DEPARTMENT OF ECONOMICS

Financial Integration, Monetary Policy and Stock Prices: Empirical Evidence for the New EU Member States

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Abstract

We provide empirical evidence on the interaction between monetary policy and stock prices in 4 new EU member states of Central and Eastern Europe by estimating a small open economy macroeconometric model (SVAR) identified by means of short-run restrictions. Our modeling choices reflect the increasing integration between the NMS and the Euro Area. Our contributions are twofold. We analyze the monetary transmission mechanism through stock prices in the NMS and we determine the extent to which financial markets in the aforementioned countries are sensitive to euro area monetary policy actions. We conclude that stock prices in the NMS are more sensitive to changes in the Euro Area interest rate than to the domestic one. Only in the Czech Republic and Poland we find a significant negative effect of contractionary monetary policy on stock prices. Moreover, we find that the volatility of stock prices in the NMS is mainly due to shocks related to exchange rate and Euro Area monetary policy shocks.

Keywords: monetary policy shocks, new EU members, stock prices, financial integration.

JEL codes: E440 E520 F360 G100

1 Introduction

Since the outburst of the financial crisis in 2007, financial markets came again under the spotlight. The crisis made clear that bad practices and consequent turmoils in the financial world can have disastrous consequences on the real economy. Hence, academics and policymakers are making efforts to formulate policies in order to prevent an event like the recent crisis to repeat itself. On the grounds of monetary policy, a great deal of attention has been devoted to the issue as to whether Central Banks should target stock market inflation. Given that asset prices are highly volatile and prone to bubbles and bursts, why shuldn't Central Banks extend their primary concern of price stability to the broader concept of asset prices stability? The debate between the advocates of explicit asset prices targeting by Central Banks[1] and the opponents[2] is still open; nevertheless, both factions agree on the fundamental assumption that monetary policy can influence stock prices. While monetary policy is carried out by Central Banks in order to achieve goals in terms of price stability, output growth and employment, the effects of changes in the instrument of monetary policy (i.e. the short-term interest rate) are only slowly passed through the real economy.[3] On the contrary, financial markets are much more reactive, in that asset prices tend to quickly incorporate new information. In particular, the stock market is the most thoroughly monitored market by investors who observe new economic developments and are quick in incorporating them in their investment decisions.

Empirical evidence focusing on the United States and other developed countries recognises the ability of monetary policy to influence stock prices.[4] We focus our attention on the new EU member states of Central and Eastern Europe. In two recent waves of enlargement, 10 Central and Eastern European countries (henceforth, CEECs) became members of the European Union, bringing the number of official members to 27.[5] The liberalization of capital markets

[1] See for example Borio and White (2004), Bordo and Jeanne (2002) and Roubini (2006). A recent study focusing on the Euro Area is performed by Cassola and Morana (2002), who conclude that a monetary policy focused on maintaining price stability in the long-run can also contribute to stock price stability.

[2] Bernanke and Gertler (2001) and Schwartz (2002)

[3] The reader is referred to Mishkin (1996) fo a survey on the channels of monetary policy transmission.

[4] Thorbecke (1997), Lastrapes (1998), Rapach (2001), Cassola and Morana (2002), Rigobon and Sack (2004), Neri (2004), Ioannidis and Kontonikas (2008).

[5] The Czech Republic, Estonia, Hungary, Latvia, Lithuania, Poland, the Slovak Republic and Slovenia joined the EU in 2004; Bulgaria and Romania in 2007.

initiated in the beginning of the 1990s and EU membership stimulated financial integration both within these countries and between them, the Euro area and the rest of the world. Since the beginning of the 1990s the new EU member states (henceforth, NMS) have been recipients of massive capital inflows, both in the form of FDI and, more recently, financial flows, originating maily in Euro Area countries such as Austria and Germany, Finland for the Baltic countries and Greece for Bulgaria and Romania.[6] Financial integration between the old and new EU member states is not only reflected in increased financial flows, but also in the behavior of the stock markets. Recent studies by Christiansen and Ranaldo (2009) and Savva and Aslanidis (2009) explore this issue by examining the extent of co-movement between stock prices in the Euro Area and in the NMS. While their methodologies differ, they both find strong linkages between the NMS and Euro Area stock markets. In particular, the connection between old and new member states is found to increase after the 2004 enlargement, and seems to be a phenomenon typical of the European Union, instead of a general tendence of increased worldwide financial integration. [7]

This study aims at exploring the sensitivity of equity markets in the NMS to changes in domestic and Euro Area monetary policy. Given the increased financial linkages between the two regions, this question becomes relevant also in light of the future membership of the NMS in the European Monetary Union (henceforth, EMU). Prospects of EMU membership caused the proliferation, in recent years, of a large amount of studies focussing on assessing the extent of similarity of the monetary transmission mechanism both among NMS and between them and the Euro Area. On one side, Elbourne and De Haan (2006) estimate a SVAR model of 10 transition countries and find substantial differences in monetary transmission among them regarding both inflation and output. Moreover, they do not find any systematic relation between financial structure and monetary transmission. On the other hand, Anzuini and Levy (2007) find that although impulse-responses for the NMS are qualitatively similar to those found for old EU members, they are quantitatively less pronounced. The authors explain this finding with the difference in the degree of financial development between the two regions. Jarocinski (2008) compares the monetary transmission mechanism between the NMS and the Euro Area estimating a

[6] We refer to Lane and Milesi-Ferretti (2007), Pirovano, Vanneste and Van Poeck (2009), for an overview of capital inflows and external asset positions in the Central and Eastern European Countries.

[7] On stock market integration in NMS see also Babetsii et al (2007), Cappiello et al (2006), Chelley-Steeley (2005), Poghosyan (2009).

SVAR with Bayesian techniques in order to correct for small sample bias. Just like Anzuini and Levy (2007), he finds similarity in the impulse-responses of output and prices to monetary policy shocks: in particular, the similarity in responses increases in more recent periods, when inflation in the NMS decreased. To the best of our knowledge, none of the studies conducted so far has taken into account the potential role of the stock market in the transmission of monetary policy in these countries. Given the heavy financial linkages between the NMS and the EA, it is reasonable to assume that a change in the foreign interest rate might have an impact on the NMS' stock markets. Hence, we estimate a macroeconometric model of a small open economy using monthly observations for the Czech Republic, Hungary, Poland and Slovenia. We identify the structural VARs by means of short-run restrictions following Neri (2004), Jarocinsky (2008) and Li, Iscan and Xu (2010). The model allows us to explore two main issues. First, improve our understanding of monetary policy transmission through the stock market in the NMS. Second, the extent to which a change in the Euro Area monetary policy influence the stock market in the NMS. We find that while the effect of domestic monetary policy on stock prices presents some differences across countries, in all countries a change in the Euro area interest rate has a significantly negative impact on the domestic stock markets.

2 Theory & Evidence

According to the well-known present value model, the current price of a stock (P_t) is defined as the discounted value of the stream of future expected cash flow (D_{t+j}) over the investment horizon (K).[8]

$$P_t = E_t\left(\sum_{j=1}^{K}\left(\frac{1}{1+R_t}\right)^j D_{t+j}\right) \tag{1}$$

Where E_t is the expectation operator based on the information set available at time t and R_t is the rate of return used by economic agent to discount the future. From this simple model we can see that monetary policy can influence stock prices in two ways. A first, direct effect is on the discount rate, if we assume that discount rates are tied to market rates, which the Central Banks is

[8] $P_t = E_t\left(\sum_{j=1}^{K}\left(\frac{1}{1+R_t}\right)^j D_{t+j}\right) + E_t\left(\left(\frac{1}{1+R_t}\right)^K P_{t+K}\right)$ Imposing the no-bubble condition $\lim_{K\to\infty} E_t\left(\left(\frac{1}{1+R_t}\right)^K P_{t+K}\right) = 0$ we obtain the well-known form as in equation 1.

able to influence. Secondly, monetary policy affects the expectations of future cash flow. A rise in the interest rate implies a higher cost of investment, which in turn decreases the expected future cash flow, leading to lower stock prices. In turn, stock prices affect the real economy by influencing financial wealth, and with it consumption and investment decisions. Changes in stock prices are then transmitted to the real economy through four channels.[9] First, an increase in stock prices increases the market value of a firm's assets relative to their replacement costs, the so-called Tobin's Q (Tobin 1969). Therefore, the cost of new capital is low relative to its replacement cost and firms are able to finance new investmets with small issuance of new shares. Second, an increase in asset prices due to a decrease in the policy rate increases the wealth of investors, thereby stimulating consumption and the economic activity. According to Modigliani's (1971) permanent income theory, individuals prefer smooth consumption paths to volatile ones: a permanent increase in wealth then translates into higher current and future consumption. The third channel relates to balance sheet effects, and it is due to the presence of asymmetric information in credit markets. Firms are constrained in the amount of bank credit they can be granted by the value of their collateral. Rises in stock prices increase the value of such collateral, allowing firms to borrow funds to finance investments and fostering economic activity. Finally, a cash flow channel arises from the fact that increases in stock prices decrease consumers' perception of finding themselves in financial distress.

Empirical studies conducted so far broadly support the notion that monetary policy is able to significantly affect stock prices: in particular, contractionary (expansive) monetary policy decreases (increases) stock prices. A number of empirical studies have been produced, relying on different statistical methods. The great majority of studies relies on the methodology of structural VAR (SVAR), and focused, at least in the beginning, on the United States' stock market and the FED monetary policy. Thorbecke (1997) estimates a VAR model identified by means of a Cholesky scheme, including monthly equity returns, output growth, inflation, and the federal funds rate. His results highlight that expansionary monetary policy has a large and positive effect on monthly stock returns. Thorbeke's results are confirmed by Rapach (2001) who identifies the shocks in his VAR model by means of long-run restrictions. Again, a negative relationship between changes in the Federal Funds rate and stock prices emerges.

[9] Mishkin (1996).

The same conclusions for the US are reached by Rigobon and Sack (2004), who identify their VAR by means of changes in the covariance between stock prices and interest rates after a (known) change in the variance of monetary policy shocks. International evidence on the effect of monetary policy on stock prices focuses mainly on developed countries. The earliest contribution is by Lastrapes (1998), who identifies a VAR model for the G7 countries and Holland by means of long-run restrictions on a model including an index of real equity prices, the interest rate, output, the price level and the nominal money stock. He finds that in many countries real equity price indices respond positively, persistently and significantly to an expansionary monetary policy shock: in particular, the effect is stronger in Holland, Italy and Japan (with peaks between 2 and 3 percent). A later paper by Neri (2004) updates the results of Lastrapes considering a sample period for the same countries from 1985 to 2000. While relying on the same methodology of structural VAR, he adopts an identification scheme based on short-run restrictions. Moreover, he models the G7 countries as small open economies, where the exchange rate plays an important role in the monetary transmission mechanism and its inclusion allows a better identification of money supply shocks. The monetary policy rule is specified such that the supply of money is a function of the monetary aggregate, the nominal exchange rate, industrial production, the consumer price index and the commodity price index. It is therefore assumed that the Central Bank reacts to contemporaneous domestic and external economic developments. His results are in line with the notion that a contractionary monetary shock produces a decrease in stock market index, in all countries considered. The study that is more close to ours is by Li, Iscan and Xu (2010). Based on the observed trade and financial linkages between Canada (a small open economy) and the United States (a relatively closed economy), they study the differences in the response of the stock market to changes in monetary policy in the two countries, modeling them differently according to their structural characteristics. Their analysis relies on a VAR model following Lastrapes (1998), imposing short-run identification restrictions. They find that while in the United States the response of stock prices to a contractionary monetary policy shock is large and persistent, in a small open economy such as Canada the response is smaller and less persistent. Moreover, unexpected changes in the U.S. Federal Funds rate have a significant impact on Canadian stock prices. Given the parallelism between the binomial U.S.-Canada and Euro Area-NMS, we perform a similar analysis adapting it to the characteristics of the NMS.

3 Methodology

3.1 A SVAR refresher

Vector Autoregressive models are extensively used in macroeconomic analysis and particularly in monetary economics, in order to analyze the effect of exogenous shocks in monetary policy on macroeconomic variables. One of the biggest challenges of modeling monetary policy is to distinguish between monetary policy actions that the market has already anticipated and those it has not. If we accept the rational expectations hypothesis, anticipated monetary policy changes will already be embedded in the current stock prices, while only unanticipated actions will have an effect on current returns. In this context, the VAR framework is the most adequate. A Vector Autoregressive model is a system of equations where each variable influences the time path of all variables in the system.[10] Hence, it allows to model dynamic inter-dependencies among variables. Each variable in the VAR model depends on its own lags, lags of the other variables in the model and (possibly) exogenous variables. In our case, the model will contain only endogenous variables. The structural form of a VAR model without exogenous variables takes the following form:

$$Bx_t = \Gamma_0 + \Gamma_1 x_{t-1} + \Gamma_2 x_{t-2} + ... + \Gamma_p x_{t-p} + \varepsilon_t \quad (2)$$

Where x_t is a $(n \times 1)$ vector of the n endogenous variables included in the VAR (in our case $n = 7$), $x_{t-1}, ..., x_{t-p}$ are $(n \times 1)$ vectors the lagged values of the variables in the VAR, B, $\Gamma_1, ..., \Gamma_p$ are $(n \times n)$ coefficient matrices, Γ_0 is a $(n \times 1)$ vector of constant terms and ε_t is a $(n \times 1)$ vector of structural innovations. In partricular, the structural innovations are assumed to have zero mean and variance-covariance matrix Σ, uncorrelated across equations. In other words, the structural innovations are white noise, uncorrelated disturbances with constant variance. From this notation it is clear that VARs in structural form allow for feedback effects between the variables in the model. To see this, note that the left-hand side of equation 2 implies that each variable in the VAR is influenced by the contemporaneous values of all the other variables in the model. Hence, the structural form of the VAR corresponds to a simultaneous equation model. The structural form of a VAR model cannot be estimated directly, because of

[10] The reader is referred to Enders (2004), Plasmans (2006) and Lütkepohl (2006) for multivariate time series techniques.

such feedback effects. Feedback implies that each variable of the model is correlated with the error terms of all the other equations, and standard estimation techniques require the regressors to be independent of the error terms. In order to get to a formulation more suitable for estimation purposes, let us rewrite the model by premultiplying both sides by B^{-1}.[11]

$$B^{-1}Bx_t = B^{-1}\Gamma_0 + B^{-1}\Gamma_1 x_{t-1} + B^{-1}\Gamma_2 x_{t-2} + ... + B^{-1}\Gamma_p x_{t-p} + B^{-1}\varepsilon_t$$

$$x_t = A_0 + A_1 x_{t-1} + A_2 x_{t-2} + ... + A_p x_{t-p} + u_t \tag{3}$$

The model in equation 3 is called a VAR in reduced form, where x_t is again our $(n \times 1)$ vector of the n endogenous variables at time t, $A_0 = B^{-1}\Gamma_0$ is a vector of constants, $A_1 = B^{-1}\Gamma_1, ..., A_p = B^{-1}\Gamma_p$ are matrices of coefficients of the p lagged values of all endogenous variables and $u_t = B^{-1}\varepsilon_t$ is the reduced form error term. From this expression it is clear that the reduced form disturbances are linear combinations of the structural form disturbances. It is possible to prove [12] that the error terms of the reduced form model have zero mean, constant variance and are individually serially uncorrelated. Nevertheless, being linear combinations of the structural form disturbance, they are correlated with each other[13]: therefore, the variance-covariance matrix of the reduced-form residuals is not diagonal.

The reduced form of a VAR is estimable, in that it overcomes the simultaneity problems of the structural form. The goal is then to estimate the parameters of the reduced form model and then to retrieve the parameters of the structural model. But there is one more problem: the reduced form model has less parameters to be estimated than the structural form model and hence it does not allow to uniquely identify the structural form parameters. Generally speaking, the reduced form model in equation 3 has $N^2p + N + \frac{N(N+1)}{2}$ parameters to be estimated: $N^2p + N$ in the equation for x_t and $\frac{N(N+1)}{2}$ unique elements in the covariance matrix. The structural form of equation 2 has $2N + N^2p$ unknown parameters, which is smaller than $N^2p + N + \frac{N(N+1)}{2}$.

It is therefore necessary to impose restrictions on the reduced form model in

[11] Assuming B is non-singular.

[12] The reader is referred to Enders (2004) p.296

[13] They are uncorrelated if and only if in each equation the coefficients of the contemporaneous values of all variables are zero.

order to identify the primitive stystem. Instead of imposing a recursive structure, we rely on economic theory to impose restrictions to recover the structural innovations from the residuals of the estimated reduced-form VAR. The structural innovations allow us to study the effect of an exogenous shock on a variable on the dynamics of the variables in the model, *ceteris paribus*. Moreover, we will compute the impulse response functions (IRFs), which graphically depict how the innovations to one variable in the model affect another variable after a certain number of periods (again, *ceteris paribus*). In order to do so, we cannot rely on the estimated reduced-form errors, because these residuals are correlated with each other: this implies that a shock in one of the variables will trigger shocks to other variables and hence the ceteris paribus condition is not met. Hence, it is not possible to establish a clear causation link between the occurrence of shock to different variables. In order to determine causation, we need to rely on the structural innovations, whose covariance matrix is diagonal. Our SVAR model will rely on economic theory to impose restrictions on the covariance matrix of the reduced form residuals in order to identify the structural shocks. Recall that equation 3 implies that:

$$\varepsilon_t = Bu_t \tag{4}$$

Where B is the matrix of contemporaneous effects. Given estimates of the u_ts, we will recover the ε_ts by imposing suitable, economically meaningful restrictions on the B matrix.

3.2 Identification scheme

The seven variables in our model are the industrial production index (ip), the domestic price level (cpi), the nominal exchange rate vis-à-vis the euro (er), the Euro Area interest rate ($eaint$), the domestic short-term interest rate (i), the monetary aggregate M2 (m) and the stock market index (smi).[14] All variables except the interest rates are in natural logarithms.[15] Hence, our SVAR model in structural form takes the form of equation 2, with:

[14] Data on the consumer price index, industrial production index and the monetary aggregare (money + quasi-money) are from the IMF's International Financial Statistics. Data on exchange rates, the money market rate (except for Hungary, where the T-bill rate is used) and the stock market index are from Eurostat. All variables except the short-term interest rate are in logarithms.

[15] Cfr. Appendix A for a detailed description of the data and of the variables in the model.

$$x_t = \begin{pmatrix} EAint_t \\ er_t \\ ip_t \\ cpi_t \\ i_t \\ m_t \\ smi_t \end{pmatrix} \tag{5}$$

and $x_{t-1}, ..., x_{t-p}$ defined analogously according to the appropriate lag. As before, the structural shocks are assumed to be serially uncorrelated and mutually independent. Again, we can rewrite the model in the reduced form of equation 3, implying that the reduced form residuals are linear combinations of the structural form residuals as in 4. In order to identify the structural shocks we impose short-run restrictions on the B matrix following Neri (2004) and Li, Iscan and Xu (2010). As a result, the long-run behavior of the model is left completely unconstrained. Alternatively, we could restrict the reduced form model by means of the so-called Cholesky decomposition, which consists in restricting the covariance matrix of the reduced form residuals to be a lower-triangular matrix and which leads to a just-identified model (i.e. a model with as many restrictions as there need to be). Although it is handy from a practical point of view, the Choleski decomposition imposes restrictions without a theoretical foundation; moreover, it imposes an ordering to the variables in the VAR which might or might not be appropriate. Hence, if there is not a theoretical basis to impose such restrictions, the impulse-resposes might be poorly identified.[16] We will impose a Choleski-type of identification as a sensitivity check (cfr. section 5). We identify our SVAR model by specifying the B matrix as follows:

$$\begin{pmatrix} \varepsilon_{t,EAint} \\ \varepsilon_{t,er} \\ \varepsilon_{t,ip} \\ \varepsilon_{t,cpi} \\ \varepsilon_{t,i} \\ \varepsilon_{t,m} \\ \varepsilon_{t,smi} \end{pmatrix} = \begin{pmatrix} 1 & 0 & 0 & 0 & 0 & 0 & 0 \\ b_{21} & 1 & 0 & 0 & 0 & 0 & 0 \\ b_{31} & b_{32} & 1 & 0 & 0 & 0 & 0 \\ b_{41} & b_{42} & b_{43} & 1 & 0 & 0 & 0 \\ 0 & b_{52} & 0 & 0 & 1 & b_{56} & 0 \\ 0 & 0 & b_{63} & b_{64} & b_{65} & 1 & 0 \\ b_{71} & b_{72} & b_{73} & b_{74} & b_{75} & b_{76} & 1 \end{pmatrix} \begin{pmatrix} u_{t,EAint} \\ u_{t,er} \\ u_{t,ip} \\ u_{t,cpi} \\ u_{t,i} \\ u_{t,m} \\ u_{t,smi} \end{pmatrix} \tag{6}$$

[16] Again, we refer the reader to Enders (2004) for a detailed description of the different ways to impose restrctions on the reduced form model and their critics.

This leads to the following specification of the equations in the system, which is derived combining the proposed identification scheme with equation 2 . The money supply equation is a feedback rule according to which the Central Bank sets the short-term interest rate. We assume that, when setting the monetary policy instrument, the Central Bank does not observe the current values of real output and the price level, which are observable only after a lag. This assumption reasonable in light of the monthly frequency of our dataset. Hence, money supply is defined as conditional on the contemporaneous values of the exchange rate and the monetary aggregate, and lagged values of the other variables in the model, as follows:

$$i_t = b_{50} + b_{52}er_t + b_{56}m_t + h(x_{t-p}) + \varepsilon_{t,int} \tag{7}$$

In particular, we expect the coefficient b_{52} to be positive: a depreciation of the exchange rate (increase in er_t) causes imports from abroad to be more expensive, hence resulting in inflationary pressures. The Central Bank is then supposed to react by tightening monetary policy. Furthermore, the inclusion of the exchange rate in the Central Bank's reaction function is motivated by the choice of exchange rate regime in the NMS. Although in the late 1990s many countries opted for flexible regimes, their Central Banks tried to keep the domestic currency to fluctuate too much, a phenomenon called "fear of floating" by Calvo and Reinhart (2002). In most recent years, while Poland adopted a free floating regime, the other countries opted for less flexible alternatives: while Hungary and Slovakia pegged their currencies to the euro with large fluctuation bands, the Czech Republic adopted a managed floating regime with the euro as a reference currency.[17] It is therefore reasonable to assume that these countries have been following, although not as their primary objective, some form of exchange rate targeting. The last term in equation 7 represents an exogenous shift in monetary policy: a positive (negative) shock to the short-term interest rate results in an unexpected tightening (relaxing) of monetary policy.

The equation representing equilibrium on the money market is a pretty standard one, linking the monetary aggregate to output, the price level and the interest rate. The resulting LM-type equation is the following:

$$m_t = b_{60} + b_{63}ip_t + b_{64}cpi_t + b_{65}i_t + h(x_{t-p}) + \varepsilon_{t,m} \tag{8}$$

[17] Toma (2007)

Where we expect a negative relationship in the (m, int) space. $\varepsilon_{t,m}$ is an exogenous shock in the demand of money.

The equation for the aggregate supply sees industrial production as dependent on contemporaneous values of the foreign interest rate and the bilateral exchange rate:

$$ip_t = b_{30} + b_{31}EAint_t + b_{32}er_t + h(x_{t-p}) + \varepsilon_{t,ip} \tag{9}$$

Given the heavy presence of foreign investors in the NMS, it is reasonable to assume that changes in the foreign interest rate will have an impact on the capital stock of domestic firms. As many (especially small) firms rely on foreign borrowing to finance their capital stock, an increase in the foreign interest rate represents an increase in their produciton costs. The bilateral exchange rate enters the equation in reason of the imported intermediate goods used in production. $\varepsilon_{t,ip}$ represents aggregate supply shocks (productivity and cost-push).

Following Li, Iscan and Xu (2010), aggregate demand is divided in domestic and external demand. Domestic demand is defined in terms of the factors affecting the domestic price level, as follows:

$$cpi_t = b_{40} + b_{41}EAint_t + b_{42}er_t + b_{43}ip_t + h(x_{t-p}) + \varepsilon_{t,cpi} \tag{10}$$

Where we assume a certain degree of price stickiness reflected in the coefficient b_{43}. The error term represents exogenous demand shocks, e.g. an exogenous change in preferences. External demand is defined by the exchange rate equation. We define the exchange rate as being conditional only on contemporaneous values of the foreign interest rate and lagged values of all other variables in the model. This approach differs from Li, Iscan and Xu (2010) in that they consider the exchange rate market to be strongly efficient, i.e. incorporating all publicly and privately available information. In their model, the exchange rate is dependent on the contemporanoues values of real output, the domestic price level, the monetary aggregate, the short-term interest rate and the foreign interest rate. We adopt a different approach, by modeling the exchange rate equation as:

$$er_t = b_{20} + b_{21}EAint_t + h(x_{t-p}) + \varepsilon_{t,er} \tag{11}$$

We motivate this specification mainly because of the large, short-term and interest-sensitive capital inflows which are typical of the region. We hypothesize that a contemporaneous shift in the foreign interest rate and hence a change in the interest rate differential produces capital flows that put pressures on the value of the currency.

The stock price equation is left unrestricted: current stock prices are influenced by current and lagged values of all variables in the model, reflecting they quick receptiveness of new information. Finally, the foreign interest rate is a contemporaneously exogenous variable.

4 Data & Estimation Results

4.1 Dataset and preliminary analysis

The SVAR model is estimated using data for the Czech Republic, Hungary, Poland and Slovenia. We exclude the Baltic States and Bulgaria because they adopted a strictly fixed exchange rate regime for most of the sample period.[18] In the beginning of the transtition period, many Central and Eastern European countries opted for fixed exchange rate regimes: also the Czech Republic, Hungary, Poland and Slovakia initially pegged their currencies to a basket of currencies. Nevertheless, despite the *de jure* regime was fixed, *de facto* it was more close to an intermediate or free floating regime.[19] Moreover, following Jarocinski (2008) we exclude Romania and Slovakia because for a large part of the sample they exhibit a discrepance between the central bank rate and the money market rate: in particular, the latter moves independently from the former. One of our main assumptions (cfr. section 2) is that the Central Bank is able to influence money market rates, and it is clearly violated. The sample period differs slightly for each country: from January 1998 to August 2009 for the Czech Republic (140 observations), from April 1995 to August 2009 for Hungary (173 observations), from May 1995 to July 2009 for Poland (171 observations) and from May 1997 to October 2005[20] for Slovenia (106 observations).

[18] Estonia, Lithuania and Bulgaria are currently operating under a currency board on the euro (previously on the Deutsche Mark or the U.S.Dollar); Latvia settled on a fixed peg on the DST and on the euro since 2005.

[19] (Toma, 2007)

[20] The sample for Slovenia is limited due to the lack of data on the monetary aggregate from October 2005 onwards.

Table 1: Lag Length Selection

	CZ	HU	PL	SI
AIC	12	3	3	12
FPE	2	3	3	2
HQC	2	2	2	1
SC	1	1	1	1
Lag chosen	4	2	4	2

Note: Country names have been denoted as follows: CZ=Czech Republic, HU=Hungary, PL=Poland, SI=Slovenia

Figure A1 in appendix A provides plots of the endogenous variables, and tables A1 and A2 summarize the results of tests for stationarity and cointegration. Both visual inspection (Figure A1) and the Phillips-Perron test for unit roots[21] confirm that roughly half of the series for each country are non-stationary. Trace and Maximum Eigenvalue tests for cointegration show the presence of cointegrating relationships in the variables, which increase in number when all seven variables are included in the model. This justifies the estimation of a VAR in levels: when cointegration relatioships are present, differencing the variables results in a misspecification in that it disregards the error correction mechanism.

We now estimate our unrestricted VAR models in levels. First, we determine the optimal lag length of our models by means of the AIC, FPE, HQIC and SBIC information criteria. In particular, we look for the number of lags that minimizes these criteria. It is important to note that while the AIC tends to overestimate the number of lags, under fairly general conditions the HQ and SC criteria estimate the correct number of lags (Lutkepohl (2006)). This procedure leads us to obtain a different number of lags for the models of the different countries, as summarized in Table 1. While for Hungary and Slovenia the 2 lags suggested by the test statistics are sufficient to yield a correctly specified model, for the Czech Republic and Poland we decide to include a number of lags equal to 4. Indeed, for these countries the inclusion of fewer lags results in a large amount of autocorrelation in the residuals of the VAR model. In order to evaluate the adequacy of the variables included in the models, we estimate the unrestricted model for each country excluding, in turn, the Euro Area interest rate, the bilateral exchange rate and the stock market index . In all cases and for all countries, the specification including the whole set of seven variables performs better, as measured by the AIC and SBC. Finally, we check for stability of our

[21] Cfr. Tables A1 and A2 in Appendix 1.

Table 3: Test for Overidentifying Restrictions

	CZ	HU	PL	SI
LR Test	2.0666*	4.9809*	2.3791*	5.4897*
p-value	0.5587	0.1732	0.6664	0.2406

*Note:** significant at the 0.95 level
Note: The models for the Czech Republic and Hungary have 3 overidentifying restrictions; those for Poland and Slovenia have 4 overidentifying restrictions

VAR models, by computing the roots of the characterisitc polynomial: despite some roots are close to one, formal testing reveals that none of them lies outside the unit circle, and our VARs are stable (cfr. Figure A2).

Given that formal testing confirmed our model specification, we impose the short-run restrictions described in the previous section and we obtain estimates of the contemporaneous coefficients, which are summarized in Table 2. While the test for overidentifying restrictions confirmed the validity of the imposed identification scheme in the models for Poland and Slovenia, the baseline specification was rejected for the Czech Republic and Hungary. Hence, in these models, we had to drop one overidentifying restriction: specifically, in these models coefficient b_{54} is not restricted to be equal to zero. This leads to a specification of the money supply equation (cfr. equation 7) where the Central Bank is able to observe the current price level and incorporate this information in the monetary policy reaction function. The overidentifying restriction tests for this alternative specification don't reject the null that the overidentifying restrictions are valid (Table 3).

4.2 Impulse-response functions

In this section, we examine the impulse-response functions and the variance decompositions obtained imposing the chosen structural identification scheme.

As our interest lies in the analysis of the impact of monetary policy actions on stock prices, we first examine the responses of the stock price index to a one standard deviation structural innovation to domestic and Euro Area monetary policies, represented in figure 1. The upper part of figure 1 depicts the response of the stock price index to a contractionary, one standard deviation shock to domestic monetary policy, i.e. an increase in the domestic interest rate. As expected, a contractionary monetary policy shock causes a decrease in the stock price index in all countries, albeit with some differences. While the negative

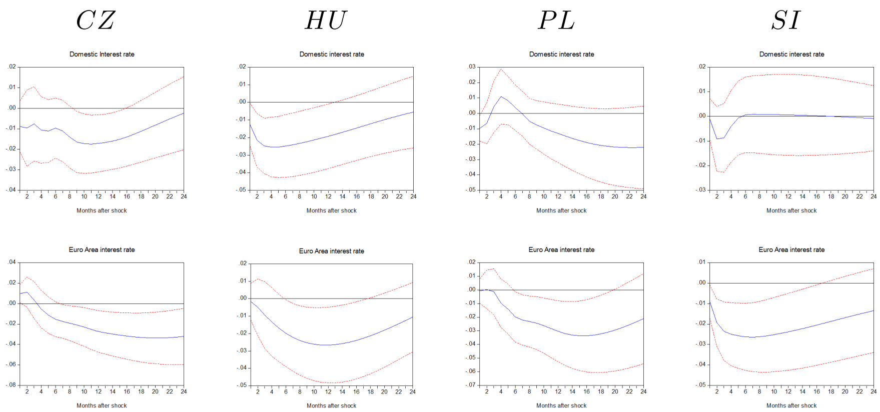
Table 2: Estimated contemporaneous coefficients

	CZ	HU	PL	SI
Exchange rate equation				
$EAint$	0.329 (0.662)	-2.147 (0.003)***	-0.585 (0.597)	0.254 (0.317)
Aggregate supply				
$EAint$	1.342 (0.383)	1.549 (0.225)	0.600 (0.668)	6.078 (0.000)***
er	0.176 (0.304)	-0.101 (0.429)	-0.006 (0.948)	-0.493 (0.403)
Domestic demand				
$EAint$	0.402 (0.234)	-0.468 (0.141)	-0.174 (0.516)	-0.402 (0.274)
er	-0.077 (0.039)**	0.022 (0.476)	-0.033 (0.072)*	0.097 (0.455)
ip	0.011 (0.543)	0.010 (0.574)	-0.003 (0.830)	0.027 (0.193)
Money supply				
er	0.068 (0.002)***	-0.080 (0.427)	0.101 (0.065)*	0.730 (0.084)*
cpi	0.139 (0.001)***	0.940 (0.316)	-	-
m	-0.135 (0.118)	0.885 (0.296)	-0.777 (0.040)**	0.446 (0.240)
Money mkt equilibrium				
ip	-0.052 (0.209)	0.381 (0.338)	0.062 (0.291)	-0.105 (0.250)
cpi	-0.148 (0.534)	-1.252 (0.043)**	-0.848 (0.025)**	-0.035 (0.929)
i	2.644 (0.090)*	-6.741 (0.372)	1.662 (0.078)*	-2.371 (0.086)*
Stock price equation				
$EAint$	7.679 (0.017)**	-6.278 (0.059)*	-1.293 (0.653)	-7.889 (0.035)**
er	-0.580 (0.126)	-2.551 (0.000)***	-1.175 (0.000)***	0.282 (0.836)
ip	0.087 (0.621)	0.439 (0.027)**	0.053 (0.733)	-0.098 (0.656)
cpi	0.458 (0.581)	2.188 (0.009)***	0.054 (0.948)	-1.675 (0.088)*
i	-7.166 (0.000)***	-2.522 (0.000)***	-0.675 (0.170)	-0.702 (0.327)
m	1.008 (0.017)**	0.496 (0.080)*	-0.726 (0.032)**	-0.188 (0.543)

*Note:** denotes sgnificance at the 0.95 level

Note: Country names have been denoted as follows: CZ=Czech Republic, HU=Hungary, PL=Poland, SI=Slovenia

Figure 1: Responses of the Stock Price Index to a one Standard Deviation Structural Innovation in the domestic and Euro Area monetary policy (±2 standard errors)



impact effect is clearly negative and statistically significant in Hungary and Poland, it is not statistically significant in the Czech Republic and Slovenia. Moreover, in the latter country, the shock is not significant for the whole time horizon. Both magnitude and persistence of the shock is greater in countries such as the Czech Republic and Hungary. In both countries, the impact effect is in the neighborhood of -0.01 (albeit in the Czech Republic it is not significant) and the stock price index remains negative for the rest of the time horizon, showing a tendency to revert towards the pre-shock equilibrium. In the Czech Republic the negative response of the stock price index becomes statistically significant only 8 months after the shock, and it reaches a peak at the 10th month after the shock (-0.017). In Hungary the negative peak of -0.025 is reached much sooner (5th month), and the response is significant until a year after the shock occurred. Moreover, the dynamic adjustment towards the pre-shock equilibrium is smooth, signalling a high degree of persistence. The reaction of the stock price index in Poland is somewhat different. Despite on impact the stock price index significantly decreases (-0.009), it exhibits a (non significant) increasing tendency until the 4th month after the shock, where it reaches a positive peak of 0.01 after steadily decreasing for the rest of the time horizon.

The bottom part of figure 1 represents the responses of the stock price index to a one standard deviation increase in the Euro Area interest rate. A first glace suffices to see an overall negative effect and a high degree of similarity between countries. The largest and statistically significant impact effect is registered in

Slovenia (-0.01), followed by a further decline in the stock price index until a negative peak of -0.026 is reached at the 7th month after the shock. Note how the persistence of this shock is much higher than in the case of a domestic monetary policy shock: here, the stock price index does not revert to its pre-shock level within 2 years after the shock. The impact effect in Hungary and Poland is very small and not statistically significant, but it becomes larger and significant in later months, signalling that a shock in the foreign interest rate is not immediately transmitted to the domestic economies. While in Hungary a negative peak (-0.026) is reached a year after the shock, in Poland it is reached only after 16 months. The dynamics of the response of the Polish stock price index stands out from the group in that the impact effect of a rise in the foreign interest rate is positive and statistically significant. Nevertheless, from the 3rd month onwards it starts decreasing and it becomes negative from the 4th month until the rest of the time horizon.

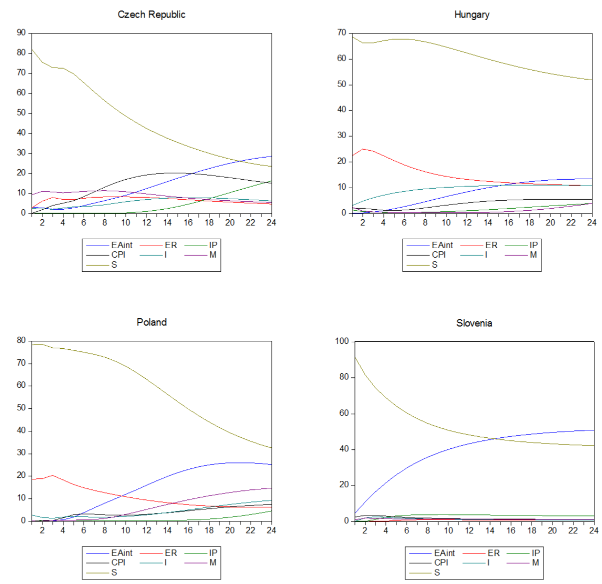
From this analysis we are led to formulate two main conclusions. First, domestic monetary policy affects stock prices in some of the countries considered, and it does so in the expected direction. Country differences in terms of magnitude and persistence of the shock's effect suggest that it is not appropriate to formulate general monetary policy directions. On one side, in Slovenia an unexpected monetary policy tightening does not have a significant impact on the stock market, neither on impact or in later periods. On the other side, in Hungary the effect is much more persistent and tends to reach its peak few months after the shock occurred. Given this lagged effect, the Central Bank should be cautious before enacting a strong contractionary policy action. In Poland, while the impact effect is significantly negative, the dynamic evolution of the shock is surrounded by a deal of uncertainty. Moreover, the swing of the stock price index from negative to positive and then back to negative is a symptom of instability, which the Central Bank might wish to take into account before unexpectively change its actions. Second, Euro Area monetary policy affects stock prices in the NMS, and the effect is negative. It is interesting to notice how in Slovenia, while domestic monetary policy does not have a significant impact on the stock price index, changes in the Euro Area interest rate do. Moreover, in all countries except Slovenia (where also the impact effect is significantly negative), the response of the stock price index seems to be transmitted to stock prices with a lag. Given that according to equation 1 the discount rate that enters the defiition of a stock's price is the domestic one, the negative effect of the Euro Area interest rate on stock prices must stem from elsewhere. We suggest an explanation related to the open economy nature of our model and the high degree of finanical integration between the NMS and the Euro Area. Given the heavy reliance of the NMS on foreign borrowing to finance investments, an increase in the Euro Area interest rate represents an increase in the cost of funds. This in turn causes an increase in firms' costs and depresses the expectations of future cash flow, decreasing stock prices.

The bottom panels of figures A3, A4, A5 and A6 in the Appendix reveal that there are other macroeconomic variables besides the domestic and foreign interest rates that affect stock prices in the NMS. An unexpected increase in the exchange rate, i.e. a depreciation, significantly decreases stock prices in all countries except Slovenia, where the response is very weak. Being small open economies, an unexpected change in the exchange rate could have two roots. First, it could be given by an unexpected stop in capital inflows, which puts downward pressure on the exchange rate. The decrease in demand for domestic capital, including domestic stocks, causes their prices to decrease to re-establish equilibrium on the market. Second, following the explanation of Li, Iscan and Liu (2010) an unexpected depreciation might be triggered by an unexpected decline in the foreign demand of domestic goods, which depresses production and expectations of future cash flow. An unexpected increase in industrial production, i.e. an aggregate supply shock, seems to increase stock prices in all countries except Slovenia, but the responses are characterized by a large degree of uncertainty. The same holds for an unexpected increase in the consumer price index. Finally, an unexpected exogenous increase in the stock price due, for example, to a sudden increase in the demand for stocks significantly increases stock prices in all countries, and the responses are quite persistent.

Despite it is not the primary focus of the analysis, we briefly describe the impulse-response functions relating to the other variables in our SVAR models, referring to figures A3, A4, A5 and A6 in the Appendix. The dynamic responses obtained for the Czech Republic (Figure A3) are broadly consistent with the theoretical expectations. An aggregate supply shock, i.e. an unexpected increase in industrial production, causes the exchange rate to appreciate, the consumer price index to decline, and the interest rate to rise (albeit not significantly), in responses of expected inflationary pressures (cfr. column 3 of figure A3). An unexpected increase in the liquidity in the economy leads the domestic interest rate to decrease, and real output and prices to increase. On the external side, the exchange rate depreciates because of the increased amount of liquidity lowers the price of domestic currency. An unexpected monetary policy contraction does not produce strongly significant effects on the real economy: the responses of prices and real output are quite weak. While the impact on real output is not significant, in the long run industrial production appears significantly decreased. As for the price level, the dynamic response is not significant at any time horizon. Moreover, an unexpected increase in the short term interest rate does not produce the desired liquidity effect, in that the monetary aggregate increases instead of decreasing, albeit not significantly. The Euro Area interest rate does not react to the unexpected domestic policy shock. Finally, an unexpected increase in the Euro Area interest rate significantly increases the domestic interest rate and produces an appreciation of the domestic exchange rate followed by a smooth depreciation, consistently with uncovered interest parity. The dynamic responses for Hungary to aggergate supply and the foreign interest rate are similar to the ones of the Czech Republic. Concerning the responses to an unexpected monetary policy shock, we clearly see a liquidity effect, in that the monetary aggregate significantly decreases. Moreover, while prices significantly decrease, real output does not seem to be significantly affected at any time horizon. The responses for Poland to shocks to the Euro Area interest rate and liquidity shocks are similar to the ones previously described, despite in this case following an unexpected increase in liquidity real output significantly increases only in the long run and the exchange rate does not significantly react. The dynamic responses for Poland differ in the cases of an aggregate supply shock (where none of the variables seems to be significantly affected) and a contractionary monetary policy shock. In the latter case, we clearly see the absence of a liquidity effect, in that the monetary aggregate significantly increases and a significant increase in the price level. The latter phenomenon, known as "price puzzle", is not extraneous to the empirical literature on monetary transmission in Central and Eastern Europe.[22] In Slovenia, we observe a strong liquidity effect after an unexpected monetary policy shock, and a significant decrease in prices in the short-run. On the other side, liquidity shocks and aggregate supply

[22] An excellent survey is provided by Egert and MacDonald (2009), who compare the most recent results on monetary transmission mechanism in the new EU member states. The empirical evidence on the effects of monetary policy on output, prices and the exchange rate is quite mixed. While some authors find evidence of price and exchange rate puzzles (e.g. Arnostova and Hurnık (2004) for the Czech Republic, Creel and Lavasseur (2005) for all countries, Héricourt (2005) for the Czech Republic, Hungary and Slovenia), others detect a decrease in prices and output following a contranctionary monetary policy shock (Jarocinski (2005), Vonnak (2005).

Figure 2: Variance Decomposition of the Stock Price Index



shocks are in most cases not significant.[23]

4.3 Variance Decomposition

In this section, we analyze the contribution of the seven variables in the model to fluctuations in the stock price index by means of variance decomposition. In Figure 2, the total variation in the stock price index is decomposed in fractions ascribed to the single shocks in the model, during a time horizon of 24 months. As it is to be expected, shocks to the stock price index itself account for the largest fraction of the variation. Nevertheless, important considerations stem from the observation of the contributions of the other six variables. First of all, we immediately notice substantial differences across countries. Secondly, it is evident that it is important to make a distinction between the short and the long-run in that shocks that do not contribute much to the variation in stock prices on impact, gain importance in the medium to long term. In the Czech Republic, shocks to the monetary base and the exchange rate account for the largest fraction of variation in stock prices while losing importance later

[23]We suspect this result is given by the limited number of time observations available for this country.

on in the time horizon, and being outperformed by the Euro Area and the domestic interest rates. In Hungary and Poland, the exchange rate accounts for the largest fraction of variation of stock prices on impact and in the short to medium run: 23% and 19% respectively. In Hungary, the domestic interest rate comes right after the exchange rate: although on impact its contribution is around 4%, it increases in the medium-run and it stabilizes around 10%. In Poland, both the domestic and the foreign interest rates contribute to the variation of stock prices quite insignificantly on impact: nevertheless, the Euro Area interest rate becomes increasingly important and in the medium-run, it accounts for more than 20% of the total variation. The picture for Slovenia is quite different from the previous ones. Here, the Euro Area interest rate dominates the scene: although on impact it accounts for only 5% of the total variation, already after 4 months it reaches 20% and keeps on increasing with time. All other variables have a negligible impact. These figures reinforce our previous conclusion according to which increasig financial openness has an impact on stock prices in the NMS. Moreover, it emerges that domestic stock prices are more sensitive to external developments than domestic ones. This can be explained by the heavy presence of foreign investors in the stock markets of the new EU members, which outnumber the proportion of domestic investors.

5 Sensitivity Analysis

In what follows we check the robustness of our results to four alternative model specifications. Table 3 provides a summary of the alternative specifications estimated in this section.

The first specification differs from the baseline in the specification of the monetary policy rule. While in the baseline specification we assumed that the Central Bank does not react to contemporaneous values of output and the price level, we now relax this assumption and estimate coefficients b_{53} and b_{54}.

In the second specification, we relax the hypothesis according to which the nominal exchange rate is contemporaneously influenced only by the Euro Area interest rate. We now assume that the exchange rate market is efficient, in the sense that it incorporates all available information: the contemporaneous coefficients of output, the price level, the domestic interest rate and the money stock are now left unrestricted, as we can read in the second line of table 3. As the exchange rate is contemporaneously affected by the contemporaneous values

Table 3: Alternative specifications

Alternative 1	$i_t = b_{50} + b_{52}er_t + b_{53}ip_t + b_{54}cpi_t + b_{56}m_t + \epsilon_{t,i}$	$\begin{pmatrix} 1 & 0 & 0 & 0 & 0 & 0 & 0 \\ b_{21} & 1 & 0 & 0 & 0 & 0 & 0 \\ b_{31} & b_{32} & 1 & 0 & 0 & 0 & 0 \\ b_{41} & b_{42} & b_{43} & 1 & 0 & 0 & 0 \\ 0 & b_{52} & b_{53} & b_{54} & 1 & b_{56} & 0 \\ 0 & 0 & b_{63} & b_{64} & b_{65} & 1 & 0 \\ b_{71} & b_{72} & b_{73} & b_{74} & b_{75} & b_{76} & 1 \end{pmatrix} \begin{pmatrix} EAint \\ er \\ ip \\ cpi \\ i \\ m \\ smi \end{pmatrix}$
Alternative 2	$er_t = b_{20} + b_{21}EAint_t + b_{23}ip_t + b_{24}cpi_t + b_{25}i_t + b_{26}m_t + \epsilon_{t,er}$	$\begin{pmatrix} 1 & 0 & 0 & 0 & 0 & 0 & 0 \\ b_{21} & 1 & 0 & 0 & 0 & 0 & 0 \\ b_{31} & b_{32} & 1 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 1 & b_{45} & b_{46} & 0 \\ 0 & b_{52} & -1 & b_{54} & 1 & 0 & 0 \\ b_{61} & b_{62} & b_{63} & b_{64} & b_{65} & 1 & 0 \\ b_{71} & b_{72} & b_{73} & b_{74} & b_{75} & b_{76} & 1 \end{pmatrix} \begin{pmatrix} EAint \\ ip \\ cpi \\ i \\ m \\ er \\ smi \end{pmatrix}$
Alternative 3	Cholesky decomposition 1	Ordering of the variables as in the baseline model
Alternative 4	Cholesky decomposition 2	Ordering of the variables as in Alternative 2

of all other variables in the model except the stock price index, its position in the ordering of the variables in the model changes, and it now correspond to the sixth row of the B matrix. As one of the main criticism to VAR models is their susceptibility to different variable ordering, it is useful to check whether our results are robust to a change in the order of the variables. As this alternative formulation of the exchange rate equation amounts to the elimination of four restrictions, we impose an additional restriction in order to give more structure to the model. In particular, in the money market equilibrium equation we constrain coefficient b_{53} to be equal to minus one. This simply implies that the demand of real money balances is a function of the contemporaneous values of output and the short-term interest rate.

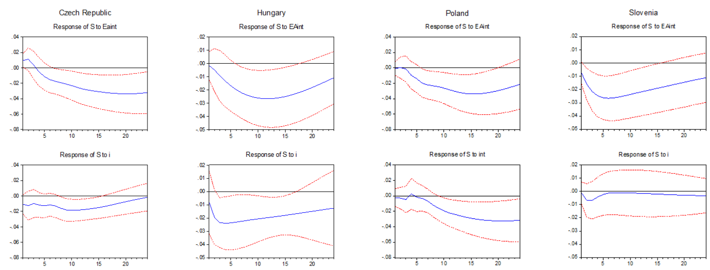
As a third alternative, we estimate the impulse-response functions by imposing a Cholesky identification scheme on the matrix B of contemporaneous coefficients. The Cholesky scheme is equivalent to imposing a recursive structure on the underlying VAR model, and results a lower triangular B matrix. Compared to the baseline model, the Cholesky scheme differs in the specification of the monetary policy equation and the money market equilibrium equation. The baseline model assumes that only the nominal exchange rate and the monetary aggregate (and the consumer price index in the Czech Republic and Hungary) contemporaneously affect the interest setting behavior of the Central Bank (cfr. equation 7). Here, the Central Banks' monetary policy actions are contemporaneously influenced by the Euro Area interest rate, the exchange rate, output and the price level. Furthermore, the money market equilibrium equation, which we defined in the baseline specification as a standard LM equation, is now poorly identified in that it includes the contemporaneous values of the Euro Area interest rate and the exchange rate.

Finally, we estimate a second Cholesky decomposition changing the ordering of the variables of the VAR. Specifically, the ordering is now the same as in alternative 2, with the exchange rate in the sixth row. Figure 3 reports the estimated dynamic responses of a one standard deviation shock to the domestic and Euro Area interest rates on stock prices for the four alternatives.

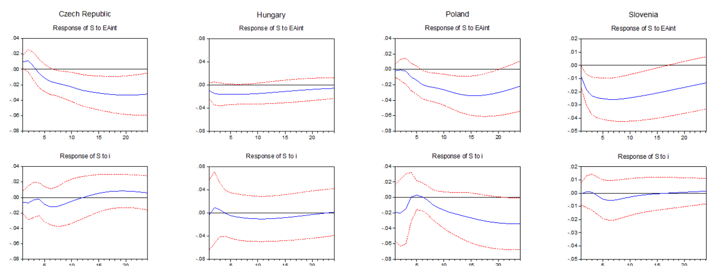
As we can see by comparing Figures 1 and 3, the responses of the stock price index to an exogenous shock to the Euro Area interest rate are robust through all specifications, both in terms of magnitude and significance. The estimated response is negative in all countries: while the negative response is negative and significant in the short-run in Slovenia, the negative effect of an unexpected foreign interest rate shock is significant in the medium or long-run in the rest

Figure 3: Responses of the Stock Price Index to a one Standard Deviation Innovation in the domestic and Euro Area monetary policy: Alternative specifications (±2 standard errors)

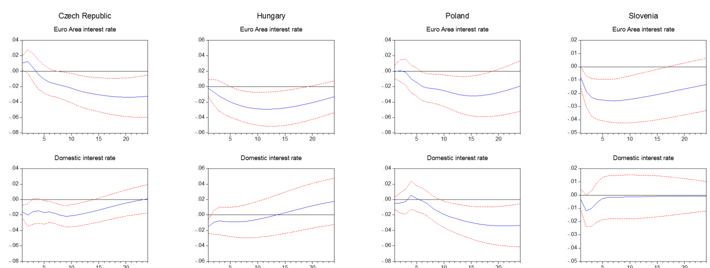
Alternative 1



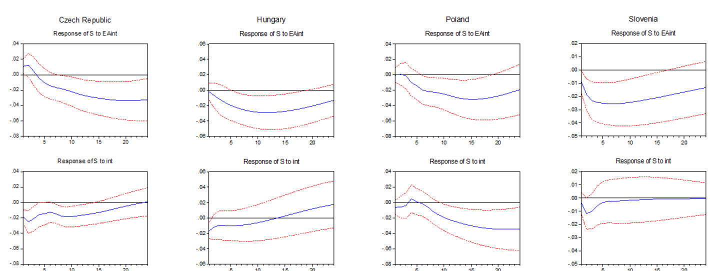
Alternative 2



Alternative 3



Alternative 4



of the countries. Concerning the responses to an unexpected, contractionary domestic monetary policy shock, we don't remark any striking differences qualitatively or quantitatively. Nevertheless, some issues of significance arise. In the first alternative specification, the estimated response in Hungary is very similar to the baseline, only the confidence bands are much wider. Despite the broadening of the confidence intervals, the negative response is still significant in the short and medium run. This does not hold in specifications 2, 3 and 4: the estimated responses are much less pronounced than in the baseline specification and are not statistically significant. Only in the last two specifications, the impact effect is significantly negative. Overall, Hungary is the country for which the estimated responses differ the most across the different model specifications. Alternative 2 seems the one generating the largest differences from the baseline and from the rest of the alternative specifications. In particular, the estimated impulse-responses to a contractionary monetary policy shock lose their significance, while conserving the same qualitative features. Nevertheless, the key features of this specification (i.e. the order of the variables and the specification of the exchange rate equation) figure also in alternative 4, which provides results very consistent with the baseline.

Overall, the set of robustness checks lead us to consider our results stable across different model specifications.

6 Conclusion

In this paper, we estimate four structural VAR models with short-run restrictions for the Czech Republic, Hungary, Poland and Slovenia. The models respresent a small-scale, small open economy model with stock prices in order to examine the interplay between financial integration, domestic monetary policy and stock prices.

The results of our analysis lead us to formulate three conclusions. First, the stock markets in the NMS are more sensitive to changes in the Euro Area interest rate than to the domestic one. This reflects the increasing financial integration between the East and West of Europe. Second, evidence on the effect of a change in domestic monetary policy on stock prices is mixed. While in Slovenia the effect is not significant at any time horizon, we find evidence of a significant negative effect in the medium and long-run in Poland and in the short and medium run in the Czech Republic. Except for Hungary, the results are robust to changes in the identification scheme and to the ordering of the variables. Hence, we feel confident to conclude that there is an effect of domestic monetary policy on stock prices in the Czech Republic and Poland, while we do not detect any in Slovenia. Given the heterogeneity of the responses in Hungary, we refrain to make any conclusion, despite two out of five specifications point towards the existence of an interplay between changes in the short-term interest rate and the stock price index. A third finding of this paper relates to the contribution of the variables in the model to the variation in stock prices. The analysis of variance revealed that overall external shocks account for the bulk of variation in the stock price index in the considered countries. While exchange rate shocks are dominant in Hungary and Poland in the short-run, shocks to the Euro Area interest rate are the main determinants in the medium to long-run in Poland, in the long-run in Hungary and during the whole time horizon in Slovenia. Hence, stock prices in the new EU member states seem to be driven by shocks related to external trade and finance.

Although this analysis is standard, in that the chosen methodology is common in the literature related to the transmission of monetary policy, it allowed us to shed a first light on an issue scarcely investigated before. The identification of some effect of monetary policy on stock prices, and of the importance of financial and trade openness open the way to further, more sofisticated research. In particular, we suggest three ways in which we could enrich our results in the future. First, we could rely on Bayesian estimation techniques in order to make our results more robust and less dependent on assumed distributions. Moreover, given the limited sample size, Bayesian techniques could help reduce the small sample bias. Secondly, we could perform a similar analysis by explicitely account for the degree of financial integration in each country, by means of the Global VAR methodology. Third, it would be interesting to investigate whether the estimated responses of the stock price index in the NMS are constant in time or evolve with the increase in financial integration. This question could be answer by a model featuring time-varying coefficients.

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Table A0: Data sources

Series	*Description*	*Source*
EAint	Euro Area money market rate (3-month)	EUROSTAT
er	nominal exchange rate: domestic currency units per euro	EUROSTAT
ip	Industrial Production Index (2005=100) Seasonally Adj.	IMF IFS
cpi	Consumer Price Index (2005=100)	IMF IFS
i	Money market interest rate (Tbill rate for Hungary)	EUROSTAT
m	M2 (HU and SI), Money + Quasi-Money (CZ and PL)	EUROSTAT and IMF IFS
s	stock price index	EUROSTAT

(32) Toma, R., 2007, "Exchange rate arrangements in Central and Eastern European countries: Evolution and Characteristics", MPRA Paper No.1967

8 Appendix A - Dataset and Variables

Figure A1: Plot of Endogenous Variables

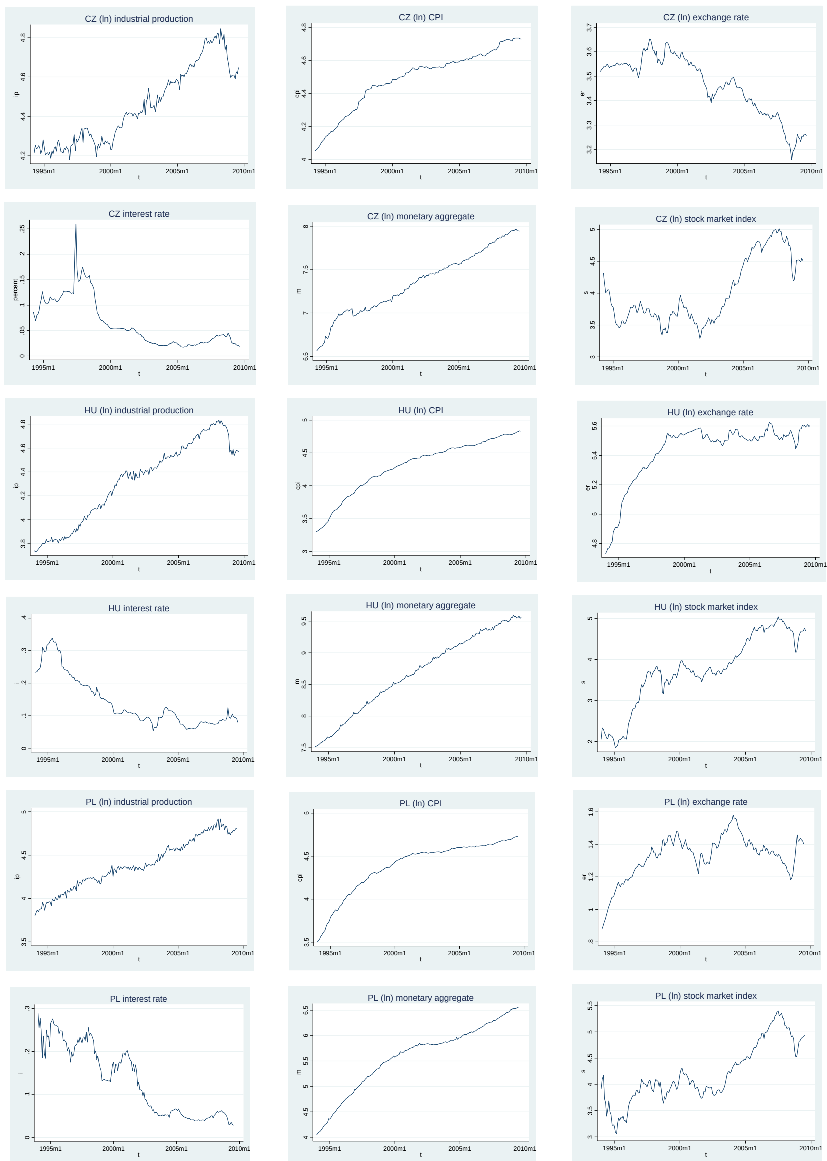

Figure A1 - continued: Plot of Endogenous Variables

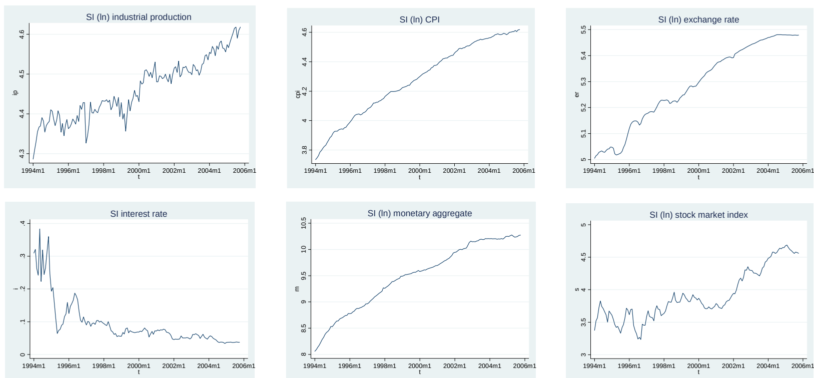

Table A1: Phillips-Perron Unit Roots Tests

	Czech Republic		Hungary	
	Specif.	PP	Specif.	PP
$EAint$	I	-3.523 (0.626)	I	-3.523 (0.626)
ip	I,T	-1.860 (0.6750)	I	-1.835 (0.3635)
cpi	I,T	-3.481 (0,0410)*	I,T	-4.243 (0.0038)*
er	I,T	-2.739 (0.2203)	I	-4.262 (0.0005)*
i	I	-5.468 (0.0000)*	I	-3.623 (0,0053)*
m	I,T	-3,680 (0,0237)*	I,T	-1.409 (0.8582)
s	I	-0.892 (0.790)	I,T	-2.562 (0.2975)
	Poland		Slovenia	
$EAint$	I	-3.523 (0.626)	I	-3.523 (0.626)
ip	I,T	-5.217 (0.0001)*	I,T	-4.933 (0.0003)*
cpi	I,T	-4.176 (0.0049)*	I	-2.806 (0.0575)*
er	I	-2.387 (0.1453)	I	-2.126 (0.2344)
i	I	-1.427 (0.5692)	I,T	-3.374 (0.0550)*
m	I,T	-3.800 (0.0165)*	I	-3.413 (0.0105)*
s	I	-1.482 (0.5424)	I	-0.613 (0.8681)

Note: I = Intercept, T = Trend. The correct specification has been chosen by visual inspection of the data and by consideration of statistical significance
Note: *Denotes rejection of the null hypothesis of unit root

Table A2: Cointegration tests

	Specification	Relevant test statistic		N. cointegrating relationships
		Trace	Max. Eigenv.	
CZ	R	12.557	11.9522	2
	F	9.863	19.8935	4
HU	R	12.365	8.264	2
	F	45.411	22.358	3
PL	R	14.811	21.612	2
	F	37.470	28.762	3-4
SI	R	49.888	14.025	1
	F	29.453	14.930	3

Note: R = Restricted, F= Full. The Full model includes all 7 endogenous variables. The Reduced model includes only the non-stationary variables resulting from Table A1
Note: Country names have been denoted as follows: CZ=Czech Republic, HU=Hungary, PL=Poland, SI=Slovenia
Note: The low number of detected cointegration relationships for Slovenia might be due to the fact that, given the limited sample size, the power of the tests might be lower.

Figure A2: Stability Check for Unrestricted VARs

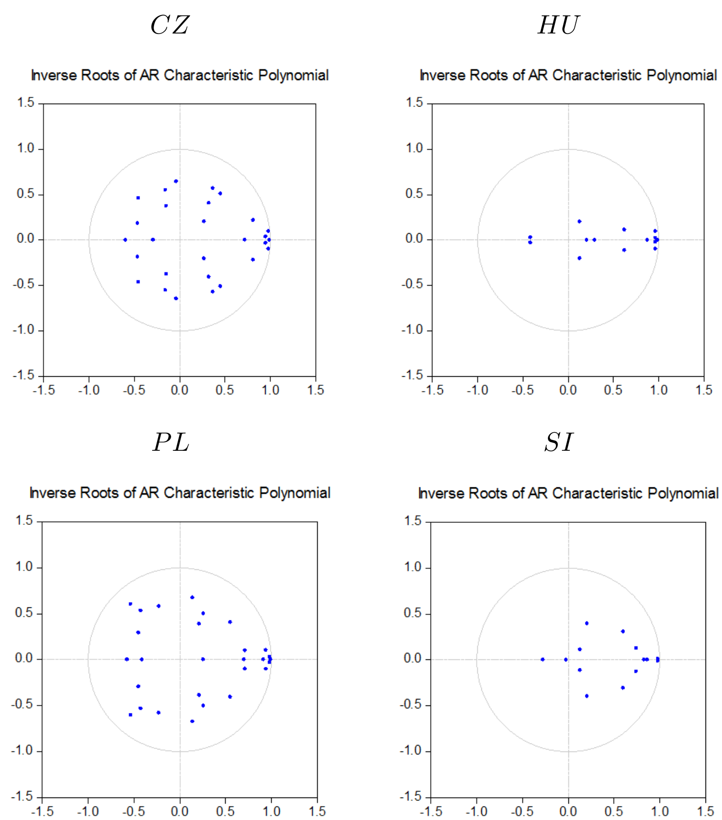

Figure A3: Czech Republic - Responses to a one Standard Deviation Structural Innovations (±2 standard errors)

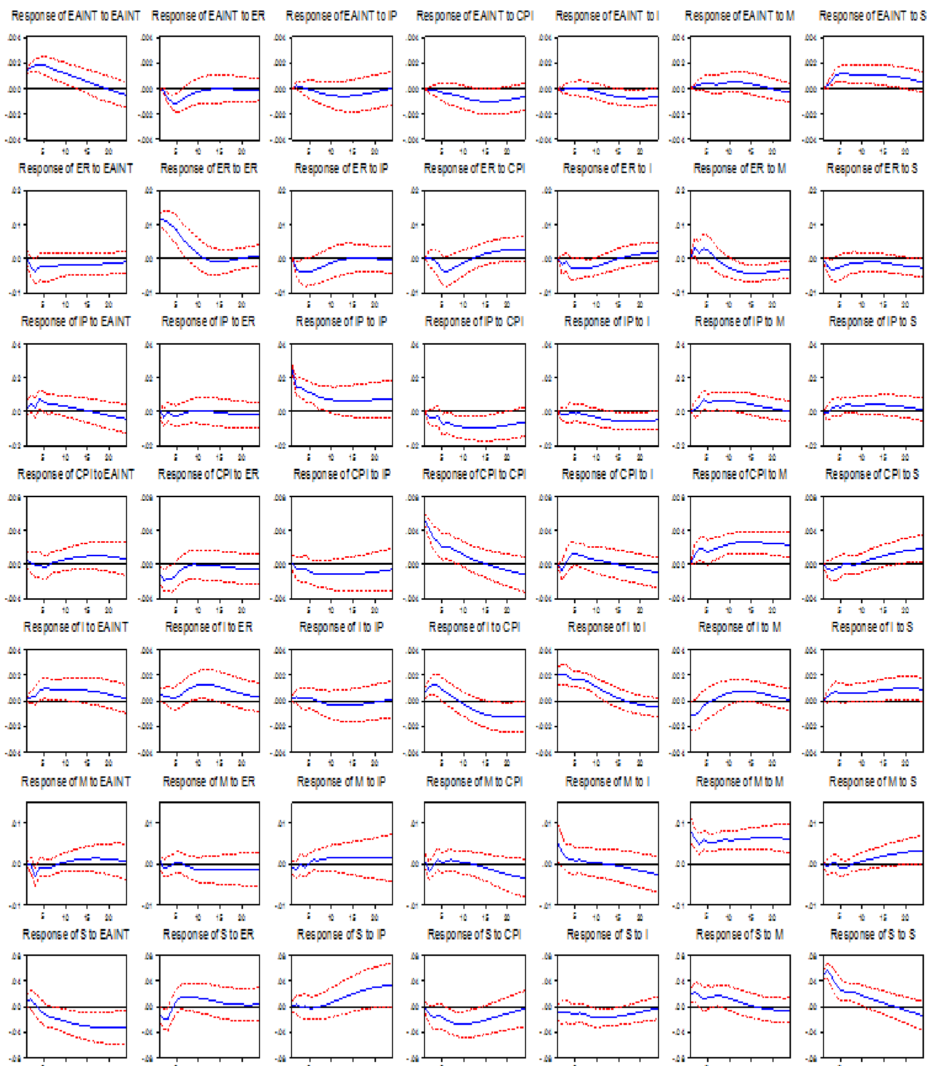

Figure A4: Hungary - Responses to a one Standard Deviation Structural Innovations (±2 standard errors)

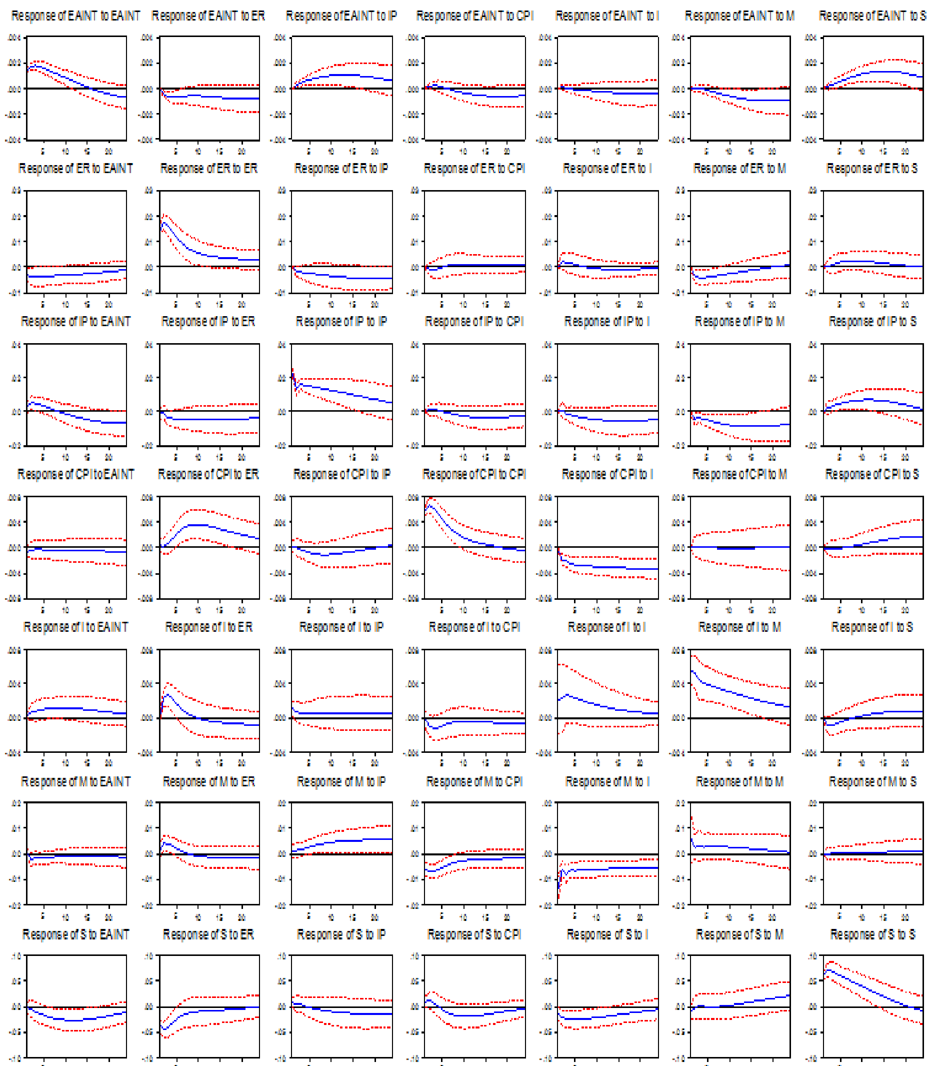

Figure A5: Poland - Responses to a one Standard Deviation Structural Innovations (± 2 standard errors)

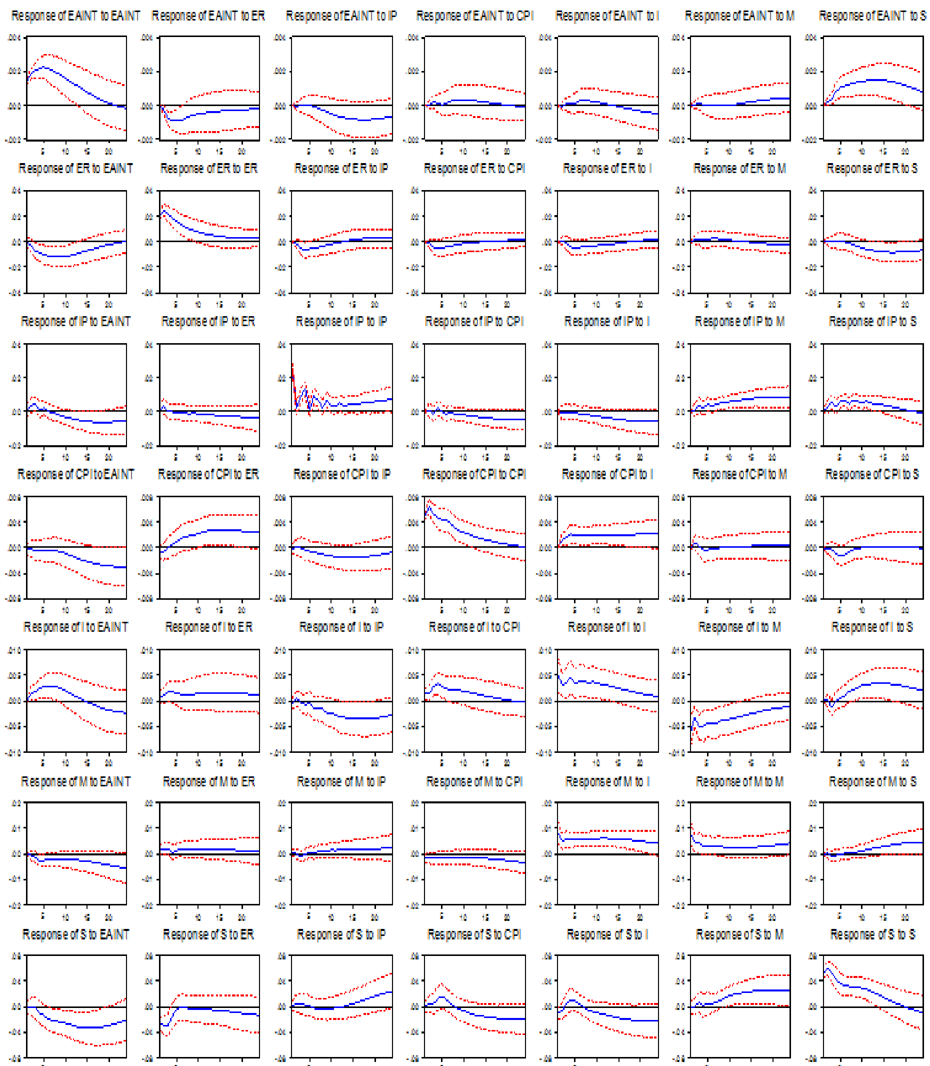

Figure A6: Slovenia - Responses to a one Standard Deviation Structural Innovations (±2 standard errors)