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ESSAYS IN FINANCE AND MONETARY POLICY: EVIDENCE FROM VISEGRAD COUNTRIES

Magdalena Morgese Borys

Dissertation

Prague, June 2009

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ABSTRACT

This dissertation consists of three empirical papers on the issues of monetary policy as well as finance in the group of four Visegrad countries, namely the Czech Republic, Hungary, Poland, and Slovakia. The first paper, entitled “Testing Multi-Factor Asset Pricing Models in the Visegrad Countries”, attempts to point to a suitable asset-pricing model that could be used to estimate the cost of equity capital in the Visegrad countries. The Capital Asset Pricing Model (CAPM) that is most often used for this purpose in developed markets has a poor empirical record and is likely not to hold in less developed and less liquid emerging markets. Various factor models have been proposed to overcome the shortcomings of the CAPM. This paper examines both the CAPM and the macroeconomic factor models in terms of their ability to explain the average stock returns using the data from the Visegrad countries. We find, as expected, that the CAPM is not able to do this task. However, factor models, including factors such as: excess market return, industrial production, inflation, money, exchange rate, exports, commodity index, and term structure, can in fact explain part of the variance in the Visegrad countries’ stock returns.

A second paper, “Size and Value Effects in Visegrad Countries”, is an extension of the previous paper. This paper has two main objectives. The first is to test for the presence of the size and book-to-market value effects in Visegrad countries, while the second is to propose a plausible model for the cost of capital estimation in the Visegrad region. Size and book-to-market effects have been found in the United States and many other developed stock markets. We demonstrate that these effects do in fact explain the expected return/cost of capital in Eastern Europe. Based on this result, we proceed by constructing regional size and book-to-market portfolios for a combined Visegrad market. Returns on these portfolios serve as factors in addition to the market portfolio. The regional three-factor model performs as well as country specific versions of the model. However, it can be estimated for a more current sample in Prague, Warsaw, Budapest, and Bratislava. Therefore it is a plausible model for the cost of capital in this region and we use it to calculate the cost of capital for the following industries: banks; capital goods; food, beverage and tobacco; materials; and utilities.

The final third paper (a joint work with R. Horvath), “The Effects of Monetary Policy in the Czech Republic: An Empirical Study”, examines the effects of Czech monetary policy on the economy within the VAR, structural VAR, and factor-augmented VAR frameworks. We document a well-functioning transmission mechanism similar to the euro area countries, especially in terms of the persistence of monetary policy shocks. Subject to various sensitivity tests, we find that a contractionary monetary policy shock has a negative effect on the degree of economic activity and the price level, both with a peak response after one year or so. Regarding prices at the sectoral level, tradables adjust faster than non-tradables, which is in line with microeconomic evidence on price stickiness. There is no price puzzle, as our data come from a single monetary policy regime. There is a rationale in using the real-time output gap instead of current GDP growth, as using the former results in much more precise estimates.

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The Effects of Monetary Policy in the Czech Republic: An Empirical Study*

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Abstract

In this paper, we examine the effects of Czech monetary policy on the economy within the VAR, structural VAR, and factor-augmented VAR frameworks. We document a well-functioning transmission mechanism similar to the euro area countries, especially in terms of the persistence of monetary policy shocks. Subject to various sensitivity tests, we find that a contractionary monetary policy shock has a negative effect on the degree of economic activity and the price level, both with a peak response after one year or so. Regarding prices at the sectoral level, tradables adjust faster than non-tradables, which is in line with microeconomic evidence on price stickiness. There is no price puzzle, as our data come from a single monetary policy regime. There is a rationale in using the real-time output gap instead of current GDP growth, as using the former results in much more precise estimates. The results indicate a rather persistent appreciation of the domestic currency after a monetary tightening, with a gradual depreciation afterwards.

Keywords: monetary policy transmission, VAR, real-time data, sectoral prices
JEL Codes: E52, E58, E31

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1 Introduction

Understanding the transmission of monetary policy to inflation and other real economic variables is of key importance if central bankers are to conduct monetary policy effectively. Not surprisingly, there is extensive theoretical as well as empirical literature studying the effects of monetary policy shocks on real economy aggregates and prices. For a small open economy such as the Czech Republic, it is vital to analyze monetary policy transmission for several reasons. First, there is somewhat mixed evidence regarding monetary policy transmission, as many studies estimate standard vector autoregression (VAR) models, mixing data from two distinct policy regimes, i.e., from the fixed exchange rate regime under which the Czech National Bank (CNB) conducted its policy until May 1997, and from the inflation targeting regime that was adopted in January 1998.[1] Not surprisingly, the identification of monetary policy shocks then becomes somewhat cumbersome and *all* these studies exhibit the price puzzle (see Table 1 in the results section).

Therefore, it is worthwhile to update previous results reflecting the monetary policy regime changes, to utilize a wider range of econometric techniques and, on top of that, to incorporate real-time and forward-looking variables into the VAR analysis. To our knowledge, real-time data has not been applied to study monetary transmission in the Czech Republic. This is in a sense paradoxical, as an important feature of monetary policy conduct is that it is based on the information set available at the time of policy-making. This implies that using *ex-post* revised data (note that these are typically more precise, but are not available at the time of monetary policy action) may contaminate the estimated effects of monetary policy (Croushore and Evans, 2006). The revisions are typical for output data.[2]

There is also no empirical evidence about monetary policy effects on sectoral prices. This is striking, because tradable prices in a small open economy may be driven to a large extent by international factors that domestic monetary policy is unlikely to affect. Our prior assumption is that as non-tradable prices are typically less exposed to international competition and more labor-intensive, the reaction of non-tradable prices is likely to be more

[1] See Coats, Laxton, and Rose (2003) and Kotlán and Navrátil (2003) for an overview of Czech monetary policy.
[2] We therefore utilize the real-time estimates of the output gap available from the Czech National Bank (CNB). Using the central bank output gap is advantageous for monetary policy shock identification, as the central bank conducts its policy based on *its* estimate of the degree of economic activity, not the estimates of other institutions or individuals. Note that price indices are not revised *ex post* by the Czech Statistical Office. An additional rationale for using the output gap is that in an environment of changing potential growth of the economy, as is the case in our sample, actual GDP growth does not necessarily give an accurate picture about the degree of economic activity.

persistent (see e.g. Barro, 1972 and Martin, 1993 for models relating the degree of competition to price rigidity).[3]

In this paper, we examine the effects of monetary policy within the vector autoregression (VAR), structural VAR (SVAR), and factor-augmented VAR (FAVAR) frameworks during the inflation targeting period in the Czech Republic. More specifically, we focus on assessing the persistence and magnitude of monetary policy shocks on output (including the real-time output gap), prices (at both the aggregate and sectoral level) and the exchange rate, controlling for a standard set of factors.

The paper is organized as follows. Section 2 discusses the related literature. The data are presented in section 3. Section 4 is focused on identification issues. Section 5 contains our results on the effects of monetary policy. We present our conclusions in section 6, and an appendix follows.

2 Related VAR Literature

Vector autoregressions (VARs), as introduced by Sims (1980), are considered to be benchmarks in econometric modeling of monetary policy transmission. It has been argued that this class of models provides a certain mix between a mere "data-driven" approach and an approach coherently based on economic theory (see Fry and Pagan, 2005 on the application of VARs for macroeconomic research). In terms of monetary policy analysis, the VAR methodology has been further developed among others by Gerlach and Smets (1995), Leeper, Sims, and Zha (1998), and Christiano, Eichenbaum, and Evans (1999). This last study provides a detailed review of the literature on this topic in the United States. Similarly, there has been extensive research undertaken in Europe to study various aspects of monetary transmission in the euro area countries (see Angeloni, Kashyap, and Mojon, 2003). The research on monetary transmission in the euro area either focuses on euro area-wide analysis (Peersman and Smets, 2001) or studies specific countries in detail (Mojon and Peersman, 2001).

The economic theory suggests that output and prices should temporarily fall after a monetary contraction. Nevertheless, as regards prices, a number of papers document that, on the contrary, prices rise after a monetary contraction. This effect has been labeled as the

[3] The negative link between the degree of competition and price rigidity is also documented empirically using microeconomic data at the price-setter level by Alvarez and Hernando (2006) for the euro area and Coricelli and Horvath (2006) for Slovakia.

"price puzzle." The literature typically argues that the price puzzle is a consequence of some model misspecification (Brissimis and Magginas, 2006 and Giordani, 2004). Meanwhile, Barth and Ramey (2001) suggest that a fall in both prices and output would indicate that monetary policy affects the economy mainly through the demand channel. On the other hand, falling output and rising price levels would point to the prevalence of the supply or cost channel.[4]

In addition, the literature examines the effect of monetary policy on exchange rate behavior. Generally, an immediate exchange rate appreciation after a monetary tightening and then a gradual depreciation of the domestic currency is expected according to uncovered interest rate parity. However, the empirical evidence is again somewhat mixed. Some authors find a rather persistent appreciation of the domestic currency ("delayed overshooting", Eichenbaum and Evans, 1995), while others report that the exchange rate actually depreciates with a monetary contraction and provide explanations for the so-called exchange rate puzzle (Kim and Roubini, 2000).

A number of approaches to dealing with model misspecification related to monetary policy shock identification have been stressed in the literature. For example, Brissimis and Magginas (2006) show that by adding forward-looking variables such as federal funds futures to a standard VAR specification, one is able to obtain responses to monetary policy that are consistent with the theory. The rationale for including federal funds futures is that they contain market expectations about future monetary policy action (this expectation element may also be found in commodity prices or money, to a certain extent).

In addition, Croushore and Evans (2006) emphasize the role of data revisions for monetary policy shock identification. Monetary policy makers react to the information set available at the time they make their decision, and it is often the case that GDP data are revised afterwards. As a result, using *ex-post* GDP data series may contaminate the estimated monetary policy effects. Also, monetary policy makers often tend to react to the output gap rather than GDP growth. In addition, Giordani (2004) shows that using the output gap instead of GDP growth alleviates the price puzzle. These concerns are especially appealing in our case. First, the CNB's main forecasting model (the so-called Quarterly Projection Model) indeed contains an output gap in its reaction function (Coats *et al.*, 2003). Second, GDP growth may still be useful as a measure of the degree of economic activity if potential output

[4] If a firm has to borrow to finance its production, interest rates enter its cost function. Consequently, a monetary policy tightening increases the firm's costs, to which the firm may react by increasing the price of the products it sells. In consequence, this argument suggests that the price puzzle does not have to be caused by model misspecification. In general, see Coricelli *et al.* (2006) for more specific explanations of the price puzzle.

growth is not changing much. However, in the case of the Czech economy, it is estimated that potential output growth sharply increased from some 2% in 1998 to around 5.5% in 2005 (Dybczak, Flek, Hájková, and Hurník, 2006).

Next, there has been a lot of research focusing on the sensitivity of the responses of aggregate variables such as aggregate inflation and output to monetary policy within the VAR framework. However, much less is known about the responses to monetary policy at the more disaggregated level. Erceg and Levin (2006) find that the durable goods sector is more sensitive to interest rate changes than the non-durable goods sector in the U.S. Based on this empirical finding, they investigate the impact of monetary policy on these two industries and find, as expected, that monetary policy effects are much stronger in the durable goods industry. Dedola and Lippi (2005) study the responses to monetary policy of various industrial sectors for a number of OECD countries. They find that the responses vary between sectors in terms of their magnitude and persistence. This result is confirmed by Peersman and Smets (2005), who find a number of significant differences between various industries in the euro area in terms of both the magnitude of the output response as well as the asymmetry of the responses over the business cycle.

Bouakez *et al.* (2005) is one of the few studies that examine the impact of monetary policy on disaggregate prices. Their results suggest that monetary transmission affects household consumption in the construction and durable manufacturing sectors the most, but the impact of a monetary policy shock vanishes relatively quickly. They also find significant differences between the sectors' inflation in terms of variance decomposition, volatility, and persistence. Bouakez *et al.* (2005) find that the response of services inflation to monetary policy shocks is relatively pronounced and also the most persistent. Boivin *et al.* (2007) study the effect of macroeconomic fluctuations on disaggregate prices within the factor-augmented VAR framework. Among other things, their results indicate that the degree of market power explains the diversity of the responses of disaggregate prices to monetary policy shocks.

Several papers study the monetary policy effects for the Czech Republic within the VAR framework.[5] Using the sample period after the adoption of inflation targeting (1998–2004), Hurník and Arnoštová (2005) find that prices respond with a peak around 5–6 quarters after a shock, although there is some evidence for a price puzzle in the first two quarters after the shock. Output falls after a monetary contraction, with a peak after one year or so. There is a delayed overshooting of the exchange rate, as it depreciates only some 4 to 5 quarters after

[5] See Coricelli, Égert, and MacDonald (2006) for a survey of the current findings on monetary policy transmission in Central and Eastern Europe, including those undertaken within the VAR framework.

the monetary policy innovation. Extending the sample back to 1994, when the fixed exchange rate regime was in use, yields less satisfactory results, as it is obviously more difficult to identify monetary policy shocks across two monetary policy regimes. In our paper, we use a similar, slightly extended time horizon (after the adoption of inflation targeting), but we opt for monthly rather than quarterly data. In addition, in our paper we include the real-time output gap in the benchmark specification, as opposed to the ex-post revised GDP used by Hurník and Arnoštová (2005). The effects of a monetary policy contraction estimated in our paper are largely in line with the responses observed in more developed economies and countries in the Eurozone, in particular. Contrary to Hurník and Arnoštová (2005), we do not find evidence for a price puzzle in the Czech Republic.

Next, there are a number of papers analyzing and comparing the effects of monetary policy in groups of Central and Eastern European countries vis-à-vis other, more advanced economies (Creel and Levasseur, 2005; Darvas, 2006; European Forecasting Network, 2004; Héricourt, 2005). Many studies find evidence of price and/or exchange rate puzzles for the Czech Republic. As argued by Coricelli *et al.* (2006), the price puzzle is generally avoided in studies that allow for changes in coefficients and in papers employing more sophisticated identification schemes. As we argue below, the price puzzle in these studies often arises because monetary policy regime changes are ignored. In our paper we consider only the period after the change of monetary policy regime, characterized by stable coefficients (as assessed by the estimation of recursive coefficients).

Among the studies that do not find evidence of a price puzzle, Jarocinski (2006) provides a Bayesian VAR analysis of monetary policy effects in Western and Central Europe. Interestingly, Jarocinski finds that monetary policy is more potent in Central Europe, despite a lower level of financial development and smaller indebtedness. Regarding the Czech Republic, he uncovers that there is a relatively strong appreciation of exchange rates as well as a larger price decline after a monetary policy innovation, as compared to other Central European countries. Elbourne and Haan (2006) study the interactions between the financial system and monetary transmission within the structural VAR framework for a group of ten Central and Eastern European countries. For the Czech Republic, they find a hump-shaped response of prices, an exchange rate appreciation, and a fall in industrial production after a monetary policy innovation. Next, financial structure is found to be of little importance for monetary transmission.

3 Data

This section contains a description of our dataset. We restrict our sample to the data from 1998 onwards, i.e., since the inflation targeting framework was adopted by the Czech National Bank (until May 1997 it had operated a fixed exchange rate regime). Our sample thus spans from 1998:1 to 2006:5 at monthly frequency. While studies in this stream of literature often employ quarterly data, given the length of our sample we decided to work at a monthly frequency. As a result, we have 101 observations. The source of our data is the CNB's public database ARAD (except for the output gap, which is only available internally within the CNB). The plots of all the series are available in Appendix 1.

We use GDP, $lgdp_t$, and the real-time output gap estimate, $outputgapreal_t$, as measures of economic activity.[6] GDP is traditionally used for this kind of exercise, but Giordani (2004) suggests using the output gap. In addition, by using the real-time output gap estimate we avoid the risk resulting from the use of *ex-post* data, which are not available to central bankers at the time of monetary policy formulation (Croushore and Evans, 2005). As GDP and the output gap are only available at quarterly frequency, we interpolate these two using the quadratic-match average procedure.[7] Note that all the other variables we use are not revised afterwards.

Next, we employ the net price index, $lnet_t$ (the net price index is the consumer price index excluding regulated prices). For our disaggregate analysis, we employ the tradable price index, $tradable_t$, and the non-tradable price index, $nontradable_t$. Note that the individual components underlying the consumer price indices are grouped into tradables and non-tradables categories in line with the internal CNB classification.

Further, the nominal CZK/EUR exchange rate, $lexrate_t$, and the three-month interbank interest rate (3M PRIBOR[8]), $pribor_t$, are used. To capture external developments, the 1-year EURIBOR, $euribor_t$, and the commodity price index, $lcommodity_t$, are utilized. The forward rate agreement rate (9*12 FRA rate), fra_t, is used to bring in an additional forward-looking element. Given that there are no futures or forwards in the Czech Republic that are directly

[6] See Coats *et al.* (2003, chapter 5) on the construction of the output gap used by the CNB. The output gap is the difference between actual and potential output, where the latter is estimated by a multivariate filter, more specifically by the Kalman filter procedure, where the system of equations is in the state-space representation.

[7] We admit that interpolation introduces information not available at the time of policy making.

[8] The actual monetary policy instrument of the CNB is the 2W repo rate. Since the repo rate is not changed continuously and is censored, we opt for the 3M PRIBOR, which is very closely linked to the 2W repo rate; its correlation stands at 0.998 in our sample. In addition, the 3M PRIBOR may capture central bank communication. See Horvath (2008) for a discussion related to the use of the monetary policy rate vs. the interbank market rate in the Czech Republic.

linked to the monetary policy rate (2W repo) as is the case in the U.S., we decided to use forwards on interbank rates, which are very closely related to the policy rate. Finally, all data are in logs except interest rates and the real-time output gap.

4 Identification

In this section, we discuss the VAR framework we adopt. The choice of variables for our VAR model is largely motivated by an open economy New Keynesian model (see for example Gali and Monacelli, 2005). The main equations of this class of models are aggregate demand, the Phillips curve, the monetary policy rule, and uncovered interest rate parity.

We estimate two benchmark models and then undertake a sensitivity analysis. The difference between these two benchmark models is that the first includes only the aggregate price index, while the second distinguishes between the tradable and non-tradable price indices. The specification of the first baseline model is the following:

$$Y_t = A(L)Y_{t-p} + B(L)X_t + u_t , \tag{1}$$

where Y_t and X_t represent endogenous and exogenous variables,[9] respectively. The data vectors are $Y_t = \{outputgapreal_t, lnet_t, pribor_t, lexrate_t\}$ and $X_t = \{euribor_t, lcommodity_t, fra_t\}$. For our second benchmark specification, $Y_t = \{outputgapreal_t, lnontradable_t, tradable_t, pribor_t, lexrate_t\}$ and X_t remains the same.

The VAR specification in (1) represents a so-called reduced-form equation. In order to identify the original shocks we can apply the recursiveness assumption by imposing restrictions on a matrix linking the structural shocks to the reduced-form disturbances. The variables are ordered in a specific way so as to represent the assumption that the monetary authorities choose the interest rate taking into account the current level of prices and output (as in Mojon and Peersman, 2001). In addition, the output gap and prices are assumed not to react immediately to the monetary policy shock, but rather with a one-period lag. Mojon and Peersman (2001) follow a recursive specification to analyze the impact of a monetary policy shock in some of the euro area countries.

[9] The inclusion of foreign variables that are considered exogenous is motivated by the need to control for foreign shocks and thus not to confuse domestic monetary shocks with the central bank's responses to external developments (Jarocinski, 2006).

We analyze the sensitivity of our benchmark models first by using GDP instead of the output gap, second by estimating a very parsimonious model without exogenous variables, and third by estimating the baseline models by structural VAR instead of recursive VAR.

As regards the first sensitivity check, actual GDP data are used instead of the output gap. The rationale for this exercise is that the output gap, as opposed to GDP, is unobservable. Our second sensitivity check is motivated by degrees-of-freedom considerations. Here, we assume that external shocks influence the Czech economy only via the exchange rate (i.e., *B(L)*=0). Admittedly, this is a simplistic specification, but its main advantage is its limited number of variables and thus its greater degree of freedom in comparison to our other models. As the third robustness check, the two baseline models are estimated by structural VAR (SVAR). SVAR represents an alternative identification scheme in order to recover the original residuals from the reduced-form VAR. For structural VAR, we apply here the AB-model of Amisano and Giannini (1997), which is defined as follows in a reduced form:

$$Y_t = A^*(L)Y_{t-p} + B^*(L)X_t + u_t, \qquad (2)$$

$u_t = A^{-1}Be_t, e_t \sim (0, I_K),$

where I is the identity matrix and K is the number of variables. A and B are $k \times k$ matrices to be estimated. In the case of our first benchmark model, they are specified as follows.

$$A = \begin{bmatrix} 1 & 0 & 0 & 0 \\ a_{21} & 1 & 0 & 0 \\ a_{31} & 0 & 1 & a_{34} \\ a_{41} & a_{42} & a_{43} & 1 \end{bmatrix} \qquad B = \begin{bmatrix} b_{11} & 0 & 0 & 0 \\ 0 & b_{22} & 0 & 0 \\ 0 & 0 & b_{33} & 0 \\ 0 & 0 & 0 & b_{44} \end{bmatrix}.$$

It follows from matrix A that a forward-looking monetary authority does not consider contemporaneous prices while deciding on monetary policy (i.e., $a_{32}=0$). However, monetary authorities are likely to react to contemporaneous output (a_{31}, as output can be regarded as an excess demand pressure indicator) and exchange rate shocks (a_{34}), which is a reasonable assumption for small open economies according to Kim and Roubini (2000). More specifically, exchange rate fluctuations influence the inflation forecast if they are deemed not to be transitory.

For our second benchmark model, in which we consider disaggregate prices (hence five variables), matrices A and B look as follows:

$$A = \begin{bmatrix} 1 & 0 & 0 & 0 & 0 \\ a_{21} & 1 & 0 & 0 & 0 \\ a_{31} & a_{32} & 1 & 0 & 0 \\ a_{41} & 0 & 0 & 1 & a_{45} \\ a_{51} & a_{52} & a_{53} & a_{54} & 1 \end{bmatrix} \qquad B = \begin{bmatrix} b_{11} & 0 & 0 & 0 & 0 \\ 0 & b_{22} & 0 & 0 & 0 \\ 0 & 0 & b_{33} & 0 & 0 \\ 0 & 0 & 0 & b_{44} & 0 \\ 0 & 0 & 0 & 0 & b_{55} \end{bmatrix}.$$

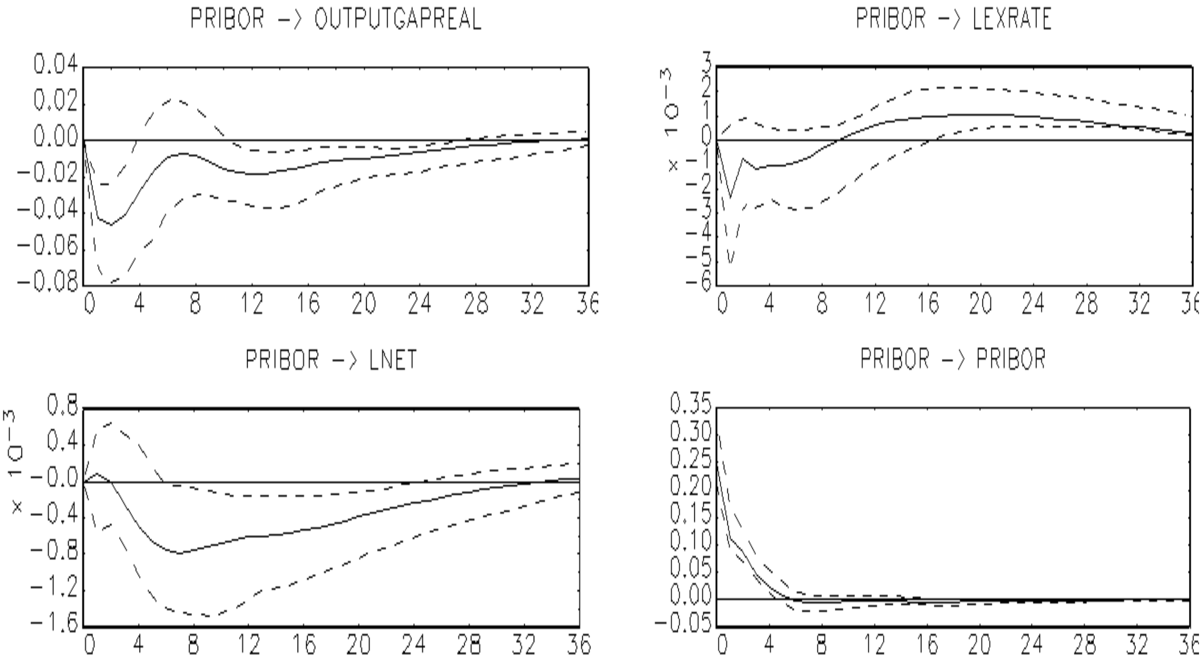
Following each VAR estimation, we perform stability checks in order to ensure the robustness of our results (the results of these tests are available upon request). It is important to note that the variables used in the VAR analysis do not need to be stationary. Sims (1980), among others, argues against differencing even if the series contain a unit root. The main goal of the VAR analysis is to analyze the co-movements in the data. What matters for the robustness of the VAR results is the overall stationarity of the system (see Lütkepohl, 2006 for details). A description of the FAVAR model is presented in Appendix 3.

5 Results

In this section, we discuss the estimated effects of Czech monetary policy within the aforementioned specifications. The number of lags has been chosen according to the Schwartz criterion and the parameter stability addressed by the CUSUM and CUSUM of squares tests and the recursive coefficient estimation (the results are available upon request).

Figure 1 presents our results regarding the effects of a contractionary monetary policy shock on several economic variables of interest to a monetary authority. These figures contain the impulse responses and the associated 95% confidence interval, which was bootstrapped using 1,000 replications according to the percentile method by Hall (1988).

Figure 1: Contractionary monetary policy shock, impulse responses



Notes: This figure shows the impulse responses to a one standard deviation contractionary monetary policy shock. Time (on the horizontal axis) is measured in months.

We find that prices fall after a monetary tightening and bottom out after one year or so. This is in line with the targeting horizon of the CNB, which is considered to be between 12 and 18 months. In terms of magnitude, our results show that a one Cholesky standard deviation of interest rates (a 30 basis point monetary policy shock) decreases the log of prices by about 0.1%.[10] Notably, there is essentially no evidence for a price puzzle.

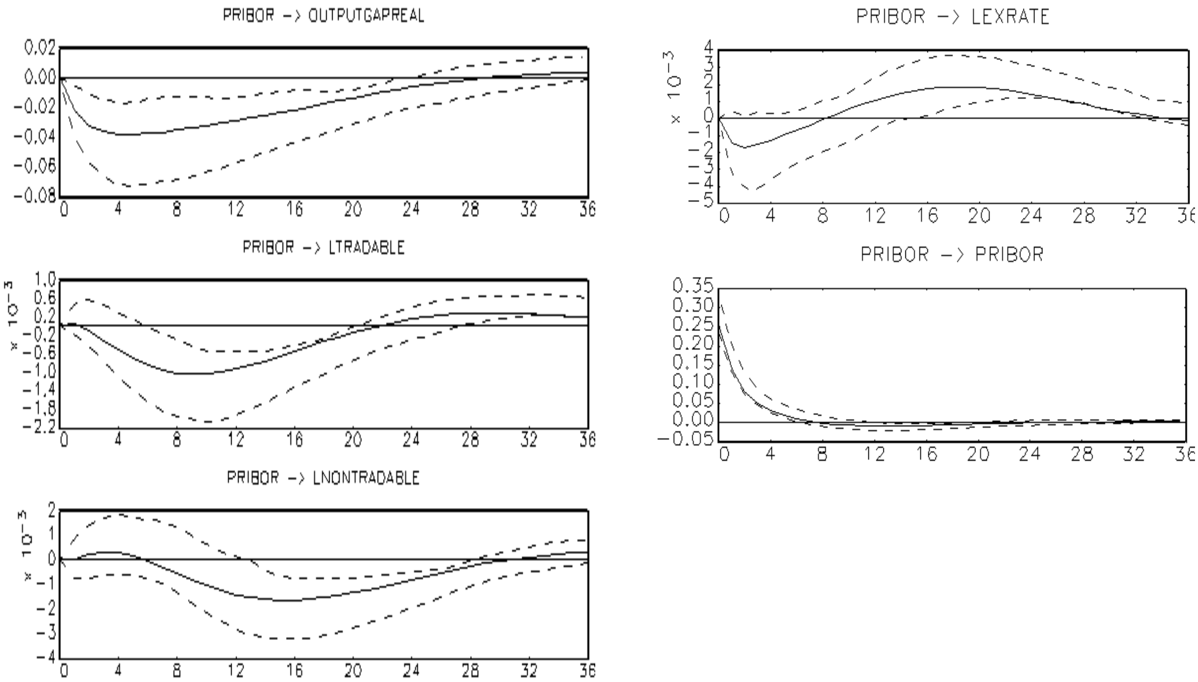
The degree of economic activity, as measured by the output gap, falls after a contractionary monetary policy shock, bottoming out after about four months (this, however, is not confirmed in our sensitivity analysis, which identifies the bottom after about twelve months, which is more sensible). The results indicate that a monetary shock of 30 basis points decreases the output gap by about 5%. The responses of output and prices to a monetary shock show no support for the cost channel of monetary policy.

Next, our results show a delayed overshooting in exchange rate behavior, i.e., a rather persistent appreciation of the domestic currency after a monetary tightening (lasting typically about 6 months) and a gradual depreciation afterwards. However, it has to be pointed out that the estimated confidence intervals are relatively wide, which brings some margins of uncertainty into interpreting the results. Nevertheless, we can see that irrespective of

[10] Several authors have raised the question of the accuracy of monetary policy shocks within VARs. See Boivin and Giannoni (2002) for a related discussion.

specification and estimation technique, the exchange rate depreciates over the longer term, which conforms to the uncovered interest rate parity hypothesis (Kim and Roubini, 2000).

Figure 2: Contractionary monetary policy shock, impulse responses: Tradable vs. non-tradable prices



Notes: This figure shows the impulse responses to a one standard deviation contractionary monetary policy shock. Time (on the horizontal axis) is measured in months.

Figure 2 contains the estimates of the effect of monetary policy shocks on tradable and non-tradable prices. Generally, tradable prices react faster than non-tradable prices to a monetary contraction. While the bottom response of tradable prices is at one year or so (even 9–10 months), the bottom response of non-tradable prices occurs only after one and a half years. This result matches the findings based on micro-level data (Alvarez and Hernando, 2006; Coricelli and Horvath, 2006), which show that the frequency of non-tradable price changes is lower (and negatively affected by the degree of competition); hence, a slower response to the monetary policy shock is to be expected. On the other hand, the reaction of non-tradable prices is more pronounced. A monetary shock of about 0.3% decreases tradable and non-tradable prices by 0.1% and 0.2%, respectively. In addition, the results in Figure 2 largely confirm the results of the effect of monetary policy on output and the exchange rate from Figure 1.

Next, we analyze the sensitivity of our benchmark models, with all the results reported in Appendix 2. First, we investigate how our results change when we include ex-post revised data (GDP) instead of the real-time output gap in our data vector. Real-time variables are part of the information set available at the time of policy-making, so by using these variables in the VAR analysis we avoid the likely contamination of the results caused by data revisions.[11] There is no statistically significant reaction of GDP to the monetary shock and it seems that GDP does not capture adequately the degree of demand pressures in an environment of sharply changing potential output growth. Thus, our results stress the importance of using the real-time output gap in the VAR specification, as it improves the precision of the empirical analysis.

Second, we estimate a very parsimonious model without exogenous variables, including the forward-looking component. The rationale behind this is merely degrees-of-freedom considerations. Interestingly, we find that a four-variable VAR is able to generate quite sensible and precisely estimated impulse responses.[12] This would suggest that economic agents during our sample period form their expectations in a rather backward-looking manner. This is somewhat surprising, but one has to consider the transition process of the Czech economy and the corresponding greater uncertainty in economic development, which could make agents rely more on current data than on forecasts.

Finally, we also estimate the benchmark models by structural VAR instead of recursive VAR, but SVAR seems to provide little value added and typically generates impulse responses close to those of VAR, but with much larger confidence intervals.

Next, we compare our results with other recent studies that analyze monetary policy shocks in the Czech Republic within the VAR approach. The comparison is summarized in Table 1. Most of the existing studies ignore the monetary policy regime change in the Czech Republic (the fixed exchange rate regime until May 1997 and the adoption of inflation targeting in January 1998). Consequently, it is not surprising that simple VAR methods have difficulty in identifying monetary policy shocks across these two regimes, i.e., they do not deliver plausible results and all exhibit the price puzzle (some of them even report a positive reaction of output to a monetary tightening).[13] This suggests that the price puzzle in these

[11] In general, the output gap should be a better measure of demand pressures (especially when potential output growth is changing), but one should keep in mind that it is unobservable and thus subject to greater uncertainty.

[12] The output gap and prices fall after a contractionary monetary policy shock, bottoming out after about twelve months. The exchange rate first appreciates, but later depreciates significantly, in line with uncovered interest rate parity (see also Eichenbaum and Evans, 1995). The results on the reaction of tradable and non-tradable prices largely comply with the benchmark case, except that non-tradable prices reach their bottom response a bit later (about two years).

[13] The exemption is Jarocinski (2006). His sample starts in June 1997, which is before the adoption of inflation targeting, but after the exchange rate turbulence and the abandonment of the fixed exchange rate regime. As a result, we code his sample in Table 1 as coming from a single monetary policy regime. Another approach to dealing with monetary policy changes is presented by Darvas (2005), who estimates a time-varying coefficient VAR. Indeed, his results suggest that the values of the estimated parameters change rather abruptly around 1997

studies is associated with the monetary policy regime change. This is further confirmed by two papers that employ data from the inflation targeting period (Elbourne and Haan, 2006 and this paper), as their results do not exhibit the price puzzle. Finally, the results in Table 1 indicate that the bottom responses of output and prices seem to be at around 4 quarters, which is in line with our findings.

Table 1: Comparison to other VAR studies on monetary transmission in the Czech Republic

	Sample period	Single monetary policy regime	Estimation technique	Reaction of output to MP shock	Reaction of prices to MP shock	Bottom reaction of output and prices
EFN (2004)	1994–2003	No	VAR	(-), sig.	(+), sig.	6Q/---
Ganev et al. (2004)	1995–2000	No	VAR	(+), n.a.	(+), n.a.	----
Creel and Levasseur (2005)	1993–2004	No	SVAR	(+), sig.	(+), sig.	----
Darvas (2005)	1993–2004	No	TVC-SVAR	(-), n.a.	n.a.	4Q/n.a.
Héricourt (2005)	1995–2004	No	VAR	(-), sig.	(+), sig.	1Q/---
Hurník and Arnoštová (2005)	1994–2004	No	VAR	insig.	insig.	8Q/6Q
Elbourne and Haan (2006)	1998–2004	Yes	SVAR	(-), sig.	(-), sig.	4Q/4Q
Jarocinski (2006)	1997–2004	Yes	Bayesian VAR	(-), sig.	(-), sig.	4Q/4Q
Gavin and Kemme (2007)	1995–2006	No	SVAR	(-), sig.	(+), sig.	----
Anzuini and Levy (2007)	1993–2002	No	VAR, SVAR	(-), sig.	insig.	4Q/8Q
This paper	1998–2006	Yes	VAR, SVAR, FAVAR	(-), sig.	(-), sig.	3Q/4Q

Note: (-) and (+) denote a statistically significant decline and increase, respectively, of the variable after a monetary policy shock. The column "Single monetary policy regime" indicates whether the sample period of the study comes from a single monetary regime or spans different regimes (the fixed exchange rate regime until May 1997 and the inflation targeting regime adopted in January 1998). Abbreviations: TVC-SVAR – time-varying coefficient SVAR, Sig. – the reaction of the variable to a monetary policy shock is statistically significant at the 5% level, and Q – quarters. If the reaction of the variable to a monetary shock does not have the correct sign, the bottom reaction of the variable is not reported (denoted as "----" in the table; n.a. indicates that the corresponding estimates were not available in the original study).

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and remain relatively stable afterwards. (This is confirmed in this study by the recursive estimation of the parameters. The results are available upon request.)

6 Concluding Remarks

In this paper, we analyze the transmission of monetary policy shocks in the Czech Republic within the VAR, SVAR, and FAVAR frameworks. In general, monetary transmission in the Czech Republic seems to be similar, in terms of the persistence of the responses of economic variables to monetary shocks, to that in more developed countries, including the euro area (see e.g. Mojon and Peersman, 2001).

All in all, subject to various sensitivity tests, we find that prices and output decline after a monetary tightening, with the bottom response occurring after about one year. This finding corresponds with the actual targeting horizon of the Czech National Bank.[14] In addition, we document that the reaction of tradable prices is faster than that of non-tradable prices. While the maximum effect of a monetary shock on tradables can be seen after a year or so, it is at least a year and a half for non-tradable prices. This result broadly confirms the microeconomic evidence on the effect of competition on price rigidity (Alvarez and Hernando, 2006; Coricelli and Horvath, 2006). We avoid a price puzzle within the system. Thus, our results support the notion that the price puzzle is associated with model misspecification rather than with the actual behavior of the economy. This is also supported in other VAR studies on monetary transmission in the Czech Republic, as all studies estimating the effects of monetary policy across different monetary policy regimes (i.e., the fixed exchange rate regime and inflation targeting regime mixed together) exhibit the price puzzle.

Next, there is a rationale for using the real-time output gap estimate instead of current GDP growth, as using the former results in much more precise estimates. The impulse responses of GDP to an interest rate shock are less precisely estimated, and thus our findings point to the importance of real-time data in monetary policy analysis. Finally, our results also indicate a persistent appreciation of the domestic currency after a monetary tightening ("delayed overshooting", Eichenbaum and Evans, 1995), although the confidence intervals are in this case rather wide, with a gradual depreciation afterwards.

[14] However, note that the targeting horizon (i.e., the horizon minimizing the loss function of the monetary authority) and the horizon at which the monetary policy impact is the most profound are not identical concepts. See Strasky (2005) for details.

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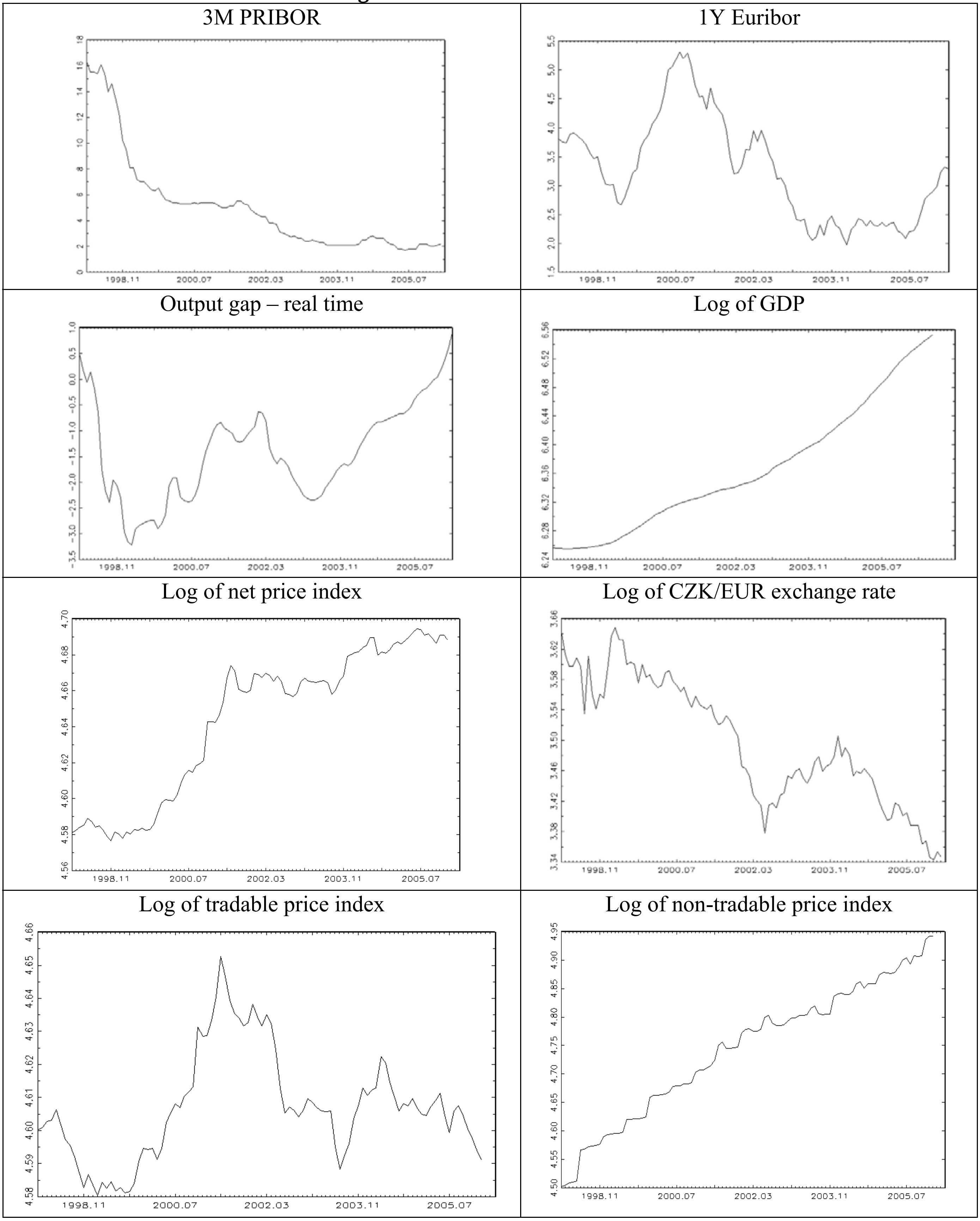
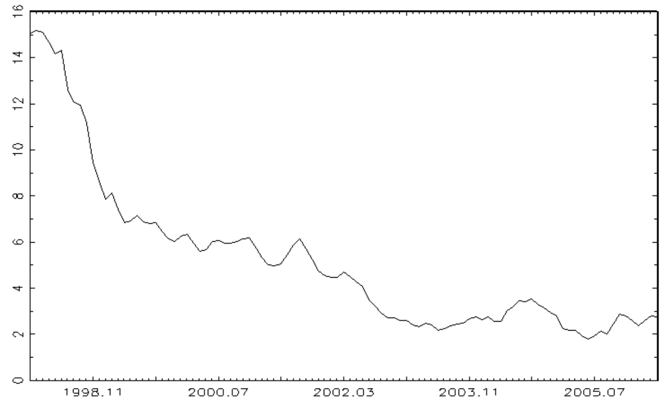
Appendix 1

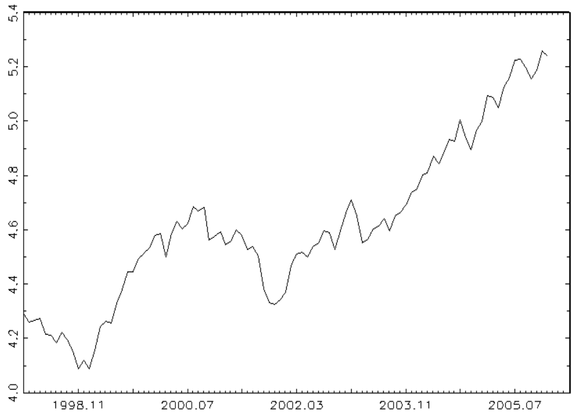
Figure 3: Time series



Forward rate agreements



Log of commodity price index



Appendix 2 – Additional Results: Impulse Responses to a Monetary Shock

Figure 4: GDP instead of output gap

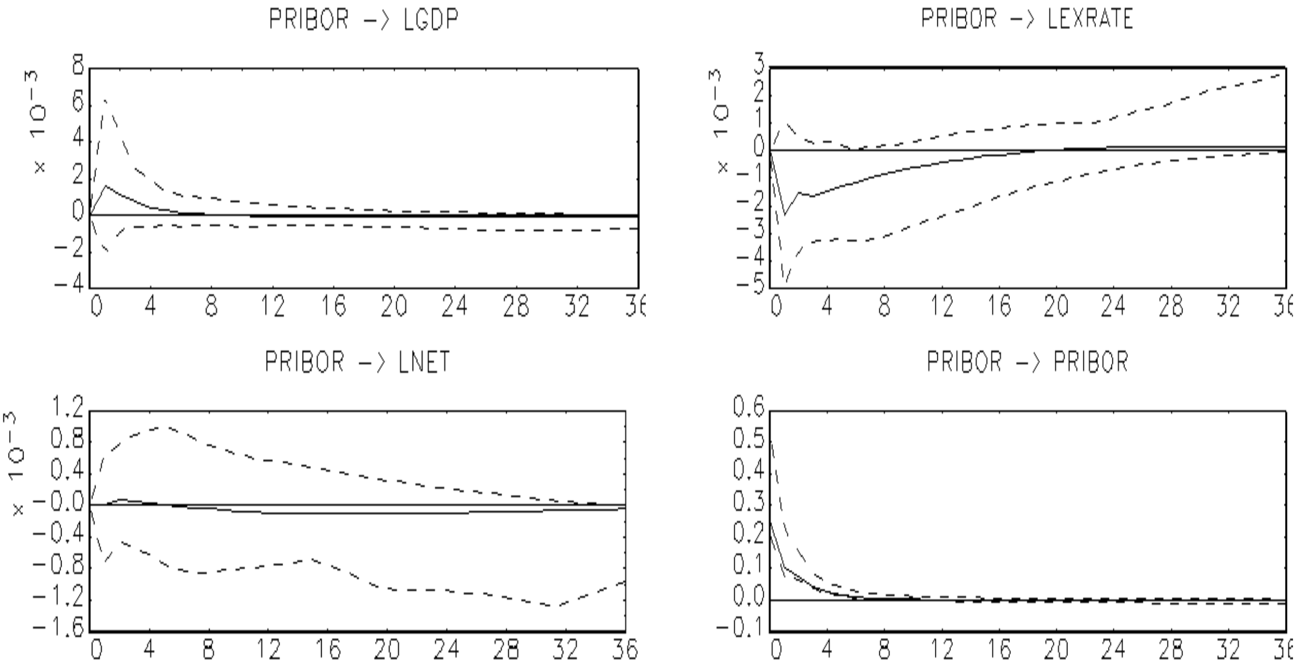


Figure 5: GDP instead of output gap, sectoral prices

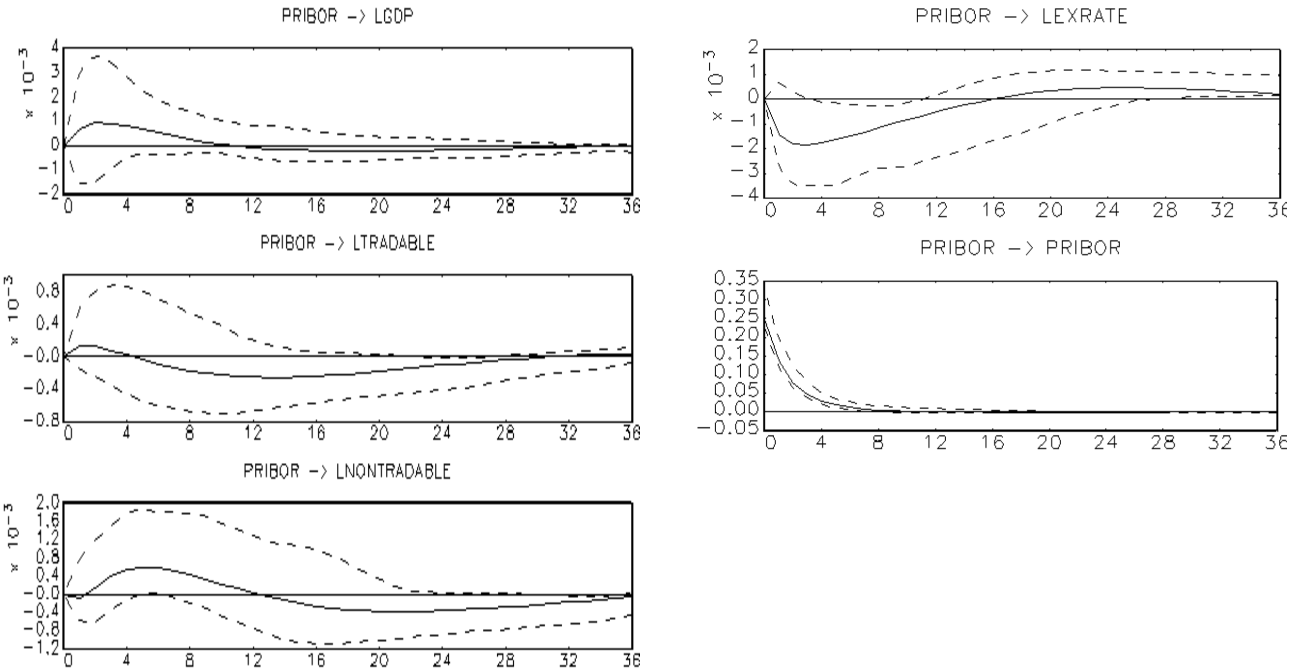


Figure 6: No exogenous variables

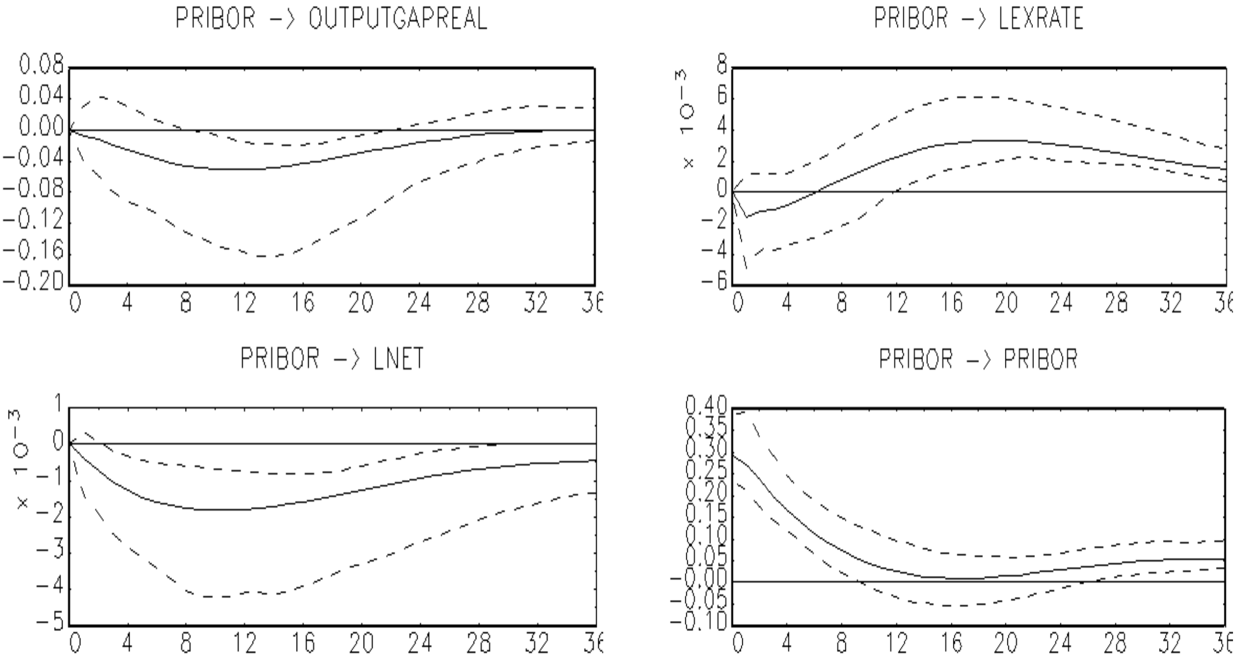


Figure 7: No exogenous variables, sectoral prices

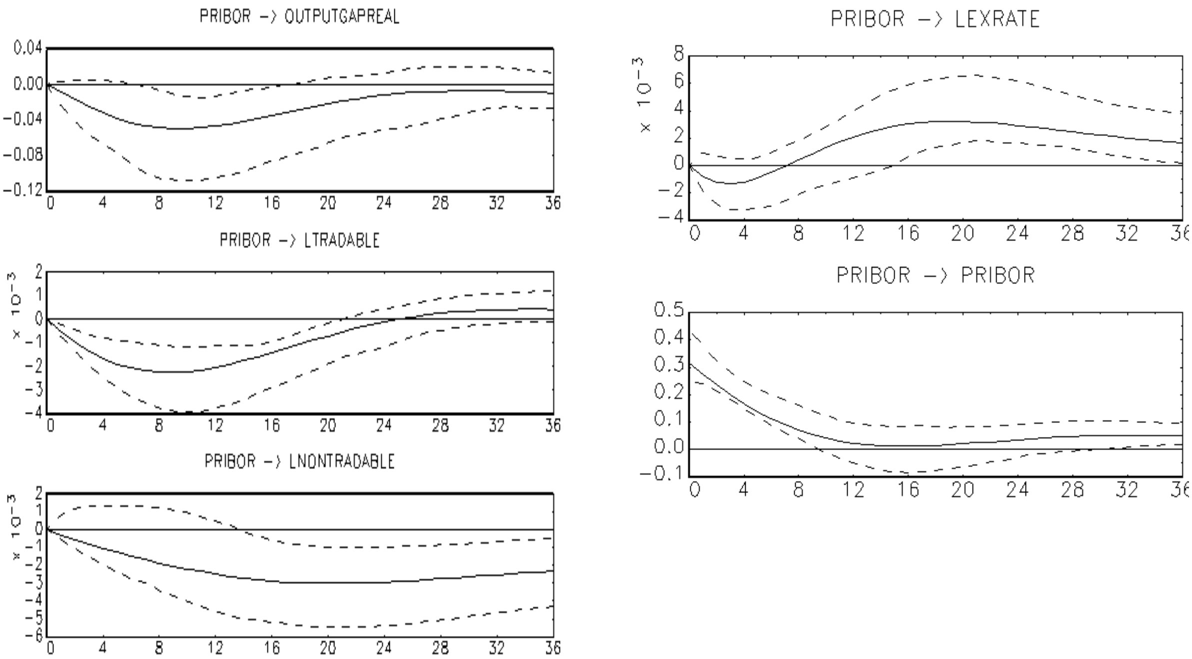


Figure 8: SVAR

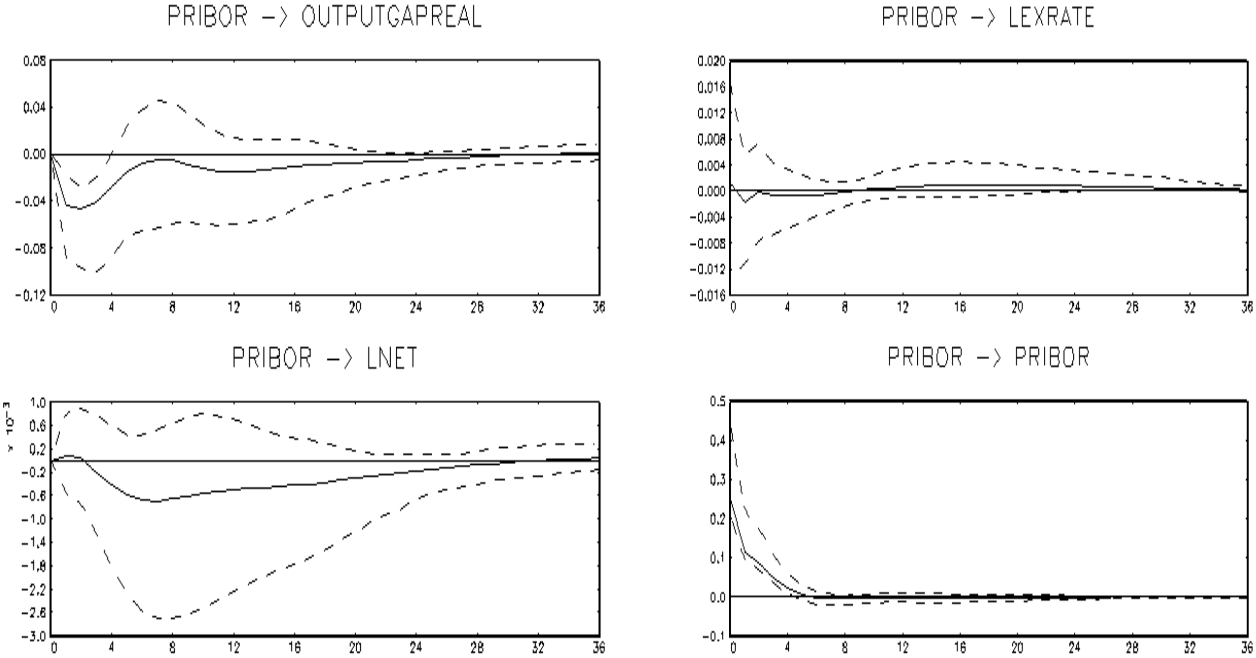
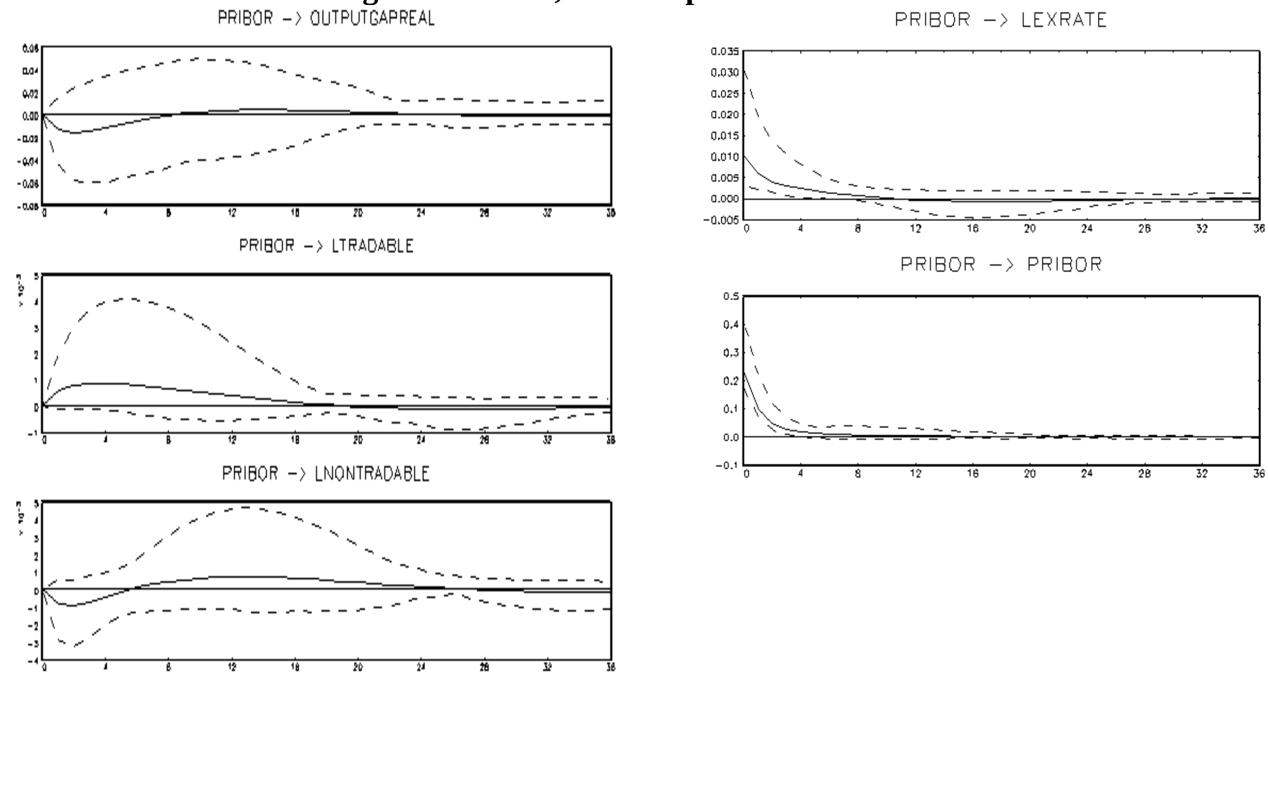


Figure 9: SVAR, sectoral prices



Appendix 3 – Factor-Augmented VAR

In this Appendix, we briefly document our attempt to study monetary policy effects within factor-augmented VAR (FAVAR). However, as documented below, we find that the results based on FAVAR are very sensitive and the confidence intervals for the impulse responses are rather large.

We follow an approach developed by Bernanke *et al.* (2005).[15] FAVAR can be represented in the following form:

$$\begin{bmatrix} F_t \\ Y_t \end{bmatrix} = \Phi(L) \begin{bmatrix} F_{t-p} \\ Y_{t-p} \end{bmatrix} + v_t \ .$$

Vector Y_t contains observable economic variables, whereas F_t represents unobserved factors that provide additional economic information not fully captured by Y_t. We estimate these unobservable factors using a principal component approach, which exploits the assumption that information about the unobservable economic factors can be inferred from a large number of economic time series X_t. Specifically, we can think of the unobservable factors in terms of concepts such as “economic activity” or “investment climate.” They can be represented not by a single economic variable, but rather by several time series of economic indicators.

The FAVAR methodology allows us not only to use a richer information set in the model specification, but also to analyze the effects of a monetary policy shock on a greater number of economic variables. There are two main approaches to estimating FAVAR: a two-step principal components approach and a one-step approach that estimates (3) and a dynamic factor model jointly. As Bernanke *et al.* (2005) do not find any particular differences between these two estimators in terms of inference, we opt for the computationally simpler two-step approach.[16]

In our FAVAR specification, X_t consists of a balanced panel of 40 series that have been transformed in order to ensure their stationarity. The description of these series and their transformations is included in Table 2. The data is at a monthly frequency and spans the period from February 1998 to May 2006. Following Bernanke *et al.* (2005), we assume that the monetary policy instrument (the 3-month interest rate) is the only observable factor, hence

[15] We followed the algorithm developed by Bernanke *et al.* (2005) to estimate FAVAR, which is available on the personal website of Jean Boivin: http://www2.gsb.columbia.edu/faculty/JBoivin/Personal/ .
[16] See Bernanke *et al.* (2005) for a more detailed discussion of principal component analysis and FAVAR.

it is the only variable included in Y_t. For identification purposes the monetary policy instrument is ordered last, which implies that latent factors do not respond contemporaneously (within a month) to innovations in monetary policy. As in Bernanke *et al*. (2005), we distinguish between "slow-moving" and "fast-moving" variables. A "slow-moving" variable is assumed not to react contemporaneously to shocks, while the "fast-moving" variables react instantaneously to changes in monetary policy or economic conditions. The classification of the variables into these two categories is included in Table 2.

In the first step of the two-step estimation, we can distinguish three stages. First, we use principal component analysis to estimate the common factors C_t from all the variables in X_t. Second, after dividing the series in X_t into slow- and fast-moving ones, we estimate the "slow-moving" factors $\widehat{F}_t^s$ as the principal components of the "slow-moving" variables. Finally, we estimate the following regression:

$$\widehat{C}_t = b_{Fs}\widehat{F}_t^s + b_Y Y_{t+e_t} .$$

Based on these estimates, $\widehat{F}_t$ is constructed as $\widehat{C}_t - \widehat{b}_Y Y_t$. In the second step, we estimate the VAR in $\widehat{F}_t$ and Y_t, using a recursive assumption.

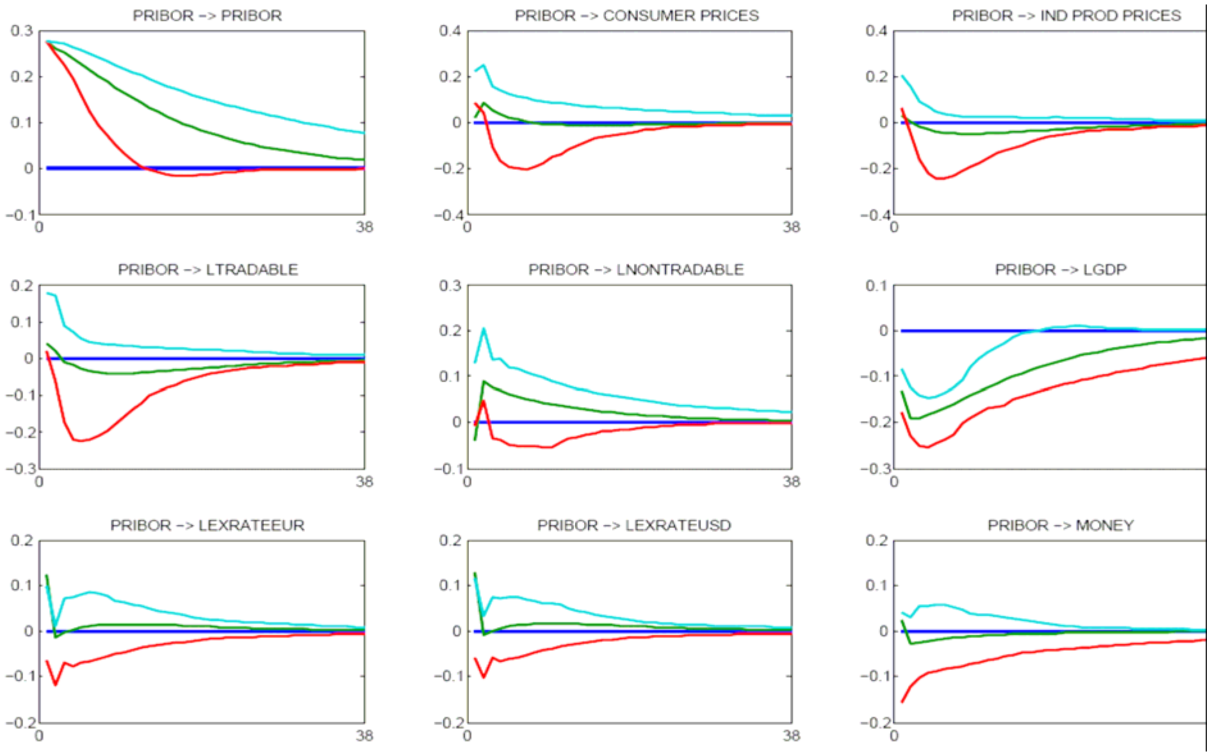
One caveat in our analysis is the fact that we have only 40 series available for principal component analysis, as compared to the 120 used in Bernanke *et al*. (2005). While this may be viewed as a weakness at first sight, it has been argued by Boivin and Ng (2006) that more series do not necessarily ensure better data quality, due to cross-correlation of idiosyncratic errors. In one of their tests, they show that the factors extracted from 40 pre-screened series may in some cases yield better results compared to using 147 series. Therefore, for the sake of size, 40 series, at least in general, should not pose a problem.

Our main results are given in Figure 9. Each panel shows the impulse responses of selected macroeconomic variables to a monetary policy shock with 90% confidence intervals. The FAVAR model in Figure 9 includes three principal factors, but the results were no different when the number of factors was changed. In the benchmark specification we use one lag. The results are highly sensitive to the numbers of lags used, with more lags resulting in highly improbable results.

As a result, the FAVAR model does not appear to properly capture the developments in the Czech economy. Most importantly, the confidence intervals are too large to infer the direction of the impact of the monetary policy change on the macroeconomic variables. The exception is actual GDP growth, $lgdp_t$, which declines after a monetary policy shock, as predicted by the theory.

There may be several reasons for the lack of significant results in the FAVAR estimation for the Czech economy. One reason is likely to do with the relatively short span of the available data; another may be data quality, as discussed by Boivin and Ng (2006). As it is at a monthly frequency, our dataset lacks variables related to consumption, housing starts and sales as well as real inventories and therefore some important economic information may be missing.

Figure 9: FAVAR results



Note: Impulse responses with 90% confidence intervals are presented.

Table 2 – Data description

VAR	DESCRIPTION	TRANSFORMATION	SOURCE
Real output and income			
var1*	Industrial Production, Index number (sa)	3	IFS
var10*	Construction output, constant prices - % (sa)	1	ARAD
var11*	Contracted construction work in enterprises with 20 employees or more - constant prices - (%) (sa)	1	ARAD
var20*	Output gap real - interpolated from quarterly values	2	ARAD
var21*	GDP - interpolated from quarterly values (sa)	3	ARAD
var30	Total agricultural goods output (sa)	3	ARAD
Employment and Hours			
var2*	Industrial Employment (sa)	3	IFS
var3*	Unemployment Rate (sa)	1	Eurostat
var12*	Registered job applicants, total (thousand persons, sa)	1	ARAD
var13*	Vacancies (thousand, sa)	3	ARAD
var14*	Newly registered job applicants (thousand persons, sa)	3	ARAD
var15*	Registered job applicants on unemployment benefit (thousand persons, sa)	3	ARAD
Industry Sales			
var6*	Total sales revenues, Index sales in industry-constant price (corresponding period of preceding year=100, sa)	3	ARAD
var7*	Mining and quarrying, Index sales in industry-constant price (corresponding period of preceding year=100, sa)	1	ARAD
var8*	Manufacturing, Index sales in industry-constant price (corresponding period of preceding year=100, sa)	1	ARAD
var9*	Electricity, gas and water supply, Index sales in industry-constant price (corresponding period of preceding year=100, sa)	1	ARAD
Exchange Rates			
var22	Foreign Exchange Rate (Czech Krown per Euro)	3	IFS
var23	Foreign Exchange Rate (Czech Krown per U.S. $)	3	IFS
Interest Rates			
var26	Treasury Bill Rate	3	IFS
var27	Deposit Rate	3	IFS
var28	Lending Rate	3	IFS
var29	Government Bond Yield	3	IFS
var31	1 day Interbank Rate PRIBOR (%)	1	ARAD
var32	7 day Interbank Rate PRIBOR(%)	1	ARAD
var33	14 day Interbank Rate PRIBOR(%)	1	ARAD
var34	1 month Interbank Rate PRIBOR(%)	2	ARAD
var35	2 month Interbank Rate PRIBOR (%)	2	ARAD
var36	6 month Interbank Rate PRIBOR(%)	2	ARAD
var37	9 month Interbank Rate PRIBOR (%)	2	ARAD
var38	1 year Interbank Rate PRIBOR(%)	2	ARAD
var41	3 month Interbank Rate PRIBOR(%); monetary policy instrument	1	ARAD
Money Aggregates			

var24	Money (sa)	3	IFS
var25	Money plus Quasi Money (sa)	3	IFS
Price Indexes			
var16*	Consumer Prices CPI (sa)	3	ARAD
var17*	Industrial Produces Prices (sa)	3	ARAD
var18*	Tradable prices (sa)	3	ARAD
var19*	Nontradable prices (sa)	3	ARAD
var40	Prague Stock Exchange Index PX50, Historical close, average of observations through period	3	IFS
Exports and Imports			
var4*	Exports	3	IFS
var5*	Imports, FOB	3	IFS

All series were tested for a unit root and when necessary were transformed to achieve stationarity. The transformation codes are: 1-no transformation, 2-first difference, and 3-first difference of logarithm.
An asterisk (*) next to the mnemonic indicates a variable assumed to be "slow-moving" in the estimation.

Testing Multi-Factor Asset Pricing Models in the Visegrad Countries*

Magdalena Morgese Borys

Abstract

There is no consensus in the literature as to which model should be used to estimate stock returns and the cost of capital in emerging markets. The Capital Asset Pricing Model (CAPM) that is most often used for this purpose in the developed markets has a poor empirical record and is likely not to hold in the less developed and less liquid emerging markets. Various factor models have been proposed to overcome the shortcomings of the CAPM. This paper examines both the CAPM and the macroeconomic factor models in terms of their ability to explain the average stock returns using the data from the Visegrad countries. We find, as expected, that the CAPM is not able to do this task. However, factor models, including factors such as: excess market return, industrial production, inflation, money, exchange rate, exports, commodity index, and term structure, can in fact explain part of the variance in the Visegrad countries' stock returns.

JEL Classification: G10, G 11, G12, G15, G31
Keywords: CAPM, macroeconomic factor models, asset pricing, cost of capital, Poland

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1 Introduction

Emerging markets have been quite extensively studied due to the large interest of investors who view them as an attractive alternative to investing in more developed markets. Emerging markets are typically characterized by relatively high returns, but also higher volatility of stock returns as compared to the developed markets. However, there is no consensus in the literature as to which model should be used to explain the returns in these markets and estimate the cost of equity capital. The aim of this paper is to propose such a model for the stock markets in the Visegrad countries: the Czech Republic, Hungary, Poland, and the Slovak Republic. More specifically, we will analyze how different models perform in explaining the variations in stock returns on the stock markets and which of these models should be used to estimate the cost of equity capital in these markets.

The cost of equity capital is crucial information that is needed in order to assess the investment opportunities and the performance of managed portfolios. The cost of equity capital is used as a discount factor when calculating the net present value (NPV) of investment projects.[1] In developed markets, the capital asset pricing model (CAPM) is commonly used by financial managers to calculate the cost of the equity capital,[2] as well as to assess the performance of managed portfolios, such as mutual funds (Fama and French 2004).[3]

[1] In principle, by using the net present value, investors want to verify whether the payoff of the investment exceeds its cost. The future payoffs expected from a particular investment need to be discounted, so that they can be compared to the costs of the investment that need to be incurred at the present time. A good discussion of the NPV methodology can be found in Brealey and Myers (1988). In short, a simple NPV formula is as follows: $NPV = C_0 + \frac{C_1}{1+r} + \frac{C_2}{(1+r)^2} + \ldots,$ where C_0 is the cash flow today (i.e. the cost of investment, a negative number), C_1 is the payoff from the investment one-period ahead, and r is the rate of return that investors demand for the delayed payment. This is the cost of capital.

[2] The rationale behind using the cost of equity capital estimated by the CAPM is the following: since the future payoffs from the investment are risky, i.e. not certain, the rate of return used to calculate the NPV of this investment should come from a comparably risky alternative investment opportunity. A good candidate for such an alternative is investment in the stock market. Therefore, the expected rate of return on the investment in the local stock market, as predicted by the CAPM, can be used as a discount rate in calculating the NPV of the investment project to be undertaken in a given market. If the estimate of the firm's beta coming from the CAPM is biased upward it may lead to a rejection of profitable investment projects, i.e. when the internal rate of return is not greater than the upward biased hurdle rate.

[3] Graham and Harvey (2001) report that 75.5 percent of the 392 respondents to their survey use the CAPM to estimate the cost of equity capital, which is then used to calculate the net present value (NPV) of investment projects, where the cost of equity capital is used as the discount rate.

The CAPM formulated first by Sharpe (1964), Lintner (1965), and Black (1972) describes the relationship between risk and expected return and is used to price risky securities. A very clear and intuitive link between an asset's risk in relation to the risk of the overall market and an asset's expected return is one of the main advantages of the CAPM and is key to understanding its widespread use. However, it is well documented in the literature that CAPM fails to explain a cross section of average stock returns.[4] There are two possible reasons proposed in the literature for the failure of the CAPM to explain the average stock returns. First, there may be a number of priced risks that a single-factor model is not able to properly account for. Second, time variation in either risk or the price of risk may cause the unconditional models to fail.

Following the first reasoning, a number of researchers explored alternative risk factors and proposed various multifactor models. Fama and French (1993, 1996) propose a three-factor model, which includes in addition to market return two factors related to the firm's size (SMB) and the firm's book value (HML).[5] They believe that their SMB and HML factors proxy for unobserved common risk in portfolios.[6]

In addition, factors related to some macro variables have also proven to be able to explain the variation in stock returns. Chen, Roll and Ross (1986) test whether additional sources of risk such as innovations in macroeconomic variables are priced in the stock market. They find that the spread between long- and short-term interest rates, expected and unexpected inflation, industrial production, and the spread between high- and low-grade bonds are significantly priced and able to explain the variations in the stock market.

Breeden (1979) developed the Consumption CAPM (CCAPM) and argued that with complete markets only the aggregate consumption risk should be priced.

[4] While Black, Jensen and Scholes (1972) and Fama and Macbeth (1973) find that the CAPM holds for the 1926–1968 period, more recent studies of the 1960's to the 2000's find otherwise. Among the first studies to report the disappearance of the simple relation between an asset's risk and the average return, as predicted by the CAPM, were Reinganum (1981) and Lakonishok and Shapiro (1986).

[5] In order to obtain these factors the stocks need to be grouped into portfolios on the basis of firm size as well as firm book-to-market value. Their three-factor model then consists of excess market return, the return on small stocks minus the return on big stocks (SMB), and the return on stocks with a high book to market ratio minus the return on stocks with a low book to market ratio (HML).

[6] Other authors have shown the HML and SMB factors may be viewed as proxies for various macroeconomic variables. Liew and Vassalou (2000) and Vassalou (2003) argue that these two factors contain news related to future GDP growth. Petkova (2006), on the other hand, shows that the Fama-French factors may proxy for more fundamental macroeconomic risks as they are correlated with innovations in term spread and default spread.

This insight has been followed by several studies that have experimented with various versions of consumption-based models. While the CCAPM is even less successful than CAPM in explaining the cross section of average asset returns over a short horizon (Campbell 1996, Cochrane 1996), it gains explanatory power over longer time horizons (Parker and Juliard 2005, Jagannathan and Wang 2005). On the other hand, several models have been proposed that include various consumption categories rather than an aggregate consumption variable (Piazzesi, Schneider and Tuzel 2007 and Yogo 2006).

As argued, the empirical failure of the CAPM may also be attributed to the static nature of this model and hence its inability to capture time varying risk premia or correlation structures. There are several models put forth in the literature that use conditioning variables in order to improve the cross sectional power of asset pricing models. Dividend yields and term spreads have been successfully used in models by Jagannathan and Wang (1996) and Fama and French (1988). Lettau and Ludvigson (2001) argue that using log consumption to aggregate wealth ratio as a conditioning variable both in the CAPM and CCAPM models improves their power in explaining the cross section of average stock returns.

Wang (2005) provides an extensive overview of both strains of the literature, including the multi-factor models and models with conditioning variables. In his paper, Wang runs a horse race among models composed of various combinations of eight factors and eight conditioning variables proposed in the literature. He finds that conditional CCAPM conditioning on lagged business income growth has the smallest pricing error in all tests.

Lewellen and Nagel (2006) and Lewellen, Nagel and Shanken (2006) provide a critique of the standard asset pricing tests of the conditional CAPM. They point to a number of problems with these tests, including relying exclusively on book-to-market and size-sorted portfolios, which are known to have a strong factor structure, ignoring theoretical restrictions in the cross-sectional slopes, and additional sampling issues. They offer a handful of proposals aimed at improving the rigor of these asset-pricing tests and contrast the results obtained with their ‘modified’ tests with the results found in influential papers in this strain of literature. None of the five influential models proposed in the literature, including the Fama-French model, perform well according to a more strict set of empirical tests

As discussed above, the classical CAPM model does not always hold in practice when used to analyze the markets in developed countries. The markets in emerging markets, including the stock markets in the Visegrad countries, are less efficient and less liquid compared to developed markets and so it is likely that the CAPM model, especially in its classical formulation, may not be suitable for estimating the cost of capital for these economies. There have been very few studies analyzing these issues in emerging markets. Harvey (1995) argues that emerging markets are characterized by low betas and the CAPM model is not able to capture the relationship between the stock returns in these countries and the market portfolio. Based on this finding there are several studies that analyze various factors that influence the stock returns in the emerging markets and propose models suitable for estimating the cost of capital in these markets. Erb, Harvey, and Viskanta (1995, 1996) find that country credit ratings are significantly related to stock returns, and they propose a model based on these indices. Similarly, Harvey (2004) argues that the country risk rating from the International Country Risk Guide impacts the expected returns in emerging markets and so he incorporates these indices in his version of the CAPM model.

The issue of the relative integration of emerging markets with global markets and the implications on the stock returns in these markets has been central in the literature. Bekaert and Harvey (1995) argue that the integration of the emerging markets with the global markets has been a dynamic process and therefore also the cost of capital should be allowed to vary over time as the relative measure of the integration with global markets changes. In a more recent paper, Bekaert and Harvey (2000) develop a model in which dividend yields are used as a measure of the equity cost of capital. They find that the cost of capital declines as the emerging markets become more integrated with global markets. In one of the few papers that study the markets in Central and Eastern Europe (CEE), Sokalska (2001) finds that the stock prices of the Czech Republic, Hungary and Poland move together. She argues that local macroeconomic fundamentals are of relatively small importance in those markets and that the key factors influencing the movements of stock prices are exogenous. Namely, she claims that it is the flow of foreign portfolio capital that can be traced to affect the movement of stock prices in those markets. De Jong and de Roon (2001) link the issue of time-varying market integration with expected returns in emerging markets.[7] They find, in line with the theory, that increasing market integration (or decreasing market segmentation) leads to lower expected returns and hence lower cost of capital.

There are some important data and methodological issues that need to be addressed in the Visegrad countries. First, the data available is of a relatively short time span, which may influence the plausibility of our results. Second, there is a limited number of stocks traded on these stock exchanges,[8] which makes some of the commonly used portfolio techniques difficult to apply. Taking these considerations into account, we used the Fama and MacBeth (1973) procedure (FMB) to estimate our models. This procedure is extensively used by researchers to estimate and test the single- and multi-factor models' predictions. It gives unbiased estimates even when there is correlation between observations of different firms in the same year. It also accounts for variation coming from both time-series and cross-section regressions, which is especially important when there is a limited number of observations, as is the case with the Visegrad countries' data.

First, we estimated the CAPM by the FMB procedure to see how this model performs in the stock markets of Visegrad countries. As expected, the market factor alone was not able to explain the average stock returns. Given these results we proceeded with the estimation of various macroeconomic factor models. It turned out that different macroeconomic factors are priced in each of the Visegrad countries. Multi-factor models perform much better than the simple CAPM in explaining the average stock returns in these countries. In addition, we also employed models with principal components as factors.

This paper is organized as follows. In Section 2 we discuss the CAPM and factor models in greater detail, as well as the testing procedures. In Section 3 we introduce the data and discuss some of its limitations, which make the use of some standard techniques impossible. Section 4 contains the empirical results from testing the CAPM and factor models in the Visegrad countries. In Section 5, we briefly summarize the findings of this paper and suggest some directions for further research.

[7] They develop a model in which expected returns depend on the degree of market segmentation, measured as a ratio of assets in a given market that cannot be traded by foreign investors. Given that the degree of segmentation changes over time, they allow the expected returns to also vary with time. Using data from 30 emerging markets, including the Visegrad countries, de Jong and de Roon provide evidence that the market segmentation has a significant effect on the expected returns.

[8] The variability in the number of stocks traded in the sample is given in Table I and Table II.

2 Methodology

The main objective of this paper is to point to a suitable asset-pricing model that could be used to estimate the cost of equity capital in the Visegrad countries. The first candidate is the Capital Asset Pricing Model (CAPM). According to this classical model specification, the expected return on the security or on the portfolio of securities should be equal to the risk-free rate plus a risk premium, which consists of the portfolio's beta multiplied by the expected excess return of the market portfolio (return on the market portfolio minus the risk-free rate). A classical, one-factor CAPM looks as follows:

$$E(r_{it})\text{-}r_t^{\,f} = \beta_i\left(E(r_t^{\,m}) - r_t^{\,f}\right), \qquad (1)$$

where E(r_{it}) is the expected *i-th* stock return (*i=1….N*), $r_t^{\,f}$ is the risk-free rate, and $E(r_t^{\,m})$ is the expected market return. This model can be empirically tested using the following regression equation:

$$r_{it}\text{-}r_t^{\,f} = \alpha_i + \beta_i^m\left(r_t^{\,m} - r_t^{\,f}\right) + \varepsilon_{it}\,, \qquad (2)$$

where r_{it} - $r_t^{\,f}$ is the excess market return on the *i-th* stock, r_{mt} is the market return, α_i is the constant term, β_i^m is the coefficient on the excess market return for each of the *i* stocks, and ε_{it} is the error term. According to the CAPM prediction, the constant term, α_i, should be statistically insignificant (i.e. equal to zero) for each of the *i* stocks. If this is the case, the pricing errors are zero and the CAPM is said to hold empirically. In addition, the slope coefficient, β_i^m, should be significantly different from zero, indicating that the excess market return is indeed priced by the stock market, i.e. it helps to explain the variation in the stock returns. Moreover this coefficient, which represents the measure of one stock's risk as compared to the risk of the overall market, should vary among the stocks.

We also considered an extension of the classical CAPM—a multi-factor model. Suppose there are *k-factors* that are believed to influence the stock returns in a given market. The *k-factor* model can be tested by using the following regression equation:

$$r_{it}\text{-}r_t^{\,f} = \alpha_i + \beta_i^m\left(r_t^{\,m} - r_t^{\,f}\right) + \beta_i^2\left(r_t^{\,2}\right) + \ldots\ldots + \beta_i^K\left(r_t^{\,K}\right) + \varepsilon_{it}\,, \qquad (3)$$

where $r_t^{\,m}$ is a local index return, $r_t^{\,k}$ is the *k-th* factor return (*k=2…K*), r_{it} is the *i-th* stock return, $r_t^{\,f}$ is the risk-free rate, α_i is the constant term, and ε_{it} is the error term. Similarly to the CAPM, this multi-factor model predicts that the constant terms, α_i, should be insignificant and the slope coefficients, β_i^m and β_i^k, should be significantly different from zero. As discussed, these factors may include Fama-French factors (FF factors) or other factors, including macroeconomic variables or the principal factors obtained by the principal component analysis.

Principal component analysis allows capturing factors that most accurately proxy the driving forces behind the economic activity in a given country. This methodology exploits the assumption that the information about the key economic factors can be inferred from a large number of economic time series. Specifically, we can think of the key factors in terms of concepts such as “economic activity” or “investment climate”. They cannot be represented by a single economic variable but rather by several time series of economic indicators. Principal component analysis may be particularly suitable for analyzing transition economies, such as Visegrad countries. In particular, given the relatively short span of data available and the need to account for structural breaks, which may limit the span of data under study even further, it is key to use a methodology that uses the information available at a given time to a maximum. In our analysis we follow the methodology developed by Stock and Watson (1998). The model is defined in the following way:

$$X_t = \Lambda_t F_t + e_t\,, \qquad (4)$$

where X_t represents a N-dimensional vector of time series, F_t is the *rx1* common factor, Λ_t is the factor loading, and e_t is a *Nx1* idiosyncratic error term. The factors summarizing multiple time series are unobservable. They can be estimated by a quasi-Maximum Likelihood Estimation (MLE), which involves the two following assumptions: i) $\Lambda_t = \Lambda_0$ and ii) e_{it} are i.i.d. and independent across series. These estimated principal factors are then used as factors augmenting the CAPM.

In the literature one can find several ways of testing capital asset pricing models. They can be divided into the following three categories: tests involving time-series regressions, tests involving cross-section regression, and tests involving a combination of the above.[9] One of the most widely used methods is the Fama-MacBeth procedure (FMB), which combines the time series and the cross section regressions. Suppose we have N firms returns for any given month t, R_t. In the first stage we regress the excess stock return on the excess market return and other *k-factors* in order to obtain the CAPM cross-section betas, $\hat{\beta}_i^m$ and $\hat{\beta}_i^k$, where i is a firm's subscript ($i = 1 \ldots N$), m stands for the market return, and k is a factor's subscript ($k = 2 \ldots K$).[10] In the second stage we run the following cross-section Ordinary Least Squares (OLS) regression for any single month t:

$$R_i = \gamma_0 + \gamma^m \hat{\beta}_i^m + \gamma^k \hat{\beta}_i^k + \eta_i \quad , \qquad (5)$$

where $R_i = (R_1, R_2, \ldots, R_N)$ is a $Nx1$ vector of cross-section excess monthly stock returns,

$\hat{\beta}_i^m$ and $\hat{\beta}_i^k$ are NxK matrices of CAPM betas (obtained in the first stage regressions),

γ^m and γ^k are vectors of cross-section coefficients for each of the *k- factors*,

γ_0 is a scalar and an estimate of intercept, and

η_i is a $Nx1$ vector of cross-section error terms.

Then we repeat this regression as in (4) for each month $t = 1, 2, \ldots T$ and obtain T estimates of γ_0, γ^m and γ^k. Finally, the time-series estimates of these parameters are tested to see if: $E(\gamma_0) = 0$ (i.e. pricing errors are zero), $E(\gamma^m) > 0$ (i.e. positive risk premium on the excess market return), and $E(\gamma^k) > 0$ (i.e. positive risk premium on

[9] A detailed discussion of these various methods, including the Fama-MacBeth procedure, can be found in Cuthbertson and Nitzsche (2001).

[10] The first stage regressions are based on 2-year window regressions, assuming that betas are relatively stable over that time period. Empirical tests developed in the literature often assume this window to be even longer, e.g. 5 years. See Lewellen and Nagel (2005) for a more detailed discussion of high frequency changes in betas.

betas for each of the k factors). Assuming the returns are i.i.d. and normally distributed, the following t-statistic is used:

$$t_\gamma = \frac{\bar{\hat{\gamma}}^k}{\sigma(\hat{\gamma}^k)} \quad \text{, where} \quad \sigma^2(\hat{\gamma}^k) = \frac{1}{T(T-1)} \sum_{t=1}^{T} (\hat{\gamma}^k - \bar{\hat{\gamma}}^k)^2 \ . \qquad (6)$$

Similarly, one can obtain t-statistics t_γ for γ^m and γ_0 and test all of the CAPM restrictions.

There is one important caveat to the FMB approach. Since the estimates of betas, $\hat{\beta}_i^m$ and $\hat{\beta}_i^k$, obtained in the first stage regressions may be measured with error, we may encounter the 'errors-in-variables' problem in the second stage regressions. Specifically, if the estimates of $\hat{\beta}_i^m$ and $\hat{\beta}_i^k$ that we use in the second stage regressions contain measurement error, then the estimates γ^m and γ^k will be biased.[11] The most common approach to minimizing this problem is to group the stocks into portfolios[12] and estimate the portfolio betas instead of the stock betas in the first stage regression. Then, in the second stage, the average excess return $\bar{r}_i - \bar{r}^f$ for each of the stocks is regressed on the appropriate portfolio beta. This approach reduces the measurement error but it does not completely resolve this problem since it still uses betas estimated in the first step in the regressions in the second step. Using portfolios rather than individual stocks can, however, improve the estimates mainly due to utilizing the portfolios' betas rather than the individual stocks' betas, which may contain structural breaks. Due to limited number of companies listed on the Visegrad stock exchanges, there may not be enough observations at any given point of time to form portfolios of individuals stocks. Instead, we proceed with Generalized Method of Moments (GMM) estimation that allows for a simultaneous estimation of betas and gammas, and therefore avoids the errors-in-variable problem altogether. The GMM estimator is defined by minimizing the following criterion function:[13]

$$\sum_t m(y_t, \theta) A(y_t, \theta) m(y_t, \theta) \ , \qquad (7)$$

[11] In the least squares regressions, the errors-in-variables are likely to cause the estimates of the slope coefficients to be biased downward and the estimate of a constant term to be biased upward.

[12] The portfolios can be formed based on the size, beta or book-to-market ratio of individual stocks obtained from running the time-series regressions.

[13] See Davidson and MacKinnon (1993) for a more detailed discussion of GMM estimation.

where $m(y_t, \theta)$ are a set of moment conditions that parameters θ should satisfy and A is a weighting matrix. While any symmetric positive definite matrix A will obtain a consistent estimate of θ, it is possible to show that in order to obtain an asymptotically efficient estimate of the θ matrix A should be equal to the inverse of the covariance matrix of the sample moments. There are various methods of estimating this covariance matrix, including White's heteroskedasticity consistent matrix and the heteroskedasticity and autocorrelation (HAC) consistent matrix. We estimate our asset pricing models by one-step GMM, in which White's optimal weighting matrix is obtained in an iterative process by sequential updating of the coefficients.

Following Cochrane (2005) we write the moment conditions in the following way:

$$gT(b) = \begin{bmatrix} E\left(R_t^{\ e} - \alpha - \beta f_t\right) \\ E\left[\left(R_t^{\ e} - \alpha - \beta f_t\right) f_t\right] \\ E\left(R^e - \beta\gamma\right) \end{bmatrix} = \begin{bmatrix} 0 \\ 0 \\ 0 \end{bmatrix} \quad . \tag{8}$$

These moment conditions are written assuming one asset and one factor but can be easily extended to include N assets and K factors. In such a system there would be *N(1+K+1)* moment conditions since for each asset *N* we would have one moment condition for the constant, *K* moment conditions for *K* factors and one moment condition that allows estimation of the gammas (asset-pricing model condition). On the other hand, there would be *N(1+K)+K* parameters and hence *n-K* overidentifying restrictions. They can be tested with a chi-square test (J-test for overidentifying restrictions).

To summarize, in this paper we estimated several alternative models, including the classical CAPM, macroeconomic factor models and principal factor model using two alternative estimation methods: the FMB and the GMM. One of the main advantages of the FMB estimation procedure is that it uses all the information available for a given data point, accounting for variation coming from both sources: time-series and cross-section. Given relatively short time spans of data available for the stock markets in the Visegrad countries, it is key to be able to utilize all the available data points to their maximum. This procedure is, however, prone to the errors-in-variables problem due to a two-stage estimation. In order to account for this problem, we obtained alternative estimates for Poland using the GMM one-step procedure, which allowed us to assess the potential importance of the errors-in-variables in our models. Unfortunately, we were not able to obtain similar estimates for other Visegrad countries due to the variance-covariance matrices not being positive definite, and hence not invertible.

3 Data

Data on individual stocks as well as the local market indices needed to test the validity of the classical CAPM in Visegrad countries were obtained from Wharton Research Data Services. Other data was obtained from IMF's International Financial Statistics Database, national banks' and ministries of finance's websites. A summary of these variables is presented in Table 1.

As argued, the classical, one-factor CAPM does not always hold empirically and therefore various multi-factor models have been proposed in the literature. FF factors are the most commonly used in the literature as they turned out to be the most successful empirically. In order to obtain these factors the stocks need to be grouped into portfolios on the basis of the firm's size as well as the firm's book-to-market value. Due to a limited number of stocks traded in the stock markets of Visegrad countries (see Chart 1), the portfolio grouping may not be optimal.

Therefore, in this paper a second best approach is used, namely the macroeconomic factor model. It has been noted that observable economic time series like inflation and interest rates can be used as measures of pervasive and common factors in stock returns. Chen, Roll and Ross (1986) argue that stock prices can be expressed as expected discounted dividends:

$$p = \frac{E(c)}{k}, \tag{9}$$

where c is the dividend stream and k is the discount factor. From this it can be deducted that the economic variables that influence discount factors as well as expected cash flows will also influence the expected returns. Chen, Roll and Ross (CRR) use the following factors: industrial production growth, a measure of unexpected inflation, changes in expected inflation, the difference in returns on low grade corporate bonds and long-term government bonds (risk premia), the difference in returns on long-term government bonds and the short-term Treasury bills (term structure), changes in real consumption, and oil prices. In our factor model, similarly to CRR, we included monthly industrial growth and the term structure. In contrast to CRR, we did not include two inflation variables in order to avoid likely correlations between them. Instead, we used only the monthly inflation. Since there is no time-series data on corporate bond grading in Visegrad countries, we did not incorporate any measure of risk premia in our model.[14] To summarize, in our baseline factor model we used the following four factors: market return, monthly growth rate of industrial production, inflation, and term structure. Changes in the level of industrial production affect the real value of cash flows. In addition, a direct link between the returns and production is specified in the business cycle models. Inflation influences the nominal value of cash flows as well as the nominal interest rate. Finally, the discount rate is affected by the changes in the term structure spreads between different maturities. In addition, we estimated alternative factor models, which included variables that we believe may be important in Visegrad countries. The additional variables included: exchange rate, all primary commodity index,[15] German industrial production, money, and exports. Given that all these countries are relatively small, open economies, the fluctuations in exchange rate, commodity prices, exports and money base are likely to have a strong impact on other macroeconomic variables. The economic situation in Germany (proxied by its industrial production), one of the most important trading partners for the Visegrad countries, may have a significant impact on the economies of these countries and therefore may also influence their stock markets. The time series of all these additional variables were obtained from the IMF International Financial Statistics Database. The summary of these variables is presented in Table 2.

In order to overcome the limitations of factor models with respect to the small number of variables that can be used in the estimation, we employed principal component analysis to extract the main factors driving the economies of the Visegrad countries. This method is mainly used for forecasting purposes. It is based on the

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[14] Omitting risk premia in the model specification is likely to result in the omitted variable bias in the coefficient on the term structure (being the variable most highly correlated with the risk premia (according to CRR). This bias is likely to be negative as, according to CRR, the correlation between these two variables is negative and the likely sign of the coefficient on the risk premia is positive. All the other variables (excess market return, inflation, and industrial production) are positively correlated with risk premia and therefore omitting risk premia is likely to create an upward bias in coefficients on these variables.

[15] CRR also consider changes in consumption and oil prices in their model but find that these variables are not significantly related to the stock returns. Due to a lack of available data on consumption we did not include it in the analysis. For the oil prices we decided to use as a proxy the all primary commodities index from the IMF.

principle that there are a few forces driving the dynamics of all macroeconomic series. Since these forces are unobservable they need to be estimated from a large number of economic time series. Given the still-transitionary character of the Visegrad stock markets, as well as the limited span of data available, principal component analysis may be very useful for explaining the stock returns in these countries. The list of variables used to obtain the principal factors is included in the Appendix. The first three principal factors, which explained most of the variance of the average stock returns, were then used as the factors in the alternative multi-factor model (principal factor model). In addition, the same number of factors as the baseline multi-factor model allowed a direct comparison of the performance of these two models. The summary statistics of the three first principal factors used in the principal factor model are presented in Table 3.

4 Estimation

The CAPM (single-factor model) was estimated using the regression equation (2) by the FMB procedure, where local market indices were used as proxies for market portfolio,[16] and monthly returns on local t-bills represented the risk-free rate. The results obtained for the four Visegrad markets are presented in Table 4.

These results indicate that the CAPM should not be rejected for Poland, as the constant term was statistically different from zero. However, also the coefficient gamma (γ^m) was not statistically different from zero, which implies that the market return was not priced. In Hungary and in the Slovak Republic the market return was priced, however, as the constant terms were also significantly different from zero the CAPM should be rejected, due to the presence of pricing errors in the model specification. The results for the Czech Republic indicated that the CAPM should be rejected, as the constant term was significantly different from zero. The rejection of the CAPM model is not surprising and is in line with the literature covering the behavior of stock exchanges in the second half of the twentieth century. Therefore,

[16] Initially we considered using the following three alternative variables as a proxy for the market portfolio: local market index, S&P 500 index and the MSCI world index. We tested these various specifications for the Polish market and found that the choice of market proxy did not influence the test of the validity of the CAPM. Our findings are consistent with Low and Nayak (2005), who show that the choice of market portfolio is irrelevant for the validity of CAPM. Therefore, we proceeded with the local market index as a proxy for the market portfolio.

we extended the single-factor model by adding additional macroeconomic factors. In the baseline factor model, we added the following three variables: industrial production, inflation, and the term structure. This extended four-factor model was also tested following the FMB procedure. The results from these regressions are presented in Table 5.

Some of the factors turned out to be significantly priced in Poland and in the Slovak Republic. In Poland, inflation was able to explain part of the variation in the average stock returns. In the Slovak Republic, two factors—inflation and the industrial production—seemed to have some explanatory power. While none of the factors turned out to be significant in the Czech Republic or Hungary, some of the t-statistics were quite high, bordering on significance at the 10 percent level (for the term structure in the Czech Republic and for the term structure and inflation in Hungary). These lower values may be due to a downward bias, caused by the presence of errors-in-variables resulting from the two-step estimation in FMB. For all four countries, the constant terms were not statistically significant, indicating that no pricing errors were present in this specification.

Given that few factors turned out to be statistically significant, we proceeded with alternative multi-factor models, in which we included additional variables such as exchange rate, German industrial production, money, commodity index, and exports. The results of these alternative estimations are presented in Table 6. For the Czech Republic, a model including money, industrial production and exports in addition to the market factor was the most promising. Not only were the three macroeconomic factors statistically significant, but also the constant term was not statistically significant, so that the model could not be rejected. For Hungary, the constant term was not statistically significant in the two models: the one with inflation and the exchange rate, in which the exchange rate was significantly priced, and the other one with excess market return, exchange rate, money, and commodity index, in which all factors but the excess market return were significantly priced. In case of Poland, two models were correctly specified, the one with inflation and money and also the alternative one with the commodity index as an additional factor (albeit not priced). In the Slovak market, the correct model appeared to be the one with industrial production, term structure and money, in which market factor, term structure and money were all significant and the constant term was not. In addition, a model with inflation, industrial production, term structure, money, and commodity index was also plausible, with inflation and term structure as the significantly priced factors. We have performed a series of tests of the added explanatory power of macroeconomic factor models and the CAPM.[17] For Hungary, Poland, and Slovakia the correct factor model appears to be the baseline four-factor model, which is superior to both the CAPM and the alternative factor models. In the Czech Republic, on the other hand, the baseline four-factor model is strongly rejected in favor of the alternative model with the market factor, industrial production, exports, and money.

In the next stage, we employed principal component analysis to obtain the key factors,[18] which we then incorporated into a factor model together with an excess market return. This four-factor model (including three principal factors/components and an excess market return) was estimated using FMB. The results of this estimation are presented in Table 7.

According to the results presented in Table 7, Poland was the only country for which the principal factor model had to be rejected, as the constant term was statistically significant. In spite of the model rejection, all factors as well as the excess market return were significant. In the other three Visegrad countries, the constant terms were not statistically significant and hence the principal factor models could not be rejected. In the Czech Republic, the second factor was significant, which was mainly driven by developments in imports and exports. In Hungary, the excess market return was significant as well as the second and third factors, which included primarily information on exchange rates, consumer prices and prices of primary commodities. Finally, in Slovakia, none of the factors turned out to be significant.

As argued, the results obtained by using FMB procedure are likely to be biased due to the errors-in-variables problem. In order to verify this hypothesis we proceeded with an alternative GMM estimation, in which all the slope coefficients

[17] These results are available upon request.

[18] Given the relatively small samples and short time series it is difficult to argue with certainty that these unobservable factors are different across the Visegrad countries. Therefore, the following results should be viewed with caution. In the Czech Republic these three principal components accounted for 56 percent of the variance of data used for the principal component analysis. The first factor was driven mainly by interest rates, the second by imports and exports, the third by prices of primary commodities. In Hungary, the three principal factors captured 82 percent of the variance. The first and third factors had similar composition to the Czech Republic, while the second factor was driven mainly by exchange rates (forint against USD and EUR) and consumer prices. In Poland, the three factors amounted to 49 percent of the variance. The first one was primarily summarizing the developments in exchange rates (zloty against USD and EUR) and in producer prices. The second factor captured again the developments in exchange rates as well as in prices of primary commodities. In Slovakia, the three principal factors captured 72 percent of the overall variance in data. The first factor summarized movements in the prices of primary commodities, the second in both consumer and producer prices, while the third contained information mainly on the changes in industrial production.

(betas and gammas) are estimated simultaneously. We obtained satisfactory confirmation of this hypothesis for Poland. For other countries, however, we were not able to obtain the GMM estimates due to data issues. Specifically, it was not possible to obtain the inverses of the variance-covariance matrices of residuals defined in (6) in these systems (these matrices were not positive definite). The estimates obtained for the four-factor model for Poland from the one-step GMM estimation are presented in Table 8.

The results presented in Table 8 support the hypothesis that the FMB estimates of the slope coefficients are likely to be biased downward. The estimates obtained by one-step GMM for the four factors are in all cases greater than the FMB estimates. More importantly, they all turned significant, as compared with only two factors: inflation and term structure being significant in the FMB case.[19]

5 Summary

Emerging markets returns have been quite extensively studied in the last decade. However, it is not clear which model should be used to explain the returns in these markets and to estimate the cost of capital. The cost of capital is important information that is needed to evaluate investment opportunities, as well as to assess the performance of managed portfolios. In developed markets, the CAPM is most often used to estimate the cost of capital, even though its empirical record is quite poor. Factor models have been developed to overcome some of the CAPM's shortcomings, namely the inability of the excess market return alone to explain the variance of the average stock returns. Factor models extend the CAPM by adding additional factors to the excess market return in order to improve the predictive power of the model.

In this paper we tested various asset-pricing models and evaluated their relative performance in explaining the average stock returns in the Visegrad countries. These models, as argued, can be potentially used to estimate the cost of capital, which is then used to evaluate investment opportunities. We began by formally estimating

[19] These results are presented for a shorter sample and are meant only for illustrational purposes. Due to the model being weakly identified, we were not able to replicate these results for all the Visegrad countries.

the CAPM by the FMB procedure using data from the Visegrad markets to see how it performs. While we were not able to reject the null hypothesis that the CAPM holds (i.e. constant terms are not significantly different from zero) for Poland, we were also not able to reject the null hypothesis that the coefficients on the factor loading (betas of the excess market return) were statistically insignificant. In contrast, we could reject the CAPM for the three other Visegrad markets. Having confirmed the low power of the CAPM in explaining the variance of the average stock returns we then proceeded to estimate a factor model.

Due to a limited number of stocks traded in the Visegrad markets we decided not to proceed with the FF factors. Another alternative is to use the so-called 'macroeconomic factor models', in which observable economic time series like inflation or interest rates are used as measures of pervasive or common factors in asset returns. We employed a macroeconomic factor model based on the factors used by CRR (1986). In our baseline model we included the following four factors: excess market return, industrial production, inflation, and excess term structure. We estimated this four-factor model using the FMB procedure. This model had some explanatory power in Poland and in Slovakia since some of the slope coefficients were significant, indicating that the factors were priced. Moreover, in all the countries, the coefficients on the constant terms were not significant; hence there were no pricing errors present in this specification. Given that most of the factors were not significantly priced we proceeded with estimating alternative macroeconomic factor models. While more of the factors were priced in these models, with the exception of the Czech Republic they could be rejected in favor of the baseline four-factor models in tests of added explanatory power. For the Czech Republic, an alternative four-factor model with industrial production, money, and exports turned out to be superior to the baseline four-factor model.

Even though these results turned out to be satisfactory, we decided to proceed with principal component analysis in order to extract the key factors that explain the variability of stock returns in these countries. In Poland all the factors as well as the market factor were statistically significant but the model was rejected, as also the constant term was significant. For the other three Visegrad countries, the model could not be rejected and various principal factors turned out to be significantly priced (apart from Slovakia, where none of the factors were significant). In the Czech Republic, the second factor, summarizing the developments in imports and exports, was statistically significant. In Hungary the excess market return was significant as well as the first and the second principal factors, driven mainly by changes in exchange rates, consumer prices, and primary commodity prices. We also performed tests of added explanatory power between the CAPM (restricted model) and the baseline four-factor model (unrestricted model). In all countries we were able to reject the CAPM in favor of the baseline four-factor model.[20] In addition, in the Czech Republic, Hungary, and Slovakia the CAPM was also rejected in favor of the alternative factor models.

Based on these results we concluded that macroeconomic factor models, rather than the capital asset pricing or the principal factor models, are suitable for estimating the cost of capital in the Visegrad countries. Our conclusion is supported by the results obtained for Poland when using the one-step GMM estimation method. These alternative estimates, free of the error-in-variables problem, resulted in all the factors turning significant, confirming that the FMB estimates are likely to be biased downward. Even though due to empirical problems we were not able to obtain similar alternative estimates for other Visegrad countries we can expect that the estimates obtained by the FMB most likely undermine the significance of macroeconomic factors in explaining the average stock returns.

[20] In the Czech Republic, Hungary, and Poland we rejected the CAPM in favor of the four-factor model at the 1 percent significance level, while in Slovakia only at the 10 percent level.

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Appendix: Description of Data Used in the Principal Component Analysis

All data was downloaded from the IMF International Financial Statistics (IFS) Database. The series were downloaded in quarterly frequency and then transformed into monthly frequency by the constant sum method (the quarterly value is divided by three and for each of the three months in this quarter the same value is repeated). The time series were imported into Eviews and automatically converted into logarithms of the original series if the values were positive. Then, unit root tests were performed and each series was differenced a sufficient number of times to achieve stationarity. These series were then used in the principal component analysis.

Series name	Units
Money Aggregates:	
Money, Seasonally Adjusted	Units of National Currency
Money Plus Quasi-Money	Units of National Currency
M1	Units of National Currency
M2	Units of National Currency
M3	Units of National Currency
Interest Rates:	
Discount Rate (End of Period)	Percent Per Annum
Money Market Rate	Percent Per Annum
Treasury Bill Rate	Percent Per Annum
Deposit Rate	Percent Per Annum
Lending Rate	Percent Per Annum
Government Bond Yield	Percent Per Annum
Prices:	
Share Price Index	Index Number
Producer Price Index	Index Number
Consumer Price Index	Index Number
Industrial Production and Employment:	
Industrial Production (sa)	Index Number
Industrial Employment	Index Number
Wages: Average Earnings	Index Number
Unemployment Rate	Percent Per Annum
Imports and Exports:	
Imports	Millions of US$
Exports	Millions of US$

Exchange Rates:	
Official Rate (per 1 USD)	Units of National Currency
Official Rate (per 1 EUR or ECU)	Units of National Currency
Other:	
German Industrial Production (sa)	Index Number
All Primary Commodities Prices	Index Number
Non-Fuel Primary Commodities	Index Number

Table 1: Summary Statistics - CAPM

Sample mean, standard deviation, maximum and minimum values are reported for the variables used in the CAPM regression. These statistics are reported for the cross sectional distribution, where the number of firms varies from 2 to 74 depending on the country. All the variables represent monthly returns in local currency. Stock_rt stands for stock return, market_rt is the local market return, and tbill is the monthly return on short-term government securities.

Variable	**Mean**	**Std. Dev.**	**Min.**	**Max.**
Czech Republic; 4942 obs; Feb 1994 – Dec 2007; No of Companies: 6-74				
stock_rt	-.0060	.1649	-.9250	1.5352
market_rt	-.0052	.0732	-.2318	.2275
local t-bill	.0070	.0028	.0014	.0129
Hungary 2558 obs; Feb 1994 – Dec 2007; No of Companies: 9-18				
stock_rt	.0173	.1619	-.9000	2.2605
market_rt	.0250	.0969	-.3606	.5809
local t-bill	.0128	.0066	.0045	.0283
Poland 4751 obs; Feb 1994 – Dec 2007; No of Companies: 9-37				
stock_rt	.0158	.1621	-.9280	1.8950
market_rt	.0204	.1064	-.3526	1.0593
local t-bill	.0110	.0066	.0032	.0278
Slovak Republic 1476 obs; Feb 1996 – Dec 2007; No of Companies: 2-20				
stock_rt	.0020	.1912	-.9811	2.1395
market_rt	-.0001	.0610	-.1708	.3582
local t-bill	.0104	.0053	.0022	.0217

Chart 1: Average Number of Stocks

The annual average number of stocks (companies) for each in the Visegrad countries is reported. The sample spans from February 1994 to December 2007. The number of stocks in each month varies from 2 to 74.

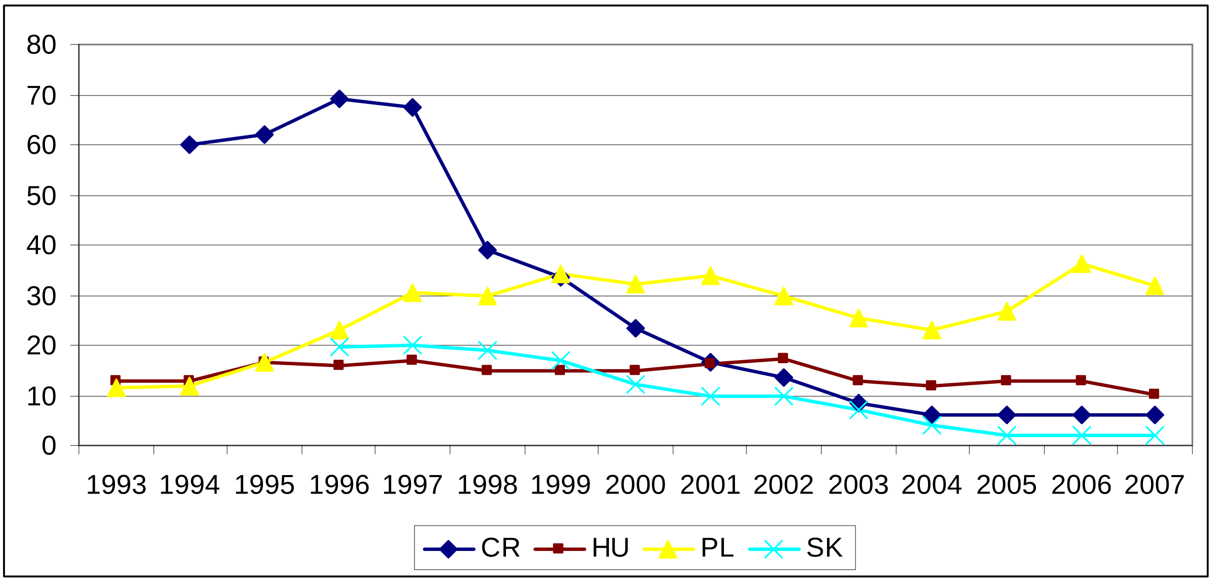

Table 2: Summary Statistics - Multi-Factor Models

Sample mean, standard deviation, maximum and minimum values are reported for the additional variables used in the multi-factor model regression (market return, stock return and local t-bill statistics are reported in Table 1). These statistics are reported for the cross sectional distribution, where the number of firms varies from 2 to 74 depending on the country. All the variables represent monthly returns or growth rates in local currency. The time series for the term structure was obtained by subtracting the monthly return on treasury bills from the monthly return on long-term government bonds. In the subsequent statistical analysis CPI, inflation and the term structure are used in first differences since their original time series contain unit roots. Indprod stands for monthly industrial production growth rate, infl represents monthly growth in inflation, ts is the term structure, exrate is the monthly appreciation/depreciation of the national currency as compared to the euro, ger_indprod stands for the monthly industrial production growth in Germany, money represents monthly growth in M1, commod stands for monthly growth in all primary commodities index, and exports is the monthly growth in exports.

Variable	Obs.	Mean	Std. Dev.	Min.	Max.
Czech Republic; Feb 1994 – Dec 2007; No of Companies: 6-74					
indprod	4936	.0088	.0935	-.2390	.2141
infl	4942	.0056	.0068	-.0078	.0402
ts	4942	.0005	.0011	-.0032	.0036
exrate	4942	.0008	.0181	-.0536	.0677
ger_indprod	4936	.0061	.0747	-.1130	.1914
money	4936	.0068	.0329	-.1049	.2437
exports	4942	.0217	.1137	-.3016	.2922
commod	4955	.0035	.0299	-.0990	.0907
Hungary; Feb 1993 – Dec 2007; No of Companies: 9-18					
indprod	2548	.0117	.0915	-.1952	.2619
infl	2558	.0093	.0089	-.0039	.0439
ts	1857	-.0014	.0012	-.0042	.0009
exrate	2558	.0056	.0201	-.0508	.1220
ger_indprod	2548	.0050	.0718	-.1130	.1914
money	2548	.0122	.0294	-.1013	.0938
exports	2538	.0284	.1287	-.2921	.4532
commod	2559	.0059	.0351	-.0990	.0907
Poland; Feb 1993 – Dec 2007; No of Companies: 9-37					
indprod	4751	.0082	.0661	-.1989	.2091
infl	4751	.0064	.0085	-.0090	.0560
ts	4624	-.0010	.0016	-.0055	.0016
exrate	4751	.0022	.0259	-.0426	.1192
ger_indprod	4719	.0052	.0706	-.1130	.1914
money	4719	.0167	.0345	-.1086	.1554
exports	4719	.0201	.0907	-.1692	.2913
commod	4752	.0068	.0375	-.0990	.0907
Slovak Republic; Feb 1996 – Dec 2007; No of Companies: 2-20					
indprod	1476	.0057	.0621	-.1441	.1621
infl	1476	. 0060	.0093	-.0037	.0569
ts	1358	.0003	.0019	-.0041	.0079
exrate	1358	.0010	.0135	-.0370	.0358
ger_indprod	1474	.0057	.0715	-.1130	.1650
money	1474	.0076	.0460	-.1158	.3779
exports	1476	.0187	.1004	-.2309	.2492
commod	1513	.0028	.0352	-.0990	.0907

Table 3: Summary Statistics - Principal Factor Models

Sample mean, standard deviation, maximum and minimum values are reported for the first three leading factors obtained from the principal component analysis. These statistics are reported for the cross sectional distribution, where the number of firms varies from 2 to 74 depending on the country.

Principal components	Mean	Std. Dev.	Min.	Max.
Czech Republic 4962 obs; Jun 1992 – Dec 2007; No of Companies: 6 -74				
Pc 1	.3872365	.879905	-14.35263	4.966973
Pc 2	.0057085	.8266291	-6.668375	6.712875
Pc 3	.0201583	.5822934	-5.526254	1.997139
Hungary 2566 obs; Jun 1992 – Dec 2007; No of Companies: 9 -18				
Pc 1	-.1417332	3.612419	-7.868495	8.045865
Pc 2	.0096652	2.542095	-11.84293	8.084753
Pc 3	.0277519	2.139743	-5.654095	11.74568
Poland 4759 obs; Jun 1992 – Dec 2007; No of Companies: 9-37				
Pc 1	-.5479955	3.648634	-10.95179	10.8843
Pc 2	.0013293	2.120843	-7.976968	5.431263
Pc 3	.0214359	1.58124	-4.862857	5.372016
Slovak Republic 1521 obs; May 1992 – Dec 2007; No of Companies: 2-20				
Pc 1	.4861237	1.011218	-15.63167	2.581877
Pc 2	-.102012	.8812424	-4.37315	8.389556
Pc 3	.0333495	1.125391	-6.032537	7.89641

Table 4: Results - CAPM

We estimated the CAPM by the FMB procedure using the following regression equation: $r_{it}\text{-}r_t^{\,f} = \alpha_i + \beta_i^m \left(r_t^{\,m} - r_t^{\,f}\right) + \varepsilon_{it}$, where r_{it} is the i-th stock return (i =1….N), $r_t^{\,f}$ is the risk-free rate, $r_t^{\,m}$ is the market return, α_i is the constant term and ε_{it} is the error term. In the first stage, we regressed the excess stock return $r_{it} - r_t^{\,f}$ on the excess market return $r_t^{\,m} - r_t^{\,f}$ in order to obtain the CAPM betas, $\hat{\beta}_i$, where *i* is a firm's subscript. These beta estimates were then used in the second stage as the independent variables in the following regression equation: $r_i - r^f = \gamma_0 + \gamma^m \hat{\beta}_i^m + \eta_i$. This regression was repeated for each month and we obtained *T* estimates of γ_0 and γ^m. Specifically, first beta estimates were obtained for the first 24 months of data and then used to calculate the gammas (γ_0 and γ^m) for the twenty-fourth month. Then, the betas were obtained for the period from second to twenty-fifth month, and used in the second stage to obtain the estimate of gammas for the twenty-fifth month. This procedure of rolling regressions with a fixed window of twenty-four months was used to cover the whole sample of data. Finally, we tested the averages of these *T* estimates to see if: $E\left(\gamma_0\right) = 0$ (i.e. pricing errors are zero) and $E\left(\gamma^m\right) > 0$ (i.e. positive risk premium on excess market return). Monthly return on a local index was used as a proxy for the market portfolio's return and the monthly return on a local t-bill was used as a risk-free rate. In the table below we report average slopes and t-statistics (in parentheses) from month-by-month regressions of excess stock returns on the betas of excess market returns. *, **, *** indicate significant differences at the 10, 5, and 1 percent levels, respectively.

Country	**Sample**	**Local index (γ^m)**	**Constant (γ_0)**
Czech Republic	Feb 1994 - Dec 2007	0.0000 (0.6305)	0.0063 (1.7814**)
Hungary	Feb 1993 - Dec 2007	0.0140 (1.5430*)	-0.0099 (-1.3281*)
Poland	Feb 1993 - Dec 2007	0.0034 (0.4008)	-0.0034 (-0.5485)
Slovak Republic	Feb 1996 - Dec 2007	-0.0822 (-2.2423**)	0.0696 (2.7071***)

Table 5: Results - Baseline Multi-Factor Models

We estimated the four-factor model by the FMB procedure using the following regression equation: $r_{it}\text{-}r_t^{\,f} = \alpha_i + \beta_i^m\left(r_t^{\,m} - r_t^{\,f}\right) + \beta_i^2\left(r_t^{\,2}\right) + \ldots\ldots + \beta_i^4\left(r_t^{\,4}\right) + \varepsilon_{it}$, where $r_t^{\,m}$ is a local index return, $r_t^{\,k}$ is the *k-th* factor return (*k=2…4*), r_{it} is the i-th stock return, $r_t^{\,f}$ is the risk-free rate, α_i is the constant term, and ε_{it} is the error term. The monthly return on a local index was used as a proxy for the market portfolio's return and the monthly return on a local t-bill was used as a risk-free rate. We considered the following four factors: excess market return, inflation, industrial production, and term structure. All series represent monthly growth rates or monthly returns. Inflation and term structure are used in first differences since the unit root tests detected nonstationarity in these series. Similarly to the CAPM, this multi-factor model predicts that the constant terms should be insignificant and the slope coefficients should be significantly different from zero. In the first stage, we regressed the excess stock return $r_{it} - r_t^{\,f}$ on the four factors in order to obtain the betas, $\hat{\beta}_i^m$ and $\hat{\beta}_i^f$, where *i* is the firm's subscript, *m* indicates the excess market return, and *f* is the factor's subscript *(f=2…4).* These beta estimates were then used in the second stage as the independent variables in the following regression equation: $r_i - r^f = \gamma_0 + \gamma^m\hat{\beta}_i^m + \gamma^f\hat{\beta}_i^f + \eta_i$. This regression was repeated for each month and we obtained *T* estimates of γ_0, γ^m and γ^f (for each of the factors *f*). Specifically, first beta estimates were obtained for the first 24 months of data and then used to calculate the gammas (γ_0, γ^m and γ^f) for the twenty-fourth month. Then, the betas were obtained for the period from the second to the twenty-fifth month, and used in the second stage to obtain the estimate of gammas for the twenty-fifth month. This procedure of rolling regressions with a fixed window of twenty-four months was used to cover the whole sample of data. Finally, we tested the averages of these *T* estimates to see if: $E\left(\gamma_0\right) = 0$ (i.e. pricing errors are zero), $E\left(\gamma^m\right) > 0$ and $E\left(\gamma^f\right) > 0$ (i.e. positive risk premium on excess market return and other factors *f*). In the table below we report average slopes and t-statistics (in parentheses) from month-by-month regressions of excess stock returns on thebetas of excess market returns, inflation, industrial production, and term structure. *, **, *** indicate significant differences at the 10, 5, and 1 percent levels, respectively.

Country	Sample	Excess Market Return (γ^m)	Inflation (γ^2)	Ind. Prod. (γ^3)	Term Structure (γ^4)	Constant (γ_0)
Czech Republic	Feb 1994 - Dec 2007	0.0016 (0.2035)	-0.0010 (-0.6687)	-0.0115 (-0.4836)	-0.0001 (-1.0935)	0.0002 (0.0301)
Hungary	Mar 1997 - Dec 2007	0.0111 (1.0264)	-0.0018 (-1.0926)	-0.0071 (-0.3117)	0.0001 (1.1913)	-0.0081 (-0.9357)
Poland	Feb 1993 - Dec 2007	0.0089 (1.1506)	0.0017 (2.0011**)	-0.0065 (-0.9165)	0.0000 (-1.0032)	-0.0054 (-0.9771)
Slovak Republic	Feb 1996 - Nov 2003	-0.0168 (-1.1553)	-0.0072 (-1.5274*)	0.0279 (1.3154*)	-0.0006 (-1.2609)	0.0196 (1.2624)

Table 6: Results - Additional Multi-Factor Models

We estimated the multi-factor model by the FMB procedure using the following regression equation: $r_{it}\text{-}r_t^{\,f} = \alpha_i + \beta_i^m\left(r_t^{\,m} - r_t^{\,f}\right) + \beta_i^2\left(r_t^{\,2}\right) + \ldots\ldots + \beta_i^8\left(r_t^{\,8}\right) + \varepsilon_{it}$, where $r_t^{\,m}$ is a local index return, $r_t^{\,k}$ is the *k-th* factor return (*k=2…8*), r_{it} is the i-th stock return, $r_t^{\,f}$ is the risk-free rate, α_i is the constant term, and ε_{it} is the error term. The monthly return on a local index was used as a proxy for the market portfolio's return and the monthly return on a local t-bill was used as a risk-free rate. Compared to the baseline four-factor model (results in Table 5), we added up to five additional variables, including: exchange rate, German industrial production, money, commodity index, and exports. All series represent monthly growth rates or monthly returns. Inflation and the term structure are used in first differences (unless otherwise indicated) since the unit root tests detected nonstationarity in these series. Similarly to the CAPM, this multi-factor model predicts that the constant terms should be insignificant and the slope coefficients should be significantly different from zero. In the first stage, we regressed the excess stock return $r_{it} - r_t^{\,f}$ on the four factors in order to obtain the betas, $\hat{\beta}_i^m$ and $\hat{\beta}_i^f$, where *i* is a firm's subscript, *m* indicates the excess market return, and *f* is the factor's subscript *(f=2…8)*. These beta estimates were then used in the second stage as the independent variables in the following regression equation: $r_i - r^f = \gamma_0 + \gamma^m\hat{\beta}_i^m + \gamma^f\hat{\beta}_i^f + \eta_i$. This regression was repeated for each month and we obtained *T* estimates of gammas (γ_0, γ^m and γ^f) for each of the factors *f*. Specifically, first beta estimates were obtained for the first 24 months of data and then used to calculate the gammas for the twenty-fourth month. Then, the betas were obtained for the period from the second to the twenty-fifth month, and used in the second stage to obtain the estimate of gammas for the twenty-fifth month. This procedure of rolling regressions with a fixed window of twenty-four months was used to cover the whole sample of data. Finally, we tested the averages of these *T* estimates to see if: $E(\gamma_0) = 0$ (i.e. pricing errors are zero), $E(\gamma^m) > 0$ and $E(\gamma^f) > 0$ (i.e. positive risk premium on excess market return and other factors *f*). In the table below we report average slopes and t-statistics (in parentheses) from month-by-month regressions of excess stock returns on the betas of excess market returns, inflation, industrial production, term structure, exchange rate, German industrial production, money and exports. *, **, *** indicate significant differences at the 10, 5, and 1 percent levels, respectively.

Excess market return	Inflation	Ind. Prod.	Term Structure	Exchange rate	German Ind. Prod.	Money	Exports	Commodity Index	Constant
Czech Republic; Feb 1994 - Dec 2007 Term structure, inflation and money in first differences									
0.0044		-0.0511				-0.0215	-0.0412		-0.0034
(0.4786)		(-2.0116**)				(-2.5540***)	(-1.3224*)		(-0.3803)
Hungary; Feb 1993 - Dec 2007 Term structure, inflation and money in first differences									
0.0055	0.0000			-0.0073					0.0003
(0.5885)	(-0.0192)			(-1.9198**)					(0.0345)
0.0137				-0.0061		-0.0153		0.0133	-0.0075
(1.2578)				(-1.6248*)		(-1.3693*)		(2.0324**)	(-1.0332)
Poland; Feb 1993 - Dec 2007 Term structure and inflation in first differences									
0.0075	0.0018					-0.0046			-0.0020
(0.9788)	(2.4678***)					(-1.3227*)			(-0.3682)
0.0087	0.0011					-0.0062		0.0030	-0.0056
(0.9623)	(1.4898*)					(-1.6624**)		(0.8225)	(-0.9906)
Slovakia; Feb 1996 - Dec 2007 Term structure, inflation, money, and exchange rate in first differences									
0.0032	-0.0122	0.0372	-0.0011			0.0319		-0.0163	-0.0122
(0.2086)	(-1.8047**)	(1.2154)	(-2.2190**)			(1.2187)		(-0.9600)	(-1.0330)
-0.0256		0.0149	-0.0012			0.0389			0.0192
(-1.7500**)		(0.7169)	(-2.3922***)			(1.4128*)			(1.2408)

Table 7: Results - Principal Factor Models

We estimated the principal model by the FMB procedure using the following regression equation: $r_{it}\text{-}r_t^{\,f} = \alpha_i + \beta_i^m\left(r_t^{\,m} - r_t^{\,f}\right) + \beta_i^2\left(r_t^{\,2}\right) + \ldots\ldots + \beta_i^4\left(r_t^{\,4}\right) + \varepsilon_{it}$, where $r_t^{\,m}$ is a local index return, $r_t^{\,k}$ is the *k-th* principal factor return (*k=2…4*), r_{it} is the i-th stock return, $r_t^{\,f}$ is the risk-free rate, α_i is the constant term, and ε_{it} is the error term. The monthly return on a local index was used as a proxy for market portfolio's return and the monthly return on a local t-bill was used as a risk-free rate. We obtained the principal factors using principal component analysis. Then, we used the first three as the principal factors in the asset-pricing model. Similarly to the CAPM, this multi-factor model predicts that the constant terms should be insignificant and the slope coefficients should be significantly different from zero. In the first stage, we regressed the excess stock return $r_{it} - r_t^{\,f}$ on the excess market return and on the three first principal factors in order to obtain the betas ($\hat{\beta}_i^m$ and $\hat{\beta}_i^f$), where *i* is a firm's subscript, *m* indicates the excess market return, and *f* is the principal factor's subscript *(f=2…4)*. These beta estimates were then used in the second stage as the independent variables in the following regression equation: $r_i - r^f = \gamma_0 + \gamma^m\hat{\beta}_i^m + \gamma^f\hat{\beta}_i^f + \eta_i$. This regression was repeated for each month and we obtained *T* estimates of γ_0, γ^m and γ^f (for each of the factors *f*). Specifically, first beta estimates were obtained for the first 24 months of data and then used to calculate the gammas (γ_0, γ^m and γ^f) for the twenty-fourth month. Then, the betas were obtained for the period from the second to the twenty-fifth month, and used in the second stage to obtain the estimate of gammas for the twenty-fifth month. This procedure of rolling regressions with a fixed window of twenty-four months was used to cover the whole sample of data. Finally, we tested the averages of these *T* estimates to see if: $E\left(\gamma_0\right) = 0$ (i.e. pricing errors are zero), $E\left(\gamma^m\right) > 0$ and $E\left(\gamma^f\right) > 0$ (i.e. positive risk premium on excess market return and other factors *f*). In the table below we report average slopes and t-statistics (in parentheses) from month-by-month regressions of excess stock returns on the betas of excess market returns and the first three principal factors. *, **, *** indicate significant differences at the 10, 5, and 1 percent levels, respectively.

Country	**Sample**	**Excess Market Return** (γ^m)	**Pc1** (γ^2)	**Pc2** (γ^3)	**Pc3** (γ^4)	**Constant** (γ_0)
Czech Republic	Jun 1992 – Dec 2007	0.0025 (0.3007)	0.1556 (0.5982)	-0.4919 (-1.5475*)	0.1262 (0.5825)	-0.0036 (-0.5204)
Hungary	Jun 1992 – Dec 2007	0.0143 (1.6145*)	-0.0961 (-0.3263)	-0.1590 (-1.3168*)	0.2186 (1.4278*)	-0.0076 (-1.2128)
Poland	Jun 1992 – Dec 2007	0.0127 (1.8266**)	0.2290 (1.3839*)	-0.2090 (-1.6572**)	-0.2424 (-2.2612**)	-0.0089 (-1.8761**)
Slovak Republic	Feb 1996 – Oct 2003	0.0092 (0.6514)	-0.4116 (-0.9029)	-0.3366 (-0.9837)	0.0182 (0.0464)	-0.0095 (-0.7632)

Table 8: Results - GMM

We estimated the GMM system, in which moment restrictions allowed for a joint estimation of betas and gammas, as specified in the FMB. Based on Cochrane (2005) we wrote the moment conditions in the following way:

$$gT(b) = \begin{bmatrix} E\left(R_{ti}{}^{e} - a_i - \beta_i{}^{m} f_t{}^{m} - \beta_i{}^{2} f_t{}^{2} - \ldots - \beta_i{}^{k} f_t{}^{k}\right) \\ E\begin{bmatrix} \left(R_{ti}{}^{e} - a_i - \beta_i{}^{m} f_t{}^{m} - \beta_i{}^{2} f_t{}^{2} - \ldots - \beta_i{}^{k} f_t{}^{k}\right) f_t{}^{m} \\ \left(R_{ti}{}^{e} - a_i - \beta_i{}^{m} f_t{}^{m} - \beta_i{}^{2} f_t{}^{2} - \ldots - \beta_i{}^{k} f_t{}^{k}\right) f_t{}^{2} \\ \cdots\cdots\cdots\cdots\cdots\cdots \\ \left(R_{ti}{}^{e} - a_i - \beta_i{}^{m} f_t{}^{m} - \beta_i{}^{2} f_t{}^{2} - \ldots - \beta_i{}^{k} f_t{}^{k}\right) f_t{}^{k} \end{bmatrix} \\ E\left(R_{ti}{}^{e} - \beta_i{}^{m}\gamma^{m} - \beta_i{}^{2}\gamma^{2} - \ldots - \beta_i{}^{k}\gamma^{k}\right) \end{bmatrix} = \begin{bmatrix} 0 \\ 0 \\ 0 \end{bmatrix},$$

where $R_{ti}{}^{e}$ is the i-th stock excess return (i-th stock return minus risk-free rate), $f_t{}^{m}$ is a local index excess return (local index return minus risk-free rate), and $f_t{}^{k}$ is the *k-th* factor return (*k=2…4).* In this system there are *N(1+K+1)* moment conditions since for each asset *N* we have one moment condition for the constant, *K* moment conditions for *K* factors and one moment condition that allows estimation of the gammas (asset-pricing model condition). On the other hand, there are N(1+K)+K parameters and hence we have n-K overidentifying restrictions. They can be tested with a chi-square test (J-test for overidentiffying restrictions). We estimated this system by GMM, in which the optimal weighting matrix was obtained in an iterative process by sequential updating of the coefficients. To test the overall model we calculated the J-statistic in the following way: $T * J = T * \left\lfloor g_T\left(\hat{b}\right)\hat{S}^{-1} g_T\left(\hat{b}\right)\right\rfloor$, where $g_T\left(\hat{b}\right)$ are the moment conditions evaluated at the estimated values of coefficients β and γ, whereas $\hat{S}^{-1}$ is the inverse of the optimal weighting matrix (variance-covariance matrix). This statistic follows approximately a χ^2 distribution with degrees of freedom equal to number of moment conditions minus the number of parameters. In our case the J-statistic was equal to *17.4*, whereas the χ^2 critical value at the 95 percent level of significance with 39 degrees of freedom was *18.5*. Since the J-statistic was less than the appropriate critical value we could not reject the model. In the table below we report average slopes and t-statistics (in parenthesis) from a one-step GMM estimation of excess stock returns on excess market returns, inflation, industrial production, and term structure for Poland with a time span of January 1993 to February 2003. *, **, *** indicate significant differences at the 10, 5, and 1 percent levels, respectively.

Excess Market Return (γ^m)	**Inflation** (γ^2)	**Ind. Prod.** (γ^3)	**Term Structure** (γ^4)
-0.0935 (-2.3566**)	-0.0073 (-1.9595*)	-0.0400 (-1.8555*)	-0.0016 (-2.3029**)

Size and Value Effects in the Visegrad Countries

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Abstract

The paper has two main objectives. The first is to test for the presence of the size and book-to-market value effects in the Visegrad countries. Such effects have been found in the United States and many other developed stock markets. The Visegrad countries consist of the Czech Republic, Hungary, Poland, and Slovakia. We demonstrate that size and value do in fact explain the expected return/cost of capital in Eastern Europe. Based on this result, we proceed by constructing regional size and book-to-market portfolios for a combined Visegrad market. Returns on these portfolios serve as factors in addition to the market portfolio. The regional three-factor model performs as well as country-specific versions of the model. However, it can be estimated for a more current sample in Prague, Warsaw, Budapest, and Bratislava. Therefore it is a plausible model for the cost of capital in this region and we use it to calculate the cost of capital for the following industries: banks; capital goods; food, beverage and tobacco; materials; and utilities.

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1 Introduction

The Capital Asset Pricing Model (CAPM) is a standard model to calculate the cost of capital. The CAPM is very intuitive, describes the trade-off between risk and variance, and is a theoretically well-founded equilibrium model. Practical implementation was very straightforward when the world stock market was essentially equivalent to the US stock market. In this case, one would use some measure of the US value-weighted market return such as the S&P 500 to calculate the risk premium, which is needed to compute the cost of capital. The situation changed in the 1970s, in part due to floating exchange rates and lifted barriers to the movement of capital across borders. This change altered the usage of the CAPM. Shultz (1995a, b) argues that instead of a local country-specific CAPM, one needs to use a global international version of the CAPM with a proxy for the global portfolio such as the Morgan Stanley Capital International Index.

Even the country-specific implementation of the CAPM can be problematic. According to the CAPM the only thing that matters to investors is market risk. However, numerous authors have found that other, non-market factors matter for stock returns. Fama and French (1992) argue that a firm's size and book-to-market equity help explain the cross-section of average returns in the US market.[1] Building on this finding, Fama and French (1993, 1996) construct two factors: SMB (a return on a portfolio of small stocks minus the return of a portfolio of big stocks) and HML (a return on a portfolio of high value

[1]The CAPM has been extensively empirically tested since the beginning of its existence. Early on, during the 1960s and 1970s, the CAPM seemed to capture well both the cross-section and time-series properties of asset returns. One of the first deviations from this conclusion was the size effect identified in Banz (1981). The nature of the size effect lies in the observation that the returns on stocks of relatively small firms tend to do better then predicted by the CAPM. In other words, the market beta is not sufficient to capture the cross-section of asset returns. Similar observations were made using other firm-specific characteristics. Bhandari (1988) finds that there is a positive relation between average returns and leverage, even after controlling for size measured by the market equity (ME). Rosenberg, Reid, and Lanstein (1985) show that the ratio of a firm's book value of equity (BE) to size is also positively related to average returns. Controlling for size and the market beta, Basu (1983) documents that the earnings-price ratio (E/P) adds to the explanatory power of the cross-section of returns. All such evidence is brought together in Fama and French (1992). In a univariate setting, they confirm a strong relationship between the average returns and the market beta, ME, leverage, E/P, and BE/ME. In multivariate regressions, it is ME and BE/ME, whose relation with average returns persists.

stocks minus the return of a portfolio of low value stocks). These factors related to size and the book-to-market ratio when used in a three-factor model together with the market return are able to explain a large part of the variance of average stock returns.

The apparent empirical success of the so-called "Fama and French factors" (FF factors) have started an on-going debate in the literature on the economic meaning of these factors. Several theories regarding the FF factors have emerged. Fama and French (1996) argue that firm size and book-to-market, which underline the SMB and HML factors, proxy for firm distress. Daniel and Titman (1997) point out that the FF factors pick up the co-movements of stocks with similar characteristics. They argue that firm characeristics such as size and book-to market explain stock returns better than the FF factors constructed using these characteristics. Rolph (2003) and Ferguson and Shockley (2003) attribute the significance of FF factors to the leverage effects they explain. On the other hand, Chung, Johnson, Herb and Schill (2004) argue that FF factors proxy for higher order co-moments and become insignificant when these co-moments are included in the model. The significance of these higher order co-moments an be explained by the risk-aversion of the investors, who take into account extreme outcomes.

There is also a strain of literature that argues that FF factors may be mistakenly taken as useful risk factors, while in fact they have no power in explaining any of the fundamental risk. Lakonishok, Shleifer, and Vishny (1994) argue that the empirical significance of the FF factors is due to sub-optimal behaviour of the investors rather than fundamental risk. Others point to the fact that explanatory power of the FF factors may be simply due to the construction of these factors. Berk (1995) argues that high book-to-market and small companies by construction will earn higher mean returns. In a similar mode, Ferson, Sarkissian, and Simin (1999) demonstrate that if stocks are sorted by some attributes that are empirically related to stock returns, such portfolios will likely appear as useful risk factors even if these attributes have nothing to do with fundamental risk. Therefore, they caution against using empirical regularities as "self-explanatory risk factors".

Some argue that a real test for the empirical importance of FF factors is to use them in an international context and to verify whether they still appear to have explanatory power. Fama and French (1998) have themselves attempted to extend their three-factor model to a global context. They find that value stocks (stocks with high ratios of book-to-market equity, earnings to price or cash flows to price) have higher returns than growth stocks (stocks with low ratios of book-to-market equity, earnings to price or cash flows to price) in twelve out of the thirteen developed markets they study. They also perform some out-of-sample tests for emerging markets and confirm that the value premium is also present in these markets. The international CAPM (ICAPM) is not able to explain this value premium in international markets but a two-factor model with global market return and a risk factor for relative distress can account for this phenomenon. Fama and French argue that this two-factor model provides a parsimonious way of summarizing the general patterns in international returns and is not meant to challenge the existing asset-pricing theory.

Vassalou and Liew (2000) use data from 10 developed markets and find that FF factors predict future GDP growth even when business cycle variables are included in the analysis. Thus, their evidence supports a risk-based hypothesis for the performance of FF factors. Daniel, Titman and Wei (2001) argue that its the firm characteristics not the factor loadings that explain stock returns in Japan. Griffin (2002) poses an interesting question of whether FF factors are global or country-specific and he argues for the latter. Specifically, he finds that country-specific versions of the Fama and French three-factor model can explain the variation in stocks returns (both portfolios and individual stocks) better than a world three-factor model. Therefore, he argues that cost of capital calculations, performance measurement, and risk analysis using models stylized on Fama-French three-factor models should be performed on a within-country basis.

The present study analyzes financial markets emerging in Eastern and Central Europe. Specifically, the focus is on the Visegrad countries, i.e. the Czech Republic, Hungary, Poland, and Slovakia. Our objectives in this case are a confirmation of stylized facts for stock returns on the four markets and a formulation of a plausible model of the cost of capital. We start with size and value effects. Similarly to Fama and French (1993) for the US market and Connor and Sehgal (2001) for the Indian market, we first construct regional size- and value-sorted portfolios. The stock markets in Prague, Warsaw, Bratislava and Budapest are fairly small in size, which in all countries expect for Poland severely reduces sample size. Therefore, we adopt a different approach and first combine all the markets in one to increase the overall market value, as well as the number of traded stocks. Only then we rank the stocks according to capitalization and book-to-market ratios using data pooled from all four Visegrad countries. This approach is likely to work for a fairly homogeneous market and also in the situations when a local market is too small to calculate the relevant factor returns without a large measurement error. The stock markets emerging in the Czech Republic, Hungary, Poland, and Slovakia are good candidates for this type of analysis. Moreover, these markets have not yet been studied from this perspective. Our investigation of the portfolio properties reveals that the size and value effects are present in our data. Even though the effects (especially the size effect) are smaller than in the case of the US data, the found patterns are qualitatively similar to those reported in Fama and French (1993). Next we conduct the Fama and MacBeth (1973) method (FMB henceforth) for individual stock returns in our sample starting in the early 1990s and demonstrate the existence of size and value effects in the four countries.

In the next step, we construct a regional multi-factor model of expected returns for the Visegrad countries. Fama and French (1998) and Griffin (2002) construct the factors within each country in his sample and then uses their value-weighted averages. We again combine the stocks in all countries first and only then construct the regional market, size, and book-to-market factors. In time series regressions with six size- and value-sorted portfolios as dependent variables, the regional CAPM is easily outperformed by the regional Fama and French three-factor model. This provides additional evidence in favor of the presence of value and size effects in the Visegrad countries. We compare the regional and country-specific factor models, with the local models being estimated using a restricted series due to data limitations. The regional CAPM does a better job at explaining the expected returns than the local version of the model. There is no difference in the performance of the regional and country-specific Fama and French three factor models. However, the regional model coefficients can be estimated for a current sample and hence provides a suitable cost of capital model for the Visegrad countries. We confirm this conclusion by running time series regressions with individual stock returns as dependent variables. The market factor is significant in a majority of the cases though the size and value factors are mainly important when a local risk free rate is used as a reference but not when a risk-free rate of an outside investor is employed (Germany's in our case). Finally, we compute the cost of capital for five major industries for all four countries to illustrate the use of our regional three-factor model and demonstrate that it works well for other than size- and value-related excess returns.

The rest of the paper is organized as follows. Section 2 discusses the relevant pricing models, Section 4 provides details of the implemented econometric methodology, Section 3 describes our data sources, Section 5 comments on our results and Section 6 concludes.

2 Pricing Models

Our primary objective is the calculation of the cost of capital in the Visegrad countries. A natural start in this context is a country-specific Capital Asset Pricing Model (CAPM):

$$E[R_i] = R_F + b_i E[MRF^C], \tag{1}$$

where R_i denotes a (stock) return on an asset i. MRF^C is the market return for a given country (hence the superscript C) in excess of the risk-free rate:

$$MRF^C = R_m^C - R_F, \tag{2}$$

with R_m^C being a return on a local stock market index, and R_F the risk-free rate. b_i is the sensitivity to the country-specific market index (beta). $E[R_i]$ is the cost of capital. Stultz (1995a,b) points out that the single-factor country-specific CAPM is valid only in a fairly closed stock market. The alternative in this case is a global CAPM where the local market index would be replaced by of a global market index as a measure of R_m.

The single-factor CAPM equation (1) is theoretically well founded but not as successful empirically. It has been demonstrated that firm-specific information is needed to achieve more accurate expected stock returns. Fama and French (1992) use all major previously used variables and provide evidence that it is mainly size and book-to-market ratio that are important in explaining the time-series and cross-sectional properties of stock returns. Based on this evidence, Fama and French (1993) construct the following country-specific three-factor model:

$$E[R_i] = R_F + b_i\ E[MRF^C] + s_i\ E[SMB^C] + h_i\ E[HML^C]. \tag{3}$$

SMB^C is a premium on small stocks vs. big stocks (*Small Minus Big*) and HML^C is a premium on the stocks with high book-to-market ratio vs. stocks with low book to market ratios (*High Minus Low*). The superscript C stresses the country-specific nature of the two additional factors.

Fama and French (1998) also propose a global version of their three-factor model (3):

$$E[R_i] = R_F + b_i\ E[MRF^G] + s_i\ E[SMB^G] + h_i\ E[HML^G]. \tag{4}$$

HML^G is again the value premium. Fama and French (1998) assume there is no size related factor in this case ($s_i \equiv 0$) since the size effect has proved to be spurious in the international context while the value effect has been shown to be more robust. A crucial point in the formulation of the global Fama and French model (4) is the calculation of $HML.^G$ They first compute a value-weighted difference in returns of value portfolios with high book-to-market ratios and growth portfolios with low book-to-market ratios within a country. HML^G is then a weighted average across countries where weights are given by the capitalization of each country within a sample of all countries. Griffin (2002) uses the model (4) with $s_i \neq 0$ and in some specifications allows for a different impact of domestic and foreign components of the three factors.

Assuming that there are value and size effects in our four countries (this assumption is confirmed below), we can propose a version of the Fama and French three-factor model. However, both the country-specific specification (3) and the global specification (4) of the model require a calculation of the country-specific factors. This is problematic for the stock markets in the Visegrad countries. Only a small number of stocks is traded on each date, which may induce unnecessary variability in the three risk premia. To circumvent this problem, we consider the following specification:

$$E[R_i] = R_F + b_i\; E[MRF^R] + s_i\; E[SMB^R] + h_i\; E[HML^R]. \tag{5}$$

We will refer to (5) as the Regional Three Factor Fama and French Model (R3FFFM). The novel feature of this model is the calculation of the regional size and value factors SMB^R and HML^R, respectively. We first combine stocks from all the Visegrad countries and form a regional stock market. Only then do we rank them by size and book-to-market ratio and compute the corresponding risk premia. The regional market excess return MRF^R is calculated as a value or equally weighted average of returns on the local market indexes. All returns are in Euros.

3 Data

The stock market variables are from the S&P Emerging Markets Database (EMDB). Individual returns are calculated from indexes for individual stocks. The index is given by (closing price*100)/intitial closing price. The closing price is the price in local currency per share of the stock at the end of the last trading day. The prices are not adjusted to changes in capitalization such as a two-for-one split. To calculate returns and capitalization in Euros, exchange rates from International Financial Statistics are used. The local index

R_m^C is computed in the same fashion as R_i using the major indexes for the Visegrad countries: PX 50 (Prague, Czech Republic), BUX (Budapest, Hungary), WIG (Warsaw, Poland), and SAX (Bratislava, Slovakia). The market capitalization or market equity (ME) is calculated for each stock as the number of shares outstanding*1000000*closing price. Then it is converted into Euros. Book equity (also referred to as net worth or book value) is simply the difference between total assets and total liabilities. In markets with high inflation, S&P may make adjustments to net worth for inflation. BE/ME is the book value per share divided by the stock price.

The number of observations is in Figure 1. The total number of observations in all countries reached 100 only in the years 1997-1999. The number of traded items has been under 20 in all countries except for Poland over the whole sample starting in 1993. Sample mean, standard deviation, maximum and minimum values are reported by a country for the variables used in the FMB and time series regressions in Table 1. These statistics are reported for the cross sectional distribution, where the number of firms in a given month varies from 2 to 74 depending on the country. The average capitalization of firms is similar across countries though the book-to-market values differ. The only country with negative mean returns for individual stocks and a market index is the Czech Republic. Volatilities of the market indexes are similar across countries. Volatilities of individual stock returns are higher and again similar across countries with the exception of Slovakia. Table 2 presents the summary statistics for the combined sample. The market return is now the value-weighted regional market return with weights given by capitalization in Euros. For excess returns used later, we need a measure of the risk-free rate. In most cases, we use the rate for German government bonds. Since they are not available for the last four months in our sample (2007:9-2007:12), we employ the LIBOR rate minus the average risk premium of the LIBOR rate over the German government bond rate until 2007:8. For the country-specific versions of the CAPM and the three-factor FF model, it is also appropriate to use local risk-free rates. We get the treasury bill rate for each country from the International Financial Statistics database maintained by the IMF (with the exception of Slovakia for which the data comes from Global Financial Data).

Fama and French (1993) introduce two additional factors in addition to the market excess return, which are motivated by the size and value effects. Similar factors are calculated for the combined stock markets in the Visegrad countries. To construct these factors, all available stocks are divided into two groups based on median market equity (size), Small (S) and Big (B). The stocks are also divided into three groups based on the 30th and 70th percentile of their median book-to-market equity ratios (BE/ME) into High (H), Medium (M), and Low (L) categories. This gives us six groups of stocks which can be used to form six corresponding portfolios. We denote them SH, SM, SL, BH, BM, and BL, respectively. The return on each of these portfolios is an equally-weighted return on stocks in the corresponding groups. The Fama and French (1993) factors are then calculated as

$$SMB_t^R = 1/3\ [R_t^{SH} + R_t^{SM} + R_t^{SL}] - 1/3\ [R_t^{BH} + R_t^{BM} + R_t^{BL}] \tag{6}$$

and

$$HML_t^R = 1/2\ [R_t^{SH} + R_t^{BH}] - 1/2\ [R_t^{SL} + R_t^{BL}], \tag{7}$$

where R_t^{SH} denotes the equally weighted return on a portfolio of stocks, which belong to the small size and high book-to-market categories. The remaining returns are denoted in a similar fashion. The returns are equally weighted, following Connor and Sehgal (2001, who cite suggestions from Lakonishok, Shliefer and Vishny 1994) that these portfolios tend to perform better in explaining the stock returns. This suggestion is also confirmed in Fama and French (1996).

The summary statistics for all these portfolio returns are reported in Table 3. There is a negative relation between size and average returns but only for high value stocks. The relation between size and average returns for low and medium value stocks is positive. There is a positive relation between value and average return irrespective of size. Therefore the size effect seems to be conditional on value, while there is a strong unconditional value effect. The fact that the size effect depends on value makes the SMB^R return negative, contrary to what is observed in the United States (see Fama and French 1993).

This outcome is consistent with studies that find the size effect spurious internationally while the value effect is robust.[2] Chan, Yasushi, and Lakonishok (1991) find the value effect to be pervasive in Japan. The first broader international evidence of the presence of value effect is documented by Capaul, Rowley, and Sharpe (1993). This evidence is then confirmed on a larger group of countries and longer time span of data by Fama and French (1998). They conduct out-of-sample tests for the value premium in international markets and conclude that this premium is present in thirteen major markets, as well as in emerging markets.

Skewness of all returns except on BL is positive (skewness is zero for a normal distribution). Kurtosis is greater than 3 (the value for a normal distribution) in all cases, which indicates that extreme values are more likely - the distribution has 'fat-tails'. This is most extreme for the BH return. HML is negatively correlated with the market return - see Table 4. The correlation between SMB^R with the market and HML^R is negative but small. We have also formed the six portfolios for all the countries individually. The premia can be reasonably calculated for the whole sample (until 2007:12) only for Poland. The sample ends in 2005:6 for the Czech Republic and in Hungary, and in 2004:6 for Slovakia. However, the stock returns from these countries are included when the regional premia are constructed.

Finally, we also construct industry portfolios. A typical number of industries considered in various studies of the US market range from 30 to 48. However, the number of observations in the four markets in our sample is not sufficient for such a fine distinc-

[2] We have also considered the price-to-earnings ratio, which is often used to distinguish between value and growth stocks similarly to the book-to-market ratio. The price-to-earnings ratio is priced and this adds another piece to the international evidence in favor of the value effect. Since our focus is on the universally accepted HML portfolio and there are some potential issues with construction of the price-to-earnings based portfolios, we do not pursue this issue further.

tion and we therefore compute the cost of capital only for five industries in each country. Our five industries are banks, capital goods, food beverage and tobacco, materials, and utilities. We calculate equally weighted returns on stocks for a given industry in a given country if there is at least one observation available.

4 Econometric Methodology

The Fama and MacBeth (1973) method is used to verify if size and book-to-market ratios in fact matter in the stock markets of the Czech Republic, Hungary, Poland, and Slovakia. Consider a cross-sectional regression:

$$R_t = \gamma_t' e + \delta_t \hat{F}_t + \varepsilon_t, \tag{8}$$

where R_t is an $N \times 1$ vector of gross stock returns, e an $N \times 1$ vector of ones, and $\hat{F}_t$ an $K \times 1$ vector of estimated factors. N is given by the number of traded companies on all the four stock markets. γ_t and δ_t are vectors of parameters conformable with e and F_t, respectively. ε_t is an $N \times 1$ vector of error terms. The factors are coefficients from the following time series regression, estimated for a two-year period ending at time t:

$$R_{it} = \phi_i + F_t' \, X_{it} + \epsilon_{it}, \quad i = 1, ..., N, \tag{9}$$

with ϵ_{it} an i.i.d. error term. X_{it} is a $K \times 1$ vector of explanatory variables. $K = 1, 2, 3$ and X includes various combinations of the regional market excess return and firm measures of size (capitalization) and of a book-to-market ratio. Our calculations show that replacing returns in (8) and (9) with excess returns does not change the results. The cross-sectional regression (8) is run at each time $t = 1, ..., T$, so we have sequences of parameter estimates $\{\hat{\gamma}_t, \hat{\delta}_t\}_{t=1}^{T}$. If the time-series average of $\hat{\delta}_i$ is statistically different from zero, the factor F_i is priced. The t-statistic for this test is given by

$$t_{\delta_i} = \frac{\tilde{\hat{\delta}}_i}{\tilde{\sigma}_i}, \tag{10}$$

where

$$\tilde{\delta}_i = \sum_{t=1}^{T} \hat{\delta}_i / T \;\text{ and }\; \tilde{\sigma}_i^2 = \frac{1}{T(T-1)} \sum_{t=1}^{T} (\hat{\delta}_i - \tilde{\delta}_i)^2 \;\; i = 1, ..., K. \tag{11}$$

It is expected that all the three factors are priced.

Estimates of coefficients from the cross-sectional regression (8) are subject to a measurement error because the factors have to be estimated from the time series regression (9). A standard way of mitigating this problem is a joint estimation of equations (8) and (9) by the Generalized Method of Moments (GMM). The moment conditions follow Cochrane (2005). For an asset i and K factors, the conditions can be written as:

$$\begin{bmatrix} E[R_{it} - \phi_i - F_t' \, X_{it}] & = & 0 \\ E[(R_{it} - \phi_i - F_t' \, X_{it}) X_{it1}] & = & 0 \\ \vdots & & \\ E[(R_{it} - \phi_i - F_t' \, X_{it}) X_{itK}] & = & 0 \\ E[R_{it} - \delta_t \, F_t] & = & 0 \end{bmatrix}. \tag{12}$$

Therefore, for N assets and K factors, we have $N(K+2)$ orthogonality conditions. There are $N(K+1)+K$ parameters to be estimated. Consequently, there are $N-K$ overidentifying restrictions for the Hansen J chi-square statistic. Another possibility for reducing the errors-in-variables problem is the use of the Shanken (1992) correction. However, Shanken and Weinstein (2006) suggest that there is a potential issue with a negative positive cross-product matrix of the OLS residuals from the time series regressions. Therefore we do not attempt to use the correction. Moreover, we run into a similar issue when we try to estimate the risk premia using the GMM approach (see Section 5).

In the next step, we follow Connor and Sehgal (2001) and run the following time series regressions of the six size and book-to-market portfolios on the market excess return and returns on our two newly-constructed factor-portfolios:

$$R_t^{XY} - R_t^F = a^{XY} + b^{XY} \, MRF_t^R + s^{XY} \, SMB_t^R + h^{XY} \, HML_t^R + \nu_t, \tag{13}$$

where X can be S or L and Y can be H, M, or L. We also follow Fama and French (1993) and estimate regression (13) for individual stocks i, i.e.

$$R_{it} - R_t^F = a_i + b_i \, MRF^R + s_i \, SMB_t^R + h_i \, HML_t^R + \eta_{it}, \quad i = 1, \ldots, N. \tag{14}$$

Finally, we use the constructed factors to calculate the cost of capital in the Visegrad countries for individual industries. The cost of capital is based on estimates of sensitivities from the following time-series regressions of industry excess returns:

$$R_{kt} - R_t^F = a_k + b_k\ MRF_t^R + s_k\ SMB_t^R + h_k\ HML_t^R + \eta_{kt}, \quad k = 1, \ldots, K. \tag{15}$$

The cost of capital is then the expected industry return:

$$E[R_{kt}] = E[R_t^F] + \hat{b}_k + \hat{b}_k\ E[MRF_t^R] + \hat{s}_k\ E[SMB_t^R] + \hat{h}_k\ E[HML_t^R], \quad k = 1, \ldots, K. \tag{16}$$

The expected results of this exercise is twofold. First, to the best of our knowledge, this is the first rigorous attempt to calculate the cost of capital in these countries. Second, this exercise will enable us to evaluate the performance of the three-factor regional model with dependent variables not sorted by value and/or size but by an industry. R^2 from regression (15) could be a good measure of this performance. The confidence interval for the cost of capital is calculated by regressing the industry excess returns on de-meaned excess returns from (15) plus a constant, and using the 95% confidence interval of the intercept.

Note that the econometric methodology is in part chosen to address the potential problem of biased regression coefficients. In the time series regressions, excess returns on size and book-to-market-sorted portfolios are regressed on excess returns on factor portfolios also related to these stock characteristics. This problem is addressed in several steps. First, the Fama and MacBeth repeated cross-sectional regressions of returns of individual stocks (i.e. not size- or value-related) on size and the book-to-market ratio document the importance of size and value effects. Second, Fama and French (1993) argue that a finer sorting for dependent excess returns implies that there would be no bias in their regressions and document this with a number of robustness checks. In our setting, there are only six size- and value-sorted dependent excess returns due to data limitations. The data limitations also prevent us from doing robustness checks of the sort done by Fama and French since splitting the sample in a number of ways is problematic for the already-small number of stocks. What we do instead is run the time series regression for individual stocks and look at the t-statistics of coefficients. Finally, we follow Fama and French (1997) and use the three-factor returns to calculate the cost of capital for industry related returns, i.e. use an entirely different sorting of the dependent variables.

5 Results

We first investigate whether capitalization and book-to-market ratios are priced. Table 5 reports the results of the Fama and MacBeth method described in the previous section. Each row corresponds to a particular combination of risk factors whose number is $K = 1, 2$ or 3. In the full model with market excess return, capitalization, and the book-to-market ratio, the two latter variables have risk premia significantly different from zero. The market variable is only priced when joined by either the size or the value variables separately. Interestingly, when capitalization and book-equity divided by market equity are considered without the market, they are not significant. There is enough evidence in Table 5 to believe that size and value matter in the Visegrad countries.

We have tried to confirm this finding by GMM to reduce the potential error-in-variables problem. For all countries, the GMM never converged due to a singular weighting matrix either using TSP or EViews. This is clearly the same problem discussed by Shanken and Weinstein (2006). They refer to OLS residuals but one just needs to realize that OLS is but a special case of GMM. In an attempt to analyze this technical issue, we have focused only on Poland. In this case, there are 66 usable stocks and 6 moment conditions for each (see 12), which amounts to 396 moment conditions. At the same time, there are only 179 time series observations. It is clear that a longer sample would be needed for a sensible usage of GMM. Nevertheless, our results regarding the role of size and value stand because the measurement error typically causes a downward bias in the estimates of cross-sectional coefficients.

In the next step, we estimate the time series regression (13) to find whether the newly constructed factors improve the performance of the CAPM. Results are reported in Table 6. First we estimate a standard version of the CAPM with the Visegrad value-weighted excess return as the explanatory variable and the six size- and and book-to-market-sorted portfolio excess returns. The market beta is strongly significant in all cases, which indicates that the joint market variable does have some explanatory power in explaining the time series variation of returns. However, adding the other two factors improves the performance of the model as measured by the adjusted R^2 in all cases. The sensitivities to the size factor are positive for small companies and negative for large companies. Fama and French (1993) consider US data and a 5×5 division of stocks based on size and value vs. our 2×3 division. They find negative betas for the size factor for firms in the biggest capitalization quintile. Similarly, we find that the value effect is negative for low and medium book-to-value ratios. Fama and French (1993) make a similar observation with the negative effect appearing in the first two quintiles of the book-to-value ratios. The intercepts are significantly negative with the exception of R^{SH} for the CAPM when the intercept is significantly positive. This is in contrast to zero for a well-specified asset pricing model where all factors are portfolio (excess) returns. Also, Fama and French (1993) report higher R^2's and mostly insignificant constants in the country-specific version of their model using US data.

The comparison of our regional model with a country-specific Fama and French US model shows that the model behaves similarly quantitatively though it is somewhat less successful empirically. More appropriate would be a comparison of our regional model with its country-specific versions. Table 7 reports adjusted R^2's from country-specific models. We make several observations. First, the regional CAPM means a significant improvement over the local CAPM in 20 out of 24 cases. Second, the inclusion of the size and value factors increases the R^2 in 24 out of 24 cases. Finally, the regional three-factor model performs better than a country-specific version only in 12 out of 24 cases. However, this means that the regional model is at least as good as the local model but it can be estimated for the whole sample for all countries. The country-specific model can be only estimated for the whole sample for Poland.

As another examination of the regional model performance, we estimate equation (14) for each stock in our sample (see Table 8). For country-specific risk-free rates, the average absolute t-statistic for the size and value factors is slightly below the 10% critical value of the t-distribution, which is 1.68. The sensitivities for the two factors are significant in 37% of the cases and the constant is significant only in 11% of the cases, both documenting fairly good performance of the Fama-French Visegrad model. The results do not support the use of the size and value factors to this extent for the German risk-free rates since the factor loadings are significant only in 14% of the cases for the size factor and 15% of the cases for the value factor.

Finally, Table 9 reports the cost of capital for the five industries for each country. The cost of capital is calculated including the intercept from the time series regression (15) for the sake of consistency with an estimation of the confidence intervals. There are not enough observations to calculate the cost of capital for Slovakia for the food, beverage, and tobacco industry and for utilities. We also do not have enough information to compute the expected return for utilities in Poland. The cost of capital is significantly negative only twice, for banks and capital-goods companies in the Czech Republic. The latter is likely a result of data issues related to a short data series ending in 2001. The negative cost of capital reflects the possibility that Czech banks are undervalued according to our model; assuming that the pricing error (the intercept) is zero, the expected return on investment in the Czech banks is 1% annually. Coefficients of the regional SMB portfolio are significant in 8 out of 17 cases and coefficients of the regional HML portfolio in 10 out of 17 cases. Apparently, the factors matter in the expected return regressions. The signs of the coefficients vary and are probably related to the size of companies in particular countries - this would be a question for future research.

6 Summary

The present study analyzes the stock markets in the Czech Republic, Hungary, Poland, and Slovakia. The first objective is to document the presence of size and value effects in the Visegrad countries. We use the Fama and MacBeth (1973) method to demonstrate that size and book-to-market equity are in fact priced in all Visegrad countries. Based on this result, we construct a three-factor regional Fama and French (1993) model. The model factors include the market, the size, and the value premia. The main innovation is constructing these variables for a regional market across the four Visegrad countries. The model behaves qualitatively similarly to the US-calibrated Fama and French model, replicating the size and value effects. However, the size effect is somewhat less pronounced and the model's R^2 is smaller as compared with the model whose coefficients are estimated using US stocks.

Next, we compare the performance of this model with a country-specific CAPM and a country-specific three-factor Fama and French model. Our dependent variables are six size- and book-to-market-sorted portfolios across our four markets. The regional model outperforms the country-specific CAPM and is comparable to the country-specific three-factor model. However, the main advantage of the regional model is the possibility of the calculation of the premia for the full sample, which is not possible for the country-specific versions of this model because the markets in the Visegrad countries are thin, shallow, and with small capitalization. In this sense the regional model is superior and can be readily used as a basis for cost-of-capital calculations. We also investigate the performance of the model using individual stock returns. The constant is almost never significant, which implies good model performance. However, the size and value premia are significant more frequently when local risk-free rates are used as opposed to the German risk free rates. Interestingly, while the regional CAPM provides a major improvement over the country-specific CAPM, this is not the case for the three-factor model. It seems that the sensitivity to size and value premia is important but it is not relevant whether the premia are local or regional.

Finally, we use the regional three-factor model to calculate the cost of capital for the following industries: banks, capital goods, food, beverage and tobacco, materials, and utilities. The regional market, SMB, and HML factors are significant in these calculations as well.

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Figure 1: Number of Observations for Stock Returns

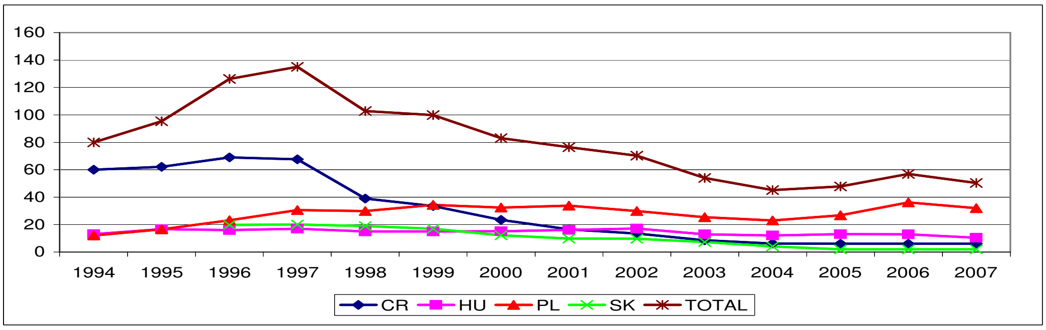

Table 1: Summary Statistics of the Stock and Market Variables by Country

All the variables are reported in euros (ECU before 1999). R_i stands for an individual stock return, R_m^C is a value-weighted country index market index, ME is market equity, BE/BE stands for book-to-market equity, and the *exrate* represents the exchange rate of the local currency against the euro.

Variable	Mean	Std. Dev.	Min	Max
Czech Republic; 4942 obs; 1994:02 - 2007:12; up to 79 stocks				
R_i	-.0067	.1679	-.9839	1.5562
R_m^C	-.0048	.0737	-.2798	.1962
$\log ME$	18.12	1.76	12.75	24.14
$\log BE/ME$	.3229	1.0601	-2.8898	3.9120
exrate	33.7806	2.7645	26.2610	38.6249
Hungary; 2559 obs; 1993:01-2007:12 up to 29 stocks				
R_i	.0118	.1651	-.9137	2.2073
R_m^C	.0157	.0906	-.2798	.5988
$\log ME$	18.78	1.90	14.59	23.28
$\log BE/ME$	-.3721	.7718	-2.5634	1.7148
exrate	222.5687	42.9413	101.7910	283.3500
Poland; 4751 obs; 1993:02-2007:12 up to 66 stocks				
R_i	.0136	.1673	-.9469	1.9149
R_m^C	.01304	.0766	-.2798	.5987
$\log ME$	19.90	2.37	14.98	29.54
$\log BE/ME$	-.6126	.7850	-4.2977	3.9120
exrate	3.7267	.5023	1.9290	4.8721
Slovak Republic; 1476 obs; 1996:02-2007:12 up to 23 stocks				
R_i	.0071	.3022	-1.0040	8.9833
R_m^C	.0023	.0668	-.2798	.1506
$\log ME$	17.17	1.78	10.07	21.18
$\log BE/ME$	.8774	1.3371	-6.1051	4.60517
exrate	39.8628	2.6817	33.1610	45.5000

Table 2: Summary Statistics for Stock Returns in the Combined Visegrad Market, 13728 obs.

All the variables are reported in euros (ECU before 1999). R_i stands for an individual stock return, R_m is a value-weighted country index market index, ME is market equity, and BE/BE stands for book-to-market equity.

Variable	Mean	Std. Dev.	Min	Max
R_i	.0052	.1866	-1.0040	8.9833
R_m	.0052	.07578	-.2728	.5076
$\log ME$	18.76	2.22	10.07	29.54
$\log BE/ME$	-.0708	1.0917	-6.1051	4.6052

Table 3: Summary Statistics for Portfolio Returns 1993:07-2007:12

S and B denote small and big capitalization, respectively. H, M, and L are high, medium, and low book-to-market ratios. R_{XY} are returns on portfolios with size X and book-to-market ratio Y.

Return	Mean	St. Dev.	Skeweness	Kurtosis
R^{SL}	-.0009	.0966	1.61	10.20
R^{SM}	-.0016	.0520	.4163	4.69
R^{SH}	.0087	.0834	2.35	16.12
R^{BL}	.0009	.0856	-.17	7.43
R^{BM}	.0042	.0829	.51	12.72
R^{BH}	.0055	.1146	6.50	99.59
R_m^R	.0047	.0743	.26	7.61
SMB^R	-.0016	.0520	.41	4.69
HML^R	.0071	.0797	.86	13.13

Table 4: Correlations for Portfolio Returns, 1993:07-2007:12

	SMB^R	HML^R
R_m^R	-0.33	-0.06
SMB^R		-0.03

Table 5: Results of Fama and MacBeth (1973) Regressions, Sample 1993:07-2007:12

R_m	$\log ME$	$\log BE/ME$	const.
0.0068	0.0332	-0.0269	-0.0054
(1.1151)	(3.2135***)	(-2.8438***)	(-1.0429)
0.0062			0.0009
(0.9685)			(0.2032)
	0.0263		0.0022
	(2.5495***)		(0.4473)
		-0.0120	0.0055
		(-0.7911)	(0.9408)
0.0091	0.0492		-0.0084
(1.5059*)	(3.9651***)		(-1.7070**)
0.0081		-0.0275	-0.0047
(1.3454*)		(-2.2818**)	(-1.0230)
	0.0117	-0.0059	0.0056
	(0.8556)	(-0.4143)	(0.8502)

Table 6: Regional Factor Models, 1993:07-2007:12

Based on the regression equation

$$R_t^{XY} - R_t^F = a^{XY} + b^{XY}\ MRF_t^R + s^{XY}\ SMB_t^R + h^{XY}\ HML_t^R + \nu_t,$$

where X is either S or L and Y is H, M, or L. $s^{XY} \equiv 0$ and $h^{XY} \equiv 0$ for CAPM. t-statistics are reported in parentheses.

Dependent Var.	a^{XY}	b^{XY}	s^{XY}	h^{XY}	Adj. R^2
$R^{SL} - R^F$	-0.0091	0.7733			0.32
	(-12.58***)	(79.48***)			
$R^{SM} - R^F$	-0.0080	0.8952			0.58
	(-16.44***)	(136.90***)			
$R^{SH} - R^F$	0.0027	0.7374			0.40
	(4.64***)	(95.12***)			
$R^{BL} - R^F$	-0.0076	1.1019			0.75
	(-18.72***)	(202.24***)			
$R^{BM} - R^F$	-0.0031	1.0650			0.78
	(-8.49***)	(219.22***)			
$R^{BH} - R^F$	-0.0003	0.9193			0.33
	(-0.37)	(81.78***)			
$R^{SL} - R^F$	-0.0044	0.8821	0.6774	-0.4097	0.54
	(-7.41***)	(104.23***)	(56.71***)	(-56.27***)	
$R^{SM} - R^F$	-0.0089	0.9670	0.2540	0.1297	0.61
	(-19.03***)	(145.10***)	(27.01***)	(22.62***)	
$R^{SH} - R^F$	-0.0018	0.9554	0.6937	0.5556	0.81
	(-5.55***)	(205.62***)	(105.78***)	(138.98***)	
$R^{BL} - R^F$	-0.0048	0.9539	-0.4820	-0.3538	0.90
	(-18.45***)	(259.93***)	(-93.06***)	(-112.04***)	
$R^{BM} - R^F$	-0.0030	0.9699	-0.3945	-0.0516	0.83
	(-9.44***)	(212.93***)	(-61.37***)	(-13.16***)	
$R^{BH} - R^F$	-0.0074	0.8807	-0.4984	0.6809	0.60
	(-11.32***)	(95.29***)	(-38.21***)	(85.63***)	

Table 7: Adjusted R^2 for Country-specific Factor Models

Based on the regression equation

$$R_t^{XY} - R_t^F = a^{XY} + b^{XY}\ MRF_t^C + s^{XY}\ SMB_t^C + h^{XY}\ HML_t^C + \nu_t,$$

where X is either S or L and Y is H, M, or L. $s^{XY} \equiv 0$ and $h^{XY} \equiv 0$ for CAPM.

Dependent Var.	Regional M. 93:7-07:12	Pol. 93:7-7:12	Czech R. 94:7-05:6	Hungary 94:7-05:6	Slovakia 96:7-04:6
			CAPM		
$R^{SL} - R^F$	0.32	0.37	0.18	0.15	0.01
$R^{SM} - R^F$	0.58	0.59	0.30	0.46	0.01
$R^{SH} - R^F$	0.40	0.61	0.26	0.29	0.00
$R^{BL} - R^F$	0.75	0.75	0.68	0.84	0.23
$R^{BM} - R^F$	0.78	0.73	0.70	0.61	0.45
$R^{BH} - R^F$	0.33	0.34	0.38	0.14	0.01
			Three-factor Model		
$R^{SL} - R^F$	0.54	0.62	0.55	0.45	0.59
$R^{SM} - R^F$	0.61	0.73	0.57	0.68	0.32
$R^{SH} - R^F$	0.81	0.84	0.83	0.83	0.91
$R^{BL} - R^F$	0.90	0.85	0.68	0.87	0.28
$R^{BM} - R^F$	0.83	0.77	0.75	0.62	0.45
$R^{BH} - R^F$	0.60	0.70	0.69	0.68	0.47

Table 8: T-statistics from Regressions of Individual Stock Excess Returns on the Visegrad Market Excess Return, and on Returns on the Size and Value Factor Portfolios

Based on the regression equation

$$R_{it} - R_t^F = a_i + b_i\,(R_{mt} - R_t^F) + s_i\,SMB_t + h_i\,HML_t + \eta_{it}, \quad i = 1, \ldots, N.$$

	t_{a_i}	t_{b_i}	t_{s_i}	t_{h_i}
Country-specific risk-free rates				
Avg. abs. values	0.83	4.18	1.46	1.50
Ratio of sig. values	0.11	0.86	0.37	0.37
German risk free rate				
Avg. abs. values	-0.01	1.01	0.1	0.06
Ratio of sig. values	0.03	0.86	0.14	0.15

Table 9: Industry Cost of Capital

Based on the time-series regressions: $R_{kt} - R_t^F = a_k + b_k\ MRF_t^R + s_k\ SMB_t^R + h_k\ HML_t^R + \eta_{kt}, \quad k = 1, \ldots, K$. The cost of capital is defined as: $E[R_{kt}] = E[R_t^F] + \hat{a}_k + \hat{b}_k\ E[MRF_t^R] + \hat{s}_k\ E[SMB_t^R] + \hat{h}_k\ E[HML_t^R], \quad k = 1, \ldots, K$. CR is the Czech Republic, HU Hungary, PL Poland, and SK Slovakia. The cost of capital is annualized, in %.

Country	Sample	$\hat{a}^k$ (se)	$\hat{b}^k$ (se)	$\hat{s}^k$ (se)	$\hat{h}^k$ (se)	Cost of capital (95% conf.int)
			Banks			
CR	94:2-07:12	-0.01	0.86	-0.38	0.07	-17
		(-1.96**)	(9.78***)	(-3.34***)	(-0.89)	(-30,-3)
HU	96:2-07:12	0.01	1.06	-0.01	-0.31	18
		(-0.89)	(6.53***)	(-0.07)	(-2.09**)	(-4,41)
PL	93:7-07:12	0.01	1.19	-0.82	-0.32	23
		(3.07***)	(22.85***)	(-11.04***)	(-7.37***)	(15,32)
SK	96:2-07:12	0.01	0.35	0.33	0.14	14
		(-0.81)	(1.92*)	(-1.29)	(-0.84)	(-12,40)
			Capital Goods			
CR	94:2-01:9	0.00	1.30	0.40	0.32	-25
		(-0.88)	(17.11***)	(4.19***)	(4.58***)	(-38,-13)
HU	93:7-03:10	-0.01	0.81	0.11	-0.03	-4
		(-1.33)	(8.41***)	(-0.76)	(-0.43)	(-23,16)
PL	93:7-03:10	-0.01	1.06	-0.54	-0.24	-11
		(-1.47)	(11.64***)	(-3.95***)	(-3.32***)	(-28,6)
SK	96:2-02:8	0.00	0.49	0.87	0.61	-26
		(-0.33)	(2.17**)	(2.86**)	(2.91**)	(-58,6)
			Food Beverage & Tobacco			
CR	94:2-02:10	0.00	0.95	0.19	0.14	-12
		(-0.41)	(8.58***)	(1.32)	(1.37)	(-31,6)
HU		0.01	1.04	0.09	0.00	13
		(0.79)	(11.26***)	(0.63)	(-0.00)	(-5,31)
PL	93:7-02:10	0.01	0.70	-0.72	-0.49	7
		(0.95)	(9.16***)	(-6.14***)	(-8.07***)	(-8,22)
			Materials			
CR	94:2-07:12	0.00	0.91	0.09	0.22	-6
		(-0.22)	(12.90***)	(0.97)	(3.46***)	(-17, 6)
HU	96:12-07:12	0.00	1.03	0.04	-0.09	9
		(0.06)	(10.59***)	(0.29)	(-1.07)	(-5,23)
PL	93:7-07:12	0.00	1.31	-0.32	-0.31	11
		(0.66)	(18.23***)	(-3.11***)	(-4.93***)	(-1,22)
SK	96:2-03:9	0.01	0.73	0.19	0.50	-1
		(0.65)	(4.71***)	(0.91)	(3.46***)	(-25,22)
			Utilities			
CR	94:2-00:10	0.01	0.96	0.05	0.10	1
		(2.07**)	(11.47***)	(0.48)	(1.33)	(-12,15)
HU	95:2-00:10	0.00	1.23	0.27	-0.40	-6
		(0.22)	(5.69***)	(0.90)	(-2.19**)	(-40,27)