding out the cointegration relationships, we then search for short-run drivers, as in (2) or (3). This allows adding the variables found to be stationary in the integration analysis step. Although stepwise procedures are strongly criticised by statistical purists as being automated procedures for data mining, we look at the problem from another angle than a pure data mining application. We suggest dividing the dataset into the focus group variables, which are the purpose of the empirical research and which are suggested by theory as "must be included", and the auxiliary ones, the inclusion of which is doubtful, but still possible on the theoretical grounds. Therefore stepwise least squares regression taken with the set of macro variables explored in prior research can be regarded as the first approximation to unknown data generating process. The focus group of the variables includes those on market trends, industrial production and sentiment, the other variables are less clear. Thus, we can search for a statistically adequate model by a general-to-specific approach. The suggested split opens interesting empirical possibilities briefly discussed in conclusive remarks section, where we suggest some other data mining methods for further investigation (weighted average least squares appears to be a nice alternative to stepwise). On the other hand, we do not end up with the stepwise results (that are made on the log-linearised equation (1)), but augment them, including the interaction terms, namely products of sentiment indices with the macro drivers. This inclusion is inspired by the idea of the natural smooth transition from one state of the economy to another.

Finally, though the markets in our dataset belong to the European area and all are qualified as emerging, they differ with their development and sophistication level and, therefore, commonality in exposures across markets at first sight is not expected. This approach seems to be useful, when searching for common patterns in the panel data models. Thus, following the assumption of regional commonality, we test whether the 10 emerging Eastern European markets are cointegrated. The exploration of the common patterns in Eastern European stock markets lead us to the panel data fixed effects model. The use of a fixed effects model, rather than random effects, is based on the fact that time invariant unobserved components (fixed effects) are correlated with observed explanatory variables. The Hausman test statistic (19.53 with 7 degrees of freedom) suggests an application of the fixed effects model that is consistent with the data. The logic of ECM models for each country is carried out to this restricted system of equations. Moreover, the results from individual regressions help to identify the typical drivers of the Eastern European market.

4. Empirical Analysis and Discussions

Since formal tests (see Appendix), as well as graphical examination of the data series, showed that the logarithmically transformed data is mostly integrated of the first order, we have obtained a theoretical justification to search for cointegrating vectors. Backward stepwise selection of auxiliary variables has been performed, and then the augmented models have been estimated. The results of this selection after the augmentation for interaction terms are shortly summarised in Table 1 below. Each of the equations is followed by some goodness-of-fit statistics: adjusted coefficient of determination, which for ECM form is rarely as high as for static equation, but shows what part of stock market returns (recall, that the first difference of the logarithm in y is actually the log-return) is actually explained by the model; Schwarz information criterion (*BIC*), which allow to compare the non-nested models; Durbin Watson (*DW*) and Ljung-Box (*Q*) statistics that test if the residual term resembles a white noise process, the Dickey-Fuller statistic to test for stationarity of the residuals and the mean average percentage error *MAPE*(6) to see how well the model forecasts 6 months out-of-sample. For the latter, we shorten the estimation period by the last 6 months and after re-estimation dynamically forecast these points to compare them with the actual data. All are included, but some intercept terms' parameters are significant at the 5 per cent significance level. Below the parameter estimates, we provide standard deviations in parenthesis.

From Table 1, we draw some interesting conclusions about the similar impacts and behavior of the stock markets. Our primary attention has been focused on the influences related to the sentiment indices. Most of the equations have the economic sentiment index as the part of the cointegration relationship, with the highest elasticity observed for the Hungarian market. In some cases the countries are forming particular pairs: Slovakia and Slovenia, and Romania and Bulgaria, which are also similar by the auxiliary variables included. For instance, good news on the changes in a turnover gap in the Slovakia-Slovenia pair of countries generates a decrease in index returns, and good news on exports of goods, and an increase in consumer confidence affects the index returns positively; whereas, news on an increase in imports of goods seems to have an opposite response on index returns. Note, however, that although the Latvian and Lithuanian markets are known as being closely related, we find that different factors explain the variation in index returns in Lithuania.

Economic sentiment index for Bulgaria, Romania, Poland, and the Czech Republic enters not only the static equation, but has a significant indirect impact on the short run dynamics through the industrial turnover gap (ITI_t / ITI_t*) that represents the state of the economy and dependence on fundamental economic activity. The inclusion of an interaction term, namely the product of economic state and market sentiment variables, allows finding that the returns are influenced by an asymmetric reaction that lowers the gaps impact on returns' variation in "good" times and significantly increases it in "bad" times. Besides in the Czech and Polish markets, both the turnover gap and industrial production prove to be important predictors of index returns in the long run. However, an increase in industrial production implies positive stock returns. Whereas the larger is the turnover gap, the lower are the stock returns. The negative impact of a turnover gap on stock returns is mitigated by an indirect positive effect of economic sentiment that enters the static equation. As a result, the Czech and Polish markets exhibit smaller negative asymmetries compared to the Bulgarian and Romanian cases.

Table 1. Individual Error Correction Models for the Eastern European markets

Country	ECM equation
Bulgaria:	$\Delta\log(Y_t) = \underset{(1.95)}{3.94} - \underset{(0.09)}{0.68}\cdot\log(Y_{t-1}) + \underset{(0.22)}{0.44}\cdot\log(ESI_{t-1}) + \underset{(0.05)}{0.43}\cdot\log(MC_t) - \underset{(0.01)}{0.05}\cdot R_t -$ $- \underset{(0.007)}{0.018}\cdot NEER_{t-1} + [\underset{(4.63)}{19.54} - \underset{(0.99)}{4.17}\cdot\log(ESI_t)]\cdot\log(ITI_t/ITI_t^*) - \underset{(0.2)}{0.66}\cdot\log(PPI_{t-1}) + \eta_t$ *adj.* $R^2 = 0.64$, $BIC = -2.57$, $DW = 1.8$, $p(Q(4)) = 0.1$, $DF(\eta_t) = -7.4$, $MAPE(6) = 8.48\%$
Czech Republic:	$\Delta\log(Y_t) = \underset{(0.39)}{-1.73} - \underset{(0.02)}{0.05}\cdot\log(Y_{t-1}) + \underset{(0.005)}{0.012}\cdot UR_{t-1} + \underset{(0.09)}{0.37}\cdot\log(IPI_{t-1}) +$ $+ \underset{(0.05)}{0.17}\cdot\Delta\log(X_t/M_t) - [\underset{(5.09)}{14.296} - \underset{(1.09)}{2.99}\cdot\log(ESI_{t-1})]\cdot\log(ITI_{t-1}/ITI_{t-1}^*) + \underset{(0.02)}{0.05}\cdot\Delta(UR_t) + \eta_t$ *adj.* $R^2 = 0.34$, $BIC = -3.26$, $DW = 1.99$, $p(Q(4)) = 0.45$, $DF(\eta_t) = -10.42$, $MAPE(6) = 5\%$
Estonia:	$\Delta\log(Y_t) = \underset{(1.06)}{-2.71} - \underset{(0.02)}{0.05}\cdot\log(Y_{t-1}) + \underset{(0.26)}{0.61}\cdot\Delta\log(ESI_t) - \underset{(0.05)}{0.12}\cdot\log(ITI_t/ITI_t^*) -$ $- \underset{(0.54)}{1.48}\cdot\log(HCPI_{t-1}) + \underset{(0.76)}{2.103}\cdot\log(PPI_{t-1}) - \underset{(0.01)}{0.05}\cdot R_{t-1} + \underset{(0.04)}{0.08}\cdot\Delta\log(X_t/M_t) + \eta_t$ *adj.* $R^2 = 0.31$, $BIC = -2.62$, $DW = 1.86$, $p(Q(4)) = 0.31$, $DF(\eta_t) = -10.45$, $MAPE(6) = 6.48\%$
Hungary:	$\Delta\log(Y_t) = \underset{(3.96)}{-17.485} - \underset{(0.09)}{0.39}\cdot\log(Y_{t-1}) + \underset{(0.21)}{0.703}\cdot\log(M2_t) - \underset{(0.002)}{0.006}\cdot ICI_t + \underset{(0.03)}{0.08}\cdot UR_t +$ $+ \underset{(0.006)}{0.021}\cdot R_{t-1} + \underset{(0.44)}{2.02}\cdot\log(ESI_{t-1}) + \underset{(0.002)}{0.005}\cdot\Delta(CCI_t) + \eta_t$ *adj.* $R^2 = 0.37$, $BIC = -1.97$, $DW = 2.13$, $p(Q(4)) = 0.72$, $DF(\eta_t) = -8.45$, $MAPE(6) = 13.49\%$
Latvia:	$\Delta\log(Y_t) = \underset{(0.31)}{-1.38} - \underset{(0.03)}{0.09}\cdot\log(Y_{t-1}) + \underset{(0.09)}{0.51}\cdot\log(ESI_t) + \underset{(0.004)}{0.026}\cdot UR_{t-1} - \underset{(0.16)}{0.39}\cdot\Delta\log(ESI_{t-1}) +$ $+ \underset{(0.84)}{2.16}\cdot\Delta\log(PPI_{t-1}) - \underset{(0.002)}{0.012}\cdot NEER_t + \eta_t$ *adj.* $R^2 = 0.33$, $BIC = -2.71$, $DW = 1.79$, $p(Q(4)) = 0.73$, $DF(\eta_t) = -9.6$, $MAPE(6)=11.47\%$
Lithuania:	$\Delta\log(Y_t) = \underset{(0.19)}{-0.43} - \underset{(0.04)}{0.22}\cdot\log(Y_{t-1}) + \underset{(0.001)}{0.003}\cdot CCI_t - \underset{(1.18)}{2.31}\cdot\Delta\log(HCPI_t) + \underset{(0.04)}{0.09}\cdot\Delta\log(M_{t-2}) -$ $- \underset{(0.002)}{0.005}\cdot REER_t + \underset{(0.08)}{0.27}\cdot\log(IPI_{t-1}) + \underset{(0.09)}{0.23}\cdot\Delta\log(Y_{t-1}) + \eta_t$ *adj.* $R^2 = 0.36$, $BIC = -2.68$, $DW = 1.89$, $p(Q(4)) = 0.58$, $DF(\eta_t) = -9.46$, $MAPE(6)=12.02\%$
Poland:	$\Delta\log(Y_t) = \underset{(1.01)}{3.05} - \underset{(0.02)}{0.07}\cdot\log(Y_{t-1}) - \underset{(0.02)}{0.09}\cdot\log(ESI_t)\cdot\log(ITI_t/ITI_t^*) + \underset{(0.11)}{0.35}\cdot\log(ITI_t) +$ $+ \underset{(0.06)}{0.1}\cdot\Delta\log(X_t) - \underset{(0.11)}{0.35}\cdot\Delta\log(IPI_t) - \underset{(0.29)}{1.01}\cdot\log(HCPI_{t-1}) + \eta_t$ *adj.* $R^2 = 0.22$, $BIC = -2.89$, $DW = 1.89$, $p(Q(4)) = 0.38$, $DF(\eta_t) = -10.55$, $MAPE(6)=5.01\%$
Romania:	$\Delta\log(Y_t) = \underset{(0.92)}{-3.22} - \underset{(0.04)}{0.19}\cdot\log(Y_{t-1}) + \underset{(0.21)}{0.69}\cdot\log(ESI_t) + \underset{(0.002)}{0.004}\cdot NEER_t - \underset{(0.002)}{0.006}\cdot R_{t-2} +$ $+ [\underset{(3.39)}{9.29} - \underset{(0.74)}{2.06}\cdot\log(ESI_t)]\cdot\log(ITI_t/ITI_t^*) - \underset{(1.46)}{5.33}\cdot\Delta\log(HCPI_t) + \underset{(0.009)}{0.02}\cdot UR_{t-1} + \eta_t$ *adj.* $R^2 = 0.27$, $BIC = -1.83$, $DW = 1.72$, $p(Q(4)) = 0.54$, $DF(\eta_t) = -9.25$, $MAPE(6)=21.75\%$
Slovakia:	$\Delta\log(Y_t) = \underset{(0.44)}{-1.19} - \underset{(0.03)}{0.09}\cdot\log(Y_{t-1}) + \underset{(0.05)}{0.13}\cdot\log(ESI_t) - \underset{(0.05)}{0.2}\cdot\log(ITI_t/ITI_t^*) + \underset{(0.001)}{0.002}\cdot CCI_{t-1} +$ $+ \underset{(0.06)}{0.32}\cdot\log(X_t) - \underset{(0.06)}{0.22}\cdot\log(M_t) + \underset{(0.08)}{0.16}\cdot\Delta\log(Y_{t-1}) + \eta_t$ *adj.* $R^2 = 0.36$, $BIC = -3.26$, $DW = 1.66$, $p(Q(4)) = 0.31$, $DF(\eta_t) = -8.84$, $MAPE(6)=2.32\%$
Slovenia:	$\Delta\log(Y_t) = \underset{(1.51)}{-4.86} - \underset{(0.04)}{0.12}\cdot\log(Y_{t-1}) + \underset{(0.28)}{0.73}\cdot\log(ESI_t) + \underset{(0.18)}{0.34}\cdot\log(IPI_t) + \underset{(0.05)}{0.14}\cdot\Delta\log(X_t/M_t) -$ $- \underset{(0.12)}{0.36}\cdot\log(ITI_t/ITI_t^*) + \eta_t$ *adj.* $R^2 = 0.23$, $BIC = -3.16$, $DW = 1.83$, $p(Q(4)) = 0.91$, $DF(\eta_t) = -6.71$, $MAPE(6)=4.46\%$

Source: Authors' calculations

For Hungary, we find a strong direct consumer confidence impact on index return dynamics – not production driven. A similar conclusion holds for the Lithuanian market; where, the estimates of the consumer confidence index and harmonised consumer price index suggest a strong index return dependence on the domestic consumption. Considering the harmonised consumer price index in the short run (inflation pressure), we observe its impact for Poland, Estonia and Romania. Findings for negative inflation rate impact on stock returns are consistent with the theoretical and the previous empirical evidence.

Macroeconomic theory tells that a weaker domestic currency (dollar/domestic currency increases) exchange rate will boost exports, which in turn will lift employment and resulting in economic growth. This creates a favourable news perception on the stock market and thereby generates subsequent positive stock returns. Our findings do not support the hypothesis of a positive relationship between exchange rate and stock returns, the sole exception being the Romanian market. We find that exchange rates and producer price index affect the index returns in the long run; whereas, interest rate and market capitalisation affect them both in the long and short run. It is worth mentioning that empirical evidence is controversial, but our findings are in line with Mukherjee and Naka (1995) and Ibrahim and Aziz (2003), who also report the negative relationships between the exchange rate and stock returns

The theory states that an increase in a money supply generates a higher inflation rate and discount rate. Consistently with Kwon et al. (1997), Mookerjee and Yu (1997), Bilson et al. (1999), Hanousek and Filler (2000), Flannery and Protopapadakis (2002), Ibrahim and Aziz (2003), Frimpong (2009), and Asmy et al. (2009), we find a significant positive money supply effect on stock returns in the short run for the Hungarian market. However, either due to a short time series or other reasons, the support for the relationship with money supply was not strong for the other markets.

We find that index returns often depend on trade balance (proxy for the market's openness) in the short-run, and sometimes in the long-run. The log difference of export and import is significantly positively related to stock returns for the Czech, Slovakian, Slovenian, and Estonian markets. Trade components also entered the equations for the Lithuanian and Polish markets.

We report positive relations between the unemployment rate and index returns in the long and short run for the Czech, Hungarian, Romanian and Latvian markets. These findings are in line with Boyd et al. (2005), who also state the positive effect on stock returns in an expanding economy. It follows that the stock market rises on bad news from the labour market, namely on the increasing unemployment rate.

The error correction parameter acquires a relatively high value for the Bulgarian stock market, and this is suggestive of a fast adjustment of index returns after macroeconomic shocks to equilibrium in the market. However, this is probably the feature of the Bulgarian market, since other countries have no such high speed of adjustment.

Generally, we may conclude the weak form of efficiency for the above-mentioned markets – with an exception to Slovakia and Lithuania, which are also dependent on the lags of log-returns. Our results are in line with the conditions for market efficiency: stock returns are cointegrated with macro variables and are not affected by their lagged values. The probable common patterns should include economic sentiment, industrial turnover index and/or its gap, trade balance and interest rates. This provided the basis for the panel data analysis made as the final step.

We use the fixed effects panel data model, hypothesising that the Central and Eastern European markets are interrelated and exhibit commonality in their movements. Panel models address a broader range of issues and tackle more complex problems than would be possible with pure time series or pure cross-sectional data alone. We employ the raw values of interest rate and real effective exchange rate; as they possess good statistical properties to be used in multiple regression analysis (see Table 2).

Table 2. Fixed Effects Model Estimation for the Eastern European Stock Returns

Variable	Coefficient	Std. Error	*t*-Statistic	Prob.
$\Delta(R_t)$	-0.01**	0.01	-1.99	0.047
C	-0.29**	0.12	-2.4	0.017
$\Delta\log(MC_t)$	0.25***	0.03	8.48	0.000
$\Delta\log(X_t/M_t)$	0.07***	0.02	3.36	0.001
$\log(IPI_{t-1})$	0.07**	0.03	2.50	0.013
$\log(ITI_t/ITI_t^*)$	6.96***	1.57	4.43	0.000
$\log(ESI_t)\cdot\log(ITI_t/ITI_t^*)$	-1.52***	0.39	-4.48	0.000
$\log(Y_{t-1})$	-0.02**	0.01	-2.37	0.018
$adj.\ R^2 = 0.23$, $BIC = -2.67$, $DW = 1.96$, $p(Hausman) = 0.007$, $MAPE(6)=13.76\%$				

Note: Asterisks refer to the level of significance: ** significance at 0.05, *** significance at 0.01.
Source: Authors' calculations

Results, in fact, generalise what we have seen for individual ECM models. In general, Eastern European markets have an asymmetric response to good and bad news, with higher volatility during recessions. Another aspect is related to the trade balance; it shows that the competitiveness of the economy is important for these emerging markets. All other common factors are related to stock fundamentals.

Consistently with Frimpong (2009) and Samitas and Kenourgios (2007), we find a significant negative interest rate impact on index returns across countries. Decline in interest rates is good news for returns, as they subsequently tend to grow. When an interest rate for money market instruments (proxy for risk free rate) is low, the investors want the higher discount rates and invest in riskier assets with higher returns. Running regressions for each country separately, we find that interest rate has an explanatory power for index returns in Hungary and Romania. However, for Hungary it has a positive effect, whereas for Romania it has a negative impact. Furthermore, the lagged interest rate effect on index returns suggests the Romanian market to be inefficient.

We report that good news on changes in market capitalisation has a positive impact on index returns across emerging countries. Running regressions for each country, we find market capitalisation to be positively related to index returns in Bulgaria; whereas, it is not included for the rest of the countries, but unbalanced panel estimates enables to overcome the difficulties with this particular variable – this is interesting to include since it represents the general characterstic of the market's level of development.

The Eastern European emerging stock markets are common in their responses to good news about the changes in exports relative to imports of goods. We also find that exports relative to imports of goods have a positive effect on index returns in the Czech, Estonian and Slovenian markets. Exports of goods positively related to stock returns in Poland and Slovakia. Imports of goods have a negative impact on stock returns in Slovakia, and the lagged imports of goods generate an increase in index returns in Lithuania. The significant effect of lagged values suggests market inefficiency for the Lithuanian market.

Industrial production index effect on stock returns is explored in Kwon et al. (1997), Bilson et al. (1999), Hanousek and Filler (2000), and Samitas and Kenourgios (2007). We use an industrial production index as a proxy for GDP, which is not available on a monthly basis. Consistently with theory and empirical evidence, it proves to be positively correlated to stock returns at 5% significance level across countries. It is also proven to have an explanatory power for the Czech Republic, Lithuania and Poland index returns. However, for Lithuania and the Czech Republic the direction of the effect is consistent with the theory; whereas for Poland we have a negative industrial production effect on stock returns.

Economic sentiment index has a significant indirect impact on the short run dynamics through the industrial turnover gap (ITI_t / ITI_t^*) across countries. Therefore, the returns are influenced by asymmetric reaction that lowers gap impact on returns' variation in "good" times and significantly increases it in "bad" times. The same effect is found for the Bulgarian, Czech, Polish and Romanian markets.

The error correction parameter (Y_{t-1}) acquires a relatively low value across most of the emerging markets and in the cross-section, except for Bulgaria and Hungary, and this is suggestive of a slow adjustment of index returns after macroeconomic shocks to the equilibrium in the market. This also explains why there may be no immediate impacts after the macro economic data is announced, the cumulative long run impact takes a much longer period of time to accumulate than the usual impact window used in previous research. For example, half life shock correction for the Czech Republic would be 1/0.05 = 20 months, but only 2-3 months for Bulgaria and Hungary. The latter may also be linked to the stock market's memory characteristic.

5. Conclusions

We find that macroeconomic news (interest rate, industrial production index, exports relative to imports of goods, industrial turnover gap and market sentiment) have a significant impact on stock returns across Eastern European stock markets. However, macroeconomic news does not influence the rapid change in the market in the short run, but has a rather prolonged influence on trends in log returns in the long-run, especially bearing in mind a slow rate of error correction. This partly explains why previous research on the subject was not very successful, since total macro news' impact can be seen within months and even years of trade with the immediate impact being relatively small, though not insignificant.

Nonetheless, the nature of macro information affects the direction of subsequent changes in index returns, both in the short and in the long-run. Allowing for variation in the economic and market conditions, we find that market sentiment affects the asymmetry in the index returns' responses to the industry turnover gap. The returns are influenced by asymmetric reaction that lowers gap impact on returns' variation in "good" times and significantly increases it in "bad" times. The asymmetry effects are found to be significant across markets and at the country level for the Bulgarian, Czech, Polish and Romanian index returns.

Although a panel error correction model and cointegration analysis for cross-sectional data are useful, the more sophisticated econometric tools can be applied to extend the findings of this research. Alternative useful methods for data mining left for further research are weighted average least squares (WALS) and Bayesian model averaging (BMA). These methods are based on the idea that it is impossible to find out a true data generating process by pretesting estimates (e.g. stepwise selection). The latter has one serious drawback, that they may introduce a quite large bias in parameter estimates, by excluding statistically insignificant coefficients. Averaging approaches reduce this bias, since they include both unrestricted and all restricted estimates. It, therefore, insures that empirically based conclusions are more consistent and reliable.

References

Adams, G., McQueen, G. and Wood, R. 2004. The Impact of Inflation News on High Frequency Stock Returns. *Journal of Business*, Vol. 77, No. 3, pp. 547-574.

Andersen, T., Bollerslev, T., Diebold, F. and Vega, C. 2007. Real-Time Price Discovery in Stock, Bond, and Foreign Exchange Markets. *Journal of International Economics*, Vol. 73, No. 2, pp. 251-277.

Asmy, M., Rohilina, W., Hassama, A. and Fouad, M. 2009. Effects of Macroeconomic Variables on Stock Prices in Malaysia: An Approach of Error Correction Model. Retrieved from http://mpra.ub.uni-muenchen.de/20970/, MPRA Paper No. 20970.

Babecky, J., Komárek, L. and Komárková, Z. 2008. Financial Integration of Stock Markets among New EU Member States and the Euro Area. *The Warwick Economics Research Paper Series (TWERPS),* No. 849, University of Warwick, Department of Economics.

Bilson, C., Brailsford, T. and Hooper, V. 1999. Selecting Macroeconomic Variables as Explanatory Factors of Emerging stock markets Returns. *Working Paper Series*. Retrieved from http://ssrn.com/abstract=201908

Boyd, J., Hu, J. and Jaganathan, R. 2005. The Stock Market's Reaction to Unemployment News: Why Bad News is Usually Good for Stocks. *The Journal of Finance*, Vol. 60, No. 2, pp. 649-672.

Conrad J., Cornell, B. and Landsman, W. 2002. When Bad News Is Really Bad News? *Journal of Finance*, Vol. 57, No. 6, pp. 2507-2532.

Fama, E.F. and French, K. 1989. Business Conditions and the Expected Returns on Bonds and Stocks. *Journal of Financial Economics*, Vol. 25, No. 1, pp. 23-50.

Flannery, M. and Protopapadakis, A. 2002. Macroeconomic Factors Do Influence Aggregate Stock Returns. *The Review of Financial Studies*, Vol.15, No. 3, pp. 751-782.

Frimpong, J.M. 2009. Economic Forces and Stock Market in a Developing Economy: Cointegration Evidence from Ghana. *European Journal of Economics, Finance, and Administrative Studies*, No. 16. Retrieved from http://www.eurojournals.com.

Gay, R.D. 2008. Effect of Macroeconomic Variables on Stock Market Returns for Four Emerging Economies: Brazil, Russia, India, and China. *International Business and Economics Research Journal*, Vol. 7, No. 3, pp. 1-8.

Hanousek, J. and Filler, R.K. 2000. The Relationships between Economic Factors and Equity Markets in Central Europe. *Economics of Transition*, Vol. 8, No. 3, pp. 623-638.

Ibrahim, M.H. and Aziz, H. 2003. Macroeconomic Variables and the Malaysian Equity Market: A View through Rolling Subsamples. *Journal of Economic Studies*, Vol. 30, No.1, pp. 6-27.

Kandir, S.Y. 2008. Macroeconomic Variables, Firm Characteristics and Stock Returns: Evidence from Turkey. *International Research Journal of Finance and Economics*, No.16, pp. 35-45.

Kwiatkowski, D., Phillips P.C.B., Schmidt, P. and Shin Y. 1992. Testing the Null Hypothesis of Stationarity against the Alternative of a Unit Root. *Journal of Econometrics*, Vol. 54, No. 1-3, pp. 159-178.

Kwon, C.S., Shin, T.S. and Bacon, F.W. 1997. The Effect of Macroeconomic Variables on Stock Market Returns in Developing Markets. *Multinational Business Review*, Vol. 5, No. 2, Fall, pp. 63-70.

Laakkonen, H. and Lanne, M. 2008. Asymmetric News Effect on Return Volatility: Good vs Bad News in Good vs Bad Times. Retrieved from http://mpra.ub.uni-muenchen.de/8296/, MPRA Paper No. 8296.

McQueen, G. and Roley, V.V. 1993. Stock Prices, News, and Business Conditions. *The Review of Financial Studies*, Vol. 6, No. 3, pp. 683-707.

Mukherjee, T.K. and Naka, A. 1995. Dynamic Relations between Macroeconomic Variables and the Japanese Stock Market: An Application of a Vector Error Correction Model. *International Business and Economics Research Journal*, Vol. 18, No. 2, pp. 223-237.

Mookerjee, R. and Yu, Q. 1997. Macroeconomic Variables and Stock Prices in a Small Open Economy: The Case of Singapore. *Pacific-Basin Finance Journal*, Vol. 5, No. 3, pp. 377-388.

Muradoglu, G., Taskin, F. and Bigan, I. 2000. Causality between Stock Returns and Macroeconomic Variables in Emerging Markets. *Russian and East European Finance and Trade*, Vol. 36, No. 6, pp. 33-53.

Pedersen, T.M. 2001. The Hodrick-Prescott Filter, the Slutzky Effect, and the Distortionary Effect of Filters. *Journal of Economic Dynamics and Control*, Vol. 25, No. 8, pp. 1081-1101.

Samitas, A.G. and Kenourgios, D.F. 2007. Macroeconomic Factor's Influence on "New" European Countries Stock Returns: The Case of Four Transition Economies. *International Journal of Financial Services Management*, Vol. 2, No. 1/2, pp. 34-49.

Veronesi, P. 1999. Stock Market Overreaction to Bad News in Good Times: A Rational Expectations Equilibrium Model. *The Review of Financial Studies*, Vol. 12, No. 5, pp. 975-1007.

Zivot, E. and Andrews, D.W.K. 1992. Further Evidence on the Great Crash, the Oil-Price Shock, and the Unit-Root Hypothesis, *Journal of Business & Economic Statistics*, Vol. 10, No. 3, pp. 251-270.

Appendix. Integration Order and KPSS and Zivot-Andrew Tests for the Levels of the Data

		Bulgaria	Czech Rep.	Estonia	Hungary	Lithuania	Latvia	Poland	Romania	Slovenia	Slovakia
CCI	ZA	-5.77	-3.29	-3.49	-4.24	-4.18	-3.30	-4.73	-4.31	-4.47	-4.29
	KPSS	0.23	0.23	0.40	0.28	0.44	0.38	0.29	0.25	0.18	0.29
ESI	ZA	-4.14	-3.50	-3.34	-3.43	-3.95	-3.70	-4.16	-3.77	-4.51	-4.56
	KPSS	0.27	0.31	0.44	0.41	0.25	0.35	0.33	0.22	0.42	0.28
HCPI	ZA	-3.48	-3.26	-3.50	-5.35	-3.07	-3.12	-3.98	-2.96	-3.25	-4.76
	KPSS	0.50	0.39	0.47	0.31	0.58	0.55	0.29	0.57	0.38	0.48
ICI	ZA	-4.64	-4.78	-4.65	-4.29	-4.42	-3.23	-4.09	-4.45	-4.02	-4.29
	KPSS	0.27	0.26	0.40	0.18	0.37	0.48	0.24	0.35	0.25	0.24
IPI	ZA	-4.84	-4.04	-4.19	-4.89	-4.71	-3.70	-3.29	-5.03	-5.05	-3.70
	KPSS	0.29	0.23	0.37	0.23	0.37	0.44	0.19	0.22	0.18	0.13
ITI	ZA	-3.36	-4.08	-3.60	-3.99	-3.51	-4.45	-3.65	-5.31	-5.58	-4.50
	KPSS	0.20	0.27	0.25	0.23	0.20	0.22	0.18	0.08	0.23	0.22
M2	ZA	-3.24	-2.65	-3.45	-6.09	-4.19	-2.86	-5.16	-4.65		-4.19
	KPSS	0.20	0.49	0.40	0.17	0.35	0.34	0.15	0.30		0.34
M	ZA	-3.26	-3.28	-2.95	-3.85	-3.65	-3.43	-3.43	-3.34	-3.67	-3.65
	KPSS	0.17	0.21	0.31	0.22	0.22	0.23	0.19	0.18	0.22	0.17
MC	ZA	-4.61	-4.10	-4.60	-3.57	-3.91	-4.55	-3.93	-3.25	-5.02	-4.02
	KPSS	0.32	0.22	0.14	0.29	0.41	0.37	0.22	0.25	0.15	0.35
NEER	ZA	-3.28	-4.75	-2.93	-3.98	-2.92	-3.25	-3.96	-4.42	-4.28	-3.86
	KPSS	0.43	0.12	0.26	0.25	0.44	0.43	0.20	0.48	0.41	0.48
PPI	ZA	-5.48	-3.40	-2.89	-4.28	-3.12	-3.03	-5.22	-3.88	-4.31	-3.45
	KPSS	0.25	0.22	0.50	0.24	0.32	0.45	0.11	0.26	0.14	0.40
R	ZA	-2.87	-2.99	-1.89	-3.44	-3.08	-3.96	-5.34	-3.86	-5.57	-3.53
	KPSS	0.34	0.45	0.51	0.14	0.50	0.31	0.50	0.52	0.18	0.34
REER	ZA	-4.67	-4.52	-3.08	-3.95	-3.35	-3.54	-3.99	-2.62	-3.71	-3.93
	KPSS	0.38	0.20	0.36	0.21	0.35	0.56	0.20	0.25	0.21	0.30
ST	ZA	-4.71	-4.02	-4.02	-3.87	-3.64	-3.64	-4.40	-3.74	-7.06	-4.44
	KPSS	0.29	0.29	0.33	0.37	0.32	0.27	0.23	0.31	0.16	0.33
UR	ZA	-3.78	-2.21	-4.39	-3.29	-4.61	-5.17	-2.86	-5.95	-4.15	-2.16
	KPSS	0.20	0.20	0.33	0.34	0.36	0.34	0.52	0.13	0.16	0.27
X	ZA	-4.44	-3.92	-3.31	-3.82	-4.26	-4.06	-3.37	-4.59	-3.48	-3.51
	KPSS	0.18	0.17	0.18	0.19	0.17	0.15	0.21	0.11	0.23	0.18
Y	ZA	-3.73	-2.65	-3.85	-2.85	-3.89	-3.48	-2.13	-3.30	-4.64	-5.70
	KPSS	0.35	0.48	0.40	0.27	0.44	0.41	0.42	0.29	0.21	0.19

Note: Critical values for ZA: 0.01= -5.57; **0.05= -5.08**; 0.1= -4.82, for KPSS: 0.1 = 0.119; **0.05 = 0.146**; 0.025=0.176; 0.01=0.216. Stationary series are shaded