

Back on the Map

Essays on Financial Markets in the Baltic States

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Abstract

This thesis consists of five self-contained papers, which are all related to the financial markets in the three Baltic States, Estonia, Latvia and Lithuania.

Paper [I] studies the impact of news from the Moscow and New York stock exchanges on the returns and volatilities of the Baltic States' stock market indices using a time series model that accounts for asymmetries in the conditional mean and variance functions. We find that news from New York has stronger effects on returns in Tallinn. High-risk shocks in New York have a stronger impact on volatility in Tallinn, whereas volatility in Vilnius is more influenced by high-risk shocks from Moscow. Riga does not seem to be affected by news arriving from abroad.

Paper [II] suggests a nonlinear and multivariate time series model framework that enables the study of simultaneity in returns and in volatilities, as well as asymmetric effects arising from shocks and exogenous variables. The model is employed to study the three Baltic States' stock exchanges. Using daily data, we find recursive structures, with returns in Riga, directly depending on returns in Tallinn and Vilnius, and Tallinn on Vilnius. For volatilities, both Riga and Vilnius depend on Tallinn.

Paper [III] studies the link between political news, and the returns and volatilities in the Baltic States' stock markets. We find that domestic and foreign non-Russian political news led, on average, to lower uncertainty in the stock markets of Riga and Tallinn in 2001-2003. At the same time, political risk from Russia increased the volatility of the stock market in Tallinn. There is a weak relationship between political risk and the stock market volatility in the Baltic countries in 2004-2007.

Paper [IV] studies the impact of market jumps on the time varying return correlations between stock market indices in the Baltic countries. An EARJI-EGARCH model facilitating direct modeling of the time varying return correlations is introduced. The empirical results indicate that there are quite a large number of identified jumps in the emerging Baltic States' stock markets. Isolated market jumps in one of the markets generally have no or small effects on the time-varying correlations. In contrast, simultaneous jumps of equal sign increase the average correlation, in some cases by as much as 100 percent.

In **Paper [V]** the hypothesis that financial development promotes economic growth is tested for the three Baltic countries using a time series approach that allows for interactions between the countries. We find that economic growth is a positive function of financial development, proxied by the amount of bank credit to the private sector, in the long run. The results also show that there is long run interaction between the three Baltic countries.

Key Words: Financial Markets, Time series, GARCH, Asymmetry, News

Acknowledgements

If, when I was studying foreign languages in Russia over ten years ago, somebody would have told me that I would write this thesis and start a career as an economist at a central bank, I would most likely have a good laugh. But as it happened, several years later, I found myself in Umeå, in my first Economics class. And what's more, I was actually enjoying it! For that I have to thank Kenneth Backlund, Jørn Stage, Olle Westerlund and several other teachers, for their inspiration and encouragement.

A couple of years later I surprised myself (and probably many others) again by joining the Ph.D. program in Economics at Umeå University. Who would have thought that after having spent some 16 years in the educational system I would be inclined to continue studying?! Even though my seemingly endless journey through academia has now (hopefully) come close to its end, there are many people, whose support and encouragement I've will treasure for many years to come.

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I've also treasured the advice from the co-author of the fourth paper of this thesis, Jörgen Hellström. Not least at times when I was puzzled with the econometrics of our paper or was trying to figure out the programming code for the model. I have also been lucky to work with Carl Lönnbark, co-author of the second paper of this thesis, who besides being a genius at computational econometrics had enough patience and

perseverance to help me estimate the “monster” model in our paper (Carl, I could not find a better word for it!).

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¹And of course, the usual disclaimer applies; the views expressed in this thesis are solely the responsibility of the authors and should not to be interpreted as reflecting the views of the Riksbank.

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Finally, I would like to express my deepest gratitude to my dear friends. Sergej and Natasha, you've become my best friends in the recent years. I can never thank you enough for being so generous, caring and understanding, even when I have to miss our evenings together because of work. I am certain that no matter what curve balls life throws at us our friendship will always remain strong. Maria, I am not sure that you will even get to this part of the thesis, since there are too many words and too little mathematical formulas. I must say that it has been a great joy to spend time and discuss things with you. Maybe this equation will help you understand: $J = \sum_{t=1}^T (D_t + \varepsilon_t)$, where J = my joy, T = time with you, D_t = discussions and ε_t = is the error term representing all the spontaneous and wicked things we come up with from time to time.

Further, I owe a great deal to Britt and Olle, Stella and Igor, Johan and Susanne. You all have been like a family to me for which I will be endlessly grateful.

Last but not least, I would like to express my sisterly love and my deepest appreciation to my sister Kristina. Dear sis, you have been there for me throughout my life and always stood behind me no matter what crazy adventures I have chosen to venture on. I am blessed to have you as a sister, not least for teaching me how to sail through deep oceans, both literally and figuratively.

Stockholm, December 2010

Albina

This thesis consist of a summary and the following five self-contained papers:

- [I] Brännäs K. and Sultanaeva A. 2006. Influence of News in Moscow and New York on Returns and Risks of the Baltic State Stock Indices. Umeå Economic Studies 696 (revised).
- [II] Brännäs K., De Gooijer J.G., Lönnbark C. and Sultanaeva A. 2007. Simultaneity and Asymmetry of Returns and Volatilities in the Emerging Baltic State Stock Exchanges. Umeå Economic Studies 725 (revised).
- [III] Sultanaeva A. 2008. Impact of Political News on Returns and Risks of the Baltic State Stock Markets. Umeå Economic Studies 735 (revised).
- [IV] Hellström J. and Sultanaeva A. 2010. What Drives the Stock Market Correlations in Baltic States: Volatility versus Jump Dynamics. Umeå Economic Studies 816.
- [V] Sultanaeva A. 2010. Financial Intermediation and Economic Growth: Evidence from the Baltic countries. Umeå Economic Studies 817.

"Models basically play the same role in economics as in fashion: they provide an articulated frame on which to show off your material to advantage ..." (Drèze, 1985)

Introduction and summary

This thesis consists of five independent, although related papers, dealing with several issues in empirical financial economics. Through the use of quantitative methods, the aim of this thesis is to empirically study relevant problems related to the functioning of the financial markets in the Baltic States. The research questions stretch all the way from cross market linkages, importance of extreme events and political news for the stock market dynamics, to a question of what role financial intermediation plays for economic growth. The papers have several things in common; they are all empirical and use time series analysis to study the relevant research questions. Also, they all provide additional knowledge on the financial markets of the three Baltic States, Estonia, Latvia, and Lithuania. Therefore, this helps us gain a deeper and more structured general understanding of these markets that are in an intermediate stage of development, or are so called markets in transition.

1 Introduction

In many financial theories and models, such as modern portfolio theory and option pricing, a trade-off between risk and return plays a prominent role. For example, the Capital Asset Pricing Model (CAPM) (e.g., Markowitz, 1959; Sharpe, 1964; Linter, 1965) postulates an investor's portfolio selection problem in terms of expected return on an asset and its variance, i.e. its risk or volatility. Similarly, the uncertainty associated with the price of the underlying asset, measured by its volatility is the most important determinant for the price of an option.

Obviously for these reasons, many studies have in addition to studying the returns of financial assets, been interested in the associated volatility, which is often regarded as a measure of risk. However, the choice of risk measure is to a large degree context dependent.² In this thesis, the focus is on market risk, which in the Financial Times Lexicon is defined as “the risk that an overall market or asset class will change in value according to economic conditions or other factors that may override any characteristics specific to a particular stock, bond, commodity or currency”. In particular, this thesis studies the factors that may affect the market risk in each of the three Baltic States’ stock markets.

Historically, a standard way of measuring risk has been through the variance of asset returns. Recent evidence however shows that the risk, also referred to as volatility, of financial assets is not constant, but changes over time. In particular, within financial markets, it can often be observed how large returns tend to be followed by large returns, and small returns tend to be followed by small returns (in both cases, either positive or negative). Hence, periods of high (low) return variance or volatility are often followed by periods when the variance of returns is high (low). Moreover, in empirical finance it is now well understood that financial time series data displays asymmetric behavior. An example of this behavior is that large negative returns appear more frequently than large positive returns. Indeed, on Friday October 24, 2008, Japan’s Nikkei index sank 9.6 percent, Stockholm’s OMX and Germany’s DAX index dropped 8.2 and over 9 percent, respectively; while we rarely observe positive daily returns of the same magnitude on most of the stock markets. Another example is that large negative returns are often a prelude to a period of substantial volatility, while large positive returns are less so. Needless to say such features should be incorporated in models used for studying the time-varying volatility. Throughout this thesis,

²For further discussion on different risk measures and views on risk, see e.g., Granger (2002).

the time series models capturing the characteristic features of financial time series data are used to describe the risk in the Baltic States' stock markets.

As illustrated above, many international stock markets in Asia, Europe, and even the US dropped by about 10 percent on Friday October 24, 2008. This is just one of many examples that illustrate the existing linkages between the financial markets across the world. In fact, academic research emphasized quite early that not only domestic but also, and more importantly, international factors play a role in the pricing of domestic securities.

In general, an accurate assessment of the degree of interdependence among international stock markets is important for several reasons. For investors that follow an international diversification strategy, the design of a well-diversified portfolio crucially depends on correctly understanding how closely international stock markets may be interlinked. Changes in international cross market linkages call for an adjustment of a portfolio. In addition, policy makers are interested in cross market linkages because of their implications for the stability of the international financial markets (e.g., Hartmann et al., 2004). To this end, Paper [I] of this thesis studies whether the return and volatility dynamics of the Baltic States' stock markets are influenced by the Russian and US stock markets.

It can also be argued that growing political and economic integration as well as technological advances in financial markets play a role for growing linkages between financial markets. Also, markets that are located in the same region may have stronger linkages than anticipated by investors (e.g., Fazio, 2007). Paper [II] suggests a model framework that enables us to study the joint evolution of stock market returns and volatilities, and applies the model on the three closely related Baltic States' stock markets.

While there is a general consensus on the importance of cross market

linkages, there is far less agreement on what causes changes in stock market volatility. Some studies show that volatility is driven by the arrival of new and unanticipated information that alters investors' expectations about stock returns (e.g., Engle and Ng, 1993). In particular, one may expect that specific news events, related, for example, to changes in the local or global economic or even political as well as regulatory environments, may explain changes in market returns or in the underlying volatility. More importantly, financial markets in their initial or intermediate stage of development are often particularly sensitive to political risk factors and news events (e.g., Bailey and Chung, 1995; Durnev et al., 2004; Goriaev and Zobotkin, 2006). Paper [III] of this thesis explores the importance of different political news for the stock market movements in the three Baltic countries.

Another important debate in empirical finance concerns the question of whether financial crises, extreme events, or just large shocks to a particular market, can increase the co-movements between international financial markets. There are numerous examples of how large shocks arising within a financial market of a particular country may be transmitted to other markets and countries. The most common examples are the Asian crises in 1997 or the Russian Bond default in August 1998, which both had repercussions for many international bond, equity, and currency markets. More recently, in the aftermath of the subprime crisis in the US, prices of many financial assets fell, which clearly illustrated that the dependence between financial assets has been underestimated. However, it may sometimes be difficult to define particular shocks, news, or other extreme events that have an impact on the co-movements or correlations between markets. To shed some light on the issue, Paper [IV] of this thesis studies the impact of extreme events, measured as large discrete changes in stock market returns, and referred to as jumps, on the co-movement between the stock markets in the three Baltic States.

In general, a better understanding of stock market developments and

co-movements between financial assets or markets is important for investors whose portfolio selection problem depends on the return and risk characteristics of assets included in the portfolio. However, in line with their development, financial markets are also becoming more important for the stability and growth of the economy (see, e.g., Pagano, 1993; Levine, 1997; Thiel, 2001; and Wachtel, 2003 for a review of the earlier literature and theoretical rationale). *The Economist* (2008) noted for example that, although stock markets play a relatively small role in the economies of East European countries, stock market declines in the aftermath of the global financial crises of 2008 may well have exacerbated the negative impact on domestic demand and economic actors' confidence, and hence on the growth in the region during this period. Still, many developing economies, including so called transition economies, have primarily a bank-based financial system (Berglöf and Bolton, 2002). Hence, Paper [V] of this thesis, instead of focusing on stock markets, examines the role the bank-based financial intermediation played for economic growth in the three Baltic countries.

In what follows, the topics introduced above are reviewed in more detail, and the contributions of this thesis are related to the existing literature. But first of all, the next section provides a brief description of the three Baltic States' stock markets.

2 The Baltic States' Stock Markets

Trading on the stock exchanges in the three Baltic States was first launched in the mid-1990s. In Estonia however, a foreign currency and securities exchange, the predecessor of the Tallinn Stock Exchange, was functioning between 1920 and March 1941, after which it was closed following Soviet occupation in Estonia. Still, unlike mature stock markets of advanced economies, the stock markets in the three Baltic countries have begun to develop rapidly, only in the last decade.

Table 1: Some basic facts about Baltic stock markets.

Year	2000	2001	2002	2003	2004	2005	2006	2007	2008	2009*
Market capitalization (MEUR)										
Riga	673	846	692	900	1207	2122	2034	2098	1166	1013
Tallinn	1923	1666	2316	3005	4627	3022	4578	4105	1403	1382
Vilnius	1481	1038	1409	2767	4753	6937	7728	6892	2608	2487
Market capitalization (% of GDP)										
Riga	7.9	9.1	7.0	9.0	10.8	16.3	12.7	9.9	5.0	4.8
Tallinn	31.5	24.1	29.9	34.6	47.9	27.2	34.9	26.9	8.8	9.4
Vilnius	12.0	7.6	9.4	16.8	26.2	33.2	32.2	24.2	8.1	8.3
Market turnover (MEUR)										
Riga	248	153	127	123	88	77	88	108	28	6
Tallinn	349	259	263	494	663	1938	766	1526	618	74
Vilnius	124	227	177	166	314	588	1607	757	332	75
Market turnover (% of Market cap)										
Riga	36.8	18.1	18.4	13.7	7.3	3.6	4.3	4.9	2.4	1.7
Tallinn	18.1	15.6	11.4	16.4	14.3	64.1	16.7	37.2	44.0	21.4
Vilnius	8.4	21.9	12.6	6.0	6.6	8.5	20.8	11.0	12.7	9.1
Number of companies										
Riga	63	63	62	56	39	45	40	41	35	34
Tallinn	21	18	15	15	14	16	16	18	18	17
Vilnius	47	46	51	43	42	43	42	40	40	40
Average company size (MEUR)										
Riga	10.7	13.4	11.2	16.1	31.0	47.2	50.9	51.2	33.3	29.8
Tallinn	91.6	92.5	154.4	200.3	330.5	188.9	286.1	228.1	78.0	81.3
Vilnius	31.5	22.6	27.6	64.3	113.2	161.3	184.0	172.3	65.2	62.2

* Indicates that statistics are for the first half of the year. MEUR= Million Euro.

Source: Guide to NASDAQ OMX Baltic Securities Market (2009).

In Table 1, the yearly turnover volume, market capitalization at the end of the year, number of listed companies and average company size are presented for the stock exchanges in Estonia (Tallinn), Latvia (Riga), and Lithuania (Vilnius) between 2000 and the first half of 2009.

The table shows that the three Baltic States' stock exchanges are still relatively small, even though they developed well in the years prior to the financial crisis in 2007-2008. The market capitalization more than doubled for all three stock exchanges over the period 2000-2006. By the end of 2006, the stock markets in the Baltic countries had a capitalization of 12 to 35 percent relative to domestic GDP, whereas for most of the developed country stock exchanges market capitalization was well above 100 percent. As such, it seems more reasonable to compare the Baltic States' stock markets to other stock markets in Central and Eastern Europe, i.e. so called transition economies, where, for example, the Warsaw stock exchange had a market capitalization to GDP of 41 percent in the end of 2006.³

Another important characteristic of stock markets is their liquidity, which is often measured by market turnover defined as the ratio of the total value of trading to market capitalization. A high ratio indicates that the market is relatively liquid. The three Baltic exchanges showed low liquidity, where the Tallinn stock exchange was the most liquid one during the considered time period. The market turnover on the Riga stock exchange has declined dramatically since 2000, which could in part be explained by the decreasing number of listed companies. Overall, it is of no surprise that the stock markets in the three Baltic States have been less liquid during the 2000s than their counterparts in developed countries. However, Estonian stock market seems to have performed well in terms of liquidity when compared to, for example, the Warsaw stock exchange, where the market turnover was about 40 percent in 2005 and 2006. In general, the lower market turnover in CEE countries, compared

³Source: The World Federation of Exchanges (WFE).

to developed markets, can be attributed to ownership concentration. Indeed, over 85 percent of investors in the Estonian securities market had investments amounting to 10 million Estonian Kroon (or over 600 thousand Euros) by the end of 2009.⁴

In addition, the stock markets in the Baltic States have a large representation of institutional investors. For example, by the end of 2009, institutional investors represented 79 percent of the total market participants of the Latvian Stock exchange, whereas in Estonia and Lithuania about 92 percent of total investments were made by the institutional investors. Also, foreign investors represent a large share of total investments in the three markets. For example, by the end of 2009 non-residents represented 34, 53, and 65 percent of total investments in the stock exchanges in Lithuania, Latvia, and Estonia, respectively. In Estonia, Nordic investors represented over 48 percent of total investments on the Tallinn stock exchange. This is not surprising, since, as Bengtsson et al. (2007) found, different performance measures clearly illustrate that the Baltic exchanges outperformed Nordic exchanges during the whole period of 2000-2006.

The first four papers of this thesis study the returns and risks on the three Baltic stock markets in more detail. In particular, this thesis considers such issues, as cross-market linkages, since they can affect the decision making process of institutional investors following an international diversification strategy.

3 Background and contribution of the thesis

Most of the studies within financial economics describe asset returns as a function of past returns⁵ and possibly other explanatory variables. In

⁴Source: The Estonian Central Register of Securities.

⁵Campbell, Lo and MacKinlay (1997) provide two main reasons for using returns rather than asset prices. First, for an average investor, return on an asset provides a complete summary of the investment opportunity independent of the size of an

other words, the returns of financial assets are commonly modeled with various Autoregressive Moving Average (ARMA) model specifications. However, it seems reasonable to assume that investors react differently to positive and negative shocks (up markets versus down markets).⁶ The Autoregressive asymmetric Moving Average (ARasMA) model of Brännäs and De Gooijer (2004), which is an extension of the asymmetric Moving Average (asMA) model proposed by Wecker (1981) is a suitable candidate for this purpose since it allows asset returns to respond asymmetrically to own past innovations (or shocks).

As noted in the Introduction, one of the prominent stylized facts of financial asset returns is that their volatility changes over time. In particular, periods of large movements in prices alternate with periods during which prices hardly change. This characteristic feature, commonly referred to as volatility clustering, can be studied using the Autoregressive Conditional Heteroscedastic (ARCH) model of Engle (1982). The ARCH model and its extensions, most notably the Generalized ARCH model (GARCH) introduced by Bollerslev (1986) has been very successful in modeling time varying volatility in financial series.

In addition, volatile periods are often initiated by a large negative shock, which suggests that positive and negative shocks may have an asymmetric impact on the conditional variance or volatility. This so called “leverage effect” was first acknowledged by Black (1976) and is one of the most common explanations for the asymmetry in time-varying volatility.⁷ Standard GARCH models cannot capture such asymmetric

investment. Second, return series have more attractive statistical properties, such as stationarity, which makes them more appropriate for most of the time series models for financial assets or markets.

⁶Morris and Shin (2004) distinguish, for example, between risk-averse long horizon traders and short horizon traders, whose incentives to sell the asset increase when asset prices fall close to their loss limit.

⁷Black (1976) suggested that a possible explanation for this phenomenon may be the way firms are financed. When the value of a stock of a particular firm falls, the debt-to-equity ratio, commonly referred to as leverage of the firm, increases, which in turn may lead to an increase in the volatility of the returns on equity.

effects of positive and negative shocks. Instead, positive and negative shocks of the same magnitude have the same effect on conditional volatility (or risk) – that is, the sign of the shock is not important. Hence, several non-linear extensions of the standard GARCH models, designed to allow for different effects of positive and negative shocks on the time-varying risk, have been developed in the literature. The extensions of the GARCH models, which allow for asymmetric effects, include, for example, the Exponential GARCH (EGARCH) model of Nelson (1991), the GJR-GARCH model introduced by Glosten, Jagannathan and Runkle (1993) and the Quadratic GARCH (QGARCH) model by Sentana (1995), and its extension to the asymmetric Quadratic GARCH (asQGARCH) model by Brännäs and De Gooijer (2004). The extended versions of these types of models are used in this thesis (Paper [I] - [IV]) to study the dynamics of the stock markets in the Baltic States.

In particular, Paper [I] of this thesis extends the ARasMA-asQGARCH model of Brännäs and De Gooijer (2004) to take into account the fact that changes in stock market prices may be driven not only by own shocks but also, by the stock market reaction to movements in other markets. The extended model allows us to capture any potential asymmetric impact of positive and negative shocks from the most influential US and closely located Russian stock markets on the return and volatility dynamics of the Baltic States' stock markets. The transmission of shocks in this paper could either be explained by the real economic, financial, or even political interrelations between the countries or may be a result of the behavior of institutional investors who rebalance their portfolio whenever new information arrives.

The presence of strong economic ties and policy coordination between countries, may also explain why markets that are located in the same region may have strong linkages (e.g., Koch and Koch, 1991; Chen et al., 2002). In fact, given that shocks are commonly interpreted as news, and that different closely related financial markets may be af-

ected by at least certain news events simultaneously, it is of interest to consider multivariate models that describe joint movements in the financial time series. Consequently, multivariate models with GARCH-type specifications for the time-varying volatility emerge as a natural extension of the univariate models (see Bauwens, Laurent and Rombouts, 2006, for a survey). An alternative motivation for multivariate models is that the construction of a well-diversified portfolio crucially depends on the co-movements between assets included in the portfolio. Indeed, the ability to capture spillovers or shocks that are transmitted across different assets and markets is the most important feature of the multivariate models (e.g., Karolyi, 1995; Bonfiglioli and Favero, 2005). To this end, Paper [II], [III] and [IV] of this thesis, study the dynamics of the Baltic States' stock markets while allowing for interaction between the markets using various multivariate time series model frameworks.

In particular, Paper [II] of this thesis suggests a nonlinear multivariate time series model framework that enables us to study simultaneity in stock market returns and volatilities using daily data. Hence, this draws a lesson from the earlier literature suggesting that stock markets can move simultaneously in part because they are exposed to common information. In addition, the information transmission between markets is particularly fast, if it is not limited by institutional constraints and other practical considerations such as common trading platform (Fleming et al., 1998). Indeed, several models allowing for simultaneity in returns models have been developed quite recently (e.g., Rigobon and Sack, 2003; De Wet, 2006; Lee, 2006). Obviously, and more interestingly, market risks, i.e. volatilities, may also move simultaneously. Gannon and Choi (1998) and Gannon (2004, 2005) have addressed this question using realized volatilities, i.e. squared returns. However, given the attractive features of the non-linear GARCH models it seems natural to extend this model framework to allow for simultaneity in volatilities. Paper [II] of this thesis is, to the authors' knowledge, the first

in the literature to propose a non-linear multivariate model that allows for simultaneity in both returns and volatilities along with asymmetric effects.

Several studies have also showed that return and volatility dynamics in different financial markets is related to the underlying information flow, including the arrival of new and unanticipated information (e.g., Andersen and Bollerslev, 1998; Chang and Taylor, 2003; Kalev et al., 2004). Hence, besides capturing the impact of various shocks on the return and volatility dynamics, it may be of interest to study the effects of different public news announcements that reflect changes in the local or global economic, political, or regulatory environments. In particular, political news has been found to play a role in the stock market dynamics of emerging markets (e.g., Bailey and Chung, 1995; Chan and Wei, 1996; Chan et al., 2001; Kim and Mei, 2001). To this end, Paper [III] of this thesis contributes to this literature by studying the importance of different political news for the stock market movements in the Baltic countries. The underlying motivation for our analysis is, in part, due to the fact that emerging and transition markets are particularly sensitive to political factors and events (e.g., Bailey and Chung, 1995; Durnev et al., 2004; Goriaev and Zabotkin, 2006). The impact of different political news on the behavior of the Baltic States' stock markets is also of special interest because of the abundance of political events and recent developments in the countries.

Although the models described above are adequate in terms of accounting for volatility clustering, cross-market linkages (or spillovers), and asymmetric effect of positive and negative shocks, they do not fully explain sudden and large discrete changes (referred to as jumps or unusual news events) often found in financial assets returns (e.g. Mahieu and McCurdy, 2004). Such features can instead be studied using a mixed GARCH-jump model, where the GARCH part explains the smooth changes in volatility while the jump part explains infrequent

large discrete movements in asset returns (e.g., Press, 1967; Chan and Maheu, 2002; Maheu and McCurdy, 2004). Moreover, a mixed GARCH-jump model can be particularly useful in explaining the changing nature of cross-market co-movements or correlations.⁸ To shed some light on the issue, Paper [IV] introduces a model that enables us to study the impact of large discrete changes in stock market returns (i.e. jumps) on time-varying return correlations. More specifically, it studies whether the correlations of returns between the Baltic States' stock markets differ when there are smooth changes in market returns and large discrete changes.

To empirically assess the sources and characteristics of changing return and volatility dynamics is essential for many financial and economic decisions, not least for portfolio composition, risk measurement, and risk management. In the case of the Baltic States, understanding the stock market dynamics is particularly important due to the large participation of institutional and international investors. However, in transition countries other aspects of financial intermediation, namely the banking sector often plays a greater role for the domestic real economy (e.g., Berglöf and Bolton, 2002). Still, despite the period of high economic growth combined with rapid expansion of the banking sector (primarily foreign-owned banks) during several years prior to the financial crises of 2007-2008, the recent experiences regarding the relationship between financial intermediation and economic growth in the three Baltic States have, to the author's knowledge, not been studied in the academic lit-

⁸For example, there is ample empirical evidence that that cross-market linkages increase during after a large shock to one country, or group of countries or during the financial crises (e.g., Karolyi and Stulz, 1996; Ramchand and Susmel, 1998; Longin and Solnik, 1995, 2001). This facteffect has been commonly referred to as contagion (e.g., Claessens, 2001; Forbes and Rigobon, 2002) between the financial markets.

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erature. Hence, while Paper [I]-[IV] focuses on the stock markets in the three Baltic States, Paper [V] of this thesis uses the Johansen (1988, 1995) time-series approach to explore the role bank-based financial intermediation plays in the economic development of the three Baltic States (see e.g., Pagano, 1993; Levine, 1997; Thiel, 2001; and Wachtel, 2003 for a review of the earlier literature and theoretical rationale).

4 Summary of the papers

Paper [I]: Influence of News from Moscow and New York on Returns and Risks of Baltic State Stock Markets

This paper studies whether the US and Russian stock markets influence the price and volatility dynamics of the Baltic States' stock markets. Obviously, the question of whether the Baltic stock markets react to the information flow from abroad becomes even more important given a large proportion of foreign institutional investors in the Baltic markets. In particular, if international investors seek to diversify their portfolio by shifting their trading into international markets, risk reduction may prove difficult if the stock markets in the Baltic States are strongly interlinked with other markets, and in this case with the US and Russian stock markets.

The US stock market is often considered as the most influential producer of information that is transmitted to other stock markets (e.g., Eun and Shim, 1989; Koch and Koch, 1991; Liu and Pan, 1997; Forbes and Chinn, 2004). This has also been the case for the stock markets in Central and Eastern European (CEE) countries (see, for example, Tse et al., 2003, for the results for the Warsaw Stock Exchange). Spillovers from the Russian stock market can on other hand be explained by economic, historical, and political ties between the countries as well as the geographical proximity (e.g., Koch and Koch, 1991; Pajuste et al., 2000; Forbes and Chinn, 2004).

This paper employs an econometric model that is designed to capture the return and volatility dynamics of the Baltic stock markets as well as impacts of foreign shocks. Furthermore, this paper allows for an asymmetric impact of both domestic and foreign shocks. For example, to study the asymmetric impact of shocks (or news) from abroad on the stock market returns, news is split into good and bad, where the term "good news" denotes positive past returns and the term "bad news" denotes negative past returns. In order to capture the asymmetric impact of volatility shocks from abroad, shocks are instead categorized as being of "high" and "low" volatility.

The overall findings suggest that there are substantial differences among Baltic States' stock markets, with respect to the market adjustment to information arriving from abroad. For example, news from New York has stronger effects on returns in Tallinn than news from Moscow. High-risk shocks in New York have a stronger impact on volatility in Tallinn, whereas the volatility of Vilnius is more influenced by high-risk shocks from Moscow. Riga does not seem to be affected by news arriving from abroad.

Paper [II]: Simultaneity and Asymmetry of Returns and Volatilities: The Emerging Baltic States' Stock Exchanges

This paper suggests a nonlinear time series model framework that enables us to study simultaneity in stock market returns and volatilities, as well as asymmetric effects arising from shocks. We argue that simultaneity is important if there are strong linkages between markets. Information processing may be in particular fast in some marketplaces, where linkages between markets are not limited by institutional constraints, and other practical considerations such as a common trading platform. Also, simultaneity may be particularly relevant for markets that may be exposed to common information, for example because they belong to the same geopolitical region.

This paper focuses on the joint modeling of the three Baltic stock markets, while allowing for simultaneity in both returns and volatilities. Furthermore, this paper allows for asymmetry in returns and volatility dynamics and the impact of exogenous variables, such as spillovers from the Russian stock market.

The estimation results indicate recursive structures with returns in Riga directly depend on returns in Tallinn and Vilnius, and Tallinn on Vilnius. For volatilities, both Riga and Vilnius depend on Tallinn. In addition, we find evidence of asymmetric effects of shocks arising in Moscow and in Baltic States on both returns and volatilities.

In addition, the paper outlines the benefits of the suggested model for portfolio allocation and value at risk (VaR) studies. Portfolio allocation results indicate that optimal portfolio weights are more sensitive to shocks when simultaneity is not accounted for. Also, VaR measures indicate that the variability in losses that may occur due to shocks in the market is larger when simultaneity is not accounted for.

Paper [III]: Impact of Political News on the Baltic Stock Markets

This paper examines the importance of different publicly available news releases for the stock market movements in the Baltic countries. More specifically, it looks at the number of political news headlines during a day, as a proxy for the information flow.

The underlying motivation for the analysis is in part due to the fact that emerging and transition markets are particularly sensitive to political factors and events (e.g., Bailey and Chung, 1995; Durnev et al., 2004; Goriaev and Zobotkin, 2006). Given the recent historical development of the Baltic States, it is interesting to study whether the origin of political risk matters for investors' perception of market risk. In particular, this paper explores whether the political risk (i.e. political events) related to Russia versus the domestic and other foreign political

issues, excluding Russia, have different impacts on the stock markets in the Baltic countries.

A priori, it seems reasonable to assume that the risk factors related to Russia have become less important over time, whereas European factors have become more important for investors' perception of the market risk in the Baltic countries. Thus, in the paper we consider two different time periods to explore whether the sensitivity to different political events has changed over time.

We employ a multivariate time series model designed to catch the impact of news on returns and volatility. Besides capturing the news impact, this model allows us to capture asymmetric effects of positive and negative shocks on volatility as well as volatility spillovers across markets. Also, the model allows us to study whether news affecting the market risk in one of the Baltic countries has any impact on the market risk in the other two countries.

Overall, our results indicate that the impact of news can depend on both the origin and the nature of political news events, but also that the sensitivity to political factors changed over time following the favorable economic and political development of the markets in transition, and the Baltic countries in particular.

Paper [IV]: The Impact of Stock Market Jumps on Time-Varying Return Correlations: Empirical Evidence from the Baltic Countries

This paper studies the impact of large discrete changes in stock market returns, i.e. jumps, on time-varying return correlations. More specifically, we study whether the correlations of stock market returns differ when there are smooth changes in market returns or jumps.

The paper contributes to the existing literature in several ways. First, this paper extends an existing model framework that jointly captures smooth changes in returns and jumps, to a specification that also

considers time-varying return correlations. Second, we utilize a data driven procedure to identify jumps, whereas in the majority of the earlier literature that considers transmission of large shocks between financial markets, shocks, extreme events, or market crashes are usually pre-defined by authors. Finally, the effect of return jumps on the time-varying return correlations is studied more directly than in the previous literature on jumps, which has primarily focused on the correlations between the jumps or on jump spillovers. We argue that high correlation between the jumps does not necessarily imply a high correlation between returns since the jumps may be in different directions (i.e. leading to lower or higher returns).

The study is performed on stock market data for the three Baltic countries. The empirical results indicate that there are quite a large number of identified jumps in the Baltic States' stock markets. The main finding is that isolated market jumps in one of the markets generally have no or small effects on the time-varying correlations. In contrast, simultaneous positive jumps increase the average correlation, in some cases with as much as 100 per cent, whereas simultaneous negative jumps have a smaller effect on the average correlation.

Paper [V]: Financial Intermediation and Economic Growth: Evidence from the Baltic countries

The purpose of this study is to contribute to the empirical evidence on the relationship between bank-based financial intermediation and economic growth in the Central and Eastern Europe (CEE) countries. In particular, we examine the relationship between financial sector development, proxied by the level of bank credit to the private sector, and economic growth in the Baltic countries over the period 1995-2008. In addition, since all three Baltic countries are likely to be closely inter-related, we use a time-series approach that allows for cross-country dependence in the empirical analysis.

Overall, the results indicate that banking sector development can cause economic growth in the long run in the Baltic States. The results also show that the three Baltic countries are indeed interrelated. Using impulse responses, we illustrate for example that economic growth in Latvia and Lithuania reacts positively to shocks in economic growth in Estonia. Also the credit development in Latvia and Lithuania responds positively to a shock in economic growth in Estonia. These interactions can in part, be explained by the trade pattern between the countries or the fact that the banking sector in three Baltic States is dominated by the same foreign-owned banks that reallocate capital over the geographical region on the basis of expected returns and risks.

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Influence of News from Moscow and New York on Returns and Risks of Baltic State Stock Markets*

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Abstract

The impact of news from the Moscow and New York stock exchanges on the daily returns and volatilities of the Baltic stock market indices is studied. A nonlinear time series model that accounts for asymmetries in the conditional mean and variance functions is used for the empirical work. News from New York has stronger effects on returns in Tallinn than news from Moscow. High-risk shocks in New York have a stronger impact on volatility in Tallinn, whereas volatility of Vilnius is more influenced by high-risk shocks from Moscow. Riga seems not to be affected by news arriving from abroad.

Key Words: Estonia, Latvia, Lithuania, Time series, Estimation, Finance.

JEL Classification: C22, C52, G10, G15.

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1 Introduction

In this paper, we study whether the two leading US (New York, NYSE Composite) and Russian (Moscow, RTS) markets influence the price and volatility dynamics of the Baltic states' stock markets of Riga, Tallinn and Vilnius. We focus on these three markets for several reasons. The stock markets in the Baltic countries have previously received little attention in the literature. Thus, portfolio or fund managers that look for diversification of their portfolios, may benefit from additional information on these stock markets, as it may reduce the uncertainty about such an investment. In addition, given a common institutional setup in terms of a common main owner and trading platform, institutional investors can trade on all three markets with relative ease. In fact, foreign and domestic institutional investors together outweigh individual investors, with about 90 percent of the market value, whereas foreign institutional investors, predominately European ones, represent 40-47 percent of the market value in the Baltic stock markets (OMX, 2007).¹

Obviously, given a large proportion of foreign institutional investors, an interesting issue is whether the three Baltic stock markets react differently to the information flow from abroad. In particular, if there are strong links between the stock markets included in a portfolio or similarities in spillovers from abroad, international investors participating in all three markets could be exposed to unhedged risk and risk reduction may prove difficult. For example, little interdependence between the Nordic and Baltic stock markets (Nielsson, 2007), could, in fact, explain why investors from the Nordic countries represent about 40 percent of total investment in the Tallinn stock exchange down to about 11 percent in Vilnius.

The research literature established quite early that not only new domestic information, but also information from other markets, can be

¹For more information about Baltic states' stock markets, see the Appendix.

incorporated in the pricing of domestic securities. In general, there are several reasons for possible information spillovers between markets. The first reason is that common information may simultaneously affect expectations in more than one market. A second reason is cross-market hedging (Fleming et al., 1998). In particular, spillovers from the Russian stock market can be explained by economic, historical and political ties between the countries (e.g., Koch and Koch, 1991). Koch and Koch (1991) also note that interdependence is stronger among countries in the same geographic region, whose trading hours overlap. Pajuste et al. (2000) find that East European countries are likely to be affected by news coming from Russia.

Information from the US stock market is useful to include, because the US market is the most influential producer of information (e.g., Eun and Shim, 1989; Koch and Koch, 1991; Liu and Pan, 1997). In this paper, the US stock market is considered a proxy for the global information that may affect the market participants' confidence or expectations. For markets in transition, Tse et al. (2003) found that the volatility of the Warsaw Stock Exchange is not influenced by past volatility in the US market. However, they also found that there is a significant return spillover from the US market on the Polish stock market. We are interested in studying whether news from the major US stock market has more influence on the Baltic States stock markets than news from the smaller but closer Russian stock market.

To study the foreign influence, news that affects return processes is split into good and bad news, where the term "good news" denotes positive past returns and the term "bad news" denotes negative past returns. Given that volatility is related to the flow of information (Ross, 1989), we also study whether volatility shocks observed in the Russian and US markets are relevant for volatility dynamics in the three Baltic stock markets. To capture the asymmetric impact of volatility shocks from abroad, shocks are categorised as being of "high" and "low" volatility.

The econometric model that is employed is designed to represent both the asymmetric international influence and asymmetric impacts of domestic innovations on the price and volatility processes. The current study builds on previous research in several ways. Previous studies (e.g., Koutmos and Booth, 1995) found that the volatility transmission is often asymmetric with respect to positive and negative innovations. Black's (1976) leverage effect, which states that volatility increases to a greater extent after a negative shock than after a positive one, is one of the most common explanations for the asymmetry in volatility. However, for emerging markets it is possible that positive innovations cause volatility to increase more than do negative innovations. Rockinger and Urga (2001) found this pattern for Hungary, and suggest that "for countries, suffering from low liquidity, one can imagine scenarios where good news can lead to increased liquidity, which in turn can lead to increased volatility as investors rebalance their portfolios". Similar results were found by Bekaert and Harvey (1997) for some emerging markets. It is also possible that the conditional mean responds asymmetrically to past innovations (Wecker, 1981; Koutmos, 1998). To capture such features, we combine the ARasMA model of Brännäs and De Gooijer (1994) for the conditional mean with an asymmetric parameterization of the conditional variance. The volatility process is modelled as an asymmetric extension of the quadratic GARCH model of Sentana (1995). The resulting ARasMA-asQGARCH model (Brännäs and De Gooijer, 2004) allows us to detect asymmetry in both the conditional mean and variance of stock return data. We extend this model to capture any potential asymmetric impact of good and bad news from the two US and Russian marketplaces.

The remainder of the paper is organized as follows. Section 2 introduces the ARasMA-asQGARCH model and presents the estimation method. Section 3 discusses the data. Section 4 gives the empirical results and presents implications for the composition of portfolios and the

measurement of risk. The major findings are summarized in the final section.

2 Model and Estimation

To account for the possibly asymmetric effects of news in Moscow (RTS) and New York (NYSE) on the stock market indices of the Baltic states, we expand the conditionally heteroskedastic ARasMA specification of Brännäs and De Gooijer (2004), hereafter BDG (see also Wecker, 1991; Brännäs and De Gooijer, 1994). The effects of news are allowed to affect both the conditional mean (return) and heteroskedasticity (volatility or risk) functions.

Let $\{u_t\}$ be a real-valued discrete-time stochastic process generated by

$$u_t = \varepsilon_t h_t \quad (1)$$

where $\{\varepsilon_t\}$ is a sequence of independent and identically distributed random variables with mean zero and unit variance, and the conditional standard deviation h_t is independent of ε_t as well as non-negative for all t . Further, let

$$u_t^+ = \max(0, u_t) = \varepsilon_t^+ h_t \quad \text{and} \quad u_t^- = \min(u_t, 0) = \varepsilon_t^- h_t$$

where $\varepsilon_t^+ = \max(0, \varepsilon_t)$ and $\varepsilon_t^- = \min(\varepsilon_t, 0)$. In an analogous way, let $x_t^+ = \max(0, x_t)$ and $x_t^- = \min(0, x_t)$ be the positive and negative return at time t , respectively, in the Moscow and/or New York return series.

The autoregressive asymmetric moving average (ARasMA) model of order (p, r, q) , is then defined as

$$\begin{aligned} y_t = & \sum_{i=1}^p \alpha_i y_{t-i} + \sum_{i=1}^r (\gamma_i^+ x_{t-i}^+ + \gamma_i^- x_{t-i}^-) + \beta_0 \\ & + u_t + \sum_{i=1}^q (\beta_i^+ u_{t-i}^+ + \beta_i^- u_{t-i}^-) \end{aligned}$$

$$\begin{aligned}
&= \sum_{i=1}^p \alpha_i y_{t-i} + \sum_{i=1}^r (\gamma_i^- x_{t-i} + \gamma_i^\nabla I(x_{t-i} \geq 0) x_{t-i}) + \beta_0 \\
&\quad + u_t + \sum_{i=1}^q (\beta_i^- u_{t-i} + \beta_i^\nabla I(u_{t-i} \geq 0) u_{t-i}). \tag{2}
\end{aligned}$$

Here, $\gamma_i^\nabla = \gamma_i^+ - \gamma_i^-$, $i = 1, \dots, r$, $\beta_i^\nabla = \beta_i^+ - \beta_i^-$, $i = 1, \dots, q$, and $I(\cdot)$ is the indicator function. Given that the values of the β_i^+ and β_i^- parameters at the i th lag may be different, the response to equally sized positive and negative shocks may be different or asymmetric. The inherent asymmetry of the asMA model was illustrated numerically by Brännäs and Ohlsson (1999). Obviously, if $\gamma_i^+ = \gamma_i^-$ for all i , the responses to positive and negative news in the Moscow and/or New York return series is symmetric.

The conditional mean (return) of y_t given past observations is

$$\begin{aligned}
E(y_t | Y_{t-1}) &= \sum_{i=1}^p \alpha_i y_{t-i} + \sum_{i=1}^r (\gamma_i^+ x_{t-i}^+ + \gamma_i^- x_{t-i}^-) + \beta_0 \\
&\quad + \sum_{i=1}^q (\beta_i^+ u_{t-i}^+ + \beta_i^- u_{t-i}^-).
\end{aligned}$$

Note that containing several rather than one x_t series in the model presents no additional difficulty.

Various models have been proposed to represent the conditional heteroskedasticity h_t^2 in (1). Sentana (1995) introduced the QGARCH(P, P) model and BDG the Asymmetric Quadratic Generalized ARCH (asQGARCH) model of order (Q, P, P). To account for asymmetric effects through a variable z_t from Moscow and/or New York we expand the latter to obtain

$$\begin{aligned}
h_t^2 &= \alpha_0 + \sum_{i=1}^R (\eta_i^+ z_{t-i}^+ + \eta_i^- z_{t-i}^-) + \sum_{i=1}^Q (\alpha_i^+ u_{t-i}^+ + \alpha_i^- u_{t-i}^-) \\
&\quad + \sum_{i=1}^P \kappa_i u_{t-i}^2 + \sum_{i=1}^P \rho_i h_{t-i}^2
\end{aligned}$$

$$\begin{aligned}
&= \alpha_0 + \sum_{i=1}^R (\eta_i^- z_{t-i} + \eta_i^\nabla I(z_{t-i} \geq 0) z_{t-i}) \\
&\quad + \sum_{i=1}^Q (\alpha_i^- u_{t-i} + \alpha_i^\nabla I(u_{t-i} \geq 0) u_{t-i}) \\
&\quad + \sum_{i=1}^P \kappa_i u_{t-i}^2 + \sum_{i=1}^P \rho_i h_{t-i}^2. \tag{3}
\end{aligned}$$

The first term of this conditional variance (risk) function accounts for asymmetric effects in either or both of the Moscow and New York series around some threshold level \bar{z} . The second block on the right-hand side describes the asymmetry in the conditional variance. The former part of u_t^+ and u_t^- may cause a problem with the positivity of h_t^2 unless parameters are constrained, e.g., such that the effects of u_{t-i} and u_{t-i}^2 are positive. In (3) positive shocks have a different effect than negative shocks. The response of the process is parabolic, though not symmetric around zero.

If $\alpha_i^\nabla = 0$ for all $i = 1, \dots, Q$ and $Q = P$, (3) reduces to an extended QGARCH(P, P). We also see that when $\alpha_i^+ = \alpha_i^- = 0$, (3) simplifies to the extended GARCH model of order (P, P) introduced by Bollerslev (1986). Note, however, that in the case of $Q = P = 1$, (3) differs from the so-called Asymmetric Threshold GARCH (asTGARCH) of order (1;1,1) of Koutmos (1999), which is an asymmetric analogue of the TGARCH(1,1) model of Zakoïan (1994).

Unconditional moments are difficult to obtain, but are given for the case of no x_t and z_t variables in BDG and for a model with constant h_t^2 by Brännäs and De Gooijer (1994) for normally distributed $\{\varepsilon_t\}$ sequences. Some related model properties for log-generalized gamma and Pearson IV distributed $\{\varepsilon_t\}$ sequences are discussed by Brännäs and Nordman (2003a,b).

2.1 Empirical Modelling Strategy

To find empirical models, we adopt a four-step procedure. First, we find the best ARasMA model for each Baltic stock exchange. Second, this ARasMA model is augmented with an asQGARCH model for conditional heteroskedasticity. Third, we expand each specification of the second step by, in turn, including Moscow and New York in both the conditional mean and conditional variance functions. This allows us to test whether Moscow and/or New York cause mean returns or volatilities. The conditional mean and variance functions are allowed to respond asymmetrically to news in the Moscow and New York series. Fourth, both Moscow and New York are incorporated in the same model. For numerical reasons, the volatility specification only contains those lags that were significantly different from zero in the previous step of the estimation procedure. In each step we employ the AIC criterion to find a parsimonious parametrization.

2.2 Estimation

Conditional on $Y_{t-1} = (y_1, \dots, y_{t-1})$ the prediction error

$$e_t = y_t - E(y_t|Y_{t-1})$$

has the distribution of $\varepsilon_t h_t$. BDG assumed $\{\varepsilon_t\}$ to be normally distributed so that the conditional density of y_t given Y_{t-1} is normal with mean $E(y_t|Y_{t-1})$ and variance h_t^2 . The log-likelihood function to be maximized with respect to the unknown parameters is then

$$\ell = \sum_{t=r}^T \ln L_t \propto - \sum_{t=r}^T \frac{1}{2} \left(\ln h_t + \frac{e_t^2}{h_t^2} \right).$$

The log-likelihood function is not continuous in the indicator function. Qian (1998) has derived the asymptotics of the maximum likelihood estimator for the parameters in a general two-regime self-exciting threshold

model in which the errors do not necessarily have a normal density. This particular model is dual to an asMA model which is a special case of (2). The same indicator functions reappear when the model contains conditional heteroskedasticity in the form of (3). Kristensen and Rahbek (2008) give consistency and asymptotic normality results for special cases of the model containing (3).

As an estimator of the covariance matrix, we use the robust sandwich form

$$Cov(\hat{\boldsymbol{\theta}}) = \left(\sum_{t=r}^T \frac{\partial^2 \ln L}{\partial \hat{\boldsymbol{\theta}} \partial \hat{\boldsymbol{\theta}}'} \right)^{-1} \sum_{t=r}^T \frac{\partial \ln L_t}{\partial \hat{\boldsymbol{\theta}}} \frac{\partial \ln L_t}{\partial \hat{\boldsymbol{\theta}}'} \left(\sum_{t=r}^T \frac{\partial^2 \ln L}{\partial \hat{\boldsymbol{\theta}} \partial \hat{\boldsymbol{\theta}}'} \right)^{-1} \quad (4)$$

where $\boldsymbol{\theta}$ is the parameter vector and the expression is evaluated at estimates $\hat{\boldsymbol{\theta}}$.

Hypotheses of symmetric responses in the conditional mean (cf. Brännäs and De Gooijer, 1994), the conditional variance, or in both jointly may be formulated as linear restrictions on the $\boldsymbol{\theta}$ vector, i.e. as $\boldsymbol{\theta}_0 = \mathbf{R}\boldsymbol{\theta} = \mathbf{0}$. Likelihood ratio tests are easy to apply in practice. Given the estimates and the covariance matrix estimator, Wald testing is also straightforward.

The RATS 6.0 package is employed for practical estimation, using robust covariance matrices throughout.

3 Data

The data used in this paper are capitalization-weighted daily stock price indices of the Estonian (Tallinn, TALSE), Latvian (Riga, RIGSE), Lithuanian (Vilnius, VILSE), Russian (Moscow, RTS) and US (NYSE Composite) stock markets.² For more information about the characteristics of Baltic stock markets, see the Appendix. All prices are in local

²All indices were collected from web sites; *www.omxgroup.com* provides the complete description of the Baltic stock market indices, while *www.rts.ru* and *www.nyse.com* give the remaining indices.

currencies, except for Estonia where stock market trading is conducted in Euros. The dataset covers January 3, 2000 to April 29, 2005, for a total of $T = 1391$ observations; see Figure 1 for a presentation and comparison of all indices. It is quite obvious that growth rates are high, except for New York, and that the variance of Moscow is much higher than for other series. The irregularity after the 400th observation in the Riga index (RIGSE) is due to a power struggle in its largest company (Latvijas Gaze) in the summer of 2001. Instead of elaborating on modelling to contain this irregular period, the Riga series starts at September 17, 2001, and contains $T = 945$ daily observations.

Due to differences in holidays for the involved countries, the series have different shares of days for which price indices are not observable. For Baltic stock market indices, the number of missed trading days in comparison with New York, which is the standard that we used, is 39 for TALSE, 49 for RIGSE, and 46 for VILSE for the entire sample. Linear interpolation was used to fill the gaps for all series, resulting in series having a common trading week throughout.

All returns are calculated as $y_t = 100 \cdot \ln(I_t/I_{t-1})$, where I_t is the daily price index. Table 1 reports descriptive statistics for the daily returns. With the exception of New York, the Ljung-Box statistics for 10 lags (LB_{10}) indicate significant serial correlations. The large kurtoses for Riga and Vilnius indicate leptokurtic densities. The returns of Moscow and/or New York serve as the x_t variables in (2). For the z_t of the conditional variance function in (3) we construct two new series for Moscow and New York, by obtaining moving variances for a window length of 10 observations. For Moscow, the sample mean is 4.65 with a variance of 28.83, while for New York, the sample moments are much lower at 1.09 and 1.57. The z_t series that enter the conditional variance function are demeaned moving variance series; the threshold is then set at zero. The z^+ then takes on positive values and is indicative of high risk, and z^- in a corresponding way takes on negative values and indicates lower risk

Table 1: Descriptive statistics for daily return series.

Exchange	Mean	Variance	Min/Max	Skewness	Kurtosis	LB ₁₀	<i>T</i>
Riga	0.10	1.57	-9.72/9.46	-0.04	15.46	57.74	945
Tallinn	0.11	1.13	-5.87/7.34	0.20	5.68	47.80	1391
Vilnius	0.09	0.76	-10.22/4.58	-0.94	16.61	60.80	1391
Moscow	0.10	4.70	-11.57/9.62	-0.42	2.78	21.28	1391
New York	0.00	1.08	-5.27/5.18	0.08	2.29	14.23	1391

Note: LB₁₀ is the Ljung-Box statistic evaluated at 10 lags.

in Moscow and/or New York.

Table 2 gives cross-correlation functions for the return series versus Moscow and New York. There are several interesting features to note. First, Riga appears autonomous with neither Moscow nor New York having any significant influence. Second, Vilnius is throughout influenced positively by both Moscow and New York. Third, for Tallinn, Moscow returns within the same day have a strong positive impact, while for New York, yesterday's returns have the strongest impact followed by the current day. This mirrors the difficulty of synchronization that we face due to differences in time zones. Note that at the time of writing, trading at New York stock exchange starts after the Baltic Stock exchanges have closed, while all three Baltic stock markets and Moscow have periods of market activity that overlap during a day. Koch and Koch (1991) find that most significant market adjustments are completed within a day for countries in the same geographic region, whose trading hours overlap. To reflect this overlap in trading hours, in our model we account for daily interaction between Baltic states and Moscow. To study whether Baltic stock markets adjust to news from Moscow within a day or continue to adjust beyond one day, we allow empirically for Moscow lags within a day as well as beyond one day. Due to the fact that trading hours for New York and Baltic stock markets do not overlap, New

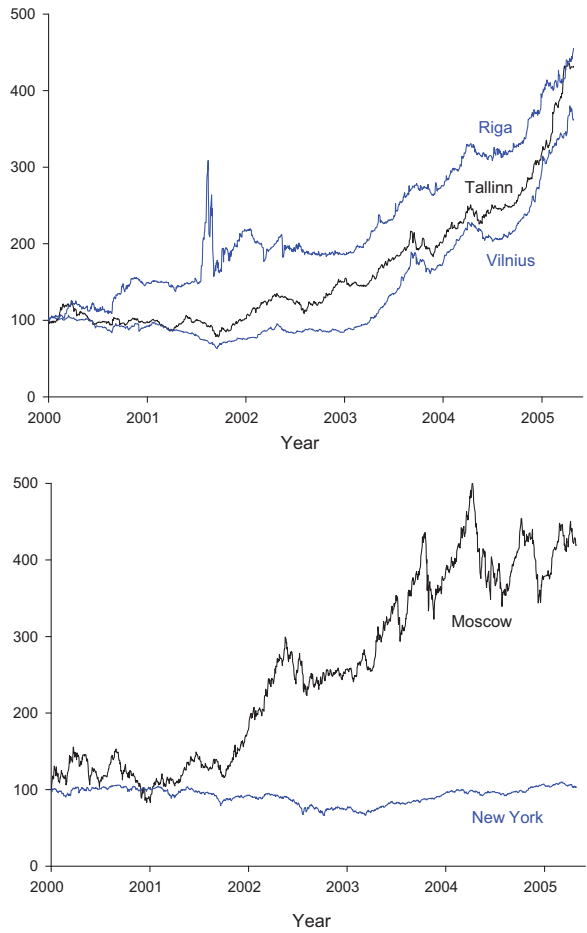


Figure 1: Indices of the Baltic stock exchanges, Moscow and New York (January 3, 2000 = 100).

Table 2: Cross-correlations for Baltic return series vs Moscow and New York.

Exchange	Lag/Lead								
	-4	-3	-2	-1	0	1	2	3	4
	vs Moscow								
Riga	-0.007	.0016	0.048	0.022	0.017	0.015	-0.012	-0.019	-0.016
Tallinn	<u>0.061</u>	<u>0.062</u>	<u>0.055</u>	0.051	<u>0.165</u>	-0.048	-0.038	-0.043	0.008
Vilnius	<u>0.081</u>	<u>0.078</u>	<u>0.059</u>	<u>0.079</u>	<u>0.076</u>	-0.028	-0.043	0.032	0.013
	vs New York								
Riga	0.032	0.004	0.051	0.015	0.020	0.020	-0.029	-0.005	0.034
Tallinn	0.027	0.039	0.033	<u>0.260</u>	<u>0.076</u>	-0.040	-0.026	-0.004	0.025
Vilnius	<u>0.081</u>	<u>0.078</u>	<u>0.059</u>	<u>0.079</u>	<u>0.076</u>	-0.028	-0.043	0.032	0.013

Note: Underlining is used to indicate significant correlations.

York is throughout incorporated with at least one lag to account for the time difference. Finally and not surprisingly given their sizes, the Baltic stock exchanges appear to exert no significant impact on either the Moscow or the New York stock returns, as shown in Table 2.

4 Results

The estimation results, which were obtained by the stepwise procedure outlined in Section 2.1 are presented in Table 3.³ The results are for models with Moscow and New York incorporated jointly.

For Riga, neither good nor bad news arriving from New York and Moscow has any significant impact; hence, such news explains little regarding the returns dynamics. It is interesting to note that negative shocks (i.e. low risk) from Moscow reduce the volatility in Riga, while negative shocks from New York have a risk-increasing effect through a

³Tables with estimation results for the intermediate steps are available on request.

Table 3: Parameter estimates for the joint conditional return and risk functions of ARasMA-asQGARCH models with both Moscow and New York in the functions (robust standard errors in parentheses).

Variable	Riga		Tallinn		Vilnius	
	Return	Risk	Return	Risk	Return	Risk
y_{t-1}			0.123 (0.080)			
y_{t-2}					0.106 (0.026)	
u_{t-1}^+	-0.157 (0.062)		0.139 (0.089)	0.169 (0.106)	0.310 (0.052)	-0.239 (0.118)
u_{t-2}^+				0.087 (0.113)		0.431 (0.100)
u_{t-3}^+			-0.075 (0.030)			
u_{t-1}^-	-0.026 (0.074)	0.153 (0.116)	-0.058 (0.090)	-0.468 (0.107)		0.261 (0.127)
u_{t-2}^-		-0.042 (0.098)		0.236 (0.111)		-0.385 (0.103)
u_{t-3}^-					0.040 (0.048)	
u_{t-4}^-			0.090 (0.038)			
u_{t-1}^2		0.512 (0.065)		0.046 (0.051)		0.295 (0.079)
u_{t-2}^2		-0.202 (0.068)		-0.087 (0.048)		-0.203 (0.067)
h_{t-1}		0.630 (0.048)		0.914 (0.014)		0.673 (0.049)

Table 3 continued

Variable	Riga		Tallinn		Vilnius	
	Return	Risk	Return	Risk	Return	Risk
$x, z_{M,t}^+$	-0.027 (0.031)		0.014 (0.019)		-0.009 (0.018)	
$x, z_{M,t-1}^+$	-0.027 (0.031)		0.000 (0.019)	-0.014 (0.007)	0.025 (0.018)	-0.017 (0.013)
$x, z_{M,t-2}^+$			-0.017 (0.018)	0.013 (0.007)	-0.006 (0.018)	0.091 (0.023)
$x, z_{M,t-3}^+$			0.040 (0.021)			-0.117 (0.021)
$x, z_{M,t-4}^+$						0.048 (0.014)
$x, z_{M,t}^-$	0.053 (0.030)		0.076 (0.017)		0.048 (0.016)	
$x, z_{M,t-1}^-$	-0.010 (0.034)	0.036 (0.009)	-0.028 (0.021)		0.017 (0.017)	0.002 (0.023)
$x, z_{M,t-2}^-$			0.059 (0.019)		0.023 (0.019)	0.007 (0.024)
$x, z_{M,t-3}^-$			0.023 (0.017)			

Table 3 continued

Variable	Riga		Tallinn		Vilnius	
	Return	Risk	Return	Risk	Return	Risk
$x, z_{NY,t-1}^+$	-0.025 (0.050)	0.015 (0.011)	0.202 (0.043)	0.186 (0.044)	0.056 (0.032)	
$x, z_{NY,t-2}^+$						-0.023 (0.031)
$x, z_{NY,t-3}^+$				-0.453 (0.121)		-0.005 (0.072)
$x, z_{NY,t-4}^+$				0.269 (0.083)		0.021 (0.043)
$x, z_{NY,t-1}^-$	0.015 (0.048)	-0.060 (0.022)	0.220 (0.037)		0.071 (0.037)	0.017 (0.093)
$x, z_{NY,t-3}^-$						0.049 (0.233)
$x, z_{NY,t-4}^-$						-0.110 (0.167)
Constant	0.235 (0.059)	0.218 (0.046)	0.149 (0.053)	-0.040 (0.007)	0.071 (0.050)	0.069 (0.030)
$\hat{\sigma}^2$ (AIC)	0.975	(0.89)	0.994	(0.91)	0.712	(0.52)
$\ln L$ (R^2)	-402.15	(0.38)	-598.95	(0.12)	-328.23	(0.06)
LB ₁₀ (LB ₁₀ ²)	11.17	(11.17)	6.87	(0.83)	7.82	(1.91)
Skew (Kurt)	0.62	(4.43)	0.71	(6.88)	-0.35	(6.17)

Note: LB₁₀ and LB₁₀² is the Ljung-Box statistic for standardized residuals and their squares at lag 10.

negative parameter value. The final model explains about 38 percent of the variation in returns.

For Tallinn, good news (i.e. market advances) from New York has a positive impact on returns, while bad news from Moscow and New York has a negative impact, through positive parameter values. The effect of bad news (market declines) from abroad is stronger. It is obvious that shocks arising in New York have larger effects than those of Moscow on the returns of Tallinn, as illustrated in Figure 2. This is consistent with earlier studies that found that the US is the major source of internationally transmitted information for developed markets (e.g., Eun and Shim, 1989; Koch and Koch, 1991; Liu and Pan, 1997). It should be noted that the adjustment to news from Moscow is prolonged. For the conditional volatility function, the results suggest that only positive, i.e. higher risk, shocks from abroad have an effect on volatility in Tallinn, with New York having a stronger influence on volatility than Moscow. About 5 percent of the variation in the conditional volatility is explained by foreign shocks. Focusing on the conditional returns of Tallinn, the estimated model explains 12 percent of the variation in returns, of which about 7 percent is explained by news from Moscow and New York.

The final columns of Table 3 report the estimates for Vilnius. Good news from Moscow has no significant impact on returns in Vilnius, whereas bad news (i.e. market declines) has a negative impact. In addition, the market reaction to bad news from Moscow is quite fast, i.e. within the same day. Moscow and New York explain jointly about 2 percent of the variation in returns in Vilnius, while $R^2 = 0.06$. For volatility spillovers, positive shocks from Moscow are the only ones to affect volatility in Vilnius. The persistence of volatility is quite low for both Riga and Vilnius.

Figures 2-3 illustrate the conditional mean and variance responses to unit positive and negative shocks in Moscow and New York. The returns of Tallinn and Vilnius are affected more by shocks that arise in

New York. High-risk shocks in New York seem to increase the risk in Tallinn, whereas high-risk shocks from Moscow have the strongest effect on the stock market index of Vilnius. It is obvious that shocks from abroad seem to quite a small effect on market returns and risks in Riga. Even though, say, shocks of one standard error in size are employed, the results of Figures 2-3 remain qualitatively unchanged.

We use Wald test statistics to test hypotheses of no asymmetry in the ARasMA-asQGARCH models. The test results are presented in Table 4. In agreement with other studies (e.g., Wecker, 1981; Koutmos, 1998), the conditional mean responses to own past innovations are asymmetric, with the exception of Riga. The same is true of the conditional variance functions. The conditional first moments of the Vilnius and Tallinn stock index returns respond asymmetrically to news from Moscow, where bad news has a stronger impact. There is no evidence for an asymmetric response to either good or bad news from New York for any of the three Baltic stock markets. For the conditional second moments, the Wald test indicates no asymmetric impact of news from New York on volatility in Vilnius and Tallinn. According to our test results, news from Moscow has an asymmetric impact on volatility in all three countries under study.

Table 4: Summary of asymmetry tests (M for Moscow and NY for New York).

Stock Market	Returns			Risk		
	<i>u</i>	NY	M	<i>u</i>	NY	M
Riga	–	–	–	–	*	*
Tallinn	*	–	*	*	*	*
Vilnius	*	–	–	*	–	*

Note: * indicates significance at the 5 percent level by Wald test statistics, cf. Section 2.

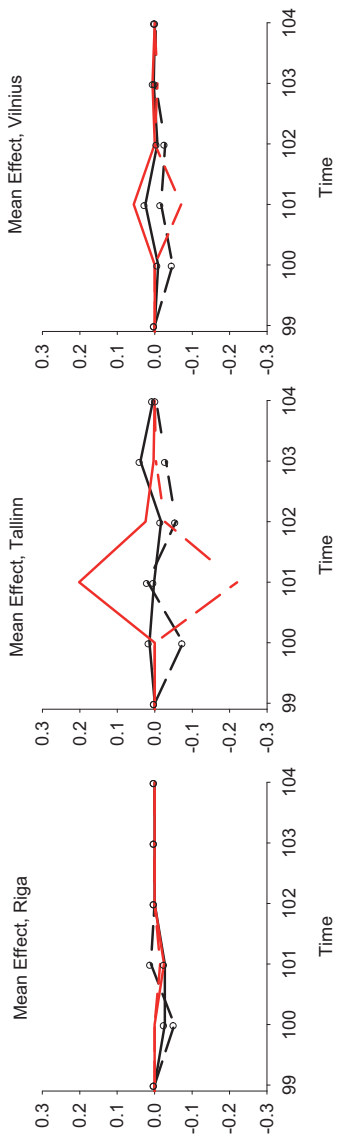


Figure 2: The effect of positive and negative shocks (at date = 100) in Moscow and New York in the conditional return functions of the Riga, Tallinn and Vilnius models. Effects are shifted to zero before the change date and $u_t = 0$ for all t (solid line with marker for positive shocks in Moscow, without marker in New York, and analogously for negative shocks in Moscow, with marker and dashed lines).

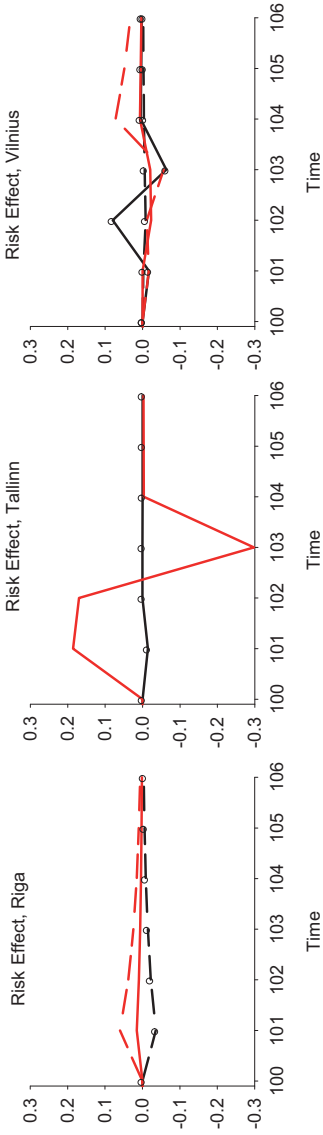


Figure 3: The effect of high- and low-risk shocks (at date = 100) in Moscow and New York in the conditional volatility functions of the Riga, Tallinn and Vilnius models. Effects are shifted to zero before the change date and $u_t = 0$ for all t (solid line with marker for positive shocks in Moscow, without marker in New York, and analogously for negative shocks and dashed lines).

To illustrate the sensitivity to different shocks in portfolio and risk management terms, we calculate the tangency portfolio weights and Value at Risk (VaR) using one-step-ahead forecasts. The results are shown in Table 5. In the base case, the estimated model is used as it is with future values for Moscow and New York set close to their values by the end of the series; $x_{M,T+1}^+ = 0.1$ and $z_{M,T+1}^- = -4$, $x_{NY,T+1}^+ = 1$ and $z_{NY,T+1}^- = -0.1$. Next, we shock the u_T for all three Baltic states (the final residual is individually multiplied by 5), and shock Moscow and New York (multiplied by 5 for shocks to return (x^+) and equally sized, but positive shock to volatility (z^+)). For the risk free rate we use 1.07, which is the level of the Euro market 10-year government bond yield by the end of the sample period. In the base case, about 58 percent of the portfolio should be placed on the Tallinn stock exchange, 25 in Riga, and about 18 in Vilnius. A shock that occurs in the stock market in question reduces the optimal portfolio weight for that market. A shock in Tallinn has the largest implication for portfolio allocation. Portfolio allocation is robust to shocks in Moscow and New York.

The VaR measure for probability 0.025 changes little for shocks in Moscow and New York. Shocks in Riga and Tallinn increase the possible loss, while a shock in Vilnius seem to reduce it. Applying the base case tangency portfolio allocation (VaR-B) reduces the possible losses throughout.

5 Concluding Remarks

We used an extended ARasMA-asQGARCH model to examine the influence of information from Moscow (RTS) and New York (NYSE) on the three Baltic state stock markets of Estonia (Tallinn), Latvia (Riga), and Lithuania (Vilnius). The hypothesis of asymmetric adjustment to a stock market's own past information and information from abroad is tested.

Table 5: Portfolio and VaR effects of shocks in innovations, Moscow and New York. The VaR is based on probability 0.025 and a portfolio with equal weights for each index (A) and with the weights obtained in the Base case (B).

	Portfolio Allocation			VaR	
	Riga	Tallinn	Vilnius	A	B
Base case	0.25	0.58	0.18	0.63	0.45
Shock-Riga	0.14	0.66	0.20	0.85	0.60
-Tallinn	0.41	0.30	0.29	0.71	0.71
-Vilnius	0.26	0.61	0.12	0.58	0.42
-Moscow (x^+)	0.25	0.57	0.18	0.64	0.45
-Moscow (z^+)	0.25	0.58	0.18	0.63	0.45
-New York (x^+)	0.25	0.58	0.18	0.63	0.45
-New York (z^+)	0.25	0.58	0.18	0.63	0.45

We found that news arriving from New York has stronger impacts on market returns in Tallinn and Vilnius than news from Moscow. The returns spillovers from the US to the stock markets in transition is consistent with the result of Tse et al. (2003), who find significant return spillovers from the US market on the Polish market. We found no evidence of asymmetric impact of good and bad news from New York on returns in Baltic states. For Vilnius, we found no volatility spillovers from New York, but we did for Tallinn. Hence, our results both confirm and reject the result of Tse et al. (2003). The Riga stock market seems to be quite autonomous to shocks from abroad. Overall, the stronger influence of New York on the stock markets of Tallinn and Vilnius may be due to the fact that these markets are larger than that of Riga. It is surprising, though, that the Riga stock market is almost independent of shock from abroad, given the fact that foreign investors represent about

46 percent of overall investments, while they represent 41 and 35 percent of investments in Tallinn and Vilnius, respectively.⁴

In addition, the conditional volatility in Tallinn responds asymmetrically to Tallinn's own past innovations, where bad news generates greater volatility. This behaviour is consistent with a partial adjustment price model that states that bad news is incorporated faster into current market prices than good news. One possible explanation for this is that the cost of failing to adjust prices downwards is higher. This result is also compatible with Black's (1976) leverage hypothesis. For Vilnius, we found that positive shocks generate greater volatility. Even though it is unexpected, this behaviour could be explained with the liquidity hypothesis of Rockinger and Urga (2001). They suggest that in an illiquid market, all news generates more liquidity and investors take advantage to dump their positions once greater liquidity has been achieved, which in turn leads to greater volatility. Another explanation for the stronger impact of positive shocks is the possibility that, given the short time series, markets have been anticipating mostly positive shocks.

The overall findings suggest that there are substantial differences among Baltic stock markets, with respect to market adjustment to information arriving from abroad. This supports the findings of Pajuste et al. (2000) that despite common characteristics, emerging markets in Central and Eastern Europe display differences in sensitivity to the risk factors that are affecting the return generating process. This behaviour may be caused by such factors as industry composition, ownership and trade structure.

The study was based on an extended version of the ARasMA-asQGARCH model of Brännäs and De Gooijer (2004). As the results indicate, this model is a suitable and flexible tool for capturing asymmetric and dynamic effects of both international and domestic influences.

⁴At the end of 2005. For more information about the characteristics of the Baltic stock markets, see Appendix A.

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Appendix

Table A: Some basic facts about Baltic stock markets.

Year	2000	2001	2002	2003	2004	2005
Market capitalization (MEUR)						
Riga	673	846	692	900	1207	2122
Tallinn	1923	1666	2316	3005	4627	3022
Vilnius	1481	1038	1409	2767	4753	6937
Market capitalization (% of GDP)						
Riga	7.9	9.1	7.0	9.0	10.8	16.3
Tallinn	31.5	24.1	29.9	34.6	47.9	27.2
Vilnius	12.0	7.6	9.4	16.8	26.2	33.2
Market turnover (MEUR)						
Riga	248	153	127	123	88	77
Tallinn	349	259	263	494	663	1938
Vilnius	124	227	177	166	314	588
Number of companies						
Riga	63	63	62	56	39	45
Tallinn	21	18	15	15	14	16
Vilnius	47	46	51	43	42	43
Average company size (MEUR)						
Riga	10.7	13.4	11.2	16.1	31.0	47.2
Tallinn	91.6	92.5	154.4	200.3	330.5	188.9
Vilnius	31.5	22.6	27.6	64.3	113.2	161.3

Source: Guide to NASDAQ OMX Baltic Securities Market (2009).

II

Simultaneity and Asymmetry of Returns and Volatilities: The Emerging Baltic States' Stock Exchanges*

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Abstract

The paper suggests a nonlinear and multivariate time series model framework that enables the study of simultaneity in returns and in volatilities, as well as asymmetric effects arising from shocks and exogenous variables. The model is employed to study the three closely related Baltic States' stock exchanges and the influence exerted by the Russian stock exchange. Using daily data we find recursive structures with returns in Riga directly depending on returns in Tallinn and Vilnius, and Tallinn on Vilnius. For volatilities both Riga and Vilnius depend on Tallinn. In addition, we find evidence of asymmetric effects of shocks arising in Moscow and in Baltic States on both returns and volatilities.

Key Words: Time series, nonlinear, multivariate, finance.

JEL Classification: C32, C51, G11, G12, G14, G15.

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1 Introduction

The main motivation for this study is the importance of simultaneity in financial assets or markets for various investment and risk management decisions. Portfolio or fund managers, for example, often invest in several markets at the same time. This investment strategy may not provide the diversification and risk reduction that managers are seeking, if there are strong linkages between markets. In addition, risk managers need to understand the nature of cross market linkages in order to appropriately assess their risk exposures and capital adequacy (Fleming et al., 1998).

Cross market linkages or information spillovers are of two types. The first is the common information that simultaneously affects expectations in more than one market. The second type of information spillovers is caused by cross-market hedging. Fleming et al. (1998) argue that information spillovers are strongest when linkages between markets are not limited by institutional constraints, and other practical considerations. These include, for example, a common trading platform, and other factors that lower the settlement risk and information costs for investors. Fazio (2007) argues that investors following an international diversification strategy may be exposed to unhedged risk when assuming that different countries are unrelated. He also finds that countries belonging to the same region are more likely to suffer from dependence in the case of extreme market movements. This implies that countries located in the same region may have stronger linkages than anticipated by investors. Also, Koch and Koch (1991) find simultaneity in returns within geographic regions but not across regions.

Another lesson from the intra-day trading literature concerning some marketplaces is that information processing is very fast (e.g., Engle and Russell, 1998). Even if there are unidirectional causations within the day, a study based on a daily sampling frequency cannot but find an average effect that may go both ways. The sampling frequency scenario

is in fact a main motivation in macro-econometrics for employing structural systems which can incorporate simultaneous endogenous effects. More recently, Rigobon and Sack (2003) and others have reported on model-based studies allowing for simultaneity in returns.

Obviously, and perhaps more interestingly from a risk management point of view, there is also reason to expect simultaneous effects in volatilities. Rigobon and Sack (2003) were the first ones to find simultaneity in volatilities. But, as in the studies of De Wet (2006) and Lee (2006), the simultaneity arises in a very restrictive way, and only as a consequence of the simultaneity in returns. Gannon and Choi (1998) and Gannon (2004, 2005) detect simultaneity for some Asian markets using realized volatilities. Engle and Kroner (1995) suggested a related framework but focus theoretically on simultaneity in returns only.

Here, our main focus will be on the joint modelling of, and the allowance for, simultaneity in both returns and volatilities along with asymmetry, and exogenous effects. The model platform is the univariate ARasMA-asQGARCH of Brännäs and De Gooijer (1994, 2004). This model combines an asymmetric ARMA model with an asymmetric and quadratic GARCH model and it is here given its first multivariate form. Notably, extensions of this type introduce additional parameters into an already richly parameterized model. Kroner and Ng (1998), De Goeij and Marquering (2005) and others discussed ways of parameterizing, in particular, the volatility functions for models to be estimable. To allow for simultaneity we will have to be restrictive in terms of correlation structure, lag lengths, and asymmetric effects. We employ the methodology to jointly study the closely related Baltic States' stock market indices and their potentially asymmetric dependence on the Russian stock market.

The paper is organized as follows. In Section 2 we introduce the model and discuss some of its properties. Section 3 presents the estimator along with the employed stepwise model specification procedure. The

section discusses testing against simultaneous, asymmetric effects, and the impact of exogenous variables. In addition, the use of the model for portfolio allocation and value at risk (VaR) studies are outlined. Section 4 introduces the empirical study and presents the data-set. The empirical findings are given in Section 5. The final section concludes and relates our findings to other studies.

2 A Structural Vector ARasMA-asQGARCH Model

2.1 The Model

Consider an m -dimensional time series $\mathbf{y}_t = (y_{1t}, \dots, y_{mt})'$. In this study $\{\mathbf{y}_t\}$ contains the variables of interest, i.e. the returns at time t of m stock market indices. The vector time series process $\{\mathbf{y}_t\}$ is assumed to be weakly stationary. Let $\mathbf{x}_t = (x_{1t}, \dots, x_{kt})'$ denote a vector of exogenous variables that may affect the process $\{\mathbf{y}_t\}$; see Section 4 for more details on these series. To introduce the asymmetric structure of the proposed model we first need to define an m -dimensional vector discrete-time stochastic process generated by $\mathbf{u}_t = (u_{1t}, \dots, u_{mt})'$ defined by

$$\mathbf{u}_t = \mathbf{H}_t^* \boldsymbol{\varepsilon}_t,$$

where $\{\boldsymbol{\varepsilon}_t\} \sim WN(\mathbf{0}, \mathbf{I})$, $\mathbf{H}_t^* = \{h_{i,j,t}^*\}$ ($i, j = 1, 2, \dots, m$), and with \mathcal{F}_{t-1} denoting the history of the time series up to and including time $t - 1$. The conditional variance is $V(\mathbf{u}_t | \mathcal{F}_{t-1}) = \mathbf{H}_t^* \mathbf{H}_t^{*'} \equiv \mathbf{H}_t$. Now a simultaneous or structural vector ARasMA model can be defined as

$$\begin{aligned} \mathbf{A}_0 \mathbf{y}_t &= \sum_{i=1}^p \mathbf{A}_i \mathbf{y}_{t-i} + \mathbf{u}_t + \sum_{i=1}^q (\mathbf{B}_i^+ \mathbf{u}_{t-i}^+ + \mathbf{B}_i^- \mathbf{u}_{t-i}^-) + \mathbf{c}_0 \\ &+ \sum_{i=0}^r (\mathbf{C}_i^+ \mathbf{x}_{t-i}^+ + \mathbf{C}_i^- \mathbf{x}_{t-i}^-), \end{aligned} \quad (1)$$

where $\mathbf{u}_t^+ = \max(\mathbf{0}, \mathbf{u}_t)$, $\mathbf{u}_t^- = \min(\mathbf{0}, \mathbf{u}_t)$, $\mathbf{x}_t^+ = \max(\mathbf{0}, \mathbf{x}_t)$, and $\mathbf{x}_t^- = \min(\mathbf{0}, \mathbf{x}_t)$. Model (1) accounts for asymmetric effects unless for all i , $\mathbf{B}_i^+ = \mathbf{B}_i^-$ and $\mathbf{C}_i^+ = \mathbf{C}_i^-$. If appropriate, the threshold level for the exogenous process $\{\mathbf{x}_t\}$ may be set at another value than $\mathbf{0}$. It is easy to see that the threshold levels in $\{\mathbf{u}_t^+\}$ and $\{\mathbf{u}_t^-\}$ can be accommodated by the vector of constants \mathbf{c}_0 .

The $m \times m$ non-symmetric matrix \mathbf{A}_0 in (1) contains the simultaneity parameters,

$$\mathbf{A}_0 = \begin{pmatrix} 1 & a_{12}^0 & \cdots & a_{1m}^0 \\ a_{21}^0 & 1 & \cdots & a_{2m}^0 \\ \vdots & \vdots & \ddots & \vdots \\ a_{m1}^0 & a_{m2}^0 & \cdots & 1 \end{pmatrix},$$

where an assumption of normalization has been imposed, i.e. coefficients along the diagonal are equal to 1. Assume \mathbf{A}_0 is nonsingular. Then the conditional mean (return) of $\{\mathbf{y}_t\}$ follows directly from the conditional reduced form of (1) as

$$\begin{aligned} E(\mathbf{y}_t | \mathcal{F}_{t-1}) &= \sum_{i=1}^p \mathbf{A}_0^{-1} \mathbf{A}_i \mathbf{y}_{t-i} + \sum_{i=1}^q \mathbf{A}_0^{-1} (\mathbf{B}_i^+ \mathbf{u}_{t-i}^+ + \mathbf{B}_i^- \mathbf{u}_{t-i}^-) \\ &\quad + \mathbf{A}_0^{-1} \mathbf{c}_0 + \sum_{i=0}^r \mathbf{A}_0^{-1} (\mathbf{C}_i^+ \mathbf{x}_{t-i}^+ + \mathbf{C}_i^- \mathbf{x}_{t-i}^-). \end{aligned}$$

Similarly, the conditional variance (volatility or risk) is given by

$$V(\mathbf{y}_t | \mathcal{F}_{t-1}) = \mathbf{A}_0^{-1} \mathbf{H}_t (\mathbf{A}_0^{-1})'$$

from which, e.g., the conditional correlation matrix can be obtained. Various options are available to specify an asymmetric model for \mathbf{H}_t ; see De Goeij and Marquering (2005). The specifications for \mathbf{H}_t suggested by these authors contain off-diagonal elements. Thus there are conditional and possibly unconditional correlations among the elements of $\{\mathbf{u}_t\}$, and consequently among those of $\{\mathbf{y}_t\}$. There is no simultaneity in conditional volatility behavior in the sense that the conditional variance

of, say, u_{it} would be a direct function of the corresponding conditional variance of u_{jt} ($i \neq j$) in the same time period.

As we wish to have simultaneity in conditional volatility as an integral part of the model we need to consider an extension of the univariate asQGARCH model. One avenue that appears feasible is to view the structures of De Goeij and Marquering (2005) as “reduced forms”. Note that structural forms may make economic sense but that only the reduced form gives the conditional variance interpretation. The situation resembles closely that of the simultaneous and reduced forms in classical macro-econometrics. Similarly, we view simultaneity to arise mainly due to the relatively low sampling frequency of one day while real trading occurs in continuous time, and partly due to common investors on different stock exchanges.

Our general simultaneous specification for the conditional variance is very much in the same spirit as model (1). Given a vector time series process $\{\mathbf{z}_t\}$ of exogenous variables, the vector asQGARCH model for $\mathbf{h}_t = \text{vech}(\mathbf{H}_t)$ is given by

$$\begin{aligned} \mathbf{D}_0 \mathbf{h}_t &= \sum_{i=1}^P \mathbf{D}_i \mathbf{h}_{t-i} + \sum_{i=1}^Q (\mathbf{F}_i^+ \mathbf{u}_{t-i}^+ + \mathbf{F}_i^- \mathbf{u}_{t-i}^-) + \sum_{i=1}^Q \mathbf{K}_i \mathbf{u}_{t-i}^{*,2} \\ &+ \mathbf{g}_0 + \sum_{i=0}^R (\mathbf{G}_i^+ \mathbf{z}_{t-i}^+ + \mathbf{G}_i^- \mathbf{z}_{t-i}^-), \end{aligned} \quad (2)$$

where \mathbf{g}_0 is an $\frac{1}{2}m(m+1) \times 1$ vector of constants, $\mathbf{z}_t^+ = \max(\mathbf{0}, \mathbf{z}_t)$, $\mathbf{z}_t^- = \min(\mathbf{0}, \mathbf{z}_t)$, and the vector $\mathbf{u}_t^{*,2}$ has elements u_{it}^2 ($i = 1, \dots, m$).

The reduced form of (2) is

$$\begin{aligned} \mathbf{h}_t &= \sum_{i=1}^P \mathbf{D}_0^{-1} \mathbf{D}_i \mathbf{h}_{t-i} + \sum_{i=1}^Q \mathbf{D}_0^{-1} (\mathbf{F}_i^+ \mathbf{u}_{t-i}^+ + \mathbf{F}_i^- \mathbf{u}_{t-i}^-) + \sum_{i=1}^Q \mathbf{D}_0^{-1} \mathbf{K}_i \mathbf{u}_{t-i}^{*,2} \\ &+ \mathbf{D}_0^{-1} \mathbf{g}_0 + \sum_{i=0}^R \mathbf{D}_0^{-1} (\mathbf{G}_i^+ \mathbf{z}_{t-i}^+ + \mathbf{G}_i^- \mathbf{z}_{t-i}^-) \end{aligned} \quad (3)$$

from which the corresponding \mathbf{H}_t matrix can be obtained. The matrix

\mathbf{D}_0 captures simultaneity, whereas the matrices \mathbf{D}_i ($i \geq 1$) are useful to represent persistence and possible cyclical features in the process $\{\mathbf{h}_t\}$. Also asymmetric effects are characterized through the matrices \mathbf{F}_i^+ (\mathbf{F}_i^-) and \mathbf{G}_i^+ (\mathbf{G}_i^-). Empirically, it is important to realize that the estimation of (3) may become infeasible with too generously parameterized specifications. Reducing lag lengths and introducing sparse matrix specifications are two ways of reducing the number of parameters; see Section 3 for a data-driven model specification procedure. Note also that the specification in (2) allows for time-varying covariances. Additional simplifications include setting these to constants, by restricting the parameter matrices.

Various moment properties, and distributional results for univariate ARasMA models have been reported by Brännäs and De Gooijer (1994) and Brännäs and Ohlsson (1999), and for univariate ARasMA-quadratic GARCH models by Brännäs and De Gooijer (2004). Since $V(\mathbf{y}_t) = \mathbf{A}_0^{-1} E_{\mathcal{F}_{t-1}}(\mathbf{H}_t)(\mathbf{A}_0^{-1})' + V_{\mathcal{F}_{t-1}}[E(\mathbf{y}_t|\mathcal{F}_{t-1})]$ by a decomposition of the variance, obtaining an explicit expression for the unconditional variance of $\{\mathbf{y}_t\}$ is a far from trivial problem.

3 Estimation and Model Use

Given a multivariate normality assumption on $\{\varepsilon_t\}$ the prediction error

$$\mathbf{y}_t - E(\mathbf{y}_t|\mathcal{F}_{t-1}) = \mathbf{A}_0^{-1}\mathbf{u}_t = \mathbf{A}_0^{-1}\mathbf{H}_t^*\varepsilon_t \equiv \mathbf{v}_t$$

is conditionally $N(\mathbf{0}, \mathbf{\Gamma}_t)$ distributed with $\mathbf{\Gamma}_t = \mathbf{A}_0^{-1}\mathbf{H}_t(\mathbf{A}_0^{-1})'$; recall (3). Here, \mathbf{H}_t is the conditional variance expression in reduced form, containing among other things the \mathbf{D}_0 matrix. Given observations up till time T , the log-likelihood function takes the form

$$\ln L \propto -\frac{1}{2} \sum_{t=s}^T \ln |\mathbf{\Gamma}_t| - \frac{1}{2} \sum_{t=s}^T \mathbf{v}_t' \mathbf{\Gamma}_t^{-1} \mathbf{v}_t$$

$$\propto (T - s) \ln |\mathbf{A}_0| - \frac{1}{2} \sum_{t=s}^T (\ln |\mathbf{H}_t| + \mathbf{u}_t' \mathbf{H}_t^{-1} \mathbf{u}_t),$$

where $s = \max(p, q, r, P, Q, R) + 1$. For practical quasi maximum likelihood estimation we use the RATS 6.0 package and employ robust standard errors.

To obtain the final model specification we advocate the following stepwise procedure.

1. Restrict all matrices in the mean and variance equations to be diagonal and select the model that minimizes AIC or some other appropriate model selection criterion. In this step we implicitly assume that there are no interactions between the series. It is equivalent to finding the "best" univariate ARasMA-asQGARCH models.
2. Take the model from step 1 and expand to non-diagonal matrices in the mean equation. First allow for simultaneity, i.e. estimate \mathbf{A}_0 . Consider thereafter the expansion of the remaining matrices. Choose the specification that minimizes AIC. The \mathbf{A}_0 is the final parameter matrix to be reduced. The volatility functions obtained in step 1 are taken as given, but $\{\hat{\mathbf{u}}_t\}$ changes in the iterative steps.
3. Take the $\{\hat{\mathbf{u}}_t\}$ -sequence from step 2 as given and expand to non-diagonal matrices in the variance equation. First allow for simultaneity, i.e. estimate \mathbf{D}_0 . Consider thereafter the expansion of the remaining matrices. Choose the specification that minimizes AIC. The \mathbf{D}_0 is the final parameter matrix to be reduced.
4. In a final step all parameters are estimated jointly.

Given the estimated model, it is of interest to test hypotheses about simultaneity and asymmetric effects in the \mathbf{x} and \mathbf{z} variables. Given the

likelihood framework and our specification procedure, Wald and likelihood ratio (LR) test statistics are relatively easy to implement.

We first consider tests of simultaneity and do so in terms of the \mathbf{A}_0 matrix. The reasoning with respect to \mathbf{D}_0 is analogous. We say that there is a simultaneous effect between markets i and j if $(\mathbf{A}_0)_{ij} \neq 0$ and $(\mathbf{A}_0)_{ji} \neq 0$. When $(\mathbf{A}_0)_{ij} \neq 0$ but $(\mathbf{A}_0)_{ji} = 0$ there is a recursive structure and causation is unidirectional from market j to market i . When $(\mathbf{A}_0)_{ij} = (\mathbf{A}_0)_{ji} = 0$ there is no causation between returns. When all off-diagonal elements equal zero $\mathbf{A}_0 = \mathbf{I}$ and the structural and reduced forms are identical.

Next we consider testing against asymmetric effects and do so in terms of the \mathbf{B}_i^+ and \mathbf{B}_i^- matrices. We may form $\mathbf{B}_i^\nabla = \mathbf{B}_i^+ - \mathbf{B}_i^-$ ($i = 1, \dots, q$), and test whether this matrix is equal to zero or whether it is nonzero. We then make no distinction between the case of both matrices having nonzero parameters $(\mathbf{B}_i^+)_{ij}$ and $(\mathbf{B}_i^-)_{ij}$ in all places and the case where, say, $(\mathbf{B}_i^-)_{ij} = 0$. Testing against asymmetric effects of exogenous variables is in terms of the parameter matrices \mathbf{C}_i^+ and \mathbf{C}_i^- ($i = 1, \dots, r$). For asymmetric effects in volatility the parameter matrices \mathbf{F}_i^+ and \mathbf{F}_i^- as well as \mathbf{G}_i^+ and \mathbf{G}_i^- are focused.

For no effects of exogenous variables on returns all matrices \mathbf{C}_i^+ and \mathbf{C}_i^- must be identical to a zero matrix, while for volatility all \mathbf{G}_i^+ and \mathbf{G}_i^- must be zero.

When we wish to use or, as here, evaluate the model in financially interesting and meaningful ways, portfolio allocation and VaR measures are of obvious interest. Two problems both stemming from the use of index series arise; how to get back to the index and what price related to the index should we consider.

First, the index is determined from the inverse of the change variable $y_{it} = 100 \cdot \ln(I_{it}/I_{it-1})$, i.e. as $I_{it} = I_{it-1} \exp(y_{it}/100)$ for stock market i . We get $E(I_{it}|\mathcal{F}_{t-1}) = I_{it-1}E(\exp(y_{it}/100)|\mathcal{F}_{t-1}) \approx I_{it-1}(1 + E(y_{it}|\mathcal{F}_{t-1})/100)$ where the first order approximation of the exponential

function is reasonable for the small values of $y_{it}/100$. Using the same first order approximation we get $V(\mathbf{I}_t|\mathcal{F}_{t-1}) = \mathbf{I}_{t-1}^\circ V(\mathbf{y}_t|\mathcal{F}_{t-1})\mathbf{I}_{t-1}^\circ/100^2$, where \mathbf{I}_t° is a matrix with elements I_{it} on the diagonal and zeroes elsewhere. These expressions are useful if we wish to forecast the index and to give its forecast variance. Second, trading is not directly in terms of the index. The presence of index funds and standard options tied to the index are reasonable justifications for using the index as a price. The chosen approach is to use the return series as is and then emphasize the return as an indicator of market risk (e.g., McNeil and Frey, 2000).

For portfolio allocation we adopt the tangency portfolio (e.g., Campbell et al., 1997, ch 5). At time $T + 1$ we have

$$\mathbf{a}_{T+1} = V^{-1}(\mathbf{y}_{T+1}|\mathcal{F}_T) \cdot [E(\mathbf{y}_{T+1}|\mathcal{F}_T) - R_f\mathbf{1}] / A,$$

where $A = \mathbf{1}'V^{-1}(\mathbf{y}_{T+1}|\mathcal{F}_T) \cdot [E(\mathbf{y}_{T+1}|\mathcal{F}_T) - R_f\mathbf{1}]$, R_f is the risk free rate, and $\mathbf{1}$ is a column vector of ones. Hence, $\mathbf{1}'\mathbf{a}_{T+1} = 1$. For the VaR-measure under normality, a time invariant allocation vector \mathbf{a} , and a probability α , Gouriéroux and Jasiak (2001, ch 16) give:

$$R_{T+1} = -\mathbf{a}'E(\mathbf{y}_{T+1}|\mathcal{F}_T) + \Phi^{-1}(1 - \alpha) [\mathbf{a}'V(\mathbf{y}_{T+1}|\mathcal{F}_T)\mathbf{a}]^{1/2},$$

where $\Phi(\cdot)$ is the standard normal distribution function. This VaR measure is in terms of returns; one in terms of indices can also be devised by simply replacing \mathbf{y}_{T+1} by \mathbf{I}_{T+1} and using the expressions given above. Using shock scenarios in terms of the \mathbf{u}_t vector or in terms of $\mathbf{x}_t^{+/-}$ and $\mathbf{z}_t^{+/-}$, the \mathbf{a}_{T+1} and R_{T+1} can be calculated and then evaluated and subjected to comparisons. To cast light on effects of simultaneity, the univariate models can be compared to the simultaneous model system in terms of the portfolio or VaR metrics either as above or over some historical period. Note, that both measures are subject to sampling variation in estimated mean return and risk functions. Britten-Jones (1999) and others have discussed the variation in allocation weights, while Christoffersen and Gonçalves (2005) among others have discussed the issue for VaR measures.

4 Empirical Study and Data

The framework described above is used to study the indices of the Baltic States' stock exchanges. There are several common features, by which the Baltic States' stock exchange indices are likely to move together simultaneously. First, these relatively small marketplaces are geographically closely located. Second, they have the same owner (Nasdaq-OMX) and share a common trading platform. In addition, many of the largest traders are common to all three marketplaces. In fact, foreign institutional investors, predominately European ones, represent 40-47 percent of the market value in the Baltic States' stock markets, whereas foreign and domestic institutional investors combined control about 90 percent of the market value.

To study the joint evolution of returns and volatilities in the Baltic States' stock markets, we use capitalization weighted daily stock price indices of the Estonian (Tallinn, TALSE), Latvian (Riga, RIGSE), Lithuanian (Vilnius, VILSE) and Russian (Moscow, RTS) stock markets. All prices are transformed into Euros from local currencies, except for Estonia where stock market trading is in Euro. Using a common currency implies that the analyzed return series also contain variation due to exchange rate movements. Hence, this paper takes an international investor perspective, when interest is in Euro returns, and the effect of these variations is therefore included in the analysis. Also, since Estonia and Lithuania joined the ERM II during 2004 and Latvia in 2005 the exchange rates have been rather stable during, at least, the later parts of the return series.¹ The data-set covers January 3, 2000 to August 16,

¹The Latvian currency was pegged to the SDR basket (the unit of accounting of the IMF) consisting of the four major currencies: the US dollar, the Euro, the British pound sterling and the Japanese yen, since February 1994. The fixed exchange rate with the Euro was implemented on January 1, 2005. The Estonian currency was pegged to the Deutsche Mark since 1992, and moved to the Euro peg after the introduction of the Euro. Lithuania introduced a US dollar-based currency board in

2006, for a total of $T = 1729$ observations, cf. Figure 1 for the three Baltic States' indices. Both indices and exchange rates are collected from DataStream. The irregularity in the summer of 2001 in the Riga index (RIGSE) is due to a power struggle in its largest company (Latvijas Gaze). Instead of elaborating on modelling to contain this irregular period, the Riga series is adjusted in the following simplistic way: For a speculation period from July 25 to September 3, 2001, observations are replaced by interpolated values.

Following Brännäs and Soultanaeva (2006), the Russian stock market index is used as a exogenous variable that may have an impact on the Baltic States' stock markets. In general, spillovers from the Russian stock market can be explained by economic, historical and political ties between the countries (e.g., Koch and Koch, 1991). Pajuste et al. (2000) find, for example, that East European countries are likely to be affected by news coming from Russia. Moreover, Brännäs and Soultanaeva (2006) demonstrated that good and bad news arriving from Russia (Moscow) have asymmetric impacts on the volatility in the Baltic States' stock markets. Thus, within the context of the empirical analysis, the time series processes $\{\mathbf{x}_t^+\}$ and $\{\mathbf{x}_t^-\}$ represent positive and negative returns at time t in the Russian Stock Exchange (RTS) index, whereas the series $\{\mathbf{z}_t\}$ will enter (2) as the demeaned moving variance series of the RTS index. In more detail, to obtain the z_t variable, we construct a new series by obtaining moving variances of Moscow returns for a window length of 10 observations and deduct the mean. The z^+ then takes on positive values and is indicative of high-risk, and z^- in a corresponding way takes on negative values and indicates a lower risk in Moscow.

Due to some differences in holidays for the involved countries the series have different shares of days for which index stock price are not observable. Linear interpolation was used to fill the gaps for all series.

1994 and changed the peg to the Euro in February 2002.

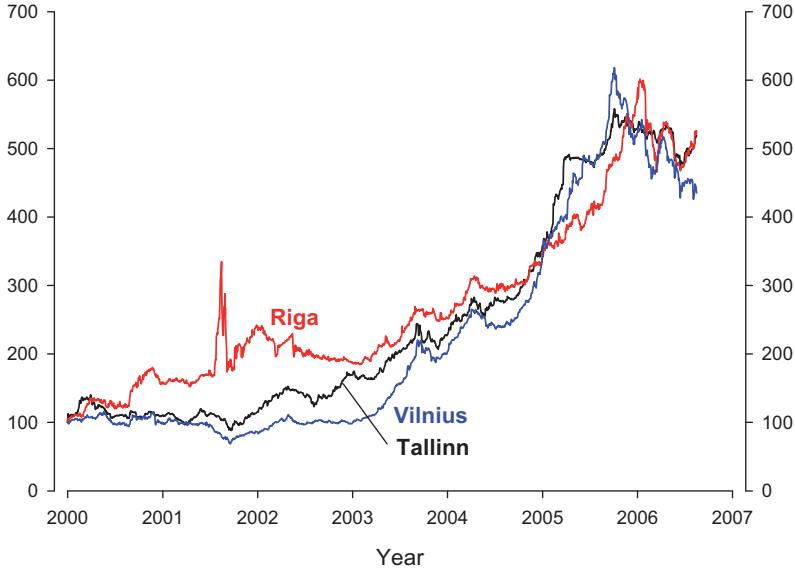


Figure 1: Indices of the Baltic stock exchanges (December 31, 1999 = 100).

The resulting series are then throughout for a common trading week. All returns are calculated as $y_t = 100 \cdot \ln(I_t/I_{t-1})$, where I_t is the daily price index. Table 1 reports descriptive statistics for the daily returns. The Ljung-Box statistics for 10 lags (LB_{10}) indicate significant serial correlations. The large kurtoses for Riga, Tallinn and Vilnius indicate leptokurtic densities. Table 2 presents cross correlations for the Baltic States' return series and for a squared returns. Table 3 gives auto and lagged cross correlations. For instance, the table indicates that Tallinn is positively affected by Vilnius both within the day and with up to three lags. There appears to be no impact from Riga.

Figure 2 gives scatterplots for pairs of returns series with a nonparametric regression line (LOWESS default settings in RATS 6.0). Visual

Table 1: Descriptive statistics for return series.

Exchange	Mean	Variance	Min/Max	Skewness	Kurtosis	LB ₁₀
Riga	0.10	1.77	-9.27/10.29	0.18	10.72	45.93
Tallinn	0.10	1.05	-5.87/12.02	1.09	15.94	51.43
Vilnius	0.09	1.05	-12.12/5.32	-0.91	13.82	46.87
Moscow	0.12	4.93	-11.92/10.23	-0.47	3.27	16.37

Note: LB₁₀ is the Ljung-Box statistic evaluated at 10 lags.

Table 2: Cross correlations for Baltic stock markets returns and squared returns.

	Returns			Squared Returns		
	Riga	Tallinn	Vilnius	Riga	Tallinn	Vilnius
Riga	1			1		
Tallinn	0.134	1		0.161	1	
Vilnius	0.141	0.208	1	0.023	0.032	1

inspection indicates that there is weak dependence between Riga and Tallinn for the majority of observations, while for the other plots there appear to be positive relationships.

5 Results

The empirical results are presented first in terms of the return function and later in terms of the volatility function. Table A in the Appendix contains estimated univariate models. The empirical specifications are obtained by the steps outlined in Section 3.

For the return function of $\{\mathbf{y}_t\}$, cf. eq (1), when returns are in the order Riga, Tallinn and Vilnius, the estimated function is

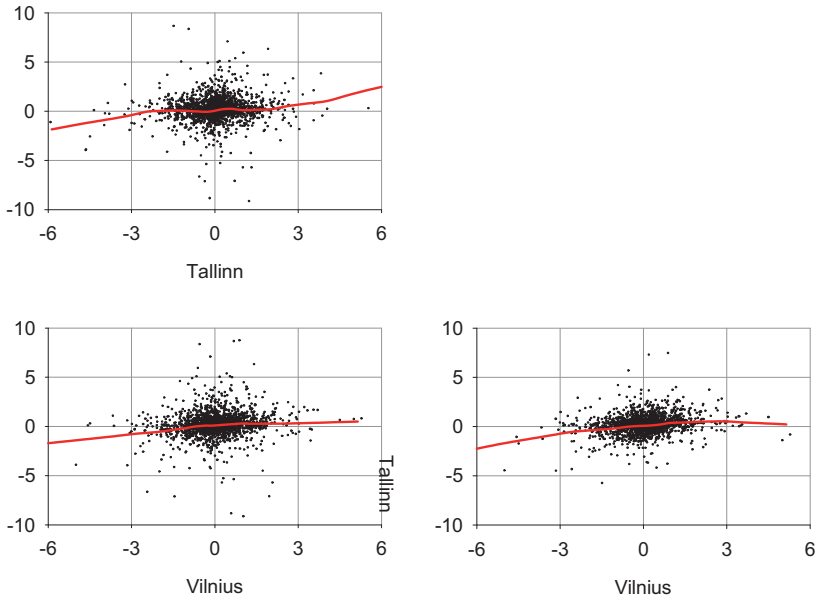


Figure 2: Cross plots for Baltic returns series. One negative outlier for Vilnius is outside the figure and three positive ones for Tallinn.

Table 3: Auto and cross correlations for Baltic stock markets returns (in the order Riga, Tallinn and Vilnius). Significant entries are indicated by signs and subindex indicates lag.

$$\begin{pmatrix} 1 & + & + \\ \cdot & 1 & + \\ \cdot & + & 1 \end{pmatrix}_0, \begin{pmatrix} - & \cdot & \cdot \\ \cdot & \cdot & + \\ \cdot & + & + \end{pmatrix}_1, \begin{pmatrix} \cdot & \cdot & + \\ \cdot & \cdot & + \\ \cdot & \cdot & + \end{pmatrix}_2, \begin{pmatrix} + & \cdot & \cdot \\ \cdot & \cdot & + \\ \cdot & \cdot & \cdot \end{pmatrix}_3, \\
 \begin{pmatrix} - & \cdot & \cdot \\ \cdot & \cdot & \cdot \\ \cdot & \cdot & + \end{pmatrix}_4$$

$$\begin{aligned}
& \begin{pmatrix} 1 & -0.06 & -0.09 \\ & (0.034) & (0.043) \\ 0 & 1 & -0.11 \\ & & (0.025) \\ 0 & 0 & 1 \end{pmatrix} \hat{\mathbf{y}}_t = \begin{pmatrix} 0 & 0 & 0 \\ 0 & 0 & 0 \\ 0 & 0 & 0.06 \\ & & (0.023) \end{pmatrix} \hat{\mathbf{y}}_{t-2} \\
& + \begin{pmatrix} -0.17 & 0 & 0 \\ & (0.050) & \\ 0 & 0.24 & 0 \\ & (0.044) & \\ 0 & 0 & 0.15 \\ & & (0.048) \end{pmatrix} \hat{\mathbf{u}}_{t-1}^+ + \begin{pmatrix} 0 & 0 & 0 \\ 0 & 0.12 & 0 \\ & (0.039) & \\ 0 & 0 & 0 \end{pmatrix} \hat{\mathbf{u}}_{t-2}^+ \\
& + \begin{pmatrix} 0 & 0 & 0 \\ 0.07 & 0.09 & 0.07 \\ & (0.023) & (0.048) & (0.026) \\ 0 & 0 & 0 \end{pmatrix} \hat{\mathbf{u}}_{t-1}^- + \begin{pmatrix} 0 & 0.12 & 0.08 \\ & (0.046) & (0.041) \\ 0 & 0 & 0 \\ 0 & 0 & 0 \end{pmatrix} \hat{\mathbf{u}}_{t-2}^- \\
& + \begin{pmatrix} 0.21 \\ & (0.043) \\ 0.12 \\ & (0.026) \\ 0.14 \\ & (0.027) \end{pmatrix} + \begin{pmatrix} 0.06 \\ & (0.021) \\ 0 \\ 0 \end{pmatrix} x_t^+ + \begin{pmatrix} 0 \\ 0 \\ 0.04 \\ & (0.018) \end{pmatrix} x_{t-1}^+ \\
& + \begin{pmatrix} 0.07 \\ & (0.024) \\ 0.09 \\ & (0.011) \\ 0.12 \\ & (0.016) \end{pmatrix} x_t^- + \begin{pmatrix} 0 \\ 0.03 \\ & (0.012) \\ 0 \end{pmatrix} x_{t-1}^- + \begin{pmatrix} 0 \\ 0.02 \\ & (0.014) \\ 0 \end{pmatrix} x_{t-2}^-.
\end{aligned}$$

With respect to simultaneity, the $\hat{\mathbf{A}}_0$ matrix indicates a recursive structure; the returns of the Riga index depends within the day positively on both the index returns of Tallinn and Vilnius, while returns in Tallinn are positively influenced by those of Vilnius. Riga returns have no impact on the returns of neither Tallinn nor Vilnius, and Tallinn returns have no influence on those of Vilnius. The only lagged influence arises for Vilnius at lag two, cf. the $\hat{\mathbf{A}}_2$ matrix.

For Riga returns, Moscow has a quite symmetric and positive effect within the day. For Tallinn, we instead find asymmetric effects spread over lags 0 – 2. The effects are much stronger for a negative shock. For

Vilnius, negative shocks out of Moscow appear to have stronger impact than positive shocks. For shocks arising in the three Baltic States' stock exchanges we find that a positive shock in Riga at lag one has a negative impact on current returns, and in addition, negative lag two shocks of Tallinn and Vilnius have negative effects. Positive shocks in Tallinn have stronger effects than equally sized negative shocks, and there are negative shocks of both Riga and Vilnius at lag 2. The off-diagonal elements of lagged shocks suggests that there are some shock-spillovers; Riga returns are negatively influenced by Tallinn and Vilnius shocks at lag two, while Tallinn is impacted by Riga and Vilnius shocks at lag one.

For the volatility function the conditional covariances are assumed time-invariant and insignificantly estimated as $\hat{\mathbf{h}}_{t,4} = 0.003$ (s.e. = 0.023), $\hat{\mathbf{h}}_{t,5} = 0.000$ (0.033) and $\hat{\mathbf{h}}_{t,6} = 0.000$ (0.025). The estimated conditional variances has the form

$$\begin{aligned} & \begin{pmatrix} 1 & -0.01 & 0 \\ & (0.004) & \\ 0 & 1 & 0 \\ 0 & 0.03 & 1 \\ & (0.016) & \end{pmatrix} \hat{\mathbf{h}}_t = \begin{pmatrix} 0.95 & 0 & 0 \\ & (0.006) & \\ 0 & 0.93 & 0 \\ & (0.009) & \\ 0 & 0 & 0.82 \\ & & (0.029) \end{pmatrix} \hat{\mathbf{h}}_{t-1} \\ & + \begin{pmatrix} -0.02 & 0 & 0 \\ & (0.018) & \\ 0 & 0 & 0 \\ 0 & -0.12 & 0.27 \\ & (0.025) & (0.035) \end{pmatrix} \hat{\mathbf{u}}_{t-1}^+ + \begin{pmatrix} 0 & 0 & 0 \\ 0 & 0.15 & 0 \\ & (0.016) & \\ 0 & 0 & 0 \end{pmatrix} \hat{\mathbf{u}}_{t-2}^+ \\ & + \begin{pmatrix} 0.37 & 0 & 0 \\ & (0.076) & \\ 0 & -0.15 & 0 \\ & (0.017) & \\ 0 & 0 & -0.26 \\ & & (0.033) \end{pmatrix} \hat{\mathbf{u}}_{t-1}^- + \begin{pmatrix} -0.29 & 0 & -0.06 \\ & (0.073) & (0.008) \\ 0 & 0 & 0 \\ 0 & 0 & 0 \end{pmatrix} \hat{\mathbf{u}}_{t-2}^- \end{aligned}$$

$$\begin{aligned}
& + \begin{pmatrix} 0.36 & 0 & 0 \\ (0.039) & & \\ 0 & 0.13 & 0 \\ & (0.037) & \\ 0 & 0.03 & 0.12 \\ & (0.007) & (0.033) \end{pmatrix} \hat{\mathbf{u}}_{t-1}^{*,2} + \begin{pmatrix} -0.31 & 0 & 0 \\ (0.037) & & \\ 0 & -0.14 & -0.01 \\ & (0.034) & (0.001) \\ 0 & 0 & -0.13 \\ & & (0.028) \end{pmatrix} \hat{\mathbf{u}}_{t-2}^{*,2} \\
& + \begin{pmatrix} 0.02 \\ (0.011) \\ -0.02 \\ (0.004) \\ 0.07 \\ (0.026) \end{pmatrix} + \begin{pmatrix} -0.002 \\ (0.000) \\ 0 \\ 0.04 \\ (0.006) \end{pmatrix} z_t^+ + \begin{pmatrix} 0.08 \\ (0.017) \\ 0.04 \\ (0.012) \\ 0 \end{pmatrix} z_t^- \\
& + \begin{pmatrix} 0 \\ 0 \\ -0.04 \\ (0.006) \end{pmatrix} z_{t-1}^+ + \begin{pmatrix} -0.08 \\ (0.016) \\ -0.04 \\ (0.012) \\ 0.01 \\ (0.004) \end{pmatrix} z_{t-1}^-.
\end{aligned}$$

Only two elements in $\hat{\mathbf{D}}_0$ are significant, the volatility of Vilnius depends negatively but weakly on that of Tallinn in the same time period, while Riga depends positively on Tallinn. As expected volatilities are quite persistent, cf. the $\hat{\mathbf{D}}_1$ -matrix estimates. The patterns for Riga and Tallinn are quite similar and asymmetric; a higher than average Moscow risk marginally reduces risk in Riga and there is no effect for Tallinn, and in both cases there is a strong negative direct effect of a lower than average Moscow risk that turns positive and then dies out. For Vilnius the direct effects are quite asymmetric and both are positive. Thereafter the effects are negative and gradually die out. The effect is an enhancing one for Vilnius.

The model evaluation phase considers formal tests against simultaneity in returns and in risk as well as tests against asymmetric effects arising from Moscow or from the innovations of the model system. As a first but informal test supporting the joint models rests on the likelihoods under the univariate models and the joint model; the likelihood ratio statistic is then $LR = 181.8$. Table 4 summarizes the Wald test results and also gives the serial correlation properties and the goodness-

Table 4: Simultaneity and asymmetry tests together with model evaluation measures.

Hypothesis	Wald	df	Measure	Riga	Tallinn	Vilnius
Simultaneity>Returns	27.0	3	LB ₁₀	10.08	5.82	22.75
Simultaneity-Risk	7.81	2	LB ₁₀ ²	11.77	1.63	1.14
Asymmetry-Return-Moscow	160.9	6	Skewness	0.47	0.54	-0.30
Asymmetry-Return-Innovation	74.4	8	Kurtosis	4.33	6.31	6.06
Asymmetry-Risk-Moscow	92.8	6	JB	1403.7	2936.8	2659.2
Asymmetry-Risk-Innovation	6033	7	R^2	0.05	0.18	0.06

of-fit for the model. The Wald tests are all significant with p -values less than 0.02. There is then evidence of simultaneity as well as of asymmetric effects. When it comes to serial correlation properties in standardized and squared standardized residuals there appears to be remaining serial correlation in only one series, the standardized residuals of Vilnius. The standardized residuals are nonnormal and leptokurtic.

Next, we consider the estimated volatility functions in some more detail in Figures 3-4. Figure 3 shows the estimated $\mathbf{H}_{t,i,i}$ functions for the final part of the series. It is quite clear from this figure that the volatilities of Riga and Vilnius are larger than those of Tallinn. This pattern reinforces the sample variance ordering of Table 1. The estimated volatility functions are positively correlated, cf. Figure 4. Since covariance estimates $\mathbf{H}_{t,i,j}$ between the innovations of stock exchanges are very small the resulting time-varying conditional correlations are also very small and always smaller than 0.05. The implied estimated conditional correlations between $\{\mathbf{y}_t\}$ variables are much larger and also positive throughout, cf. Figure 5. Average conditional correlations are relatively close to the sample correlations of Table 2.

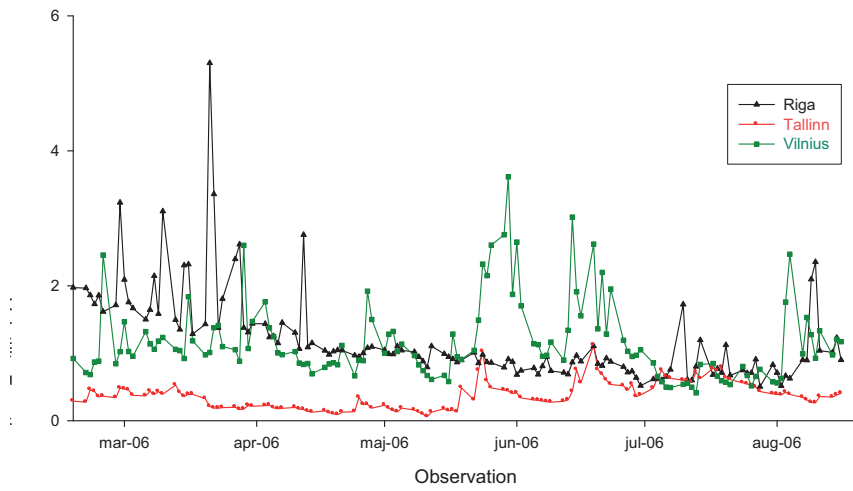


Figure 3: Estimated volatility functions for the final part of the sample period.

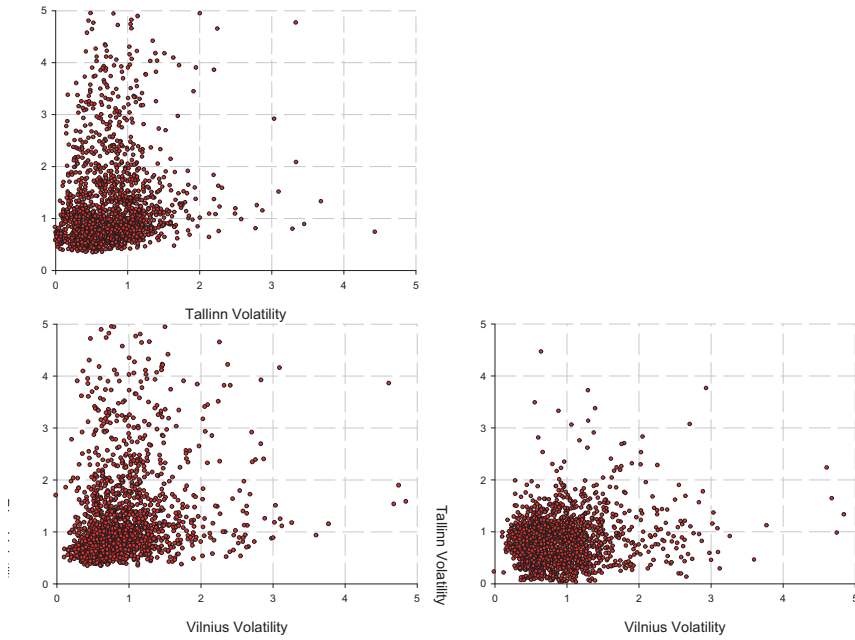


Figure 4: Plots of estimated volatilities (some outlying volatilities fall outside the graphs).

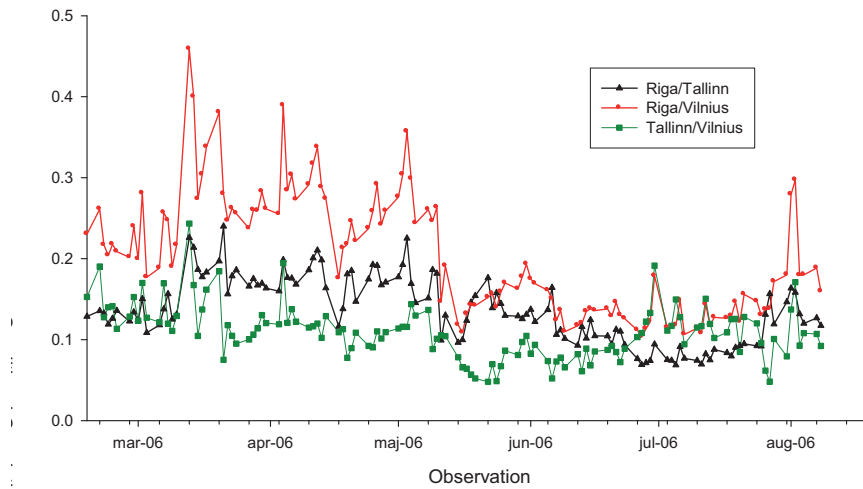


Figure 5: Estimated conditional correlations between the returns of the stock markets for the final part of the sample period.

Table 5: Portfolio and VaR effects of shocks in innovations and Moscow (Joint), together with a univariate model (Single) case. The VaR is based on probability 0.025 and a portfolio with weights 0.333 for each index (VaR-A) and with the weights obtained in the Base case (VaR-B).

	Portfolio Allocation						VaR			
	Joint			Single			A		B	
	Riga	Tallinn	Vilnius	Riga	Tallinn	Vilnius	Joint	Single	Joint	Single
Base case	0.24	0.66	0.10	0.32	0.50	0.18	1.23	0.91	1.66	0.83
Shock-Riga	0.27	0.64	0.09	0.19	0.60	0.21	1.19	1.15	1.65	0.98
-Tallinn	0.30	0.54	0.16	0.35	0.45	0.19	1.23	0.99	1.64	1.06
-Vilnius	0.26	0.72	0.02	0.37	0.58	0.05	1.42	0.99	2.02	0.81
-Moscow (x)	0.23	0.67	0.10	0.31	0.51	0.18	1.25	0.90	1.67	0.82
-Moscow (z)	0.24	0.62	0.14	0.27	0.50	0.23	1.36	1.07	1.77	1.03

Portfolio allocations and VaR measures one-step-ahead are shown in Table 5. These measures are based on forecast equations

$$\begin{aligned}
 E(\mathbf{y}_{T+1}|\mathcal{F}_T) &= \hat{\mathbf{A}}_0^{-1} \left[\hat{\mathbf{A}}_2 \mathbf{y}_{T-1} + \sum_{i=1}^2 \left(\hat{\mathbf{B}}_i^+ \hat{\mathbf{u}}_{T+1-i}^+ + \hat{\mathbf{B}}_i^- \hat{\mathbf{u}}_{T+1-i}^- \right) \right. \\
 &\quad \left. + \hat{\mathbf{c}}_0 + \sum_{i=0}^2 \left(\hat{\mathbf{C}}_i^+ \mathbf{x}_{T-i}^+ + \hat{\mathbf{C}}_i^- \mathbf{x}_{T-i}^- \right) \right] \\
 V(\mathbf{y}_{T+1}|\mathcal{F}_T) &= \hat{\mathbf{A}}_0^{-1} \hat{\mathbf{H}}_{T+1} (\hat{\mathbf{A}}_0^{-1})'
 \end{aligned}$$

and depend on the histories of \mathbf{y}_t , $\hat{\mathbf{u}}_t$, and \mathbf{x}_t for the conditional return and additionally on the histories of \mathbf{H}_t and \mathbf{z}_t for the conditional volatility. Since the impact of Moscow is in the same period we set future values (x_{T+1} and z_{T+1}) for Moscow close to their values at the end of the series, i.e. as $x_{T+1}^+ = 0.1$ and $z_{T+1}^- = -4$. This is the Base case design. For the portfolio allocation exercise the risk free rate is set at 1.07, which is the level of the Euro market government bond yield by the end of the sample period.

The allocation for the Tallinn stock exchange is 0.66, while 0.24 of the portfolio should be placed in Riga and 0.10 in Vilnius. Using the same setup but using instead the univariate models (Single) of Table A, gives a much lower allocation for Tallinn and higher ones for both Riga and Vilnius. The two model forms differ in simultaneity but also with respect to other features of the dynamic model. Therefore, we cannot infer with certainty that the differences are due solely to simultaneous effects. The VaR measures for probability 0.025 are for the simultaneous model with equal weights 1.23 and for the univariate models 0.91. For the weights obtained with the weights of the Base case we get 1.66 and 0.82, respectively.

To study the sensitivity of the Base case results we next shock the individual elements of $\hat{\mathbf{u}}_T$ (the final residuals are individually multiplied by a factor 3). Note that the underlying sizes of residuals in the univariate models have not been changed but shocks are throughout in the

direction of the joint model. For shocks in the Tallinn and Vilnius stock markets the allocations for these markets are reduced. Figure 6 illustrates this for an increasingly negative shock in Tallinn. With a decrease in the Tallinn weight comes relatively more weight for Riga than for Vilnius. The allocations obtained using the univariate models differ from those based on the joint model, mainly such that the weights for Riga and Vilnius are larger and those for Tallinn are smaller.

We also consider shocks arising in Moscow returns (x_{T+1}^+ is set to 1). This appears to have only minor impact. For Moscow risk we change from $z_{T+1}^- = -4$ to $z_{T+1}^+ = 4$ and note an increase for Vilnius and a reduction for Tallinn allocations.

The VaR measure changes little for shocks in Tallinn but responds more to shocks in Vilnius and in Moscow risk. The VaR:s based on the univariate models are smaller than the corresponding measures for the joint model. When the weights of the Base case are used the VaR:s increase markedly throughout. Figure 6 studies the impacts on VaR of Moscow shocks in more detail. Changes in risk have rather small effects, while Moscow return changes have a more sizeable and asymmetric effect.

6 Conclusion

The paper has introduced simultaneity into a multivariate and non-linear time series model framework to study jointly the indices of the Baltic States' stock exchanges. Unlike previous studies (e.g., Rigobon and Sack, 2003, De Wet, 2006, Lee, 2006), we allow for simultaneity in returns and volatility separately. The model allows us to capture "within a day" information transmission between the stock markets under study. Since information transmission between markets is virtually instantaneous (e.g., Engle and Russell, 1998) a study based on daily sampling frequency should take into account simultaneous reactions to

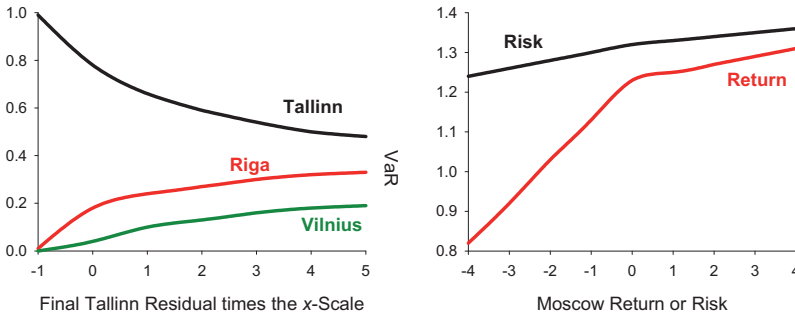


Figure 6: Allocations after shocking the final negative residual for Tallinn (left exhibit, a value on the x-scale larger than +1 means a larger negative shock). VaR effects of shocks to Moscow return, for given $z^- = -4$ and risk for given $x^+ = 0.1$ (right exhibit).

movements in other relevant assets or markets. Moreover, the model is able to capture asymmetric impacts of lagged positive and negative shocks on returns and volatility processes. We argue that measuring simultaneous and asymmetric spillovers is important for a number of reasons, including optimal portfolio allocation and risk management.

In summary, the empirical analysis provides support for the simultaneity in return and volatility. Accounting for simultaneity is of particular importance for markets located in the same geographic region or closely related due to institutional structure or other practical considerations as for example common trading platform. Given the fact that investors diversify their holdings across markets in order to reduce the risk of the portfolio, accounting for information which simultaneously alters the expectations of different markets is important for asset allocation and risk management strategies.

Empirically, we illustrate the importance of simultaneity with respect to Baltic States' stock markets. In these closely related markets

simultaneity is likely to arise due to geographical proximity, common institutional setup as well as common large traders, among other things. We found strong evidence of simultaneous effects in both returns and volatility. In returns, Riga is dependent on the indices of Tallinn and Vilnius, Tallinn is dependent on Vilnius, while Vilnius is not influenced by the other two markets. For volatility, we find within a day spillovers from Tallinn to both Riga and Vilnius. In addition, we found asymmetric effects of Moscow returns on the index returns in the Baltic States' exchanges, and asymmetric effects of Moscow risk on volatilities.

To illustrate the importance of simultaneous interaction between markets we obtain the portfolio allocations and value at risk measures for the multivariate and univariate models. Portfolio allocation results indicate that optimal portfolio weights are more sensitive to shocks when simultaneity is not accounted for. VaR measures indicate that the variability in losses that may occur due to shocks to the market are larger when simultaneity is not accounted for.

The simultaneous and dynamic econometric model generalizes previous univariate models by allowing for simultaneity but also for cross-effects of innovations. As in any simultaneous model we can therefore talk about direct, indirect and total effects in the return and volatility functions. The direct effects can be seen in the estimation results, while the portfolio and value at risk results build on total effects. To estimate the model we employ full information maximum likelihood. The suggested stepwise specification procedure resulted in a model with important deviations from corresponding univariate models. Estimation of the final model does not result in numerical problems despite the fact that the model is quite richly parametrized.

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Appendix

Table A: Estimation results for univariate models.

Variables	Riga		Tallinn		Vilnius	
	Return	Risk	Return	Risk	Return	Risk
y_{t-2}					0.057	0.021
u_{t-1}^+	-0.146	0.048	-0.072	0.018	0.252	0.041
u_{t-2}^+			0.117	0.037	0.021	0.022
u_{t-3}^+						
u_{t-1}^-			0.394	0.071	0.119	0.046
u_{t-2}^-			-0.283	0.064	-0.190	0.181
h_{t-1}			0.944	0.005	0.917	0.009
w_{t-1}^2			0.389	0.034	0.113	0.034
w_{t-2}^2			-0.322	0.031	-0.135	0.031
x, z_t^+	0.050	0.021	-0.001	0.001		
x, z_{t-1}^+					0.046	0.0167
					-0.032	0.005

Table A continued

Variables	Riga		Tallinn		Vilnius							
	Return	Risk	Return	Risk	Return	Risk						
x, z_t^-	0.105	0.021	0.121	0.0167	0.120	0.011	0.046	0.012	0.126	0.015	0.007	0.004
x, z_{t-1}^-			-0.114	0.0167	0.046	0.012	-0.050	0.012				
x, z_{t-2}^-					0.029	0.013						
Constant	0.177	0.033	0.079	0.012	0.114	0.027	-0.035	0.004	0.141	0.027	0.004	0.020
AIC	2086.9				1164.5				1446.8			
$\ln L, R^2$	-1029.5		0.03		-566.87		0.16		-709.41			0.06
LB10	10.84		8.83		7.01		1.53		21.57			1.53
Skew, Kurt, JB	0.43	5.60	2303.5		0.439	6.86	3446.6		-0.23	6.48		3030.7

III

Impact of Political News on the Baltic State Stock Markets*

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Abstract

This paper studies the link between political news releases, and the returns and volatilities in the stock markets of Riga, Tallinn and Vilnius. Political news releases are viewed as proxies for political risk. The results indicate that political news events regarding domestic and foreign political issues, excluding Russia, led, on average, to lower uncertainty in the stock markets of Riga and Tallinn in 2001-2003. At the same time, political risk from Russia increased the volatility of the stock market in Tallinn. We found that there is only a weak relationship between political risks of different origins and the stock market volatility in the Baltic states in 2004-2007. In addition, we found a significant "Monday effect", consistent with the trading behavior of institutional investors.

Key Words: Public information arrival, political risk, volatility, multivariate GARCH

JEL Classification: C32, G10, G14, G15

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1 Introduction

The question of what drives asset price movements has been a subject of interest in many empirical studies. One of the well-established empirical facts is the link between public information and changes in asset prices. To this end, we intend to examine the importance of different publicly available news releases for the stock market movements in the Baltic countries. More specifically, we look at the number of political news headlines during a day, as a proxy for the information flow.

It is a commonly held belief that stock prices equal present discounted values of rationally forecasted future dividends. If news announcements affect either expectations about future dividends or discount rates, or both, the news affects the daily stock price movements (McQueen and Roley, 1993). Consequently, as new information arrives, investors adjust their expectations about the market conditions, which in turn should be reflected in the equilibrium asset prices. Equity prices should increase if a news announcement leads to an upward revision of investors' expectations and vice versa (Tan and Gannon, 2002). Thus, asset prices in equilibrium may reflect *ex ante* premia for political risk.¹ The volatility is also related to the rate of information flow to the market (Ross, 1989). The brief description of the theoretical models regarding the effects of the public news releases is provided by Äijö (2008). In general, the information arrival can increase the volatility level of the market due to more information faced by investors, their divergent interpretations of the news, or higher market uncertainty if news is considered as bad or is highly unexpected. The information arrival can also lower the level of volatility due to a reduced degree of market uncertainty followed by the news announcement, or if news is considered as good. That

¹Howell and Chaddick (1994) define political risk as the "possibility that political decisions, events, or conditions in a country, including those that might be referred to as social, will affect the business environment such that investors will lose money or have a reduced profit margin".

is, depending on the state of the economy and given the diverse scope and nature of political news, there is a possibility that "no news is good news".

In this paper, we empirically study the impact of different political news on stock market returns and risks in the Baltic countries. The underlying motivation for our analysis is, in part, the fact that emerging and transition markets are particularly sensitive to political factors and events (see e.g., Bailey and Chung, 1995; Durnev et al., 2004; Goriaev and Zabotkin, 2006). The impact of different political events on the behavior of the Baltic stock markets is of special interest not only because of the abundance of political events, but also due to the history and recent development of the markets. Despite their small sizes, the Baltic stock markets developed well in recent years, both in term of returns, market capitalization, and increasing accessibility to international investors. As for the historical development, the Baltic countries have experienced the consequences of restructuring in the geopolitical context: from the dissolving of the Soviet Union in the early 1990s, to gaining access to the European markets as well as joining the European Union and NATO. Obviously, given the historical development, it is interesting to study whether the origin of political risk (i.e. political events) matters for investors' perception of a market risk. In particular, we are interested in studying whether the political risk (i.e. political events) related to Russia as well as to the domestic and other foreign political issues, excluding Russia, have different impacts on the stock markets in the Baltic countries. Given the economic and historical ties between Russia and the Baltic countries, it is likely that political confrontations between the countries may affect the expectations about future economic activity, and should, therefore, be reflected in the stock market prices. However, it is likely that risk factors related to Russia have become less important over time due to favorable economic development in recent years, admission into the EU and NATO in the spring of 2004, and changes in the

Russian foreign political agenda; see Morozov (2004) for more details. In a similar way, it is reasonable to assume that due to general economic and political developments related to, for example, the enlargement of the EU, European factors have become more important for investors' perception of market risk. This was previously indicated by Pajuste et al. (2000), who found that European risk factors are becoming increasingly important in Central and Eastern European emerging stock markets, with the preparations for joining the EMU, among other things. Gorjaev and Zobotkin (2006) found also that, while major economic and political news events had an impact on the prices of the Russian stock market, the political risk factors mattered the most until a certain level of corporate governance was reached. So far, we are not aware of any scientific study that analyzes the effect of political events on the asset price movements in Baltic stock markets.

In this paper, the news variables do not cover all the sources of information flow, but account for the political risk only. Mateus (2004) noted that not only political changes, but also economic spillovers from Russia have an important role in the explanation of returns in the EU accession counties, including the three Baltic countries. Earlier studies (see e.g., Pajuste, 2002; Brännäs and Sultanaeva, 2006; Brännäs et al., 2007) found that, given tighter economic and political links with Russia, as well as the geographic proximity, the Baltic countries are influenced by the stock market performance in Russia. Including the Russian stock market index (RTS) as an explanatory variable in return and volatility expressions allows us to study whether there are any spillovers from the Russian stock market, after taking the political news events into account.

We employ a multivariate time series model designed to catch the impact of news on returns and volatility. The conditional mean or return follows a vector ARMA process allowing for news effects, while the model for conditional variance or risk is based on the GJR model of Glosten et al. (1993). In the univariate framework, the GJR model is an extended

version of the GARCH model of Bollerslev (1986) and is able to capture asymmetric effects of positive and negative shocks on volatility. Hoti et al. (2002) extended the GJR model to a multivariate framework in order to incorporate volatility spillovers across markets. In this paper, we use the multivariate GJR model of Hoti et al. (2002, 2005) to capture the impact of political news. The model allows us to study whether news affecting the market risk in one of the Baltic countries has any impact on the market risks in the other two countries. Accounting for volatility links is important for portfolio management decisions, since the risk exposure of an international portfolio depends on the cross-market correlations of volatility changes; see also Fleming et al. (1998). Notably, extensions of this type introduce additional parameters into an already richly parameterized model. Kroner and Ng (1998), De Goeij and Marquering (2005) and others discussed ways of parameterizing, in particular, the volatility functions for models to be estimable. To allow for news impacts, we have to be restrictive in terms of correlation structure, lag lengths, and spillover effects.

We find that the arrival of domestic and foreign political news, excluding Russia, resolved some uncertainty, i.e. lowered the risk in the stock markets of Riga and Tallinn during years 2001-2003. During the same period, Russia-related political news had a risk-increasing effect on the stock market of Tallinn. It is interesting to note that despite the abundance of political events during years 2004-2007, the sensitivity of the stock markets in Tallinn and Riga to political events was lower during this sample period. Such a phenomenon could be explained by the changes in the investors' view on the general economic, financial conditions and political stability of the Baltic countries, after admission into the EU/NATO. In addition, we find that there is a weekend-effect consistent with the behavior of informed institutional investors. Overall, our results indicate that the impact of news could depend on both the origin and the nature of political news events, but also specific features

of the stock markets in the transition economies.

The remainder of the paper is organized as follows. First, we provide a brief description of previous literature that studied the impact of political news announcements. Section 2 provides background information on the news data and stock market data, while Section 3 describes the employed econometric model and presents the estimation technique. Section 4 gives the empirical results. The major findings are summarized in the final section.

1.1 Related literature

Most of the prior research used scheduled macroeconomic, firm specific, and/or other economic news announcements as proxies for public information. Many studies found significant relationships between different news announcements and equity, interest rate, or foreign exchange markets (see e.g., Ederington and Lee, 1993; Andersen and Bollerslev, 1998; Chang and Taylor, 2003; Kalem et al., 2004) in both intraday and daily data. Mitchell and Mulherin (1994) found that the flow of public information, measured as the daily number of headlines released by Dow Jones, is only weakly correlated with the volatility of several US indices. Using intraday data, Berry and Howe (1994) found no significant relationship between public information and price volatility. They referred to public information as firm or industry specific, political and macroeconomic news releases relevant to the US market.

Several other papers studied the importance of political news for the performance of financial markets focusing mainly on the emerging markets. For example, Chan and Wei (1996) studied the impact of political news on the stock market of Hong Kong. They found that the impact of political news depended on whether the index consisted of stocks controlled by Hong Kong or by the Chinese state-owned enterprises, where China related stocks were considered as a safe heaven. Consequently, the index composition may play a role in sensitivity to political news. Bailey

and Chung (1995) noted also that political risk can have an impact on stock prices and that "firms whose cash flows are particularly sensitive to general economic conditions may be exposed to political risks due to their broad impact on the economy". Chan et al. (2001) studied the impact of salient political and economic news on the volatility of the Hong Kong stock exchange. Interestingly, they found that political news had a distinct impact on market activity when compared to economic news, owing to differences in the information quality and investors' perceptual biases. Focusing on the same stock market, Kim and Mei (2001) used a volatility filter in order to measure the impact of political announcements on returns and volatilities. They found that large market movements were often associated with major political news. Studying the Asian crises, Kaminsky and Schumukler (1999) found that during this period political news, including political events and talks, and international agreements, had a significant impact on the changes in the stock market prices in some Asian countries. In addition, they found that not only local news, but also news of foreign origin mattered for the stock market movements.

Note that some previous studies (e.g., Chan and Wei, 1996; Kim and Mei, 2001) used different classification methodologies in order to divide political news into good and bad (and neutral). Chan and Wei (1996) noted though that some news headlines were difficult to classify. In this study, we find that the available data source and changes in the political climate during the considered time period (2001-2007) would make it difficult classify news in a way that would ensure objectivity. However, we utilize a data approach that allows us to assess whether more news events during a day reflect an increase in political risk; see Section 2 for more details.

2 Data

In this paper we use publicly available political news announcements as the proxy for public information. Although this proxy yields an imperfect treatment of the information available to market participants, it provides a reasonably broad, observable variable, that allows us to address the question about the impact of political risk on the studied stock markets. Daily news announcements are collected from the Russian News and Information Agency "RIA Novosti". To get as broad a news database as possible, news announcements were collected from the Russian language version of the web site (Politics section), generating a sample period from October 16, 2001 to October 1, 2007. News in English and other languages are available as well, but for a shorter time period. News announcements are not observable during a period from November 3 to December 31, 2003. Instead of elaboration on modeling the missing data period, the total sample period is divided into two sub-periods. However, the split in data coincides with the period after which the Baltic countries joined the EU and NATO. Consequently the two subperiods could be used to illustrate a possible change in the political atmosphere.

The following criteria were imposed in selecting the political news. We collect all the news with a clear reference to, at least, one of the Baltic countries in the headline, to start with. In general, political news in the database includes (i) agreements, (ii) political events and conflict, (iii) talks and statements about current and future policy actions as well as political conflicts. As we intend to study the impact of political risk of different origins, in the second step we separate the political news in the initial database into two categories. First, we select the political news related to Russia and Baltic countries. The main recent controversies were related to such issues as signing border treaties, advocating the rights of the Russian-speaking population in the Baltic countries, as well as the political and economic arrangements related to the EU

and NATO enlargement to the Baltic countries. The most recent news releases cover the political crisis over Estonia's relocation of a Soviet war memorial, and the pipeline agreement between a German gas company and Russia's natural gas monopoly Gazprom on transporting Russian gas to Germany via the Baltic Sea, i.e. bypassing the Baltic states. Second, we consider news related to each one of the Baltic countries (i.e. domestic political activity), and relations between Baltic countries, EU/NATO, and other countries, excluding Russia. That is, we create two news variables that allow us to study whether political news related to Russia has a different impact on the Baltic stock markets compared to domestic or other foreign news events, excluding Russia. Given the nature of political news, we do not attempt to classify political news events into "good" and "bad". The adjustment of the stock market prices to new information usually depends not only on the content of the news, but also the investors' interpretation of the news, as well as on the extent to which investors are caught by surprise (e.g., Kim, 2003). Consequently, in this paper, we examine the average response of the markets to different types of political news without assigning any valuations to the specific political news events; in part this is because we are uncertain whether our judgement would coincide with the market participants' perception of the content of news.

To avoid the double counting of news, if a chain of news headlines, regarding exactly the same issue, event, or statement appeared on consecutive days, or several times within a day, only the first-day-headline in the sequence of news is considered a news event. Consequently, our news selection criteria is based on the assumption that investors only react to new information and that news already known to investors are priced in the market. Note that one potential weakness in our analysis is that while we avoid the double counting of news headlines, we lose information about the relative importance of different news stories. However, Mitchell and Mulherin (1994) showed that using the total num-

ber of news stories released by Dow Jones & Company as a proxy for importance of public information did not improve the results.

In a final step, to facilitate estimation of the impact of news on all three markets, the news releases are transformed into count data variables. For each Baltic country, the variables take a value of 1 when there is one headline, 2 if there are two news announcements during a day, and so on. In a similar way, if there is one news release related to two or all three of the Baltic countries at the same time, the news variable for the particular countries takes on a value of 1. Consequently, we use a daily news count rather than a dummy variable approach in order to assess whether a greater number of news announcements, i.e. more information faced by investors, induces greater return variability. That is, despite the fact that we do not classify news based on the importance, we are able to assess whether more news events during a day reflects an increase in political risk.

2.1 Description of the News Data

In this section we study the political news variables. Table 1 reports the number of news headlines in each category for both weekdays and weekends. The lowest number of news is found for Vilnius. We find also that weekend news events account for 8-11 percent of all news, indicating that there may be a potential weekend-effect on the stock markets.

Table 2 reports some basic statistics for the news variables. The average number of news varies between 0.23 and 0.37 depending on news category and country, where Vilnius accounts for lowest number of news. The maximum number of news headlines varies between 4 and 5 for Riga and Vilnius, while it is more than twice as high for Tallinn. The peak of 13 news headlines regarding Tallinn and Russia relations (as well as 9 for other Tallinn related news) coincides with the political crisis over the relocation of a Soviet war memorial, during late April and early May of 2007. Interestingly, there is more clustering of news during the second

Table 1: Descriptive statistics for daily news series. Based on political news for the whole sample period, excluding missing observations ($T = 1512$).

News category	Number of headlines		Weekend ratio (%)	Frequency of days with no news	
	Weekdays	Weekends		Weekdays (%)	Weekends (%)
Riga vs Moscow	518	47	8.3	73.6	97.8
Riga	436	40	8.4	75.7	97.7
Tallinn vs Moscow	511	60	10.5	75.2	97.3
Tallinn	554	70	11.2	73.4	96.5
Vilnius vs Moscow	352	35	9.0	81.9	98.1
Vilnius	358	43	10.7	81.0	97.4

Table 2: Descriptive statistics for daily news series. Based on political news for the whole sample period, excluding missing observations ($T = 1512$).

News category	Mean	St.Dev.	Max	Skewness	Kurtosis	LB ₁₀ ¹	LB ₁₀ ²
Riga vs Moscow	0.342	0.649	5	2.243	6.350	18.87	249.45
Riga	0.288	0.564	4	2.228	5.900	10.81	109.88
Tallinn vs Moscow	0.338	0.803	13	6.237	69.986	11.93	801.13
Tallinn	0.366	0.769	9	4.117	31.080	20.33	602.73
Vilnius vs Moscow	0.233	0.564	4	3.057	11.648	31.14	151.29
Vilnius	0.237	0.544	4	2.751	9.440	22.19	197.07

Note: LB₁₀¹ and LB₁₀² are the Ljung-Box statistic evaluated at 10 lags for the first ($T=535$) and second ($T=977$) sample periods, respectively.

sample period as indicated by the Ljung-Box statistic presented in Table 2.

The Spearman correlation between news categories is reported in Table 3. We find the highest correlation coefficient within each news category. For example, the correlation between Moscow related news in Baltic countries range between 24 and 29 percent. In a similar way, the correlation between other domestic or foreign news (excluding Russia) is about 25 percent for Riga vs Tallinn, Riga vs Vilnius, and Tallinn vs Vilnius.² These results can be explained by an overlap in the news headlines, in the sense that the same issue, as for example admission to NATO or EU, or regarding Russian minorities in the Baltic countries, may be brought up simultaneously for all three Baltic States. In addition, news about political relations within the Baltic countries, are likely to have an impact on the correlation between news categories. Another explanation is the possibility of spillover effects of news (Janssen,

²Note that the results presented in Table 3 cause no multicollinearity problems as indicated by the eigenvalues of the correlation matrix.

Table 3: Spearman correlation between news categories. Based on political news for the whole sample period, excluding weekends and missing observations ($T = 1512$).

News category	Riga vs Moscow	Riga	Tallinn vs Moscow	Tallinn	Vilnius vs Moscow	Vilnius
Riga vs Moscow	1.000					
Riga	0.031	1.000				
Tallinn vs Moscow	<u>0.290</u>	<u>0.054</u>	1.000			
Tallinn	0.045	<u>0.258</u>	<u>0.111</u>	1.000		
Vilnius vs Moscow	<u>0.245</u>	<u>0.065</u>	<u>0.237</u>	0.027	1.000	
Vilnius	<u>0.067</u>	<u>0.256</u>	0.045	<u>0.247</u>	<u>0.101</u>	1.000

Note: Underlining is used to indicate significant correlations.

2004). That is, the political conflicts between, for example, Moscow and Tallinn, could affect the political activity between Moscow and other Baltic countries as well.

2.2 Stock Index Data

The stock index data used are capitalization weighted daily stock price indices of the Estonian (TALSE), Latvian (RIGSE), Lithuanian (VILSE), and Russian (RTS) stock markets. All prices are in Euro.³ The dataset covers October 16, 2001 to October 1, 2007, for a total of $T = 1555$ observations.

Due to differences in holidays, the country series have different shares of days for which index stock price are not observable. Linear interpolation was used to fill the gaps for all series. The resulting series are then throughout for a common trading week. All returns are calculated as $y_t = 100 \cdot \ln(I_t/I_{t-1})$, where I_t is the daily price index. Table 4 reports descriptive statistics for the daily returns. The Ljung-Box statistics for 10 lags (LB₁₀) indicate significant serial correlation. The large excess kurtosis indicates leptokurtic densities.

2.3 Model and Estimation

The primary purpose of this paper is to model the relationship between asset return, volatility movements, and news arrival. To account for the news effect we expand the asymmetric VARMA-GARCH (or VARMA-AGARCH) model of Hoti et al. (2002, 2005). The model is a multi-

³Note that the Euro prices are assumed to reflect the risk faced by market participants, in part, because foreign institutional investors, that need to take currency risk into account, represent a large part of investors. In addition, all three Baltic countries have fixed exchange rate regimes, where the Estonian kroon had been pegged to Euro since its introduction as electronic currency in 1999. Moreover, all trading on the Tallinn stock market is in Euro. The Latvian and Lithuanian currencies were pegged to the SDR basket (which includes Euro) and the US dollar until January 2005 and February 2002, respectively, when they switched to a Euro peg.

Table 4: Descriptive statistics for daily return series.

Exchange	Mean	St.Dev.	Min/Max	Skewness	Kurtosis	LB ₁₀	<i>T</i>
Riga	0.07	1.09	-7.86/6.97	-0.30	12.30	17.92	1554
Tallinn	0.13	0.92	-5.87/7.18	-0.06	10.80	68.91	1554
Vilnius	0.14	1.01	-13.52/11.87	-0.33	40.83	73.22	1554
Moscow	0.12	1.69	-9.91/17.71	1.43	17.91	40.61	1554

Note: LB₁₀ is the Ljung-Box statistic evaluated at 10 lags.

variate generalization of the asymmetric GARCH (or GJR) model of Glosten et al. (1993) that takes into account asymmetries in financial data. Consider the following specification for the return:

$$\begin{aligned}
 \mathbf{y}_t = & \mathbf{a}_0 + \sum_{i=1}^p \mathbf{A}_i \mathbf{y}_{t-i} + \sum_{j=1}^r \mathbf{B}_j \mathbf{d}_{jt} + \mathbf{B}_m \mathbf{d}_{mt} \\
 & + \sum_{i=1}^s \mathbf{C}_i \mathbf{x}_{t-i} + \mathbf{u}_t + \sum_{i=1}^q \mathbf{F}_i \mathbf{u}_{t-i}
 \end{aligned} \tag{1}$$

where $\{\mathbf{y}_t\}$ is a $N \times 1$ weakly stationary time series sequence, and $\{\mathbf{x}_t\}$ denotes a sequence of exogenous variables that may affect the process $\{\mathbf{y}_t\}$. In this paper, \mathbf{x}_t represents return at time t in the RTS index. The \mathbf{d}_{jt} is a $N \times 1$ vector of variables for news category $j = 1, 2$. In the empirical study, the variable $\mathbf{d}_{1,t}$, represents news related to Russia, and $\mathbf{d}_{2,t}$ includes domestic and/or other foreign news events, excluding Russia. The \mathbf{d}_{mt} is a dummy variable to capture a possible Monday effect, i.e. the elements of \mathbf{d}_{mt} variable takes on a value of 1 for Mondays and zero otherwise.⁴ The \mathbf{a}_0 is a vector of constants, \mathbf{A}_i , \mathbf{B}_j , \mathbf{B}_m , and \mathbf{C}_i

⁴The underlying motivation for the use of a Monday dummy is based, in part, on earlier studies (see e.g., Ederington and Lee, 1993; Beriment and Kiyamaz, 2001, 2003; Edmonds and Kutan, 2002; Kim, 2003; Kalev et al., 2004) that found significant day-of-the-week effects, with the lowest return and the highest volatility on Mondays.

are diagonal matrices of dimension $N \times N$. Equation (1) incorporates effects across equations, and hence spillovers in returns, through off-diagonal elements in \mathbf{F}_i .

Further, $\{\mathbf{u}_t\}$ is a stochastic $N \times 1$ vector process such that $E\mathbf{u}_t = \mathbf{0}$. The \mathbf{u}_t is conditionally heteroskedastic and generated by:

$$\mathbf{u}_t = \mathbf{H}_t^* \boldsymbol{\varepsilon}_t \quad (2)$$

where $\{\boldsymbol{\varepsilon}_t\}$ is an i.i.d. discrete time, vector error process with $E\boldsymbol{\varepsilon}_t \boldsymbol{\varepsilon}_t' = \mathbf{I}$ and $V(\mathbf{u}_t | \mathcal{F}_{t-1}) = \mathbf{H}_t^* \mathbf{H}_t^{*'} \equiv \mathbf{H}_t$, where \mathcal{F}_t denotes the past information up through time t . To specify \mathbf{H}_t various alternative asymmetric models are possible (Hoti et al., 2002; De Goeij and Marquering, 2005). We specify the conditional variance model as $\mathbf{h}_t = \text{diag}(\mathbf{H}_t)$, and treat off-diagonal elements of \mathbf{H}_t as time-invariant. The model for conditional variances is given by

$$\begin{aligned} \mathbf{h}_t = & \mathbf{g}_0 + \mathbf{G}_1 \mathbf{h}_{t-1} + \sum_{j=1}^P \mathbf{V}_j \mathbf{d}_{jt} + \mathbf{V}_m \mathbf{d}_{mt} + \sum_{i=1}^R \mathbf{W}_i \mathbf{z}_{t-i} \\ & + \left(\sum_{i=1}^S \mathbf{K}_i + \sum_{i=1}^Q \mathbf{K}_i^- \mathbf{I}_{t-i} \right) \mathbf{u}_{t-i}^2 \end{aligned} \quad (3)$$

from which the corresponding \mathbf{H}_t matrix can be obtained. The \mathbf{g}_0 is a vector of constants, \mathbf{G}_1 , \mathbf{V}_j , \mathbf{V}_m , and \mathbf{W}_i are diagonal $N \times N$ matrices, while \mathbf{K}_i , \mathbf{K}_i^- are $N \times N$ matrices with typical elements k_{ij} and k_{ij}^- , respectively, for $i, j = 1, \dots, N$. The $\mathbf{h}_t = (h_{1t}, \dots, h_{Nt})'$ and $\mathbf{u}_t = (u_{1t}, \dots, u_{Nt})'$. The vector \mathbf{u}_t^2 has elements u_{it}^2 ($i = 1, \dots, N$). The \mathbf{z}_t series entering the conditional variance function is the Moscow stock market (RTS) moving variance series for a window length of 10 observations. Equation (3) incorporates multivariate effects across equations, and hence spillovers in volatility through off-diagonal elements in \mathbf{K}_i and \mathbf{K}_i^- . Thus, h_{it} contains past information from u_{it}^2 and u_{jt}^2 for $i, j = 1, \dots, N$, $i \neq j$, but not from $u_{it} u_{jt}$. Hoti et al. (2002) define h_{it}

to contain past information from u_{it} , u_{jt} , h_{it} and h_{jt} for $i, j = 1, \dots, N$, $i \neq j$. The indicator variable $\mathbf{I}_t = \text{diag}(I_{it}, \dots, I_{nt})$ is a $N \times N$ matrix with

$$I_{it} = \begin{cases} 1, & \text{if } u_{it} \leq 0 \\ 0, & \text{otherwise} \end{cases}$$

This threshold term is designed to capture the asymmetric nature of volatility responses to positive and negative shocks to the market. In the empirical results section, the variable u_{it}^2 captures the impact of both positive and negative shocks, while the variable $u_{it}^{2,-} = I_{it}u_{it}^2$ captures the volatility responses to negative shocks only. The conditional correlations among the elements of $\{\mathbf{u}_t\}$ can be calculated as $\rho_{ij,t} = \mathbf{H}_{ij,t} / (\mathbf{H}_{i,t}\mathbf{H}_{j,t})^{1/2}$.

It is important to note that the empirical estimation of (3) may become infeasible with too generously parameterized specifications. Reducing lag lengths and introducing sparse matrix specifications are two ways of reducing the number of parameters. Section 3.1 below describes a data-driven model specification procedure.

2.4 Empirical Modelling Strategy

The assessment of the relationship between news arrival and asset price movements is done in three steps. In each step we employ the AIC criterion to find a parsimonious parameterization. First, we estimate univariate ARMA-GJR models for each Baltic stock exchange containing specifications for both mean return and conditional variance and accounting for spillovers from the stock market in Moscow through variables \mathbf{x}_t and \mathbf{z}_t . Thus, in the first step we implicitly assume that there is no interaction between the series. In the second step, we expand to a multivariate specification including non-diagonal matrices in the volatility expression. A paper by Brännäs et al. (2007) indicated that there are significant spillovers between the Baltic stock markets. Consequently, we relax the assumption of no interaction between the series

and test for spillovers through off-diagonal elements in matrices \mathbf{K}_i and \mathbf{K}_i^- . However, to make the estimation of (3) feasible, we are restrictive in terms of correlation structure, lag length, and spillover effects. For this purpose, we assume that h_{it} depends only on u_{it-1}^2 , u_{jt-1}^2 and h_{it-1} for $i, j = 1, \dots, N$, $i \neq j$, but not on cross terms, $u_{it}u_{jt}$. In the third and final step, we include news variables \mathbf{d}_{jt} and a Monday dummy \mathbf{d}_{mt} , accounting for a possible Monday effect, both in the conditional mean and conditional variance functions. This allows us to test whether news contributes significantly to the return and volatility dynamics.

2.5 Estimation

Assuming a multivariate normality of $\{\varepsilon_t\}$ the prediction error is given as:

$$\mathbf{y}_t - E(\mathbf{y}_t | \mathcal{F}_{t-1}) = \mathbf{u}_t = \mathbf{H}_t^* \varepsilon_t \quad (4)$$

which is i.i.d., $N(\mathbf{0}, \mathbf{H}_t)$. Here, \mathbf{H}_t is the conditional variance expression. Given observations up through time T , the log-likelihood function takes the form:

$$\ln L \propto -\frac{1}{2} \sum_{t=s}^T \ln |\mathbf{H}_t| - \frac{1}{2} \sum_{t=s}^T \mathbf{u}_t' \mathbf{H}_t^{-1} \mathbf{u}_t \quad (5)$$

where $s = \max(p, r, s, q, P, R, S, Q) + 1$. For practical estimation the RATS 6.0 package is employed, using robust covariance matrices throughout. There was no parameter restrictions imposed during the estimation. However, the likelihood function in equation (5) requires the conditional variance matrix to be positive definite in order for the estimation to be feasible.

3 Results

The empirical results for the first and second sample periods are presented in Table 5. We start with discussing the results for the first sample period, covering October 16, 2001 to November 3, 2003.

Looking at the conditional mean equation, we find that news related to domestic and foreign political issues, excluding Russia, has a negative impact on returns in Riga, as indicated by the parameter estimate for the news variable $d_{2,t}$. This indicates that these news events are considered as unfavorable by the investors on the stock market of Riga. There is no significant impact of either the Moscow related news variable ($d_{1,t}$) or the other political news variable ($d_{2,t}$) on returns in the stock markets of Tallinn and Vilnius. Although our measure of political news does not appear to be significant for the market returns in the stock markets of Tallinn and Vilnius, it does not necessarily mean that political news is not a significant variable, in part because we made no attempt to separate news into good and bad (McQueen and Roley, 1993). We also find that the RTS index (\mathbf{x}_{t-1}) has a small, but positive impact on the returns in Tallinn. This implies that the information priced on the Russian stock market will have an effect on the stock market prices in Tallinn, despite the insignificant impact of political news related to Russia. The parameter values for the d_{mt} variable indicate that the returns on Mondays are lower than the mean return across all weekdays combined. The largest weekend-effect is found for the stock markets of Riga and Vilnius. The negative weekend effect is consistent with earlier studies (see e.g., Chang et al., 1998; Berument and Kiyamaz, 2001; Kim, 2003). Sias and Starks (1995) argued that the weekend effect is primarily driven by institutional investors (and/or discretionary liquidity traders), and it will therefore be stronger on markets dominated by the informed institutional investors. For the Baltic stock markets, the institutional investors outweigh individual investors with up to 90 percent of the market value (OMX Guide to Baltic Markets, 2007).

Turning to conditional volatility, we find that domestic and foreign political news events, excluding Russia, lower the risk in the stock exchanges of Riga and Tallinn. That is, the market participants' perception of news, regarding relations between the Baltic states, EU, NATO,

and other western countries or institutions, seems to be towards reduced uncertainty on the markets in Tallinn and Riga. This is consistent with the idea that if there is some uncertainty before the actual announcement, the volatility will decrease if the news release resolves the uncertainty of market investors (see e.g., Äijö, 2008). Interestingly, political news events related to Russia seem to have a risk-increasing impact on the volatility in Tallinn. This could be due to, for instance, investors' divergent interpretations of the news or that news is unexpected. In addition, we find significant volatility spillovers from Vilnius to Tallinn. A positive shock in Vilnius lowers the volatility in Tallinn, whereas a negative shock increases the risk in Tallinn. The conditional covariances are significantly estimated as $H_{12,t} = 0.086$ (s.e. = 0.037), $H_{13,t} = 0.062$ (0.032), and $H_{23,t} = 0.122$ (0.030). The resulting time-varying correlations between $\{\mathbf{y}_t\}$ variables are positive throughout and vary between 0.01 and 0.5.

The results for the second sample period, presented in Table 5, cover a period from January 1, 2004 to October 1, 2007. For the conditional return function, we note that there is a negative weekend-effect on returns in the stock markets of Riga and Tallinn. Moscow related political news events lower the returns in Riga. However, the effect is smaller (even though significant) than those in earlier years, i.e. the first sample period. There is no impact of either Moscow related political news ($d_{1,t}$) or other political news ($d_{2,t}$) on returns in the stock markets of Tallinn and Vilnius. Note that the index composition could be one possible explanation to why the Riga stock market reacts stronger to Russia related political news. For example, the largest companies in the Riga index belong to the energy sector, while the indices of the stock markets in Tallinn and Vilnius are more diversified.

The results for the conditional volatility function indicate that there is a positive Monday-effect on the stock markets of Riga and Vilnius, implying that market uncertainty increases on Mondays. Similar results

Table 5: Parameter estimates for the joint conditional return and risk functions for the first and second sample periods. The first sample period (Period 1) includes $T = 535$ observations, covering October 16, 2001 to November 3, 2003. The second sample period (Period 2) includes $T = 977$ observations, covering January 1, 2004 to October 1, 2007.

Variable	Period 1			Period 2		
	Rīga, $y_{t,1}$	Tallinn, $y_{t,2}$	Vilnius, $y_{t,3}$	Rīga, $y_{t,1}$	Tallinn, $y_{t,2}$	Vilnius, $y_{t,3}$
y_{t-1}	-0.121 **	0.240 ***	0.113 **	-0.039	0.214 ***	0.179 ***
y_{t-2}			0.092 *			0.063
x_{t-1}	0.044	0.053 **	0.021	0.023	-0.007	0.005
d_{mt}	-0.189 **	-0.022	-0.203 **	-0.123	-0.105 **	0.006
$d_{1,t}$	-0.115	0.152	-0.008	-0.075 **	0.002	-0.023
$d_{2,t}$	-0.136 **	-0.042	0.089	0.075	-0.021	0.001
Constant	0.161 ***	0.088 *	0.146 ***	0.146 ***	0.079 ***	0.129 ***

Table 5 continued

	Period 1			Period 2		
	Volatility function			Volatility function		
	Riga, $h_{t,1}$	Tallinn, $h_{t,2}$	Vilnius, $h_{t,3}$	Riga, $h_{t,1}$	Tallinn, $h_{t,2}$	Vilnius, $h_{t,3}$
$w_{1,t-1}^2$	0.248 ***			0.064 ***		0.025 **
$w_{2,t-1}^2$	0.118			0.099 ***		
$w_{1,t-1}^2$		0.226 ***			0.287 ***	
$w_{2,t-1}^2$		-0.007			-0.158 ***	
$w_{3,t-1}^2$		-0.095 ***	0.187 ***			0.174 ***
$w_{3,t-1}^2$		0.268 ***	-0.007			-0.002
h_{t-1}	0.615 ***	0.503 ***	0.702 ***	0.799 ***	0.806 ***	0.763 ***
z_{t-1}	0.000	0.012	0.000	0.000	-0.000 **	0.004 **
d_{mt}	-0.118	-0.078	-0.097	0.180 **	0.008	0.245 ***
$d_{1,t}$	0.085	0.199 **	-0.005	-0.027 **	-0.008	0.033
$d_{2,t}$	-0.150 **	-0.191 ***	-0.017	-0.009	0.018 *	-0.020
Constant	0.209 ***	0.260 ***	0.108 ***	0.062 ***	0.015 **	-0.002

Table 5 continued

	Period 1			Period 2		
	Riga, $h_{t,1}$	Tallinn, $h_{t,2}$	Vilnius, $h_{t,3}$	Riga, $h_{t,1}$	Tallinn, $h_{t,2}$	Vilnius, $h_{t,3}$
LB ₁₀	12.46	9.89	9.42	15.80	12.59	12.10
LB ₁₀ ²	6.79	15.14	9.52	17.01	3.57	10.80
Skewness	0.22	0.04	-0.04	0.43	0.86	1.16
Kurtosis	3.69	1.55	1.12	4.36	11.96	14.99

Note: ***, **, * indicates significance at 1, 5 and 10 percent level, respectively.

The variable $d_{1,t}$, represents news related to Russia, while $d_{2,t}$ incorporates other foreign and domestic news. The variable d_{mt} is a Monday dummy. The variables x_{t-1} and z_{t-1} capture the return and volatility spillovers from the Russian stock market, respectively.

LB₁₀ and LB₁₀² is the Ljung-Box statistic for standardized residuals and their squares at lag 10. Kurtosis is the excess kurtosis of standardized residuals.

were found by Kiyamaz and Berument (2003) for the stock markets of Germany and Japan. French and Roll (1986) suggested that, given that more public information arrives during normal business hours, variances for days following an exchange holiday are larger than for other days, due to the trading strategies of informed traders.

Turning to the impact of political news, we find that Russia related political news events slightly lower the risk in the Riga stock market. Despite the abundance of political events during the second sample period, there seems to be little or no impact of the political news variables on returns and volatility in Tallinn. The lower impact of political news related to Russia during the second period could reflect the changes in the investors' view on the general political climate due to ongoing European integration. Consequently, the European integration, including admission to the EU and NATO, could have outweighed the effect of political crises with Russia. However, McQueen and Roley (1993) also suggested that the response coefficient may be biased towards zero, if some type of news is considered good in some states of the economy and bad in others. For Tallinn, we also find that there is an asymmetric impact of own positive and negative shock, where positive shocks have larger impact on the volatility. This phenomenon is consistent with findings of, e.g. Rockinger and Urga (2001), Brännäs and Soultanaeva (2006) and Brännäs et al. (2007). Studying the stock markets of several transition economies, Rockinger and Urga (2001) noted that in markets that suffer from low liquidity, positive shocks can lead to higher volatility. This can be explained by the fact that positive shocks open up new trade opportunities for investors that avoid selling in an illiquid market, which provides investors an opportunity to rebalance their portfolios once greater liquidity has been achieved, leading, in turn, to increased volatility. We also find significant volatility spillovers from Riga to Vilnius for the second sample period. The conditional covariances for the second sample period are $\mathbf{H}_{t,1,2} = 0.054$ (s.e. = 0.018), $\mathbf{H}_{t,1,3} = 0.049$

(0.026) and $\mathbf{H}_{t,2,3} = 0.048$ (0.016).

Overall, the volatility persistence is quite low for all three stock markets, for both sample periods. Kalev et al. (2004) and Janssen (2004) found that the inclusion of news variables leads to a substantial reduction in volatility persistence, partly due to a serial correlation pattern in the rate of information arrival. However, for the Baltic stock markets the volatility persistence is quite low, even in models with no news variables.⁵ The volatility and return spillovers from the Russian stock market (parameter values for x_{t-1} and z_{t-1} variables) are small and insignificantly estimated, except in the case of Tallinn during 2001-2003.⁶

The log-likelihood value increases significantly (the LR test rejects the null hypothesis at the 5 percent level) when news variables and the Monday dummy are jointly included in return and volatility expressions.⁷ To assess the specification of the final models, we calculate the time series of standardized residuals and their squares, for both sample periods. Summary statistics of the standardized residuals and their squares are presented in Table 5. If the models are correctly specified we expect the standardized residuals to be close to being i.i.d. distributed. We find no significant autocorrelations of standardized residuals and squared standardized residuals, as indicated by the Ljung-Box statistics.

In our study, we did not account for the possible pre-announcement effect of news. Given that the studied markets are dominated by informed institutional investors, it is possible that market participants

⁵The estimation results for the models with no news are available on request.

⁶Due to the overlapping periods of market trading activity in the Baltic and Russian stock markets we tested empirically for spillovers within a day through x_t and z_t variables (see also Koch and Koch, 1991). However, accounting for within a day spillovers from the Russian stock markets did not change the results.

⁷For the first sample period, the log-likelihood value increases significantly when only the political news variables are included in return and volatility expressions. However, for the second sample period, the log likelihood increases significantly only when the news variables and the Monday dummy are entered jointly.

trade on the information (i.e. rumors) before the actual announcements of news. To scrutinize the possibility for the pre-announcement effect on the Baltic stock markets, we run a regression of stock returns on lagged returns and leads of news variables.⁸ The findings indicate that there is no significant pre-announcement effect of political news in the stock markets under study, except for Vilnius, in the case of news related to Russia, in 2004-2007. Chan et al. (2001) also found that political news has no trading impacts on pre-event days. They suggested that the nature of political events makes it more difficult for investors to interpret the content of political events. Therefore, the investors are more likely to wait until the actual announcement before resuming active trading of stocks.

4 Concluding Remarks

This paper has studied the impact of public information on the stock market returns and volatility in Riga, Tallinn, and Vilnius. We adopted political news announcements as a proxy for public information. Under the assumption of efficient markets, each new information arrival will establish a new price equilibrium (see e.g., Kalev et al., 2004). That is, the market participants rely on all available information in forming their expectations of risk and return in the stock markets. Earlier studies (e.g., Durnev et al., 2004) showed that political factors are more important in emerging and transition economies, such as the three Baltic countries, than in developed economies. Political news events can therefore reflect a country's political risk, that can have an impact on the stock markets. The asset market in equilibrium should yield risk premiums due to exposure to such risks, if the effects of political events do not vanish in

⁸We run a regression of form $y_{i,t} = \alpha_0 + \alpha_1 y_{i,t-1} + \beta_{1,i} d_{1,i,t+1} + \beta_{2,i} d_{2,i,t+1} + \varepsilon_t$, where the index i stand for stock markets in Riga, Tallinn and Vilnius. The regression results are available on request.

well-diversified portfolios (Bailey and Chung, 1995). Therefore, understanding what drives asset prices is crucial for the analysis of the value of financial assets, and for various investment and risk management decisions. The main question of interest is whether the political news related to Russia has a different impact than news related to domestic and other foreign political events, excluding Russia.

Using a sample of index returns in the three Baltic countries, our study revealed that domestic and foreign political news, excluding Russia, lowered the risk in the stock markets of Riga and Tallinn during years 2001-2003. During the same period, political news related to Russia had a risk-increasing effect on the stock market of Tallinn. Political news had a smaller (even though significant) impact on the stock markets of Riga and Tallinn in 2004-2007. Vilnius does not seem to be affected by political news in either sample period. The overall findings suggest that the sensitivity of the Baltic stock markets to political events decreased over time. Gorjaev and Zabolotkin (2006) found that the relative importance of different risk factors for the Russian stock market varied over time, where political risk became less important after a certain level of market development is passed. In addition, the Baltic countries entered as EU and NATO members during the spring 2004, which coincides with the second sample periods. The lower sensitivity to political risk factors during the second sample period, could therefore reflect the investors' view on the general economic, financial conditions and political stability of the Baltic countries, after the admission into the EU/NATO.

Our results indicate that, despite common characteristics, there are substantial differences among Baltic stock markets, with respect to market adjustments to political news. This could be explained by the quality of information and investors' perceptual biases, regarding political news related to each one of the Baltic countries. That is, the stock market movements depend not only on the rate of information arrival, but also on differences in investors' opinions and interpretations of news

announcements (Kalev et al., 2004). Furthermore, the displayed differences in sensitivity to political risks may be caused by the industry composition, ownership and trade structure (see e.g., Pajuste et al., 2000). Bailey and Chung (1995) noted, for instance, that firms whose cash flows are especially sensitive to general economic conditions may be exposed to political risks. In addition, they argued that industries involved in international transactions are relatively more exposed to political changes. For Baltic stock markets, we can note that the companies with the largest weight in the Riga index belong to the energy sector, which could explain the sensitivity of returns to the political news related to Russia. The indices of the stock markets in Tallinn and Vilnius are more diversified, as the largest companies belong to different sectors including telecommunication, energy, industrials, and utilities.

Last but not least, consistent with earlier studies (see e.g., Chang et al., 1998; Berument and Kiyamaz, 2001, 2003; Kim, 2003) we find that the Monday return is on average lower than returns across all other weekdays. The volatility is, on other hand, higher on the days following the exchange holiday. Given the fact that institutional investors represent about 90 percent of the market value of the stock markets in Riga, Tallinn and Vilnius, this is consistent with the idea that Monday effect is primarily driven by the trading behavior of informed institutional investors (see e.g., French and Roll, 1986; Sias and Starks, 1995).

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IV

The Impact of Stock Market Jumps on Time-Varying Return Correlations: Empirical Evidence from the Baltic Countries*

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Abstract

In this paper we study the impact of market jumps on the time varying return correlations between stock market indices in the Baltic countries. An EARJI-EGARCH model facilitating direct modelling of the time varying return correlations is introduced. The empirical results indicate that there is a quite large number of identified jumps in the emerging Baltic stock markets. The main finding is that isolated market jumps in one of the markets generally have no or small effects on the time-varying correlations. In contrast, simultaneous jumps of equal sign increase the average correlation, in some cases with as much as 100 percent.

Keywords: Correlated jumps, contagion.

JEL: C32, C52.

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1 Introduction

Does international portfolio diversification give investors protection from extreme movements (or jumps) in local stock market prices? An answer to this question obviously depends on the correlation between returns in different markets included in the portfolio.

Empirical evidence concerning financial asset and market returns (e.g., Fustenberg and Jeon, 1989; Koch and Koch, 1991; Erb et al., 1994) often points towards time-varying correlation structures that tend to increase during unstable periods. Karolyi and Stulz (1996), Ramchand and Susmel (1998) and Longin and Solnik (1995, 2001), among others, found that the correlations between major stock markets rose during periods of high volatility or during market crises. In general, the literature on contagion¹ (e.g., Claessens, 2001; Forbes and Rigobon, 2002) often finds that the cross-market correlation coefficients in a stable environment are statistically different (lower) from the correlation coefficients during unstable periods. Even though international stock market correlations have received a lot of attention in the literature, due to its importance in portfolio and risk management (e.g., Fazio, 2007; Knif and Pynnönen, 2007; Campbell et al., 2008), less is known about correlation responses to large shocks (jumps) in stock market returns.

To shed some light on this issue, the current paper studies the impact of large discrete changes in stock market prices (jumps) on time-varying return correlations. More specifically, we study whether the correlations of stock market returns differ when there are smooth changes in market returns or large discrete changes (jumps). The study is performed on stock market data for the Baltic countries (Estonia, Latvia and Lithuania) which have previously received little attention in the financial literature. Despite the benefits of diversifying into emerging markets (see, Bekaert and Harvey, 2002), portfolio managers often shy

¹Contagion is commonly defined as an increase in stock market co-movements after a shock or a financial crisis.

away from these markets due to the high volatility. Thus, improved knowledge about emerging stock markets, in this case the Baltic stock markets, may reduce the uncertainty about such an investment.

In the empirical analysis a bivariate Exponential Autoregressive Jump Intensity (EARJI)-EGARCH model, based on Chan (2004), is introduced to identify stock market jumps as well as to estimate time varying return correlations. The possible effect of the identified stock market jumps on the time varying return correlations are then analyzed in separate regressions.

This paper contributes to the existing literature in several ways. First, in contrast to the earlier literature on contagion, where shocks or periods of financial turmoil usually are pre-defined by the authors, we utilize a data driven procedure to identify large shocks labelled as jumps. Thus, the results of this paper are, in this regard, more general and pertain to any market shocks rather than to prespecified shocks or to periods of financial turmoil. Second, the effects of market jumps on the dynamics of time-varying return correlations have not previously, to the authors knowledge, been addressed as directly as in this paper. In the literature on jump spillovers, Asgharian and Bengtsson (2006) found earlier that the correlation structure between the jump processes of returns differs significantly from the correlation between the regular return components (i.e. the parts that are not jumps). In comparison to Asgharian and Bengtsson (2006), we focus directly on the question of how jumps affect the return correlations, whereas they are mainly concerned with the correlation between jumps, as well as, spillovers of jumps between international markets. Note that a high correlation between jumps does not necessarily imply a high correlation between returns, since jumps may be in different directions, i.e. correspond to large positive and negative changes in stock market prices. Third, in this paper the class of mixed GARCH-jump models (e.g., Chan and Maheu, 2002) is extended to a specification with a time varying return correlation. The earlier multi-

variate models by e.g., Chan (2004), have instead implicitly modelled time varying return correlation through correlated jump components. Fourth, empirical evidence on the Baltic stock markets, which are less studied in the financial literature, is provided.

In Section 2 of the paper the econometric framework is outlined. Section 3 describes the data set used in the empirical study. Section 4 reports on the empirical results, while the final section discusses the results and concludes the paper.

2 A bivariate EARJI-EGARCH model with time-varying correlation

2.1 Background

The model used in the empirical analysis belongs to the class of GARCH-Jump mixture models originating from Press (1967), who introduced a jump model, where the arrival of jumps is governed by a Poisson distribution.² The early version of the model assumed that there is a constant number of large discrete price movements (jumps) within a fixed time interval. The average number of jump events in a time interval is called the jump intensity. Chan and Maheu (2002) extended the model to include time-varying (ARMA) jump intensities, whereas Hellström et al., (2008) introduced an exponential version of the time-varying jump intensity to account for asymmetric responses to jump innovations. Chan (2003, 2004) considered bivariate extensions of the model with correlated jump dynamics. However, their study, based on the multivariate GARCH parametrization (BEKK) is limited to the analysis of the correlation between jump components only. In contrast to the earlier studies, we focus directly on the correlation between market returns through a reparame-

²The basic jump model by Press has been used in mainly financial applications by, e.g., Jorion (1988), Vlaar and Palm (1993) and Nieuwland et al., (1994).

terization of the covariance matrix for the market returns (e.g., Tsay, 2002, ch. 10). Thus, the covariances and correlations between market returns capture the co-movements driven by both the innovations not associated with jumps and by the jump innovations. This allows us to directly study the impact of market jumps on the return correlation. To explore the effect of market jumps on the time-varying return correlations, actual jumps are first identified and then in a second step used to explain the time-varying correlation.

2.2 A bivariate EARJI-EGARCH model

To study the time-varying correlation between the stock market returns r_{1t} and r_{2t} a bivariate model based on Chan (2004) is outlined. The bivariate model structure, opposed to a trivariate structure, is chosen to simplify the identification of the time-varying second order moments.³ Given the information set at time $t - 1$, $\Phi_{it-1} = \{r_{it-1}, \dots, r_{i1}\}$ for $i = 1, 2$,⁴ the model is specified as:

$$\mathbf{R}_t = \boldsymbol{\mu}_t + \boldsymbol{\epsilon}_{1t} + \boldsymbol{\epsilon}_{2t}. \quad (1)$$

Here \mathbf{R}_t , $\boldsymbol{\mu}_t$, $\boldsymbol{\epsilon}_{1t}$ and $\boldsymbol{\epsilon}_{2t}$ are 2×1 vectors denoting the returns, the conditional mean functions specified as $\mu_{it} = \alpha_{0i} + \sum_{l=1}^L \alpha_{li} r_{it-l}$, the random disturbances, and the jump innovations, respectively.

The jump innovation component is defined as:

$$\boldsymbol{\epsilon}_{2t} = \begin{bmatrix} \sum_{k=1}^{n_{1t}} Y_{1t,k} - E_{t-1} \left(\sum_{k=1}^{n_{1t}} Y_{1t,k} \right) \\ \sum_{l=1}^{n_{2t}} Y_{2t,l} - E_{t-1} \left(\sum_{l=1}^{n_{2t}} Y_{2t,l} \right) \end{bmatrix}, \quad (2)$$

³Since the model is already highly parameterized in its univariate form and is estimated by integrating over the jump distribution, a trivariate structure would be numerically more complicated to estimate compared to the bivariate structure.

⁴Throughout the rest of the paper $i = 1, 2$.

where each of the jump size variables $Y_{it,j}$ is assumed to be independent and normally distributed with mean θ_i and variance δ_i^2 . The jumps, n_{it} , in the market returns are assumed to be generated independently of each other by an independent bivariate Poisson distribution with time-varying jump intensity parameters λ_{it} . The parameter λ_{it} is the expected conditional number of jumps, n_{it} over the time interval $(t-1, t)$, i.e. $\lambda_{it} \equiv E[n_{it} | \Phi_{it-1}]$. The bivariate Poisson density is specified as:

$$\Pr(n_{1t} = k, n_{2t} = l | \Phi_{t-1}) = \frac{\exp(-\lambda_{1t})\lambda_{1t}^k}{k!} \frac{\exp(-\lambda_{2t})\lambda_{2t}^l}{l!}. \quad (3)$$

To allow the jump intensities to vary over time, λ_{it} is specified in an Exponential Autoregressive Jump Intensity (EARJI) form given by:

$$\ln(\lambda_{it}) = \gamma_{0i} + \gamma_{1i} \ln(\lambda_{it-1}) + \gamma_{2i} \xi_{it-1}.$$

In this specification, the parameter γ_{1i} measures the persistence and γ_{2i} measures the possible asymmetric effect of shocks to the jump intensity (λ_{it}). That is, a positive parameter value for γ_{2i} indicates that a positive shock produces a larger impact on the conditional jump intensity than a negative shock of an equal magnitude. The ξ_{it-1} represents the innovation to λ_{it-1} , measured *ex post*. This measurable shock (intensity residual), which is the unpredictable component affecting the jump intensity is given by:

$$\xi_{it-1} = E[n_{it-1} | \Phi_{it-1}] - \lambda_{it-1} = \sum_{\eta=0}^{\infty} \eta \times \Pr(n_{it-1} = \eta | \Phi_{it-1}) - \lambda_{it-1}.$$

$E[n_{it-1} | \Phi_{it-1}]$ is the *ex post* assessment of the expected number of jumps that occurred from $t-2$ to $t-1$, whereas, λ_{it-1} is the conditional *ex ante* expectation of n_{it-1} , given the information set Φ_{it-2} , and $\Pr(n_{it-1} = \eta | \Phi_{it-1})$ is the *ex post* distribution of n_{it-1} , given the information at time $t-1$. Having observed r_{it} and using Bayes' rule, the *ex post* probability that η jumps occurred at time t is given by:

$$\Pr(n_{it} = \eta | \Phi_{it}) = \frac{f(r_{it} | n_{it} = \eta, \Phi_{it-1}) \Pr(n_{it} = \eta | \Phi_{it-1})}{f(r_{it} | \Phi_{it-1})}, \quad (4)$$

$$\eta = 0, 1, 2, \dots$$

Here, $f(r_{it} | n_{it} = \eta, \Phi_{it-1})$ is the marginal conditional density function for r_{it} given that η jumps occurred, $\Pr(n_{it} = \eta | \Phi_{it-1})$ is the marginal Poisson density function for $n_{it} = \eta$ implied by eq.(3), and $f(r_{it} | \Phi_{it-1})$ is the conditional density function for r_{it} . The conditional density function for r_{it} is specified and discussed in Section 2.2.

2.3 Time-varying return correlation

Given that the random disturbances in ϵ_{1t} and the jump innovation components in ϵ_{2t} are contemporaneously independent of each other, i.e. $E(\epsilon_{1it}\epsilon_{2jt}) = 0$ for $i, j = 1, 2$, the covariance matrix of returns may be expressed as:

$$\text{Var}(\mathbf{r}_t | \Phi_{it-1}) = \text{Var}(\epsilon_{1t} | \Phi_{it-1}) + \text{Var}(\epsilon_{2t} | \Phi_{it-1}).$$

The disturbances in ϵ_{1t} are assumed to be normal i.i.d. mean-zero innovations with the conditional covariance matrix:

$$\tilde{\mathbf{H}}_t = \text{Var}(\epsilon_{1t} | \Phi_{it-1}) = \begin{bmatrix} \sigma_{1t}^2 & \sigma_{12t}^{\epsilon_1} \\ \sigma_{12t}^{\epsilon_1} & \sigma_{2t}^2 \end{bmatrix}.$$

The disturbances are specified as $\epsilon_{1it} = \sigma_{it}z_t$, where $z_t \sim N(0, 1)$, and σ_{it} is assumed to follow an EGARCH(1,1) process (Nelson, 1991). The process for the conditional variance is specified as:

$$\ln(\sigma_{it}^2) = \omega_{0i} + \omega_{1i}\psi_{it-1} + \omega_{2i} \ln(\sigma_{it-1}^2) + \omega_{3i} \left(|\psi_{it-1}| - \sqrt{2/\pi} \right),$$

where $\psi_{it} = \epsilon_{1it}/\sigma_{it}$ is the normalized residual. The $\sigma_{12t}^{\epsilon_1}$ is the covariance between ϵ_{11t} and ϵ_{12t} .

Conditional on k jumps in stock market index 1 and l jumps in stock market index 2, the jump innovations in ϵ_{2t} (cf. eq. (2)) have a bivariate normal distribution with zero mean and a covariance matrix given by:

$$\bar{\mathbf{H}}_t = \text{Var}(\epsilon_{2t} | \Phi_{it-1}) = \begin{bmatrix} k\delta_1^2 & \sigma_{12t}^{\epsilon_2} \\ \sigma_{12t}^{\epsilon_2} & l\delta_2^2 \end{bmatrix}.$$

The conditional variances for the jump components are given by $k\delta_1^2$ and $l\delta_2^2$ while the covariance between ε_{21t} and ε_{22t} is denoted with $\sigma_{12t}^{\varepsilon_2}$. Note that $\sigma_{12t}^{\varepsilon_2}$ is thought to capture the possible covariance between either the jump sizes (θ_i) or jump intensities (λ_{it}), or both. In contrast, Chan (2004) lets n_{it} , the frequency of jumps between $t-1$ and t , be correlated through a bivariate Poisson distribution (through trivariate reduction). Hence, the jump intensities, λ_{it} , are allowed to be positively correlated. Chan (2004) assumes that the jump sizes, θ_i , have a constant correlation across contemporaneous equations and are zero across time. An advantage with the approach used in this paper, compared to modelling correlation through underlying parameters, i.e. correlation between σ_{it}^2 , and/or through correlated jump intensities λ_{it} and/or jump sizes θ_i , is that we avoid increasing the numbers of parameters in an already richly parameterized model. Also, the main interest of this paper is on the correlation between market returns rather than between the underlying components.

The covariance matrix for returns is given by:

$$\begin{aligned} \mathbf{H}_t &= \tilde{\mathbf{H}}_t + \bar{\mathbf{H}}_t \\ &= \begin{bmatrix} \sigma_{1t}^2 & \sigma_{12t}^{\varepsilon_1} \\ \sigma_{12t}^{\varepsilon_1} & \sigma_{2t}^2 \end{bmatrix} + \begin{bmatrix} k\delta_1^2 & \sigma_{12t}^{\varepsilon_2} \\ \sigma_{12t}^{\varepsilon_2} & l\delta_2^2 \end{bmatrix} \\ &= \begin{bmatrix} \sigma_{1t}^2 + k\delta_1^2 & \sigma_{12t} \\ \sigma_{12t} & \sigma_{2t}^2 + l\delta_2^2 \end{bmatrix}, \end{aligned} \quad (5)$$

where the covariance between the market returns, $\sigma_{12t} = \sigma_{12t}^{\varepsilon_1} + \sigma_{12t}^{\varepsilon_2}$, is the sum of the covariance between the random disturbances ($\sigma_{12t}^{\varepsilon_1}$) and the jump innovations ($\sigma_{12t}^{\varepsilon_2}$). To derive a model with time-varying return correlation, the covariance for the returns is reparameterized in the spirit of Tsay (2002, ch. 10), as $\sigma_{12t} = \rho_t \sqrt{\sigma_{1t}^2 + k\delta_1^2} \sqrt{\sigma_{2t}^2 + l\delta_2^2}$. Thus, the covariance is replaced by a parameter, ρ_t , for the time-varying correlation times the standard deviation of each return series. To accommodate

that $|\rho_t| < 1$, we use the reparametrization⁵:

$$\rho_t = \frac{\check{\rho}_t}{\sqrt{1 + \check{\rho}_t^2}}.$$

The $\check{\rho}_t$ is parameterized as:

$$\check{\rho}_t = \beta_0 + \beta_1 \varepsilon_{11t-1}^* \varepsilon_{12t-1}^* + \beta_2 \rho_{t-1},$$

where $\varepsilon_{11t-1}^* \varepsilon_{12t-1}^* = \varepsilon_{11t-1} \varepsilon_{12t-1} / \sqrt{\sigma_{1t-1}^2 \sigma_{2t-1}^2}$ and β_2 measures the persistence of the correlation over time. Note here that we use the lagged normal disturbances, ε_{1it-1} , instead of the total residuals, $\varepsilon_{it-1} = \varepsilon_{1it-1} + \varepsilon_{2it-1}$, in this specification. The reason for this is that identification of the jump parameters (θ_i, λ_{it}) is based on the normal disturbances, ε_{1it-1} , cf. eq.(7).⁶

To study the contemporaneous effects of jumps on the return correlations, the *ex post* probability is used in the identification of actual jumps.⁷ Following Maheu and McCurdy (2004), we consider actual jumps to have occurred if the *ex post* probabilities of at least one jump is larger than 0.5, i.e. $\Pr(n_{it} \geq 1 | \Phi_{it}) = 1 - \Pr(n_{it} = 0 | \Phi_{it}) > 0.5$. The identified jumps are then related to the estimated time-varying return correlations using regression analysis.

2.4 Estimation

The probability density function for \mathbf{R}_t , given k independent jumps in stock market index 1 and l independent jumps in stock market index 2 is given by:

$$f(R_t | n_{1t} = k, n_{2t} = l, \Phi_{t-1})$$

⁵Tsay (2002) restricts ρ by a Fisher transformation given by $\rho = (\exp(\tilde{\rho}) - 1) / (\exp(\tilde{\rho}) + 1)$. Baur (2006) reports that the Fisher transformation is more restrictive than the transformation used in this paper and thus less adequate.

⁶Direct testing of the effect of jump innovations on the time-varying return correlation dynamics resulted in unstable models with poor convergence properties.

⁷This approach is similar to that used by Asgharian and Bengtsson (2006).

$$= \frac{1}{2\pi^{N/2}} |\mathbf{D}_{ijt} \boldsymbol{\rho}_t \mathbf{D}_{ijt}|^{-1/2} \exp[-1/2 \boldsymbol{\epsilon}_{1t} \mathbf{D}_{ijt} \boldsymbol{\rho}_t \mathbf{D}_{ijt} \boldsymbol{\epsilon}_{1t}], \quad (6)$$

where

$$\boldsymbol{\epsilon}_{1t} = \mathbf{R}_t - \boldsymbol{\mu}_t - \boldsymbol{\epsilon}_{2t} = \begin{bmatrix} r_{1t} - \mu_{1t} - k\theta_1 + \lambda_{1t}\theta_1 \\ r_{2t} - \mu_{2t} - l\theta_2 + \lambda_{2t}\theta_2 \end{bmatrix}, \quad (7)$$

$$\mathbf{D}_t = \begin{bmatrix} \sqrt{\sigma_{1t}^2 + k\delta_1^2} & 0 \\ 0 & \sqrt{\sigma_{2t}^2 + l\delta_2^2} \end{bmatrix}, \quad (8)$$

and $\boldsymbol{\rho}_t$ is the conditional correlation matrix

$$\boldsymbol{\rho}_t = \begin{bmatrix} 1 & \rho_t \\ \rho_t & 1 \end{bmatrix}.$$

The conditional density of returns is defined by:

$$\begin{aligned} \Pr(\mathbf{R}_t | \Phi_{t-1}) &= \sum_{k=0}^{\infty} \sum_{l=0}^{\infty} f(R_t | n_{1t} = k, n_{2t} = l, \Phi_{t-1}) \\ &\times \Pr(n_{1t} = k, n_{2t} = l | \Phi_{t-1}) \end{aligned} \quad (9)$$

and the corresponding log likelihood function is simply the sum of the log conditional densities:

$$\ln L = \sum_{t=1}^T \ln \Pr(\mathbf{R}_t | \Phi_{t-1}).$$

In practice, the maximum number of jumps may be truncated to a large value τ , so that the probability of τ or more jumps is zero. In the empirical estimation $\tilde{\tau} > \tau$ is investigated to ensure that the likelihood and parameter estimates do not change. In the estimations reported in the results section, $\tau = 15$.

3 Data

The data used in this paper are capitalization weighted daily stock price indices of the Estonian (Tallinn, TALSE), Latvian (Riga, RIGSE)⁸ and

⁸There is an irregularity in the summer of 2001 in the Riga index (RIGSE), due to a power struggle in its largest company (Latvijas Gaze). Instead of modelling this

the Lithuanian (Vilnius, VILSE) stock markets. All prices are expressed in Euro.⁹ The data set, obtained from Datastream, covers January 3, 2000 to July 9, 2007, for a total of $T = 1960$ observations. Due to some differences in holidays for the involved countries, the series have different shares of days for which index stock prices are not observable. Linear interpolation was used to fill the gaps for all series, where resulting series are then throughout for a common trading week. All returns are calculated as $y_t = 100 \cdot \ln(I_t/I_{t-1})$, where I_t is the daily price index. Table 1 reports descriptive statistics and cross correlations for the daily return series. The Ljung-Box statistics for 10 lags (LB_{10}) indicate significant serial correlations. The large kurtoses for Riga, Tallinn, and Vilnius indicate leptokurtic densities. Cross-correlations indicate that the largest unconditional correlation is between Tallinn and Vilnius return series.

4 Empirical results

4.1 Basic models

In the empirical investigation of the correlation structure between the Baltic stock market indices, a number of different model specifications were estimated, including different lag structures for the mean, conditional variance and autoregressive jump intensity functions. Although, in some specifications with more lags, the Akaike Information Criteria

irregular period, the observations from July 25 to September 3, 2001, are replaced by interpolated values in the same way as in Brännäs et al., (2008).

⁹This implies that the analyzed return series also contain variation due to exchange rate movements. Since the paper is written from an international investor perspective, i.e. we are interested in Euro returns, the effects of these variations are included in the analysis. Also, the currencies of the considered countries have been pegged (Latvian to a basket of major currencies since 1994, and to the Euro from 2005; Estonian to the Deutsche Mark since 1992, and later to the Euro after its introduction; Lithuania had a US dollar-based currency board arrangement since 1994, moved to the Euro peg in 2002) under the period of study and have been rather stable during, at least, the later parts of the sample period.

Table 1: Descriptive statistics and unconditional correlations between return series.

Exchange	Mean	Variance	Min/Max	Skewness	Ex. Kurtosis	Riga	Tallinn	Vilnius
Riga	0.09	1.64	-9.27/10.29	0.18	11.29	1		
Tallinn	0.10	1.06	-5.87/12.02	0.66	14.86	0.134	1	
Vilnius	0.09	1.00	-12.12/5.32	-0.95	13.68	0.145	0.242	1

(AIC) and autocorrelation properties were slightly improved, the identification of the parameters in the time-varying correlation function became numerically unstable with more elaborate lag structures. Hence, as the focus of this paper is mainly on the correlation functions, the more simple model specifications were favored and utilized in the analysis. Overall, the EGARCH specification for the conditional variance was favored in terms of AIC compared to corresponding GARCH specifications.

Initially, we considered models without jumps (for the purpose of comparison), i.e. with residuals specified as $\epsilon_t = \mathbf{R}_t - \boldsymbol{\mu}_t$. Table 2 reports on the estimation results for this model specification with time-varying correlations.

The results indicate that the average correlation between Tallinn-Riga, Tallinn-Vilnius, and Vilnius-Riga is 0.120 (s.d. 0.065), 0.217 (s.d. 0.106), and 0.115 (s.d. 0.065), respectively.¹⁰ Note that for each series, we obtain two sets of parameter estimates due to the bivariate structure of the models. That is, for the Riga series, we obtain one set of estimates from the bivariate model with Tallinn, and another set of estimates from the bivariate model for Riga and Vilnius. However, the estimates for the same series do not differ much between the models. Figure 1 displays the time-varying correlations. Notably, there is a number of sharp spikes, both positive and negative, in the time-varying correlations for all considered indices, possibly due to market jumps. This is most pronounced for the time-varying correlation between the Tallinn and Vilnius stock market returns. The persistence in the time-varying correlations are high, as indicated by the significant lagged correlation parameters (β_2 in Table 2) that takes on values above 0.9 for all the models.

¹⁰Constant correlation models without jumps gave similar parameter estimates and correlations close to the mean of the time-varying correlations.

Table 2: Estimation results for time-varying return correlation models without jump component (robust standard errors in parantheses).

	Tallinn	Riga	Tallinn	Vilnius	Vilnius	Riga
	$\mu_{it} = \alpha_{0i} + \alpha_{1i}r_{it-1}$					
α_0	0.054* (0.018)	0.158* (0.024)	0.056* (0.018)	0.123* (0.020)	0.134* (0.020)	0.174* (0.023)
α_1	0.113* (0.008)	-0.061* (0.025)	0.109* (0.004)	0.093* (0.025)	0.104* (0.026)	-0.062* (0.014)
	$\ln(\sigma_{1it}^2) = \omega_{0i} + \omega_{1i}\psi_{it-1} + \omega_{2i}\ln(\sigma_{1it-1}^2) + \omega_{3i}(\psi_{it-1} - \sqrt{2/\pi})$					
ω_0	0.017* (0.003)	0.097* (0.009)	0.017* (0.002)	0.007 (0.007)	0.001 (0.007)	0.094* (0.008)
ω_1	0.002 (0.007)	-0.024 (0.015)	0.004 (0.006)	-0.076* (0.012)	-0.076* (0.013)	-0.020 (0.013)
ω_2	0.974* (0.004)	0.800* (0.012)	0.975* (0.003)	0.864* (0.014)	0.853* (0.016)	0.807* (0.010)
ω_3	0.244* (0.012)	0.411* (0.025)	0.241* (0.012)	0.386* (0.027)	0.385* (0.028)	0.412* (0.022)

Table 2 continued

	Tallinn	Riga	Tallinn	Vilnius	Vilnius	Riga
			$\rho_t = \check{\rho}_t / \sqrt{1 + \check{\rho}_t^2}, \check{\rho}_t = \beta_0 + \beta_1 \varepsilon_{1t-1} \varepsilon_{2t-1} + \beta_2 \check{\rho}_{t-1}$			
β_0	0.003 (0.006)		0.013* (0.005)		0.001 (0.001)	
β_1	-0.000 (0.002)		0.028* (0.008)		0.005* (0.003)	
β_2	0.978* (0.049)		0.915* (0.028)		0.980* (0.014)	
$\bar{\rho}$	0.120 (0.065)		0.217 (0.106)		0.115 (0.065)	
Log-L	-5572.18		-5117.01		-5639.70	
AIC	11.154		10.252		11.297	
LB ₁₀	16.243	20.325	17.803	35.132	36.078	23.157
LB ₁₀ ²	15.821	78.653	16.337	29.321	32.223	83.728

* Significant at the 5 percent level. $\bar{\rho}$ = mean of the time-varying correlations.

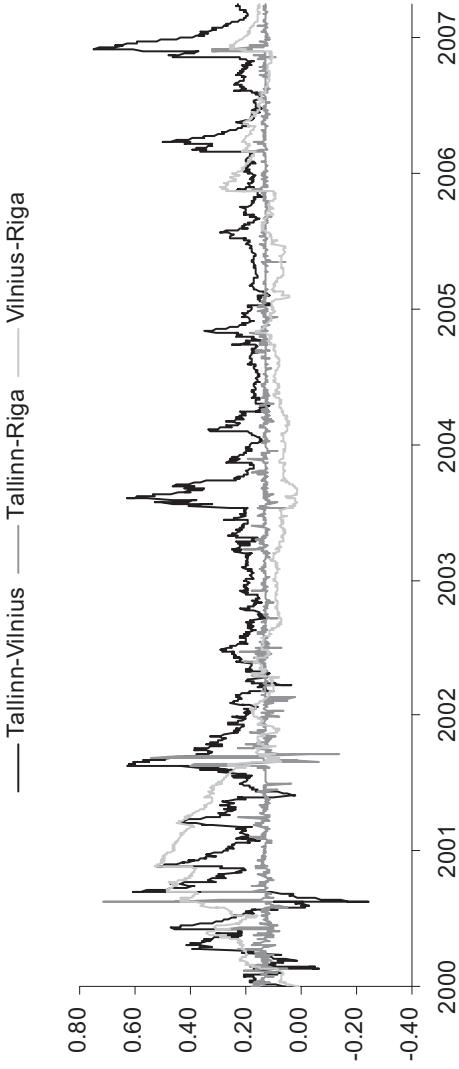


Figure 1: Time-varying return correlations.

Table 3 report estimates for constant correlation models including the jump component. Including a jump component in the models notably improves the AIC, compared to the models with constant correlation and no jump component.¹¹ It is worth noting that the estimated mean jump sizes (θ_i) are small, and significant only for Riga in the bivariate model for the Vilnius and Riga series, as the estimated standard deviations are, in general, quite large. However, the estimated jump parameters (θ_i, λ_{it}) are jointly significant, as indicated by a LR test, when comparing with models with no jumps and constant correlations.¹² The parameter estimates for the conditional mean jump intensities (λ_t) indicate that the persistence in jump intensity is high (and statistically significant) both for Riga (0.974, 0.978) and Vilnius (0.986, 0.991), while it is lower for Tallinn (0.481, 0.485). The inclusion of the jump component in the models also removed some of the autocorrelations present in the models without a jump component, as indicated by the Ljung-box statistics (LB_{10} and LB_{10}^2). However, there is little autocorrelation remaining in the final models.¹³

Table 4 reports estimates for the time-varying correlation models including the jump components. Since the parameter estimates for the mean, EGARCH, and jump components are similar to that reported in Table 3, only the parameters pertaining to the specification of the time-varying correlation are reported.

For the model specification with the time-varying correlation, the AIC improves slightly for all models. However, LR tests indicate that there are doubts about whether including time-varying correlations improve the model fit for the bivariate model for Tallinn and Riga series. The

¹¹The estimation results available from authors upon request.

¹²The LR test statistics are 951, 639 and 855 for the Tallinn-Riga, Tallinn-Vilnius and Vilnius-Riga sample, respectively.

¹³Other lag structures for the mean, conditional variance, and the autoregressive jump intensity have been tried without fully removing the autocorrelations. The more parsimonious lag structures, reported in the paper, were therefore chosen.

Table 3: Estimation results for models including jump components and constant correlation (robust standard errors in parentheses).

	Tallinn	Riga	Tallinn	Vilnius	Vilnius	Riga
	$\mu_{it} = \alpha_{0i} + \alpha_{1i}r_{it-1} + \varepsilon_{2it}$					
α_0	0.056* (0.016)	0.070* (0.020)	0.065* (0.014)	0.104* (0.018)	0.102* (0.012)	0.075* (0.014)
α_1	0.197* (0.022)	-0.084* (0.020)	0.186* (0.021)	0.120* (0.022)	0.130* (0.015)	-0.078* (0.019)
	$\ln(\sigma_{1it}^2) = \omega_{0i} + \omega_{1i}\psi_{it-1} + \omega_{2i}\ln(\sigma_{1it-1}^2) + \omega_{3i}\left(\psi_{it-1} - \sqrt{2/\pi}\right)$					
ω_0	-0.064* (0.014)	-0.258* (0.060)	-0.069* (0.016)	-0.308* (0.111)	-0.406* (0.113)	-0.238* (0.020)
ω_1	-0.027 (0.016)	0.029 (0.025)	-0.013 (0.016)	-0.040 (0.026)	-0.026 (0.029)	0.031 (0.018)
ω_2	0.982* (0.008)	0.715* (0.062)	0.974* (0.008)	0.733* (0.086)	0.679* (0.080)	0.732* (0.021)
ω_3	0.198* (0.029)	0.342* (0.042)	0.204* (0.032)	0.317* (0.048)	0.336* (0.048)	0.341* (0.029)
	$\ln(\lambda_{it}) = \gamma_{0i} + \gamma_{1i}\ln(\lambda_{it-1}) + \gamma_{2i}\xi_{it-1}$					
γ_0	-0.705* (0.361)	-0.039 (0.025)	-0.760 (0.426)	-0.012 (0.010)	-0.014 (0.011)	-0.048* (0.021)
γ_1	0.485* (0.204)	0.978* (0.015)	0.481* (0.241)	0.991* (0.007)	0.986* (0.009)	0.974* (0.012)
γ_2	1.264* (0.256)	0.552* (0.166)	1.148* (0.284)	0.365* (0.100)	0.373* (0.073)	0.587* (0.050)

Table 3 continued

	Tallinn	Rīga	Tallinn	Vilnius	Vilnius	Rīga
θ	0.063 (0.079)	0.120 (0.101)	0.076 (0.075)	0.054 (0.067)	0.048 (0.054)	0.140* (0.067)
δ	1.203* (0.058)	1.867* (0.068)	1.261* (0.063)	1.197* (0.067)	1.106* (0.055)	1.966* (0.029)
ρ	0.133* (0.024)		0.236* (0.024)		0.187* (0.024)	
Log-L	-5096.69		-4804.70		-5219.01	
AIC	10.217		9.633		10.462	
LB ₁₀	5.931	10.123	6.375	14.278	20.451	12.011
LB ₁₀ ²	8.968	31.136	10.995	11.280	20.662	30.576

* Significant at the 5 percent level.

Table 4: Estimation results for models including jump components and time-varying return correlations (robust standard errors in parantheses).

	Tallinn (1) - Riga (2)	Tallinn (1) - Vilnius (2)	Vilnius (1) - Riga (2)
	$\check{\rho}_t = \beta_0 + \beta_1 \epsilon_{11t-1}^* \epsilon_{12t-1}^* + \beta_2 \rho_{t-1}$		
β_0	0.069* (0.029)	0.013 (0.007)	0.001 (0.001)
β_1	0.014 (0.008)	0.025* (0.009)	0.008* (0.002)
β_2	0.968* (0.201)	0.924* (0.034)	0.987* (0.004)
Log-L	-5095.29	-4793.32	-5210.01
AIC	10.213	9.609	10.448

* Significant at the 5 percent level.

LR test value is 2.7832, 22.7752, and 21.2336 for the model with Tallinn and Riga, Tallinn and Vilnius, and Vilnius and Riga series, respectively. The persistence parameter for the time-varying correlation specification is quite high and ranges between 0.924 and 0.987. A number of different specifications for the time-varying correlation was tried during estimation. None of these specifications, including jump residuals, $\hat{\epsilon}_{2t-1} = E[\epsilon_{2t-1} | \Phi_{it-1}]$, the *ex post* assessment of the expected number of jumps $E[n_{it-1} | \Phi_{it-1}]$, as well as the lagged conditional variance, σ_{t-1}^2 , improved the fit of the model and mostly rendered numerically unstable models. Hence, to study the effect of market jumps on the time-varying return correlations, we instead turn our attention towards the identification of actual jumps.

4.2 The effect of jumps on time-varying return correlations

Actual jumps are determined to have occurred if the *ex post* probability of at least one jump is larger than 0.5, i.e. $\Pr(n_{it} \geq 1 | \Phi_{it}) =$

$1 - \Pr(n_{it} = 0 | \Phi_{it}) > 0.5$.¹⁴ The *ex post* jump probabilities during 2006-2007 for the Riga stock market are displayed in Figure 2 along with the daily return series. Over this period, 23 return observations are determined to be jumps according to the chosen criteria. Using the above criterion to identify actual jumps, we find that there are 270 and 95 jumps for the bivariate model for Tallinn and Riga, 195 and 296 jumps for the model of Tallinn and Vilnius, and 353 and 95 jumps for the Vilnius and Riga model.¹⁵ Of these, there are 12, 33, and 21 simultaneous jumps in the three corresponding bivariate models. Based on the signs of the return series, we determine that there are 2, 9, and 8 simultaneous negative jumps while there are 5, 10, and 4 simultaneous positive jumps for the models for Tallinn and Riga, Tallinn and Vilnius, and Vilnius and Riga series. Thus, on a number of occasions, there are simultaneous jumps in opposite directions.

To examine the impact of the identified jumps on the time-varying return correlations, we run different linear regression models for the estimated time-varying correlation, $\hat{\rho}_{ijt}$, on a number of dummy variables that reflect the number of jumps in the series. The dummy variables for individual jumps in each series are $djump_{it}$, $djump_{it}^+$ and $djump_{it}^-$, which take on a value of one if there is a jump at time t in series i , a

¹⁴To study the sensitivity of the results to the chosen criteria the analysis was repeated with $\Pr(n_{it} \geq 1 | \Phi_{it}) = 1 - \Pr(n_{it} = 0 | \Phi_{it}) > 0.7$. This specification did not change the results.

¹⁵As mentioned before, due to the bivariate structure of the model, we obtain two estimated series of jump probabilities (and series of identified jumps) for each return series. For example, for Tallinn we identify one series of jump probabilities based on the bivariate model with Riga and another series based on the bivariate model of Tallinn and Vilnius. The correlations (Spearman's rho) between the two identified series of jump probabilities for each return series are 0.97, 0.99, and 0.99 for Tallinn, Riga and Vilnius, respectively. The Spearman's rho for the actually identified jump series are 0.83, 0.96, and 0.86 for the Tallinn, Riga, and Vilnius series. This indicates that the level of the identified jump probabilities differs to some degree depending on the combination of the series in the model. This also explains the difference in the number of actual identified jumps (depending on combination) for the same series.

positive jump in series i , or a negative jump in series i , respectively, and zero otherwise. The dummy variable controlling for simultaneous jumps are $dsimjump_{ijt}$, $dsimjump_{ijt}^+$, $dsimjump_{ijt}^-$, and $dsimjump_{ijt}^{+/-}$, which take on a value of one if there are simultaneous jumps in series i and j , simultaneous positive jumps in series i and j , and simultaneous negative jumps in series i and j , and simultaneous jumps of opposite sign in series i and j , respectively, and zero otherwise.¹⁶ Table 5 report estimation results for the model specification with the highest adjusted R^2 .

From Table 5 we see that a jump in one series, controlling for simultaneous jumps, contributes positively (when significant) to the time-varying correlations. On average, the time-varying correlation between Tallinn and Vilnius series increases with 0.007 when there are positive jumps on the stock market in Tallinn. Both positive as well as negative jumps on the Riga stock market increase the correlation between Tallinn and Riga by on average 0.015. Negative jumps on the Vilnius stock exchange on average increase the correlation between both the Tallinn and Vilnius (with 0.017), as well as the Vilnius and Riga (with 0.004) series. Overall, the effects of the individual jumps, when controlling for simultaneous jumps, is rather small. For example, the correlation between the Tallinn-Vilnius return series increase on average from 0.228 to 0.245 when there are isolated jumps on the Vilnius stock market.

¹⁶The lagged correlation $\hat{\rho}_{t-1}$ is also included in all regressions to control for serial correlation.

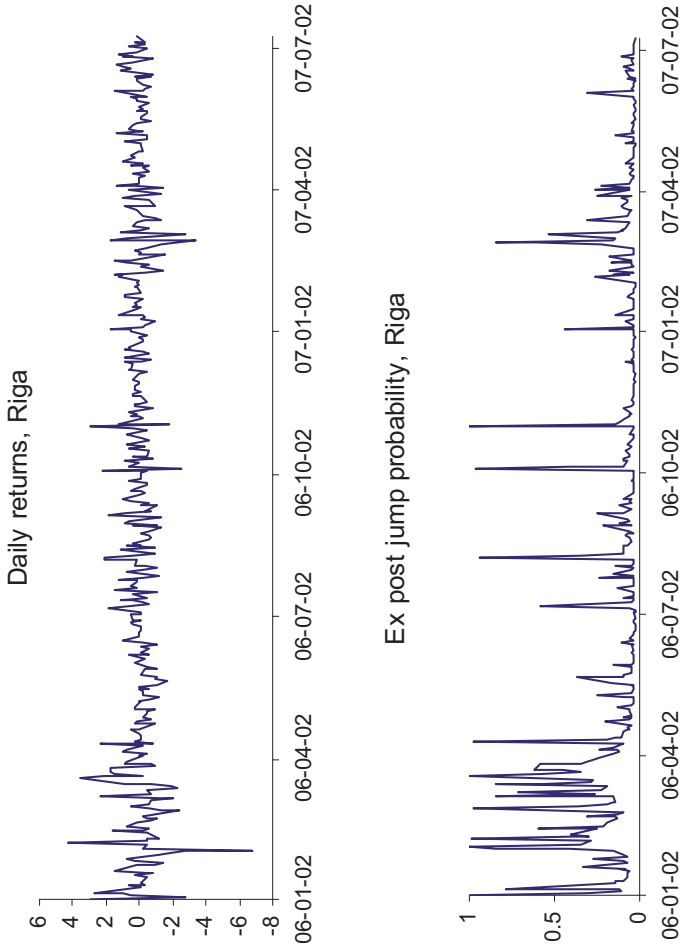


Figure 2: Return and *ex post* jump probability, Riga.

For simultaneous jumps in the series, the effect on the time-varying correlations depends on the direction of the jumps. For example, the time-varying correlation increases on average with 0.129 for simultaneous positive jumps, and with 0.061 for simultaneous negative jumps in the Tallinn-Riga model. Thus, the average correlation almost doubles (compared to the model with constant correlation) on days when there are simultaneous positive jumps. Notably the impact on the correlation between the Tallinn and Riga series is much larger when markets are jointly rising, compared to when markets are jointly falling. The opposite is true for Tallinn-Vilnius and Vilnius-Riga models, where the correlation increases on average with 0.052 and 0.017 for simultaneous positive jumps, and 0.083 and 0.026 for simultaneous negative jumps. These changes correspond to a correlation increase ranging from 11 to 34 percent. Note that these results could be related to the contagion literature, where positive contagion is defined as an increase in the correlation caused by positive shocks, while an increase in the correlation due to negative shocks is usually referred to as negative contagion (Baur and Fry, 2005).

On the occasions when there are simultaneous jumps in opposite directions, the correlations decreases on average with 0.077, 0.094, and 0.033 for the Tallinn-Riga, Tallinn-Vilnius, and Vilnius-Riga correlation series, respectively. Including $d\text{simjump}_{ijt}$, i.e. controlling for simultaneous jumps with no regard to the direction of the jumps, in general, yields an insignificant impact on the time-varying correlations.

5 Concluding remarks

The results of this paper show a strong support for models including a jump component (compared to the EGARCH-alternatives), as well as support for time-varying return correlations over constant correlation models. In general, the number of identified jumps during the period

Table 5: Effect of identified jumps on time-varying return correlations (robust standard errors in parantheses).

	Tallinn (1) - Riga (2)	Tallinn (1) - Vilnius (2)	Vilnius (1) - Riga (2)
<i>const</i>	0.051* (0.002)	0.008* (0.001)	0.001 (0.000)
$\hat{\rho}_{t-1}$	0.908* (0.017)	0.959* (0.005)	0.994* (0.002)
$djump_{1t}^+$	0.001 (0.002)	0.007* (0.002)	0.001 (0.001)
$djump_{1t}^-$	0.004 (0.002)	-0.002 (0.003)	0.004* (0.001)
$djump_{2t}^+$	0.015* (0.003)	0.004 (0.002)	0.000 (0.001)
$djump_{2t}^-$	0.015* (0.004)	0.017* (0.002)	0.001 (0.002)
$dsimjump_t^+$	0.129* (0.011)	0.052* (0.008)	0.017* (0.005)
$dsimjump_t^-$	0.061* (0.024)	0.083* (0.009)	0.026* (0.004)
$dsimjump_t^{+/-}$	-0.077* (0.011)	-0.094* (0.007)	-0.033* (0.003)
R^2	0.446	0.949	0.993
DW	1.961	1.899	1.966

* Significant at the 5 percent level.

may seem large, at least, compared to results for developed markets. For example, Bollerslev et al., (2008) find on average 7 major jumps in equity market indices for a number of developed countries during the period 2001-2005. Kim and Mei (2001), however, report 71 identified price jumps for the Hong Kong stock market during 1989-1993. Thus, our results are in line with the idea that the emerging stock markets are, in general, more volatile and have empirical return distributions with fatter tails than more developed markets (e.g., Harvey, 1995; Bekaert and Harvey, 2002). A possible explanation to the large number of identified jumps is that, the markets under study are relatively small with a few large institutional traders active on all three markets. Thus, a number of these jumps may be driven by liquidity motivated trading.

The time-varying return correlations increase slightly when there are individual market jumps (i.e. conditional on being non-simultaneous jumps) for some of the markets. For simultaneous jumps, we find that the effect of these on the return correlations depend on the jump signs. This is particularly important to keep in mind when studying jump correlations (e.g., Chan, 2004; Asgharian and Bengtsson, 2006), as a positive jump correlation, i.e. the correlation between jump intensities with no regard to the sign of a jump, often is taken as a sign of increasing return correlations. This becomes even more important for emerging markets, where more jumps in both directions could be expected. In this paper, we find that on average 58 percent of the simultaneous jumps (over all samples) are of the same sign, and as many as 42 percent are of opposite sign. In addition, we find that the correlation increases by as much as 100 percent on average due to simultaneous positive jumps (for the Tallinn-Riga model), but by 47 percent due to simultaneous negative jumps.¹⁷

Overall, we find that stock market return correlations increase mainly

¹⁷However, since the results are based on a few observations, as there is only a small number of simultaneous jumps, conclusions should be interpreted with some caution.

due to simultaneous market jumps, that may depend on other factors than market crises, while individual (non-simultaneous jumps) only have small effects. The underlying model could also be of use for studies of the correlation between stock and bond markets, thus, of the so called flight-to-quality effect. For example, if there is a negative jump in the stock market together with a decrease in the correlation coefficient, this may indicate the flight-to-quality from stocks to bonds. Similar patterns on two stock markets are harder to interpret, as investor's preferences could also be affected by the liquidity on the markets. However, studying the possible impact of market jumps on return correlation dynamics and, in particular, how these effects may differ between financial assets and markets, is useful for risk management and portfolio diversification.

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V

Financial Intermediation and Economic Growth: Evidence from the Baltic countries*

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Abstract

The hypothesis that financial development promotes economic growth is largely supported by empirical studies. This hypothesis is tested for the three Baltic countries using a time series approach that allows for interactions between the three countries. We find that economic growth is a positive function of financial development, proxied by banking credit available to private sector, in the long run. The results also show that there are long run interactions between the three Baltic countries.

Key Words: Cointegration; Spillovers; Financial development; Emerging markets

JEL Classification: O16; C32; F43

*The comments of Thomas Aronsson and Ulf Holmberg are gratefully acknowledged.

1 Introduction

The relationship between financial development and economic growth has essentially become a commonly accepted fact. In general, there are several channels through which financial development can affect economic growth (for a survey see, e.g., Pagano 1993; Levine, 1997). First, the financial sector promotes accumulation of capital, which is an important condition for economic growth. In practice, this means that a more efficient financial system reduces the loss of resources required to allocate capital, i.e. lowers the transaction costs. Second, along with its effect on capital accumulation, there are a number of channels through which financial development can raise the productivity of capital, i.e. contribute to technological progress. These channels are related to the function of financial intermediaries: (i) to evaluate and select the most profitable investment projects; (ii) to provide liquidity, which creates incentives to invest a larger share of savings in more profitable long term projects; and (iii) to provide a possibility for portfolio diversification, which allows individual agents to undertake riskier and more specialized investment projects.

Even though, the existence of a finance-growth relationship is generally recognized, the empirical results vary considerably across countries, depending on the institutional characteristics, market size, and the level of initial development (e.g., Rousseau and Wachtel, 1998; Fink et al., 2005).¹ The findings on the contribution of financial developments to economic growth in Central and Eastern Europe (CEE) are ambiguous (cf. Fink et al., 2009, for a survey). However, many previous studies used cross-country analysis that don't necessarily take into account the different country characteristics. According to Arestis et al. (2001), time series methods can provide useful insights into the differ-

¹The results on the finance-growth relationship may also vary due to estimation methods, variables, and data sets for different periods and countries; see, e.g., Levine (1997), Thiel (2001), and Beck (2008) for a review of the earlier literature.

ences in the finance-growth relationship across individual countries and may highlight important details that are hidden in averaged-out results from cross-country regressions. Also, Rousseau and Wachtel (2005) call for more studies on individual countries' experiences to gain insight into the role financial development plays for economic growth.

The purpose of this study is to contribute to the empirical evidence on the finance-growth relationship in the CEE countries. In particular, this study focuses on the three Baltic states that have received little attention in the previous literature; this being despite the fact that they experienced a period of high economic growth and rapid credit expansion facilitated primarily by foreign-owned banks during the several years prior to the financial crises of 2007-2008. We utilize a time-series approach to examine the relationship between financial sector development, proxied by the level of bank credit to the private sector, and economic growth in the Baltic countries over the period 1995-2008. In addition, since all three Baltic countries are likely to be closely interrelated, through, for example, trade between the countries and exposure to the same (primarily foreign-owned) banks, we allow for cross-country (i.e. cross-sectional) dependence in the empirical analysis.

2 Econometric Method and Data

Following Arestis et al. (2001) the empirical investigation is carried out using the Johansen method, see Johansen (1988, 1995). It is based on the vector error correction representation of a VAR(p) model:

$$\Delta \mathbf{X}_t = \Pi \mathbf{X}_{t-1} + \sum_{i=1}^{p-1} \Gamma_i \Delta \mathbf{X}_{t-i} + \Psi_t \mathbf{D}_t + \varepsilon_t, \quad (1)$$

where \mathbf{X}_t is an $n \times 1$ vector of $\mathbf{I}(1)$ variables, Δ is the first difference operator, Γ_i for $i = 1, \dots, p-1$, is an $n \times n$ parameter matrix, and \mathbf{D}_t is a set of $\mathbf{I}(0)$ deterministic variables such as a constant and seasonal

dummies. The ε_t is a vector of i.i.d. errors with zero mean and constant variance. The variables employed in the empirical analysis are measured as in Arestis et al. (2001). That is, real economic growth is measured by the logarithm of real GDP ($\ln Y$). The banking system's development, used as a proxy for financial development or financial depth, is measured by a logarithm of the ratio of commercial bank credit to the private sector to nominal GDP ($\ln C$).² We employ quarterly data on output and indicators for credit growth for Latvia, Estonia, and Lithuania during 1995Q1 -2008Q1.³ Seasonality in data is dealt with using (centered) seasonal dummies.⁴

Also, evidence from the panel unit root/cointegration literature suggests that ignoring cross-sectional dependence can have a major influence on the statistical properties of the estimators and test statistics, see, e.g., Breitung and Pesaran (2008). Hence, we allow for cross-country (i.e. cross-sectional) dependence. In the empirical analysis the three countries are modelled jointly, i.e. $\mathbf{X}_t = [\ln Y_t^{Lat}, \ln Y_t^{Est}, \ln Y_t^{Lit}, \ln C_t^{Lat}, \ln C_t^{Est}, \ln C_t^{Lit}]'$, where the superscripts stand for Latvia, Estonia, and Lithuania, respectively.

The Johansen (1988) test procedure can be used to test the null hypothesis of r cointegration relations. If $r = 0$, then $\mathbf{\Pi} = \mathbf{0}$, and the variables are not cointegrated. If there exists r cointegration relations,

²Several other indicators of financial sector depth have been utilized in the previous literature, including, for instance, deposit based measures that are primarily applicable for countries in their first stage of development (e.g., Hondroyannis et al., 2005). Credit-based variables are chosen in this study since in many emerging countries, including the Baltic states, the banking sector is often the only provider of financial intermediation, in contrast to the developed economies that have a wide range of market oriented institutions (e.g., Wachtel, 2003).

³In view of the small data sample, other explanatory variables are not included to save degrees of freedom.

⁴For instance, let $d_{j,t}$ be a centered seasonal dummy, then it takes on the value of 0.75 in quarter j and -0.25 in the other quarters, and has therefore mean zero over a full year.

$0 < r < n$, it implies that $\mathbf{\Pi}$ is rank-deficient and can be decomposed into two matrices, $\boldsymbol{\alpha}$ ($n \times r$) and $\boldsymbol{\beta}$ ($r \times n$), such that $\mathbf{\Pi} = \boldsymbol{\alpha}\boldsymbol{\beta}'$. Equation (1) can be rewritten as:

$$\Delta \mathbf{X}_t = \boldsymbol{\alpha}\boldsymbol{\beta}'\mathbf{X}_{t-1} + \sum_{i=1}^{p-1} \boldsymbol{\Gamma}_i \Delta \mathbf{X}_{t-i} + \boldsymbol{\Psi}_t \mathbf{D}_t + \boldsymbol{\varepsilon}_t, \quad (2)$$

where the rows of $\boldsymbol{\beta}$ can be interpreted as the distinct cointegrating vectors, i.e. the long run relationships between the variables in \mathbf{X}_t . The coefficients in the $\boldsymbol{\alpha}$ matrix indicate the speed of adjustment toward the long run equilibrium. Finally, if $r = n$ then $\mathbf{\Pi}$ is of full rank, and the variables in \mathbf{X}_t are stationary, i.e. $\mathbf{I}(0)$.

Using the above framework we test also whether there are any causal flows in the long run relationship between the variables in \mathbf{X}_t . A test of zero restrictions on the $\boldsymbol{\alpha}$ in Equation (2) corresponds to a test of weak exogeneity, which in a cointegrated systems equals the long run causality.⁵ Due to the small data sample a bootstrap procedure is utilized for inference purposes.⁶

3 Results

The examination of the long run relationship between the variables is carried out in several steps. First, since a VAR framework depends on the time series characteristics of the data set, we test for the presence of a unit root. The Augmented Dickey-Fuller (ADF) tests suggest that all variables are $I(1)$. Second, a VAR (4) model in levels is estimated and Schwarz Information Criterion (SC) is used to decide upon lag length. In order to allow for any deterministic seasonality, centered quarterly

⁵The null hypothesis of $\alpha_{ij} = 0$ can be tested by an LR test which follows a standard χ^2 -distribution in large samples. If the null hypothesis of $\alpha_{ij} = 0$ is rejected, then there is long run causality, see, e.g., Granger and Lin (1995).

⁶See, e.g., Li and Maddala (1997), who demonstrate that the bootstrap can provide significant improvement, as the cointegration test has poor small sample properties.

dummies are included throughout the estimation. SC indicates that the second-order VAR model (in levels) is appropriate, which implies one lag in differences. A visual inspection of the autocorrelation functions of the estimated VAR residuals show no remaining serial correlation, indicating the adequacy of the VAR lag length.

Second, the existence and number of cointegration vectors is tested using the Johansen maximum likelihood approach. The results of the sequential likelihood ratio (LR) tests are presented in the Table 1, where recursive bootstrap p -values (as in Li and Maddala, 1997) are displayed. The LR test results and the bootstrap p -values indicate that there exist three cointegrating vectors.

Table 1: Test statistics and cointegration results (the p -values are the recursive bootstrap p -values as in Li and Maddala, 1997).

Null	Alternative	LR	p -value
$r = 0$	$r = 6$	220.75*	<0.001
$r \leq 1$	$r = 6$	143.34*	<0.001
$r \leq 2$	$r = 6$	80.42	0.003
$r \leq 3$	$r = 6$	38.44	0.128
$r \leq 4$	$r = 6$	17.46	0.193
$r \leq 5$	$r = 6$	23.29	0.376

r indicates the number of cointegrating vectors.

* indicates rejection of the null hypothesis at 5% level.

Next, we estimate the long run relationship in all three countries simultaneously. In order to find the long run relationship, each cointegrating vector is normalized on the economic growth variables ($\ln Y$). As noted earlier, during the modelling procedure we allow for cross-country effects in the long run relationship. However, we exclude parameters that are close to zero and not significantly different from zero. The final

specification for the β matrix, as favored by an LR test is presented in Equation (3) below.⁷ The LR test is χ^2 distributed with 3 df with the bootstrapped p -value of 0.185.

$$\mathbf{\Pi} = \boldsymbol{\alpha}\boldsymbol{\beta}' = \begin{pmatrix} -0.90 & 1.01 & -0.57 \\ (0.08) & (0.12) & (0.06) \\ 0 & 0 & 0 \\ -0.48 & 0.94 & -0.49 \\ (0.14) & (0.19) & (0.09) \\ 0 & 1.64 & 0 \\ & (0.28) & \\ 0.56 & 0 & 0 \\ (0.10) & & \\ 0 & 0 & 0 \end{pmatrix} \begin{pmatrix} 1 & 0 & 0 \\ 0 & 1 & 0 \\ 0 & 0 & 1 \\ -0.32 & -0.22 & 0 \\ (0.02) & (0.01) & \\ -0.30 & -0.15 & 0 \\ (0.03) & (0.02) & \\ 0.26 & 0 & -0.44 \\ (0.02) & & (0.03) \end{pmatrix}' \quad (3)$$

Finally, given the estimated β matrix we evaluate the causal relationship between the financial development and economic growth by weak exogeneity tests. The results are reported in Table 2. Note that for each country, we test the weak exogeneity of economic growth (financial development) in all three countries. For example, if the null hypothesis of weak exogeneity of $\ln Y$ is rejected for Latvia, then financial development in all three Baltic countries affects the economic growth in Latvia in the long run. The bootstrap p -values for the weak exogeneity test of $\ln C$ for Latvia and Estonia are 0.02, indicating that some of the parameters may be zero. Hence, in the final specification we exclude parameters that are not significantly different from zero. The final specification (without presenting deterministic terms, short run dynamics and error terms), as favored by the LR test is displayed in Equation (3). The joint LR test of all the restrictions on $\boldsymbol{\alpha}$ yields a test statistic of 29.53, which is asymptotically χ^2 with 10 df. The bootstrap p -value is 0.113, which indicates that the restrictions cannot be rejected.

For the long run relationship, the estimation results yield negative

⁷Other model specifications, including the block diagonal β , were tested, but rejected using the asymptotic and the bootstrapped p -values.

Table 2: Results of weak exogeneity tests (the p -values are the recursive bootstrap p -values as in Li and Maddala, 1997).

Null hypothesis:	ln Y is weakly exogenous		ln C is weakly exogenous	
	LR	p -value	LR	p -value
Latvia	34.88*	0.01	20.79*	0.02
Estonia	1.60	0.78	16.42	0.02
Lithuania	24.07*	<0.001	12.80*	0.06

* indicates rejection of the null hypothesis at 5% level.

parameter estimates of the diagonal elements in the β matrix (i.e. β_{11} , β_{22} and β_{33}). This is consistent with the theoretical reasoning and several empirical studies (e.g., Arestis et al., 2001; Hondroyannis et al., 2005) that found a positive long run relationship between economic growth and the level of financial intermediation. Also, the results indicate that there is some long run interaction between the three Baltic countries, as indicated by the off-diagonal elements in the β matrix. However, the long run relationship in Lithuania seems to be autonomous with respect to two other Baltic countries. This fact is supported by the findings of Fadejeva and Melihovs (2008) who find that even though Baltic countries share a common pattern in GDP growth, Lithuania exhibits some discrepancies in the economic structure. Economic growth in Estonia is, on the other side, affected by the credit development in the own country as well as in Latvia in the long run. Economic growth in Latvia is determined in the long run by the level of credit in all three countries.

Turning to weak exogeneity tests, our results show that the null hypothesis of weak exogeneity of economic growth (ln Y) in Latvia and Lithuania is rejected at the 5 percent significance level. This means that the level of credit in all three Baltic states affects economic growth in the

long run in those two countries. The results are consistent with several previous studies that found evidence for the "finance causes growth" view (e.g., Calderon and Liu, 2003; Kenourgios and Samitas, 2007; Caporale et al., 2009). However, in Estonia, economic growth is not affected by the credit growth in the three Baltic countries (i.e. is determined outside the model). Our findings of no causality from credit growth to the economic growth in Estonia could, in part, depend on the fact that compared to Latvia and Lithuania. the credit to private sector consisted to a greater extent of household lending (primarily mortgages) rather than resources allocated to productive investments (see Caporale et al., 2009, for more details).

For the weak exogeneity of credit variables, the null hypothesis of weak exogeneity of financial intermediation cannot be rejected at the 5 percent significance level for Lithuania. Thus, the results suggest that economic growth in the Baltic States does not cause financial development in Lithuania in the long run. In other words, for Lithuania, we find only a unidirectional causal relationship from financial development ($\ln C$) to economic growth ($\ln Y$) in the long run.

Next, in order to illustrate the effect of the economic integration (i.e. spillovers) between the three Baltic countries and short run dynamics in more detail, we consider impulse responses. In Figure 1, some selected impulse responses based on Equation (3) are displayed. These are responses to non-factorized one unit standard deviations shock in economic growth, $\ln Y$, in Estonia. According to the impulse responses, economic growth in Latvia and Lithuania reacts positively to a shock in economic growth in Estonia. This could in part depend on the trade pattern between the countries. For instance, by the end of 2008, about 17 percent of the Estonian exports was to the other two Baltic countries. The corresponding number for Lithuania was 19 percent, whereas for Latvia about 30 percent of exports was to the two neighboring countries. Also the credit development in Latvia and Lithuania responds

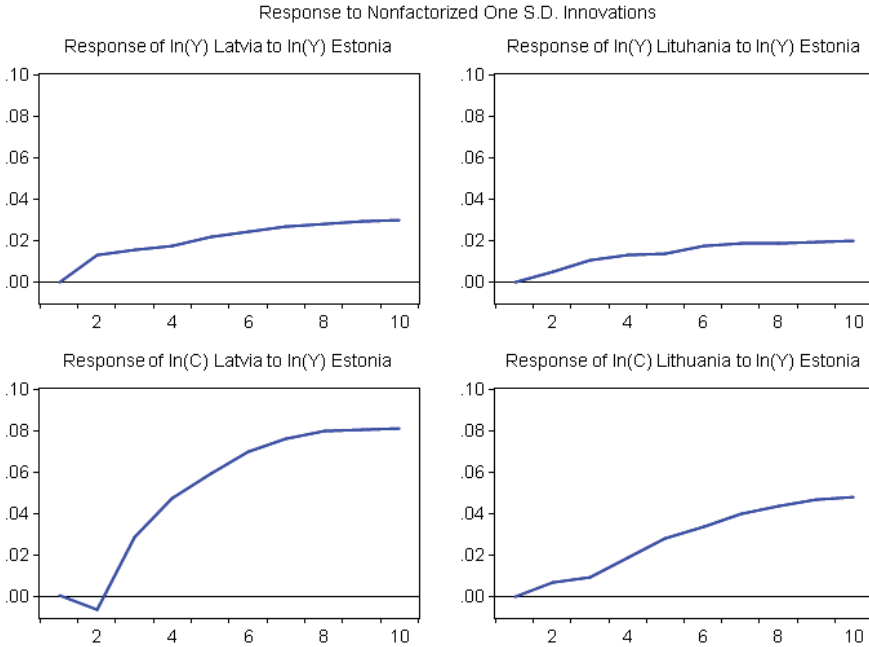


Figure 1: Impulse responses to one unit standard deviation shock in economic growth in Estonia.

positively to a shock in economic growth in Estonia, where the impact on credit development in Latvia is the largest. This can be explained by the fact that foreign-owned banks reallocate capital over different geographical regions on the basis of expected returns and risks. Since economic growth in Latvia and Lithuania reacts positively to a shock in economic growth in Estonia, such a shock may induce the subsidiaries of foreign banks to expand their activity in the other two countries as well.

4 Concluding Remarks

Previously, several studies found that there is no strong relationship between financial development and economic growth for economies in transition during the early 1990s (e.g., Berglöf and Bolton, 2002; Dawson, 2003). However, this relationship can vary depending on the level of financial development in a county or region (e.g., Rioja and Valev, 2004). During the several years prior to the financial crises of 2007-2008, many of the Central and Eastern European (CEE) countries, including the Baltic countries, experienced a period of high economic growth and a rapid expansion of credit facilitated primarily by foreign-owned banks. Considering the economic development in the Baltic countries during 1995-2008, we find support for the view that the banking sector development can cause economic growth in the long run (cf. Kenourgios and Samitas, 2007; Caporale, et al., 2009, for similar results for a selection of CEE countries). This is also consistent with the findings of Bonin et al. (2005), that suggest that in transition countries foreign-owned banks provide better service and are more cost-efficient than other banks, and hence, can have a larger impact of capital accumulation or productivity of capital.

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