# International Financial Reporting Standards and Noise Trading: Evidence from Central and Eastern European countries

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## Abstract

This paper examines whether the mandatory adoption of International Financial Reporting Standards (IFRS) in 2005 has produced an impact on the level of noise trading in Central and Eastern European (CEE) markets. Our results show that noise trading was mostly significant prior to the IFRS introduction, with its significance dissipating following the implementation. These findings are consistent with the notion that IFRS adoption has the potential to enhance the stability and informational efficiency of capital markets by promoting information-based trading and reducing the impact of noise traders. Overall, our results yield important insight into the impact of IFRS adoption on the overall market quality and investors' behaviour and bear important implications for the accounting professions and market regulators alike.

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## 1. Introduction

In recent years the adoption of International Financial Reporting Standards (IFRS) has gained considerable momentum around the world (Brüggemann *et al.*, 2010). The European Union (EU) Parliament, for instance, passed a new legislation in 2002 which requires all listed companies in the EU to adopt IFRS starting from the 1st January, 2005. Given this recent development, there has been a surge of academic interest in the area with an increasing number of studies investigating the economic consequences of adopting IFRS.[1] The general finding from this growing literature is that, after the harmonization of accounting standards, there appears to be a significant improvement in the quality of accounting information.[2]

Nonetheless, scarce evidence exists in the current literature on the direct effect (if any) of IFRS adoption on the behaviour of investors.[3] This is surprising given that the proponents of higher quality financial reporting have often maintained that the harmonization in accounting systems should improve the overall informational environment which in turn would attract greater participation of sophisticated investors from both domestic and foreign markets. Intuitively, to the extent that the introduction of IFRS improves financial statement quality and transparency, the informational efficiency of markets is expected to increase as a result of enhanced information-based trading. This increased transparency should also reduce the level of noise trading and enhance the speed at which new information is incorporated into prices.

However, to date, there exists no research aiming at directly addressing the above issues. Against this backdrop, we aim at examining for the first time the effect of accounting

[1] Soderstrom and Sun (2007) and Bruggemann et al. (2010) provide excellent reviews of the recent literature.
[2] See, amongst others, Barth et al. (2008), Wang et al. (2008), and Horton et al. (2009).
[3] In a study related to this paper, Florou and Pope (2009) examine whether the mandatory introduction of IFRS leads to an increase in institutional ownership of equities. In the microstructure setting, they show that the equity ownership by sophisticated users of financial statements (e.g., institutional investors) significantly increases in the years following IFRS adoption.

standards' harmonization upon the level of noise trading and volatility of three major transition markets (the Czech Republic, Hungary, and Poland) in the Central and Eastern European (CEE) area that began the implementation of IFRS as of January 2005 in line with the aforementioned EU legislation. More specifically, utilising the Sentana and Wadhwani (1992; thereafter, SW) feedback trading model that allows the existence of both rational informed investors and trend-chasing traders, we seek to address the following questions:

- Does the adoption of IFRS promote or inhibit noise trading in stock markets?
- Has the volatility of stock returns been reduced upon the introduction of IFRS?
- Are there any differences in the speed at which new information is incorporated into prices following the implementation of IFRS?

The results of our analyses are of direct interest to regulators and policy-makers in evaluating the potential benefits and costs of mandatory IFRS adoption and to financial statement users who invest on the basis of company fundamentals, treating them as principal indicators for future market movements.[4] Furthermore, this study adds to the growing literature studying the effects of mandatory IFRS adoption in a number of ways. First, we complement the recent research of Florou and Pope (2009) and DeFond et al. (2010) who test for the link between the accounting standards harmonization and institutional ownership of equities, and Bruggemann et al. (2009) who examine the association between individual investor ownership and IFRS adoption in 31 countries. Their results suggest that mandating a uniform set of accounting standards improves financial statement quality which in turn attracts greater investments not only by professional institutional investors, but also by individual investors. Our paper addresses this connection using a theoretical framework that allows for the existence of both groups of investors in an attempt to provide new insights to this issue.

[4] The U.S. Securities and Exchange Commission (SEC) has recently proposed a roadmap which, if approved, could lead to the required use of IFRS by U.S. issuers from 2014 onwards.

Secondly, our study also contributes to the understanding of *why* feedback trading might take place. Although a number of reasons have been put forward in explaining the presence of trend-chasing behaviour, such strategies are usually associated with noise or uninformed traders whose demands are expected to vary with the informational environment.[5]

The main findings of our investigation can be summarized as follows. First, our results suggest that, the IFRS adoption reduces the intensity of the positive feedback trading, suggesting that informed traders are playing a greater role in the markets during the post-IFRS period. Moreover, we find that the level (and persistence) of stock return volatility has greatly decreased after the implementation of IFRS. Specifically, we observe an improvement in the speed of adjustment to equilibrium after the arrival of new information in the markets. This implies that IFRS adoption contributes positively to the market stability and efficiency. Additional analysis indicates that the above results survive an array of robustness checks. Taken together, our findings are consistent with the view that IFRS adoption has the potential to enhance the stability and informational efficiency of capital markets by promoting information-based trading and reducing the impact of noise traders who exhibit positive feedback, or trend-chasing, behaviour.

The remainder of the paper is organized as follows. Section 2 briefly outlines the theoretical framework and model specifications employed. Section 3 describes the data, presents our testable hypotheses and delineates the research design used in our investigation. The results and the sensitivity analyses are then presented and discussed in Section 4. Finally, Section 5 provides concluding remarks and discusses the implications of the findings.

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[5] However, it should be noted that feedback trading can also be the result of many ‘rational’ motivations such as trading on extrapolative expectations, activation of stop-loss orders, and portfolio insurance strategies.

## 2. Theoretical Framework and Model Specifications

Whether noise traders affect stock prices is a question of long-standing interest to economists. Shiller (1984), for example, argues that social norms or fashions can influence asset price movements. Black (1986) introduces the concept of noise traders and offers a formal definition of 'noise trading' as trading on noise (or non-information) as if it were information. He argues that noise traders may not be eliminated from the market because rational arbitrage against them is very costly or limited and, therefore, uninformed noise traders could significantly influence stock price dynamics.[6] Many researchers have since advocated the use of noise-trader models as an alternative in explaining asset price dynamics (see, e.g., Shiller, 1984; De Long et al., 1990), showing that noise traders cause prices to deviate from the fundamental values. In this paper, we employ the theoretical framework proposed by Sentana and Wadhwani (1992; SW) to examine the extent to which IFRS adoption inhibits or promotes the level of noise trading in the market place.

SW assume two distinct groups of investors: one consisting of 'smart-money' investors who invest rationally on the basis of rational forecasts of future returns subject to their wealth limitations; the other is a group of 'noise traders' who do not base their investment decisions on fundamental values but rather react to previous price changes instead (feedback traders). More specifically, the demand for shares by the first group (rational investors) in period $t$, $D_{1,t}$, is given by:

$$D_{1,t} = [E_{t-1}(R_t) - \kappa] / \theta(\sigma_t^2) \qquad (1)$$

where $E_{t-1}(R_t)$ is the expected return in the period t-1, κ is the risk-free rate of return, $\sigma^2_t$ is the conditional variance in period $t$ and $\theta$ is the fixed coefficient measuring the degree of risk

[6] Economists have long debated the effects of noise traders on the equilibrium market prices. Some argue that their existence is destabilizing, causing inefficiency and instability in asset prices. It should, however, be recognized that noise traders can be beneficial as they provide the market with liquidity (De Long et al. 1990).

aversion. Assuming $\theta$ is positive, the product $\theta(\sigma^2_t)$ is the required risk premium at time t.[7] This demand function is consistent with the maximization of expected mean-variance utility and implies that rational investors' demand for risky assets is a positive function of the expected excess return, $E_{t-1}(R_t)$ - κ, but is inversely related to the degree of risk aversion, $\theta$.

The second group of investors, 'noise traders', is assumed to follow a feedback trading strategy whose demand for shares, $D_{2,t}$ , depends *solely* on the previous period's return:

$$D_{2,t} = \lambda R_{t-1} \tag{2}$$

where $R_{t-1}$ denotes the actual return in the previous period. The value of the parameter λ allows us to discriminate between two types of feedback traders: λ > 0 refers to the case of positive feedback traders who buy (sell) after a price rise (fall), while λ < 0 indicates negative feedback traders, who adhere to a 'buy low, sell high' investment strategy.[8] Feedback traders of either type have the detrimental effect of moving prices away from their fundamentals. Hence, investigation of whether the IFRS adoption has had an impact over feedback trading should be of interest to both developed and emerging markets considering introducing IFRS. In equilibrium all shares in the market must be held, thus:

$$D_{1,t} + D_{2,t} = 1 \tag{3}$$

Therefore, substituting (1) and (2) into (3) and rearranging gives:

$$E_{t-1}(R_t) - \kappa = \theta(\sigma_t^2) - \lambda[\theta(\sigma_t^2)]R_{t-1} \tag{4}$$

Then, assuming the rational expectation [i.e., $R_t = E_{t-1}(R_t) + \varepsilon_t$ ], equation (4) becomes:

$$R_t = \kappa + \theta(\sigma_t^2) - \lambda[\theta(\sigma_t^2)]R_{t-1} + \varepsilon_t \tag{5}$$

[7] Note that if all investors are rational 'smart-money' investors (i.e., $D_{1,t}$ = 1), market equilibrium yields the standard capital asset pricing model (CAPM): $E_{t-1}(R_t)$ - κ = $\theta(\sigma^2_t)$.

[8] Evidence of this type of behaviour can be found in both individual and institutional investors (see, e.g., Nofsinger and Sias, 1999). As noted earlier feedback trading need *not* be irrational or, noise trading in the sense of Black (1986) and De Long et al. (1990). It is consistent with, for example, portfolio insurance strategies and stop-loss orders. Nonetheless, as Shleifer (2000) points out, the interaction of feedback traders and rational investors could lead to price movements that are not warranted by their fundamental value.

where $\varepsilon_t$ is an independently and identically distributed error term. To test the hypothesis that IFRS adoption has significantly affected the level of feedback trading we convert and re-parameterize equation (5) into an empirical regression model:

$$R_t = \alpha + \theta(\sigma_t^2) + (\varphi_0 + \varphi_1 \sigma_t^2) R_{t-1} + \varepsilon_t \qquad (6)$$

where $\alpha = \kappa$; $\varphi_1 = -\lambda\theta$. Thus, the presence of positive (negative) feedback trading implies that $\varphi_1$ is negative (positive) and statistically significant. The coefficient $\varphi_0$ is added to capture the autocorrelation induced by potential market frictions (e.g., non-synchronous trading).[9]

It is clear from equation (6) that, to complete the model, it is necessary to specify a process for the evolution of the conditional variance $\sigma^2_t$. As it is now well established in the literature that stock returns are characterised by conditional heteroscedasticity, a GARCH specification is employed.[10] In particular, we assume an asymmetric GJR-GARCH (1,1) process given by:

$$\sigma_t^2 = \alpha_0 + \alpha_1 \varepsilon_{t-1}^2 + \beta \sigma_{t-1}^2 + \delta I_{t-1} \varepsilon_{t-1}^2 \qquad (7)$$

where $\sigma^2_t$ is the conditional volatility in period $t$, $\varepsilon_{t-1}$ is the innovation in period $t-1$ and $I_{t-1}$ is an indicator which assumes a value of one in response to bad news ($\varepsilon_{t-1} < 0$) and 0 otherwise. If $\delta$ is positive and statistically significant, it indicates that a negative shock has a greater impact on future volatility than a positive shock of the same size. $\alpha_1$ is the news coefficient capturing the impact of the most recent innovation, $\beta$ is a measure of persistence, and $\alpha_0$ represents unconditional volatility.[11] The model given by equations (6) and (7) is tested on daily return series from three major CEE markets that have implemented IFRS in 2005 so as to examine the potential link between accounting standards harmonization and noise trading.

[9] For recent applications of this theoretical model, see, e.g., Antoniou *et al.* (2005), Chau *et al*. (2008), Kallinterakis and Kaur (2010), Salm and Schuppli (2010), and Schuppli and Bohl (2010).
[10] Many studies have shown that stock returns are heteroscedastic. See Teräsvirta (2009) for an excellent survey.
[11] Given the initial values for $\varepsilon_t$ and $\sigma^2_t$, the parameters of mean and variance equations can be estimated simultaneously by Quasi-maximum likelihood method. WinRats 7.2 software was used and for numerical optimisation the Newton-Raphson and Broyden–Fletcher–Goldfarb–Shanno (BFGS) method were employed.

## 3. Data, Testable hypotheses and Research Design

The stock markets in Central and Eastern European (CEE) region have been expanding dramatically since the early 1990s and have begun their integration into the European area (Kaszuba, 2010). As shown in Figure 1, the stock market index prices in each sample country have increased substantially since the early transition period. Although several transition economies in CEE emerged from communism in the early 1990s, we focus on the Czech Republic, Hungary, and Poland, which have the three largest stock markets of the zone and the most active and developed markets in the region.[12] In addition, our sample choice is also motivated by the fact that these countries share similarities in their post-communism institutional evolution (see, e.g., Helmenstein, 1999).

**[FIGURE 1 ABOUT HERE]**

Therefore, our empirical analysis includes the main equity indices of these three CEE markets, namely the Czech Republic (PX GLOBAL), Hungary (BUX), and Poland (WIG). With each of these indices launched at a different point in the 1990s, the start-date for each market is bound to be different: the Czech Republic (6/4/1994), Hungary (2/1/1991), and Poland (16/4/1991). The end-date of our analysis for all six markets is the 31st of December 2007; this choice was made in order to mitigate any effects from the global financial crisis that has been ongoing since 2008. Daily closing prices were collected from DataStream, and returns were then calculated as the first-difference of the natural logarithms of prices, $R_t = \ln(P_t / P_{t-1}) \times 100$.

[12] According to the statistics of the World Federation of Exchanges, at the end of 2010, the market capitalization was US$ 80 billion in the Czech Republic, 28 billion in Hungary, and 190 billion in Poland.

Given their growing significance in global financial markets and their unique characteristics, CEE markets provide a unique opportunity to investigate the effect of IFRS on noise trading. Compared to other mature and emerging markets, CEE stock markets remain small and relatively illiquid, deterring large international investors from trading in these markets (Korczak and Bohl, 2005). However, after the harmonization of accounting standards in 2005, the improved financial statement transparency and comparability is expected to encourage a greater participation of foreign investors in these markets. Indeed, Florou and Pope (2009) and DeFond et al. (2010) show that the mandatory adoption of IFRS in the EU has led to a surge of cross-border investments, especially from the institutional investors. With institutional investors having been coined as the natural candidates for the "rational" informed investors (Barber and Odean, 2008), an increase in their participation in the market is expected to reduce the level of noise trading and enhance the speed at which new information is incorporated into prices. Additionally, the reduction in the impact of feedback noise traders should further enhance market stability by reducing the stock price volatility (Antoniou *et al.*, 2005).

Based on the above discussion, we propose the following hypotheses:

$H_1$. By providing better information to investors, the introduction of IFRS promotes information-based trading and reduces the impact of noise traders in the market.

$H_2$. Following the adoption of IFRS and an associated improvement in information quality, the speed at which new information is incorporated into prices should also increase.

$H_3$. The reduction in the impact of noise trading and improvement in information efficiency after IFRS adoption tends to lessen the overall stock market volatility.

In this paper, we explore the above hypotheses by defining two distinctive sub-periods for each market: (i) the 'Pre-IFRS' period extending from each market's start-date to 31/12/2004 and (ii) the 'Post-IFRS' period from 03/01/2005 to 31/12/2007. For each sub-period, we run the noise trading model (as specified in equations (6) and (7)) following the estimation procedure described in Section 2; comparisons are then made of the estimated coefficients. Although attention will be paid to the coefficients describing the conditional variance process, $\alpha_0$, $\alpha_1$, $\beta$, and $\delta$, our analysis will focus on the change (if any) in the value of the parameter governing the level of feedback trading, $\varphi_1$, during the pre- and post-IFRS periods. To illustrate the change in persistence and the asymmetric nature of the volatility process, we calculate the Half-life in line with Harris and Pisedtasalasai (2006) as $\ln(0.5)/\ln(\alpha_1 + \beta + \delta/2)$ and the asymmetric ratio as $(\alpha_1 + \delta)/\alpha_1$.

Furthermore, in order to formally test the shift in the feedback trading parameter and volatility dynamics, we follow Antoniou et al. (2005) and estimate the following 'augmented' feedback trading model for the whole sample period.

$$R_t = \alpha + \theta(\sigma_t^2) + [(\varphi_{0,pre} D_t + \varphi_{0,post}(1 - D_t)] R_{t-1} + +[(\varphi_{1,pre} D_t + \varphi_{1,post}(1 - D_t)] \sigma_t^2 R_{t-1} + \varepsilon_t \quad (8)$$

$$\sigma_t^2 = \alpha_{0,pre} D_t + \alpha_{0,post}(1 - D_t) + \alpha_1 \varepsilon_{t-1}^2 + \beta \sigma_{t-1}^2 + \delta I_{t-1} \varepsilon_{t-1}^2 \quad (9)$$

where $D_t$ is a dummy variable which assumes the value of one post-IFRS and zero pre-IFRS. We then test for the hypotheses $H_{01}$: $\varphi_{0,pre} = \varphi_{0,post}$ , $H_{02}$: $\varphi_{1,pre} = \varphi_{1,post}$ , and $H_{03}$: $\alpha_{0,pre} = \alpha_{0,post}$ on the basis of the likelihood ratio (LR) test statistic. In addition, we consider various modifications to the original SW model given by equations (6) and (7) in order to check the sensitivity of our results. First, it can be argued that feedback trading may be more intense during market declines (Koutmos, 1997). We therefore estimate a variant of equation (6) to give consideration to this possibility of asymmetric feedback trading:

$$R_t = \alpha + \theta(\sigma_t^2) + (\varphi_0 + \varphi_1 \sigma_t^2) R_{t-1} + \varphi_2 |R_{t-1}| + \varepsilon_t \quad (10)$$

Second, in line with Salm and Schuppli (2010) and Antoniou *et al.* (2011), we modify the original feedback trading model by including a second-order lag to allow for the possibility that some feedback traders might maintain a longer memory in their investment decisions and respond to the last few previous periods' price movements:

$$R_t = \alpha + \theta(\sigma_t^2) + (\varphi_0 + \varphi_1\sigma_t^2)R_{t-1} + +(\varphi_0^* + \varphi_1^*\sigma_t^2)R_{t-2} + \varepsilon_t \qquad (11)$$

## 4. Empirical Results

We estimate the empirical models outlined in Section 3 using the Quasi-maximum likelihood method of Bollerslev and Wooldridge (1992) which is robust to non-normality. Table 2 summarises the results of the maximum likelihood estimates for the original version of SW feedback trading model as specified in equations (6) and (7) which was designed to examine whether there exists a significant change in the level of noise trading following the IFRS introduction. Estimation results for the 'augmented' model (as given by equations (8) and (9)) during the whole sample period are reported in Table 3. Finally, Table 4 presents our robustness checks' results from implementing two different specifications to equation (6).

### 4.1 Descriptive Statistics

As a useful starting point, we present the descriptive statistics of the six markets' index-returns in Table 1, Panel A. The table shows that all return series display a skewed and leptokurtic pattern, contributing to the clear departures from normality (as indicated by the significant Jarque-Bera statistics). The Ljung-Box (LB) statistics suggest that there are temporal dependencies in the first and second moments of the return distribution. Results for the ARCH (1) and JOINT tests confirm the presence of significant heteroscedasticity and asymmetries in the conditional volatility of all returns.[13] Furthermore, to gauge the initial idea on the intensity of feedback trading in these markets, we estimate an autoregressive model.[14] The results reported in Panel B of Table 1 show that there are significant autocorrelations and the coefficients are mostly positive. Nonetheless, as shown in Section 2, the interaction of rational investors and feedback traders will give rise to return patterns that are more complex than a simple autoregressive model can capture. It is, therefore, interesting and informative to further investigate the significance of feedback trading in these markets, and whether the intensity of such trading behaviour changes as a result of the IFRS adoption.

**[TABLE 1 ABOUT HERE]**

### 4.2 Evidence on Feedback Trading and Volatility

*(i) Volatility*

We begin with the presentation of our results for the original SW model reported in Table 2, starting first from the estimates of the conditional variance equation. At first glance, we notice that $\delta$ remains overall insignificant (with the exception of the post-IFRS period in Czech Republic), thus denoting the near-absence of volatility asymmetries. The latter is further confirmed when calculating the asymmetric ratio ($\alpha_1 + \delta / \alpha_1$), which assumes relatively moderate values in absolute terms, except for the Czech Republic post-IFRS. This is somewhat surprising given that the presence of significant volatility asymmetry (i.e., responses to bad news generally lead to greater volatility than do responses to good news) is often interpreted as indirect evidence for the presence of noise traders (see Chau et al., 2008).

[13] The JOINT test is refers to Engle and Ng's (1993) test for the potential asymmetries in conditional volatility. The test statistic is a F-statistic for the null hypothesis of b1=b2=b3=0 of the following regression:

$$Z_t^2 = a + b_1 S_t^- + b_2 S_t^- \varepsilon_{t-1} + b_3 S_t^+ \varepsilon_{t-1} + v_t$$

where $Z_t^2$ is the square standardized residuals, $(\varepsilon_{t-1}/\sigma_t)^2$, $St^-$ is a dummy variable that takes a value of unity if $\varepsilon_{t-1} < 0$ and zero otherwise; and $St^+$ is a dummy variable that takes a value of unity if $\varepsilon_{t-1} > 0$ and zero otherwise.

[14] Common perception is that the positive (negative) feedback trading leads to positive (negative) autocorrelation of returns. To investigate this possibility we estimate a simple autoregressive model, AR(5).

Additionally, volatility appears highly persistent as indicated by the overwhelmingly significant (at the 5 percent level) $\beta$ coefficient, reflecting significant temporal dependencies in the conditional volatility process whose magnitude exhibits signs of decline after the IFRS adoption as illustrated by the substantial reduction of half-life values in the post-IFRS period. In all three cases the $\alpha_1$ coefficient appears overtly significant pre-IFRS, yet its post-IFRS estimates indicate a notable drop in terms of both its absolute size and significance (it appears significant only in Hungary), thus suggesting a reduced impact of news over volatility following the introduction of IFRS which could be indicative of a slower aggregation of information flow in the market.

Taken as a whole, the evidence presented thus far in relation to the changes in $\alpha_1$ and $\delta$ coefficients from the pre- to post-IFRS period provides some initially surprising results, indicating that post-IFRS news has had a reduced impact over volatility and that asymmetric responses of volatility to news both before and after the harmonization of accounting standards in our sample countries appear non-existent. However, consideration of changes in $\beta$ and half-life values suggests that the old innovations have become less persistent post-IFRS. Nonetheless, to assess the extent of which the introduction of IFRS inhibits / promotes noise and feedback trading in stock markets, further investigation is required.

**[TABLE 2 ABOUT HERE]**

*(ii) Feedback Trading*

Consider next the conditional mean equation estimates, especially on the change (if any) in the value of the parameter governing the level of feedback trading, $\varphi_1$, from pre- and post-IFRS periods. In each of the markets there is clear evidence that the overall impact of IFRS-adoption on the market efficiency has been positive: in Hungary and Poland the pre-IFRS significance of the $\varphi_0$ coefficient dissipates post-IFRS; in the Czech Republic $\varphi_0$ remains significant both before and after the IFRS-adoption, yet its size exhibits a notable decline in absolute terms post-IFRS, indicating a lower stock return autocorrelation in these markets. This lends strong support to the proponents of high quality financial reporting who often maintain that the harmonization in accounting systems would improve the financial statement quality and transparency, which in turn should lead to an improvement in market efficiency.

Turning our attention to the focus of this paper, consider the impact of the IFRS adoption on feedback noise trading in stock markets. As shown in Panel A of Table 2, the significance of the feedback parameter, $\varphi_1$, seems to be confined solely to the pre-IFRS period. Specifically, the Czech Republic and Hungary present us with significant (5 percent level) positive feedback trading prior to the adoption of IFRS; the aftermath of the IFRS-adoption sees this significance evaporating, with $\varphi_1$ appearing uniformly insignificant across all three markets.[15]

**[FIGURE 2 ABOUT HERE]**

**[TABLE 3 ABOUT HERE]**

To determine whether the above estimates are significantly different pre- versus post-IFRS we run the set of equations (8) and (9) for the whole sample period. Results in Table 3 confirm the previously documented rise in efficiency as well as the dissipation of feedback trading significance post-IFRS, although the significance in the difference between the $\varphi_0$ and $\varphi_1$ estimates pre- versus post-IFRS is statistically confirmed only in the Czech Republic. More specifically, in all three cases the estimated coefficients for autocorrelation, $\varphi_{0,pre}$ , and feedback trading, $\varphi_{1,pre,}$ for pre-IFRS are highly significant, with its significance dissipating following the implementation of IFRS (as indicated by the insignificance of $\varphi_{0,post}$ and $\varphi_{1,post}$).

[15] To further examine the change in feedback noise trading behaviour over time, we compute the conditional return autocorrelation implied by the feedback trading model (6)-(7) for a full sample period $\rho_t^{implied} = \hat{\varphi}_0 + \hat{\varphi}_1 \sigma_t^2$. As can be seen from Figure 2, conditional return autocorrelation exhibits huge variations over time and, in particular, downward spikes coinciding with pronounced drops in market prices. Similar figures for the conditional return autocorrelation can also be found in Salm and Schuppli (2010) and Chau et al. (2011).

## 4.3 Robustness and Additional Tests

In sum, the results of the above analysis suggest that the introduction of IFRS has enhanced the efficiency of these transition markets by improving the quality of their disclosure and rendering them more transparent, thus promoting fundamentals-based investments and reducing the scope for (and as our results indicate, the significance of) feedback trading. To check the robustness of our results, further estimations and additional tests were undertaken.

First, a series of studies (Koutmos, 1997; Koutmos and Saidi, 2001; Antoniou et al, 2005) have demonstrated that positive feedback trading appears more pronounced during market slumps as opposed to market upswings and have rationalized this through the use of strategies such as portfolio insurance and stop-loss orders whose activation during market downturns amplifies feedback tendencies in the market. To examine whether this is the case in our sample markets, we run the set of equations (10) and (7). Results are reported in Table 4, Panel A and for reasons of brevity include only the estimates of the feedback parameter ($\varphi_1$) and the feedback asymmetry parameter ($\varphi_2$). These results confirm the pre-IFRS presence of significant positive feedback trading in the Czech Republic and Hungary yet reflect the complete absence of any asymmetric feedback behaviour, as the $\varphi_2$ coefficient appears overwhelmingly insignificant in all tests. To control for the possibility of feedback trading entailing a longer memory in its structure, we run the set of equations (7) and (11), with results being reported in Table 4, Panel B. According to the estimated results for $\varphi_1''$, there exists no evidence of significant feedback trading at the second lag.[16] Overall, taking these additional set of tests together, it appears that the main findings and the general conclusions deriving from the study are robust.

[16] This is however inconsistent with the recent findings of Antoniou et al. (2011) who find that, in support of Shiller's (1990) hypothesis, feedback traders exhibit long memory / persistence in their feedback mechanism.

## 5. Conclusion

This paper has examined the effects of the adoption of International Financial Reporting Standards (IFRS) on the level of noise trading and volatility dynamics in three major CEE markets. Intuitively, the introduction of IFRS by mandating a uniform set of accounting standards and improving financial statement quality and transparency, should not only promote higher information-based trading in the market place, but also improve the overall efficiency and enhance the speed at which new information is reflected in prices.

Our results demonstrate the following. First, the intensity of feedback trading has decreased during the post-IFRS period. That is, the introduction of IFRS appears to have enhanced the efficiency of the stock markets by improving the quality of corporate disclosure and accounting transparency, in turn promoting information-based investments and has reduced the scope for (and the impact of) extrapolative behaviours by noise traders. Second, the level and persistence of volatility has decreased during the IFRS period, suggesting that the speed of adjustment has increased upon the arrival of new information. This implies that the accounting standards synchronization has enhanced the efficiency of the arbitrage process, such that disequilibria persist shorter than prior to the IFRS implementation. Finally, when compared to pre-IFRS periods, the autocorrelation of stock returns exhibits a notable decline in absolute terms post-IFRS, indicating an improvement in the efficiency of the markets.

Overall, the results suggest the IFRS adoption has the potential of enhancing the stability and informational efficiency of capital markets by promoting information-based trading and reducing the impact of noise traders. These findings are important in understanding the effect of IFRS adoption on the information environment and market dynamics and bear some important implications for the corporate managers, accounting professions and policy makers.

On the one hand, the documented dissipation in noise trading post-IFRS implies that the activity in our sample markets is dominated by fundamental investors for whom financial reports constitute part of their decision-making input. Since the presence of these investors enhances efficiency in securities' pricing, it is in the interest of the regulatory authorities that the informational environment is kept up to the standards of these investors' expectations. With the overall documented investors' reaction to IFRS being positive internationally (Armstrong et al., 2010), it is crucial that the enforcement of IFRS constitutes a key priority of local authorities as a contributing factor to the transparency of the market environment. This is a particularly important issue as regards transition economies, whose institutional structures are following an evolutionary trajectory and are interested in seeing their capital markets attracting more sophisticated investors as a means towards efficiency and stability.

On the other hand, our results suggest that listed firms in transition economies have an incentive in adopting IFRS in their financial reporting as a quality-signal of their disclosure practices in view of the increased presence of rational investors in these markets post-IFRS. This is particularly important in the case of overseas institutional traders, since information about transition economies at the company level can often appear ambiguous to foreign investors; thus, applying IFRS can boost a company's visibility towards such investors as it will indicate its commitment to pursuing internationally established reporting standards.

In addition, the findings of the paper present a clear message to accountants wishing to understand more precisely how the harmonization of accounting standards has affected the quality of company accounts and investors' trading behaviour. They imply that, for a more effective communication with the domestic and foreign groups of users, accountants should give high priority to adopting internationally accepted accounting standards such as IFRS.

Finally, our results may help settle the controversy surrounding the benefits and costs of IFRS adoption (Ball, 2006). Many proponents argue that the primary reason of moving towards the international accounting standards is to improve the quality and reliability of financial statements so the users of financial accounts can make more informed decisions. The findings of our paper suggest that to some extent this is being achieved in the Eastern European markets, providing a useful reference for many other countries which have recently introduced and/or been considering switching to IFRS as their mandatory accounting standards. They should help the accounting professions and regulators, such as the U.S. SEC, to make a decision on whether to incorporate these new standards into their financial reporting system.

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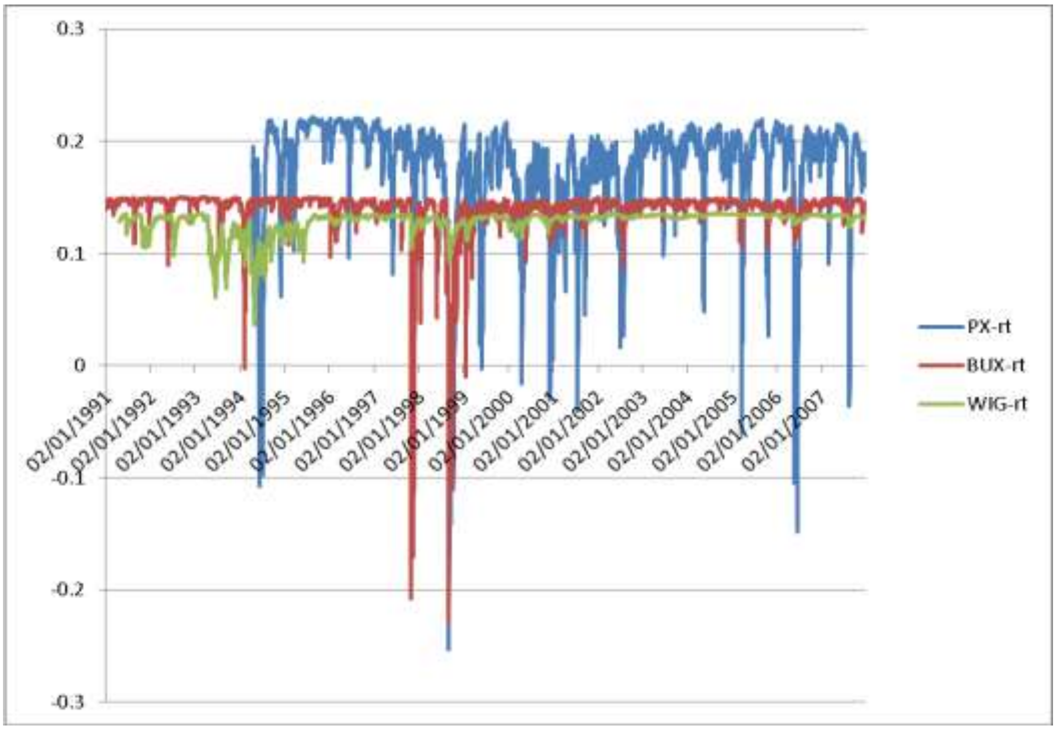
**Figure 1: Central and Eastern European (CEE) Stock Market Prices**


**Figure 2: IFRS Adoption and Conditional Return Autocorrelation**




**Notes:**

PX = Prague SE index, BUX = Budapest SE index, and WIG = Warsaw SE index

PXρt, BUXρt, and WIGρtare the conditional return autocorrelation, $\rho_t^{implied} = \hat{\varphi}_0 + \hat{\varphi}_1 \sigma_t^2$ implied by the original Sentana and Wadhwani's (1992) feedback model for the whole sample period.

## Table 1: Descriptive Statistics and Autocorrelation of Stock Market Returns

| | Czech Rep. | Hungary | Poland |
|---|---|---|---|
| **Panel A: Summary Statistics** | | | |
| **μ** | 0.017 | 0.074 | 0.092 |
| **σ** | 1.177 | 1.581 | 1.945 |
| **S** | -0.307 | -0.777 | -0.015 |
| **K** | 3.057 | 14.232 | 7.252 |
| **JB** | 1451.18 *** | 37855.69 *** | 9552.07 *** |
| **LB(12)** | 86.900 *** | 90.917 *** | 148.311 *** |
| **LB²(12)** | 370.606 *** | 1648.157 *** | 2452.859 *** |
| **ARCH(1)** | 75.197 *** | 431.996 *** | 172.843 *** |
| **JOINT** | 104.120 *** | 420.212 *** | 173.112 *** |
| **Panel B: Autocorrelation** | | | |
| **$b_0$** | 0.014 | 0.067 *** | 0.073 *** |
| | (0.713) | (2.807) | (2.488) |
| **$b_1$** | 0.110 *** | 0.084 *** | 0.135 *** |
| | (6.554) | (5.590) | (8.883) |
| **$b_2$** | 0.057 *** | 0.039 *** | 0.069 *** |
| | (3.415) | (2.602) | (4.490) |
| **$b_3$** | 0.007 | -0.044 *** | 0.002 |
| | (0.429) | (-2.939) | (0.162) |
| **$b_4$** | 0.024 | 0.024 | 0.008 |
| | (1.447) | (1.580) | (0.510) |
| **$b_5$** | -0.021 | -0.006 | 0.006 |
| | (-1.255) | (-0.415) | (0.428) |
| **F-test** | 13.189 *** | 9.723 *** | 23.217 *** |
| | <0.000> | <0.000> | <0.000> |
| **Durbin-Watson** | 1.999 | 1.999 | 1.999 |

**Notes:**

μ = sample mean; σ = standard deviation; S = skewness; K = Excess Kurtosis; JB = Jarque-Bera test for normality.

LB(n) and LB²(n) are the Ljung-Box Q test of serial correlation for the level & squared stock returns, respectively; the test statistics are distributed as $\chi^2$ with n degree of freedom where n is the number of lags. ARCH(1) is the Lagrange Multiplier LM test for ARCH effects and distributed as a $\chi^2$ with 1 degree of freedom. The test results for JOINT are Engle and Ng's (1993) test for the potential asymmetries in conditional volatility. The test statistic is a F-statistic for the null hypothesis of b1=b2=b3=0 of the following regression:

$$Z_t^2 = a + b_1 S_t^- + b_2 S_t^- \varepsilon_{t-1} + b_3 S_t^+ \varepsilon_{t-1} + v_t$$

where $Z_t^2$ is the square standardized residuals, $(\varepsilon_{t-1}/\sigma_t)^2$, $St^-$ is a dummy variable that takes a value of unity if $\varepsilon_{t-1} < 0$ and zero otherwise; and $St^+$ is a dummy variable that takes a value of unity if $\varepsilon_{t-1} > 0$ and zero otherwise.

The unconditional autocorrelation ($b_i$) estimates are obtained using the following autoregressive equation:

$$R_t = b_0 + \sum_{i=1}^{5} b_i R_{t-i} + u_t$$

The t-statistic and p-value are reported in ( ) and < >, respectively. *** indicates statistical significance at the 5% level.

## Table 2: Maximum Likelihood Estimates of the Feedback Model, Pre- vs. Post-IFRS

| | Czech Rep. | | Hungary | | Poland | |
|---|---|---|---|---|---|---|
| | Pre-IFRS | Post-IFRS | Pre-IFRS | Post-IFRS | Pre-IFRS | Post-IFRS |
| **Panel A: Mean Equation** | | | | | | |
| $\alpha$ | 0.0490 | 0.0482 | 0.0044 | 0.1669 | 0.0185 | 0.2308 *** |
| | (1.611) | (0.859) | (0.148) | (1.287) | (0.477) | (2.642) |
| $\theta$ | -0.0294 | 0.0236 | 0.0227 | -0.0474 | 0.0162 | -0.0968 |
| | (-1.228) | (0.438) | (1.512) | (-0.655) | (1.323) | (-1.583) |
| $\varphi_0$ | 0.3018 *** | 0.1083 *** | 0.1637 *** | 0.0842 | 0.1609 *** | 0.0872 |
| | (8.375) | (1.982) | (6.668) | (1.044) | (6.661) | (1.336) |
| $\varphi_1$ | -0.0612 *** | -0.0136 | -0.0056 *** | 0.0006 | -0.0041 | -0.0161 |
| | (-5.021) | (-1.020) | (-2.160) | (0.020) | (-1.588) | (-0.529) |
| **Panel B: Variance Equation** | | | | | | |
| $\alpha_0$ | 0.0279 *** | 0.1381 *** | 0.1217 *** | 0.1239 *** | 0.0700 | 0.0497 |
| | (2.083) | (2.412) | (2.029) | (2.279) | (1.787) | (1.800) |
| $\alpha_1$ | 0.0874 *** | -0.0180 | 0.1546 *** | 0.0663 *** | 0.0869 *** | 0.0530 |
| | (5.119) | (-0.515) | (5.380) | (2.083) | (3.450) | (1.686) |
| $\beta$ | 0.8755 *** | 0.7462 *** | 0.7770 *** | 0.8378 *** | 0.8978 *** | 0.8910 *** |
| | (40.758) | (11.781) | (12.310) | (19.579) | (27.430) | (45.997) |
| $\delta$ | 0.0399 | 0.2929 *** | 0.0582 | 0.0597 | -0.0073 | 0.0407 |
| | (1.564) | (2.548) | (1.010) | (1.359) | (-0.461) | (0.825) |
| **Panel C: Miscellaneous** | | | | | | |
| **Half-life** | 40.069 | 5.175 | 17.288 | 10.144 | 36.230 | 19.094 |
| **Asym. Ratio** | 1.457 | -15.272 | 1.376 | 1.900 | 0.916 | 1.768 |

**Notes:**
This table presents maximum likelihood estimates for the original Sentana and Wadhwani (1992) feedback trading model [i.e., the model given by equations (6) and (7)] for three national stock market indices during the pre- and post-IFRS periods. In particular, the estimated mean equation is

$$R_t = \alpha + \theta(\sigma_t^2) + (\varphi_0 + \varphi_1 \sigma_t^2) R_{t-1} + \varepsilon_t \qquad (6)$$

The variance equation is given by

$$\sigma_t^2 = \alpha_0 + \alpha_1 \varepsilon_{t-1}^2 + \beta \sigma_{t-1}^2 + \delta I_{t-1} \varepsilon_{t-1}^2 \qquad (7)$$

The estimated t-statistics (shown in parentheses) are robust to autocorrelation and heteroscedasticity using Bollerslev and Wooldridge (1992) standard errors. *** indicates statistical significance at the 5% level.

Half-life, calculated as $\ln(0.5)/\ln(\alpha_1 + \beta + \delta/2)$, represents the time it takes the shocks to reduce its impact by one-half. Asymmetric Ratio, given by $(\alpha_1 + \delta)/\alpha_1$, may be greater than, equal to, or less than 1 indicating negative asymmetry, symmetry, and positive asymmetry respectively.

## Table 3: Maximum Likelihood Estimates of the 'augmented' Feedback Model

| | **Czech. Rep** | **Hungary** | **Poland** |
|---|---|---|---|
| **Panel A: Mean Equation** | | | |
| $\alpha$ | 0.0590 *** | 0.0175 | 0.0598 |
| | (2.351) | (0.664) | (1.912) |
| $\theta$ | -0.0253 | 0.0199 | 0.0074 |
| | (-1.263) | (1.439) | (0.620) |
| $\varphi_{0, pre}$ | 0.2928 *** | 0.1655 *** | 0.1593 *** |
| | (9.785) | (6.871) | (12.905) |
| $\varphi_{0, post}$ | 0.0517 | 0.0958 | 0.1175 |
| | (1.613) | (1.589) | (1.741) |
| $\varphi_{1, pre}$ | -0.0569 *** | -0.0061 *** | -0.0040 *** |
| | (-5.969) | (-2.397) | (-2.755) |
| $\varphi_{1, post}$ | -0.0001 | -0.0002 | -0.0246 |
| | (-0.044) | (-0.012) | (-0.809) |
| **Panel B: Variance Equation** | | | |
| $\alpha_{0, pre}$ | 0.0399 *** | 0.1254 *** | 0.0308 *** |
| | (5.318) | (6.034) | (5.817) |
| $\alpha_{0, post}$ | 0.0361 *** | 0.1155 *** | 0.0638 *** |
| | (7.937) | (8.074) | (25.162) |
| $\alpha_1$ | 0.0878 *** | 0.1365 *** | 0.0807 *** |
| | (6.570) | (7.579) | (53.313) |
| $\beta$ | 0.8582 *** | 0.7915 *** | 0.9037 *** |
| | (92.214) | (215.475) | (1077.856) |
| $\delta$ | 0.0602 *** | 0.0581 *** | -0.0041 |
| | (3.515) | (2.762) | (-1.289) |
| **Panel C: Miscellaneous** | | | |
| **LR1** | 30.743 *** | 1.202 | 0.402 |
| **LR2** | 17.479 *** | 0.100 | 0.452 |
| **LR3** | 0.165 | 0.231 | 28.437 *** |
| **Half-life** | 28.654 | 15.789 | 38.924 |
| **Asym. Ratio** | 1.686 | 1.426 | 0.949 |

**Notes:**

This table presents maximum likelihood estimates for the 'augmented' Sentana and Wadhwani (1992) feedback trading model [i.e., model given by equations (8) and (9)] for three national stock market indices during the whole sample period to test for the parameter changes in pre- and post-IFRS periods.

In particular, the estimated mean equation is

$$R_t = \alpha + \theta(\sigma_t^2) + [(\varphi_{0,pre}D_t + \varphi_{0,post}(1-D_t)]R_{t-1} + +[(\varphi_{1,pre}D_t + \varphi_{1,post}(1-D_t)]\sigma_t^2 R_{t-1} + \varepsilon_t \quad (8)$$

The variance equation is given by

$$\sigma_t^2 = \alpha_{0,pre}D_t + \alpha_{0,post}(1-D_t) + \alpha_1\varepsilon_{t-1}^2 + \beta\sigma_{t-1}^2 + \delta I_{t-1}\varepsilon_{t-1}^2 \quad (9)$$

The estimated t-statistics (shown in parentheses) are robust to autocorrelation and heteroscedasticity using Bollerslev and Wooldridge (1992) standard errors. *** indicates statistical significance at the 5% level.

Dt is a dummy variable that is equal to 1 in post-IFRS and 0 in pre-IFRS.

LR is the likelihood ratio test statistic for $H_{01}$: $\varphi_{0, pre} = \varphi_{0, post}$ (LR1), $H_{02}$: $\varphi_{1, pre} = \varphi_{1, post}$ (LR2), $H_{03}$: $\alpha_{0, pre} = \alpha_{0, post}$ (LR3).

Half-life, calculated as $\ln(0.5)/\ln(\alpha_1 + \beta + \delta/2)$, represents the time it takes the shocks to reduce its impact by one-half. Asymmetric Ratio, given by $(\alpha_1 + \delta)/\alpha_1$, may be greater than, equal to, or less than 1 indicating negative asymmetry, symmetry, and positive asymmetry respectively.

## Table 4: Maximum Likelihood Estimates of the Feedback Model, Robustness Tests

| | | Czech Rep. | | Hungary | | Poland |
|---|---|---|---|---|---|---|
| **Panel A: Original Model with Asymmetry** | | | | | | |
| **Pre** | $\varphi_1$ | -0.0584 | *** | -0.0056 | *** | -0.0041 |
| | | (-4.813) | | (-2.110) | | (-1.446) |
| | $\varphi_2$ | 0.0467 | | 0.0652 | | -0.0185 |
| | | (1.313) | | (1.506) | | (-0.621) |
| **Post** | $\varphi_1$ | -0.0053 | | -0.0002 | | 0.0003 |
| | | (-0.323) | | (-0.004) | | (0.009) |
| | $\varphi_2$ | 0.1027 | | -0.0064 | | 0.0801 |
| | | (1.397) | | (-0.094) | | (1.271) |
| **Panel B: Original Model with Long Memory** | | | | | | |
| **Pre** | $\varphi_1$ | -0.0634 | *** | -0.0055 | *** | -0.0046 |
| | | (-5.295) | | (-1.98) | | (-1.468) |
| | $\varphi_1''$ | 0.0087 | | -0.0005 | | 0.0039 |
| | | (0.697) | | (-0.228) | | (1.336) |
| **Post** | $\varphi_1$ | -0.0117 | | 0.0011 | | -0.0187 |
| | | (-0.586) | | (0.040) | | (-0.564) |
| | $\varphi_1''$ | 0.0209 | | -0.0149 | | 0.0483 |
| | | (0.758) | | (-0.445) | | (1.342) |

**Notes:**
This table summarises the maximum likelihood estimates for the Sentana and Wadhwani (1992) feedback trading model [i.e., the model given by equations (6) and (7)] for three national stock market indices during the pre- and post-IFRS periods. In particular, the estimated mean equations are:

$$R_t = \alpha + \theta(\sigma_t^2) + (\varphi_0 + \varphi_1\sigma_t^2)R_{t-1} + \varphi_2\left|R_{t-1}\right| + \varepsilon_t$$ (Original Model with Asymmetry: (10))

$$R_t = \alpha + \theta(\sigma_t^2) + (\varphi_0 + \varphi_1\sigma_t^2)R_{t-1} + +(\varphi_0'' + \varphi_1''\sigma_t^2)R_{t-2} + \varepsilon_t$$ (Original Model with Long Memory: (11))

The variance equation is given by

$$\sigma_t^2 = \alpha_0 + \alpha_1\varepsilon_{t-1}^2 + \beta\sigma_{t-1}^2 + \delta I_{t-1}\varepsilon_{t-1}^2$$

The estimated t-statistics (shown in parentheses) are robust to autocorrelation and heteroscedasticity using Bollerslev and Wooldridge (1992) standard errors. *** indicates statistical significance at the 5% level.