Charles University in Prague

Faculty of Social Sciences
Institute of Economic Studies

MASTER THESIS

The Exchange Rate Pass-Through in Central and Eastern Europe

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Academic Year: **2013/2014**

Bibliographic record

Mirková, B. (2014). *The Exchange Rate Pass-Through in Central and Eastern Europe*. (Master thesis). Charles University in Prague.

Volume: 97 069 characters

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Acknowledgments

I would like to express my greatest gratitude to my supervisor doc. Roman Horváth Ph.D. for his guidance and valuable comments. Apart from that, I would like to thank the people closest to me for their immense support and encouragement.

Abstract

This thesis examines the exchange rate pass-through into consumer prices in Central and Eastern Europe. The study is based on quarterly data of 12 countries from 2003 to 2013. Estimations are conducted using heterogeneous panel cointegration methods, namely the mean group and the pooled mean group estimators. Fixed effects are used as a reference. The thesis provides short-run and long-run estimates of the exchange rate pass-through for the individual countries and for the region as a whole. Based on the results, we conclude that the exchange rate pass-through is highly variable across Central and Eastern Europe. We find that there is no clear distinction between the pass-through rates in euro area countries, EU countries not using the euro and non-EU countries. Further, we find that the generally accepted concept of higher exchange rate pass-though in developing countries does not hold in this region.

JEL Classification	C23, E31, E52, F31
Keywords	exchange rate pass-through, pooled mean group, mean group, heterogeneous panel cointegration

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Abstrakt

Diplomová práce zkoumá exchange rate pass-through, tedy vliv změn směných kurzů na spotřebitelské ceny, ve střední a východní Evropě. Práce je založena na čtvrtletních datech 12 zemí od roku 2003 do roku 2013. Odhady jsou provedeny pomocí metod heterogenní panelové kointegrace, konkrétně mean group a pooled mean group estimátorů. Fixed effects jsou použity pro srovnání. Diplomová práce obsahuje krátkodobé a dlouhodobé odhady exchange rate pass-through pro jednotlivé země i region jako celek. Výsledky ukázaly, že se hodnoty exchange rate pass-through velmi liší pro jednotlivé státy střední a východní Evropy. Došli jsme však k závěru, že hodnoty exchange rate pass-through se neliší v závislosti na tom, jestli jsou země členy eurozóny, členy EU mimo eurozónu nebo zeměmi mimo EU. Naše výsledky ukázaly, že ve střední a východní Evropě neplatí obecný názor, že exchange rate pass-through je v rozvojových zemích vyšší než v rozvinutých.

Klasifikace JEL	C23, E31, E52, F31
Klíčová slova	směnné kurzy, spotřebitelské ceny, heterogenní panelová kointegrace, pooled mean group, mean group

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Acronyms

ADF	Augmented Dickey-Fuller test
AIC	Akaike information criterion
ARDL	Autoregressive distributive lag
BIC	Bayesian information criterion
CEE	Central and Eastern Europe
CPI	Consumer price index
DFE	Dynamic fixed effects
DOLS	Dynamic ordinary least squares
EC	Error correction
ERPT	Exchange rate pass-through
EU	European Union
FE	Fixed effects
FMOLS	Fully modified ordinary least squares
GDP	Gross domestic product
IFS	International Financial Statistics
IID	Independent and identically distributed
IMF	International Monetary Fund
IPS	Im, Pesaran and Shin test
IR	Interest rate
LM	Lagrange multiplier
LR	Long run
MG	Mean group
NEER	Nominal effective exchange rate
OECD	Organisation for Economic Co-operation and Development
OLS	Ordinary least squares

PMG	Pooled mean group
PPI	Producer price index
SR	Short run
SUR	Seemingly unrelated regressions
US	United States
VAR	Vector autoregression
VECM	Vector error correction model

Master Thesis Proposal

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Proposed Topic:

The Exchange Rate Pass-Through in Central and Eastern Europe

Topic Characteristics:

The exchange rate pass-through is an important indicator for many economic and monetary policy decisions. These include decisions about inflation targeting, currency devaluation in countries with fixed exchange rates, entering the monetary union and others. It is, therefore, a topic of high interest and importance.

There is an extensive empirical research on exchange rate pass-through in industrialized countries, mainly done by single-equation and vector autoregression models. In recent years, several studies using the panel cointegration methods have appeared. The Central and Eastern Europe is, however, not well covered in the literature. There are some studies that include countries from this region but these are usually a part of a broader worldwide study or are based only on few selected countries, often the members of the European Union. To my knowledge, a comprehensive study of exchange rate pass-through in the Central and Eastern European has not been conducted. The contribution of this thesis will lie not only in covering the so far somewhat neglected region but also in using the heterogenous panel cointegration methods, which are not yet commonly used in this context.

The data on effective exchange rates of the selected countries will be obtained from the Bank of International Settlements. Price indices will most likely be acquired from the national statistical bureaus.

Hypotheses:

1. What is the exchange rate pass-through in the Central and Eastern Europe?
2. The domestic prices are influenced by the changes in the exchange rate more in developing countries than the developed ones.
3. There are significant differences between the exchange rate pass-through for the euro area countries, countries that have not yet adopted the Euro and countries that are not members of the European Union.

Methodology:

The estimations of the exchange rate pass-through will be done using the mean group estimator developed by Pesaran and Smith (1995) and pooled mean group estimator developed by Pesaran, Shin and Smith (1999). Fixed-effects estimator will be used as reference. The mean group estimator and the pooled mean group estimator will be used to account for the heterogeneity that is likely to appear when conducting a panel data model across countries in different stages of economic development, in various phases of European integration, with different institutions etc.

The models differ in the degree to which they allow for heterogeneity. In the fixed-effects model only the intercepts are allowed to vary. The slopes are assumed to be the same for all countries. In the pooled mean group estimator, long-run coefficients are fixed, while short-run coefficients are allowed to vary. The mean group estimator allows for the largest heterogeneity; both intercepts and slopes are allowed to vary.

Outline:

1. Introduction
2. Exchange rate pass through
3. Methodology
4. Data description
5. Estimation and results
6. Conclusion

Core Bibliography:

Asteriou, D. (2006). *Applied econometrics : a modern approach using EViews and Microfit*. 1st. ed. New York: Palgrave Macmillan.

Beirne, J. and Bijsterbosch, M. (2011). Exchange rate pass-through in central and eastern European EU Member States. *Journal of Policy Modeling* [online]. 33, p.241-254. Available from: <http://www.sciencedirect.com/science/article/pii/S0161893810001067>. [Accessed 5 June 2013].

Blackburne, E. F. and Frank, M. W. (2007). Estimation of nonstationary heterogeneous panels. *The Stata Journal* [online]. 7 (2), p.197-208. Available from: <http://www.stata-journal.com/sjpdf.html?articlenum=st0125>. [Accessed 11 April 2013].

Campa, J. M. and Goldberg, L. S. (2002). *Exchange rate pass-through into import prices: Macro or micro phenomenon?*. (Research paper No 475). IESE Business School, Centro International de Investigacion Financiera, Barcelona. Available from: <http://www.iese.edu/research/pdfs/di-0475-e.pdf>. [Accessed 9 June 2013].

Cheikh, N. B. (2011). *Long run exchange rate pass-through: Evidence from new panel data techniques*. (MPRA Paper No. 39663). Université de Rennes 1, Centre de Recherche en Économie et Management,. Available from: <http://mpra.ub.uni-muenchen.de/39663/>. [Accessed 6 April 2013].

Pesaran, M. H., Shin, Y. and Smith R. P. (1999). Pooled mean group estimation of dynamic heterogeneous panels. *Journal of the American Statistical Association* [online]. 94 (446), p.621-634. Available from: <http://www.jstor.org/stable/2670182>. [Accessed 1 May 2013].

Pesaran, M. H. and Smith R. P. (1995). Estimating long-run relationships from dynamic heterogeneous panels. *Journal of Econometrics* [online]. 68 (1), p.79-113. Available from: <http://www.sciencedirect.com/science/article/pii/030440769401644F>. [Accessed 1 May 2013].

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1 Introduction

This thesis estimates the exchange rate pass-through in selected countries of Central and Eastern Europe. The exchange rate pass-through expresses to which extent changes in exchange rate affect domestic prices. Such dynamics are of high interest as their knowledge is important for decisions about several policy issues. These include a choice of exchange rate regime, adjustment in trade balances, international transmission of shocks and last but not least a set-up of monetary policy (Cheikh, 2013). The last applies especially to the exchange rate pass-through into consumer prices. The timing and the magnitude of the exchange rate pass-through plays an important role in inflation forecasting. Being able to properly forecast inflation dynamics is crucial for appropriate decisions on monetary policy. The Central and Eastern European region brings specific reasons for the examination of the exchange rate pass-through. These are mostly connected to the fact that majority of the countries from the region have either already joined the European Union or are planning to become members in the future. One of the questions connected with the accession to the European Union and the subsequent joining of the euro area, which can be answered by the knowledge of the exchange rate pass-through of the countries in question, is, whether the inflation convergence required by the Maastricht criteria is sustainable in the long run (Beirne and Bijsterbosch, 2011).

Even-though the benefits of the knowledge of the exchange rate pass-through dynamics in the Central and Eastern Europe are high, adequate attention of the exchange rate pass-through research has not been paid to the region. There are some studies that cover the Central and Eastern European countries, but the amount is significantly lower than the number of studies focusing on Western Europe and other more advanced economies. Moreover, majority of the available estimates are outdated. This thesis aims to provide up-to-date estimates of the exchange rate pass-through for selected countries of this somewhat neglected region. Specifically, we cover 12 Central and Eastern European countries from the first quarter of 2003 to the first quarter of 2013. Furthermore, this thesis uses the advanced methods of heterogeneous panel cointegration, namely the mean group and the pooled mean group estimator.

The aim of this thesis is to provide short-run and long-run estimates of the exchange rate pass-through both for individual countries and for the region as a whole, and to compare them with previous results obtained by other authors. Finally, two hypotheses are tested. First hypothesis states that the domestic prices are influenced by the changes in the exchange rate more in developing countries than the developed ones. The second hypothesis is that there are significant differences in the exchange rate pass-through among the euro area countries, countries that have not adopted the euro yet and countries that are not members of the European Union.

The remainder of the thesis is structured as follows. Chapter 2 reviews the existing literature and summarizes the previous exchange rate pass-through estimates for the countries of the Central and Eastern Europe. Chapter 3 discusses the theoretical approach to the exchange rate pass-through and Chapter 4 provides an overview of the methods of estimation. Chapter 5 includes a data description, empirical estimation and discussion of results. Chapter 6 concludes.

2 Literature review

2.1 Origins of exchange rate pass-through studies

The exchange rate pass-through started to become a topic of interest in the 1970s. One of the main reasons for the increased focus on the relationship of prices and exchange rates was interest in the impact of devaluation of domestic currency on the country's trade balance. Other reasons drove the increasing focus on the topic of prices and exchange rates as well – among others a desire to assess the validity of the theories of purchasing power parity and the law of one price and to investigate the impact of depreciation or appreciation of currency on domestic inflation (Goldberg and Knetter, 1996). The amount of conducted studies grew during the following decade. As both Menon (1995) and Goldberg and Knetter (1996) state, during the 1980s the research of the exchange rate pass-through focused mainly on pass-through into prices in the United States. Menon's (Menon, 1995) literature survey is an illustration of this. In the 43 studies he examined, there were 78 exchange rate pass-through estimates and out of these 78 estimates 27 belonged to the United States. The rest covered mainly other major industrialized countries, namely Japan, Australia, Germany, Canada, the United Kingdom, South Korea, France, the Netherlands, Belgium, Italy and few others with only one estimate. Developing countries were covered very seldom. Based on these studies, it is evident that in the early years of the exchange rate pass-through research the by far most common method of pass-through estimation were the ordinary least squares. Estimation of the exchange rate pass-through using the ordinary least squares can be problematic since in many cases the time series character of the data does not fulfill all assumptions of the ordinary least squares, especially the assumption of stationarity (Menon, 1995). In the studies reviewed by Menon (1995), exchange rate pass-through was calculated as pass-through into import prices, domestic producer prices or export prices.

2.2 Current research

The research of exchange rate pass-through has evolved over the past few decades and is currently much more diverse and sophisticated than it was in the 1980s and the first half of the 1990s. Following sections look at this diversity and development in terms of country coverage, data, methodology and findings. Overview of selected studies can be found in Table 1: Overview of selected literature.

Table 1: Overview of selected literature

Study	Data	Model	Method
Babecká (2009)	monthly data (1996 - 2006) for the Czech Republic	ERPT into consumer prices	VAR, VECM
Barhoumi (2005)	annual data (1980 - 2003) for 24 developing countries (none from CEE)	ERPT into import prices	non-stationary panel, PMG, MG, FMOLS, DOLS
Beirne and Bijsterbosch (2011)	monthly data (1/1995 - 4/2008) for nine central and eastern EU member states	ERPT to consumer prices	five-variate cointegrated VAR, impulse response derived from VECM
Billmeier and Bonato (2002)	monthly data (1/1994 - 1/2001) for Croatia	ERPT into consumer prices	cointegrated VAR
Bitāns (2004)	monthly data (1/1993 - 6/2003) for 13 East European countries	ERPT into producer and consumer prices	recursive VAR
Campa and Goldberg (2002)	quarterly data (1975 - 1999) for 25 OECD countries	ERPT into import prices	OLS
Ca'Zorzi, Hahn and Sánchez (2007)	quarterly data (period starting 1975 - 1993 and ending 2003 - 2004) for 12 emerging markets in Asia, Latin America and Central and Eastern Europe	ERPT into CPI and import prices	VAR
Cheikh (2011)	quarterly data (1/1994 - 4/2010) for 27 OECD countries	ERPT into import prices	FMOLS, DOLS
Coricelli, Jazbec and Masten (2006)	monthly data (1993 - 2002) for four EMU acceding countries	ERPT into CPI inflation	single equation and cointegrated VAR
Dabusinskas (2003)	quarterly data (1/1995 - 1/2003) for Estonia	ERPT into import, producer and consumer prices, total and disaggregated	OLS, SUR
Darvas (2001)	quarterly data (1993 - 2000) for 4 EU candidate countries	ERPT into consumer prices	VAR, OLS and simultaneous time-varying error correction framework
Holmes (2006)	monthly data (4/1972 - 6/2004) for 12 EU countries	ERPT of US dollar ER into consumer prices	panel data cointegration, DOLS

Holmes (2008)	annual data (since 1971) for 19 African countries	ERPT into import prices	panel data cointegration, FMOLS
Korhonen and Wachtel (2005)	monthly data (1/1999 - 11 or 12/2004) for 7 countries of Commonwealth of Independent States	ERPT into consumer prices	VAR without error correction terms
Kozluk, Banerjee and de Bandt (2008)	1995 - 2005 data for euro area countries	ERPT into import prices, disaggregate data - industry level	panel data cointegration
McCarthy (1999)	quarterly data (1/1976 - 4/1998) for 9 industrialized countries	ERPT into producer and consumer prices	VAR
Mumtaz, Oomen and Wang (2006)	quarterly data (1/1984 - 1/2004) for the United Kingdom	ERPT into import prices, disaggregate data - industry level	OLS
Schröder and Hüfner (2002)	monthly data (1981 - 2001) for 5 countries of the euro area	ERPT into consumer prices	VECM
Vonnák (2010)	monthly data (starting between 1995 and 1997) for the Czech Republic, Hungary and Poland	ERPT into consumer prices	VAR

Source: author’s compilation of information from the listed studies

2.2.1 Country coverage

The structure of the covered countries changed significantly. Naturally, the exchange rate pass-through for the United States and other major industrialized countries remained a topic of interest of many studies, but small developed economies and transitional and developing countries have been receiving growing attention. There are studies that focus only on one specific country (for example Mumtaz, Oomen and Wang (2006) who estimated the exchange rate pass-through into import prices in the United Kingdom in 1984 – 2004; Billmeier and Bonato (2002) who estimated the exchange rate pass-through to consumer prices in Croatia; Dabusinskas (2003) who estimated the exchange rate pass-through to import, producer and consumer prices in Estonia), but many studies have a multicountry character. These often focus on a certain type of countries or a region. Among the works dealing with the more industrialized economies, there are quite common studies covering selected OECD countries (e.g., Campa and Goldberg (2002), Cheikh (2011)) or selected European Union member states (e.g., Holmes (2006) – estimations for selected member states; Kozluk, Banerjee and de Bandt (2008) – estimations for members of the euro area; Beirne and Bijsterbosch (2011) – estimations for central and eastern EU member states). Exchange rate pass-through estimates for developing countries from various world regions are provided for example in Barhoumi (2005) (Africa, Asia, South and Central America) or Ca'Zorzi, Hahn and Sánchez (2007) (Asia, South and Central America, Central and Eastern Europe). Some authors focus on a comparison of exchange rate pass-through in a certain region. For example, Holmes (2008) studied 19 African countries and Korhonen and Wachtel (2005) calculated the exchange rate pass-through for seven countries of the Commonwealth of Independent States[1] and few other developing countries as a benchmark. The results of the later study are of an interest to us, since some of the countries of the Commonwealth of Independent States and some of the countries from the benchmark belong to Central and Eastern Europe (CEE). From the Central and Eastern European region, the countries most represented in the literature are members of the European Union or were EU candidates at the time the study was conducted. The number of estimates is highest for the Czech Republic, Hungary and Poland, followed by the Slovak Republic, Slovenia and Romania. These countries are often included in the multicountry studies along both major industrialized economies (estimates for the Czech Republic, Hungary and Poland can be found in Campa and Goldberg (2002); estimates for the Czech Republic, the Slovak Republic and Poland are available in Cheikh (2011)) and developing countries (estimates for the Czech Republic, Hungary and Poland can be found for example in Ca'Zorzi, Hahn and Sánchez (2007)). There are also studies specifically dedicated to the more developed countries of Central and Eastern Europe. Darvas (2001) and Coricelli, Jazbec and Masten (2006) estimated the exchange rate pass-through for the Czech Republic, Hungary, Poland and Slovenia. The work covering the most extensive sample of CEE countries (12 countries) is Bitāns (2004).

2.2.2 Data

The character of the data used for the exchange rate pass-through estimation is either a time series or a panel. The frequency of the data used ranges in the overviewed literature from monthly to annual. The studies using monthly data often use proxies for some of the variables. There seems to be a tradeoff between the use of precise variables and availability of more frequent data and thus more observations. In terms of the type of prices used currently in the literature, import prices and domestic consumer prices are the ones most commonly used. Unlike in the 1980s and the 1990s, the exchange rate pass-through into export prices does not currently seem to be of an interest of researchers. None of the 19 reviewed studies estimated this type of pass-through. This is a significant change from the 10 out of 43 studies solely on export prices and 13 out of 43 on both export and import prices that Menon (2005)

[1] Regional organization composed of former Soviet Republics

reports. Similarly to the past, some researchers (Mumtaz, Oomen and Wang (2006); Kozluk, Banerjee and de Bandt (2008); Dabusinskas (2003)) are not interested only in the overall level of the exchange rate pass-through in a country, but investigate deeper by estimating the values for each industry. The disaggregated approach is claimed to be more accurate, since it provides more detailed information, but for the sake of a comparison of the exchange rate pass-through across the region, the aggregate approach is appropriate and sufficient.

2.2.3 Methodology

The ordinary least squares are still used as an estimation method for exchange rate pass-through; however, the frequency of application is now significantly lower. Mumtaz, Oomen and Wang (2006) and Campa and Goldberg (2002) used ordinary least squares as their sole method of estimation. Others (e.g. Darvas (2001)) used it only as a one of the estimation methods, allowing for a comparison across different methods. Vector autoregression is a method of choice for many researchers dealing with the exchange rate pass-through. These include Darvas (2001), Bitāns (2004), Korhonen and Wachtel (2005), Ca'Zorzi, Hahn and Sánchez (2007), Vonnák (2010) and Beirne and Bijsterbosch (2011). Lately, methods of panel cointegration are becoming more popular in estimating the exchange rate pass-through. Holmes (2006) used the dynamic ordinary least squares (DOLS) on a sample of 12 members of the European Union; Holmes (2008) used Fully Modified Ordinary Least Squares on 19 African countries. Cheikh (2011) used both Dynamic and Fully Modified Ordinary Least Squares to estimate the pass-through for a number of OECD countries, including some countries relevant for this work – namely the Czech Republic, the Slovak Republic and Poland. In the reviewed literature, only Barhoumi (2005) used the pooled mean group (PMG) estimator and the mean group (MG) estimator – methods that are used to estimate the exchange rate pass-through for Central and Eastern European countries in this thesis. He, however, did not include the CEE region in his study.

2.2.4 Findings

Since this thesis is focusing on the countries of Central and Eastern Europe, the following paragraphs compare the exchange rate pass-through estimates for the countries from this region. The findings are compared across countries, time and methods used.

2.2.4.1 Exchange rate pass-through and its dynamics

As illustrated by Table 2: Findings for the CEE countries, incomplete exchange rate pass-through is a very common phenomenon in the Central and Eastern Europe. Out of the 198 estimates available in the reviewed literature, only nine are close to a complete pass-through (with pass-through between 0.9 and 1.1[2]). These are the VAR estimates for Poland and Slovenia in Coricelli, Jazbec and Masten (2006), the estimate for Poland in Campa and Goldberg (2002), the estimate for the Czech Republic and Poland in Cheikh (2011), the VAR estimates for Estonia and Latvia in Beirne and Bijsterbosch (2011), the eight quarter consumer price estimate for Hungary in Ca'Zorzi, Hahn and Sánchez (2007) and 24-month first period estimate for Lithuania in Bitāns (2004). Large pass-through (between 0.8 and 0.9) can be observed in four other estimates. There are further four estimates larger than 1.1. Three of these occur in Ca'Zorzi, Hahn and Sánchez (2007), which can point to an imperfect specification of the model.

There are several exchange rate pass-through estimates approaching zero (between 0 and 0.1). These are the both single equation estimates for Hungary in Coricelli, Jazbec and Masten (2006), the estimate for the Slovak Republic in Cheikh (2011), the six month shock response estimates for Hungary, the Slovak Republic and Estonia in Beirne and Bijsterbosch (2011), both estimates for the Czech Republic, Poland, Hungary and the Slovak Republic in Korhonen and Wachtel (2005), the short run estimate for Poland in Darvas (2001) and the 3-month first period estimates for the Czech Republic and Latvia, the 3-month second period estimates for the Czech Republic and Romania and the 6-month second period estimate for the Czech Republic in Bitāns (2004). There are 24 more very low estimates (in the range 0.1-0.2). In the reviewed literature, there were two instances (Coricelli, Jazbec and Masten (2006) single equation short and long run estimates for Poland) where the exchange rate pass-through was negative. Klein (1990) explains that this can occur when the exchange rate depreciation is associated with a decline in a conditional expectation of income and unobserved domestic price level that outweigh the effect of higher value of foreign costs in the domestic currency. This, however, does not necessarily have to be the explanation for the negative results in this case.

Development of the exchange rate pass-through over time can be observed in Bitāns (2004). He calculated the exchange rate pass through for two periods for all covered countries with the exception of Croatia, where only data for the more recent period were available. Vast majority of the estimates decreased from the first period to the second. The only exceptions were some of the 3-month estimates. Besides the results

[2] In this category the only estimate larger than one has a value of 1.01 (VAR estimate for Slovenia in Coricelli, Jazbec and Masten (2006)). Despite the problematic nature of estimates larger than one, such a small plus variation from one can be considered as complete pass-through.

published by Bitāns (2004), there does not seem to be a clear pattern of development of the exchange rate pass-through over time. The reason for this might be that the examined periods of a majority of the studies overlap each other. Another reason can be the general variability of results that occurs in this sample.

2.2.4.2 Exchange rate pass-through across countries

There are major differences in the extent of the exchange rate pass-through among countries. The country-specific averages across all studies and estimates range from 0.26 for the Slovak Republic to 0.64 for Russia. In addition to the Slovak Republic, Croatia, the Czech Republic, Lithuania, Macedonia, Slovenia, Hungary, Latvia and Ukraine fall below the average of the range. Poland, Romania, Estonia, Moldova and Bulgaria belong alongside Russia to the upper half.[3] Despite the rather large difference between the lowest and highest average estimate, nine out of the fifteen countries fall in the range 0.4 - 0.5. The differences in estimates for different countries within one study and type of estimate are even larger than the differences in the average values. For example, Cheikh (2011) quotes the pass-through rate of 0.07 for Slovakia while the Czech Republic reaches 0.95 and Poland 0.98. In Korhonen and Wachtel (2005), the estimates range from 0.03 for the Czech Republic both in 12 and 14 months to 0.88 and 1.12 for Romania in 12 and 14 months respectively.

2.2.4.3 Exchange rate pass-through across studies for one country

It is clear from Table 2: Findings for the CEE countries that there are not only differences between countries but also between estimates for one country. These differences can occur for several reasons; the two main ones being a different model specification and a use of a different methodology.

The main difference in the model specification is the type of prices into which is the exchange rate pass-through measured. The two main types of prices used in the reviewed literature are import and consumer prices[4]. Estimates for both import and consumer prices are available only for five countries (the Czech Republic, Poland, Hungary, the Slovak Republic and Estonia). From the estimates for the first three countries, it is clearly visible that the exchange rate pass-through into consumer prices is generally lower than the one into import prices. This is in line with the

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[3] The countries appear in the order of their average exchange rate pass-through estimate from the lowest to the highest.

[4] Producer prices appeared in only two studies focusing on the Central and Eastern Europe, one of them covering only Estonia.

general expectations. The difference between the pass-through into import and consumer prices is believed to exist for several reasons (Bacchetta and van Wincoop, 2002):

i) The imported goods incorporate some value added in the distribution sector of the domestic market and thus the consumer prices are less sensitive to changes in the exchange rate. Alternatively, the distribution costs play a role in lowering the pass-through into consumer prices.
ii) The final goods sold to customers in the domestic market are a mix of imported intermediate goods and domestic goods. Therefore, the resulting price is set in two stages – by the producers of intermediate goods and the final goods producers.
iii) If domestic firms face a strong competition from domestic producers of final goods, they tend to set prices in domestic currency. On the other hand, foreign firms prefer pricing in the currency of the exporter. This leads to the exchange rate pass-through being complete for import prices and zero for consumer prices.

Based on the estimates provided in the reviewed studies, the Slovak Republic and Estonia seem to be an exception to this rule. The estimates of the exchange rate pass-through into consumer prices are for all countries mostly low, many close to zero and only very few approaching the complete pass-through.

Different methods can also yield significantly different results. This is apparent especially in the studies that estimated the pass-through by more methods, since the specification of the model is most likely the same or comparable and thus the differences cannot be attributed to it. Good examples of the possible differences in estimates for one country, within one study, across more methods are the estimates for Hungary and Slovenia in Coricelli, Jazbec and Masten (2006). For Hungary, the long run single equation estimate equals to 0.06, while the long run cointegrated vector autoregression estimate equals to 0.8. For Slovenia, the long run single equation estimate equals to 0.19, while the long run cointegrated vector autoregression estimate equals to 1.01. The variance of results based on the method used can be further illustrated by the 0.47 spread between the fully modified ordinary least squares estimate (Cheikh, 2011) and the vector autoregression estimate (Ca'Zorzi, Hahn and Sánchez, 2007) for the Czech Republic.

The data also suggest that the exchange rate pass-through is larger in the long term than in the short term. For all Central and Eastern European estimates in Campa and Goldberg (2002), Beirne and Bijsterbosch (2011), Korhonen and Wachtel (2005) and

Darvas (2001) and all estimates for consumer prices in Ca'Zorzi, Hahn and Sánchez (2007), it holds that the estimates of the exchange rate pass-through into the same prices using the same method are equal or larger in the long term than in the short term. In addition to the four studies mentioned above, in three more studies (Coricelli, Jazbec and Masten (2006); Bitāns (2004); and the import-price part of Ca'Zorzi, Hahn and Sánchez (2007)) only one country does not satisfy this condition.

The previous paragraphs lead to a conclusion, that the specification of the pass-through equation and the selection of the estimation method are crucial and that results are not easily comparable across studies.

Table 2: Findings for the CEE countries

Study	Coricelli, Jazbec and Masten (2006)			Campa and Goldberg (2002)		Cheikh (2011)	Beirne and Bijsterbosch (2011)				
Time period	1993 - 2002			1975 - 1999		1994 - 2010	1995 - 2008				
Method of estimation	Single equation		Cointe-grated VAR	OLS		FMOLS	Shock response				Cointe-grated VAR
Prices	Consumer			Import		Import	Consumer				
Short run / Long run	SR	LR	LR	SR	LR	LR	6 months	12 months	24 months	48 months	LR
Bulgaria							0.20	0.21	0.32	0.36	0.70
Croatia											
Czech Republic	0.20	0.22	0.46	0.38	0.61	0.95	0.25	0.38	0.41	0.43	0.50
Estonia							0.06	0.16	0.57	0.60	0.93
Hungary	0.05	0.06	0.97	0.58	0.85		0.09	0.24	0.37	0.40	0.63
Latvia							0.36	0.44	0.51	0.62	0.97
Lithuania							0.15	0.21	0.34	0.46	0.44
Macedonia											
Moldova											
Poland	-0.02	-0.02	0.80	0.50	0.99	0.98	0.27	0.36	0.40	0.40	0.47
Romania							0.14	0.18	0.23	0.34	0.44
Russia											
Slovak Republic						0.07	0.05	0.18	0.39	0.39	0.37
Slovenia	0.22	0.19	1.01								
Ukraine											

Study	Korhonen and Wachtel (2005)		Darvas (2001)		Ca'Zorzi, Hahn and Sánchez (2007)				Dabusinskas (2003)	
Time period	1999 - 2004		1993 - 2000		CZ: 1993 - 2004; PL: 1991 - 2003; HU: 1988 -2003				1995 -2003	
Method of estimation	VAR				VAR				SUR	OLS
Prices	Consumer		Consumer		Import		Consumer		Import	Consumer
Short run / Long run	12 months	24 months	SR	LR	4 quarters	8 quarters	4 quarters	8 quarters	SR & LR	
Bulgaria										
Croatia										
Czech Republic	0.03	0.03	0.10	0.15	0.72	0.48	0.61	0.77		
Estonia									0.30	no significant ERPT
Hungary	0.06	0.06	0.10	0.40	1.26	1.77	0.48	0.91		
Latvia										
Lithuania										
Macedonia										
Moldova	0.49	0.49								
Poland	0.09	0.09	0.00	0.20	0.86	1.30	0.31	0.56		
Romania	0.88	1.12								
Russia	0.63	0.64								
Slovak Republic	0.05	0.05								
Slovenia	0.16	0.18	0.20	0.40						
Ukraine	0.42	0.42								

Study	Billmeier and Bonato (2002)	Bitāns (2004)							
Time period	1994 - 2001	Varying from 1993-1994 to 1997-2001				Varying from 1997-2001 to 2003			
Method of estimation	Cointegrated VAR	Recursive VAR				Recursive VAR			
Prices	Consumer	Consumer				Consumer			
Short run / Long run	LR	3 months	6 months	12 months	24 months	3 months	6 months	12 months	24 months
Bulgaria		0.80	0.98	1.11	1.19	0.44	0.47	0.47	0.48
Croatia	0.30					0.18	0.23	0.35	0.36
Czech Republic		0.02	0.14	0.21	0.30	0.05	0.09	0.13	0.14
Estonia		0.59	0.67	0.72	0.73	0.21	0.35	0.34	0.34
Hungary		0.14	0.24	0.27	0.28	0.15	0.20	0.21	0.21
Latvia		0.07	0.47	0.51	0.58	0.13	0.25	0.26	0.26
Lithuania		0.22	0.44	0.73	1.00	0.22	0.24	0.28	0.39
Macedonia		0.26	0.44	0.66	0.73	0.21	0.29	0.30	0.30
Moldova									
Poland		0.29	0.47	0.51	0.51	0.28	0.31	0.34	0.34
Romania		0.58	0.70	0.7	0.71	0.09	0.14	0.23	0.24
Russia									
Slovak Republic		0.35	0.47	0.47	0.47	0.21	0.21	0.23	0.23
Slovenia		0.55	0.71	0.71	0.74	0.17	0.23	0.32	0.33
Ukraine									

Source: author's compilation of data from the listed studies

3 Relationship between prices and exchange rates

3.1 Exchange rate pass-through

The relationship between prices and exchange rates can be described as a correlation between these two variables. The correlation takes a form of

$$\beta = \frac{cov\,(p, e)}{var(e)} \qquad (1)$$

where p is a logarithm of the price denominated in the domestic currency and e is a logarithm of the nominal exchange rate (in units of the currency of the importer per one unit of currency of the exporter). This relationship is, however, strictly statistical. It lacks an economic interpretation and ignores endogeneity of the variables (Campa, Goldberg and Gonzales-Mínguez, 2005).

Alternatively, the relationship between prices and exchange rates can be expressed as the exchange rate pass-through. The exchange rate pass-through expresses weather changes in exchange rates have an impact on prices of traded goods or on producer markups and to what extent. In earlier studies, the main focus lied on exchange rate pass-through into import and export prices. As Darvas (2001) points out, this is natural since it allows to study pricing practices of firms and also because changes in the exchange rate usually impact first the import and export prices and only after that the consumer prices.

The exchange rate pass-through is defined (Campa and Goldberg, 2002) as the percentage change in import prices denoted in local currency resulting from a one percent change in the exchange rate between the importing and exporting country. A basic equation for estimation of the exchange rate pass-through is

$$mp_t = \gamma e_t + \varepsilon_t \qquad (2)$$

where e is defined as above, mp is a logarithm of import price denominated in the currency of the importer and ε is the error term. The coefficient γ represents the exchange rate pass-through. If $\gamma = 1$, it is said that the exchange rate pass-through is complete; if $\gamma < 1$, it is said that the exchange rate pass-through is incomplete

(Goldberg and Knetter, 1996). This equation represents, however, again only a statistical relationship that does not have, as it is, a meaningful economic interpretation.

To better understand the dynamics of the exchange rate pass-through into import prices, it is suggested (Campa, Goldberg and Gonzales-Mínguez, 2005) to take into consideration the micro-foundations of exporter's pricing behavior. The import prices for a certain country (MP_t) can be defined as the export prices of its trading partners (XP_t) transformed by the exchange rate in units of the currency of the importer per one unit of the currency of the exporter (E_t):

$$MP_t = E_t \cdot XP_t \tag{3}$$

The logarithmic form of this transformation is

$$mp_t = e_t + xp_t \tag{4}$$

where the lower case denotes a logarithm of the variable.

The export prices are further composed of the exporter's markup ($XMKUP_t$) and marginal costs (XMC_t):

$$XP_t = XMC_t \cdot XMKUP_t \tag{5}$$

rewritten in a logarithmic form as

$$xp_t = xmc_t + xmkup_t \tag{6}$$

where the lower case again denotes a logarithm of the variable.

Thus the relationship between import prices and exchange rate is given by equation:

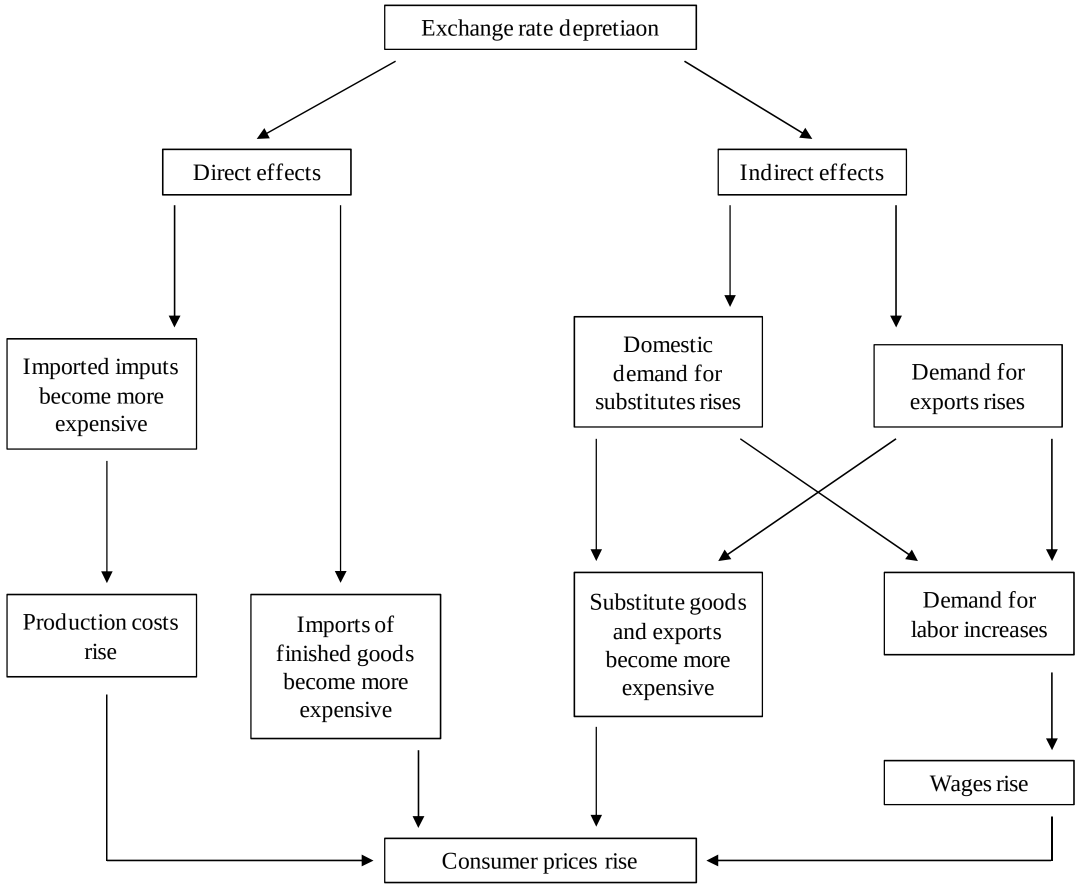
$$mp_t = e_t + xmc_t + xmkup_t \tag{7}$$

This equation is a base for the estimated equations in a large portion of the exchange rate pass-through literature. Generally, the exporter's markup and the marginal costs are further broken down or substituted by a proxy in order to be able to find suitable data. For example, Hooper and Mann (1989) define the markup as a variable that is influenced both by domestic competitive pressures and by demand pressures in all markets combined. They further assume that the domestic competitive pressure is represented by the gap between the competitor's prices in the domestic market and the production costs in the foreign market. The demand pressures are measured in their study by capacity utilization. The exporter's marginal costs are broken down for example in Campa, Goldberg and Gonzales-Mínguez (2005). They assume the exporter's marginal costs to be a positive function of three variables: demand in the importing country, marginal costs of production (represented by wages) and commodity prices in foreign currency. The variety of variables used in the exchange rate pass-through literature to explain the exporter's markups and marginal costs is quite large. The data for some of these variables are not easily accessible. They often either have a lower frequency than required or they are not available at all. This holds especially for developing countries, which are not members of large economic organizations or unions (e.g., OECD, the European Union) that publish a wide range of statistics about their members. Some authors (e.g., Hooper and Mann, 1989) solve this problem by constructing proxies, substituting missing data by data from neighboring countries, etc. This does not necessarily mean that the results of such estimations are significantly different from what they would be if the precise data were available; however, this work is dealing from a large part with countries that are not members of the OECD or the European Union and thus the extent of substitutions would need to be quite large. Most importantly, import price indices are available only for a few CEE countries, each of the price indices being defined in a slightly different way. There are import unit value indices available for several countries but these are considered an unsuitable substitution to import price indices since they misrepresent the price changes (Silver, 2007). The shortage of suitable import price data and the fact that consumer prices are considered to be the variable that is more interesting to policymakers (Beirne and Bijsterbosch, 2011), led us to opt for a model of exchange rate pass-through into consumer prices.

3.2 Exchange rate pass-through into consumer prices

There are two channels through which changes in the exchange rate are passed on to consumer prices – a direct channel and an indirect channel. Currency appreciation or depreciation reflects directly in the import prices. If producers change their prices proportionally with the shift in the prices of imported goods, the changes of the prices of imported goods are then passed on to domestic producer and consumer prices. This is called the direct channel. The indirect channel is not dependent on passing the effect via prices but rather through changes in the composition of demand and in the level of wages. If the exchange rate increases, the imports become more expensive for domestic buyers and the domestic products become relatively cheaper for foreign buyers. This leads to a higher demand for substitute domestic products by domestic buyers and a higher demand for exports by foreign buyers. This increase in the aggregate demand pushes the domestic price level up. The increasing demand for domestic products eventually leads to a higher production and therefore a higher demand for labor, which can be followed by an increase in wages. This increase puts further upward pressure on the domestic prices (Schröder and Hüfner, 2002; Laflèche, 1997). The mechanism of direct and indirect exchange rate pass-through can be seen in Figure 1.

Figure 1: Exchange rate pass-through into consumer prices



Source: Laflèche (1997)

To assess the extent of the exchange rate pass-through into consumer prices, McCarthy (1999) suggests a model of pricing along a distribution chain. This model includes inflation shocks at previous stages of the distribution chain. For consumer prices, this means including shocks in import and producer prices. The model further includes domestic supply and demand shocks on inflation. The composition of the variables thus allows the model to consider both direct and indirect effects of the exchange rate changes.

The model used in this work is based on previous research and is constructed in the following way. It is partially based on McCarthy's model of pricing along the distribution chain (McCarthy, 1999), but due to the previously mentioned reasons, import prices are not included in the model. Other researchers faced the setback of missing import prices for the countries of the Central and Eastern Europe as well. Bitāns (2004) was also not able to include the import prices in his analysis and thus implicitly assumed that there is a complete exchange rate pass-through into import prices. He based the assumption on the observation that in this region there is only a small degree of local currency pricing and thus a majority of foreign trade is invoiced in foreign currencies. Including the producer prices partially preserves the distribution chain character of the model, as it not only allows for direct effects of changes in the exchange rate on consumer prices but for indirect effects as well. Oil prices are included as a proxy for supply shocks and gross domestic product (GDP) as a proxy for demand shocks. Short-term interest rates serve as an indicator of central bank policy.

4 Estimation of non-stationary heterogeneous panels

In this work, two methods suitable for the estimation of models that exhibit heterogeneity, non-stationarity and cointegration relationships among variables are used. First method is the mean group estimator of Pesaran and Smith (1995) and the second is the pooled mean group estimator of Pesaran, Shin and Smith (1999). Traditional fixed effects estimator is used as a reference.

The fixed effects estimator – similarly to other traditional pooled estimators such as random effects estimator – constrains the slope coefficients and the error variances to be the same across groups. The only coefficients allowed to vary across groups are the intercepts. If there is, however, heterogeneity in the slope coefficients the estimation using fixed effects can yield inconsistent and misleading results. The mean group estimator is the other extreme. The equation for each group is fitted separately and the mean of the coefficient estimates is then considered. Since the model is estimated for each group separately, the number of time periods (T) needs to be large enough to allow for it. The pooled mean group estimator lies between the fixed effects and the mean group estimator; it combines pooling and averaging. The estimator constrains only the long-run coefficients to be the same across groups while it allows the intercepts, short-run coefficients and error variance to vary.

The description of the models in this chapter is based on Asteriou (2006), Blackburne and Frank (2007), Pesaran and Smith (1995) and Pesaran, Shin and Smith (1999).

4.1 Fixed effects

Fixed effects (FE) are one of the linear panel data models. In this model, the slope coefficients and error variances are fixed to be the same across groups, while the intercept is allowed to vary from group to group. To enable such variance in the constant, it includes dummy variables - one for each group. The model can be written as (Asteriou, 2006):

$$Y_{it} = a_i + \beta_1 X_{1it} + \beta_2 X_{2it} + \cdots + \beta_k X_{kit} + u_{it} \tag{8}$$

or in a matrix notation

$$\boldsymbol{Y} = \boldsymbol{D\alpha} + \boldsymbol{X\beta}' + \boldsymbol{u} \tag{9}$$

where $\boldsymbol{Y} = \begin{pmatrix} Y_1 \\ Y_2 \\ \vdots \\ Y_N \end{pmatrix}_{NT\times 1}$, $\boldsymbol{\alpha} = \begin{pmatrix} \alpha_1 \\ \alpha_2 \\ \vdots \\ \alpha_N \end{pmatrix}_{N\times 1}$, $\boldsymbol{\beta} = \begin{pmatrix} \beta_1 \\ \beta_2 \\ \vdots \\ \beta_N \end{pmatrix}_{k\times 1}$

$$\boldsymbol{D} = \begin{pmatrix} i_T & 0 & \cdots & 0 \\ 0 & i_T & & 0 \\ \vdots & \vdots & & \vdots \\ 0 & 0 & \cdots & i_T \end{pmatrix}_{NT\times N}, \boldsymbol{X} = \begin{pmatrix} x_{11} & x_{12} & \cdots & x_{1K} \\ x_{21} & x_{22} & & x_{2k} \\ \vdots & \vdots & & \vdots \\ x_{N1} & x_{N1} & \cdots & x_{Nk} \end{pmatrix}_{NT\times k}$$

4.2 Mean group and pooled mean group estimator

The mean group and the pooled mean group estimators were developed by Pesaran and Smith (1995) and Pesaran, Shin and Smith (1999) respectively. For clarity, the notation in this chapter is mainly based on Blackburne and Frank (2007), who are the authors of the Stata command xtpmg, which is later used to estimate the model.

Both the mean group and the pooled mean group estimators are derived from the autoregressive distributive lag (ARDL) dynamic panel. Assume a model specified in the following way:

$$y_{i,t} = \sum_{j=1}^{p} \lambda_{ij} y_{i,t-j} + \sum_{j=0}^{q} \delta'_{ij} X_{i,t-j} + \mu_i + \epsilon_{it} \tag{10}$$

where $i = 1,2,\ldots,N$ is the number of groups; $t = 1,2,\ldots,T$ is the number of periods; $X_{i,t}$ is a $k \times 1$ vector of explanatory variables; $\delta_{i,t}$ are the $k \times 1$ coefficient vectors, λ_{ij}, the coefficients of the lagged dependent variables, are scalars; and μ_i is the group-specific effect. As stated before, T needs to be large enough so that the equation for each group can be estimated separately. It is possible for p and q to vary across groups and for the panel to be unbalanced.

If variables are cointegrated, they react to any deviation from a long run equilibrium. Therefore, in the presence of cointegrated variables, an error correction model is appropriate. The reparametrization of the equation (10) into the error correction equation has the following form:

$$\Delta y_{it} = \phi_i\big(y_{i,t-1} - \theta_i' X_{it}\big) + \sum_{j=1}^{p-1} \lambda_{ij}^* \Delta y_{i,t-1} + \sum_{j=0}^{q-1} \delta_{ij}'^* \Delta X_{i,t-j} + \mu_i + \epsilon_{it} \quad (11)$$

where $\phi_i = -\big(1 - \sum_{j=1}^{p} \lambda_{ij}\big)$

$$\theta_i = \frac{\sum_{j=0}^{q} \delta_{ij}}{(1 - \sum_k \lambda_{ik})}$$

$$\lambda_{ij}^* = -\sum_{m=j+1}^{p} \lambda_{im} \quad j = 1,2, \dots, p-1$$

$$\delta_{ij}^* = -\sum_{m=j+1}^{q} \delta_{im} \quad j = 1,2, \dots, q-1$$

The parameter ϕ_i is the equilibrium or error-correction parameter; $\phi_i = 0$ suggests no evidence for a long-run relationship. The parameter θ_i is the long-run parameter.

The difference between the mean group estimator and the pooled mean group estimator lies in the restriction of the long-run parameters in the case of the pooled mean group estimator. The model for pooled mean group estimator therefore restricts θ to be the same across groups. The equation then takes the following form:

$$\Delta y_{it} = \phi_i\big(y_{i,t-1} - \theta' X_{it}\big) + \sum_{j=1}^{p-1} \lambda_{ij}^* \Delta y_{i,t-1} + \sum_{j=0}^{q-1} \delta_{ij}'^* \Delta X_{i,t-j} + \mu_{\mathrm{i}} + \epsilon_{it} \quad (12)$$

To estimate the equations (11) and (12), Pesaran, Shin and Smith (1999) suggest the use of a maximum likelihood estimator. They develop a maximum likelihood estimator, where the likelihood is expressed as the product of each cross-section’s likelihood and then the log yields are taken. The form of the maximum likelihood estimator is as follows:

$$l_T(\theta', \varphi', \sigma') = -\frac{T}{2}\sum_{i=1}^{N} \ln(2\pi\sigma_i^2) - \frac{1}{2}\sum_{i=1}^{N} \frac{1}{\sigma_i^2}\{\Delta y_i - \phi_i \xi_i(\theta)\}' H_i \{\Delta y_i - \phi_i \xi_i(\theta)\} \qquad (13)$$

for i=1,2,…,N

where $\xi_i(\theta) = y_{i,t-1} - X_i\theta_i$

$H_i = I_T - W_i(W_i' W_i)W_i$

$W_i = \left(\Delta y_{i,t-1}, \dots, \Delta y_{i,t-p+1}, \Delta X_i, \Delta X_{i,t-1}, \dots, \Delta X_{i,t-q+1}\right)$

I_T is an identity matrix of order T

5 Data sources and empirical estimation

5.1 Data description

The sample covers 12 Central and Eastern European countries, namely Bulgaria, Croatia, the Czech Republic, Estonia, Hungary, Latvia, Lithuania, Macedonia, Poland, Romania, Russia and Slovenia. The data were collected on quarterly basis and for all countries range from the first quarter of 2003 to the first quarter of 2013.

The exchange rates used are the nominal effective exchange rates. The primary source of the nominal effective exchange rates is the International Financial Statistics database of International Monetary Fund (IMF IFS). This database, however, does not contain the nominal effective exchange rates for all of the countries of interest. The data are therefore supplemented by nominal effective exchange rates provided by the Bank for International Settlements. The sample includes only 12 Central and Eastern European countries, because even after merging the data on nominal effective exchange rates from these two sources, data for only 14 countries were available, two of which had to be excluded for other reasons. Consumer price indices as well as a majority of the producer price indices come from the IMF IFS database. For Croatia, a wholesale price index (from IMF IFS) had to be used instead of the producer price index.[5] The producer prices for Russia were obtained from the Federal State Statistics Service of the Russian Federation. The basic source of the data for the gross domestic products is Eurostat. It, however, does not include GDP for Russia, so the data are supplemented from the IMF IFS database. The GDPs are denominated in national currencies. All of the series for euro area countries come from Eurostat and are "euro fixed". In the surveyed literature, the short-term interest rates are commonly the money market rates or the treasury bill rates (Hüfner and Schröder, 2002; Vonnák, 2010; Ca'Zorzi, Hahn and Sánchez, 2007). Neither of these rates was available for all of the analyzed countries and their combination did not cover the sample either. Ca'Zorzi, Hahn and Sánchez (2007) overcame the problem of unavailability of

[5] In the IMF IFS database, the producer price index and wholesale price index both come from line 63. In construction of the line 63, preference is given to the producer price index, if it is available. In the opposite case, the wholesale price index is included.

neither money market nor treasury bill rates for some countries by using the bank deposit rates. The set of deposit rates in the IMF IFS database covers the highest number of countries and it is therefore used as the main measure of short-term interest rates in this work. For three of the countries with unavailable deposit rates (namely Lithuania, Poland and Slovenia), money market rates are used. The oil price used for our analysis is the UK Brent price from the IMF IFS database.

The sample of countries includes countries with both fixed and floating exchange rate regimes. The countries with fixed exchange rates are not excluded from the sample, because their elimination would cause the sample to be too small to be estimated using the mean group and pooled mean group methods. The mean group and pooled mean group models are suitable for panels with large number of both time periods (T) and groups (N). Without exclusion of the countries with fixed exchange rate, the panel used in this work has 41 time periods and 12 cross-sectional groups. Such number of cross-sections is borderline acceptable and thus elimination of the fixed-exchange-rate countries would impair the quality of the results. Furthermore, Pesaran, Shin and Smith (1999) note, that it is desirable for the number of time periods and the number of groups to be of the same order of magnitude. The non-desirable widening of the difference in magnitudes is another reason to preserve the sample of countries as it is.

5.2 Panel data tests

Prior to the estimation of the model, data are tested for unit-roots and cointegration. In the following subsections, panel unit root and cointegration tests and their results are described.

5.2.1 Panel unit root tests

When dealing with longer time series, variables are likely to be non-stationary. Our dataset covers 41 time periods and we thus employ tests to check for unit roots. We employ two different tests – the test suggested by Im, Pesaran and Shin (2003) and the test suggested by Hadri (2000).

5.2.1.1 Im, Pesaran and Shin test

The test proposed by Im, Pesaran and Shin (2003) is commonly used to test for non-stationarity in heterogeneous panels. In this work, the test is described based on Baltagi (2005) and Asteriou (2006). The Im, Pesaran and Shin test (denoted the IPS test) is based on a model of the following form:

$$\Delta Y_{i,t} = a_i + \rho_i Y_{i,t-1} + \sum_{k=1}^{n} \phi_k \Delta Y_{i,t-k} + \delta_i t + \theta_t + u_{it} \qquad (14)$$

The null and alternative hypotheses are

$$H_0: \rho_i = 0 \text{ for all } i$$

$$H_A: \rho < 0 \text{ for at least one } i$$

It therefore tests the null hypothesis of non-stationarity of all series against the alternative hypothesis that at least one of the series is stationary.

The $\bar{t}$ statistic in the IPS test is the average of the individual Augmented Dickey-Fuller (ADF) test statistics. It is specified as

$$\bar{t} = \frac{1}{N} \sum_{i=1}^{N} t_{\rho_i} \qquad (15)$$

where t_{ρ_i} is the individual ADF t-statistic for testing that $\rho_i = 0$ for all *i*.

Im, Pesaran and Shin (2003) constructed a test statistic that has an asymptotic N(0,1) distribution as T → ∞ followed by N → ∞ sequentially. The IPS statistic is defined as[6]

$$t_{IPS} = \frac{\sqrt{N}\left(\bar{t} - \frac{1}{N}\sum_{i=1}^{N} E[t_{iT}|\rho_i = 0]\right)}{\sqrt{\frac{1}{N}\sum_{i=1}^{N} \text{var}[t_{iT}|\rho_i = 0]}} \qquad (16)$$

5.2.1.2 Hadri test

The residual-based Lagrange multiplier (LM) stationarity test derived by Hadri (2000) is believed to have an important advantage over the Im, Pesaran and Shin test. It assumes stationarity under the null hypothesis and therefore avoids the lack of power of the unit-root based tests (Barhoumi, 2005). Hadri (2000) considers two models:

[6] For more information see Baltagi (2005).

$$y_{it} = r_{it} + \epsilon_{it} \tag{17}$$

and

$$y_{it} = r_{it} + \beta_i t + \epsilon_{it} \tag{18}$$

where $t = 1,\ldots,T$

$i = 1,\ldots,N$

r_{it} is a random walk: $r_{it} = r_{it-1} + u_{it}$

ϵ_{it} is a mutually independent normal, IID across i and over t, $E[\epsilon_{it}] = 0$, $E[\epsilon_{it}^2] = \sigma_\epsilon^2 > 0$

u_{it} is a mutually independent normal, IID across i and over t, $E[u_{it}] = 0$, $E[u_{it}^2] = \sigma_u^2 \geq 0$

Using back substitution, the model can be rewritten in the following way:

$$y_{it} = r_{i0} + \beta_i t + \sum_{t=1}^{T} u_{it} + \epsilon_{it} = r_{i0} + \beta_i t + e_{it} \tag{19}$$

where $e_{it} = \sum_{t=1}^{T} u_{it} + \epsilon_{it}$

The null hypothesis is a hypothesis of stationarity. That is $\sigma_u^2 = 0$ and then $e_{it} = \epsilon_{it}$.

The LM statistic has the following form:

$$LM = \frac{\frac{1}{N}\sum_{i=1}^{N}\frac{1}{T^2}\sum_{t=1}^{T} S_{it}^2}{\hat{\sigma}_\epsilon^2} \tag{20}$$

where S_{it} is the partial sum of residuals $S_{it} = \sum_{j=1}^{t} \hat{\epsilon}_{ij}$

$\hat{\sigma}_\epsilon^2$ is a consistent estimator of σ_ϵ^2 under H_0

The LM statistic can be adjusted to allow for heteroscedasticity in disturbance terms across i. The statistic is then given as

$$\widehat{LM} = \frac{1}{N}\sum_{i=1}^{N}\left(\frac{\frac{1}{T^2}\sum_{t=1}^{T} S_{it}^2}{\hat{\sigma}_{\epsilon,i}^2}\right) \quad (21)$$

The test statistic is described by the following formula:

$$Z = \frac{\sqrt{N}\left(\widehat{LM} - \xi\right)}{\zeta} \Rightarrow N(0,1) \quad (22)$$

where $\xi = \frac{1}{6}$ and $\zeta^2 = \frac{1}{45}$ if testing for the null of level stationarity and

$\xi = \frac{1}{15}$ and $\zeta^2 = \frac{11}{6300}$ if testing for the null of trend stationarity

5.2.1.3 Unit root test results

Two tests were conducted to assess the stationarity of the variables. In order to deal with a problem of cross-sectional dependence, the tests were applied on demeaned data (as suggested by Barhoumi (2005)). The only variable that was not demeaned is the variable *oil*, as the data are the same across cross-sections and removing the mean would thus result in removing all information from the variable.

The results of the IPS panel unit root tests are displayed in Table 3, the corresponding critical values in Table 4.

Table 3: IPS panel unit root test results

Variables	Statistics for levels	Statistics for first difference
cpi	-1.2082	-5.5059
neer	-1.9617	-5.2032
ppi	-1.4477	-4.7331
ir	-1.5355	-5.0265
gdp	-1.6729	-7.6518
oil	-1.4030	-4.7767

Source: author's calculation

Table 4: Critical values for the IPS panel unit root test

Significance level	1%	5%	10%
Critical value	-2.04	-1.90	-1.81

Source: StataCorp (2011b).

For variables in levels, the null hypothesis of non-stationarity cannot be rejected at the 5% (or even 10%) level of significance for any of the variables, except for the nominal effective exchange rate (*neer*), for which it is possible to reject the hypothesis at the 5% level. For the first differences, it is possible to reject the null hypothesis of non-stationarity for all variables at the 1% level of significance. The results of the Hadri test are shown in Table 5. In first differences, it is not possible to reject the null hypothesis at the 5% (or even 10%) level of significance for any of the variables except for the consumer price index (*cpi*). Stationarity of *cpi* can be rejected at the 5% level of significance.

Table 5: Hadri panel unit root test results

Variables	Levels		First difference	
	Statistic	P-value	Statistic	P-value
cpi	85.3988	0.0000	2.2404	0.0125
neer	51.4684	0.0000	-1.6569	0.9512
ppi	77.8169	0.0000	-1.9427	0.9740
ir	24.3865	0.0000	-0.1861	0.5738
gdp	83.2787	0.0000	-3.0793	0.9990
oil	62.3108	0.0000	-2.9940	0.9986

Source: author's calculation

For most variables, the results of both tests coincide in suggesting that the variables are non-stationary in levels but stationary in first differences. There are only two instances, where the results are not clear. First of them is the stationarity of the levels of *neer*. According to the IPS test, the non-stationarity of the variable can be rejected at the 5% level of significance. However, since it can be rejected at the 1% level and since the Hadri test strongly rejects the null hypothesis of stationarity, we regard the variable as non-stationary. Similarly, the null hypothesis of the Hadri test, that the first differences of *cpi* are stationary, can be rejected at the 5% level. It, however, cannot be rejected at the 1% level and the IPS test strongly rejects the null hypothesis of non-stationarity. We therefore consider *cpi* to be stationary in first differences.

Combining the results of both tests, we therefore come to a conclusion that all of the variables are integrated of order one.

5.2.2 Cointegration tests

Since the variables are non-stationary, a spurious regression could occur. To investigate this problem and to assess the suitability of the selected estimation methods, the data are tested for cointegration. The Pedroni (1997, 1999 and 2000) tests were selected as a testing method of choice, as they allow for significant heterogeneity across groups (Asteriou, 2006).

5.2.2.1 Pedroni cointegration tests

Pedroni (1997, 1999 and 2000) developed seven tests to test for cointegration. The description in this chapter is based on Asteriou (2006) and Barhoumi (2005).

Consider a model:

$$Y_{i,t} = a_i + \delta_t + \sum_{m=1}^{M} \beta_{mi} X_{mi,t} + e_{i,t} \quad (23)$$

Pedroni (1997, 1999 and 2000) proposed seven test statistics, all with the null hypothesis of no cointegration. They can be split into two groups. The first group includes four test statistics that are all based on a within panel estimator. The second group consists of three test statistics, based on pooling along the between dimension.

The first group consists of the following statistics (also known as panel statistics)

1. The panel ν statistic (a variance ration test)

$$T^2 N^{\frac{3}{2}} Z_{\hat{v}_{N,T}} = \frac{T^2 N^{\frac{3}{2}}}{\sum_{i=1}^{N} \sum_{t=1}^{T} \hat{L}_{11i}^{-2} \hat{e}_{i,t}^2} \quad (24)$$

2. The panel ρ statistic (a panel version of the Phillips-Perron statistic)

$$T\sqrt{N} Z_{\hat{\rho}_{N,T}} = \frac{T\sqrt{N}(\sum_{i=1}^{N} \sum_{t=1}^{T} \hat{L}_{11i}^{-2} \left(\hat{e}_{i,t-1}^2 \hat{e}_{i,t}^2 - \hat{\lambda}_i\right)}{\sum_{i=1}^{N} \sum_{t=1}^{T} \hat{L}_{11i}^{-2} \hat{e}_{i,t}^2} \quad (25)$$

3. The panel t statistic (non-parametric)

$$Z_{t_{N,T}} = \sqrt{\tilde{\sigma}_{N,T}^2 \sum_{i=1}^{N} \sum_{t=1}^{T} \hat{L}_{11i}^{-2} \hat{e}_{i,t-1}^2} \left(\sum_{i=1}^{N} \sum_{t=1}^{T} \hat{L}_{11i}^{-2} \left(\hat{e}_{i,t-1}^2 \hat{e}_{i,t}^2 - \hat{\lambda}_i\right) \right) \quad (26)$$

4. The panel ADF *t* statistic (parametric)

$$Z_{t_{N,T}} = \sqrt{\tilde{\sigma}_{NT}^{*2} \sum_{i=1}^{N} \sum_{t=1}^{T} \hat{L}_{11i}^{-2} \hat{e}_{i,t-1}^{*2}} \left(\sum_{i=1}^{N} \sum_{t=1}^{T} \hat{L}_{11i}^{-2} \left(\hat{e}_{i,t-1}^{*2} \hat{e}_{i,t}^{*2} - \hat{\lambda}_i\right) \right) \quad (27)$$

The following statistics belong to the second group and are called the group mean statistics:

5. The group ρ statistic (parametric)

$$T\sqrt{N}\tilde{Z}_{\tilde{\rho}_{N,T}} = T\sqrt{N}\frac{\sum_{t=1}^{T}\left(\hat{e}_{i,t-1}^{2}\Delta\hat{e}_{i,t-1}^{2} - \hat{\lambda}_{i}\right)}{\sum_{i=1}^{N}\left(\sum_{t=1}^{T}\hat{e}_{i,t-1}^{2}\right)} \quad (28)$$

6. The group t statistic (non-parametric)

$$\sqrt{N}\tilde{Z}_{t_{N,T-1}} = \sqrt{N}\sum_{i=1}^{N}\left(\sqrt{\tilde{\sigma}_{i}^{2}\sum_{t=1}^{T}\hat{e}_{i,t-1}^{2}}\right)\sum_{t=1}^{T}\left(\hat{e}_{i,t-1}^{2}\hat{e}_{i,t}^{2} - \hat{\lambda}_{i}\right) \quad (29)$$

7. The group t statistic (parametric)

$$\sqrt{N}\tilde{Z}_{t_{NT-1}}^{*} = \sqrt{N}\sum_{i=1}^{N}\left(\sqrt{\tilde{s}_{i}^{*2}\sum_{t=1}^{T}\hat{e}_{i,t-1}^{*2}}\right)\sum_{t=1}^{T}\left(\hat{e}_{i,t-1}^{*2}\Delta\hat{e}_{i,t}^{*2}\right) \quad (30)$$

5.2.2.2 Cointegration test results

The Pedroni tests are used to assess the cointegration of variables for four different model specifications. These are the specifications, for which the exchange rate pass-through is estimated in the following chapter. The results of the tests are given in Table 6.

Table 6: Pedroni cointegration test results

Model specification	*cpi, neer, ppi, ir, gdp, oil*		*cpi, neer, ppi, ir, gdp*	
	Statistic	P-value	Statistic	P-value
Panel ν statistic	3.0566	0.0011***	4.6660	0.0000***
Panel ρ statistic	-0.2388	0.4056	-1.6812	0.0464**
Panel PP statistic	-2.4864	0.0065***	-3.8001	0.0001***
Panel ADF statistic	-2.0243	0.0215**	-3.1032	0.0010***
Group ρ statistic	1.8881	0.9705	0.7797	0.7822
Group PP statistic	-0.4486	0.3268	-1.3211	0.0932*
Group ADF statistic	-0.7781	0.2183	-1.7457	0.0404**

Model specification	*cpi, neer, ppi, gdp*		*cpi, neer, ppi, ir*	
	Statistic	P-value	Statistic	P-value
Panel ν statistic	4.3657	0.0000***	4.0396	0.0000***
Panel ρ statistic	-1.4829	0.0691*	-1.7000	0.0446**
Panel PP statistic	-2.7078	0.0034***	-3.0880	0.0010***
Panel ADF statistic	-4.5660	0.0000***	-4.0481	0.0000***
Group ρ statistic	0.4344	0.6680	0.0499	0.5199
Group PP statistic	-1.0809	0.1399	-1.7733	0.0381**
Group ADF statistic	-3.3208	0.0004***	-3.2564	0.0006***

Source: author's calculation

The results of the tests are mixed for all four specifications. For the first specification with a full set of variables (*cpi, neer, ppi, ir, gdp, oil*), the null hypothesis of no cointegration is rejected at the 5% level of significance for three out of the seven test statistics. This number increases after dropping some of the variables. For the *cpi, neer, ppi, ir, gdp* specification, five statistics reject the null hypothesis at the 5% level of significance (six at the 10% level). In the case of the specification *cpi, neer, ppi, gdp,* four statistics reject the null hypothesis at the 5% (and even 1%) level, with one more rejecting the null at the 10% level of significance. Finally, for the specification *cpi, neer, ppi, ir* the null hypothesis is rejected at the 5% level in six cases.

The mixed results are likely given by the insufficient power of the tests. Due to the number of observations, the tests are not able to reject the null hypothesis of no cointegration. This applies especially to the specifications with a higher number of variables. Despite the unclear results, we conclude that cointegration is present in the data and proceed to the estimation using the mean group and pooled mean group estimators.

5.3 Empirical estimation

Since we concluded that the variables are I(1) and cointegrated, we proceed to the estimation of the models using the mean group and pooled mean group estimation methods. Estimations were carried out for 28 model specifications. There are four specification groups based on the selection of variables, each estimated with zero to six lags of the short-run variables. The specifications are for the mean group estimation given by the following modifications of the equation (11):

First specification group (*cpi neer ppi ir gdp oil*):

$$\begin{aligned}\Delta cpi_{it} = \phi_i\big(&cpi_{i,t-1} - \theta_{1i}neer_{it} - \theta_{2i}ppi_{it} - \theta_{3i}ir_{it} - \theta_{4i}gdp_{it} \\ &- \theta_{5i}oil_{it}) + \sum_{j=1}^{p-1}\lambda_{ij}^{*}\Delta cpi_{i,t-1} + \sum_{j=0}^{q-1}\delta_{1\mathrm{ij}}\Delta neer_{i,t-j} \\ &+ \sum_{j=0}^{q-1}\delta_{2\mathrm{ij}}\Delta ppi_{i,t-j} + \sum_{j=0}^{q-1}\delta_{3\mathrm{ij}}\Delta ir_{i,t-j} + \sum_{j=0}^{q-1}\delta_{4\mathrm{ij}}\Delta gdp_{i,t-j} \\ &+ \sum_{j=0}^{q-1}\delta_{5\mathrm{ij}}\Delta oil_{i,t-j} + \mu_i + \epsilon_{it}\end{aligned} \tag{31}$$

Second specification group (*cpi neer ppi ir gdp*):

$$\begin{aligned}\Delta cpi_{it} = \phi_i\big(&cpi_{i,t-1} - \theta_{1i}neer_{it} - \theta_{2i}ppi_{it} - \theta_{3i}ir_{it} - \theta_{4i}gdp_{it}\big) \\ &+ \sum_{j=1}^{p-1}\lambda_{ij}^{*}\Delta cpi_{i,t-1} + \sum_{j=0}^{q-1}\delta_{1\mathrm{ij}}\Delta neer_{i,t-j} + \sum_{j=0}^{q-1}\delta_{2\mathrm{ij}}\Delta ppi_{i,t-j} \\ &+ \sum_{j=0}^{q-1}\delta_{3\mathrm{ij}}\Delta ir_{i,t-j} + \sum_{j=0}^{q-1}\delta_{4\mathrm{ij}}\Delta gdp_{i,t-j} + \mu_i + \epsilon_{it}\end{aligned} \tag{32}$$

Third specification group (*cpi neer ppi gdp*):

$$\begin{aligned}\Delta cpi_{it} = \phi_i\big(&cpi_{i,t-1} - \theta_{1i}neer_{it} - \theta_{2i}ppi_{it} - \theta_{3i}gdp_{it}\big) \\ &+ \sum_{j=1}^{p-1}\lambda_{ij}^{*}\Delta cpi_{i,t-1} + \sum_{j=0}^{q-1}\delta_{1\mathrm{ij}}\Delta neer_{i,t-j} + \sum_{j=0}^{q-1}\delta_{2\mathrm{ij}}\Delta ppi_{i,t-j} \\ &+ \sum_{j=0}^{q-1}\delta_{3\mathrm{ij}}\Delta gdp_{i,t-j} + \mu_i + \epsilon_{it}\end{aligned} \tag{33}$$

Fourth specification group (*cpi neer ppi ir*):

$$\Delta cpi_{it} = \phi_i\left(cpi_{i,t-1} - \theta_{1i}neer_{it} - \theta_{2i}ppi_{it} - \theta_{3i}ir_{it}\right) + \sum_{j=1}^{p-1}\lambda_{ij}^{*}\Delta cpi_{i,t-1} + \sum_{j=0}^{q-1}\delta_{1\mathrm{ij}}\Delta neer_{i,t-j} + \sum_{j=0}^{q-1}\delta_{2\mathrm{ij}}\Delta ppi_{i,t-j} + \sum_{j=0}^{q-1}\delta_{3\mathrm{ij}}\Delta ir_{i,t-j} + +\mu_i + \epsilon_{it} \quad (34)$$

where $p \in \{0,1, \dots ,6\}$ and $q \in \{0,1, \dots ,6\}$

Overview of the four specification groups is given in Table 7.

Table 7: Model specification groups

Variables	Denotation	Specifications			
Consumer price index	*cpi*	×	×	×	×
Nominal effective exchange rate	*neer*	×	×	×	×
Producer price index	*ppi*	×	×	×	×
Interest rate	*ir*	×	×		×
Gross domestic product	*gdp*	×	×	×	
Price of oil	*oil*	×			

The following two sections cover overall as well as detailed results of our estimations. We, however, do not want to overemphasize the precision of the estimates. One of the main reasons to interpret the results with caution is the number of observations. As discussed previously, the mean group and pooled mean group models are appropriate for panels with large number of both time periods (*T*) and cross-sections (*N*). The panel used in this work has 41 time periods but only 12 cross-sectional groups, which may not be a sufficiently large amount. Furthermore, according to Pesaran, Shin and Smith (1999), the number of time periods and the number of cross-sections should be of the same order of magnitude. Our sample does not satisfy this condition. We attempted to split the covered time period into two (approximately in half, which would roughly correspond to the pre-financial-crisis and post-financial-crisis time period), but the estimations did not yield results that would exhibit significantly better qualities. The *T* and *N* were more similar in magnitude (20 or 21 vs. 12) but this improvement may have been outweighed by the overall lower number of observations. The results presented in the following sections therefore correspond to the panel covering the whole time period (i.e. from the first quarter of 2003 to the first quarter of 2013).

5.3.1 Overall empirical results

The empirical results of the pooled mean group, mean group and fixed effects estimations of the long-run exchange rate pass-through (θ_1) and the error correction parameter (ϕ) are presented in Table 8. It was, however, not possible to obtain results for all of the specifications, as Stata was not able to execute some of the calculations. The most common problem was nonconvergence of some of the models, which occurred despite the use of the maximum number of iterations. In other instances, the estimation resulted in a numerical overflow – it became too large for Stata to deal with.

Table 8: Long-run exchange rate pass-through and error correction estimates

	cpi neer ppi ir gdp oil			*cpi neer ppi ir gdp*		
	MG	PMG	DFE	MG	PMG	DFE
			0 lags			
ERPT	0.09	0.19	N/A	0.01	0.18*	N/A
EC	-0.23***	-0.07***		-0.22***	-0.07***	
			1 lag			
ERPT	**-0.21**	-0.18**	-0.23*	N/A	0.18***	N/A
EC	**-0.30*** **	-0.08**	-0.09***		-0.09***	
			2 lag			
ERPT	N/A	0.20***	-0.29**	0.62	0.26***	N/A
EC		-0.06**	-0.08***	-0.31***	-0.06*	
			3 lag			
ERPT	**N/A**	0.27***	-0.21	**0.78**	7.19*	-0.24
EC		-0.16**	-0.75***	**-0.42*** **	-0.01**	0.07***
			4 lag			
ERPT	0.85	**0.10*** **	-0.06	**N/A**	**-0.53**	0.08
EC	4.82	**-0.24**	-0.08***		**-0.07**	-0.08***
			5 lag			
ERPT	N/A	N/A	-0.32	-1.48	-0.14***	-0.03
EC			-0.08***	-2.78*	-0.69**	-0.07***
			6 lag			
ERPT	N/A	N/A	0.35	N/A	N/A	-0.07
EC			-0.07***			-0.06***

	cpi neer ppi gdp			*cpi neer ppi ir*		
	MG	PMG	DFE	MG	PMG	DFE
			0 lags			
ERPT	0.08	0.09	N/A	0.19	0.21***	-0.15
EC	-0.19***	-0.10***		-0.25***	-0.11***	-0.08***
			1 lag			
ERPT	-0.24	0.12**	N/A	0.21	0.28***	-0.03
EC	-0.25***	-0.11***		-0.32***	-0.09***	-0.09***
			2 lag			
ERPT	-0.89**	0.14***	-0.05	0.35	0.31***	-0.08
EC	-0.23***	-0.09**	-0.07***	-0.32***	-0.85**	-0.07***
			3 lag			
ERPT	0.06	0.08	-0.02	0.26	6.99***	-0.10
EC	-0.30***	-0.11*	-0.07***	-0.36***	-0.01*	-0.07***
			4 lag			
ERPT	-0.50	0.15***	0.06	0.70**	0.33***	-0.01
EC	-0.43***	-0.20*	-0.08***	-0.66***	-0.22*	-0.08***
			5 lag			
ERPT	**0.74**	0.13***	-0.02	**0.74****	0.26***	-0.09
EC	**-0.54****	-0.24*	-0.0745***	**-1.16*****	-0.40	-0.07***
			6 lag			
ERPT	-0.02	**-0.17*****	-0.02	-12.84	**0.24*****	-0.14
EC	-0.96***	**-0.14**	-0.07***	-1.24**	**-0.51**	-0.06***

***, **, * denote significance at the 1%, 5% and 10% level of significance, respectively
N/A signifies that no results are available
Bold denotes the best (or the best available) specification based on information criteria

Source: author's calculation

To determine the suitable lag length, Akaike information criterion (AIC) and Bayesian information criterion (BIC) are used. The information criteria are consistent in selecting the most appropriate model specification. In some cases, however, the preferred model specification is one of the specifications with unavailable results. The preferred lag lengths, based on the information criteria, are displayed in Table 9. In the case, when the results of the preferred model specification are not available, the most appropriate lag length among the available results is given in parenthesis.

Table 9: Optimal lag lengths based on AIC and BIC

	cpi neer ppi ir gdp oil		*cpi neer ppi ir gdp*	
	MG	PMG	MG	PMG
Number of lags preferred by information criteria	3 (1)	4	4 (3)	4

	cpi neer ppi gdp		*cpi neer ppi ir*	
	MG	PMG	MG	PMG
Number of lags preferred by information criteria	5	6	5	6

(#) denotes the best specification among specifications with available results

Source: author's calculation

As described before, the mean group estimator allows the long run coefficients to vary across countries, while the pooled mean group estimator restricts these coefficients to be the same. To evaluate the homogeneity of the long run coefficients and therefore the suitability of the use of the pooled mean group estimator, Hausman test is employed as suggested by Pesaran, Shin and Smith (1999). The MG estimates of the mean of the long run coefficients are consistent. They are, however, inefficient in the case of slope homogeneity. If the long run slopes are homogeneous, the PMG are consistent and efficient. Hausman test can be applied to the difference between the MG and PMG estimators to assess the effect of heterogeneity on the means of the coefficients. The test statistic then has the following form (Asterious, 2005):

$$H = \hat{q}'[var(\hat{q})]^{-1}\hat{q} \sim \chi_k^2 \qquad (35)$$

where $\hat{q}$ is a $k \times 1$ vector of the difference between MG and PMG estimates

$var(\hat{q})$ is the corresponding covariance matrix

The null hypothesis of the Hausman test is the equivalence between the two estimators. Instead of testing the difference between MG and PMG estimators, the test can be also used for the difference between MG and FE estimators.

Table 10: Hausman test for MG and PMG

Specification		Test statistic	P-value
cpi neer ppi ir gdp oil	1 lag	4.12	0.5323
cpi neer ppi ir gdp	3 lags	-2.50	N/A
cpi neer ppi gdp	5 lags	0.44	0.9324
cpi neer ppi ir	5 lags	1.63	0.6528

Source: author's calculation

For first, third and fourth specification we can easily say, that we cannot reject the null hypothesis that there is not a systematic difference between the MG and PMG estimators. The sign of the Hausman statistic for the second specification is negative and p-value is therefore not available. The negative sign can be interpreted as a strong evidence that the null hypothesis cannot be rejected (StataCorp, 2011a).

Since the results of the Hausman test suggest that both MG and PMG are suitable methods of estimation, we can proceed to compare the results of individual specifications.

For the MG estimator of the first specification group (*cpi neer ppi ir gdp oil*), three lags were suggested by the information criteria. The estimates of this specification are, however, not available and therefore the best specification among the available ones (one lag) is considered. Both the MG estimate of the exchange rate pass-through (ERPT) for this specification and its PMG equivalent have a negative sign. This is not in line with our expectations, because it does not make a general economic sense since it would indicate that the domestic currency depreciation leads to a decrease in domestic consumer prices. Contrary to expectations, negative estimates of ERPT do, however, appear in empirical literature (e.g. Mihaljek and Klau (2009) found negative ERPT for Lithuania). Returning to the first specification group of our model and taking the PMG estimate for the specification that was originally chosen by the information criteria as the most appropriate one for the MG estimator (3 lags) or the specification chosen as the most appropriate for the PMG estimator (4 lags), we obtain ERPT estimates that are appropriately signed and statistically significant at the 1% level. They differ in their magnitude (0.28 and 0.10 respectively) and since the 4-lag specification was found to be the most suitable for the PMG estimator, we prefer the second result to the first one. This means that a 1% depreciation of the domestic currency in the Central and Eastern European region results in a 0.10% increase in domestic consumer prices.

The preferred lag length for the second specification group (*cpi neer ppi ir gdp*) is four lags for both estimators. The MG result for this lag length is, however, not available; the most appropriate specification of those with available results (3 lags) is thus considered. The MG estimation of the ERPT for this specification is 0.78, but this result is not statistically significant. The estimate of the ERPT of the PMG equivalent is 7.19, which is considerably higher than any estimates available in the reviewed literature. The estimate is, however, not significant at the 5% level. The PMG estimate of the ERPT for the specification preferred by the information criteria (4 lags) is not statistically significant either, which is appropriate, since the coefficient has an opposite sign than expected. The ERPT estimates for the selected specifications in this group are in conflict not only with the results from previous studies, but with general economic logic as well. Supported by the statistical insignificance of the results, we conclude that the second specification group is a poor fit for the estimation of the exchange rate pass through into consumer prices.

For the third specification group (*cpi neer ppi gdp*) the information criteria recommend the use of a model with five lags for the MG estimator and a model with six lags for the PMG estimator. The MG estimate of ERPT from the preferred model with five lags is 0.74 and statistically not significant. The PMG equivalent is significant at the 1% level and is equal to 0.13. This is a comparable result to the PMG estimate from the *cpi neer ppi ir gdp oil* 4-lag specification of the model discussed above. The results of these two are similar not only in the ERPT estimate but also in the estimate of the error correction parameter (-0.24 for both the *cpi neer ppi ir gdp oil* 4-lag specification and the *cpi neer ppi gdp* 5-lag specification). Moving to the 6-lag specification preferred by the information criteria for the PMG estimator, the ERPT estimate keeps the significance but the magnitude drops again to negative values.

The preferred lag lengths for the fourth specification group (*cpi neer ppi ir*) are the same as for the third group. That is five lags for the MG estimator and six lags for the PMG estimator. The MG estimate of the ERPT for the 5-lag specification is equal to 0.74. In magnitude, this result is comparable with the MG results for the *cpi neer ppi ir gdp* 3-lag specification (0.78) and the *cpi neer ppi gdp* 5-lag specification (0.74) but unlike these two is statistically significant at the 5% level. Moreover, the error correction parameter in this model is appropriately signed and significant at the 1% level. Moving to the PMG results for the same specification, we receive an ERPT estimate of 0.26 significant at the 1% level. This result is consistent with the estimate for the 6-lag specification preferred by the information criteria (0.24).

The estimates for the error correction parameter (which can be also described as the speed of adjustment) vary both in magnitude and in significance, but are negatively signed for all of the preferred specification as expected. The expectation of negativity stems from the assumption that the variables tend to return to a long-run equilibrium. In the case that the parameter would be equal to zero, there would be no evidence of a long-run equilibrium (Blackburne and Frank, 2007).

Since the estimation results for the fourth specification group are the most plausible, consistent and significant, we select it for further, more detailed investigation. Before proceeding to do so, the overall results of the dynamic fixed effects (DFE) estimator are discussed.

Table 11: Hausman test for MG and DFE

Specification		Test statistic	P-value
cpi neer ppi ir gdp oil	1 lag	6.71	0.2430
cpi neer ppi ir gdp	3 lags	0.05	0.9996
cpi neer ppi gdp	5 lags	0.01	0.9996
cpi neer ppi ir	5 lags	0.07	0.9950

Source: author's calculation

Nearly all of the DFE estimates of the ERPT (25 out of 28) are negative and only one is statistically significant at the 5% level. The Hausman test for the four preferred and available specifications[7] suggests that we cannot reject the null hypothesis that there is not a systematic difference between the MG and DFE estimators (see Table 11), but since the negative ERPT estimates are against the general economic logic and the significance of these results is negligible, we regard the dynamic fixed effects as an unsuitable method of estimation of the exchange rate pass-through in our case. The error correction parameters are on the other hand all with only one exception appropriately signed and all significant at the 1% level. Their magnitude, however, varies substantially.

5.3.2 Detailed results of the selected specification

As mentioned above, the fourth specification group (models including the variables consumer prices, nominal effective exchange rate, producer price index and interest rate) was selected for further investigation. First, the individual country MG estimates of the exchange rate pass-through and the speed of adjustment (error correction parameter) for the 5-lag model are discussed. Second, the short run effects of

[7] Specifications selected by the information criteria for the MG estimator and the specifications selected among the ones with available results if the preferred ones were not available.

exchange rate pass-through for individual countries for the same specification are presented.

The estimated exchange rate pass-through varies substantially among the individual Eastern and Central European countries. The results for two of the twelve examined countries (Poland and Romania) suggest a negative exchange rate pass-through. As discussed previously, this goes against the general logic since prices are expected to increase when currency depreciates. Surprisingly, the ERPT estimate for Romania is significant at the 1% level. Another two estimates do not lie within the expected interval from zero to one. The estimate of ERPT for Bulgaria is equal to 1.28 and the Lithuanian estimate to 4.52. Both of these estimates are significant at least at the 10% level. The pass-through rates larger than one may be caused by the limited amount of variables included in this specification of the model. Coricelli, Jazbec and Masten (2006) note, that ERPT might exceed one if no controls for real supply and demand effects are included. In our models, the GDP and the price of oil are used as the proxies of demand and supply shocks, but they are not included in this specification. The remaining countries obtained exchange rate pass-through estimates in the expected interval and can be split into two groups based on the magnitude of the estimates. The ERPT estimates of the first group of countries are rather low, ranging from 0.15 for the Czech Republic to 0.34 for Russia. The other countries in this group, ordered from the lowest ERPT to the highest, are Slovenia, Macedonia and Estonia. These estimates are, with the exception of Slovenia, not statistically significant. The second group with higher pass-through estimates consists of Latvia, Hungary and Croatia (listed from the smallest to the largest estimate). The estimates of the ERPT differ only slightly for these countries, from 0.78 to 0.81, but vary in significance. The speed of adjustment coefficient is negative for all countries with the exception of Lithuania and Poland. For more details on the magnitudes and significance of both ERPT and EC estimates, refer to Table 12.

Table 12: Individual country MG long-run exchange rate pass-through and error correction estimates for the *cpi neer ppi ir* 5-lag specification

	cpi neer ppi ir	- 5 lags
	ERPT	EC
Bulgaria	1.28*	-1.26***
Croatia	0.81***	-3.75***
Czech Republic	0.15	-0.30
Estonia	0.32	-1.34***
Hungary	0.79**	-0.50**
Latvia	0.78	-0.19
Lithuania	4.52***	0.50
Macedonia	0.27	-1.43**
Poland	-0.39	0.21
Romania	-0.19***	-1.85***
Russia	0.34	-0.50***
Slovenia	0.23***	-3.55***

Source: author's calculation

The exchange rate pass-through estimates discussed so far were estimates of the effects of changes in the nominal effective exchange rate on consumer prices in the long run. These effects were the main focus of the work, but the mean group and pooled mean group methods of estimation also provide information on the short run effects. Since the model under scrutiny includes five lags of all explanatory variables, it is possible to assess the impact of changes in the exchange rate either immediately or in one to five quarters (coefficients $\delta_{1i0},\dots,\delta_{1i5}$). The short run coefficients are displayed in Table 13. The results for the short-run pass-through are rather unsatisfactory. Majority (98 out of 144) of the estimates have a negative (that is the "wrong") sign and the significance of the results is very low. The PMG estimator seems to give slightly better results – it provided less negative estimates (52 for MG and 46 for PMG) and more of the estimates are statistically significant, some even at the 1% level. Unfortunately, the higher significance does not apply only to the positive estimates but to the negative ones as well. Even though the short-term results are in general poor and hardly interpretable, there are some countries for which the estimates make sense. For example the MG estimates for Romania increase from 0.49 in the current quarter to 0.72 two quarters later (current and second quarter estimates significant at the 1% level, first quarter estimate significant at the 5% level). From the third quarter on, the magnitude of the pass-through diminishes (ending at 0.12 in the fifth quarter) and also loses its significance.

Table 13: Individual country MG and PMG estimates of short-run exchange rate pass-through for the *cpi neer ppi ir* 5-lag specification

	cpi neer ppi ir - 5 lags					
	Short run coefficients for *neer*					
Response (in)	immediately		1 quarter		2 quarters	
	MG	PMG	MG	PMG	MG	PMG
Bulgaria	-1.16	0.45	-1.39	-0.24	-1.09	0.18
Croatia	-1.70**	0.37***	-1.24***	-0.37**	-0.69	-0.37**
Czech Republic	0.05	-0.01	-0.09	-0.17***	-0.05	-0.12**
Estonia	-0.59	-0.22*	-0.69	-0.35**	-0.64	-0.18
Hungary	-0.18	0.16***	-0.25	-0.08	-0.16	-0.05
Latvia	0.24	0.42***	-0.35**	-0.43***	0.10	0.09
Lithuania	1.70	0.11	1.17	-0.50***	1.22*	0.05
Macedonia	-0.12	0.64***	0.21	0.66***	-0.23	-0.01
Poland	-0.01	0.02	-0.07	-0.02	-0.10	-0.04*
Romania	0.49***	0.04	0.56**	-0.15*	0.72***	0.05
Russia	-0.05	-0.12*	-0.37	-0.47***	-1.11***	-0.82***
Slovenia	-0.09	-0.22**	-0.19	-0.32***	-0.37*	-0.48***

	cpi neer ppi ir - 5 lags					
	Short run coefficients for *neer*					
Response (in)	3 quarters		4 quarters		5 quarters	
	MG	PMG	MG	PMG	MG	PMG
Bulgaria	-0.95	-0.14	-0.64	0.13	-0.07	0.57***
Croatia	-0.49	0.01	-0.21	0.11	-0.51**	0.02
Czech Republic	0.01	-0.05	0.05	0.01	-0.10	-0.14***
Estonia	-0.90***	-0.45***	-0.77**	-0.48**	-0.42*	-0.29**
Hungary	-0.16	-0.03	-0.14	0.06	-0.22*	-0.03
Latvia	-0.12	0.10	-0.44*	-0.23*	-0.38**	-0.04
Lithuania	0.59	-0.18	0.52	-0.05	0.05	0.05
Macedonia	-0.57	-0.35**	-0.10	0.50***	0.13	0.76***
Poland	-0.09	-0.06**	-0.09	-0.05*	-0.09	-0.11***
Romania	0.50*	-0.18**	0.21	0.01	0.12	-0.05
Russia	-1.07***	-0.22	-0.85***	-0.11	-0.31**	-0.06
Slovenia	-0.18	-0.25***	-0.26	-0.31***	0.10	0.05

Source: author's calculation

Similar behavior can be also observed in the MG estimates for Croatia. For some other countries (the Czech Republic and Hungary for both MG and PMG estimates), the values drop before the first quarter then grow and eventually start declining. Since the results are, however, overall strongly unsatisfactory, we should be very careful in interpreting these dynamics as general trends.

5.4 Results discussion

This section compares our results with the results from previous studies and discusses the hypotheses that a) the domestic prices are influenced by the changes in the exchange rate more in developing countries than the developed ones, and b) there are significant differences between the exchange rate pass-through for the euro area countries, countries that have not adopted the euro yet and countries that are not members of the European Union.

5.4.1 Results comparison

The results of our estimates are compared with the results from previous studies. Both long-term and short-term exchange rate pass-through estimates for individual countries are discussed. The division into the short-term and long-term estimates is clear in our case and for several authors, who estimated either short-term or long-term exchange rate pass-through or both and labeled them so. For studies where the authors did not denote the estimates to be either short-term or long-term but rather provided estimates for different time periods in months or quarters, we consider the long term to start at 24 months / eight quarters. The decision to set the long term to start at two years was based on the division into the short term and the long term in our model, where short-term estimates end in five quarters. Period of eight quarters was the closest larger time period present in the results by other authors and at the same time appeared to be sufficiently larger than the five quarters to be called long-term.

Since majority of our short-run estimates for Bulgaria are negative, there are some significant differences from the results available in the literature. The negative estimates are hardly comparable with the previous results and we thus focus only on the positive estimates. The immediate response based on PMG is 0.45, which is very similar to the three-month 1997-2003 estimate equal to 0.44 provided by Bitāns (2004). We do not have a positive estimate for one quarter available so we are not able to compare the reaction delayed by one quarter; however, we have PMG estimates for 2 quarters (0.18) and 4 quarters (0.13) which are smaller but comparable to the results by Beirne and Bijsterbosch (2011) (0.20 and 0.21 respectively), who moreover estimated ERPT for a similar time period (1995-2008). Thus, we can conclude that there are some similarities in the Bulgarian short-run ERPT estimates in this thesis and previous works. The long-run estimate (1.28) is similar to the 24-month 1993-1997 estimate by Bitāns (2004) (1.19), but since there is a decade between the two time periods and also the more current results by Bitāns (2004) and Beirne and Bijsterbosch (2011) are significantly lower, strong conclusions should not be drawn.

A comparison of Croatian results with the previous studies faces the same problem as the comparison in the case of Bulgaria – the negativity of a majority of the short-run ERPT estimates. We therefore focus again only on the positive – and thus meaningful – results. As in Bulgaria, they are all pooled mean group estimates. The short-term estimates are not in line with the previous research. The immediate response is moderate (0.37) but drops down for three, four and five-quarter estimates. The response delayed by four quarters is by far the largest of these three (0.11) but is still 12 percentage points lower than the equivalent estimate provided by Bitāns (2004). In the long run, our results suggest a high exchange rate pass-through of 0.81, which is substantially more than the previous estimates by Bitāns (2004) and Billmeier and Bonato (2002) (0.36 and 0.3 respectively). The estimation periods of this thesis and the two other studies nearly do not overlap, which might be a reason for the significantly different results.

A number of the short-run estimates of the exchange rate pass-through for the Czech Republic are negative. The remainder (mostly MG estimates) is positive, but very low. None of the short-run estimates exceed 0.05. There is a variety in the short-run ERPT magnitudes obtained in previous studies from very low ones (0.02 – Bitāns (2003)) to rather high ones (0.61 - Ca'Zorzi, Hahn and Sánchez (2007)). Our results are similar to the 3-month 1993-1997 estimate (0.02) and the 3- and 6-month 1998 – 2003 estimates (0.05 and 0.09 respectively) by Bitāns (2004) and the 12-month estimate (0.03) by Korhonen and Wachtel (2006). Furthermore, Darvas (2001) provides a short-run estimate of 0.10, which is larger than our results, but still within the lower range. The long-run ERPT estimate for the Czech Republic obtained from our model (0.15) is equal to the long-run estimate by Darvas (2011) and nearly equal (difference 0.01) to the second period estimate by Bitāns (2004).

Our model did not generate any positive short-run ERPT estimates for Estonia, so it is not possible to compare them with the results available in the reviewed literature. The estimate of the long-run ERPT is on the other hand positive and being equal to 0.32, it is very close to the 24-month estimate for the period 1998-2003 provided by

Bitāns (2004). The difference is only 0.02 as the estimate by Bitāns (2004) is equal to 0.34. The remaining long-run estimates for Estonia from the previous studies are higher than these two.

Disregarding the negative results, the ERPT estimates for Hungary are in line with some of the previous studies. The immediate response to changes in the exchange rate based on PMG is 0.16. This is very close to the 3-month estimates for both periods (0.14 for the period 1993-2001 and 0.15 for the period 2001-2003) provided by Bitāns (2004). Our 4-quarter PMG estimate of 0.06 is equal to the 12-month estimate by Korhonen and Wachtel (2006) and comparable to the short-run estimate by Coricelli, Jazbec and Masten (2006) (0.05). The long-run ERPT estimate (0.79) differs from results available in the reviewed literature but it is of a similar order of magnitude as the long-run estimates provided by Beirne and Bijsterbosch (2011) and Coricelli, Jazbec and Masten (2006) and the 8-quarter estimate by Ca'Zorzi, Hahn and Sánchez (2007) (0.63, 0.97 and 0.91 respectively).

The short-run ERPT estimates for Latvia go against the previous results. Considering only the positive values, our short-run pass-through is the highest immediately (MG estimate 0.24, PMG estimate 0.42) and decreases to 0.09 or 0.10 in two to three quarters. In the previous literature, the dynamic is opposite. ERPT starts at low values and grows with the increase in time between the change in the exchange rate and the resulting effect. The long run estimate (0.78) agrees with the results by Beirne and Bijsterbosch (2011) in that the ERPT is rather large (0.62 for 48-month estimate and 0.97 for a general long run estimate), but the exact magnitudes differ.

Even though Lithuania is the first country with majority of positive short-run estimates, the results are overall very poor. In the short-run the MG and PMG estimates differ significantly (for example the difference in immediate response is 1.59) and neither of the estimators provides results in line with previous research. Since the long-run ERPT comes up to 4.52, it also cannot be compared with any of the previous results, where the largest estimate is equal to one.

The short-run estimates for Macedonia also vary between the two estimators, the differences between the positive ones are, however, not as large as in the case of Lithuania. The effect of the change in the exchange rate delayed by one quarter based on MG (0.21) is equal to the estimate provided by Bitāns (2004) for the period 1999-2003 and similar to the estimate for the period 1994-1998 (0.26). The long-run exchange rate pass-through for Macedonia is equal to 0.27, which is in line with the 24-month 0.30 estimate for the more current period provided by Bitāns (2004).

For Poland, there is only one positive short-term estimate of the exchange rate pass-through. It is the MG estimate of the immediate response of prices to the changes in exchange rate and is equal to 0.02. This value is in line with some of the previous studies that suggest very low levels of ERPT for Poland. Korhonen and Wachtel (2006) provided a 12-month estimate equal to 0.09, Darvas (2001) estimated a zero short-term pass-through and Coricelli, Jazbec and Masten (2006) even obtained a negative short-term ERPT of -0.02. The remainder of available short-run estimates lies, however, in the 0.27-0.51 range. The long-run ERPT is significantly negative (-0.39) and thus contradicts both the general economic logic and previous results.

There are strong differences between the MG and PMG short-run ERPT estimates for Romania. The PMG estimates are rather low and do not correspond to the findings of other researchers. The MG estimates are, on the other hand, in line with some of the previous results. The 1-quarter estimate (0.56) is lower by 0.02 and the 2-quarter estimate (0.72) is larger by 0.02 than the equivalent estimates provided by Bitāns (2004) for the period 1993-1999. The effect in four quarters (0.21) is by 0.02 smaller than the effect for equivalent delay estimated by Bitāns (2004) for the period 1998-2003). The estimate of the long-run ERPT is for Romania negative and therefore does not exhibit similarity with any of the previous results.

The short-run ERPT estimates for Russia are all negative and therefore, we cannot compare them to the previous findings. The long-run ERPT estimate is equal to 0.34, which is slightly over half of the magnitude of the only long-run pass-through estimate available in the reviewed literature (Korhonen and Wachtel, 2006). Since Korhonen and Wachtel (2006) cover different time period than this thesis and provide the only available estimates, conclusions should not be drawn from this difference.

Majority of the short-run estimates is negative also for Slovenia. The two available estimates are 5-quarter estimates and are equal to 0.10 (MG) and 0.05 (PMG). These are lower than all other short-term estimates presented in the reviewed literature, but the mean group estimate comes close to the 0.16 short-run estimate by (Korhonen and Wachtel, 2006). With the value of 0.23, the long-run ERPT estimate is similar to the long-run estimates by Coricelli, Jazbec and Masten (2006) (0.19) and Korhonen and Wachtel (2006) (0.18).

It is clear that there are strong differences in the degree to which are our results in line with the previous findings. There are countries for which the results are in general poor and therefore hard to compare to estimates available in literature; such country is for example Russia with all short-run estimates being negative. For some other countries (e.g. Croatia or Lithuania), meaningful estimates are available but in contrast with the previous research. For majority of the countries either short-run or long-run estimates are in line with the research, while the other is not. The countries that exhibit the highest degree of similarity to the findings in the reviewed literature are Bulgaria, the Czech Republic and Macedonia.

5.4.2 Hypotheses evaluation

It is generally believed (Cheikh, 2011), that the exchange rate pass-through should be higher for developing countries. To test this hypothesis, we consider the long-run ERPT estimates. The choice of the long-run estimates is based on several factors. Primarily, the long-run effects are of a higher interest since these changes in prices are likely to be more permanent than the changes in the short run. Further factors include the better characteristics exhibited by the long-run estimates and their higher simplicity. Both allow for easier and more precise testing of the hypothesis that the domestic prices are influenced by the changes in the exchange rate more in developing countries than in the developed ones.

Table 14: ERPT and stage of development

	ERPT	Regarded as advanced since*
Bulgaria	1.28	-
Croatia	0.81	-
Czech Republic	**0.15**	**2009**
Estonia	**0.32**	**2011**
Hungary	0.79	-
Latvia	0.78	2014
Lithuania	4.52	-
Macedonia	0.27	-
Poland	-0.39	-
Romania	-0.19	-
Russia	0.34	-
Slovenia	**0.23**	**2007**

ERPT estimates are from the *cpi neer ppi ir* 5-lag specification

* according to IMF

Bold denotes countries considered developed for the purpose of this thesis

Source: author's calculation

IMF (2006; 2007; 2008; 2009; 2010; 2011; 2013; 2014)

The pass-throughs and the stages of development for each country are summarized in Table 14. The stage of development is based on the World Economic Outlook by International Monetary Fund (IMF). According to IMF (International Monetary Fund, 2006; 2007; 2008; 2009; 2010; 2011; 2013; 2014), currently four of the reviewed countries (the Czech Republic, Estonia, Latvia and Slovenia) are considered to be advanced. Latvia gained this status in 2014, which is after the end of the model's sample period, and Latvia is therefore considered as emerging or developing for the purpose of this thesis. All of the three remaining countries moved from the group of developing countries to the group of developed ones over the course of the examined time period. They were thus not considered advanced from the beginning but we consider them as such, since even though they were not fully developed yet, they certainly were further along the way than the other ones.

The ERPT estimates for the three developed countries are low, with the highest one being 0.32. This is in line with our hypothesis, but the results for the developing countries are not so straightforward. Bulgaria and Lithuania obtained very high ERPT values (larger than one) and the estimates for Croatia, Hungary and Latvia also lie rather high (between 0.78 and 0.81). The result for Latvia is already problematic since Latvia is considered the most developed of this sample of developing countries. For this reason, according to the hypothesis, the ERPT for Latvia would be expected to lie somewhat lower, closer to the values for developed countries. In contrast to the hypothesis, two of the least developed countries (Macedonia and Russia) obtained ERPT estimates very similar to the estimates of developed countries. Lastly, the ERPT estimates for Poland and Romania are negative, but since such results do not have a strong economic meaning, we do not include them in the consideration of the validity of the hypothesis. Regardless, in the case of Central and Eastern Europe, we cannot accept the hypothesis that the changes in the exchange rate have a higher effect on the domestic prices in the developing countries than in the developed ones.

The second hypothesis to be tested is a hypothesis that there are significant differences among the levels of the exchange rate pass-through for the euro area countries, countries that have not yet adopted the euro and countries that are not members of the European Union (EU). The assessment of the validity of the hypothesis is based on the long-run exchange rate pass-through estimates, as in the previous hypothesis. The information on ERPT and the EU and euro area membership is summarized in Table 15. The five highest ERPT estimates (two estimates larger than one and three in the 0.78-0.81 range) belong to countries from one group, namely to the countries that became EU member states and did not adopted euro before the end of our sample period. The remaining three countries from this group have, however, low ERPT values (two of them even negative ones). The ERPT estimates for the two countries that adopted euro prior to 2013 and the two countries that are not members of the European Union are all of very similar values.

Table 15: ERPT, EU membership and euro adoption

	ERPT	EU member since	Euro adoption
Bulgaria	1.28	2007	-
Croatia	0.81	2013	-
Czech Republic	0.15	2004	-
Estonia	0.32	2004	2011
Hungary	0.79	2004	-
Latvia	0.78	2004	2014
Lithuania	4.52	2004	-
Macedonia	0.27	-	-
Poland	-0.39	2004	-
Romania	-0.19	2007	-
Russia	0.34	-	-
Slovenia	0.23	2004	2007

ERPT estimates are from the *cpi neer ppi ir* 5-lag specification

Source: author's calculation, European Central Bank (2014)

This is in contrast to the expectation that estimates for these two groups would differ significantly, since the characteristics of the two groups are very distinct. The variety of magnitudes among the non-euro-area EU member states and the similarity between the euro area members and countries outside of the EU lead to a conclusion that it is necessary to reject the hypothesis of significant difference of the exchange rate pass-through among the three groups.

6 Conclusion

This thesis provides short-run and long-run estimates of the exchange rate pass-through in twelve Central and Eastern European countries – Bulgaria, Croatia, the Czech Republic, Estonia, Hungary, Latvia, Lithuania, Macedonia, Poland, Romania, Russia and Slovenia. The estimation is based on quarterly data covering the period from the first quarter of 2003 to the first quarter of 2013. Panel cointegration methods, namely the mean group and pooled mean group estimators, are used for the estimation. In addition to these two methods, fixed effects are used as reference. The model in this thesis is partially based on the model of pricing along the distribution chain by McCarthy (1999).

We conducted the estimation of the exchange rate pass-through for several specifications differing in the explanatory variables and lag lengths used. Long-run overall estimates of the exchange rate pass-through are provided for four specification groups each of them with zero to six lags. Therefore, there are 28 models for the overall long-run exchange rate pass-through, each of them estimated using three different methods. The specification groups differ by the set of explanatory variables. The full set includes the nominal effective exchange rate, producer price index, short-term interest rates, gross domestic product and oil prices. For each of the three remaining specification groups some of the explanatory variables were excluded. A specification with five lags including the nominal effective exchange rate, producer price index and interest rates as explanatory variables was selected as the most appropriate and used for further estimations and discussions. The selection was based on Akaike and Bayesian information criteria as well as on the results of the individual models.

The results show, that the pass-through significantly differs across the region. The mean group long-run estimates for most countries range between 0.15 and 0.81, with estimates of five countries lying below 0.35 and three above 0.78. The variation is even larger in the short run. The results for some of the countries do not exhibit very good properties (Russia) or are in contrast with previous research (Croatia, Lithuania) while other align exceptionally well with estimates by other researchers (Bulgaria, the Czech Republic, Macedonia). The results do not support either of the two hypotheses. We therefore can accept neither the hypothesis that the exchange rate pass-through is higher in developing countries than in the developed ones, nor the hypothesis that there are significant differences among the pass-through rates in the euro area countries, the European Union members outside of the euro area and the countries that are not members of the European Union.

This thesis contributes to the existing literature on the exchange rate pass-through for several reasons. Firstly, it provides estimates for a region that has not been covered well by previous research. Our sample of the Central and Eastern European countries is one of the largest in the existing literature and includes some of the countries (e.g., Macedonia and Russia) that are examined most seldom. Furthermore, this thesis provides estimates for a more current period than the previous studies. The second major contribution is the use of panel cointegration methods, which not only take into account the time series character of the data, but allow for the use of the cross-sectional information as well. Previous studies on the exchange rate pass-through almost exclusively estimate separate models for each country, therefore their results omit any cross-sectional information. Moreover, many authors use estimation methods, which do not consider the time series properties of the data such as non-stationarity and cointegration. In the few cases when panels are used in the previous studies to estimate the exchange rate pass-through, they are usually homogenous panels. We, on the other hand, use heterogeneous panel estimators and thus are able to provide individual estimates for each country.

Even though the mean group and pooled mean group estimators are currently considered to be the most suitable methods of estimation of the exchange rate pass-through, the results exhibit some shortcomings. There might be several reasons for this. First, the Central and Eastern Europe has been going through rapid changes. Many of the countries are still transitioning and their economic environment has altered over the course of the examined period. Such changes might be the cause for the inconsistency of the results. Second possible reason is the size of the sample. The mean group and pooled mean group estimators are supposed to provide the best results, when the number of both countries and time periods is large and of a similar magnitude. Our model estimated the pass-through for 12 countries, which is a borderline amount. We originally planned to include more countries, but the availability of the necessary data is very poor in the Central and Eastern European region. For several countries, it was not possible to obtain nearly any useful data. Other countries only started to collect such data recently and the available information was therefore not sufficient to be included in the models.

Even though we faced a shortage of data, this thesis is still an addition to the existing literature as we covered a range of countries from a region that is not frequently present in the previous studies while using some of the more advanced methods of estimation. As stated before, Central and Eastern Europe is still developing and thus it would be interesting to estimate the exchange rate pass-through models again in few years, when the economic development is more stabilized and the available dataset is larger.

Bibliography

Asteriou, D. (2006). *Applied Econometrics: A Modern Approach using EViews and Microfit*. 1st. ed. Hampshire: Palgrave Macmillan.

Babecká-Kucharčuková, O. (2009). Transmission of exchange rate shocks into domestic inflation: the case of the Czech Republic. *Czech Journal of Economics and Finance*. 59(2), pp.137-152.

Bacchetta, P. and van Wincoop, E. (2002). *Why do consumer prices react less than import prices to exchange rates?*. (Working paper 9352). National Bureau of Economic Research, Cambridge, MA.

Baltagi, B. H. (2005). *Econometric Analysis of Panel Data*. 3rd. ed. Chichester: John Wiley & Sons Ltd.

Barhoumi, K. (2005). Long Run Exchange Rate Pass−Through Into Import Prices In Developing Countries: An Homogeneous or Heterogeneous Phenomenon?. *Economics Bulletin*. 6(14), pp.1-12.

Beirne, J. and Bijsterbosch, M. (2011). Exchange rate pass-through in central and eastern European EU Member States. *Journal of Policy Modeling*. 33, pp.241-254.

Bilmeier, A. and Bonato L. (2002). *Exchange rate pass-through and monetary policy in Croatia*. (WP/02/109). International Monetary Fund, Washington, DC.

Bitāns, M. (2004). *Pass-through of exchange rates to domestic prices in East European countries and the role of economic environment*. (Working paper 4/2004). Bank of Latvia, Riga.

Blackburne, E. F. and Frank, M. W. (2007). Estimation of nonstationary heterogeneous panels. *The Stata Journal*. 7 (2), pp.197-208.

Campa, J. M. and Goldberg, L. S. (2002). *Exchange rate pass-through into import prices: Macro or micro phenomenon?*. (Research paper No 475). IESE, University of Navarra, Barcelona.

Campa, J. M., Goldberg, L. S. and Gonzáles-Mínguez, J. M. (2005). *Exchange rate pass-through to import prices in the euro area*. (Working paper 11632). National Bureau of Economic Research, Cambridge, MA.

Ca'Zorzi, M., Hahn, E. and Sánchez, M. (2007). *Exchange rate pass-through in emerging markets*. (Working paper No 739). European Central Bank, Frankfurt am Main.

Cheikh, N. B. (2011). *Long run exchange rate pass-through: Evidence from new panel data techniques*. Université de Rennes 1, Centre de Recherche en Économie et Management, Rennes Cedex, France.

Cheikh, N. B. (2013). *The Pass-Through of Exchange Rate Changes to Prices in the Euro Area: An Empirical Investigation*. Ph.D. thesis, University of Rennes 1.

Coricelli, F, Jazbec, B. and Masten, I. (2006). Exchange rate pass-through in EMU acceding countries: Empirical analysis and policy implications. *Journal of Banking*. 30, pp.1375-1391.

Darvas, Z. (2001). *Exchange rate pass-through and real exchange rate in EU candidate countries*. (Discussion paper 10/01). Deutsche Bundesbank, Frankfurt am Main.

Dabusinskas, A. (2003). *Exchange rate pass-through to Estonian prices*. (No 10, 2003). Bank of Estonia, Tallinn.

European Central Bank (2014). *Euro area 1999 – 2014* [online]. Available from: <http://www.ecb.europa.eu/euro/intro/html/map.en.html>. [Accessed 5 May 2014].

Goldberg, P. K. and Knetter, M. M. (1996). *Goods prices and exchange rates: What have we learned?*. (Working Paper 5862). National Bureau of Economic Research, Cambridge, MA.

Hadri, K. (2000). Testing for stationarity in heterogeneous panel data. *Econometrics Journal*. 3, pp.148-161.

Holmes, M. (2006). Is a low-inflation environment associated with reduced exchange rate pass through?. *Finish Economic Papers*. 19(2), pp.58-68.

Holmes, M. (2008). Long-run pass-through from exchange rate to import prices in African countries. *Journal of Economic Development*. 33(1), pp.97-111.

Hooper, P. and Mann, C. L. (1989). Exchange Rate Pass-Through in the 1980s: The Case of U.S. Imports of Manufactures. *Brookings Papers on Economic Activity*. 1989(1), pp.297-337.

Im, K. S., Pesaran, M. H. and Shin, Y. (2003). Testing for unit roots in heterogeneous panels. *Journal of Econometrics*. 115, pp.53-74.

IMF (2006). *World economic outlook*. Washington, DC.

IMF (2007). *World economic outlook*. Washington, DC.

IMF (2008). *World economic outlook*. Washington, DC.

IMF (2009). *World economic outlook*. Washington, DC.

IMF (2010). *World economic outlook*. Washington, DC.

IMF (2011). *World economic outlook*. Washington, DC.

IMF (2013). *World economic outlook*. Washington, DC.

IMF (2014). *World economic outlook*. Washington, DC.

Klein, M. W. (1990). Macroeconomic aspects of exchange rate pass-through. *Journal of International Money and Finance*. 9, pp.376-387.

Korhonen I. and Wachtel, P. (2006). A note on exchange rate pass-through in CIS countries. *Research in International Business and Finance*. 20, pp.215-226.

Kozluk, T., Banerjee, A. and de Bandt, O. (2008). Measuring long-run exchange rate pass-through. *economics - The Open-Access, Open-Assessment EJournal*. 2, pp.1-36.

Laflèche, T. (1997). The impact of exchange rate movements on consumer prices. *Bank of Canada Review*. Winter 1996-1997, pp.21-32.

McCarthy, J (1999). *Pass-through of exchange rates and import prices to domestic inflation in some industrialised economies*. (Working paper No. 79). Bank for International Settlements, Basel.

Menon, J. (1995). Exchange rate pass-through. *Journal of Economic Surveys*. 9 (2), pp.197-231.

Mihaljek, D. and Klau, M (2009). Catching-Up and Inflation in the Baltics and Southeastern Europe: the Role of the Balassa-Samuelson Effect. *In: Workshops – Proceedings of OeNB Workshops No.15.*

Mumtaz, H., Oomen, Ö. and Wang, J. (2006). *Exchange rate pass-through into UK import prices*. (Working paper no. 312). Bank of England, London.

Pedroni, P. (1997). *Panel cointegration: Asymptotic and finite sample properties of pooled time series with application to the PPP Hypothesis: New results*. (Working paper). Indiana University, Indiana.

Pedroni, P. (1999). Critical values of cointegration tests in heterogeneous panels with multiple regressors. *Oxford Bulletin of Economics and Statistics*. Special issue, 62, pp.653-670.

Pedroni, P. (2000). Fully modified OLS for heterogeneous cointegrated panel. *Advances in Econometrics*. 15, pp.93-130.

Pesaran, M. H., Shin, Y. and Smith R. P. (1999). Pooled mean group estimation of dynamic heterogeneous panels. *Journal of the American Statistical Association*. 94 (446), pp.621-634.

Pesaran, M. H. and Smith R. P. (1995). Estimating long-run relationships from dynamic heterogeneous panels. *Journal of Econometrics*. 68 (1), pp.79-113.

Schröder, M. and Hüfner, F. P. (2002). *Exchange rate pass-through to consumer prices: a European perspective*. (Discussion Paper No. 02-20). Center for European Economic Research, Mannheim.

Silver, M. (2007). *Do unit value export, import, and terms of trade indices represent or misrepresent price indices?*. (Working paper 07/21). International Monetary Fund, Washington, DC.

StataCorp (2011a). *Stata base reference manual*. Release 12. Texas: Stata Press.

StataCorp (2011b). *Stata: Release 12*. Statistical Software. College Station, TX: StataCorp LP.

Vonnák, B. (2010). *Risk premium shocks, monetary policy and exchange rate pass-through in the Czech Republic, Hungary and Poland*. (MNB Working Papers 2010/1). Magyar Nemzeti Bank, Budapest.

Data sources

Bank for International Settlements (2014). *BIS effective exchange rate indices* [online]. Available from: <http://www.bis.org/statistics/eer/index.htm>. [Accessed 8 March 2014].

Eurostat (2014). *GDP and main components - Current prices* [online]. Available from: <http://appsso.eurostat.ec.europa.eu/nui/show.do?dataset=namq_gdp_c>. [Accessed 1 April 2014].

Federal State Statistics Service (2014). *Producer price indices for industrial goods, Russian Federation, 2001-2014* [online]. Available from: <http://www.gks.ru/bgd/free/b00_25/isswww.exe/stg/d000/I000750R.HTM>. [Accessed 4 April 2014].

International Monetary Fund (2014). *International Financial Statistics* [CD-ROM].

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