Explaining inflation dynamics: the role of price setting rules, expectations and exchange rate

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Abstract

Inflation is one of the key macroeconomic variables. Not only central banks, but also the public pays attention to its fluctuation. This thesis aims to contribute to the better understanding of what drives changes of the inflation rate. The first chapter provides descriptive statistics of store-level pricing practices in Hungary and explores whether changes in the average size or frequency of price changes are behind the observed volatility of the inflation rate. In the second chapter I analyze the linkages between inflation and inflation expectations and show that changes in household expectations may help to predict changes in the inflation rate. The third chapter is about the drivers of asymmetry in exchange rate pass-through to import prices. I show that the asymmetry in exchange rate pass-through is positively correlated with the inflation rate of imported goods.

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Introduction

Chapter 1[1] uses Hungarian micro CPI-data between December 2001 and June 2007 to provide empirical evidence on store-level pricing practices in Hungary. Our focus is threefold. First, in supplement to a large number of empirical studies from all over the world we calculate simple descriptive statistics about price setting – frequency and average size of price changes – in Hungary, where the inflation rate was relatively high and more volatile during the sample period. Our paper is the first one to provide empirical analysis of price setting in Hungary based on a wide range of products.

Second, given the relatively high volatility of Hungarian inflation rate, we decompose inflation variation to variation in the frequency and average size of price changes. We use the method of Klenow and Kryvtsov (2008) for this decomposition. They argue that this exercise is informative about the question whether price adjustment takes place on the extensive or intensive margin, and helps us distinguishing between the two major sticky-price model families, time-dependent and state-dependent models. Calvo-type time-dependent models imply that all adjustment is on the intensive margin, while state-dependent models imply that some of the adjustment takes place on the extensive margin.

Third, using a similar decomposition into frequency and size effects, we estimate the inflation effect of three major Value-Added Tax (VAT) changes during our sample period.

We find that in terms of price flexibility, Hungary is in between the Euro area and the US: the frequency and average size of price adjustment is larger than in the Euro area, and smaller than in the US. Further, inflation variation is mainly driven by changes in the price increase and price decrease frequencies (i.e. in the relative share of price increases and decreases within all price changes), which is in line with the results reported by Gagnon (2007) in Mexico. Finally, the short-term inflation effect of a unit VAT-increase is estimated to be substantially larger than that of a unit VAT-decrease.

Chapter 2 explores the linkages between inflation expectations and inflation. Inflation expectations influence prices through numerous channels. Investors need reliable inflation projections to make well-founded investment decisions. Firms need to determine their expected inflation rate in order to set prices, to make capital investment and deciding on borrowing and liquidity needs. Expected inflation is crucial in contracts which are not continuously renegotiated, like wages. Consumers use information on the future inflation rate when allocating consumption between today and tomorrow. If inflation expectations are high consumers tend to consume today, which may increase prices further.

[1] The first chapter is based on Gabriel-Reiff (2010), which was a joint work with Adam Reiff.

Although expectations are important because of several reasons, the empirical literature on inflation expectation formation had a quite limited scope in general. Most of the literature analyzed whether expectations (of households, producers or investors) are unbiased and efficient predictors of future inflation rates.

This paper explores whether changes in expectations have an impact on other macroeconomic variables in three countries (Czech Republic, Hungary and the United Kingdom). The selection of countries was motivated by that all countries have inflation targeting monetary regimes, so managing expectations is in the focus of monetary policy and for all countries relatively long time series are available about inflation perceptions and expectations. In this paper I quantify qualitative survey responses about inflation perceptions and expectations and use a SVAR framework to identify the effect of changes in the expected inflation rate. Previous papers usually used simple ordering assumptions to put inflation expectations into VARs. The main contribution of this paper is that it proposes a SVAR framework with sign restrictions, which can be more appropriate to identify the effect of expectation shocks. Nominal wages are also included among the variables of the SVAR, to underpin one of central banks' main concerns, namely that non-anchored expectations may have an impact also on wage setting. The results show that an increase in inflation expectations raises prices and nominal wages in all the three countries. By comparing impulse responses I also evaluate how anchored expectations are in the three countries. Expectations are the most anchored in United Kingdom and the least in Hungary.

Chapter 3 discusses how perceived trends in nominal exchange rate and foreign prices affects exchange rate pass-through to import prices. Asymmetric pass-through has been analyzed by many theoretical and empirical papers. The theoretical literature provides two main mechanism which are potentially important regarding asymmetric pass-through. Knetter (1994) argued, that if firms operate under capacity constraints, which limits the magnitude of production on the short run, it is not worthwhile to have low prices. Hence, a depreciation of the exporter's currency might result in a lower pass-through than an appreciation, for which the capacity constraint is not binding. The other argument provides an explanation of asymmetries in the other direction. If firms compete strategically for market shares, an appreciation of the exporter's currency will result in firms adjusting by reducing the markup, while during a depreciation they will maintain the markup and allow prices to fall.

In this paper I provide an alternative explanation for asymmetric exchange rate pass-through. I estimate short run exchange-rate pass-through for seven European economies and I provide evidence that in some countries exchange rate pass-through is indeed asymmetric. I show that there is a positive correlation between the measured asymmetry in exchange rate pass-through and the average inflation of import prices. Next I calibrate a small menu cost model to analyze the exchange rate pass-through. The most closely related paper is Flodén and Wilander (2006), which also calibrated a menu cost model to analyze exchange rate pass-through. However they examined pass-through only at the individual firm level. In this paper I analyze the problem in a general equilibrium setup, so strategic interactions among firms also play a role. I show that measured pass-through is quite sensitive to different anticipated trends in exogenous variables, which may potentially explain the surprising diversity of pass-through estimates in time and cross-section in emerging economies. Although the model is very stylized, I could replicate many stylized facts on exchange rate pass-through: pass-through is incomplete, pass-through is higher if inflation is higher, asymmetry increases if inflation is higher.

Chapter 1

Price Setting in Hungary – A Store-Level Analysis

This paper uses Hungarian micro CPI data between December 2001 and June 2007 to provide descriptive statistics of store-level pricing practices in Hungary. First we present the frequency and average size of price changes, the duration distribution of price spells and calculate mean durations for different product categories. Then we decompose the observed variations in the inflation rate to variations in frequencies and sizes. Finally we estimate the inflation effects of three general VAT-rate changes during our sample period.

1.1 Introduction

This paper uses Hungarian micro CPI-data between December 2001 and June 2007 to provide empirical evidence on store-level pricing practices in Hungary. Our focus is threefold. First, in supplement to a large number of empirical studies from all over the world (e.g. Dhyne et al. (2006) for the Euro area, Bils and Klenow (2004), Nakamura and Steinsson (2008) and Klenow and Kryvtsov (2008) for the US), we calculate simple descriptive statistics about price setting – frequency and average size of price changes, and mean duration of prices – in Hungary, where the inflation rate was relatively high and more volatile during the sample period (Figure 1). Our paper is the first one to provide empirical analysis of price setting in Hungary based on a wide range of products.[1]

[1] Ratfai (2007) presents microeconomic evidence about the frequency and size of price changes in Hungary, but his data set is limited to 14 meat products, observed between 1993-96,

Second, given the relatively high volatility of Hungarian inflation rate, we decompose inflation variation to variation in the frequency and average size of price changes. We use the method of Klenow and Kryvtsov (2008) for this decomposition. They argue that this exercise is informative about the question whether price adjustment takes place on the extensive or intensive margin, and helps us distinguishing between the two major sticky-price model families, time-dependent and state-dependent models. Calvo-type time-dependent models (Calvo (1983)) imply that all adjustment is on the intensive margin, while state-dependent models (e.g. Sheshinsky and Weiss (1977), Dotsey et al. (1999), Golosov and Lucas (2007)) imply that some of the adjustment takes place on the extensive margin.

Third, using a similar decomposition into frequency and size effects, we estimate the inflation effect of three major Value-Added Tax (VAT) changes during our sample period. The Hungarian fiscal authorities reduced the number of different VAT-rates with a series of VAT-changes between 2004 and 2006[2], which provides a unique natural experiment to study the inflation effects of general, easily identifiable cost-push shocks.

We find that in terms of price flexibility, Hungary is in between the Euro area and the US. The frequency and average size of price adjustment is larger than in the Euro area, and smaller than in the US. The mean duration of prices is somewhat smaller than the implied mean duration in the Euro area, and similar to mean duration estimates in the US. Further, inflation variation is mainly driven by changes in the price increase and price decrease frequencies (i.e. in the relative share of price increases and decreases within all price changes), which is in line with the results reported by Gagnon (2007) in Mexico. Finally, the short-term inflation effect of a unit VAT-increase is estimated to be substantially larger than that of a unit VAT-decrease. We also find that VAT-changes influenced the prices of items not directly affected by these changes.

The remaining of the paper is organized as follows. After describing the data set in section 2, sections 3-4 present simple descriptive statistics about the frequency and average size of price changes. Then we discuss overall and sectoral duration distributions, and direct and indirect mean duration estimates in section 5. In section 6 we decompose the inflation variation into frequency and size effects. Section 7 analyzes the effect of general VAT-rate changes to the inflation rate and the stores' pricing practices. Section 8 concludes.

in a period with much higher inflation.

[2]This intention of simplifying the VAT-system was not communicated in advance, and become clear to decision makers only ex post.

1.2 Data

To analyze price setting at store level usually two types of datasources are used. The first one is the monthly price quotes collected by statistical offices, the second one is scanner data, which is produced by big retail chains. The advantage of CPI dataset is that the sample of stores is representative in geographical terms and also regarding the types of the outlets. On the other hand CPI datasets have some drawbacks. First the number of products is limited compared to scanner datasets. Using scanner data price indices can be calculated basically for all products sold in the outlet. Second, scanner datasets provide information about the value of products sold in the outlet. Using values of products sold as weights, the calculated price indices can be more informative about the increase in the cost of a representative consumer's consumption basket. Despite these advantages of scanner data, statistical offices tend to use price quotes collected by themselves. The main reason is that scanner data is rarely representative geographically. There are some exceptions. For example since August 2005, the Norwegian statistical office has used scanner data information in order to compute the sub-index for food and non-alcoholic beverages in the Consumer Price Index (Rodriguez and Haraldsen (2006)). However in Norway the outlets that are capable of delivering scanner data to the statistical office via store chain headquarters represents over 90 percent of the total turnover in the grocery market. This is not typical in other countries. According to our knowledge in Hungary scanner datasets are not available for research.

To analyze price setting at sore level this paper uses a data set containing store-level price quotes, originally collected by the Hungarian Central Statistical Office (CSO) for Consumer Price Index (CPI) calculations.[3] Field agents of the CSO record price quotes about several hundred narrowly defined products in a wide range of outlets.

The time span of the data set is between December 2001 and June 2007, with price observations in 67 months.

Regarding the cross-sectional dimension, the best coverage (in terms of the number of items) is achieved in 2006, when the sample was constructed. In this year, we have data about 770 representative items of 896 on the item list. The data construction method implies that representative items disappearing from the item list before 2006 are not in the data set, and therefore the coverage in earlier years is generally weaker (Table1.1). The data set was updated in 2007, when the CSO added 17 new representative items to the item list, and discon-

[3]We thank Borbála Mináry and Beáta Kollár from the Hungarian Central Statistical Office for insightful discussions about the Hungarian CPI data set.

tinued data collection of 40 representative items.[4] Therefore the total number of representative items is 787, but the maximum number in any year is 770.

year	no. of items	CPI-weight	no. of observations
2002	718	66.855	805,630
2003	732	69.148	828,152
2004	739	69.087	841,282
2005	769	70.735	848,188
2006	770	70.122	879,561
2007	742	68.991	415,479
TOTAL	**787**	–	**4,618,292**

Table 1.1: Coverage of the data set by years

In 2006, the total CPI-weight of the 770 representative items in the sample is 70.122 percent. The missing items have either regulated prices (e.g. kindergarten and school catering, electric energy, pipeline gas, highway toll stickers) or the data collection methodology of the CSO makes it impossible to investigate price quotes of identical products over time (e.g. new and used cars). Table 1.2 contains the 2006 coverage of the data set by consumption categories.[5] Table 1.2 indicates that mainly non-energy industrial goods and services are missing, along with some (regulated) energy products.

	CPI-basket		Sample	
Consumption category	**Weight**	**Items**	**Weight**	**Items**
Unprocessed food	5.665	53	5.665	53
Processed food	19.865	140	19.865	140
Energy (oil) products	13.203	16	6.350	8
Non-energy industrial goods	27.077	497	20.515	436
Services	34.190	190	17.726	133
TOTAL	**100.000**	**896**	**70.122**	**770**

Table 1.2: Coverage of the data set in 2006 by consumption categories

The data set can be regarded as 787 mini "panels"of the same number of representative items. For example, for the representative item "bony pork rib

[4]This would imply 770+17-40=747 representative items in 2007, but prices are not recorded between January and June for five representative items (peach, grapes, plums, and two types of theater season tickets), so in fact we only have price data about 742 representative items in the first half of 2007.

[5]To ease comparison, these consumption categories are defined as in Dhyne et al (2006).

with tenderloin" there are 8,667 observations from 162 different outlets. This implies that for this representative item the average number of price quotes per outlet is (8,667/162=) 53.5. Moreover, for 95 stores (out of 162) we have a complete sequence of price quotes, i.e. we observe the price of the item during the whole observation period (67 months). The list of recorded outlets is typically unchanged for other representative items as well, which makes it possible to investigate store-level price developments and the pricing behavior of individual stores.

On average, there are approximately 5,952 observations per representative item in the data set, implying a total number of observations close to 4.7 million (4,684,289).[6] For each observation, we have the following information: price; month of observation; product code (5-digit representative item code); store code[7]; "change code". The change code variable indicates whether the observed price change was due to sales or it was a normal price increase/decrease, and it also identifies price imputations, forced store and/or product replacements, changes of suppliers, changes in product outfits, and mistakes in previous months' quotes. We do not have information about store characteristics (e.g. type of outlet, size, and whether the store is operated in a city or not etc).

Following Baudry et al. (2004), we call an uninterrupted sequence of price quotes of the same product in the same outlet as a price trajectory.[8] Within price trajectories, uninterrupted sequences of price quotes *with the same price* are called price spells.

We determine the number of price trajectories assuming that the following events started a new price trajectory: change of product (both forced and unforced); change of store (both forced and unforced); change of supplier; change of product outfit. According to this definition, we found a total of 272,549 different price trajectories in the data set, with an average length of 17.2 months.

1.2.1 Specific data issues: censoring, sales, imputed prices and VAT-changes

Censoring: for a given product in a given store (i.e. for a given price trajectory), we treat the first observation as left-censored, and the last observation as

[6] In Table 1.1, 65,997 observations from 2001 December are not reported.

[7] The store code also contains a location identifier in terms of county.

[8] Klenow and Kryvtsov (2008) use the notion of quote-lines. The difference between price trajectories and quote-lines is that price trajectories are uninterrupted, while quote-lines may contain periods with no price observations. Our definition of price trajectories implies that for seasonal products (e.g. gloves, cherries), every year's data belongs to a different trajectory.

right-censored. This means that (1) the ages of prices in a price trajectory are unobserved until the first price changes takes place; (2) the duration of the first and last price spell in any price trajectory is unobserved; (3) when estimating frequencies and hazards, the last observation of each price trajectory is not taken into account (as it is unobserved whether the price actually changed or not).

In case of *sales*, we have the change code variable indicating whether the quoted price was actually a sales price or not. However, we do not think that this change code variable is reliable, as the number of beginning-of-sales flags is much higher than the number of end-of-sales flags. As we did not want to use a somewhat arbitrary sales-filter (applied either simultaneously or separately from the change code variable), we used the "raw"data without any sales filtering. That is, we treated sales-induced price changes as "normal"price changes. This may bias the frequency and size estimates upward.[9]

The change code variable also indicates if a specific price was *imputed* by the CSO. When handling missing price data the CSO applied the method provided by Eurostat. The essence of the method is to replace the price actually surveyed in the previous month but missing in the reference month by an estimated (imputed) price, using an index calculated from monthly county prices of the given representative item. The method can be used for two months, following this period of time the price of the product or service that is similar to the earlier one must be collected.

These imputed prices are not observed prices, thus should be treated differently.[10] Therefore we replaced all imputed prices with the previous months's price quotes (i.e. we used the "carry forward"approach). This leads to downward bias in the frequency estimates only if there were price changes in both the month of imputation and the month after.

VAT-rates in Hungary	**Lower**	**Middle**	**Top**
– Dec 31, 2003	0%	12%	25%
Jan 1, 2004 – Dec 31, 2005	5%	**15%**	25%
Jan 1, 2006 – Aug 31, 2006	5%	15%	**20%**
Sep 1, 2006 –	5%	**20%**	20%

Table 1.3: VAT rates in Hungary

[9] Size estimates are likely to be biased upwards as sales-induced price changes tend to be relatively large.

[10] 1.6 percent of price observations are imputed prices. More than 50 percent of the imputed prices differs from the prices observed one month before the imputation. This confirms that most of the price changes are not real ones.

VAT-changes: relative to other EU member states, in Hungary there have been quite frequent changes in the general VAT-rates recently (Table 1.3). In January 2004, the middle and lower rates increased from 12 to 15 percent and from 0 to 5 percent, respectively, while in January 2006 the top rate was reduced from 25 to 20 percent. Finally, from September 2006, the middle rate was again increased to 20 percent, which means that now there are only two different VAT-categories in Hungary.

1.3 Frequency of price changes

Stores adjust prices only infrequently, but there is substantial variation among products (Figure 2). In the whole sample the (weighted) average frequency of price changes is 21.5 percent, and the (weighted) median is 14.0 percent. The minimum frequency is 0 percent for "acupuncture treatment"and the maximum is 100 percent for "currency exchange". In the Euro area the average frequency is 15.1 percent (Dhyne et al, 2005), smaller than in Hungary. In the US, Bils-Klenow (2004) report an average frequency of 26.1%, while Klenow-Kryvtsov (2007) finds 36.2% (on a different sample).

Product-level frequencies are extremely heterogenous, which is also the case at the level of main product categories (Table 1.4). Price change frequencies are highest for food and energy items, while the rate of price adjustment is the lowest in case of services. In the whole sample, price decreases are almost as frequent as price increases, so our results - similarly to previous findings - do not support downward nominal rigidity. Also in this respect there is substantial heterogeneity across sectors. Services' prices rarely change downwards, whereas in case of durable goods and clothes price decreases are even more frequent than increases.[11]

There is relatively strong correlation between the frequency of price increases and decreases (Figure 3 in the Appendix). This can be the result of volatile cost factors or pricing strategies of stores (e.g. price discrimination through randomizing prices).

As in some sectors the frequency of price changes may follow a seasonal pattern, we calculated average frequencies in the different months of the year for each sector (Figure 1.4 in the Appendix). We tested seasonality also formally by regressing monthly price change frequencies on monthly and VAT dummies. According to this, the frequency of price change is seasonal in the food, clothing and services sectors, but does not follow any seasonal pattern in the durable

[11]This result about items in the clothing category is heavily influenced by the seasonal pattern of many of them: end-of-season sales are taken into account, while higher prices of new collections (after several months of missing data) are not.

CPI category	Freq of change Mean	Freq of change Median	Freq of increase Mean	Freq of increase Median	Freq of decrease Mean	Freq of decrease Median
Food, alcohol, tobacco	0.249	0.208	0.144	0.136	0.104	0.079
Unprocessed food	0.504	0.466	0.232	0.222	0.273	0.212
Processed food	0.192	0.176	0.125	0.123	0.067	0.058
Proc. food excl. alc, tob	0.195	0.191	0.124	0.119	0.070	0.075
Clothing	0.116	0.111	0.045	0.044	0.071	0.057
Durable goods	0.122	0.108	0.050	0.048	0.072	0.058
Other goods	0.111	0.089	0.067	0.054	0.045	0.032
Energy	0.630	0.877	0.385	0.524	0.245	0.342
Services	0.080	0.069	0.067	0.062	0.013	0.005
TOTAL	**0.215**	**0.140**	**0.128**	**0.083**	**0.088**	**0.046**

Table 1.4: Frequency of price changes by product categories

goods and other non-industrial goods categories. Table 1.5 shows the seasonality in the former product categories.[12] In the unprocessed food category seasonality is influenced by the availability of fresh fruits and vegetables, in the processed food category the probability of price increase is the highest in February. In the clothing category the frequency of price changes shows an especially strong seasonal pattern. The effect of the two end-of-season sales is very transparent. Prices decrease temporarily in January and February (winter sales) and also in July and August (summer sales). Also the services sector is characterized with a very strong seasonal pattern in the adjustment frequencies: most price changes take place in the first few months of the year, and towards the end of the year the probability of price increase is low.

1.3.1 Frequency of price changes and dispersion of price observations

The CPI dataset contains information on the price level of products, so it is possible to analyze, how deviation from the average price level affects the probability of price change. However although the sampling procedure of the CSO ensures that in the same store the price quotes for two consecutive months should belong to the same product (with the exception of imputation), in two different stores the products for the same product category do not have to be identical even at the lowest disaggregation level. The product categories are relatively narrowly

[12] The numbers are differences from the average monthly frequency of price increases (decreases) measured in percentage points. Blank cells indicate that the difference was not significant.

	J	F	M	A	M	J	J	A	S	O	N	D
Price increase frequencies												
	J	F	M	A	M	J	J	A	S	O	N	D
Unprocessed food									7.8			
Processed food		3.8										
Clothing			3.1				-2.3	-2.4	2.8	2.2		
Services	13.4	3.6						-4.0			-3.9	-2.8
Price decrease frequencies												
	J	F	M	A	M	J	J	A	S	O	N	D
Unprocessed food						6.6						
Processed food												
Clothing	4.5	5.2					2.9	6.4		-3.0	-3.6	
Services												

Table 1.5: Seasonality in price change frequencies

defined, but brands may be different, which may cause permanent deviation in prices. Table 1.6 shows that prices are indeed quite disperse in all product categories[13]. On the other hand the impact of the deviation from the average price in a given county still have the expected sign on price change probabilities. If the price of a product is higher than the country average, then the probability of price increase is lower and the probability of price decrease is higher.[14]

	Standard dev. of prices	**The effect of the rel. price on the prob. of price inrease**	**price decrease**
Unprocessed food	0.24	-0.96	0.59
Processed food	0.24	-0.42	0.14
Clothing	0.34	-0.05	0.02
Durable goods	0.32	-0.08	0.05
Other goods	0.33	-0.13	0.05
Services	0.44	-0.03	0.01
Whole sample	**0.31**	**-0.29**	**0.12**

Table 1.6: Probabilty of price change and relative prices

Table 1.7 shows that taking the whole sample the dispersion of prices are negatively correlated with price change frequencies. This is consistent with simple

[13]Dispersion of prices were estimated at the county level for all products. the numbers in Table 1.6 are averages for main product categories.

[14]Numbers in Table 1.6 show the effect of 1 percentage point price increase from the county average on the probabilies of price increase and decrease.

menu cost models. If menu cost of price change is higher, then price change frequency is lower and price dispersion is bigger. However the correlation does not negative for all product categories, so the negative correlation for the whole sample can be caused also by differences among the main product categories. The third column of Table 1.7 shows that the dispersion of prices is positively correlated with the dispersion of store-level price change frequencies. If products are less homogeneous within a product category, both price change frequencies and prices are more disperse. The positive correlation indicates that products may differ considerably within a product category.

	Correlation of price dispersion with	
	price change frequencies	**standard deviation of price change frequencies**
Unprocessed food	0.309	-0.090
Processed food	-0.181	0.132
Clothing	0.195	0.226
Durable goods	-0.375	0.274
Other goods	-0.137	0.118
Services	-0.196	0.241
Whole sample	**-0.304**	**0.242**

Table 1.7: Correlation between dispersion of prices and price change frequencies

1.4 Size of price changes

In Hungary, the (weighted) average size of price changes is 12.25 percent, and the (weighted) median is 11.68 percent. The average size of price increases is 11.15 percent (with a median of 10.85 percent), and the average size of price decreases is 13.62 percent (with a median of 12.76 percent). These numbers are somewhat larger than in the Euro area, where the average size of price increase and decrease is 8.2 percent and 10 percent (Dhyne et al., 2005). In the US, the average size of price changes is 14 percent (Klenow-Kryvtsov, 2007).

Similarly to the frequencies, the average size of price changes also differs considerably among products (Figure 1.5 in the Appendix). While the average size (across items) is 12.25 percent, the maximum is 30.6 percent for "natural medical therapy"and the minimum is 1.4 percent for "currency exchange".[15]

[15] We took into account only those products for which we had observations for the whole sample period. Figure 13 in the Appendix contains all representative items, irrespective of the periods of observation.

Some heterogeneity also prevails at the main product category level (Table 1.8), but it is less pronounced than in case of frequencies. Representative items in the Energy category – mostly fuels – experience small price changes, and the price of Clothes change by relatively large amounts, but the average size of price changes in the other CPI-categories mostly fluctuates between 10-15 percent. Although the importance of factors like menu costs, pricing strategies and cost factor volatility may differ for explaining the differences in frequencies and magnitude of price changes, it is natural to assume that the lower is the frequency of price change, the bigger is its magnitude. The size of price changes is the smallest for energy, and the highest is for clothing. Although on average price increases are more frequent than decreases, its effect on inflation is partly offset by the higher magnitude of price decreases.

	Size of change		**Size of increase**		**Size of decrease**	
CPI category	**Mean**	**Median**	**Mean**	**Median**	**Mean**	**Median**
Food, alcohol, tobacco	12.54	11.56	11.63	10.79	13.64	12.75
Unprocessed food	17.75	13.60	15.39	12.72	18.84	14.33
Processed food	11.39	11.23	10.80	10.58	12.49	12.59
Proc. food excl. alc, tob	12.11	12.13	11.36	11.11	13.39	14.05
Clothing	20.61	21.77	15.71	16.21	24.06	25.36
Durable goods	10.53	10.22	9.07	9.09	11.78	11.34
Other goods	12.39	12.33	11.10	10.83	14.41	14.08
Energy	3.82	2.92	3.81	2.86	3.82	3.02
Services	12.86	12.05	12.68	11.85	14.22	13.12
TOTAL	**12.25**	**11.68**	**11.15**	**10.85**	**13.62**	**12.76**

Table 1.8: Size of price changes by product categories

The correlation between the magnitude of price increases and decreases is strong (Figure 1.6 in Appendix), which can be a sign of symmetric product specific menu costs.

Similarly to the frequencies,we examined the seasonality of the size of price changes in the different product categories (Figure 1.7 in the Appendix). We tested seasonality also formally by regressing monthly average size of price increases (decreases) on monthly and VAT dummies (Table 1.9).[16] The magnitude of price changes is much less seasonal than the frequency: while we had seasonal variation in many product categories in the frequencies, we found seasonal variation in the sizes only in the clothing and unprocessed food categories. In the

[16]The numbers are differences from the average size of price increases (decreases) over the whole sample measured in percentage points. Blank cells indicate that the difference was not significant.

unprocessed food category the seasonality reflects the changes in the availability of fresh fruit and other similar representative items. In the Clothing category the size of price increases and decreases are both strongly seasonal, reflecting seasonal sales for these types of products.

	Average size of price increases											
	J	F	M	A	M	J	J	A	S	O	N	D
Unprocessed food												
Clothing			5.9						6.8			
	Average size of price decreases											
	J	F	M	A	M	J	J	A	S	O	N	D
Unprocessed food						2.8	2.3					
Clothing	6.8						5.7					

Table 1.9: Seasonality in the average size of price changes

1.5 Mean duration of price spells

We begin this section by reporting the duration distribution of observed price spells in all product categories (Figure 1.8). Similarly to many European countries, the mode of this distribution is at 1 month. However, there are two factors that bias our results towards shorter price spells: one the one hand, in our sample there are more spells from stores with shorter average durations, thus we oversample shorter spells. On the other hand, due to left and right censoring, we lose longer spells with larger probability, which again biases our observed distribution towards the shorter spells.[17]

Sectoral duration distributions (Figures 1.9–1.10 in the Appendix) are generally quite similar to the overall duration distribution. The directly observed distributions in the clothing, consumer durables and other goods sectors are almost identical, the only difference being perhaps the high frequency of durations of 2 months in case of the clothes. For the unprocessed food items, the proportion of spells with a duration of 1 month is more than 60 percent, while the same number in the processed food category is less than 40 percent. In the energy sector, more than 80 percent of the prices last for only 1 month, reflecting frequent changes in fuel prices. In terms of the shape of the observed duration distribution, the only exception is the services sector, where the mode is at 12 months, reflecting time-dependent pricing with price revisions mostly taking place in January.

[17] If we have $T = 67$ months of observations, then we have 66 "generations"of 1-month long spells, and only 60 "generations"oh 7-month long spells.

The second column of Table 1.10 contains the means of the observed duration distributions by CPI-categories. The average duration of price spells (when calculated directly from the observed duration distribution) is 6.44 months.[18] However, as we discussed before, these mean duration estimates are biased downwards because of two factors: censoring and heterogeneity.

There are plenty of possible ways to eliminate these bias in the direct mean duration estimations. The most frequently used method in the empirical literature is the frequency approach (Baudry et al. (2004), Veronese et al. (2005)). The frequency approach derives an implied mean duration measure by calculating the frequency of price change at product level, which can be inverted to obtain the mean duration of price spells. The product-level implied durations are aggregated up with using the CPI weights. Using this method no explicit treatment of censoring is called for, what explains its popularity. The third column of Table 1.8 shows that this implied mean duration estimate is approximately 9 months in Hungary. The same estimates for the whole CPI are 7.9 months in Israel (Baharad-Eden (2004)), 8.38 months in France (Baudry et al. (2004)), 9 months in Italy (Veronese et al. (2005), 9.71 months in the Netherlands (Jonker et al. (2004) and 11.8 months in Germany (Hoffmann–Kurz-Kim (2006)).

An alternative solution to correct for the potential bias is proposed by Baudry et al. (2004). The main result of their paper is that when they weight each spell by the inverse of the number of spells observed at the particular store, then the resulting weighted mean duration estimate will be robust to both censoring and cross-sectional heterogeneity. The fourth column of Table 1.10 reports the calculated mean duration estimates with this method (weighted duration). Note that the bias-corrected mean duration estimates are bigger in all CPI-categories than the direct mean duration estimates, supporting our earlier statement about the downward bias in the latter. The difference between the estimated mean durations in the second and fourth columns of Table 1.10 shows, that if we account for the downward bias caused by censoring and cross-sectional heterogeneity, then the estimated mean duration increases from 6.44 months to 8.03 months. The weighted mean duration is similar to the value calculated for France (7.24) by Baudry et al. (2004).

[18] This number is calculated by taking the weighted average of product level mean durations, with the weights being the product-specific CPI-weights.

CPI category	Direct est.	Freq. approach	Weighted duration
Food, alcohol, tobacco	5.54	6.83	6.76
Unprocessed food	2.57	2.85	3.29
Processed food	6.05	7.51	7.35
Clothing	6.83	13.00	9.59
Durable goods	6.68	11.52	9.18
Other goods	6.62	11.26	9.14
Energy	1.73	1.84	1.78
Services	11.27	16.86	13.35
TOTAL	**6.44**	**9.27**	**8.03**

Table 1.10: Mean duration of price spells by product categories

1.6 Inflation variation: decomposition into frequency and size effects

Given that the inflation rate equals the product of the frequency and the average size of price changes,[19] it is natural to ask whether it is changes in the frequencies or sizes that is behind the observed variation in the inflation rate.[20] In Hungary, this question is of particular interest, since the inflation variation is much bigger than in other developed countries (see Figure 1 in the Appendix): in our sample period of five and half years, inflation varied between 2.3 percent and 9 percent, including a period (April 2006–March 2007) when the yearly inflation rate increased by 6.7 percentage points within a mere 11 months.

In addition to its empirical relevance, this question is potentially important from a theoretical point of view as well. In the sticky prices literature models distinguish between two types of price rigidities: they assume either time-dependent pricing (TDP) or state-dependent pricing (SDP). In the TDP models, firms are given the opportunity to re-price exogenously, and this opportunity typically depends on the time elapsed from the last price change. In the SDP models, firms' re-pricing decisions are endogenous, and depend on the (either firm-specific or aggregate) shocks that hit them. Therefore in time-dependent models the frequency of price change does not vary, even after relatively large shocks. On the other hand, under state-dependent pricing, variation in the frequency of price changes (or the extensive margin) is an important channel of price adjustment after shocks. Therefore, by exploring the source of variation in the inflation rate

[19]A formal proof of this can be found e.g. in the Appendix of Hoffmann–Kurz-Kim (2006).

[20]In the previous section we used the term "average size"for the average *absolute* size of price changes. In this section we use the same term for the average sizes of price changes with negative sign for price decreases, and positive sign for price increases.

(i.e. whether it is variation in frequencies or not) we may be able to decide which model family is more relevant empirically.

To judge whether it is volatility in frequencies or average sizes behind the volatility of inflation rates, first it may be useful to have a look at the time-series of monthly inflation rates, price change frequencies and average sizes. Figure 1.11 in the Appendix indicates that the average size of price changes exhibits stronger co-movement with the inflation rate than the frequency of price changes; but as we shall see, a deeper analysis is necessary to give a definitive answer to this question.

We use the decomposition of Klenow-Kryvtsov (2007) to more formally analyze the driving forces of the inflation variation. Klenow and Kryvtsov show that if we take the first-order Taylor series expansion of the identity $\pi_t = fr_t \cdot dp_t$ (where π_t, fr_t and dp_t are inflation, frequency and average size of price change at time t) around the sample means $\overline{fr}$ and $\overline{dp}$, then we can express the inflation volatility as

$$var(\pi_t) = var(dp_t) \cdot \overline{fr}^2 + var(fr_t) \cdot \overline{dp}^2 + 2 \cdot \overline{fr} \cdot \overline{dp} \cdot cov\,(fr_t, dp_t) + O_t. \quad (1.1)$$

In the right-hand side, the first term $(var(dp_t) \cdot \overline{fr}^2)$ stands for the inflation volatility in the intensive margin (i.e. same frequency, varying size of price adjustments), and captures all the inflation volatility in TDP models (where fr_t is constant over time). So the fraction $var(dp_t) \cdot \overline{fr}^2 / var(\pi_t)$ (the "time-dependent part"of the inflation variation) reflects how closely are the TDP models to empirically observed inflation variations.

Table 1.11 contains this time-dependent fraction $var(dp_t) \cdot \overline{fr}^2 / var(\pi_t)$ of inflation volatility for each product categories separately. Having seen Figure 1.11, perhaps it is not surprising that when calculated for all product categories, the time-dependent part accounts for 75.2 percent of all inflation variation. This number is somewhat smaller than what Klenow–Kryvtsov (2007) found for the US: they reported the time-dependent part to be between 86 percent and 113 percent (depending on the sample used).

As Table 1.11 makes it clear, there is huge sectoral heterogeneity in both the inflation volatility, and the time-dependent part of it. The volatility of inflation itself shows high variation across product categories: the volatility is biggest – perhaps not surprisingly – for unprocessed foods, clothing and energy items. The time-dependent part of the volatility is bigger than average for unprocessed food, durable goods, other goods and energy items, reflecting relatively constant price change frequencies (over time) for these. While this is true, this constant frequency is huge for unprocessed food and energy, and it is relatively low for the durable goods and other non-industrial goods.

At the other extreme, the time-dependent part of the inflation volatility is

CPI category	$var(\pi_t)$	Time-dependent part of $var(\pi_t)$	TDP %
Food, alcohol, tobacco	0.636	0.427	67.1%
Unprocessed food	7.858	7.200	91.6%
Processed food	0.330	0.122	37.0%
Proc. food excl. alc, tob	0.459	0.156	34.0%
Clothing	2.380	1.470	61.7%
Durable goods	0.082	0.070	85.1%
Other goods	0.149	0.120	80.7%
Energy	3.950	3.043	77.1%
Services	0.365	0.069	18.9%
TOTAL	**0.235**	**0.177**	**75.2%**

Table 1.11: Decomposition of inflation volatility by main CPI categories

relatively low for clothing, and extremely low for processed food and services. This would indicate state-dependence and adjustment to shocks on the extensive margin (i.e. through volatile adjustment frequencies). However – as we discussed previously –, the source of frequency volatility in these categories is mainly seasonal variation. The fact that the decomposition of equation 1.1 cannot distinguish between state-dependence and simple seasonal variation, is a big drawback of this method.[21]

Results in Table 1.11 indicate that it is predominantly the size of price changes that drives inflation, and not the frequency. But the fact that the price change frequency is relatively stable and independent from the inflation rate may arise simply because when the inflation rate increases, the increasing price increase frequencies and the decreasing price decrease frequencies offset each other. To investigate this possibility (and still following Klenow–Kryvtsov 2007) we also analyzed a more precise decomposition of the inflation rate:

$$\pi_t = fr_t^+ \cdot dp_t^+ + fr_t^- \cdot dp_t^-, \tag{1.2}$$

where fr_t^+ and fr_t^- are the frequencies of price increases and decreases, and dp_t^+ and dp_t^- are the average sizes of price increases and decreases.[22] We also introduce $pos_t = fr_t^+ \cdot dp_t^+$ and $neg_t = fr_t^- \cdot dp_t^-$, which are the contributions of

[21] Nevertheless, Klenow–Kryvtsov (2007) use this decomposition, without discussing the effects that seasonal variation may have on the results.

[22] Our sign convention is that dp_t^+ is positive by construction, and dp_t^- is negative by construction. Of course, the frequencies of price increases and decreases are always non-negative by definition.

price increases and decreases to the overall inflation rate, respectively.[23]

Having calculated the terms in equation 1.2, in Table 1.12 we report the simple correlation coefficients of these terms with the inflation rate. In columns 1-2 of Table 1.12 we report the correlations between $fr_t = fr_t^+ + fr_t^-$ and π_t, and between dp_t and π_t. These correlation coefficients are in line with our previous findings: if we use the simple inflation decomposition of $\pi_t = fr_t \cdot dp_t$, then we find that it is mostly the average size of price changes (and not the frequency) that drives the inflation rate.

CPI category	fr_t	dp_t	fr_t^+	fr_t^-	dp_t^+	dp_t^-	pos_t	neg_t
Food, alcohol, tobacco	0.71	0.94	0.81	-0.65	-0.02	0.56	0.89	0.73
Unprocessed food	-0.04	0.99	0.43	-0.77	0.45	0.70	0.67	0.90
Processed food	0.89	0.89	0.95	-0.43	-0.31	0.24	0.96	0.51
Proc. food excl. alc, tob	0.89	0.86	0.94	-0.44	-0.51	0.26	0.96	0.50
Clothing	-0.63	0.96	0.71	-0.88	0.50	0.69	0.68	0.95
Durable goods	-0.39	0.90	0.36	-0.52	-0.22	0.48	0.12	0.86
Other goods	0.15	0.93	0.65	-0.52	0.22	0.35	0.81	0.77
Energy	-0.02	0.99	0.89	-0.90	-0.10	0.10	0.85	0.87
Services	0.91	0.60	0.95	0.19	0.09	-0.22	0.99	-0.34
TOTAL	**0.61**	**0.96**	**0.78**	**-0.30**	**0.04**	**0.16**	**0.88**	**0.60**

Table 1.12: Correlation of terms in equation 1.2 with the inflation rate

In columns 3-6 of Table 1.12, however, we see that variations in the frequencies *do have* an important role in inflation volatility. The frequency of price increases is always strongly positively correlated with the inflation rate, and the frequency of price decreases is almost always negatively correlated with the inflation rate.[24] Obviously, the overall frequency (in column 1) did not show extraordinarily high correlation with the inflation rate because these two effects indeed offset each other. On the other hand, average size of price increases and decreases do not show strong co-movement with the inflation rate. We may have found strong correlation (in column 2) for the average size of *all* price changes because the size of both increases and decreases tends to have a positive correlation with the overall inflation rate.[25]

[23] Our sign convention then implies that pos_t is always positive, and neg_t is always negative.

[24] The only exception is services, where price decreases are very rare, and their frequency is apparently independent from the inflation rate.

[25] Again, our sign convention implies that whenever π_t increases, dp_t^- becomes a smaller negative number (in absolute terms), so the correlation between π_t and dp_t^- is in fact positive.

Finally, columns 7-8 of Table 1.12 are informative about the relative importance of price increases ($pos_t = fr_t^+ \cdot dp_t^+$) and decreases ($neg_t = fr_t^- \cdot dp_t^-$) in the inflation variation. According to the results, we can assert that price increases seem to be more important from the point of view of inflation than price decreases. Important exceptions are unprocessed food, clothing and durable goods; these are those product categories in which the measured inflation rate in our data set is negative. Therefore it is more appropriate to say that when the inflation rate is positive (negative), then the price increases (decreases) are the primary sources of inflation variation.

1.7 Inflation effects of Value Added Tax changes

Relative to other countries, there have been quite frequent changes in the general VAT-rates in Hungary during the sample period (Table 1.3). In January 2004, the middle and lower rates increased from 12 to 15 percent and from 0 to 5 percent, respectively, while in January 2006 the top rate was reduced from 25 to 20 percent. Finally, from September 2006, the middle rate was again increased to 20 percent, meaning that the middle and top VAT-rates were unified.

In this section we analyze the inflation effects of these VAT-changes.

1.7.1 Data

This analysis is carried out on a narrower data set than what was used in the previous sections: out of the 770 representative items with a CPI-weight of 70.122 percent in our original sample (Table 1), we dropped further 220. So the data set in this section contains only 550 representative items with a CPI-weight of 45.346 percent in 2006. The items that were dropped include fuels, alcoholic beverages and tobacco products, where highly volatile external factors (like world oil prices) and extremely frequent changes in indirect taxes (which are similar in effect to VAT-changes) make it difficult to identify the effects of VAT-changes separately. A second main reason of exclusion was that the maximum length of price spells was low: these are the products for which the CSO began data collection towards the end of the sample period,[26] rendering the estimates about the inflation effects of VAT-changes to be unreliable. Finally, we also dropped those representative items from the data set whose prices are only observed in certain months of the year.[27] Missing data makes it difficult to estimate VAT-effects, as price observations may not be available in the months of VAT-changes.

[26] Examples are LCD TV-s, MP3 players, memory cards.

[27] Examples of these are gloves, cherries, etc.

The coverage of the data set that is used in this section (labelled as "VAT-sample") is in Table 1.13.[28] In the VAT-sample, the number of price observations is 3,950,962, the number of store-product cells 178,534, so the average length of the price trajectories is 22.1 months. In the original sample we had 4,684,289 observations, 272,549 price trajectories with an average length of 17.2 months.

	CPI-basket		**Original sample**		**VAT-sample**	
Category	**Weight**	**Items**	**Weight**	**Items**	**Weight**	**Items**
Unprocessed food	5.665	53	5.665	53	4.151	34
Processed food	19.865	140	19.865	140	12.723	112
Energy (oil) products	13.203	16	6.350	8	0.723	1
Non-energy ind. goods	27.077	497	20.515	436	14.562	309
Services	34.190	190	17.726	133	13.187	94
TOTAL	**100.000**	**896**	**70.122**	**770**	**45.346**	**550**

Table 1.13: Coverage of the data set in VAT-effect calculations

Of course, the VAT-sample is not representative: the frequency of price increases and decreases are 10.2 percent and 6.2 percent, lower than in the original sample (12.8 percent and 8.8 percent). Also, the average size of price increases and decreases is also significantly different: 11.7 percent and 14.2 percent, higher than in the original sample (11.2 percent and 13.6 percent). Nevertheless, dropping some representative items is necessary to obtain reliable estimates for the inflation effects of VAT-changes.

1.7.2 Methodology

To study the inflation effects of VAT-changes, we decompose changes in inflation to changes in frequencies and average sizes of price changes. We do this both for the affected and non-affected products. Similarly to the previous section, we depart from the following inflation decomposition equation (for details, see Hoffmann and Kurz-Kim 2006):

Average price change = Inflation =
= (Proportion of price increasing stores) * (Average size of price increases)
– (Proportion of price decreasing stores) * (Average size of price decreases)
Formally:

$$\pi = p^{+}\mu^{+} - p^{-}\mu^{-}, \tag{1.3}$$

[28] There is a single Energy item remaining in the data set (propan butan gas). In this section this will be included into the Non-energy industrial goods category.

where π denotes inflation, p^+ and p^- are the proportions of price increasing and decreasing stores, and μ^+ and μ^- are the average sizes of price increases and decreases.

To quantify the inflation effect of VAT-increases, suppose that price increase and decrease frequencies are p^{+V} and p^{-V} when there is a specific VAT-increase, and p^+ and p^- would be the same frequencies in the absence of any VAT-change. Similarly, with the specific VAT-increase investigated, the average size of price increases and decreases are μ^{+V} and μ^{-V}, otherwise they would be μ^+ and μ^-. Then if there is a VAT-increase, the inflation rate is $p^{+V}\mu^{+V} - p^{-V}\mu^{-V}$, whereas it would be $p^+\mu^+ - p^-\mu^-$ in the absence of the VAT-increase. So the overall inflation effect of the VAT-increase is

$$\left(p^{+V}\mu^{+V} - p^+\mu^+\right) - \left(p^{-V}\mu^{-V} - p^-\mu^-\right). \tag{1.4}$$

In this expression the first term is the inflation effect due to higher willingness to increase prices, and the second term is the inflation effect due to lower willingness to decrease prices.[29] We use this equation to quantitatively estimate the inflation effects of VAT-changes.

We do this by setting up a latent-variable selection model on the sizes of store-level price increases and decreases. The advantage of this method is that this way we can use the full panel in estimation.

Suppose, first, that we would like to estimate the first term of equation 1.4, i.e. the inflation effect through price increases. Then our first goal is to estimate a regression on the size of price increases, and evaluate the marginal effect of various VAT-change dummies on the expected value of the size of price increases. Let us denote the size of the desired price increase for firm i at time t by $Y_{1,it}$, and the explanatory variables by X_{it}. The regression that we want to estimate is

$$Y_{1,it} = \beta' X_{it} + U_{1,it}. \tag{1.5}$$

The problem is, however, that we do not always observe $Y_{1,it}$: stores do not always adjust prices, and in fact we only observe the desired size of price increases whenever there is a price increase. If we define the latent variable $Y_{2,it}$ as the variable that determines whether there is a price increase or not, then we can write

$$Y_{2,it} = \gamma' Z_{it} + U_{2,it}, \tag{1.6}$$

where Z_{it} are the determining factors of whether there is a price increase or

[29]Note that if the frequency of price decreases declines when there is a VAT-increase (and the average size of decreases does not change), then the second term is negative, meaning a positive inflation effect.

not. We can then say that $Y_{1,it}$ in equation 1.5 is observed whenever $Y_{2,it} > 0$ in equation 1.6.

With appropriate distributional assumptions on $U_{1,it}$ and $U_{2,it}$,[30] we can estimate these equations with maximum likelihood. In particular, we can express the conditional expected value of price increases by

$$E(Y_{1,it} \mid Y_{2,it} > 0, X_{it}) = \beta' X_{it} + \sigma\rho E\left[\frac{f(\gamma' Z_{it})}{F(\gamma' Z_{it})} \mid X_{it}\right], \tag{1.7}$$

where $f(\cdot)$ and $F(\cdot)$ are the pdf and cdf of a standard normally distributed random variable. Then in the first step, we can directly estimate p^{+V} and p^{+} in equation 1.4 from equation 1.6, by estimating a probit model and calculating the probabilities of $Y_{2,it}$ being positive for both with and without a VAT-change. (See also Rátfai (2006).) Then in the second step we estimate μ^{+V} and μ^{+} in equation 1.4 using equation 1.7 (i.e. equation 1.5 augmented with an extra term estimated in the first step, according to equation 1.7).

A similar model can be estimated for the average size of price decreases to estimate the terms in the second bracket in equation 1.4.

After these equations are estimated for each item separately and the product-level inflation-effects are calculated, we obtain the aggregate inflation effect by aggregating item-specific effects. We use the CPI-weights to calculate inflation effects for higher levels of aggregation (i.e. for the level of the whole VAT-sample, or the consumption categories).

1.7.3 Inflation effects of the 2004 January and 2006 September VAT-increases

Effective from January 2004, the middle VAT-rate of 12 percent increased to 15 percent, and from September 2006, the same middle VAT-rate further increased to 20 percent. These measures affected 170 representative items in the VAT-sample, with an overall CPI-weight of 21.250 percent (Table 1.14). These items mostly belong to one of the Food categories: there are 121 affected food items with a total CPI-weight of 14.343 percent.

Figures 1.12-1.17 depict the time series of the frequency and size of price changes, increases and decreases among the items affected and not affected by these VAT-increases. Figure 1.12 shows that the frequency of price increases exploded in January 2004 and September 2006: from the usual level of 10-20

[30] More precisely, the distributional assumption is $\binom{U_1}{U_2} \sim N\left[\binom{0}{0}, \Sigma\right]$, with $\Sigma = \begin{bmatrix} \sigma^2 & \rho\sigma \\ \rho\sigma & 1 \end{bmatrix}$.

Category	Affected items CPI-Weight	Affected items Items	Not affected items CPI-Weight	Not affected items Items
Unprocessed food	4.151	34	0.000	0
Processed food	10.192	87	2.531	25
Non-energy ind. goods	1.561	17	13.724	293
Services	5.347	32	7.841	62
TOTAL	**21.250**	**170**	**24.096**	**380**

Table 1.14: Products affected by the 2004 January and 2006 September VAT-increases

percent, it jumped to above 50 percent. At the same time, there is no obvious effect on the frequency of price decreases. Moreover, according to Figure 1.13, average sizes of price changes decline in the months of the VAT-increases, and this decline can be attributed to the decreasing average size of price increases. This may be because the size of VAT-increases was small relative to the usual average size of price increases. In the subgroup of products not affected by the VAT increases, we also see a moderate increase in the price increases frequencies (from the usual 5-10 percent to 12-15 percent), and a slight decrease in the average size of price *increases*.

To numerically estimate the inflation effects of the VAT-changes, we use the decomposition equation 1.4 by estimating equations 1.6-1.7. Table 1.15 reports the estimated inflation effects of the 2004 January and the 2006 September VAT-increases by main consumption categories in the affected sub-sample. The overall inflation effect of the 3 percentage point increase in 2004 January is 2.13 percent, and of the 5 percentage point increase in 2006 September is 3.71 percent. These are smaller than what would be implied by "automatic"pass-through: 2.68 percent (115/112-1) in 2004 January and 4.35 percent (120/115-1) in 2006 September.[31]

For both VAT-increases, the highest inflation effect can be observed in the (unprocessed and processed) food categories: in both cases we find almost complete pass-through of the VAT-increase into the prices. Inflation effects are somewhat smaller, but still substantial for the non-energy industrial goods and services categories.

Comparing the two VAT-increases, the inflation effect is always larger for the 2006 September VAT-increase, when the increase itself was larger. The effects are

[31] Then the implied pass-through of a unit VAT-increase is 79 percent and 85 percent in 2004 January and 2006 September, respectively.

Category	Weight	Inflation effect	Through increase	Through decrease
2004 January				
Unprocessed food	4.151	2.20	1.94	0.26
Processed food	10.192	2.58	2.49	0.09
Non-energy ind. goods	1.561	1.43	1.69	-0.26
Services	5.347	1.40	1.44	0.04
TOTAL	**21.250**	**2.13**	**2.06**	**0.06**
2006 September				
Unprocessed food	4.151	4.17	3.86	0.31
Processed food	10.192	3.71	3.73	-0.02
Non-energy ind. goods	1.561	3.08	3.01	0.07
Services	5.347	3.52	3.57	-0.05
TOTAL	**21.250**	**3.71**	**3.66**	**0.04**

Table 1.15: Inflation effect of the 2004 January and 2006 September VAT-increases by main consumption categories, *affected products*

however non-linear (i.e. more than double for a less than double VAT-increase) in the services and the non-energy industrial goods categories.

If we decompose the overall inflation effect to the inflation effect of higher willingness to increase prices and of lower willingness to decrease prices (columns 4-5 in Table 1.15), we see that the vast majority of the total inflation effect of the VAT-increases is through the higher willingness of stores to increase prices, whereas there is not much effect in the price decrease side. In sum, we find that most of the inflation effect is through the primary channel.[32]

If there are close substitutes between products affected and not affected by the VAT-increases, we may expect that these tax increases also affect the prices of those products that are not directly affected. We therefore investigate the inflation effect also for the not affected sub-sample. The results of these are summarized in Table 1.16.

The overall inflation effect among the non-affected items is positive: it is 0.42 percent in 2004 January and 0.75 percent in 2006 September on the aggregate level, and they are almost always positive in the main consumption categories. Also, we estimate the largest inflation effects in those product categories (processed food, services) where the distribution of affected and non-affected products is relatively even. This may be consistent with our hypothesis that the

[32] We use the term "primary channel" as the price increase side in case of a VAT-increase, and the price decrease side in case of a VAT-decrease.

Category	Weight	Inflation effect	Through increase	decrease
2004 January				
Unprocessed food	0.000	n.a.	n.a.	n.a.
Processed food	2.531	0.85	1.11	-0.26
Non-energy ind. goods	13.724	0.23	0.13	0.10
Services	7.841	0.62	0.64	-0.02
TOTAL	**24.096**	**0.42**	**0.40**	**0.03**
2006 September				
Unprocessed food	0.000	–	–	–
Processed food	2.531	0.81	0.91	-0.11
Non-energy ind. goods	13.724	0.45	0.45	0.01
Services	7.841	1.26	1.24	0.02
TOTAL	**24.096**	**0.75**	**0.76**	**0.00**

Table 1.16: Inflation effect of the 2004 January and 2006 September VAT-increases by main consumption categories, *not affected products*

relative price of close substitutes may depend on *relative* tax rates, which also change for those products whose *absolute* tax rates did not change.

A further decomposition of this positive overall effects to effects through price increases and decreases reveals that out of the total effect of 0.42 and 0.75 percent, 0.40 and 0.76 percent are through the larger willingness of stores to increase their prices, so again the primary channel is far more important than the secondary channel.

1.7.4 Inflation effect of the 2006 January VAT-decrease

In January 2006, the top VAT-rate decreased from 25 percent to 20 percent. Though this affected a different set of products than the VAT-increases just discussed, still it may be interesting to compare the inflation effects of this VAT-decrease with the inflation effects of the VAT-increases. The comparison is perhaps better with the 2006 September VAT-increase, since there is only 8 months between these VAT-changes, and the size of the change (5 percentage points in both cases) is also similar.

The VAT-decrease mostly affected the non-energy industrial goods, but we can also find several food and services items in the affected sub-sample (Table 1.17).

Figures 1.18-1.23 depict the time series of the frequencies and average sizes of price changes among products affected and not affected by the VAT-decrease. In the subgroup of products affected by the VAT decrease the frequency of price

Category	Affected items CPI-Weight	Affected items Items	Not affected items CPI-Weight	Not affected items Items
Unprocessed food	0.000	0	4.151	34
Processed food	2.531	25	10.192	87
Non-energy ind. goods	13.724	293	1.561	17
Services	6.877	52	6.310	42
TOTAL	**23.132**	**370**	**22.214**	**180**

Table 1.17: Products affected by the 2006 January VAT-decrease

changes rises, and the average size of price changes falls in the month of the VAT-decrease. Figure 1.20 shows that the increased frequency is due to the increase in the price *decrease* frequencies, while the price *increase* frequency is apparently not smaller than in any regular January. As far as sizes are concerned, according to Figure 1.22, the decrease in average size of price changes is caused by a huge decrease in the average size of price *decreases*, while the average size of price *increases* does not seem to be different from other months.

In the *non-affected sub-sample* the frequency of price changes is the same as usual . The average size of price *decreases*, however, is a bit lower in the non-affected sub-sample, a similar phenomenon to what we find in the affected sample.

To quantitatively estimate the inflation effect of the VAT-decrease, we again use the decomposition equation 1.4, and estimate it with equations 1.6-1.7. The results show that the average inflation effect of the 5 percentage point VAT-decrease is -1.08 percent (Table 1.18). This is much smaller than what the effect would be if all stores automatically decreased their prices according to the tax change, -4 percent (120/125-1). Moreover, this inflation effect is also much smaller in absolute terms than the one estimated for a similar tax increase. In 2006 September, although in a different set of products, the inflation effect of a similar VAT-increase was 3.71 percent.

Category	Weight	Inflation effect	Through increase	Through decrease
Unprocessed food	0.000	–	–	n.a.
Processed food	2.531	-2.10	-0.56	-1.53
Non-energy ind. goods	13.724	-1.46	-0.06	-1.40
Services	6.877	0.05	0.28	-0.23
TOTAL	**23.132**	**-1.08**	**-0.02**	**-1.06**

Table 1.18: Inflation effect of the 2006 January VAT-decrease by main consumption categories, *affected products*

The largest effect is in the processed food category, where it is estimated to be -2.10 percent. A similar, moderately large inflation effect is measured in non-energy industrial goods, while apparently there is no inflation effect at all in services.[33]

Decomposing this inflation effect to "primary"and "secondary"sources, i.e. inflation through the increased willingness to decrease prices and through decreased willingness to increase prices, we see that much of the effect is through the primary channel. This finding seems robust for VAT-increases and decreases, and can partly explain the very small inflation effect in the services category.

Finally, we estimate the inflation effect of the 2006 January VAT-decrease for those items that were not affected by this VAT-change (Table 1.19). Here the estimated effects are again non-negligible: the average effect is -0.70 percent, with relatively high effects in the Processed food and Services categories, i.e. where the number of close substitutes across the affected and non-affected sub-samples is the highest. This gives further support to our hypothesis that in close substitutes the relative tax rate may matter. Interestingly, in this non-affected sub-sample the secondary channel, i.e. lower willingness to increase prices is the one that is more important.

			Through	
Category	**Weight**	**Inflation effect**	**increase**	**decrease**
Unprocessed food	4.151	-0.78	-0.45	-0.33
Processed food	10.192	-0.57	-0.33	-0.24
Non-energy ind. goods	1.561	-0.30	0.09	-0.39
Services	6.310	-0.75	-0.71	-0.04
TOTAL	**22.214**	**-0.70**	**-0.47**	**-0.23**

Table 1.19: Inflation effect of the 2006 January VAT-decrease by main consumption categories, *not affected products*

1.7.5 Longer-run inflation effects of the tax changes

So far we have studied the immediate (i.e. 1-month) inflation effects of VAT-changes. The method we applied, however, is appropriate to estimate the inflation effects of the tax changes over longer horizons as well. For example, if we want to estimate the extra inflation a VAT-increase caused within the first two months, we have to do the same analysis in a bimonthly data set. Here the frequency of

[33] There can be several reasons for this. First, the VAT-effect is mainly through the "primary channel", which means that most of the VAT-decrease should affect prices through price decreases. But as there are basically no price decreases in the services category, this channel may be weak. Another explanation can be that the VAT-decrease in 2006 January coincided with an almost 10 percent increase in the minimum wage, which could hit badly the unskilled labor-intensive Services sector, preventing it from significant price reductions.

price changes should be understood as the fraction of stores changing its price in any two-month period, and the average absolute size of the price changes is defined accordingly. Tables 1.20-1.21 contain the longer-run inflation effects of the 2006 September VAT-increase and the 2006 September VAT-increase.[34]

CPI category	-2m	-1m	1m	2m	3m	4m
Unprocessed food	1.12	0.93	4.17	2.30	0.33	-1.88
Processed food	1.94	0.84	3.83	3.70	3.81	2.65
Non-energy ind. goods	0.64	1.05	3.12	1.80	0.19	0.52
Services	0.03	0.09	2.81	3.37	3.11	3.26
TOTAL	**1.43**	**0.77**	**3.71**	**3.24**	**2.77**	**1.69**

Table 1.20: Longer-run inflation effect of the 2006 September VAT-increase by main CPI-categories, affected products

CPI category	-2m	-1m	1m	2m	3m	4m
Unprocessed food	0.36	0.19	-1.23	-1.30	-1.22	-1.50
Processed food	-0.17	0.01	-1.63	-2.42	-3.30	-3.36
Non-energy ind. goods	0.18	-0.11	-1.36	-1.92	-2.19	-2.58
Services	0.19	0.04	0.11	-0.46	-1.87	-1.56
TOTAL	**0.05**	**-0.04**	**-1.08**	**-1.62**	**-2.22**	**-2.47**

Table 1.21: Longer-run inflation effect of the 2006 January VAT-decrease by main CPI-categories, affected products

Results indicate that in case of a tax increase, we have almost immediate inflation effect, and not much additional effect in the following 3 months. In contrast, when the VAT-rate decreases, its inflation effect is more gradual: the cumulative effect grows over time. As a result, asymmetry in the inflation effect decreases over time.

As before, we can decompose the overall inflation effect to effect through price increases and price decreases. Tables 1.22-1.23 contain this decomposition, together with the effects on price increase and price decrease frequencies.[35] Results indicate that in case of tax increases, the immediate inflation effect is almost exclusively through price increases (in particular, the increased frequency of price increases). But when the tax rate decreases, the immediate effect is through

[34] Results for 2004 January are available upon request. The overall inflation effect (in all CPI categories) after after 1, 2 and 4 months is 2.14, 2.55 and 2.01 percent , respectively.

[35] Tables 22-23 only contain the aggregate figures. However, results are robust across sectors; these sectoral results are available upon request from the authors.

price decreases (the frequency of price decreases jumps), and the later effect is through price increases (the frequency of price increases goes down, i.e. stores refrain from their "usual"price increases).

	-2m	**-1m**	**1m**	**2m**	**3m**	**4m**
Inflation effect	1.43	0.77	3.71	3.24	2.77	1.69
through increases	*1.46*	*0.68*	*3.66*	*3.22*	*2.81*	*1.77*
through decreases	*-0.03*	*0.09*	*0.04*	*0.03*	*-0.03*	*-0.07*
Effect on price increase frequencies	0.13	0.08	0.42	0.37	0.32	0.22
Effect on price decrease frequencies	-0.02	0.01	0.00	0.00	0.01	0.02

Table 1.22: Effect of the 2006 September VAT-increase through price increases and decreases, and on price change frequencies

	-2m	**-1m**	**1m**	**2m**	**3m**	**4m**
Inflation effect	0.05	-0.04	-1.08	-1.62	-2.22	-2.47
through increases	*0.15*	*0.05*	*-0.02*	*-0.42*	*-1.04*	*-1.23*
through decreases	*-0.09*	*-0.08*	*-1.06*	*-1.20*	*-1.18*	*-1.24*
Effect on price increase frequencies	0.01	0.00	0.00	-0.04	-0.08	-0.10
Effect on price decrease frequencies	0.02	0.02	0.17	0.19	0.18	0.17

Table 1.23: Effect of the 2006 January VAT-decrease through price increases and decreases, and on price change frequencies

Tables 1.20-1.21 also contain information about the initial inflation effect of tax changes (i.e. 1-2 months *before* the actual change happened). In case of the tax decrease, clearly there was no inflation effect before the tax change (not even within the different CPI categories). In contrast, before the tax increase we see some positive inflation effect before the actual change happened (0.77 percent in 2006 August, 1.43 percent in 2006 July-August). This result, however, is not robust across CPI categories: in fact more than 2/3 of the overall effect is caused of four individual products (bread, pastry, sugar, cold meat).[36]

1.8 Summary

This paper analyzes price setting in Hungary. First it provides descriptive statistics about frequencies and sizes of price changes and mean duration of prices. Then it decomposes inflation variation to variations in frequencies and sizes. Finally, it estimates the inflation effects of three Value Added Tax changes.

[36] We do not see any systematic effect before the 2004 January tax change either.

The results are the following:

The frequency of price changes in Hungary is 21.5 percent. Prices change more frequently than in the Euro area, and less frequently than in the US. About 60 percent of these price changes are price increases. Thus prices are not rigid downwards. The exception is Services, where only 12 percent of price changes is a price decrease. The average size of price changes is 12.3 percent. This is larger than in the Euro area, but smaller than in the US. Similarly to other countries, the average size of price increases (11.2 percent) is smaller than the average size of price decreases (13.6 percent).The mean duration of price spells is approximately 8 months, that is somewhat smaller than the implied mean duration in the Euro area, and similar to mean duration estimates in the US.

A decomposition of the inflation variation to price increase frequencies and price increase sizes, and to price decrease frequencies and price decrease sizes reveals that inflation variation is mostly driven by price increase and price decrease frequencies. The overall frequency of price changes, however, is rather stable and is unrelated to the inflation rate. This would be consistent with the Calvo-type sticky price model, but adjustments to VAT-shocks show just the opposite.

Adjustment to VAT-shocks takes place mostly on the extensive margin (i.e. frequencies). The short-term inflation effect of a unit VAT-increase is estimated to be substantially larger than that of a unit VAT-decrease. The inflation effect of a 3 percentage point VAT-increase in January 2004 among the affected products is estimated to be 2.13 percent, and of a similar 5 percentage point VAT-increase in September 2006 is 3.71 percent. The inflation effect of a 5 percentage point VAT-decrease in January 2006 is -1.08 percent in affected group. VAT-changes also affect the price level of those products which are not affected by them.

Bibliography

[1] Álvarez, Luis J. and Ignacio Hernando (2004): "Price Setting Behavior in Spain: Stylized Facts Using Consumer Price Micro Data." *ECB Working Paper* 416 (November 2004).

[2] Aucremanne, Luc and Emmanuel Dhyne (2004): "How Frequently Do Prices Change? Evidence Based on the Micro Data Underlying the Belgian CPI." *ECB Working Paper* 331 (April 2004).

[3] Baudry, Laurent, Hervé Le Bihan, Patrick Sevestre and Sylvie Tarrieu (2004): "Price Rigidity - Evidence from Consumer Price Micro-Data." *ECB Working Paper* 384 (August 2004).

[4] Bils, Mark and Peter Klenow (2004): "Some Evidence on the Importance of Sticky Prices." *Journal of Political Economy*, 2004, Vol. 112, No. 5., pp. 947-985.

[5] Calvo, Guillermo A. (1983): "Staggered Prices in a Utility-Maximizing Framework." *Journal of Monetary Economics*, Vol. 12 (1983), pp. 383-98.

[6] Dhyne, Emmanuel, Luis J. Alvarez, Marco M. Hoeberichts, Claudia Kwapil, Hervé Le Bihan, Patrick Lünnemann, Fernando Martins, Roberto Sabbatini, Harald Stahl, Philip Vermeulen and Jouko Vilmunen (2005): "Sticky Prices in the Euro Area. A Summary of New Micro Evidence." *ECB Working Paper* 563 (December 2005).

[7] Dhyne, Emmanuel, Luis J. Alvarez, Hervé Le Bihan, Giovanni Veronese, Daniel Dias, Johannes Hoffmann, Nicole Jonker, Patrick Lünnemann, Fabio Rumler and Juoko Vilmulen (2006): "Price Changes in the Euro Area and the United States: Some Facts from Individual Consumer Price Data." *Journal of Economic Perspectives*, Vol. 20, No. 2 (Spring 2006), pp. 171-92.

[8] Dias, Mónica, Daniel Dias and Pedro D. Neves (2004): "Stylized Features of Price Setting in Portugal: 1992-2001." *ECB Working Paper* 332 (April 2004).

[9] Dotsey, Michael, Robert King and Alexander Wolman (1999): "State-Dependent Pricing and the General Equilibrium Dynamics of Money and Output." *Quarterly Journal of Economics*, Vol. 114 (1999), pp. 655-90.

[10] Gabriel, Peter and Reiff, Adam (2007): "Estimating the Extent of Price Stickiness in Hungary: a Hazard-Based Approach." *Unpublished paper* (September 2007).

[11] Gabriel, Peter and Reiff, Adam, 2010. "Price setting in Hungary-a store-level analysis," Managerial and Decision Economics, John Wiley & Sons, Ltd., vol. 31(2-3), pages 161-176.

[12] Gagnon, Etienne (2007): "Price Setting During High and Low Inflation: Evidence from Mexico." *International Finance Discussion Papers*, Board of Governors of the Federal Reserve System, No. 896 (2007).

[13] Golosov, Mikhail and Robert E. Lucas Jr. (2007): "Menu Costs and Phillips Curves" *Journal of Political Economy*, Vol. 115, No. 2 (April 2007), pp. 171-99.

[14] Hoffmann, Johannes and Jeong-Ryeol Kurz-Kim (2006): "Consumer Price Adjustment under the Microscope: Germany in a Period of Low Inflation." *ECB Working Paper* 652 (July 2006).

[15] Jonker, Nicole, Carsten Folkertsma and Harry Blijenberg (2004): "An Empirical Analysis of the Price-Setting Behavior in the Netherlands in the Period 1998-2003 Using Micro Data." *ECB Working Paper* 413 (November 2004).

[16] Klenow, Peter and Oleksiy Kryvtsov (2008): "State-Dependent or Time-Dependent Pricing: Does It Matter for Recent U.S. Inflation?" *Quarterly Journal of Economics*, Vol. 123, No. 3 (August 2008), pp. 863-904.

[17] Nakamura, Emi and Jón Steinsson (2008): "Five Facts About Prices: A Reevaluation of Menu Cost Models" *Quarterly Journal of Economics*, Vol. 123, No. 4 (November 2008), pp. 1415-64.

[18] Rátfai, Attila (2006): "Linking Individual and Aggregate Price Changes." *Journal of Money, Credit and Banking*, Vol. 38, No. 8 (December 2006), pp. 2199-2224.

[19] Rátfai, Attila (2007): "The Frequency and Size of Price Adjustment: Microeconomic Evidence." *Managerial and Decision Economics*, Vol. 28 (2007), pp. 751-62.

[20] Rodriguez, J., F. Haraldsen (2006), The Use of Scanner Data in the Norwegian CPI: The New Index for Food and Non Alcoholic Beverages. Economic Survey 4: 21-28.

[21] Sheshinski, E. and Yoram Weiss (1977): "Inlfation and Costs of Price Adjustment."*Review of Economic Studies*, Vol. 44 (1977), pp. 287-303.

[22] Veronese, Giovanni, Silvia Fabiani, Angela Gattulli and Roberto Sabbatini (2005): "Consumer Price Behavior in Italy: Evidence from Micro CPI Data."*ECB Working Paper* 449 (March 2005).

APPENDIX

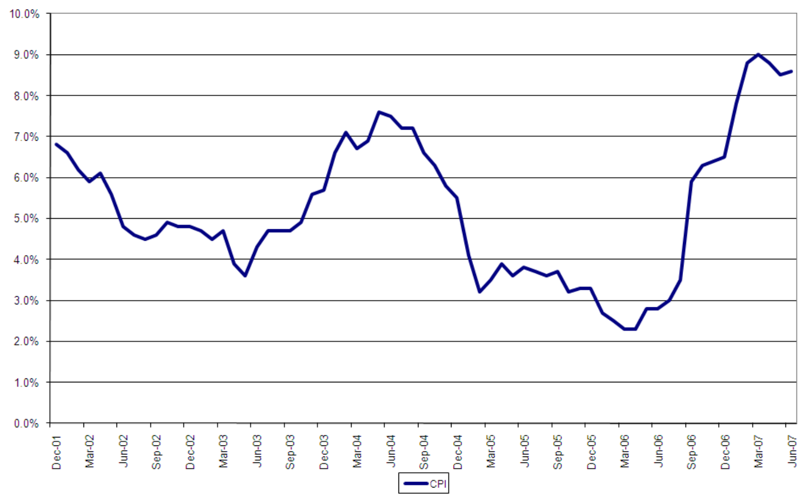


Figure 1.1: The Hungarian CPI in the sample period

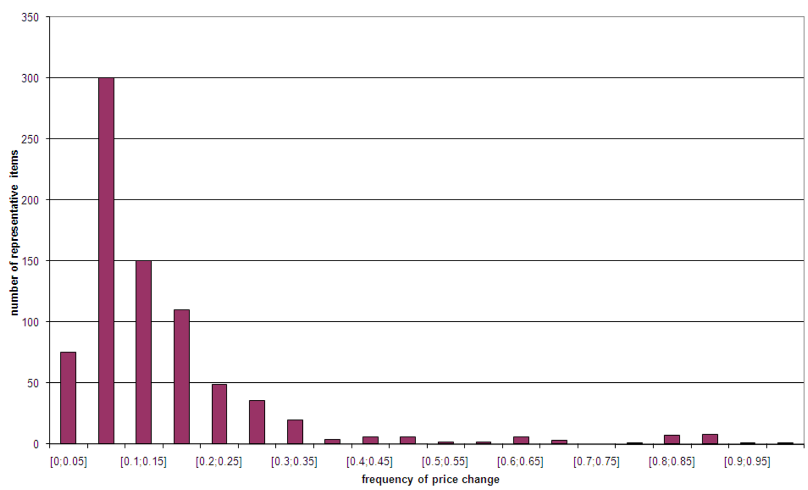


Figure 1.2: Distribution of the frequency of price changes across repr. items

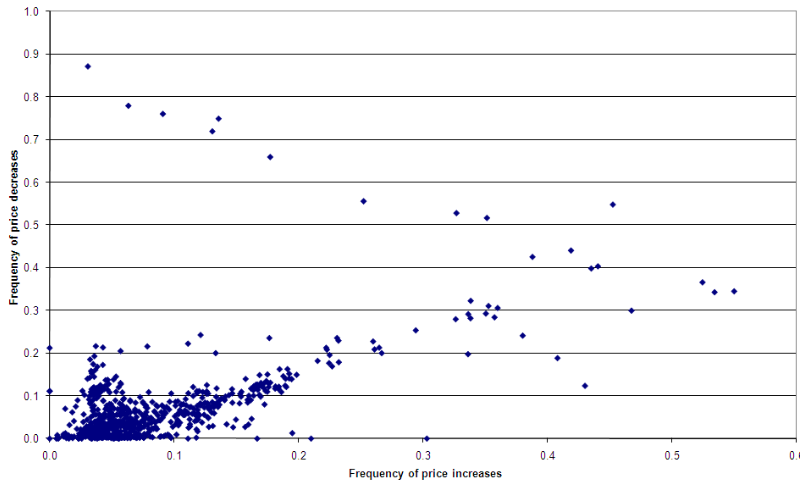


Figure 1.3: Correlation between frequency of price increase and price decrease

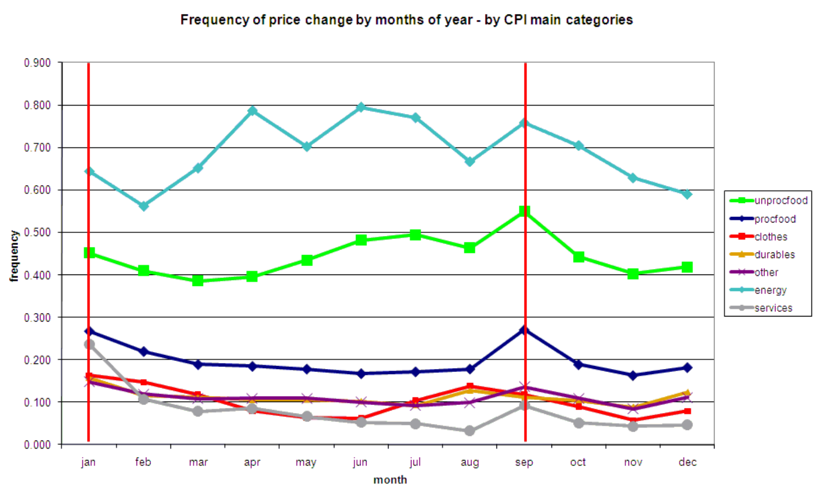


Figure 1.4: Seasonality in frequencies: frequency of price change by month

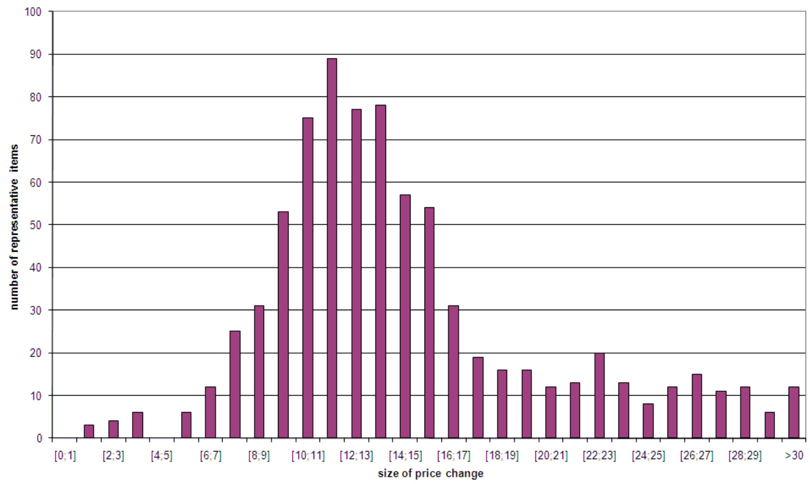


Figure 1.5: Distribution of the size of price changes across representative items

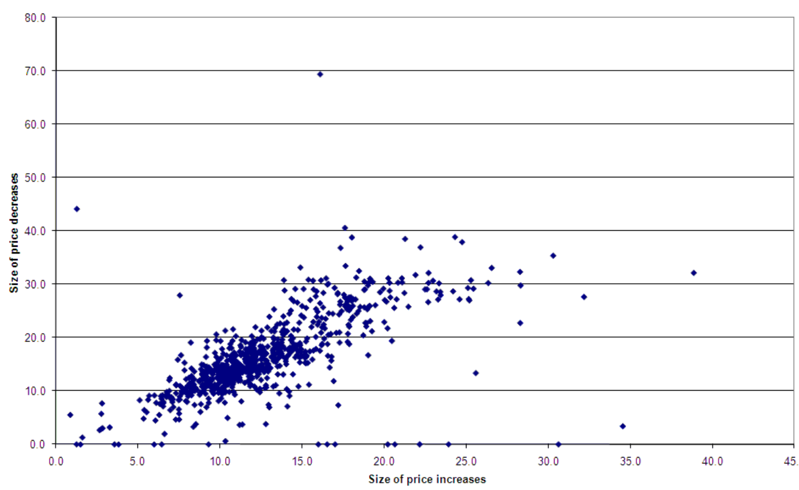


Figure 1.6: Correlation between size of price increases and price decreases

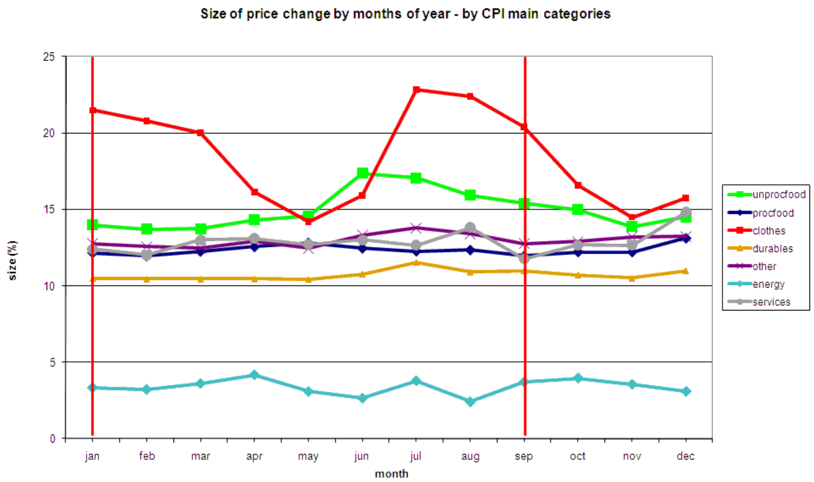


Figure 1.7: Seasonality in sizes: size of price change by month

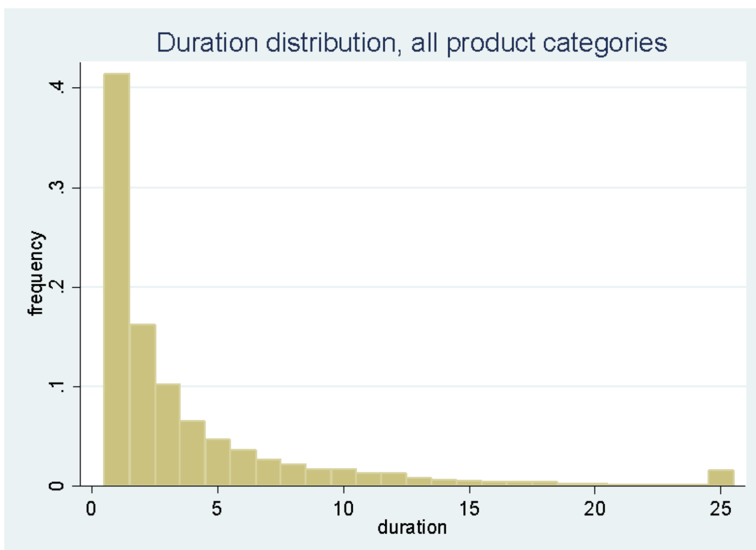


Figure 1.8: Duration distribution of uncensored spells, whole sample

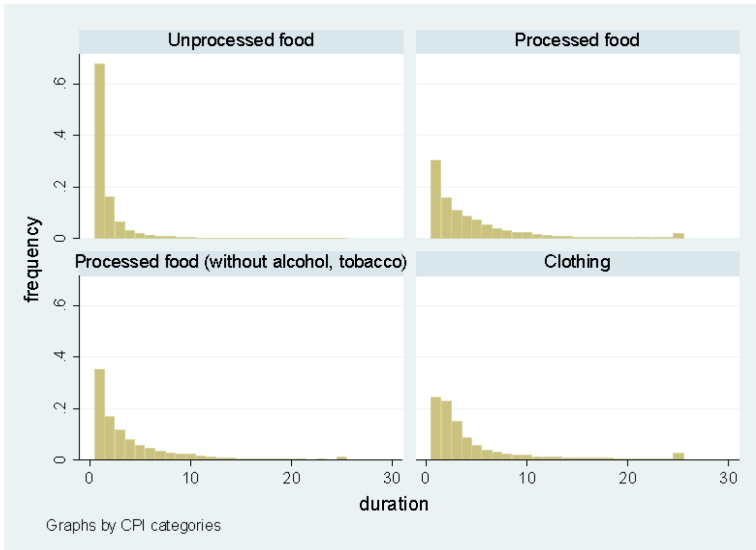


Figure 1.9: Duration distribution by CPI-categories, part 1

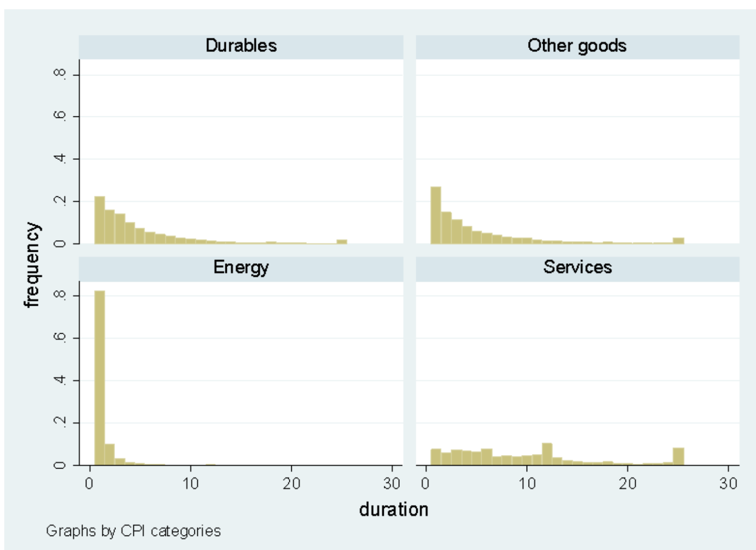


Figure 1.10: Duration distribution by CPI-categories, part 2

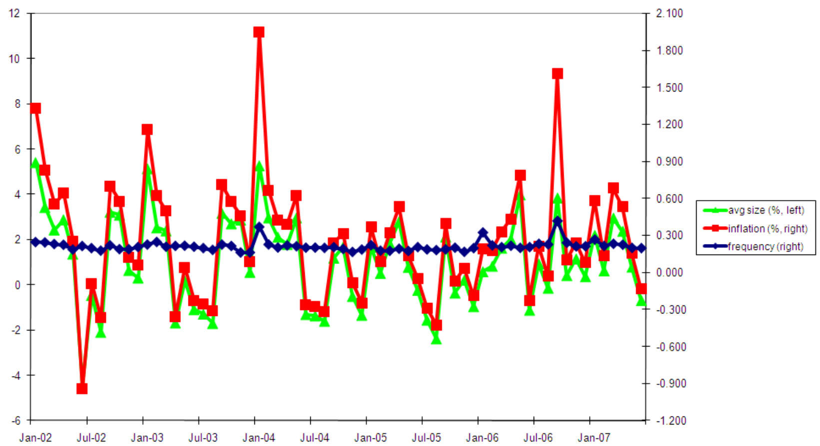


Figure 1.11: Inflation rate: decomposition to frequency and size effects

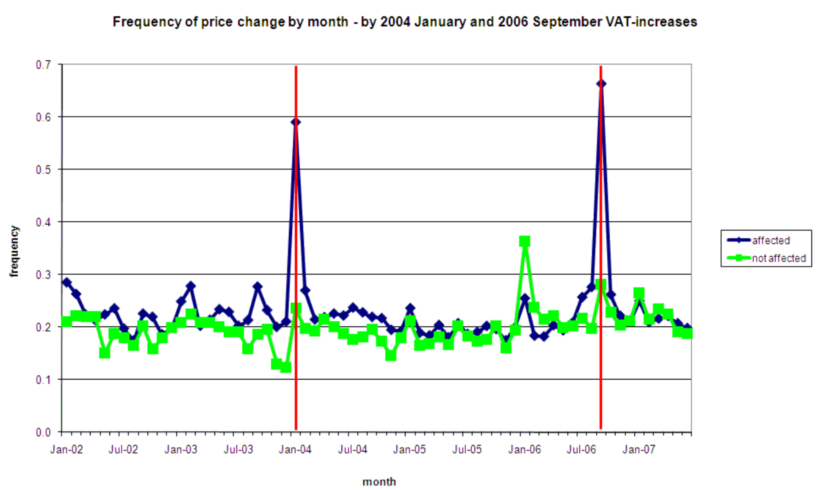


Figure 1.12: Frequency of price changes – by being affected by VAT-increases

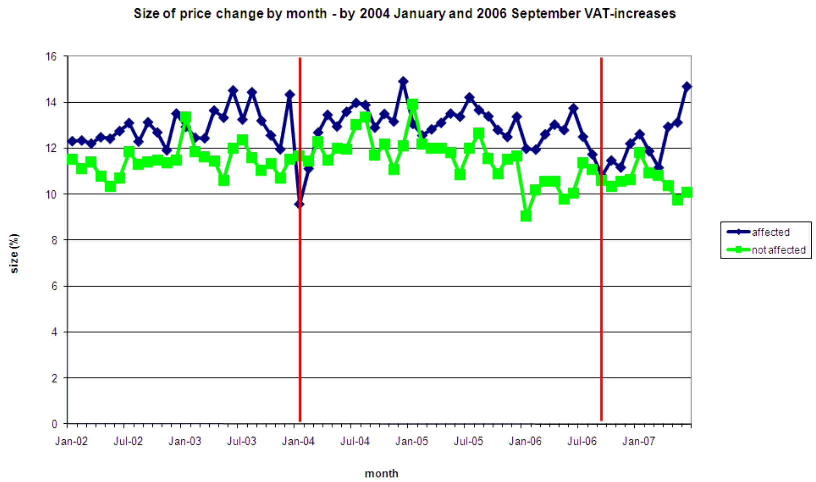


Figure 1.13: Size of price changes – by being affected by VAT-increases

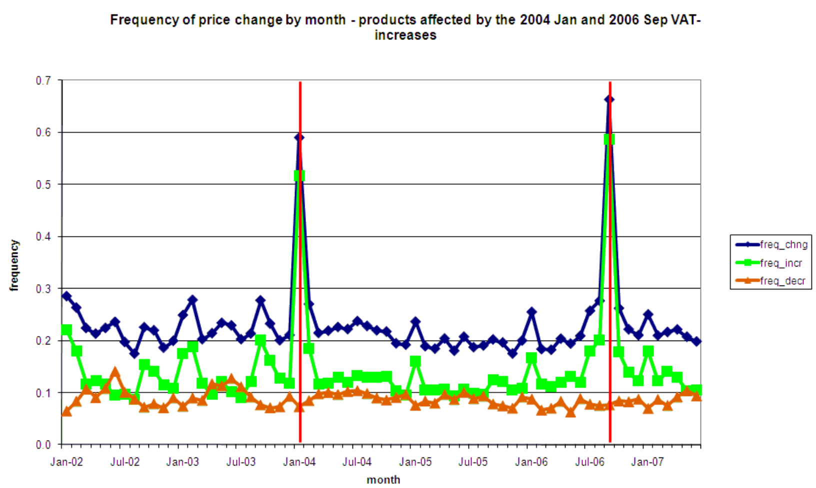


Figure 1.14: Frequency of price changes – products affected by VAT-increases

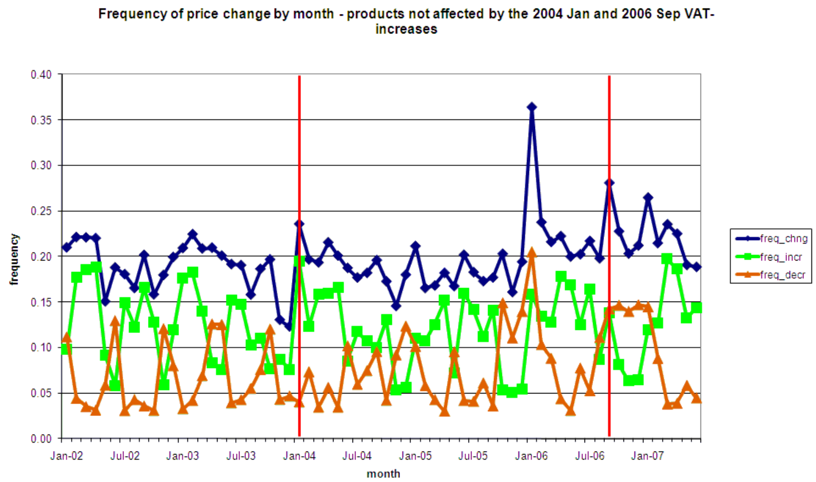


Figure 1.15: Frequency of price changes – products not affected by VAT-increases

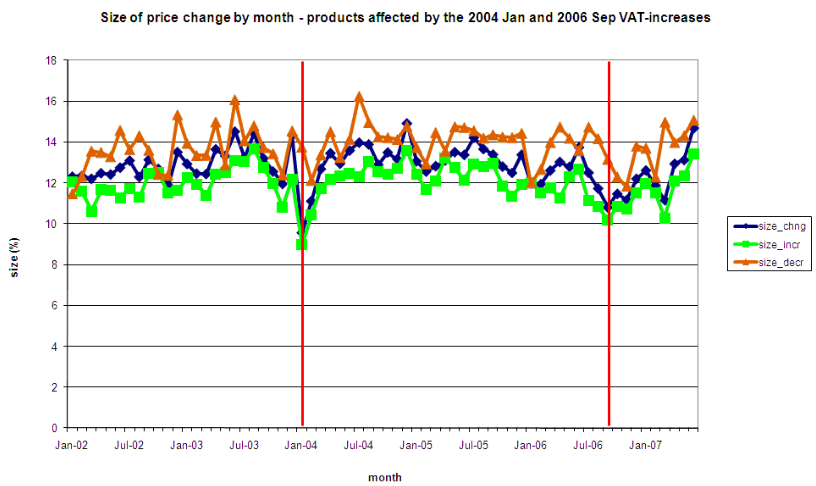


Figure 1.16: Size of price changes – products affected by VAT-increases

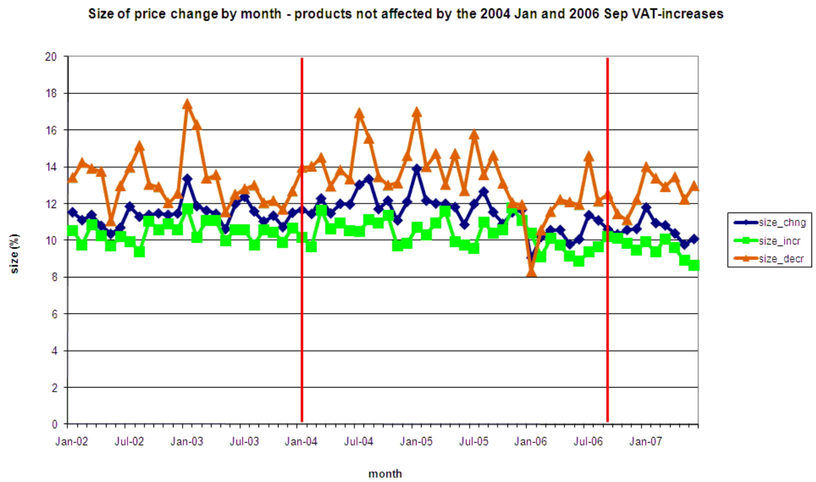


Figure 1.17: Size of price changes – products not affected by VAT-increases

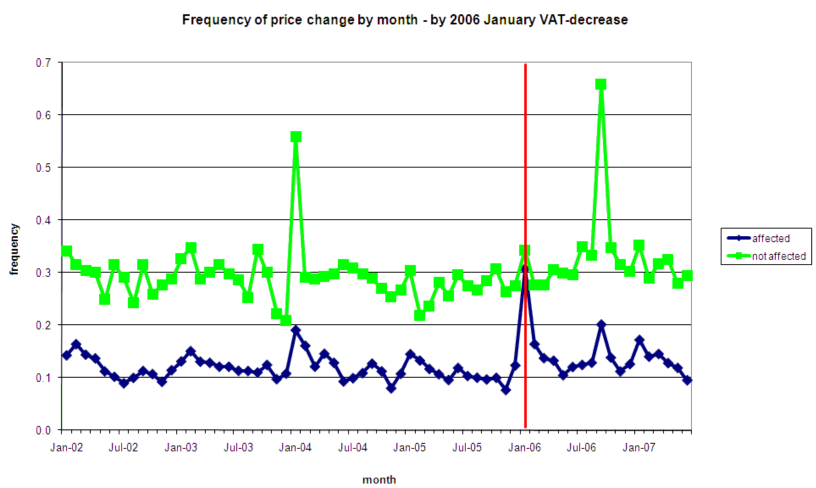


Figure 1.18: Frequency of price changes – by being affected by VAT-decrease

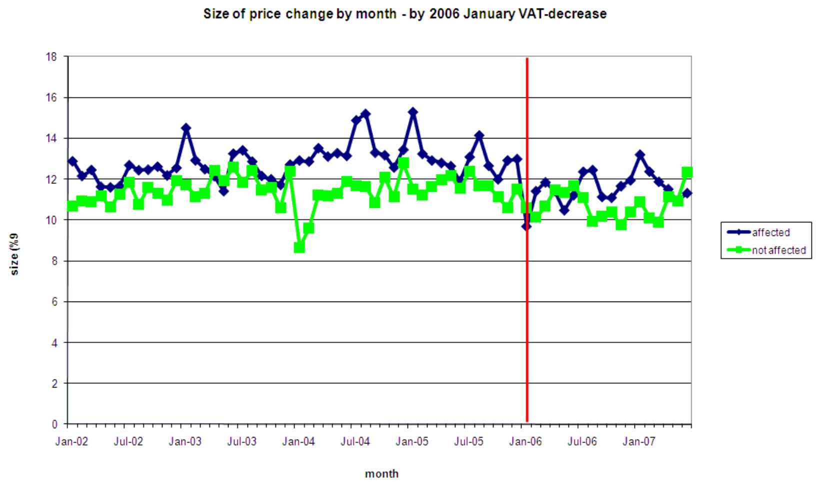


Figure 1.19: Size of price changes – by being affected by VAT-decrease

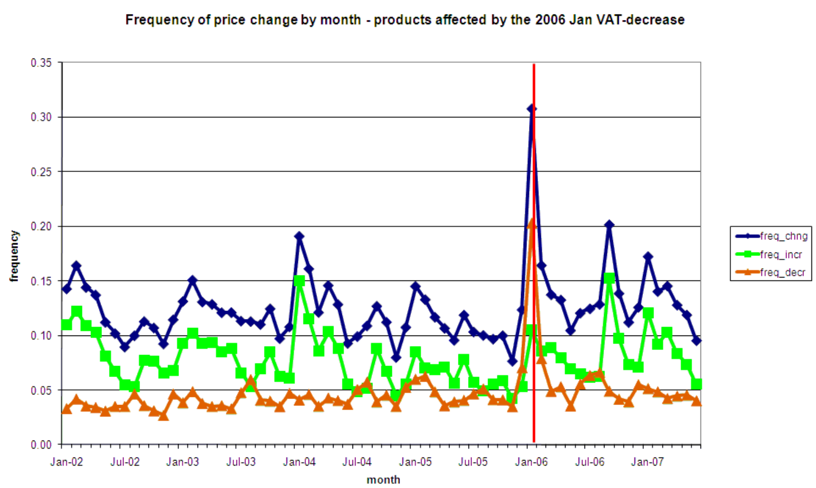


Figure 1.20: Frequency of price changes – products affected by the VAT-decrease

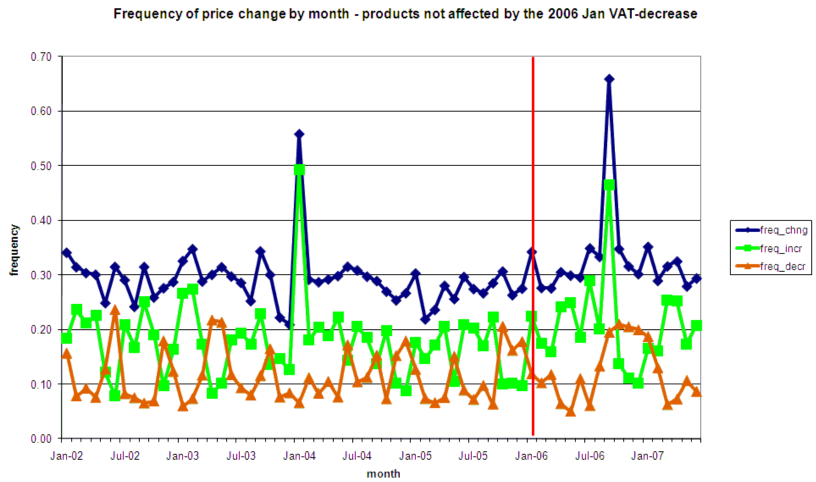


Figure 1.21: Frequency of price changes – products not affected by VAT-decrease

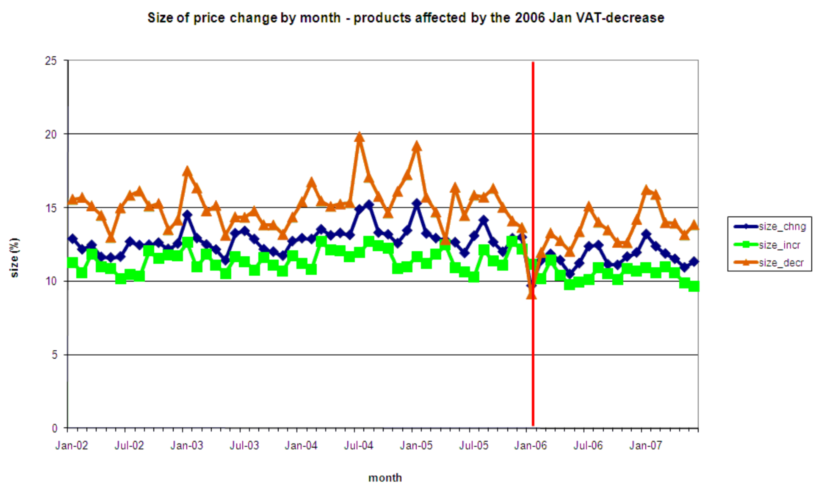


Figure 1.22: Size of price changes – products affected by VAT-decrease

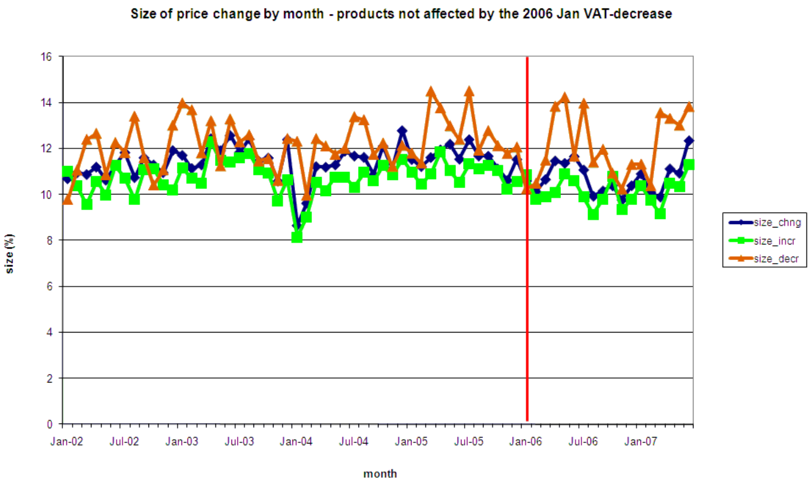


Figure 1.23: Size of price changes – products not affected by VAT-decrease

Chapter 2

Household inflation expectations and inflation dynamics

Although in modern monetary economics it is usually assumed that inflation expectations play a prominent role when economic agents set prices and wages, the empirical evidence for this link is scarce. This paper aims to identify the effect of changes in inflation expectations on prices and wages in an SVAR framework for three inflation targeting countries (Czech Republic, Hungary and United Kingdom). The results show that in all countries the effect is significant. In comparison with the United Kingdom and the Czech Republic, inflation expectations in Hungary are more volatile and less anchored, which can be an important source of the high volatility of the inflation rate.

2.1 Introduction

Managing inflation expectations is crucial in modern monetary policy. Although in the short run demand and supply effects (e. g. government spending shock or increasing commodity prices) are the main causes of changes in the inflation rate, in the long run expectations gain in importance and have the dominant role in determining inflation (Stock and Watson (2006), Cogley et al (2007)).

Inflation expectations influence prices through numerous channels. Investors need reliable inflation projections to make well-founded investment decisions. Firms need to determine their expected inflation rate in order to set prices, to make capital investment and deciding on borrowing and liquidity needs. Expected inflation is crucial in contracts which are not continuously renegotiated, like wages. Consumers use information on the future inflation rate when allocating consumption between today and tomorrow. If inflation expectations are high consumers tend to consume today, which may increase prices further.

Although expectations are important because of several reasons, the empirical literature on inflation expectation formation had a quite limited scope in general. Most of the literature analyzed whether expectations (of households, producers or investors) are unbiased and efficient predictors of future inflation rates. For the euro area Forsells and Kenny (2002) and Dias et al (2008), for the United Kingdom Mitchell and Weale (2007) tested the rationality of household expectations. The finding of these papers was that expectations are biased. Although expectations do not satisfy the rationality assumption, they may convey valuable information. For the US Ang et al. (2006) showed that household inflation forecasts outperformed many alternative methods of forecasting. A second wave of papers tried to explore more carefully how expectations are formed (e.g. Brachinger (2008), Lein and Maag (2008)). The results showed that expectations are backward looking to a great extent. Furthermore consumers do not weight products and services the same way as statistical offices. They tend to overweight frequently purchased products and expectations overreact big and transparent price increases.

On the other hand a promising new strand of literature demonstrates, that the deviation from the full information rational expectations and the introduction of different learning mechanism in macroeconomic models can help to explain economic fluctuations. As expectations are not observable, survey data can be important as a proxy variable. Del Negro and Eusepi (2009) and Orphanides and Williams (2005) used survey data about inflation expectations to inform the choice of different learning mechanisms. Ormeno (2009) and Milani (2010) moved even further and used survey data to estimate general equilibrium models.

Although the assumed linkage between expectations and other macroeconomic variables is crucial, it was rarely tested in simple multiple-equation context. Among the examples Millet (2006) found evidence that expectations influence inflation dynamics. Benkovskis (2008) examined inflation expectations in the new member states of the European Union and found that changes in expectations have a long-run effect on inflation.

This paper explores whether changes in expectations have an impact on other macroeconomic variables in three countries (Czech Republic, Hungary and the United Kingdom). The selection of countries was motivated by that all countries have inflation targeting monetary regimes, so managing expectations is in the focus of monetary policy, and for all countries relatively long time series are available about inflation perceptions and expectations. In this paper I quantify qualitative survey responses about inflation perceptions and expectations and use a SVAR framework to identify the effect of changes in the expected inflation rate. Previous papers usually used simple ordering assumptions to put inflation expectations into VARs. The main contribution of this paper is that it proposes a different SVAR framework with sign restrictions, which can be more appropriate to identify the effect of expectation shocks. Nominal wages are also included among the variables of the SVAR, to underpin one of central banks' main concerns, namely that non-anchored expectations may have an impact also on wage setting. The results show that an increase in inflation expectations raises prices and nominal wages in all the three countries. By comparing impulse responses I also evaluate how anchored expectations are in the three countries. Expectations are the most anchored in United Kingdom and the least in Hungary.

The paper is organized as follows. First I discuss the quantification method of qualitative survey responses and quantify inflation expectations. Second I describe the variables used in the SVAR. Third I discuss the SVAR framework I use to identify the effect of changes in expectations, then I summarize the results of the estimated SVAR. In the last section I conclude.

2.2 The extended Carlson–Parkin method of quantification

I use qualitative data about inflation perceptions and expectations from the Business and Consumer Surveys of the European Commission. The surveys are conducted on monthly basis in all EU countries, including the new member states. Monthly data are available from 1995 for the Czech Republic and from 1993 for Hungary and United Kingdom. Although in general the survey responses available monthly, several data points are missing so I quantify quarterly perceptions and expectations by taking the simple average of available monthly survey responses.

The survey questions on the perceived and expected inflation rate are the following:

- "How do you think that consumer prices have developed over the last 12 months?" The possible response categories are: (1) "risen a lot", (2) "risen moderately", (3) "risen slightly", (4) "stayed about the same", (5) "fallen" and (6) "don't know".
- "Given what is currently happening, do you believe that over the next 12 months prices will: (1) rise faster than at present, (2) rise at the same rate, (3) rise more slowly, (4) stay at their present level, (5) go down, (6) difficult to say".

The proportions of respective survey responses for the three countries are shown in Figure 2.1 – 2.6. Some differences among countries are worth to mention. Although perceptions and expectations seem to be disperse in all countries, in the United Kingdom the proportions are more stable than in the other two countries. The proportion of respondents perceiving and expecting an increase in the inflation rate is the highest in Hungary and the lowest in the United Kingdom.

Response:	Czech Republic	Hungary	United Kingdom
(1) “risen a lot”	16.1%	30.7%	12.4%
(2) “risen moderately”	29.1%	32.8%	23.0%
(3) "risen slightly"	29.4%	23.5%	35.2%
(4) “stayed about the same”	17.1%	7.9%	22.7%
(5) “fallen”	4.5%	3.9%	4.8%
(6) “don’t know"	3.7%	1.2%	1.9%
Average fraction of respondents choosing respective responses			

Table 2.1. Survey data on perceived price changes

Response:	Czech Republic	Hungary	United Kingdom
(1) "rise faster than at present"	23.2%	27.9%	17.1%
(2) "rise at the same rate"	42.5%	52.0%	35.8%
(3) "rise more slowly"	6.7%	13.1%	20.0%
(4) "stay at their present level"	9.9%	1.4%	18.7%
(5) "go down"	1.9%	0.5%	3.4%
(6) "difficult to say"	15.7%	5.1%	5.1%
Average fraction of respondents choosing respective responses			

Table 2.2. Survey data on expected price changes

The quantification of qualitative survey responses is done with the extended Carlson-Parkin method (henceforth called the Carlson-Parkin method), proposed by Berk (1999). The quantification is done in two steps. First the perceived inflation rate is quantified, then using perceived inflation expected inflation can be also determined.

During the quantification it is assumed that the aggregate distribution of both perceived and expected inflation rate is normal. This assumption has some theoretical motivations. The necessary and sufficient conditions for the existence of the aggregate normal distributions are that the individual, subjective distributions are independent and have finite mean and variance. In this case the Central Limit Theorem applies.[1]

Although the normal distribution assumption is the most frequently used, empirical studies have mixed results about the normality of inflation expectations.[2]

[1]For a more technical discussion about the assumptions necessary for the Carlson-Parkin method see Pesaran (2005).

[2]Normality cannot be tested on qualitative surveys. The papers cited above used cross sectional quantitative data about inflation expectations.

Normality was rejected by Carlson (1975) for the US, by Batchelor (1981) for Germany, France and UK, and more recently by Murasawa (2009) for Japan. The main findings of these papers were that compared to the normal distribution survey responses are more centrally peaked and tend to be distributed asymmetrically. On the other hand Balcombe (1996) found no evidence to support the hypothesis that price expectations are drawn from skewed or kurtotic distributions using a survey on inflation expectations in New Zealand. Although the normality assumption is questionable, Berk (1999) and Liziak (2003) found that different distribution assumptions (normal, central t, non-central t and uniform) caused relatively small changes in the quantified inflation rate.

Figure 2.7 illustrates the aggregate distribution of the perceived inflation rate. If the "moderate" inflation level is taken as given, there are four unknown parameters of the distribution and four independent proportions, so the unknown parameters are uniquely determined. Algebraically the quantification of the perceived inflation rate is done as follows. Define a, b, c, d and e, as:

$$a = P(\pi^p < -t) = F(-t) \tag{2.1}$$

$$b = P(-t < \pi^p < t) = F(t) - F(-t) \tag{2.2}$$

$$c = P(t < \pi^p < \pi^m - s) = F(\pi^m - s) - F(t) \tag{2.3}$$

$$d = P(\pi^m - s < \pi^p < \pi^m + s) = F(\pi^m + s) - F(\pi^m - s) \tag{2.4}$$

$$e = P(\pi^m + s < \pi^p) = 1 - F(\pi^m + s) \tag{2.5}$$

where $F(.)$ is the normal distribution function, m is the average and σ is the standard deviation of the aggregate distribution of perceived inflation rate. The moderate rate of price increases is denoted by π^m, and parameters t and s show how wide the ranges are, where respondents consider inflation rate to be zero or moderate. After solving the equations (2.1) - (2.5), one can show that

$$m = \frac{\pi^m(C + D)}{C + D - (A + B)} \tag{2.6}$$

$$\sigma = \frac{-2\pi^m}{C + D - (A + B)} \tag{2.7}$$

$$s = \frac{\pi^m(B + A)}{C + D - (A + B)} \tag{2.8}$$

$$t = \frac{\pi^m(D - C)}{C + D - (A + B)} \tag{2.9}$$

where $A = N^{-1}(1-e)$, $B = N^{-1}(1-d-e)$, $C = N^{-1}(1-c-d-e)$, $D = N^{-1}(1-a)$.

The aggregate distribution of the expected inflation rate is illustrated in Figure 2.8. The respondents of the survey have to compare price increases in the future to the inflation "...rate as at present". I assume that this rate is equal to the quantified perceived inflation rate. By taking this rate as given the quantification of the expected inflation rate is basically the same as the quantification of the perceived rate of inflation. Define f, g, h, i and j, as:

$$f = P(\pi^e < -u) = F(-u) \tag{2.10}$$

$$g = P(-u < \pi^e < u) = F(u) - F(-u) \tag{2.11}$$

$$h = P(u < \pi^e < m - v) = F(m - v) - F(u) \tag{2.12}$$

$$i = P(m - v < \pi^e < m + v) = F(m + v) - F(m - v) \tag{2.13}$$

$$j = P(m + v < \pi^e) = 1 - F(m + v) \tag{2.14}$$

Let n and ω denote the mean and the standard deviation of the aggregate distribution of expected inflation rate π^e. The average perceived inflation rate is denoted by m as before and parameters u and v show how wide the ranges are where respondents consider expected inflation rate zero or equal to the perceived inflation rate. After solving the equations (2.10) - (2.14), one can show that

$$n = \frac{m(H+I)}{H+I-(F+G)} \tag{2.15}$$

$$\omega = \frac{-2m}{H+I-(F+G)} \tag{2.16}$$

$$v = \frac{m(G-F)}{H+I-(F+G)} \tag{2.17}$$

$$u = \frac{m(I-H)}{H+I-(F+G)} \tag{2.18}$$

where $F = N^{-1}(1-j)$, $G = N^{-1}(1-i-j)$, $H = N^{-1}(1-h-i-j)$, $I = N^{-1}(1-f)$.

The crucial assumption to quantify inflation perceptions - and indirectly inflation expectations - is how the "moderate" rise in prices is defined. In the related literature the "moderate" inflation rate is assumed to be constant. There are two ways to choose its constant level. The first one is to take the average inflation rate through a longer period as the "moderate" inflation rate. The second is to set the "moderate" level so, that on average perceptions are equal to the actual inflation rate. However these approaches are acceptable only if the inflation rate is stationary, otherwise assuming a constant level for the "moderate" inflation rate can be misleading. The inflation rate was stationary in United Kingdom, but not in the other two countries.

To quantify perceptions I assumed that the perceived inflation rate is equal to the headline inflation rate on average. In case of Czech Republic and Hungary I relaxed the constant "moderate" inflation rate assumption. In both countries at the beginning of the sample period the inflation rate fluctuated at a relatively high level, which was followed by a disinflation period and then the inflation rate leveled out at a lower level. I assumed that the households' perceptions about the "moderate" inflation rate followed similar pattern. Technically I divided the sample into three subperiods. In the first and the third subperiod the moderate rate of inflation was assumed to be constant, and in the second subperiod it assumed to decrease gradually. The constant level in the first and third period and the beginning and the end of the second period was set by minimizing the squared difference of the perceived and actual inflation rate on the whole sample. According to the results of minimizing squared differences the length of the second period is much longer in Hungary than in Czech Republic, which is acceptable as the disinflation was also more gradual.[3]

The quantified perceived and expected inflation rates for the three countries are shown in Figure 2.9-2.11. The perceived and expected inflation rates are the least persistent in Hungary, but relative to the headline inflation rate they are quite volatile also in the other two countries (Table 2.3). Focusing on the end of the sample periods it is apparent, that during the oil and food price shock beginning in 2007 inflation perceptions and expectations remained below the headline inflation rate in Czech Republic and United Kingdom – although perceptions and expectations rose considerably also in these two countries – but in Hungary they exceeded the official numbers. These stylized facts may indicate that inflation expectations are the least anchored in Hungary. This feature of

[3] As a result of the chosen method to set the "moderate" inflation rate in the second subperiod the "moderate" and actual inflation rate decreases parallel. It can be argued that perceptions follow the disinflation only with some delay and the decrease of the perceived "moderate" inflation rate should be slower. To address this issue I used an alternative method to set the "moderate" inflation level. I set the level in each period so, that if the households perceptions of the "moderate" inflation level had have been that constant during the previous three years, the perceived and actual inflation rate would have been equal on average during that period. By using this method the perceived "moderate" inflation level is much more backward looking. The results discussed later proved to be robust to the different methods of quantifying inflation rates. This is a consequence of that the results depend on the short run dynamics of the quantified inflation rate, which is not sensitive to the selection between alternative methods.

expectations will be discussed in a more formal way later.

	Czech Republic	Hungary	United Kingdom
Average quaterly absolute change of the headline inflation (percentage points)	0.82	1.46	0.19
Average quaterly absolute change of the perceived inflation (percentage points)	0.45	0.84	0.15
Average quaterly absolute change of the expected inflation (percentage points)	1.03	1.18	0.30

Table 2.3. Relative volatility of inflation perceptions and expectations

2.3 Data for the SVAR estimation

After deriving the time series of the quantified inflation expectations and perceptions, the next step is estimating the SVAR to investigate the relationship between inflation expectations and other macroeconomic variables, most importantly the headline inflation rate. In the SVAR four endogenous variables are used: inflation expectations, CPI, average nominal wage in the private sector and the volume of retail sales. I included CPI and nominal wage, as my main interest is to explore whether changes in expectation influence prices and wages. I added retail sales as a business cycle indicator, which helps to identify structural shocks. Estimating the SVAR I used the log differences of CPI, nominal wage and retail sales and simple differences of inflation expectations.

The data for the SVAR was picked from different sources. The time series of expected inflation is the result of the Carlson-Parkin quantification method as previously discussed. The average level of the quantified expected inflation rate depends on the value of the "moderate" inflation level, so it is somewhat arbitrary. For the analyzed three countries quantitative survey data is also available, although only for relatively short period. Figure 2.12-2.14 show that quantitative survey expectations on average are relatively close to the quantified ones in Czech Republic and above them in Hungary and United Kingdom. However the expectations from the two data sources tend to move parallel. As the results presented below are affected only by the dynamic of the quantified inflation rate and not by its average value, the assumption on the "moderate" inflation rate does not bias the results.

All other data I use is publicly available. The volume of retail sales, CPI and average nominal wage in the private sector for Hungary are published by the Central Statistical Office.[4] In case of UK and Czech Republic the volume of retail sales is from the national statistical offices' websites, CPI is from the IFS database and nominal wage in the private sector is from the Eurostat database. Before estimating the SVAR the time series of retail sales, nominal wage and CPI data were seasonally adjusted.

I also included VAT change dummies as exogenous variables in the SVAR. The importance of this variable will be discussed later. The dates of VAT changes are available from the European Commission's homepage.[5]

To estimate the SVAR I used shorter sample than what was available. The reason was that in the full sample some variables were not stationary, so the estimation of the SVAR did not provided plausible parameters in the sense that without shocks the endogenous variables did not converge to plausible values. To ensure the validity of the estimation I chose the sample period 1999Q1-2009Q1 for the Czech Republic, 1996Q1-2009Q1 for Hungary and 1993Q1-2009Q1 for United Kingdom.

The robustness of the results was checked extensively. Instead of retail sales I used also consumption as business cycle indicator. As in Hungary the monetary regime was different in the first half of the sample I estimated the SVAR also on a shorter sample, which included observations between 2001Q1 and 2009Q1. For all specifications the qualitative results remained the same.

2.4 Identifying the SVAR parameters

For the identification of the effect of expectations on prices and wages I defined three shocks: a demand, a supply and an expectation shock. Using demand and supply shocks is quite standard in SVARs. The supply shock can be interpreted as a technology shock, demand shock can be regarded as a preference or a government spending shock. Because of the nature of these shocks, after they occur inflation expectations are expected to move parallel with or follow the headline inflation rate. Hence these shocks are accounting for the periods when inflation expectations were mainly adaptive which seemed to be the case in most of the sample periods. However expectations sometimes seemed to be less adaptive and

[4] The wage data for Hungary is quite volatile and occasionally distorted by government measures. These measures included tax changes, actions to battle bogus contracts and the tightening of inspections by the tax authority. Tax changes also led to change in the timing of bonus payments. Eppich and Lőrincz (2007) developed several methods to filter out the effect of government measures from the wage dynamics. The estimation was repeated with this filtered data, but the results remained unchanged.

[5] http://ec.europa.eu/taxation_customs/resources/documents/taxation...
/vat/how_vat_works/rates/vat_rates_en.pdf

inflation and expectation dynamics were different. These episodes could happen because of various reasons.

First of all households' expectations occasionally may be influenced by news from the media. Lamla and Lein (2008) showed that newspaper articles and TV news have an important role in forming consumers' inflation expectations. They showed that higher intensity of reporting makes consumers more likely to update their expectations and brings them closer to the full information rational forecast. On the other hand eventually the media causes bias in consumer expectations. The media tends to use exaggerated and incomplete information, and consuming these reports distorts the consumers' forecasts.

One empirical example for the bias caused by the media was the period before the EU accession, when in the new member states expectations rose considerably although economic theory suggested that a drop in inflation was to be expected. Expectations also rose rapidly in the eurozone countries before introducing the euro.

Consumers' expectations may become also biased, because they tend to overweight frequently purchased items, like food and gasoline. Also products with regulated prices can be overweighted, because their prices are more transparent and attract more media attention. Furthermore consumers are more sensitive to price increases than decreases, so during the period of big relative price changes consumers tend to perceive a price increase on average.

Although in the previous examples consumers did not act as professional forecasters of the headline inflation rate, their expectations may still play an important role in price and wage setting decisions and can prove to be self-fulfilling. The third shock, the expectation shock is defined to account for these types of changes in inflation expectations and to explore this self-fulfilling nature of expectations. As it was discussed before expectation shocks are hard to connect to specific macroeconomic variables and they can be taken as the "noise" when looking at the time series of inflation expectations. On the other hand exactly this feature of expectation shocks make them helpful to asses whether expectations affect the price index.

It is important to note, that expectations lead inflation also in case of previously announced changes of consumption taxes, but obviously did not cause the change in the inflation rate. During the sample periods the VAT rate changed in each country. As I am interested in whether changes in expectations affect prices and wages, it is essential to control for these episodes. Technically I used two dummy variables to filter out the effect of each VAT rate change. The first one was used to control for the effect of the tax change on the endogenous variables – most importantly on expectations – one period before the tax change, the second was introduced to control for the effect on endogenous variables – most importantly on inflation – in the period when the tax change happened.

Formally I impose the following zero and sign restrictions.[6] I assume that after the demand shock – for example a government spending shock – both retail sales and inflation rises and I do not impose any restriction on the average nominal wage and inflation expectations. Theoretically the effect of the demand shock is expected to be positive on both variables. A positive supply shock – for example a positive technology shock – is assumed to increase retail sales and decreases the inflation rate. There are no restrictions on inflation expectations and on the nominal wage. The expected sign of supply shock on inflation expectations is negative. After the supply shock a real wage increase is expected, but as inflation is falling the effect of the shock on nominal wage is ambiguous. I assume that the expectation shock affects immediately only inflation expectations, and the other variables are affected only one period later. There are no restrictions on the effect on retail sales, CPI and nominal wage after the first period.

	demand shock	supply shock	expectation shock
retail sales	+	+	0
CPI	+	-	0
nominal wage	?	?	0
expected inflation	?	?	+

Table 2.4. Zero and sign restrictions in the SVAR

2.5 Estimating the SVAR with sign restrictions

First a VAR(2) model was fitted to quarterly observations for the three countries. The number of lags in a VAR is usually determined by the Akaike or Swartz criteria. In the estimated VARs the value of the Akaike and Swartz criteria are very similar in case of using one or two lags. Because of bigger flexibility I chose two lags.

The identification of the SVAR parameters was done as the following. Consider a general reduced form VAR(p) model [7]: $y_t = A(L)y_t + e_t$, where $A(L) = A_1L + ...A_pL$ is a p^{th} order matrix polynomial and L is the lag operator.

The MA representation is $y_t = (I - A(L))^{-1}e_t = B(L)e_t$. Assume that ε_t is the shock in the structural form of the VAR, $E[\varepsilon_t\varepsilon_t'] = I$ and $e_t = H\varepsilon_t$. Then $E[e_te_t'] = HH' = \Sigma$. As H has more free parameters than Σ, restrictions

[6] The sign restrictions are imposed for four periods. The number of restricted periods was chosen to distinguish between noise and meaningful shocks in the data. However using less restricted periods does not change the results qualitatively.

[7] The exogenous variables play a role only in the estimation of the reduced form of the VAR. For simplicity I left these variables out throughout the discussion of the estimation method.

should be imposed. I use both zero and sign restrictions. Note that by putting the expectation shock last the Choleski decomposition of Σ satisfies the zero restrictions. Denote this decomposition as $\Sigma = PP'$, where P is a lower triangular matrix.

Let Q be an orthonormal matrix, that is a matrix with the property $QQ' = Q'Q = I$. Since $\Sigma = PP' = PQQ'P'$, PQ is also a potential candidate for the relationship between shocks of the reduced form and the structural form of the VAR, if all restrictions are satisfied.

The estimation of the structural parameters was done with the following algorithm:

1. First the reduced form is estimated with two lags.
2. Take the Choleski decomposition ($\Sigma = PP'$) of the estimated covariance-variance matrix.
3. Generate a random orthonormal matrix of the form:

$$Q = \begin{bmatrix} q_{11} & q_{21} & q_{41} & 0 \\ q_{21} & q_{22} & q_{42} & 0 \\ q_{31} & q_{23} & q_{43} & 0 \\ q_{41} & q_{24} & q_{43} & 1 \end{bmatrix}.$$

   Q is the rotation matrix. PQ is a candidate matrix for the structural relationships and satisfies the zero restrictions.

4. Calculate the impulse responses, and check sign restrictions. If they are satisfied store PQ.
5. Repeat the steps from the third to take enough draws of Q to scan the space of possible impulse responses that satisfy the restrictions.
6. The draws that satisfy the restrictions provide a range of responses. To summarize the result one should calculate mean and probability bands.

2.6 Results

Figure 2.15-2.17 show the mean and – as quite common in the VAR literature - the 16^{th} and the 84^{th} quantiles of the impulse responses for each horizon between zero and 20. Although sign restrictions are considered to be weak in general some results are quite significant on shorter horizons.[8],[9]

The sign of the estimated impulse responses are broadly in line with theoretical considerations in case of all countries. Demand and supply shocks have an effect until 4-8 quarters in general. In case of demand and supply shocks the sign of responses was restricted for retail sales and inflation. Considering the unrestricted impulse responses demand shocks had relatively small effect on wages and a quite persistent positive effect on expectations in case of Czech Republic and Hungary. In case of supply shocks wages remain more or less unchanged in all countries, so real wage increases. This is consistent with what we expect after a technology shock. Supply shocks also have a strong effect on expectations up to more than one year. Expectation shocks played an important role in inflation and wage dynamics. The effect on inflation and nominal wage is quite persistent in general.

To evaluate and interpret the magnitude of the responses I compared the impulse responses for the three countries.[10] Figure 2.18-2.20 compare the effect of the expectation shock on expectations, inflation and wages. Expectation shocks had the biggest impact in Hungary. The effect of the expectation shock on inflation expectations and inflation is quite persistent for all countries. Central banks tend to focus also on the connection between inflation expectations and wage bargaining. If expectations are high, then employees demand higher wages which may be hard to resist for employers. Higher wages may lead to loss in competitiveness and higher prices. The results show that expectation shocks increase nominal wages in all countries. In case of Czech Republic and Hungary the effect is also quite persistent.

According to Figure 2.21-2.22 the effect of demand shocks on inflation is quite big in all countries, but the impact on expectations is relatively small. The

[8]The impulse responses show the effect of the structural shocks of the size of one standard deviation. In case of inflation expectations the effect is measured in percentage points. In case of retail sales, inflation and nominal wage the impulse responses show the effect on the quaterly indexes.

[9]Sign restrictions are considered to be weak restrictions in general. The ratio of the rejected random rotation marices to total draws can be informative on the strength of the restrictions. In case of the expectation shock the ratio is about 50 percent for each country. In case of demand and supply shocks the ratio of rejected rotation matrices is about 10 percent for Hungary and United Kingdom. These numbers indicate that for these countries the restrictions are indeed not that strong. On the other hand for the Czech Republic 99 percent of the rotation matrices is rejected for demand and supply shocks, so these restrictions can be considered somewhat stronger.

[10]In Figure 18-24 in case of retail sales, inflation and nominal wages the impulse responses show the impact on the annualized quaterly indexes, so the effects on expectations and on the other variables are directly comparable.

effect on inflation expectations and inflation is the most persistent in the Czech Republic, in the other two countries demand shock are less important in the evolution of inflation expectations and inflation. The small and delayed effect of the demand shock on expectations also indicates that expectations are updated gradually, and even on the long run the change in the inflation rate and inflation expectations can differ considerably.

Figure 2.23-2.24 show the effect of supply shocks on inflation expectations and inflation. The size of the effect is similar in case of expectations and inflation. This can be a sign of that households are more sensitive to supply than to demand shocks. This is consistent with empirical findings, which showed that households are particularly sensitive to energy prices. In the selected countries an increase in energy prices can be considered as a negative supply shock and during the sample periods these prices were quite volatile. Relative to the increase in the inflation rate, the increase in expectations is the biggest in Hungary, which may indicate that the Hungarian households' expectations are the least anchored.[11]

Using the mean impulse responses I also did a shock decomposition exercise for the evolution of the inflation rate. Calculating the contribution of shocks to inflation volatility for the three countries, it's apparent that inflation was the most stable in the UK and the contribution of expectation shocks is low (Table 2.5). In the Czech Republic inflation volatility is higher, but the contribution of expectation shocks is still relatively low. In Hungary the inflation volatility caused by the identified shocks is also high and the contribution of expectation shocks is much more important than in the other countries.

[11] Using this simple SVAR framework does not allow to compare the stability of inflation expectations in different countries directly. There is no obvious way to distinguish anchored expectations from not-anchored expectations in periods with small aggregate shocks. To overcome this problem one can proxy the size of the shock with its effect on inflation. If the effect of the one standard deviation shock on inflation is similar for countries A and B, but the impact on expectation is bigger in A than B, then one can argue that expectation are less anchored in country A.

	Czech Republic	Hungary	United Kingdom
absolute contribution of expectation shocks (quaterly average)	0.15	0.20	0.12
absolute contribution of demand shocks (quaterly average)	0.13	0.13	0.10
absolute contribution of supply shocks (quaterly average)	0.26	0.17	0.13
absolute contribution of VAT shocks (quaterly average)	0.15	0.31	0.05
total absolute contribution of shocks (quaterly average)	0.78	0.81	0.56

Table 2.5. Contribution of shocks to the variation of the inflation rate

The decomposition of changes in inflation expectations reinforce the previous statement that expectations seem to be the most anchored in the United Kingdom and the least in Hungary (Table 2.6).[12] The volatility of expectations are the highest in Hungary and comparable to the volatility of the inflation rate. In the other two countries the volatility of the expected inflation rate is less than the half of the volatility of the headline inflation rate. Expectation shocks play an important role in all countries. In Hungary changes in the VAT rate are also important sources of the variability in expectations and may explain why monetary policy was relatively less successful in anchoring expectations.

	Czech Republic	Hungary	United Kingdom
absolute contribution of expectation shocks (quaterly average)	0.14	0.23	0.11
absolute contribution of demand shocks (quaterly average)	0.05	0.02	0.04
absolute contribution of supply shocks (quaterly average)	0.09	0.14	0.07
absolute contribution of VAT shocks (quaterly average)	0.05	0.30	0.02
total absolute contribution of shocks (quaterly average)	0.33	0.70	0.24

Table 2.6. Contribution of shocks to the variation of inflation expectations

[12] Inflation expectations are referring to annual inflation indexes. To make the values comparable to Table 5 I divided the average contributions with 4 as the inflation rate is a quarterly index.

2.7 Summary

This paper explored the connection between inflation expectations, inflation and nominal wages. To quantify household survey responses about perceived and expected inflation the extended version of the Carlson-Parkin method was applied. I argued that increases in inflation expectations can be self-fulfilling and used an SVAR framework to provide empirical evidence for Czech Republic, Hungary, and United Kingdom, that changes in expectations influence prices and wages considerably. The results may be taken as an empirical support to the new strand of modeling literature, which relaxes the full information, rational expectations hypothesis. The SVAR methodology enables to evaluate the stability of the households' inflation expectations. Comparing the impulse responses inflation expectations seems to be the most anchored in United Kingdom and least in Hungary.

Bibliography

[1] Ang, Andrew, Bekaert, Geert and Wei, Min (2007): "Do macro variables, asset markets, or surveys forecast inflation better?", Journal of Monetary Economics, 54, issue 4, p. 1163-1212,

[2] Balcombe, K. (1996) "The Carlson and Parkin method applied to NZ price expectations using QSBO survey data", Economic Letters, vol. 51

[3] Benkovskis, Konstantins (2008): "The Role of Inflation Expectations in the New EU Member States: Consumer Survey Based Results," Czech Journal of Economics and Finance, vol. 58

[4] Berk, Jan Marc (1999): "Measuring Inflation Expectations: a Survey Data Approach", Applied Economics, vol. 31

[5] Brachinger, Hans Wolfgang (2008): "A new index of perceived inflation: Assumptions, method, and application to Germany," Journal of Economic Psychology, Elsevier, vol. 29(4)

[6] Carlson, J. (1975): "Are price expectations normally distributed?", Journal of the American Statistical Association, vol. 70

[7] Carlson, J. and Parkin, M. (1975): "Inflation expectations", Economica, vol. 42

[8] Cogley, Timothy & Primiceri, Giorgio E. & Sargent, Thomas J. (2008): "Inflation-Gap Persistence in the U.S," NBER Working Papers 13749, National Bureau of Economic Research, Inc.

[9] Del Negro, M., and S. Eusepi, (2009): "Modeling Observed Inflation Expectations", mimeo, Federal Reserve Bank of New York

[10] Dias, F., Duarte, C., and Rua, A. (2008): "Ination expectations in the euro area: Are consumers rational?" Banco de Portugal, Working Paper, 23.

[11] Forsells, M., Kenny, G. (2002): „The rationality of consumers' inflation expectations: Survey-based evidence for the euro area" European Central Bank Working Paper Series, No. 163.

[12] Lamla, Michael J. & Lein, Sarah M. (2008): "The Role of Media for Consumers' Inflation Expectation Formation," KOF Working papers 08-201, KOF Swiss Economic Institute

[13] Lein, Sarah M. & Maag, Thomas (2008): "The Formation of Inflation Perceptions – Some Empirical Facts for European Countries," KOF Working papers 08-204, KOF Swiss Economic Institute

[14] Liziak, Tomasz (2003): "Consumer Inflation Expectations in Poland", ECB Working Paper, No. 287

[15] Milani, Fabio (2010): „Expectation shocks and learning as drivers of the business cycle", EABCN/CEPR Discussion Paper 52/2010

[16] Millet, Fabien Curto (2006): "Finding the Optimal Method of Quantifying Inflation Expectations on the Basis of Qualitative Survey Data", Narodowy Bank Polski conference materials

[17] Mitchell, J., & Weale, M. (2007): "The rationality and reliability of expectations reported by British households: Micro evidence from the British household panel survey" NIESR WP.

[18] Murasawa, Yasutomo (2009): "Measuring Ination Expectations from Interval-Coded Data", Osaka Prefecture University, Working Paper

[19] Ormeno, A., (2009): "Disciplining Expectations: Using Survey Data in Learning Models", mimeo, Universitat Pompeu Fabra

[20] Orphanides, A., and J. C. Williams (2005): "The Decline of Activist Stabilization Policy: Natural Rate Misper-ceptions, Learning, and Expectations", Journal of Economic Dynamics and Control, Vol. 29

[21] Pesaran, M. Hashem & Weale, Martin (2005): "Survey Expectations," IEPR Working Papers 05.30, Institute of Economic Policy Research (IEPR)

[22] Stock, James H. & Watson, Mark W. (2006): "Why Has U.S. Inflation Become Harder to Forecast?, "NBER Working Papers 12324, National Bureau of Economic Research, Inc.

[23] Woodford, Michael (2005): "Central-Bank Communication and Policy Effectiveness", NBER Working Paper Series 11898

APPENDIX

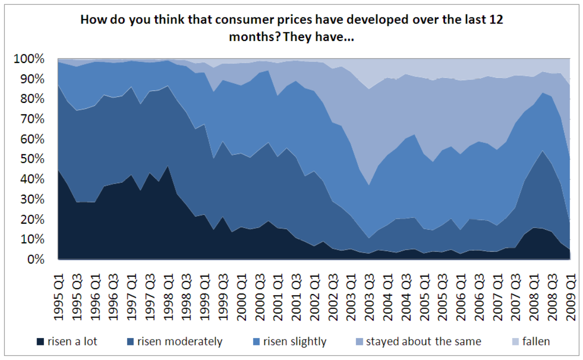


Figure 2.1. Inflation perceptions in the Czech Republic

Fraction of respondents choosing respective responses, sample Q1 1995 – Q1 2009

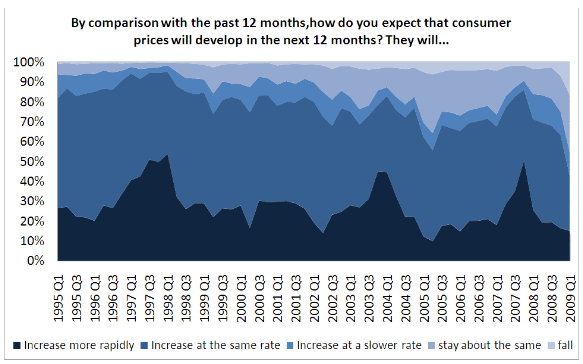


Figure 2.2. Inflation expectations in the Czech Republic

Fraction of respondents choosing respective responses, sample Q1 1995 – Q1 2009

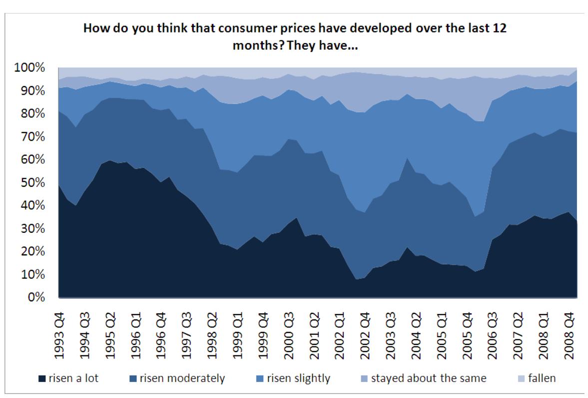


Figure 2.3. Inflation perceptions in Hungary

Fraction of respondents choosing respective responses, sample Q1 1993 – Q1 2009

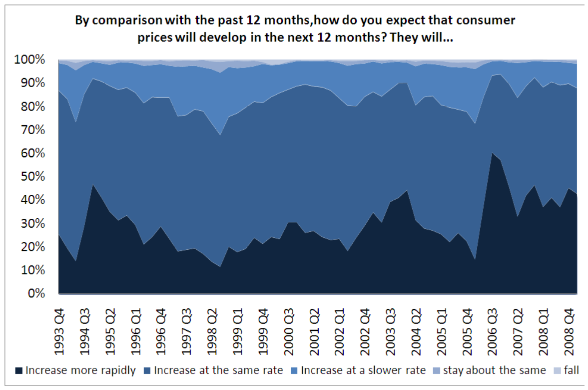


Figure 2.4. Inflation expectations in Hungary

Fraction of respondents choosing respective responses, sample Q1 1993 – Q1 2009

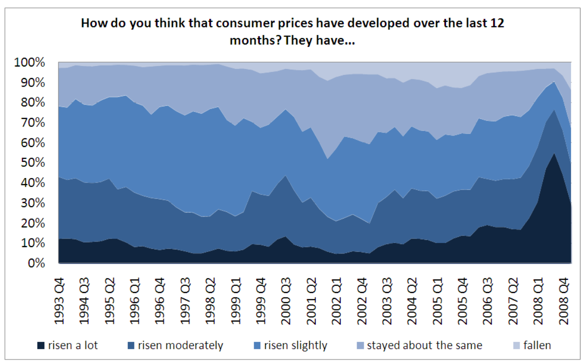


Figure 2.5. Inflation perceptions in the United Kingdom

Fraction of respondents choosing respective responses, sample Q1 1993 – Q1 2009

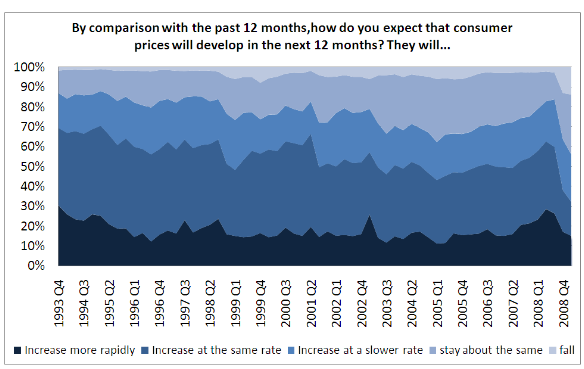


Figure 2.6. Inflation expectations in the United Kingdom

Fraction of respondents choosing respective responses, sample Q1 1993 – Q1 2009

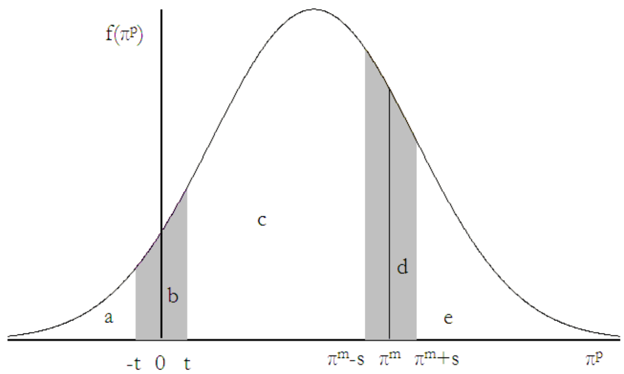


Figure 2.7. Aggregate distribution of the perceived inflation rate

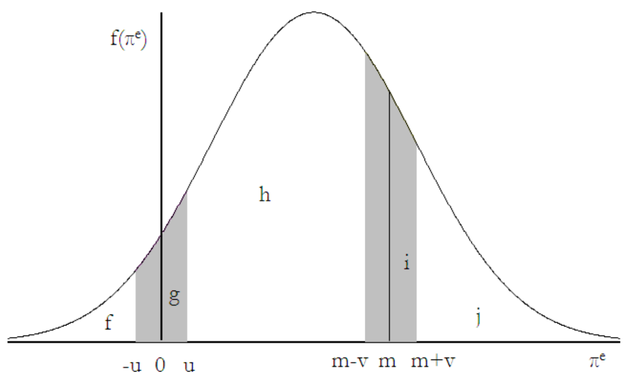


Figure 2.8. Aggregate distribution of the expected inflation rate

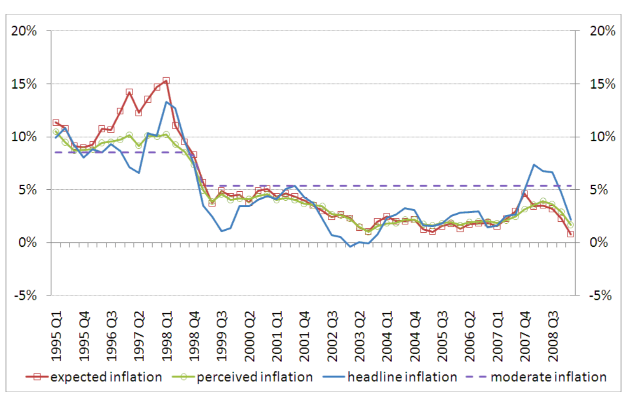


Figure 2.9. Quantified inflation perceptions and expectations in Czech Republic

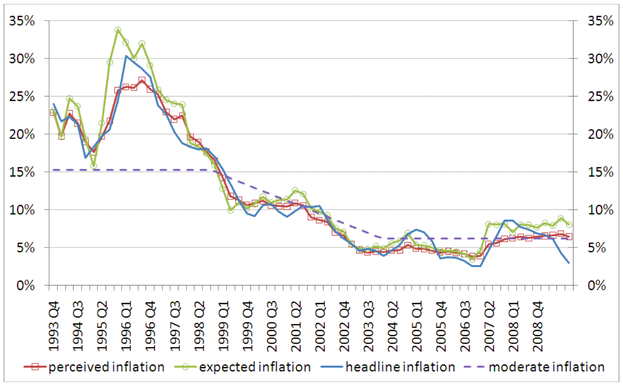


Figure 2.10. Quantified inflation perceptions and expectations in Hungary

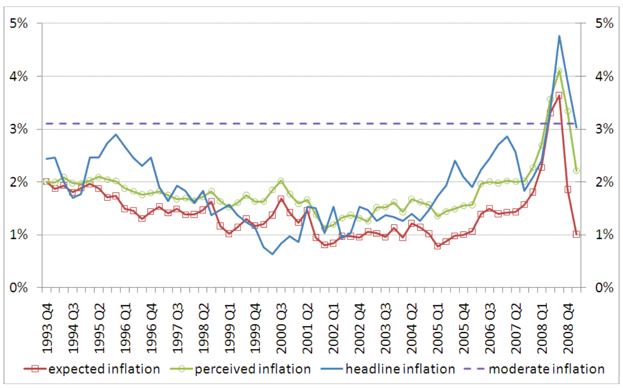


Figure 2.11. Quantified inflation perceptions and expectations in United Kingdom

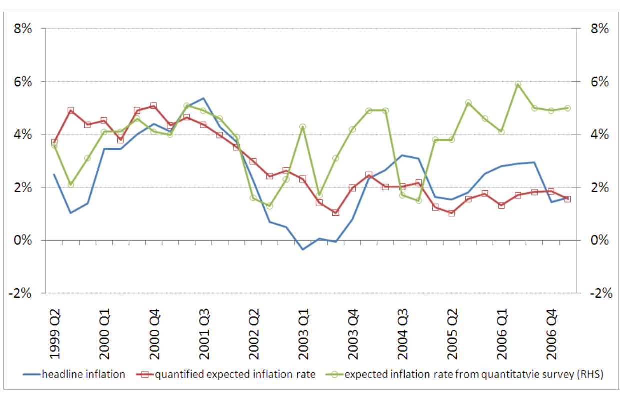


Figure 2.12. Quantitative and quantified inflation expectations in Czech Republic

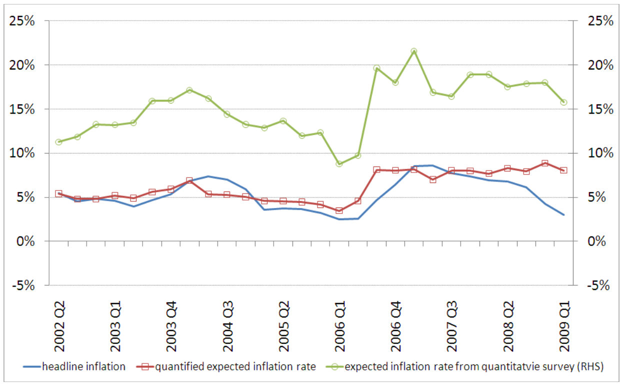


Figure 2.13. Quantitative and quantified inflation expectations in Hungary

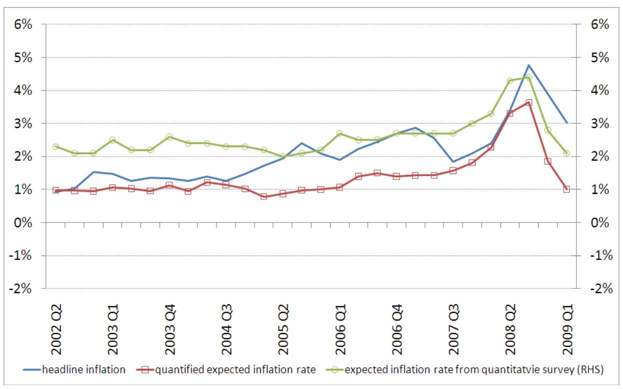


Figure 2.14. Quantitative and quantified inflation expectations in United Kingdom

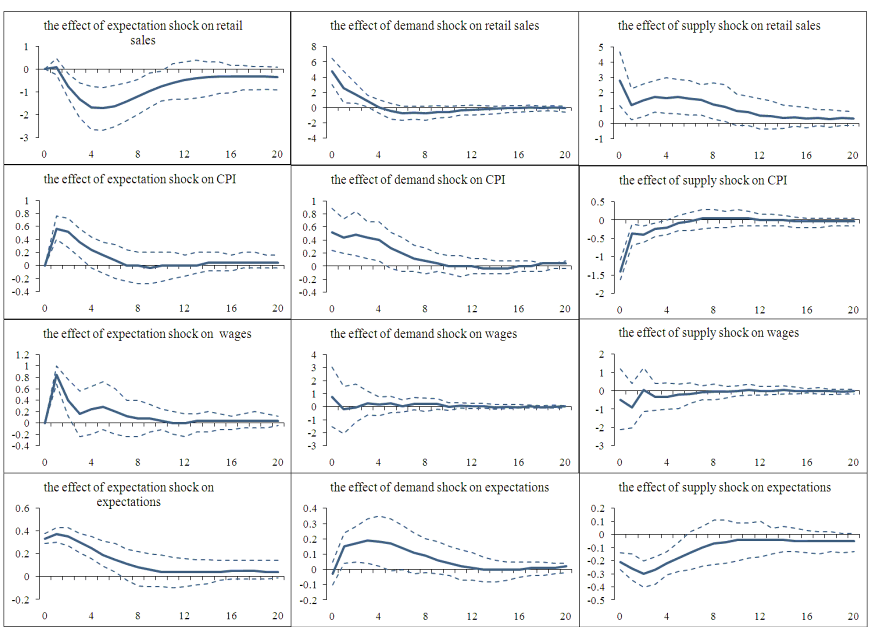


Figure 2.15. Impulse responses estimated for Czech Republic

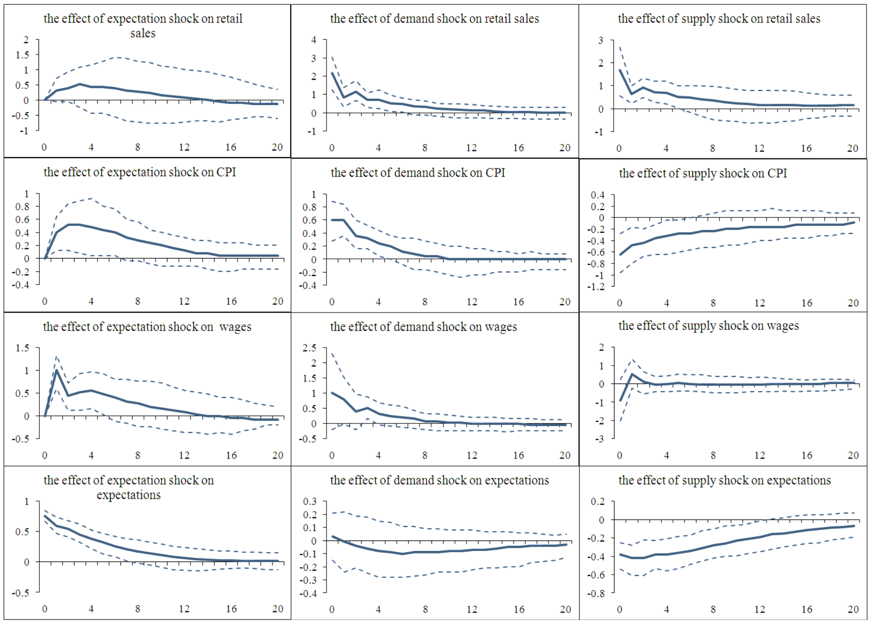


Figure 2.16. Impulse responses estimated for Hungary

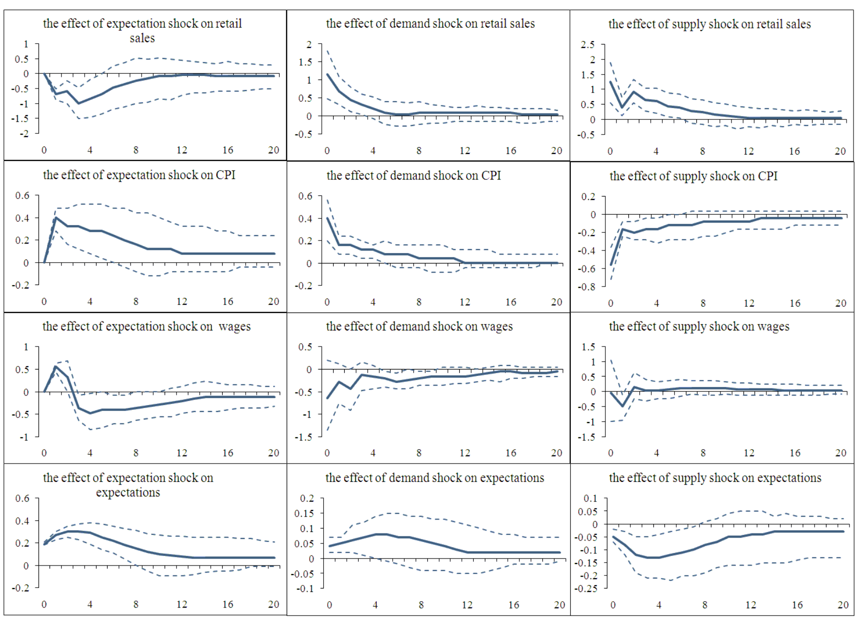


Figure 2.17. Impulse responses estimated for United Kingdom

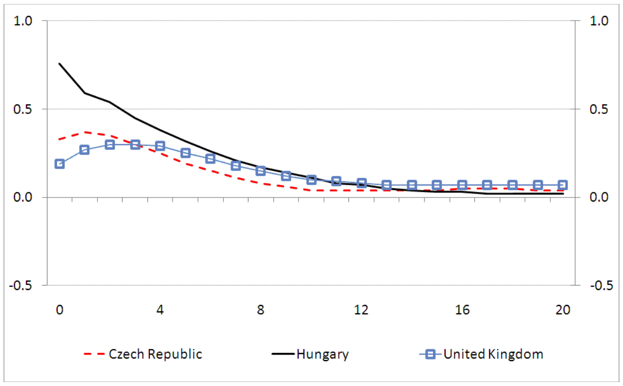


Figure 2.18. The effect of the expectation shock on inflation expectations

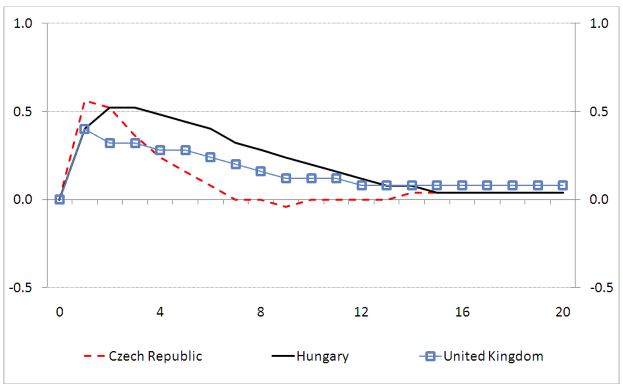


Figure 2.19. The effect of the expectation shock on inflation

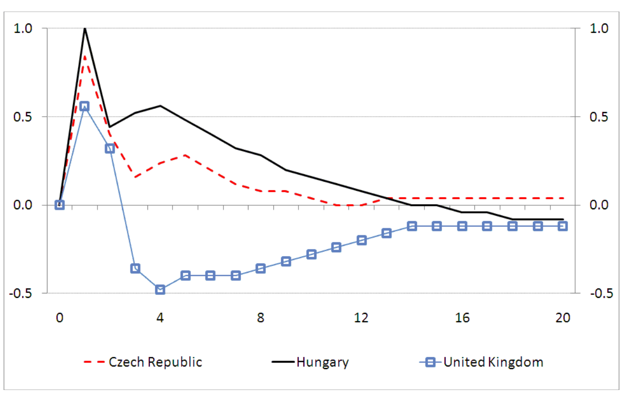


Figure 2.20. The effect of the expectation shock on nominal wage

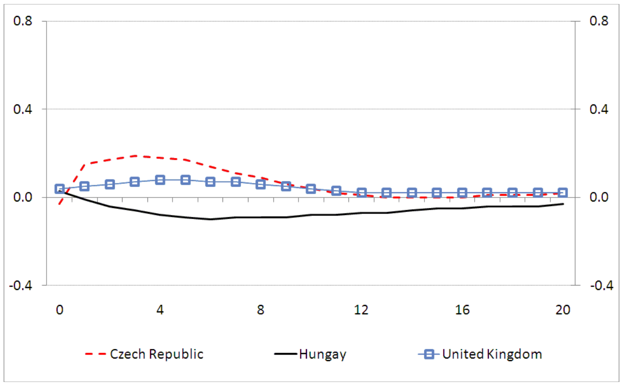


Figure 2.21. The effect of the demand shock on inflation expectations

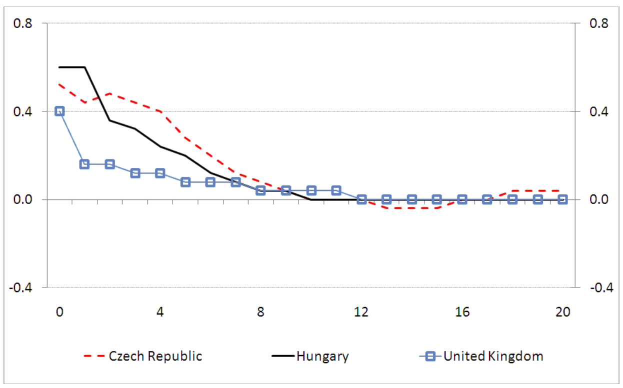


Figure 2.22. The effect of the demand shock on inflation

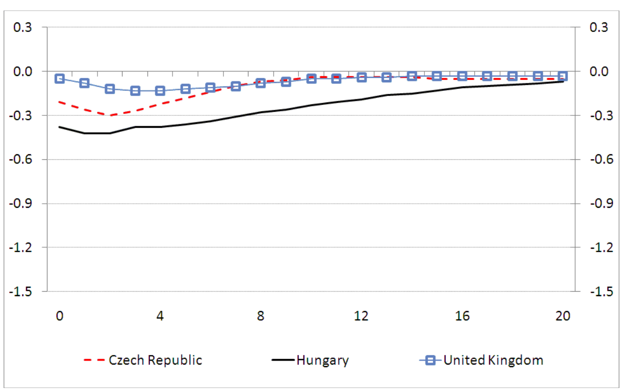


Figure 2.23. The effect of the supply shock on inflation expectations

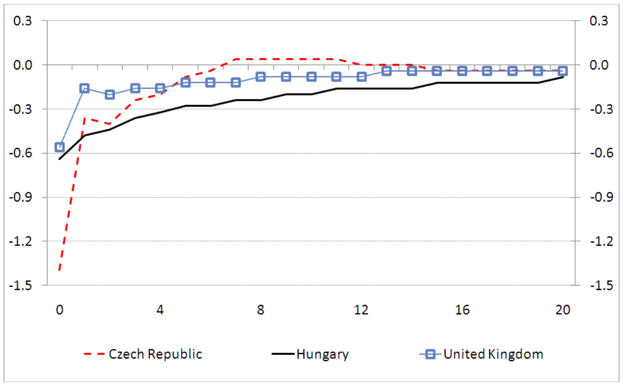


Figure 2.24. The effect of the supply shock on inflation

Chapter 3

Asymmetric exchange rate pass-through in a small menu costs model

Understanding exchange rate pass-through is crucial for small open economies. Empirical studies showed that pass-through is incomplete and bigger changes in exchange rate tend to imply higher pass-through. On the other hand empirical evidence on the asymmetry of exchange rate pass-through is mixed. This paper aims to contribute to the better understanding of asymmetry in exchange rate pass-through to import prices. First empirical evidence is provided that asymmetry is positively correlated with import price inflation. If import price inflation is high, depreciation have bigger impact on import prices than appreciations. Second I calibrate a small menu cost model for European, non-eurozone economies and calculate asymmetry in exchange rate pass-through using simulated data. Although the model is very stylized, it can reproduce also other findings of empirical studies on short run exchange rate pass-through: pass-through is incomplete; the higher the inflation rate, the higher the pass-through; the more volatile the exchange rate the smaller the pass-through.

3.1 Introduction

Understanding the behavior of import prices is crucial for small open economies. Import goods have a significant share in domestic consumption, so import price inflation is an important determinant of domestic inflation. Of particular importance is understanding the relationship between import prices and nominal exchange rates. The degree to which changes in exchange rates are reflected in local currency import prices is referred to as the degree of exchange rate pass-through. If pass-through is complete, a one percent depreciation of the domestic currency will cause import prices to increase by one percent. The pass-through of exchange rate changes to import prices has important implications for macroeconomic policy. Low pass-through implies that the effect of an exchange rate depreciation on consumer price inflation is moderated. The understanding of potential nonlinearity or asymmetry in exchange rate pass-through can help to predict changes in domestic inflation rate.

A vast empirical literature exists on how fluctuation in the exchange rate affects different price indexes. Although the studies measured pass-through with different estimation technics and used different price indexes (import, producer or consumer prices), the results show some robust findings about exchange rate pass-through. First pass-through is incomplete even on the long run, what is usually explained by pricing to market behavior of exporting firms or with the role of domestic inputs. Studies investigating price adjustment on the short run showed that pass-through is nonlinear, big changes in the exchange rate imply higher pass-through. The empirical evidence on the asymmetry of exchange rate pass-through is mixed. Several studies found that pass-through is indeed asymmetric, but the direction varied. In more developed economies appreciations in developing economies depreciations tend to have bigger impact on prices.

Exchange rate pass-through is of special interest in emerging economies, as exchange rates tend to be more volatile, pass-through tends to be higher than in developed economies, so exchange rate fluctuation affects domestic prices more. On the other hand empirical estimates of the magnitude, linearity and asymmetry of exchange rate pass-through differs considerably across countries and sample periods.

This paper aims to contribute to the better understanding of exchange rate pass-through to import prices. My particular interest is the connection between trends in exchange rate and import prices and exchange rate pass-through. The first part of the paper provides empirical evidence that there is a positive correlation between asymmetry in exchange rate pass-through and import price inflation. In the second part I calibrate a small menu cost model to analyze short-term pass-through of the exchange rate to import prices. I show that the measured pass-through is quite sensitive to different anticipated trends in exogenous variables, which may explain the surprising diversity of pass-through estimates in time and cross-section in emerging economies. Although the model is very stylized, I could replicate many stylized facts on short run exchange rate pass-through in emerging economies: pass-through is incomplete; the higher the inflation rate the higher the pass-through; if inflation rate is higher, pass-through becomes asymmetric, so depreciations have bigger impact on the price level than appreciations.

The rest of the paper is organized as follows. First I give a brief review of the literature. Second I derive the import price equation used for estimating exchange rate pass-through, describe the data and calculate an asymmetry measure for exchange rate pass-through. Third I develop a small menu cost model for import price adjustment, discuss the drivers of asymmetry in exchange rate pass-through and calibrate the model for several economies. In the last section I summarize.

3.2 Review of the literature

This paper provides evidence on the asymmetry in exchange rate pass-through and analyses the drivers of asymmetry in a menu cost model. Hence it is connected to the vast empirical and theoretical literature of exchange rate pass-through and also to a narrower group of papers, which tried to calibrate menu cost models to match empirical findings on pass-through of shocks – not necessarily exchange rate shocks – into prices.

The theoretical exchange rate pass-through literature focused on the degree of price adjustment and factors that affect pass-through. The tests on variety of goods and across countries rejected the assumption of the absolute or relative Purchasing Power Parity.

Many papers stressed the role of market organization, product segmentation, pricing to market and competition as factors explaining incomplete pass-through. Dornbusch (1987) pointed out that firms operating under imperfect competition may adjust mark-ups in response to the exchange rate shock. Devereux and Yetman (2002) used a cross-section time series model to show that there is a non-linear relationship between the degree of pass-through and inflation rate. As inflation rises, pass-through increases too but at a decelerating rate.

Another strand of the exchange rate pass-through literature emphasized that pass-through is affected by monetary policy. In a seminal paper Taylor (2000) suggested that price responsiveness to the exchange rate is positively related to the inflation rate. Bailliu and Fujii (2004) found that a transition to a low inflation environment induced by a shift in monetary policy is a cause of a declining pass-through. Devereux and Yetman (2002) built a model in which sticky prices result from menu cost of price adjustment. Monetary policy determines the average rate of inflation and exchange rate volatility. They argued that for a given menu cost, the higher the rate of inflation and the more volatile the exchange rate, the higher the pass-through effect. Campa and Goldberg (2004) showed that higher inflation and exchange rate volatility are weakly associated with higher pass-through.

Linearity and asymmetry are investigated less frequently, although they may be important drivers of inflation dynamics. Linear response to exchange rate changes means that prices react in the same way to "big" and "small" changes in the exchange rate. If there are some menu costs of price adjustment, then relatively small changes in the exchange rate may not be passed onto prices. Symmetry in exchange rate pass-through means that the magnitude of price reaction does not depend on the direction of price changes.

The theoretical literature provides explanation for asymmetric pass-through in both direction. If firms compete strategically for market shares, an appreciation of the exporter's currency will result in firms adjusting by reducing the markup, while during a depreciation they will maintain the markup and allow prices to fall. Froot and Klemperer (1989) showed that if market share matters for future profits, then firms facing an exchange rate appreciation will decide whether to raise current or future profits, depending on their perception of durability of the appreciation. Knetter (1994) argued, that if firms operate under capacity constraints, which limits the magnitude of production on the short run, it is not worthwhile to have low prices. Hence, a depreciation of the exporter's currency might result in a lower pass-through than an appreciation, for which the capacity constraint is not binding.

Turning to empirical studies on nonlinearity and asymmetry in exchange rate pass-through Mann (1986) and Goldberg (1995) found asymmetries in consumer prices' reaction to appreciations and depreciations. Pollard and Coughlin (2003) examined the symmetry in response of import prices in manufacturing to appreciations and depreciations as well as to the size of the change of the exchange rate in the US. In many industries the response was asymmetric, but the sign of the asymmetry varied. A more robust finding was that pass-through is positively related to the size of the change in the exchange rate. Herzberg et al. (2003) used aggregate data for the UK. They did not find evidence of non-linearity in the reaction of import prices. Bussiere (2007) investigated whether export and import prices of G7 countries exhibit linear and symmetrical reaction to the exchange rate movements. He found that a non-linear effect cannot be neglected, although the direction of the asymmetries and the magnitude of nonlinearity vary across countries. Wickremasinghe and Silvapulle (2004) found that in Japan import prices of manufacturing goods display a statistically significant difference in their adjustment to appreciations and depreciations and that the reaction to the former is bigger than to the latter.

There is less evidence on nonlinearity and asymmetry in import price reactions to exchange rate changes in the emerging markets. Alvarez et al. (2008) reported a weak asymmetry between the reaction of import prices to appreciations and depreciations in Chile. They found that the response after depreciations is higher than after appreciations, but the difference was only marginally significant for two out of three product categories. Przystupa and Wrobel (2009) found that the response of polish import prices to exchange rate shocks is nonlinear and symmetric, consumer price responses depend on the magnitude of the shock and the reaction after depreciation is bigger. According to Webber (1999) pass-through was higher when the importer's currency depreciated than when it appreciated in five of seven Asian countries. Mihaljek and Klau (2005) found mixed results for 14 emerging economies, but on average depreciation had bigger effect on prices.

To sum up the findings of the empirical literature on exchange rate pass-through into import prices, pass-through seems to be incomplete even on the long run, the bigger the shock the bigger the response, asymmetry is documented in many papers, but the direction varies.

The literature applying menu cost models to analyze asymmetric adjustment is also growing. In their frequently cited papers Tsiddon (1993) and Ball and Mankiw (1994) have proposed that asymmetric price adjustment to nominal shocks can be explained by the combination of price adjustment (menu) costs and drift in the nominal price level. The drift can be caused by inflation or trends in demand or input costs. More recently Karady and Reiff (2010) replicated empirical findings on asymmetric price responses after VAT changes in a menu cost framework. Menu cost models were applied also to examine the adjustment of import prices. Gopinath and Itskhoki (2009) calibrated a menu cost model for the US import sector. Flodén and Wilander (2006) used a menu cost framework to analyze exchange rate pass-through to import prices.

3.3 Deriving the import price equation

In this section I derive the import price equation which serves as the starting point for the empirical analysis. The analytical framework is the markup model, which is routinely used in studies analyzing pass-through. Assume that a representative foreign producer sets prices on n markets. The price of exports (P_i^f) to a particular importing country i as a destination specific markup (μ_i) over its marginal costs of production (C):

$$P_i^f = \mu_i C, i = 1, 2, ..., n$$

P_i^f and C are both measured in the exporting country's currency. The import price in the currency of the importing country (P_i) is obtained by multiplying the export price (P_i^f) by the nominal bilateral exchange rate (S_i):

$$P_i = S_i P_i^f = S_i \mu_i C \tag{3.1}$$

The markup is assumed to respond to competitive pressures and demand pressures in the importing country. Abstracting from competition between foreign exporters in market i, the markup can be specified as

$$\mu_i = K_i [P_i^d / S_i C]^{\lambda_i} [DP_i]^{\eta_i}, \tag{3.2}$$

where K_i is a constant, P_i^d is the price of import competing goods, and DP_i is a measure of demand pressure in market i. Both λ_i and η_i are expected to be positive. After some straightforward algebra and taking natural logarithms (denoted by lower-case letters) the import price is given by:

$$p_i = \kappa_i + (1 - \phi_i)(c + s_i) + \phi_i ph_i + \nu_i dp_i \tag{3.3}$$

where $\kappa_i = \frac{k_i}{1-\lambda_i}, \phi_i = \frac{\lambda_i}{1-\lambda_i}$ and $\nu_i = \frac{\eta_i}{1-\lambda_i}$. The pass-through coefficient, defined as the partial elasticity of the import price with respect to the exchange rate, is $(1-\phi_i)$. As long as $\phi_i > 0$, pass-through is incomplete and there is a pricing to market effect due to the presence of P_i^d in the import price equation. In the limiting case where $\phi_i = 0$ changes in the exchange rate (and foreign costs) are passed through completely and the price of competing goods has no effect on import prices.

Since C is not observable I take a geometric average of the export prices to the n markets (P_i^f) which relates to C by

$$P^f = \prod_{i=1}^{n}(P_i^f)^{\omega_i} = \prod_{i=1}^{n}(\mu_i C)^{\omega_i}, 0 \leq \omega_i \leq 1, \sum_{i=1}^{n}\omega_i = 1$$

where ω_i is the weight of market i. Taking logarithms and substituting into the previous equation the import price is

$$p_i = \kappa_i + (1-\phi_i)(p^f + s_i) + \phi_i ph_i + \nu_i dp - (1-\phi_i)\sum_{i=1}^{n}\omega_i \ln \mu_i \tag{3.4}$$

Turning to time series data for one market, introducing a time subscript and adding a stochastic disturbance term u_t I get

$$p_t = \kappa' + (1-\phi)(p_t^f + s_t) + \phi ph_t + \nu dp_t + u_t \tag{3.5}$$

Using the average export price as a proxy for marginal costs implies that the foreign exporter's (unobserved) markups in all n markets are contained in the disturbance term. At least two limitation of this model should be highlighted. First it is static and hence, does not allow for import prices to adjust gradually to changes in the explanatory variables. Second the model is a partial equilibrium model, so the coefficients are all interpretable as partial elasticities. The import price equation imposes the same rate of pass-through of exchange rates and foreign costs as well as unit homogeneity in $p^f + s$ and ph. In practice, however, these restrictions need not hold. In the short-run, exchange rates are more variable than costs, and exporters are probably more willing to absorb into their markups changes in exchange rates than changes in costs, which are likely to be permanent. Furthermore the price of import competing goods reflects also the domestic demand pressure, so empirical studies usually use only one of the two variables in the regressions used to measure exchange rate pass-through. In this study I assume that import prices are determined by foreign prices, the nominal effective exchange rate and producer prices. The last variable is used for controlling for both prices of import competing goods and demand pressure.

$$p_t = \beta_0 + \beta_1 p_t^f + \beta_2 s_t + \beta_3 ph_t + u_t \tag{3.6}$$

3.4 Data

I calculate asymmetry in exchange rate pass-through for seven European, non-eurozone economies. Taking the derived import price equation as the starting point for the empirical analysis I need data for import prices, import shares, exchange rates, foreign export prices and domestic prices of import competing products. The length of the sample period is limited by the availability of the import price indices. The index is available from 1996Q1 for Czech Republic, from 1995Q1 for Hungary, Poland, Slovakia and from 1994Q1 for the other countries. The enddate of the sample is 2009Q4 for all countries.

Import price index For import price index I use the survey based import price index from Eurostat. In empirical studies two types of indices are used for import prices: unit values and survey based measures. Unit value indices are a by-product of the collection of trade data by customs authorities, long and detailed time series are available. However, unit value indices are recognized as being prone to bias. The unit value indices record price changes not only due to price changes, but also due to quality changes and changes in the mix of products. Survey-based import price indices have the advantage that they reflect pure price changes. On the other hand the number of products used for calculating the index is smaller than in the case of unit value indices. The import price index of the Eurostat includes mostly manufactured goods, but also commodities. Statistical offices monitor the prices of these goods on a periodic basis and weight their price changes according to their trade shares, primarily based on nominal trade value shares from customs data.

Import prices in the import price index include cost, insurance and freight at the national border of the importing country, but not duties or import taxes. They are actual transaction prices (not list prices) including e.g. discounts. They are measured in the currency of the importing EU country. Transactions in other currencies have to be converted. As the price of some goods may be set in foreign currency, the import price index may reflect changes in the exchange rate even if actual prices are unchanged. This may bias the magnitude of the estimated exchange rate pass-through, however the measured asymmetry, which is the main interest of this study, is not affected.

Import shares The import shares are country weights, which were calculated using the IMF's Direction of Trade Statistics (DOT) database. Each country's weight is the imports from this country as a share of total imports from the countries in the index. The weights are calculated for every year in the sample period. The country shares are relatively stable over the sample period.

	Czech Republic	Hungary	Poland	UK	Norway	Sweden	Slovakia
Austria	5.2	7.6	2.1	0.7	0.9	1.1	5.2
Belgium	2.3	2.3	3.0	4.8	2.3	3.9	1.4
CHINA	3.4	4.9	3.2	3.5	4.0	2.4	1.7
Czech Republic	0.0	2.7	3.4	0.5	0.6	0.7	19.7
Denmark	0.7	0.6	1.8	1.2	7.2	8.3	0.5
Finland	0.6	1.0	1.6	1.1	3.5	5.7	0.6
France	4.3	4.6	5.9	8.0	4.2	5.5	3.5
Germany	31.0	26.3	26.2	13.1	13.5	18.2	21.6
Hungary	1.9	0.0	1.6	0.4	0.4	0.5	3.7
Ireland	0.5	0.5	0.5	3.7	1.4	1.4	0.3
Italy	5.1	6.3	8.1	4.4	3.6	3.3	5.4
Japan	2.1	3.4	1.6	4.1	3.5	2.5	1.3
Netherlands	3.5	3.3	4.4	6.5	4.3	6.6	2.0
Norway	0.5	0.1	1.1	3.3	0.0	7.8	0.1
Poland	0.0	2.9	0.0	0.6	1.3	1.9	3.7
Russia	5.7	7.6	7.7	1.1	2.1	1.9	12.7
Slovakia	7.3	2.3	1.6	0.1	0.1	0.2	0.0
Sweden	1.2	1.2	2.7	2.1	15.3	0.0	0.8
UK	3.2	2.7	4.1	0.0	8.0	8.3	1.9
US	3.0	2.8	2.9	11.3	6.2	5.0	1.7

Table 3.1. Import shares (percentage points, sample average)

Exchange rates The exchange rate indices are nominal effective import weighted exchange rates. The nominal exchange rate time series are taken from the IFS database and measured as average quarterly values. As in case of the import price index the trade-weighted exchange rate index is not necessarily optimal. To get a true measure of the extent of exchange rate fluctuations faced by exporters one should instead use a currency-contract-weighted exchange rate, but these weights are rarely available.

Foreign prices As foreign marginal costs are not directly observable, I use import weighted foreign price indices as proxy variables. Previous studies have employed proxies such as foreign export price indices, foreign producer price indices, foreign unit labour costs and foreign consumer price indices. Most studies collected proxy variables for a limited number of trading partners. This makes harder to interpret the pass-through parameters, so I collected proxy variables for all those trading partners, whose data was available. This means also that foreign price indices have different sources. I used the following hierarchy in selecting the foreign price proxy variable. I used export price index from the Eurostat database when it was available. These indices are calculated by national statistical offices in the same way as the import price indices. If the export price index was not available I used export unit values from the IFS database. If neither the export price index nor the unit values were available I used the producer price index as the proxy variable. The import weighted foreign export price index is constructed with the same set of weights as the exchange rate index.

Data sources	Countries
Eurostat, export price index	Austria, Belgium, Bulgaria, Croatia, Cyprus, Denmark, Estonia, Finland, France, Germany, Hungary, Ireland, Italy, Japan, Latvia, Lithuania, Netherlands, Norway, Poland, Portugal, Romania, Slovak Republic, Slovenia, Spain, Switzerland, United Kingdom, United States
CSO, export price index	Czech Republic
IFS, unit values	Australia, Brazil, Canada, Hong Kong, Greece, Israel, Singapore, Sweden, Turkey
IFS, producer price index	Indonesia, Kazakhstan, Korea, Malaysia, Mexico, Philippines, South Africa, Thailand
CSO, producer price index	Russian Federation

Table 3.2. Datasources for the foreign price index

Domestic prices of import competing products As a proxy for the domestic price of import competing goods I use producer prices. The price indices were taken from Eurostat.

The variables I use differ from the most commonly used variables in empirical studies analyzing exchange rate pass-through in several ways. First the data coverage for the foreign price and nominal effective exchange rate index is high, in terms of import share it is around 90 percent in case of the foreign price and more than 90 percent in case of the exchange rate. Second most studies use the nominal effective exchange rates published by the IMF or Eurostat. However the weights used for these indices are based on averages of export and import volumes, so these indices may be a noisy proxy for the exchange rate relevant for import prices. Because of that I opted for constructing an own import weighted index. Third many studies consider import prices statistics low quality data and use producer or consumer prices to analyze exchange rate pass-through. In contrast with these studies I use import prices as the dependent variable. The advantage of that is that pass-through to import prices is quick, the assumption of exogenous exchange rate and foreign price is more plausible, so measures of asymmetric pass-through from simple regressions can be appropriate.

	NEER			Foreign price		
	average	min	max	average	min	max
Czech Republic	91%	83%	94%	87%	80%	91%
Germany	94%	88%	95%	88%	79%	90%
Hungary	97%	96%	98%	90%	85%	94%
Poland	96%	93%	97%	90%	83%	93%
UK	91%	89%	93%	85%	80%	91%
Norway	95%	92%	96%	90%	86%	93%
Sweden	95%	87%	97%	91%	85%	95%
Slovakia	97%	95%	98%	90%	58%	95%

Table 3.3. Data coverage

Although in the empirical analysis I concentrate on short run asymmetries in exchange rate pass-through, it is useful to look at the long run characteristics of the data briefly. On the long run exchange rate and foreign price trends seem to determine import prices. The sum of the average annual change in foreign prices and exchange rate is close to the average yearly change in import prices.[1] This implies that long run pass-through of the exchange rate and foreign price is close to one. This is an important stylized fact to motivate the menu cost model discussed later, as the model implicitly assumes full long run pass-through.

	CZ	HU	PL	UK	NO	SW	SK
NEER	-64%	43%	18%	0%	-16%	3%	-42%
foreign price	49%	52%	51%	28%	34%	30%	85%
import price	-12%	81%	78%	17%	31%	51%	55%
producer price	45%	125%	100%	32%	115%	36%	90%

Table 3.4. Cumulative changes of price indexes over the sample

3.5 Measure of asymmetry in exchange rate pass-through

Empirical studies chose different ways to estimate exchange rate pass-through. The first approach is to interpret the Equation 3.6 as a long-run cointegrating relationship. Import price determination in the short run will be explained by a dynamic model where changes in import prices depend on deviations from the long-run relationship and lagged changes in the explanatory variables. The appropriateness of this method depends on the length of the available time series and the consistency of data from different sources. Usually the quality of import and export price statistics is questioned. These statistics are considered to be useful in analyzing the dynamics of the prices, but often biased in the sense that they consistently overestimates the true increases in prices.

The other approach is to estimate a simple regressions using the log differences of the variables. This approach implicitly assumes that foreign prices and the exchange rate are exogenous and it is valid to use the current and lagged log differences as explanatory variables. Empirical studies generally found quite significant short run pass-through parameters, but the significancy of lagged variables decreases rapidly if the number of lags increases. Hence the total effect of exchange rate changes is difficult to be estimated in a robust way.

This study focuses on asymmetric price adjustment after exchange rate changes. This has at least two important consequences. First as asymmetric adjustment is a short run phenomena, the long run relationship between import prices and

[1] Of course it is arbitrary to say, that the difference is small. On the other hand the difference is not systematically below or above zero, so at least the data does not provide evidence against full pass-through in the long term.

exchange rate is of less interest. Second the number of estimated coefficients doubles as separate coefficients have to be estimated for positive and negative exchange rate changes. Because of these and the relatively short samples I use a very simple measure of asymmetry in exchange rate pass-through. I estimate the following import price equation:

$$\begin{aligned} p_t^{cyc} &= \beta_0 + \beta_1^+ p_t^{f,cyc,+} + \beta_1^- p_t^{f,cyc,-} + \beta_2^+ s_t^{cyc,+} \\ &\quad + \beta_2^- s_t^{cyc,-} + \beta_3 ph_t^{cyc} + u_t \end{aligned} \tag{3.7}$$

where I filtered all variables and used the cyclical components in the regression. This approach explicitly assumes that price adjustment is asymmetric only on the short run. The measure of asymmetry is the difference between the effect of price changes for stronger and weaker exchange rate relative to the trend.

The estimated coefficients are measures of average pass-through of the cyclical component of the exchange rate. As Figure 3.15. illustrates the estimated coefficients can be understood as the average ratio of the cyclical component of import price index and the nominal effective exchange rate assuming that no other variables affect import prices. This measure does not tell much about the dynamic of exchange rate pass-through.[2] If the response is sluggish then the estimated measure of pass-through will be smaller. However as the estimated coefficients measure the pass-through of the cyclical component of exchange rate changes this issue is of less importance. If prices do not react to the cyclical component of exchange rate on the short run, then it is plausible to assume that they won't react at all. The quick response of import prices is apparent in the time series of calculated cyclical components and also according to the results of other empirical studies the pass-through is quick in case of export prices. To test the speed of price response more formally I estimated Equation 3.7 with different lags of the explanatory variables. The explanatory power of the regression (Schwarz criterion) for each economy was the highest when I used the immediate changes of the explanatory variables.[3]

[2] In Figure 15. pass-through of appreciation can be understood as the average value of A/B, and the pass-through of depreciation as the average value of C/D. The measure of asymmetry is the difference of these two ratios. This measure does not tell much about the dynamic of exchange rate pass-through, and as it is illustrated in Figure 15. delay in price adjustment may influence its value. Fortunately in case of import price the price response is quick, so the proposed measure is able to capture asymmetry in price response.

[3] Also producer prices may affect import prices asymmetrically. I introduced different parameters for producer price above and below its trend, and re-estimated equation 3.7. The new parameters were insignificant and other parameters were only negligibly affected.

	NEER depreciation	NEER apreciation	foreign price increase	foreign price decrease	producer price
Czech Republic	0.28	0.55	0.91	0.80	0.16
	(0.15)	(0.16)	(0.31)	(0.36)	(0.22)
Hungary	0.81	0.53	0.16	0.87	0.27
	(0.12)	(0.13)	(0.26)	(0.22)	(0.24)
Poland	0.52	0.32	-0.40	0.06	0.26
	(0.19)	(0.15)	(0.42)	(0.64)	(0.52)
UK	0.49	0.54	0.69	0.70	0.59
	(0.04)	(0.05)	(0.14)	(0.11)	(0.21)
Norway	0.63	0.53	0.53	0.54	0.04
	(0.09)	(0.09)	(0.12)	(0.13)	(0.04)
Sweden	0.41	0.30	(0.78)	1.03	0.27
	(0.05)	(0.06)	(0.10)	(0.12)	(0.09)
Slovakia	0.29	0.10	0.49	1.02	0.29
	(0.13)	(0.15)	(0.15)	(0.25)	(0.15)

Table 3.5. Pass-through into import prices

Table 3.5. shows the estimated parameters for each economies.[4] According to the estimated coefficients the pass-through of the cyclical component of NEER to import prices is incomplete. Empirical studies found that pass-through of foreign price changes tend to be higher than the pass-through of the exchange rate, which is explained by the bigger volatility of the exchange rate. The estimated coefficients of the import price equations are only partly in line with these findings. In Hungary, Poland and Norway the pass-through of exchange rate is as big as or even bigger than pass-through of foreign prices.

The exchange rate pass-through is asymmetric in most of the economies. Depreciation has bigger effect on prices in case of Hungary, Poland, Norway, Sweden and Slovakia, in case of Czech Republic and United Kingdom it was the other way. The asymmetry in exchange rate pass-through seems to be positively correlated with both the average annual growth rate of import prices and average NEER depreciation. Where the import price index was high and NEER depreciated on average, exchange rate depreciation tended to have bigger impact on prices than appreciation (Figure 3.16-3.17).[5]

[4] Standard errors are in parenthesis below the estimated parameters.

[5] The cyclical component was calculated with the Hodrick-Prescott filter. The results may be sensitive to the filtering method, so I calculated the cyclical components also with the Baxter-King filter. The cyclical components were quite similar to what I got with the HP filter. I also re-estimated equation 3.7 with the new cyclical components. The correlations of asymmetry in exchange rate pass-through with annual growth rate of import prices and average NEER depreciation remained positive as illustrated in Figure 3.18-3.19.

3.6 Simple menu cost model for analyzing asymmetric pass-through

To explore asymmetric exchange rate pass-through I develop a simple menu cost model for the import sector. The model is similar to the model in Flodén and Wilander (2006). However they examined pass-through only at the individual firm level. In this paper I analyze the problem in a general equilibrium setup, so strategic interactions among firms also play a role. The most important feature of the model is that economic agents have relatively sophisticated perceptions about the exchange rate process. They assume that exchange rate dynamics have a permanent and a temporary element and they set prices taking that into account.

In the import sector there are two types of firms. The importers have an extremely simple production technology. They buy foreign goods abroad and sell them to the aggregate import good producer. The aggregate import good producer produces the aggregate good with the following packaging technology:

$$Y_t^F = \left(\int_0^1 (Y_t(f))^{\frac{\varepsilon_M - 1}{\varepsilon_M}} df\right)^{\frac{\varepsilon_M}{\varepsilon_M - 1}} \tag{3.8}$$

For sake of simplicity the demand for the aggregate import good (Y_t^F) is assumed to be constant. The relative demand for foreign goods is determined by their relative price:

$$Y_t(f) = \left(\frac{P_t^F}{P_t(f)}\right)^{-\varepsilon_M} Y_t^F, \tag{3.9}$$

where P_t^F is the price index of the foreign goods.

$$P_t^F = \left[\int_0^1 P_t(f)^{1-\varepsilon_M} df\right]^{\frac{1}{1-\varepsilon_M}} \tag{3.10}$$

Importers of foreign goods are facing exchange rate shock, aggregate and idiosyncratic foreign price shocks, and given these shocks they maximize profit. Prices are set in home currency and can be changed only with menu cost. The aggregate foreign price is assumed to be a random walk process with drift:

$$\log(P_t^*) = \log(P_{t-1}^*) + N(\mu_{agg}^F, \sigma_{agg}^F) \tag{3.11}$$

Importers face also idiosyncratic foreign price shocks, which follow an AR(1) process, so the idiosyncractic import price will be given by:

$$\log(P_t^*(f)) = \log(P_t^*) + \mu_t(f), \tag{3.12}$$

where the idiosyncratic import price shock ($\mu_t(f)$) follows an AR(1) process

$$\mu_t(f) = \rho^F_{t-1}(f) + N(0, \sigma^F) \tag{3.13}$$

The nominal exchange rate process (S_t) is assumed to follow a trend (given by the drift μ_S) and to have also a temporary element (μ^S_t). The exchange rate process is driven by exogenous exchange rate shocks. Due to these shocks real exchange rate may deviate from its steady state value. However it is assumed that on the long run these deviations disappear. This happens partly due to the reaction of prices to exchange rate movements, and partly as temporary exchange rate shocks decay over time.

$$\log(S_t) = \log(S_{t-1}) + \mu_S + \mu^S_t, \tag{3.14}$$

where the exchange rate shock (μ^S_t) follows an AR(1) process

$$\mu^S_t = \rho_S \mu^S_{t-1} + N(0, \sigma^S) \tag{3.15}$$

The problem of the foreign good importer is to set price given the aggregate foreign price, idiosyncratic foreign price and the exchange rate. Price can be changed only with menu cost (κ), which is expressed in terms of the imported good.

$$\max E_t \left\{ \sum_{\tau=t}^{\infty} \beta^{\tau-t} \Pi_t(f) \right\} \tag{3.16}$$

$$\Pi_t(f) = P_t(f)Y_t(f) - S_t P^*_t(f) Y_t(f) - \kappa S_t P^*_t(f) I_t(f) \tag{3.17}$$

Dividing by the price level (P^F_t), we get the real profit function. As it is assumed that on the long run the law of one price holds, the real profit follows a stationary process.

$$\Pi^R_t(f) = \frac{\Pi_t(f)}{P^F_t} = \frac{P_t(f)}{P^F_t} Y_t(f) - \frac{S_t P^*_t(f)}{P^F_t} Y_t(f) - \kappa \frac{S_t P^*_t(f)}{P^F_t} I_t(f) \tag{3.18}$$

The state space of the firm's problem is infinite dimensional since the evolution of the price level of imported foreign goods depend on the entire joint distribution of all firms' prices and on the exchange rate. Following Nakamura and Steinsson (2009), to make the problem tractable it is assumed that the firms perceive the evolution of the price level as being a function of a small number of moments of this distribution. Specifically, I assume that firms perceive that

$$\frac{P^F_t}{P^F_{t-1}} = \Gamma\left(\frac{S_t P^*_t}{P^F_{t-1}}\right) \tag{3.19}$$

Given these assumptions the firm's optimization problem can be written recursively in the form of the Bellman equation:

$$\begin{aligned} & V(\frac{P_{t-1}(f)}{P^F_{t-1}}; \mu^S_t; \frac{S_t P^*_t}{P^F_{t-1}}; \frac{P^*_t(f)}{P^*_t}) \\ = \;& \max(\Pi^R_t(f) + E_t[D^R_{t,t+1} V(\frac{P_t(f)}{P^F_t}; \mu^S_{t+1}; \frac{S_{t+1}P^*_{t+1}}{P^F_t}; \frac{P^*_{t+1}(f)}{P^*_{t+1}})]), \end{aligned} \tag{3.20}$$

where the state variables are the relative price (of the output), the temporary element of the nominal exchange rate, the real exchange rate and the relative price of the imported good (input).

The equilibrium of this economy can be characterized with the stochastic process of the endogenous price variable that is consistent with profit maximization, and the evolution of the exogenous variables S_t, P^*_t, $P^*_t(f)/P^*_t$. I use the following iterative procedure to solve for the equilibrium:

1) I specify a finite grid of points for the state variables, $P_{t-1}(f)/P^F_t$; μ^S_t; $S_t P^*_t / P^F_{t-1}$ and $P^*_t(f)/P^*_t$.

2) I propose a function $\Gamma(S_t P^*_t / P^F_{t-1})$.

3) Given the proposed $\Gamma()$, I solve for the policy function F by value function iteration on the grid.

4) I check whether $\Gamma()$ and F are consistent. If so, I stop and use $\Gamma()$ and F to calculate other features of the equilibrium. If not, I update $\Gamma()$ and go back to step 3.

The numerical solution of the model is quite standard. The only thing what makes it somewhat more complicated is that the processes of two state variables (the real exchange rate and the temporary part of nominal exchange rate changes are not independent, as the value of S_{t+1} affected also by μ^S_{t+1}.

3.7 The drivers of asymmetry and its consequences

Before calibrating the model, it is worth to make some simulation exercises to make it clear how the model works, and what drives asymmetry in the exchange rate pass-through. In the baseline scenario I assume that the aggregate foreign price have no trends, the yearly depreciation rate of the nominal exchange rate is two percent and all other parameters are the averages of the calibrated parameters for each economies.[6] Additional to the benchmark parametrization I make sensitivity analysis for some of the parameter. After solving the model for each parameter sets I simulate the model economies and estimate asymmetry in the exchange rate adjustment. For the simulation exercise I generated random idiosyncratic input price shocks for 1000 importers, exchange rate shocks and aggregate input price shocks for 1500 periods. Using the policy function from the model I derived individual prices for each firm for every period and calculated the aggregate import price level. I cut off the first 500 observations and estimated asymmetry in price adjustment using the time series of the aggregate variables. The asymmetry is measured exactly the same way as it was done in the empirical section. First I detrend the simulated time series of the variables with Hodrick-Prescott filter, then I estimate pass-through using the cyclical components.

The asymmetry is driven by the trend in the aggregate price level of the final import good. This trend is determined by the trend in nominal exchange rate and aggregate foreign price. Table 3.6 illustrates the effect of trend in the nominal exchange rate. The average pass-through is the lowest, when the trend is zero. If the absolute value of the trend increases, the magnitude of the average pass-through also increases. The asymmetry is positive if the nominal exchange rate depreciates and negative if it appreciates.

This simple model is able the replicate the stylized fact, that the volatility of inflation is higher if the inflation rate is higher. This feature is attributed to that the higher the inflation rate , the quicker prices change after shocks and even the temporary ones will affect prices considerably.[7]

	trend in nominal exchange rate						
	-0.02	-0.01	0	0.01	0.02	0.03	0.04
average pass-through	0.83	0.70	0.54	0.67	0.82	0.90	0.92
asymmetry	-0.07	-0.04	-0.02	0.04	0.05	0.07	0.08
standard deviation of inflation	1.8%	1.4%	1.1%	1.4%	1.7%	2.0%	2.1%

Table 3.6. Consequences of trend in the nominal exchange rate

Exchange rate pass-through is also affected by how temporary are the changes in the exchange rate. If exchange rate changes are not considered to be permanent, then economic agents may opt to hold prices unchanged. This decreases the

[6] The calibration is explained in the next section.

[7] Considering welfare some studies regard the high volatility of inflation more harmful than its level, so the positive correlation of inflation rate and its volatility have important policy implication. If central bank ensure low inflationary environment, it can decrease also volatility.

magnitude of average of pass-through. However (if the exchange has a depreciating trend) the pass-through of appreciations decreases more than pass-through of depreciations, so asymmetry increases. If shocks are less persistent, then the volatility of inflation also decreases.

	persistency of exchange rate shocks (rho_S)							
	0.55	0.6	0.65	0.7	0.75	0.8	0.85	0.9
average pass-through	0.39	0.45	0.48	0.50	0.59	0.61	0.64	0.70
asymmetry	0.15	0.14	0.13	0.11	0.09	0.07	0.06	0.04
standard deviation of inflation	0.9%	1.0%	1.0%	1.1%	1.2%	1.2%	1.2%	1.3%

Table 3.7. Consequences of less persistent exchange rate shocks

If menu cost is low then import prices react quickly to exchange rate shocks and pass-through is high. In this case the expected path of the exchange rate is less important, so asymmetry also decreases. If menu costs are higher then adjustment becomes asymmetric.

	Size of menu cost				
	0.01	0.02	0.03	0.04	0.05
average pass-through	0.79	0.67	0.57	0.51	0.45
asymmetry	-0.03	0.00	0.06	0.08	0.19
standard deviation of inflation	0.9%	1.0%	1.0%	1.1%	1.2%

Table 3.8. The consequencies of menu costs

This simple model can replicate many findings of empirical research on exchange rate pass-through in Central-East European economies (e.g. Darvas (2001)).[8] In these economies pass-through decreased considerably parallel to the decrease in inflation rate and after the depreciating trends in the exchange rates had been broken. The volatility of the inflation rate also decreased parallel to the decrease in exchange rate pass-through.

3.8 Apply the model to explain different estimates of pass-through

To evaluate how this simple model can explain the differences in asymmetry of pass-through among the examined seven countries I calibrated the model for each economies. The parameters of the exchange rate and foreign price process for each country were set according the result of simple regressions. The trend

[8]Most studies focused on exchange rate pass-through to consumer prices, so only the qualitative results can be compared.

of the exchange rate process was set to be equal with the average quarterly depreciation in the sample (Table 3.4). The autoregressive parameter and the standard deviation of the temporary exchange rate shocks were estimated using the cyclical component of the nominal exchange rate process for each economy:

$$s_t^{cyc} = c + \rho s_{t-1}^{cyc} + u_t \tag{3.21}$$

The trend of the foreign price process was set to be equal with the average quarterly increase in the sample. In case of foreign prices the standard deviation of foreign price shocks was calibrated to be equal to the standard deviation of the error term in the following regression. The estimated φ parameter is close to 1 in all countries, which underpins the previous assumption, that the foreign price process is a random walk with drift.

$$p_t^f = c + \varphi p_{t-1}^f + v_t \tag{3.22}$$

The elasticity of demand was set to 4, which was chosen as it is close to the average of values used by other studies, which calibrated menu cost models.[9]

			CZ	HU	UK	PL	SW	NO	SK
calibrated parameters	NEER	trend	-1.2%	0.6%	0.0%	0.3%	0.0%	-0.3%	-0.9%
	NEER (cyclical)	AR parameter	0.71	0.71	0.83	0.76	0.74	0.66	0.69
	NEER (cyclical)	standard deviation	3.0%	2.9%	2.6%	3.7%	2.3%	2.3%	1.9%
	foreign price	trend	0.9%	0.7%	0.4%	0.8%	0.4%	0.5%	1.0%
		standard deviation	1.3%	1.1%	1.1%	1.1%	1.2%	1.4%	1.4%
	demand elasticity		4	4	4	4	4	4	4

Table 3.9. Calibrated parameters

Three parameters, the menu cost, the autoregressive parameter and the standard deviation of the idiosyncratic foreign price shock were chosen to match the autoregressive parameter and standard deviation of import prices and the value of average pass-through of the cyclical component of NEER.

[9] Berry, Levinsohn, and Pakes (1995) and Nevo (2001) find that markups vary a great deal across firms. The value of θ I choose implies a markup similar to the mean markup estimated by Berry, Levinsohn, and Pakes (1995) but slightly below the median markup found by Nevo (2001). Broda and Weinstein (2006) estimate elasticities of demand for a large array of disaggregated products using trade data. They report a median elasticity of demand below three. Also, Burstein and Hellwig (2006) estimate an elasticity of demand near five using a menu cost model. Midrigan (2006) uses $\theta = 3$, whereas Golosov and Lucas (2007) use $\theta = 7$.

		CZ	HU	UK	PL	SW	NO	SK
import price (cyclical)	AR parameter	0.78	0.73	0.85	0.35	0.86	0.84	0.79
	standard deviation	2.0%	2.5%	1.4%	3.6%	1.4%	1.4%	1.6%
pass-through	average value	0.41	0.71	0.58	0.41	0.41	0.62	0.20

Table 3.10. Matched statistics

To set the values of the menu cost, the autoregressive parameter and the standard deviation of the idiosyncratic foreign price shock the model was simulated with different parameter sets. That parameter set was chosen for each economy which minimized the sum of normalized squared difference between simulated and actual statistics. The calibrated values for menu cost, the autoregressive parameter and the standard deviation of the idiosyncratic foreign price shock are plausible and similar to values used by other studies.

		CZ	HU	UK	PL	SW	NO	SK
menu cost		0.04	0.02	0.04	0.04	0.04	0.02	0.03
idiosyncratic foreign price	AR parameter	0.70	0.75	0.80	0.70	0.75	0.90	0.70
	standard deviation	0.02	0.02	0.02	0.02	0.02	0.03	0.02

Table 3.11. Parameters calibrated to match statistics

The simulated statistics are relatively close to the actual ones, although the model tends to produce less volatile import prices. Taking into account that the model is quite simple and import prices may be influenced by factors not directly addressed here, the underestimation of import price volatility is acceptable.

		CZ	HU	UK	PL	SW	NO	SK
import price (cyclical)	AR parameter	0.83	0.75	0.86	0.77	0.86	0.83	0.86
	standard deviation	1.4%	2.0%	1.3%	2.3%	1.0%	1.3%	0.6%
pass-through	average value	0.49	0.71	0.54	0.67	0.45	0.61	0.18

Table 3.12. Simulated statistics

Studies using menu cost models report routinely frequencies and average size of price changes. According to the calibrated models 11-22 percent of firms changes price quarterly and the average size of price changes is between 6 and 11 percentage point

	CZ	HU	UK	PL	SW	NO	SK
Frequency of price change	0.167	0.220	0.138	0.190	0.140	0.184	0.105
Average size of price changes	0.094	0.088	0.098	0.114	0.084	0.091	0.058

Table 3.13. Frequency and price statistics (percentage points)

As I do not have data on firm level import prices, it is difficult to assess, wether the calculated frequency and price statistics have plausible values. Comparing the frequencies and sizes of price changes to values found by other studies may provide some guidance. The selected studies reported statistics on different product categories. As in the sample manufacturing goods are overrepresented, probably statistics on producer prices or prices of non-food non-energy consumer goods can serve as appropriate benchmarks. The average size of price changes calculated from the calibrated models seems to be within the range of statistics reported by other studies, while calculated frequencies seem to be lower.[10] Taking into account that I used quarterly data to calibrate the model, the difference between frequencies are quite natural. The temporary changes in NEER or foreign prices may have shorter impact on prices than three months, so quarterly frequencies calculated from monthly frequencies can be much higher, than frequencies from the simulated model.

	frequency of price change (%)	average size of price change (pp.)
Nakamura-Steinsson (2008, US)		
producer prices (US)	20.1	7.7
regular consumer prices	28.3	8.5
Dhyne et al. (2005, eurozone)		
non-energy industrial goods	25.1	9.4
Vermeulen et al. (2007, eurozone)		
non-food non-energy consumer goods (producer price)	31.9	3.7
non-food non-energy consumer goods (consumer prices)	24.6	10.7

Table 3.14. Frequency and size statistics for US and the eurozone

The main interest of this study was to examine, whether a small, calibrated menu cost model is able to reproduce asymmetries in price adjustment after exchange rate changes. According the results it can. Except United Kingdom the calculated asymmetry from the simulated models are very similar to the estimated ones.

	CZ	HU	UK	PL	SW	NO	SK
fx depreciation	0.42	0.83	0.62	0.83	0.53	0.63	0.23
fx appreciation	0.53	0.60	0.47	0.52	0.37	0.59	0.13
asymmetry	-0.11	0.23	0.14	0.31	0.17	0.04	0.10
estimated asymmetry	-0.28	0.27	-0.05	0.20	0.12	0.10	0.19

Table 3.15. Exchange rate pass-through and its asymmetry calculated from the calibrated models

[10] Monthly frequency (x) reprted by the studies is converted into quarterly frequency (q) by using the formula: $q = 1 - (1 - x)^3$.

3.9 Summary

This study explored the relationship between import price inflation and asymmetric exchange rate pass-through. First I provided empirical evidence that the asymmetry in exchange rate pass-through is positively correlated with import price inflation. Comparing pass-through in seven non-eurozone, European countries I showed that the higher the import price inflation rate, the bigger the difference between price adjustment after depreciations and appreciations. In the second part of the paper I calibrated small menu cost model for each countries and showed that the correlation between asymmetry in price adjustment and import price inflation is positive if price adjustment is costly. Furthermore changing trend of the nominal exchange rate can explain the stylized facts that inflation rate is positively correlated with both the volatility of the inflation rate and average exchange rate pass-through, which were documented by many empirical papers.

Bibliography

[1] Roberto Álvarez & Patricio Jaramillo & Jorge Selaive, 2008. "Exchange Rate Pass-Through into Import Prices: The Case of Chile," Working Papers Central Bank of Chile 465, Central Bank of Chile.

[2] Eiji Fuji & Jeannine Bailliu, 2004. "Exchange Rate Pass-Through and the Inflation Environment in Industrialized Countries: An Empirical Investigation," Computing in Economics and Finance 2004 135, Society for Computational Economics.

[3] Ball, Laurence & Mankiw, N Gregory, 1994. "Asymmetric Price Adjustment and Economic Fluctuations," Economic Journal, Royal Economic Society, vol. 104(423), pages 247-61, March.

[4] John Beirne & Martin Bijsterbosch, 2009. "Exchange Rate Pass-through in Central and Eastern European Member States," Working Paper Series 1120, European Central Bank.

[5] Ariel Burstein & Martin Eichenbaum & Sergio Rebelo, 2005. "Modeling Exchange Rate Passthrough After Large Devaluations," RCER Working Papers 514, University of Rochester - Center for Economic Research (RCER).

[6] Matthieu Bussiere, 2007. "Exchange rate pass-through to trade prices - the role of non-linearities and asymmetries," Working Paper Series 822, European Central Bank.

[7] Campa, José Manuel & Goldberg, Linda S, 2004. "Exchange Rate Pass-Through into Import Prices," CEPR Discussion Papers 4391, C.E.P.R. Discussion Papers.

[8] Darvas, Zsolt, 2001. "Exchange rate pass-through and real exchange rate in EU candidate countries, Discussion Paper Series 1: Economic Studies 2001,10, Deutsche Bundesbank, Research Centre.

[9] Michael B. Devereux & Charles Engel, 2002. "Exchange Rate Pass-Through, Exchange Rate Volatility, and Exchange Rate Disconnect," NBER Working Papers 8858, National Bureau of Economic Research, Inc.

[10] Michael B. Devereux & James Yetman, 2002. "Price Setting and Exhange Rate Pass-Through," Working Papers 222002, Hong Kong Institute for Monetary Research.

[11] Devereux, Michael B. & Yetman, James, 2010. "Price adjustment and exchange rate pass-through," Journal of International Money and Finance, Elsevier, vol. 29(1), pages 181-200, February.

[12] Rudiger Dornbusch, 1987. "Exchange Rates and Prices," NBER Working Papers 1769, National Bureau of Economic Research, Inc.

[13] Flodén, Martin & Wilander, Fredrik, 2004. "State Dependent Pricing and Exchange Rate Pass-Through," Working Paper Series 174, Sveriges Riksbank (Central Bank of Sweden).

[14] Floden, Martin & Wilander, Fredrik, 2006. "State dependent pricing, invoicing currency, and exchange rate pass-through," Journal of International Economics, Elsevier, vol. 70(1), pages 178-196, September

[15] Peter Karadi & Adam Reiff, 2010. "Inflation asymmetry, menu costs and aggregation bias: A further case for state dependent pricing," MNB Working Papers 2010/3, Magyar Nemzeti Bank (The Central Bank of Hungary).

[16] Froot, Kenneth A & Klemperer, Paul D, 1989. "Exchange Rate Pass-Through When Market Share Matters," American Economic Review, American Economic Association, vol. 79(4), pages 637-54, September.

[17] Gita Gopinath & Oleg Itskhoki, 2008. "Frequency of Price Adjustment and Pass-through," NBER Working Papers 14200, National Bureau of Economic Research, Inc.

[18] Valerie Herzberg & George Kapetanios & Simon Price, 2003. "Import prices and exchange rate pass-through: theory and evidence from the United Kingdom," Bank of England working papers 182, Bank of England.

[19] Knetter, Michael M., 1994. "Is export price adjustment asymmetric?: evaluating the market share and marketing bottlenecks hypotheses," Journal of International Money and Finance, Elsevier, vol. 13(1), pages 55-70, February.

[20] Catherine L. Mann, 1986. "Prices, profit margins, and exchange rates," Federal Reserve Bulletin, Board of Governors of the Federal Reserve System (U.S.), issue Jun, pages 366-379.

[21] Dubravko Mihaljek & Marc Klau, 2008. "Exchange rate pass-through in emerging market economies: what has changed and why?," BIS Papers chapters, in: Bank for International Settlements (ed.), Transmission mechanisms

for monetary policy in emerging market economies, volume 35, pages 103-130 Bank for International Settlements.

[22] Monacelli, Tommaso (2005) "Monetary Policy in a Low Pass-through Environment." Journal of Money, Credit and Banking, 37(6), 1047-66.

[23] Emi Nakamura & Jón Steinsson, 2008. "Monetary Non-Neutrality in a Multi-Sector Menu Cost Model," NBER Working Papers 14001, National Bureau of Economic Research, Inc.

[24] Patricia S. Pollard & Cletus C. Coughlin, 2004. "Size matters: asymmetric exchange rate pass-through at the industry level," Working Papers 2003-029, Federal Reserve Bank of St. Louis.

[25] Przystupa, Jan & Wróbel, Ewa, 2009. "Asymmetry of the exchange rate pass-through: An exercise on the Polish data," MPRA Paper 17660, University Library of Munich, Germany.

[26] Taylor, John B., 2000. "Low inflation, pass-through, and the pricing power of firms," European Economic Review, Elsevier, vol. 44(7), pages 1389-1408, June.

[27] Tsiddon, Daniel, 1993. "The (Mis)Behaviour of the Aggregate Price Level," Review of Economic Studies, Blackwell Publishing, vol. 60(4), pages 889-902, October.

[28] Webber, A., 1999. "Dynamic and Long Run Responses of Import Prices to the Exchange Rate in the Asia-Pacific," Economics Working Papers WP99-11, School of Economics, University of Wollongong, NSW, Australia.

[29] Guneratne Banda Wickremasinghe & Param Silvapulle, 2004. "Exchange Rate Pass-Through to Manufactured Import Prices: The Case of Japan," International Trade 0406006, EconWPA.

APPENDIX

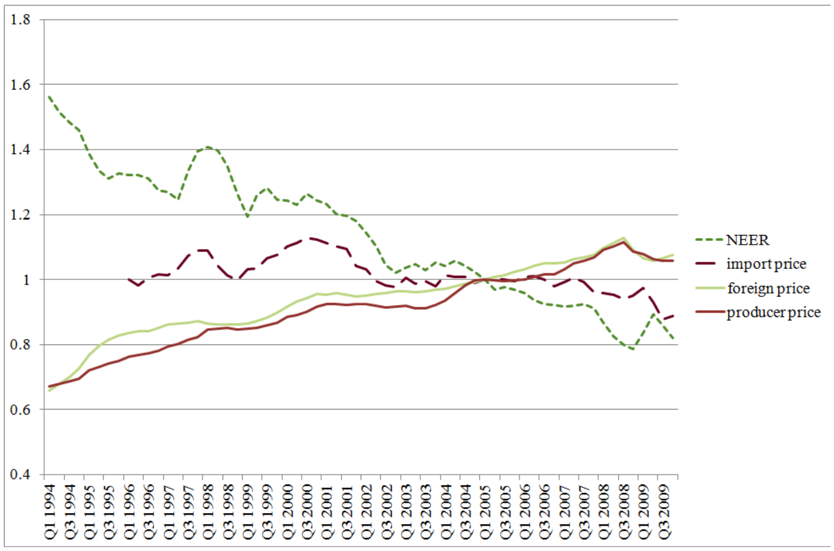


Figure 3.1. NEER and price indexes - Czech Republic

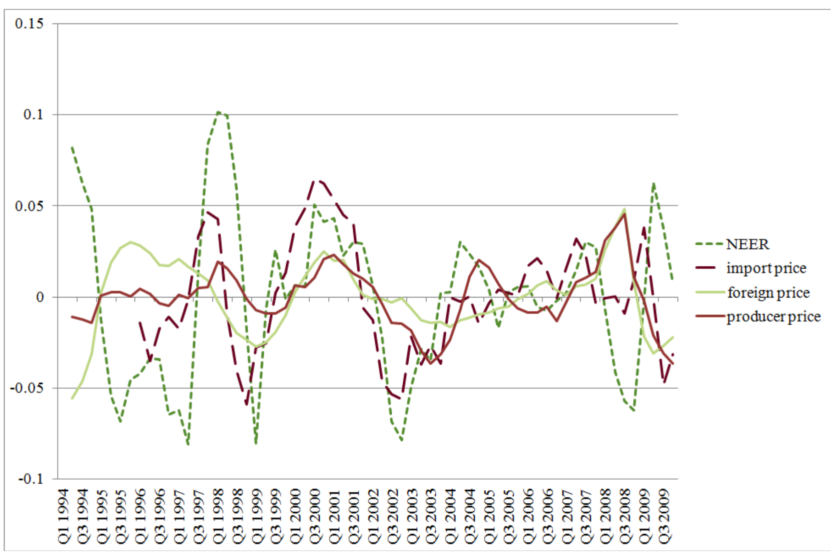
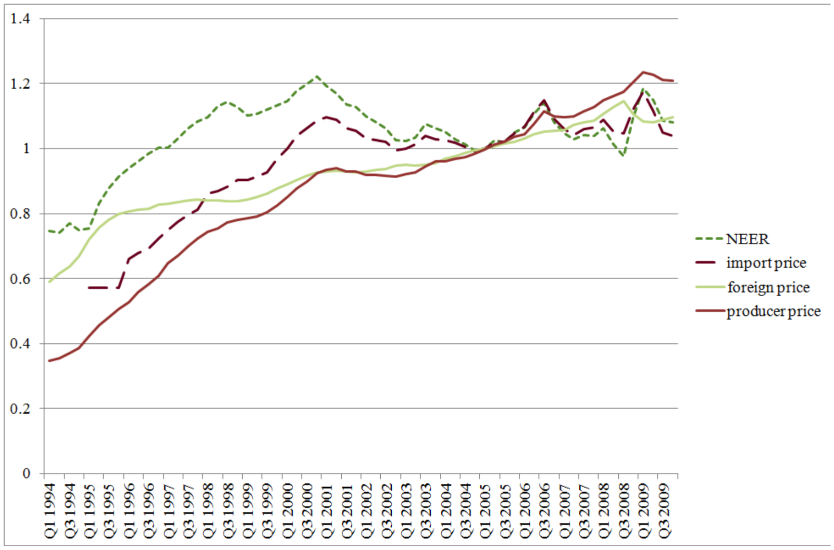


Figure 3.2. Cyclical component - Czech Republic



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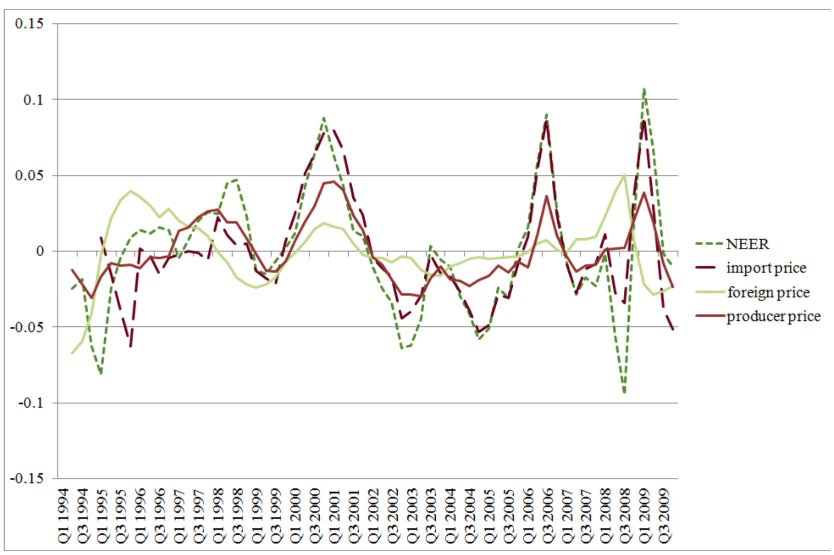

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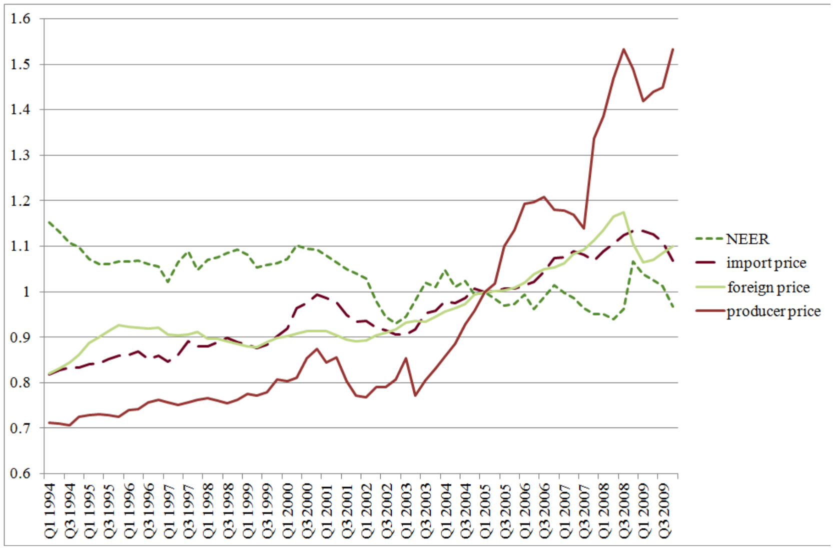


Figure 3.5. NEER and price indexes - Norway

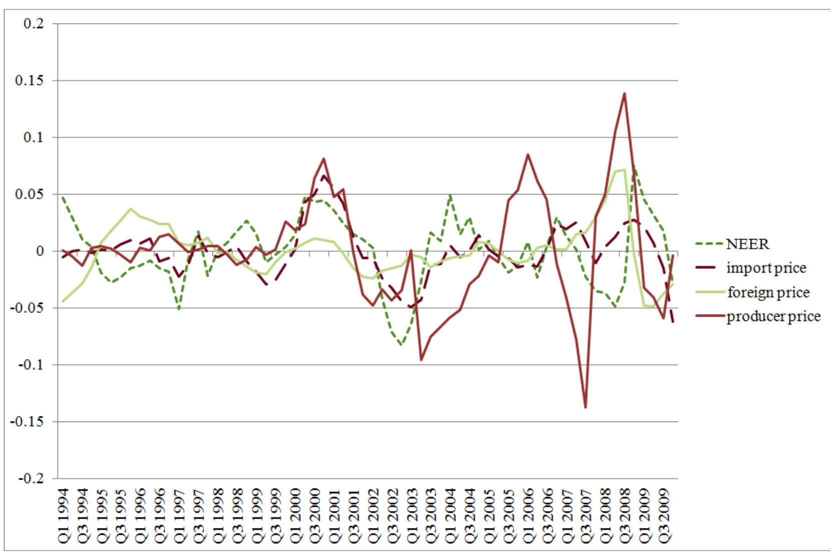


Figure 3.6. Cyclical component - Norway

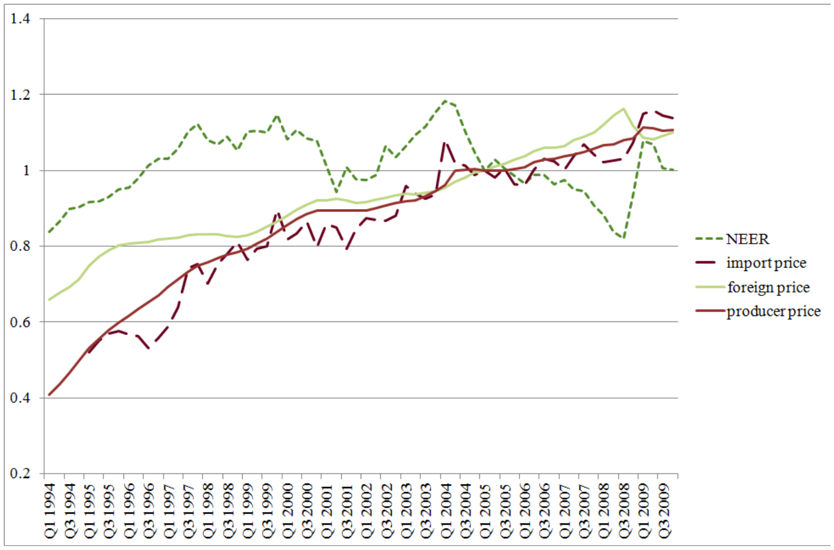
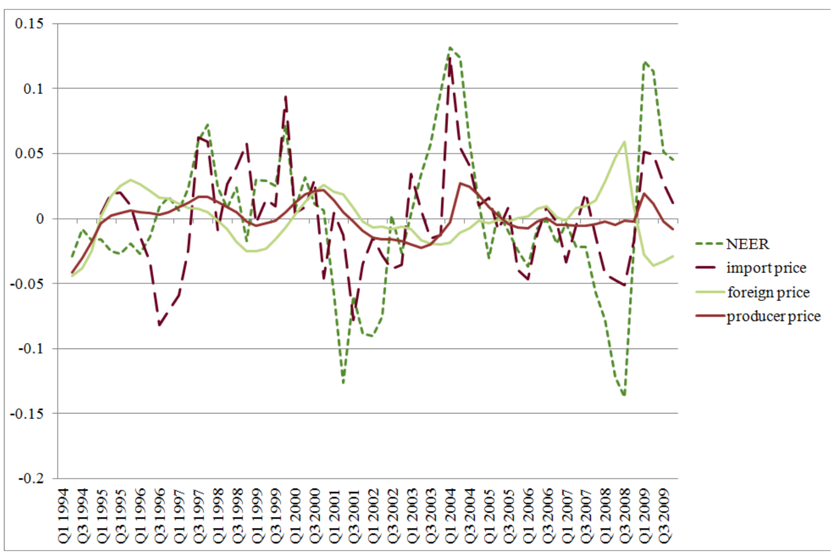


Figure 3.7. NEER and price indexes - Poland



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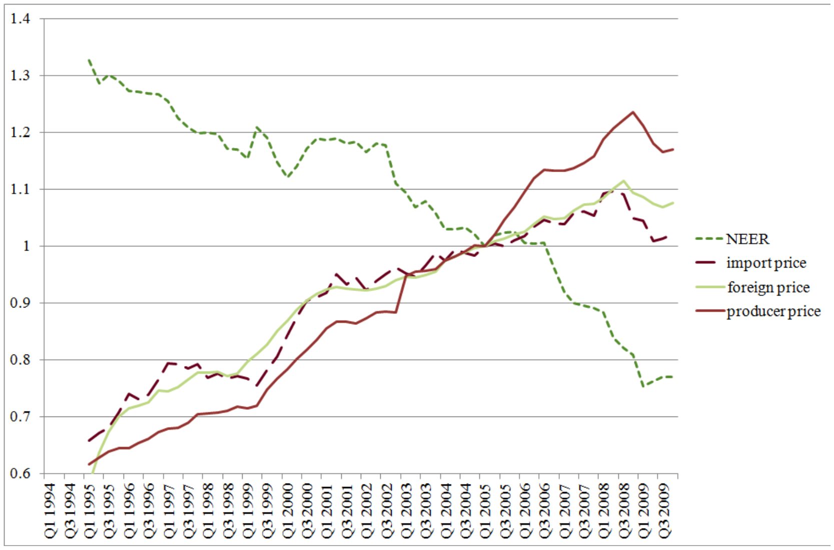

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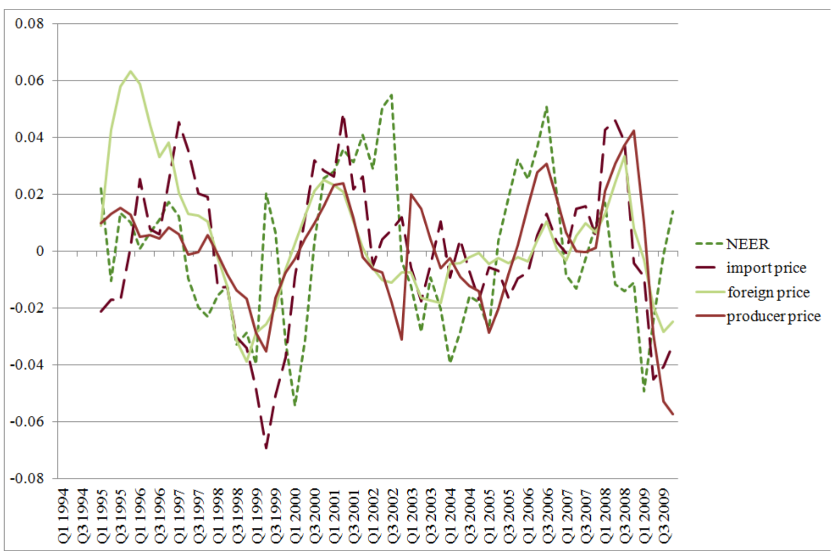

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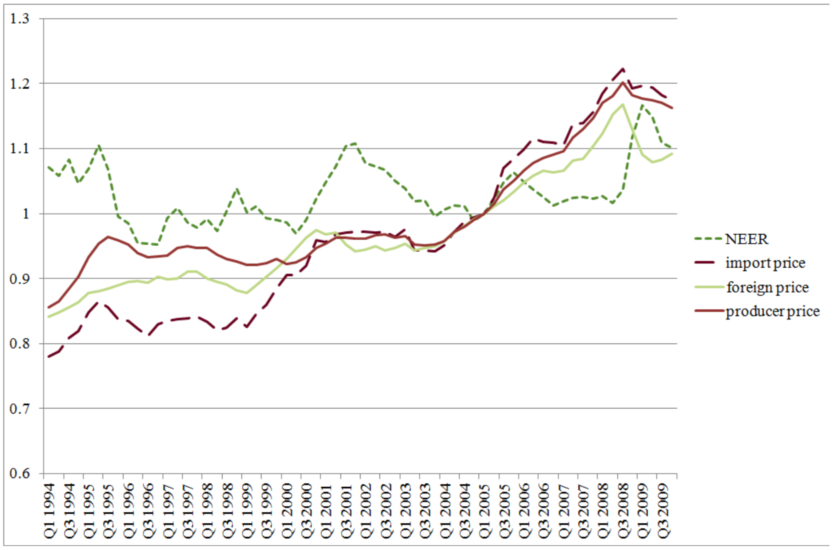


Figure 3.11. NEER and price indexes - Sweden

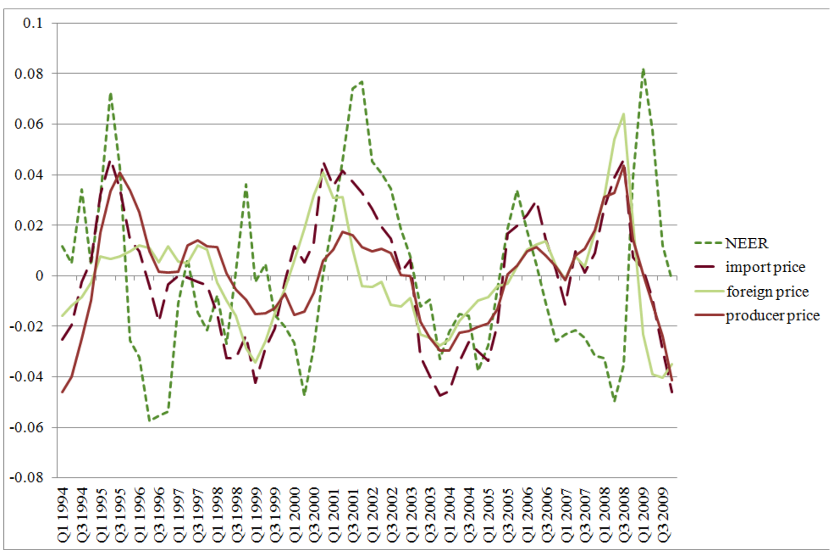
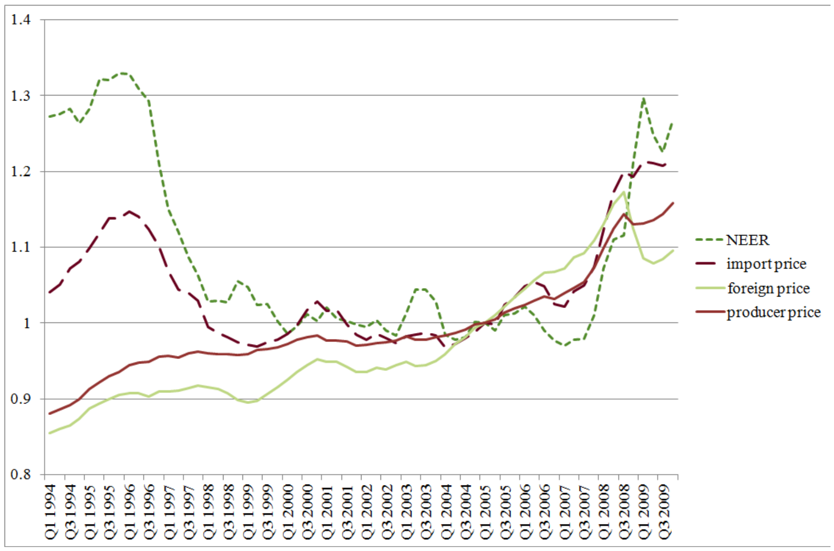


Figure 3.12. Cyclical component - Sweden



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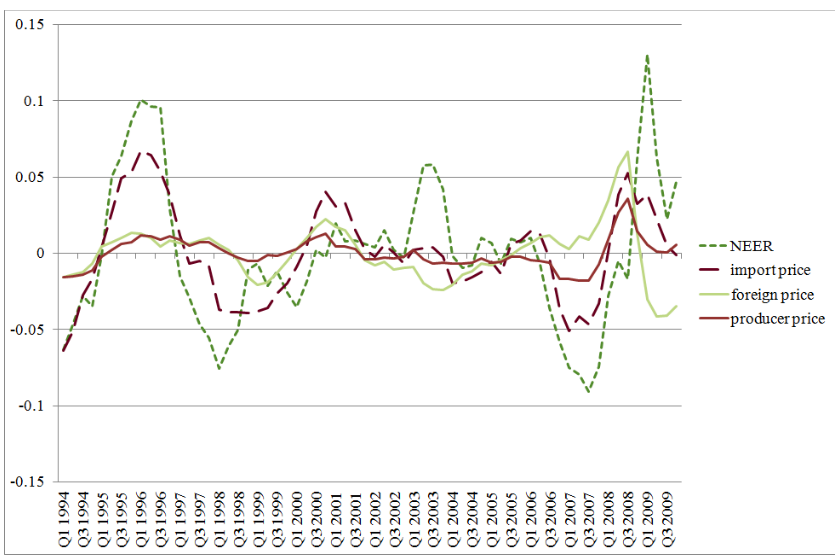

Figure 3.14. Cyclical component - United Kingdom

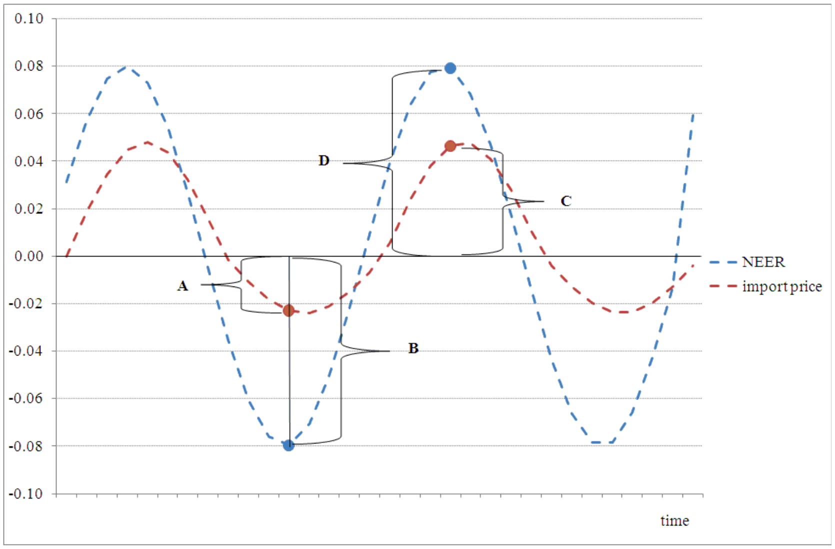

Figure 3.15. Measure of asymmetry in pass-through

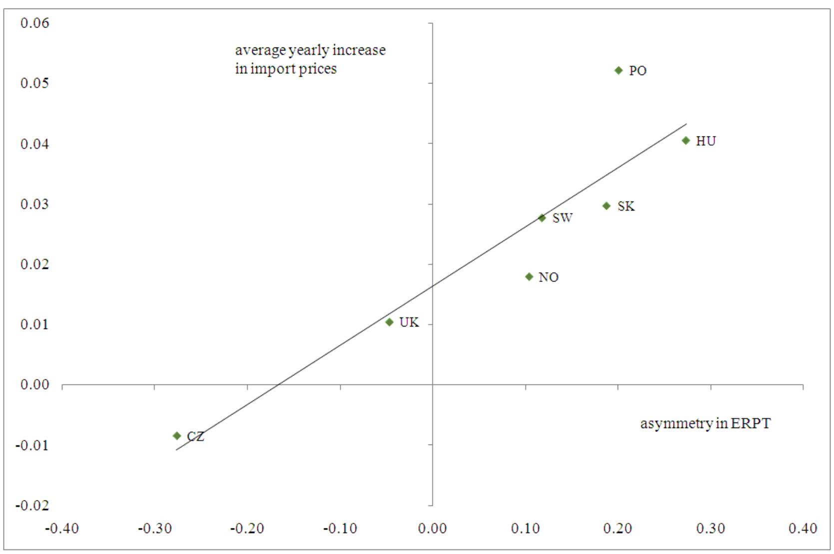

Figure 3.16. Import price inflation and asymmetric pass-through

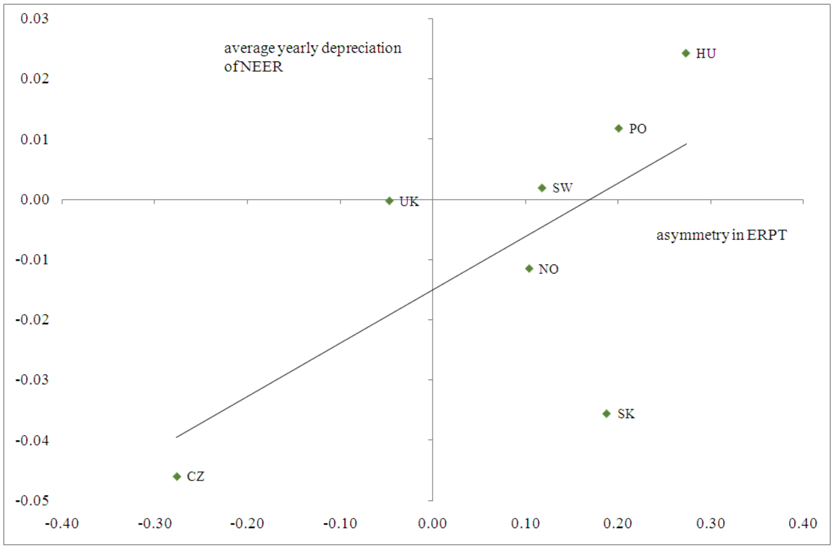

Figure 3.17. NEER and asymmetric pass-through

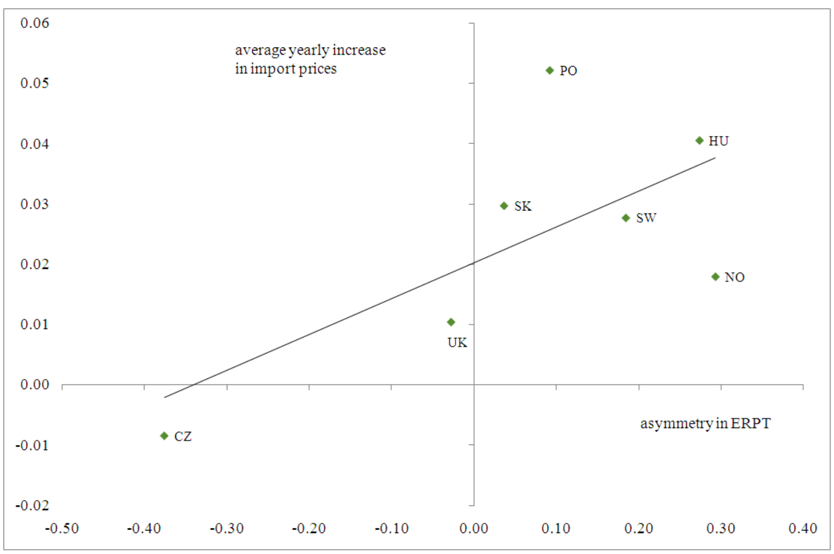

Figure 3.18. Import price inflation and asymmetric pass-through (using the Baxter-King filter)

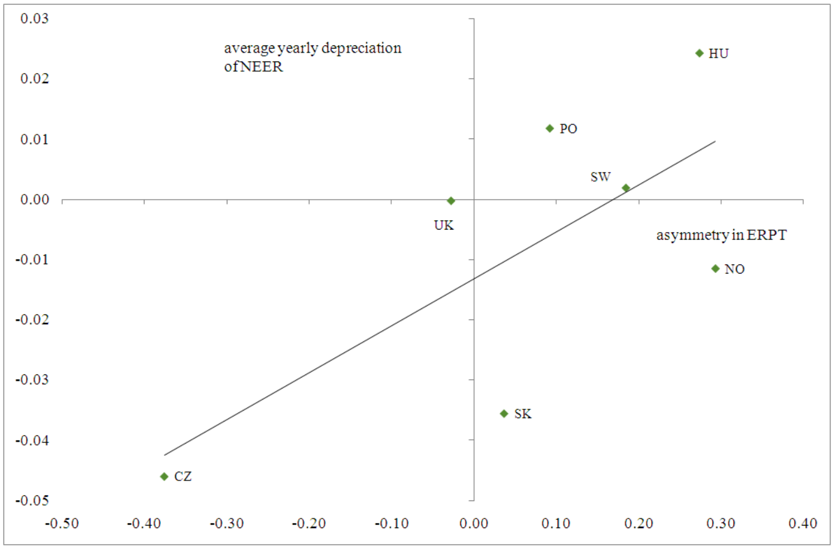

Figure 3.19. NEER and asymmetric pass-through (using the Baxter-King filter)