Mihály Hajnal, György Molnár, Judit Várhegyi

Exchange rate pass-through after the crisis: the Hungarian experience

MNB Occasional Papers 121

2015

Mihály Hajnal, György Molnár, Judit Várhegyi

Exchange rate pass-through after the crisis: the Hungarian experience

MNB Occasional Papers 121

2015

The views expressed are those of the authors and do not necessarily reflect the official view of the central bank of Hungary (Magyar Nemzeti Bank).

MNB Occasional Papers 121

Exchange rate pass-through after the crisis: the Hungarian experience*

(Árfolyam-begyűrűzés a válság után: magyarországi tapasztalatok)

Written by Mihály Hajnal, György Molnár**, Judit Várhegyi

Budapest, September 2015

Published by the Magyar Nemzeti Bank

Publisher in charge: Eszter Hergár

Szabadság tér 9., H-1054 Budapest

www.mnb.hu

ISSN 1585 5678 (on-line)

*We are grateful for comments and suggestions received from several readers of this paper. In particular, we are grateful to Gábor Pellényi, Ádám Reiff and to participants of the ECB NIPE workshop.

**Corresponding author at: Magyar Nemzeti Bank, Email address: molnargy@mnb.hu

Contents

Abstract

Exchange rate movements influence prices through numerous channels. In this paper we provide empirical evidence on pass-through of exchange rate movements into consumer prices. The pass-through depends on a number of factors, and its size may vary over time. In recent years, prices have responded less to a depreciation of the exchange rate than would have been warranted by estimates conducted before the crisis. Before the crisis a 1 per cent change in the exchange rate resulted in a 0.3 per cent change in price level after two years. Currently, the pass-through is estimated to be in the range of 0.1–0.2 per cent over a two-year horizon. Both cyclical (subdued demand) and structural (decline in level of inflation) factors have contributed to the weakening of the relationship.

JEL: C32, E31, F31.

Keywords: exchange rate pass-through, inflation, time series models.

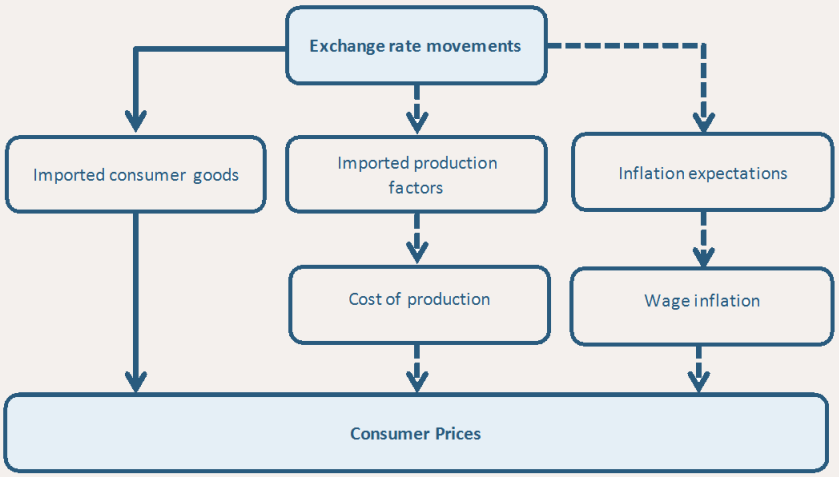
Összefoglaló

Az árfolyam változása számos csatornán befolyásolja a fogyasztói árakat. Tanulmányunkban az árfolyam inflációra gyakorolt hatását elemezzük. Az árfolyam begyűrűzés számos tényező függvénye és mértéke időben változhat. Az elmúlt időszakban azt tapasztalhattuk, hogy az árak kisebb mértékben emelkedtek az árfolyam gyengülését követően, mint ami a válság előtt becsült összefüggésekből következne. A válság előtti időszakban két év alatt nagyjából az árfolyamgyengülés 30 százaléka jelenhetett meg az árakban. Aktuálisan becsléseink szerint az árfolyam megváltozása két éves időhorizonton 0,1-0,2 megfelelő mértékben módosíthatja a fogyasztói árakat. A kapcsolat gyengülése mögött ciklikus (visszafogott kereslet) és strukturális (infláció szintjének csökkenése) tényezők egyaránt állhatnak.

1 Introduction

Hungary is a small open economy where the exchange rate has functioned as the strongest channel in shaping inflation. Exchange rate movements influence consumer prices through numerous channels (Figure 1). A weakening of the exchange rate pushes consumer prices up through the ensuing rise in import prices. Part of the impact materialises directly through the rising prices of imported consumer goods. In addition, an increase in the price of imported production factors indirectly affects the prices of consumer goods produced by domestic companies. Moreover, the exchange rate may also affect the pricing of domestic products competing with imported goods. The depreciation of the exchange rate may have second-round effects as well: higher inflation tends to induce wage increases which may lead to further increases in the consumer prices. Therefore, it is important for inflation targeting central banks to examine the impact of the exchange rate movements on inflation.

Figure 1
Exchange rate channel in a small open economy



When interpreting results about the magnitude of exchange rate pass-through (EPRT) it is important to keep in mind that the pass-through depends on a number of factors, and its size may vary over time and across countries. Its evolution may be influenced by structural factors, such as the level of inflation and the extent to which inflation expectations are anchored, as well as cyclical factors, such as the business cycle.

In recent years, prices have responded less to a depreciation of the exchange rate than would have been warranted by estimates conducted before the crisis as indicated on Figure 3 (right panel). Late recognition of the reduction of ERPT may have contributed to systematic upward bias in the forecast of non-energy industrial goods inflation. Against this background, it is advisable to reinvestigate the size of ERPT, and evaluate how we can make improvements to our current bottom-up forecast framework by taking into account recent developments.

At the same time we would like to disentangle the reasons for the decline of ERPT. One line of explanation suggests that cyclical factors (business cycle, import prices) are responsible for its decline, while others may point to structural reasons (decline in inflation and adjustment of inflation expectations). It is therefore important to identify the underlying reasons and to determine their persistence. If structural factors have contributed to the decline then pass-through is projected to remain diminished even after demand picks up.

The structure of the analysis is as follows. First, we briefly discuss related literature and some empirical results. In the next section we present historical background, some stylized facts and the motivation for our analysis. In section 3 we evaluate the magnitude of exchange rate pass-through to consumer prices in Hungary, and discuss its evolution over time. In the fourth section, we outline the factors that may change the impact of the exchange rate on inflation. Finally, we summarize the results obtained.

2 Related literature

Over the past decades a large economic literature has been devoted to examining exchange rate pass-through. Empirical studies conducted both for developed and for emerging market economies find evidence of incomplete pass-through. There are however considerable differences across countries, which leads to the question of the underlying determinants of ERPT.

The responsiveness of prices to an exchange rate shock may be attributed to structural factors. In particular, Taylor (2000) put forward the hypothesis that the magnitude of pass-through is lower in periods/countries characterized by a persistently low inflation. The rationale is that a change to a lower inflation environment, for example due to a change in the monetary policy regime, can lead to a decrease in the persistence of cost shocks. In this context exchange rate movements are perceived to be transitory, therefore firms might respond to a lesser extent via price-adjustments.

The market power of domestic firms may also have been significantly affected by globalization, because intensified competition with imports affects producers pricing. The competitive pressure stemming from globalization may force businesses to absorb temporary costs due to cost-push shocks thereby also reducing the transmission of exchange rate fluctuations to consumer prices. Furthermore, according to Campa–Goldberg (2005), the degree of exchange rate pass-through to consumer prices also reflects the structure of imports. According to their findings, the inflationary impact of exchange rate movements has declined where the import composition has shifted from primary products towards manufactured goods.

The business cycle may also impact the magnitude of EPRT. In an overheated economy, companies can more easily pass on costs induced by exchange rate movements to consumers than in an economy characterized by restrained demand. Using a cross section of 71 countries, Goldfajn–Werlang (2000) found that the output gap has a considerable impact on the estimated results. Based on Brazilian data, Correa–Minella (2006) also found that the exchange rate pass-through is lower during an economic downturn.

The results suggest that prices respond more sharply to the weakening of the exchange rate than to an appreciation, which may be indicative of the downward rigidity of prices. Another form of asymmetry discussed in the literature argues that ERPT is also affected by the magnitude and volatility of exchange rate changes. According to Correa-Minella (2006) responsiveness is higher when the exchange rate depreciates above a certain threshold – a phenomenon possibly related to the theory of menu costs. Prices may also respond more strongly when the volatility of the exchange rate is lower, as in this case, economic agents may perceive a shift in the exchange rate to be more permanent, prompting companies to adjust prices.

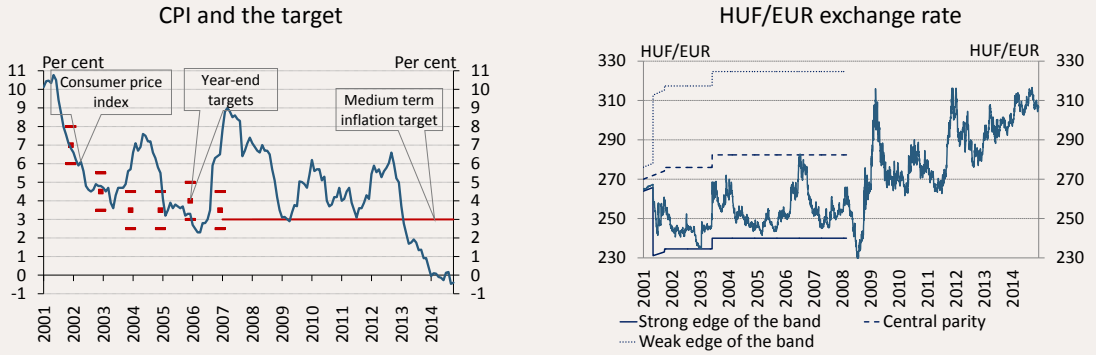
Regional experiences indicate that ERPT has declined in the Czech Republic and Poland. A study conducted by the Czech National Bank (Kucharčuková et al., 2013) indicates that pass-through was lower during the years following the outbreak of the crisis. The paper argues that the change is predominantly related to the subdued demand. Recent data, however, suggests that this was a temporary phenomenon and that ERPT has returned to its pre-crisis level. Arratibel and Michaelis (2013) examined the time evolution of pass-through using Polish data pertaining to the period of 1996–2012. ERPT was found to be twice as large in the 90s as in the past decade. The authors attributed the decline to the adoption of inflation targeting in 1998, the transition to a floating exchange rate regime in 2000, and to the lower level of inflation achieved during the review period.

Regarding the earlier results for Hungary, Darvas (2001) estimated an error correction model with time varying parameters to measure the effect of exchange rate on consumer prices in EU candidate countries. He found that in 2000 the instantaneous pass through was around 10 per cent, while in the long term 40 per cent of an exchange rate change appeared in consumer prices. Vonnák (2010) estimated a structural vector autoregression model to investigate the role of monetary policy between 1995 and 2006. He identified monetary policy and risk premium shocks. The estimates of exchange rate pass-through are slightly different across shocks. A one per cent depreciation of the exchange rate due to consecutive risk premium shocks increases consumer prices 0.2 and 0.3 per cent by the end of first and second year, respectively. The pass through in the case of monetary policy shocks is 0.1 and 0.2 per cent.

3 Historical background and stylized facts in Hungary

Disinflation has been one of the key challenges for policy makers in our region since transition. International experience shows that disinflation from double-digit price increases to moderate levels (around 8-10 per cent) was relatively straightforward, but the transition from moderate inflation to price stability was more time consuming (Figure 2). In Hungary following the transition to market economy inflation jumped to high levels. A crawling peg exchange rate regime was adapted to bring down inflation. As a result inflation decelerated from its peak of 35 per cent in 1991 to 10 per cent in 1999.

Figure 2
CPI inflation and the HUF/EUR exchange rate



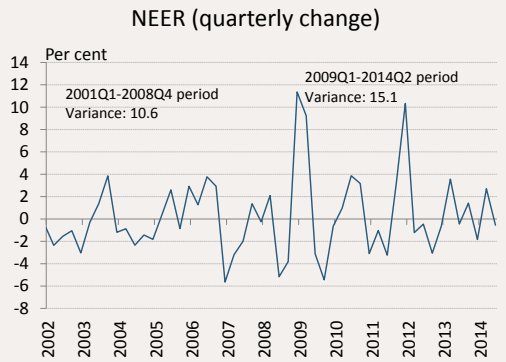
June 2001 marked the introduction of inflation targeting in Hungary. This move was a response to various unsuccessful attempts to further moderate inflation in an environment characterised by large shocks and strong exchange rate pass–through to domestic prices. Nevertheless, the new regime could not be regarded as a fully-fledged IT system from the onset, as the exchange rate band was not abolished in parallel. The transition to inflation targeting initially contributed to further disinflation, but the inflation target could not be met for a prolonged period. Overall, disinflation was fragile and inflation remained relatively high and volatile. The hybrid monetary policy regime was abandoned in 2008, when the MNB adopted a floating exchange rate system.

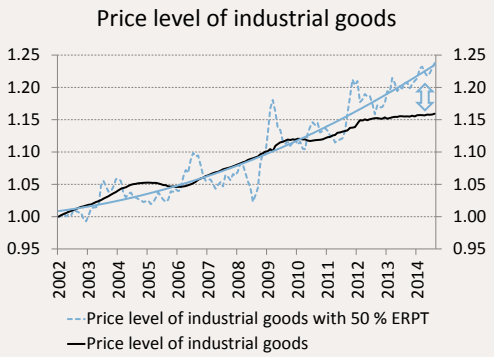
Looking at the Hungarian data, before the 2007-2008 crisis, ERPT was estimated to be around 30 per cent both at the level of the consumer price index and core inflation. In recent years, however, prices have responded less to a depreciation of the exchange rate than would have been warranted by estimates conducted before the crisis. This phenomenon is most evident in the price developments of non-energy industrial goods. Despite the fact that a substantial part of this product group is imported, the depreciation of the exchange rate was reflected in prices only to a limited degree (Figure 3). To illustrate this we compare the price level of industrial goods in the euro area adjusted by the HUF/EUR exchange rate assuming a pass-through of 50 per cent. Before 2010 the two series closely move together, however in the subsequent period the relationship breaks down. The economic downturn, the decline in the level of inflation, and increased exchange rate volatility following the outbreak of the financial crisis may have changed the rate and speed of ERPT.

The volatility of the exchange rate has increased since the crisis (Figure 3, left panel). In this situation, economic agents may perceive the shifts in the exchange rate as temporary, leading to a sluggish adjustment of prices. We analyse this phenomenon

Figure 3
Nominal effective exchange rate and prices of industrial goods

(Price level of industrial goods in the euro area adjusted by HUF/EUR exchange rate (blue line), and price level of industrial goods in Hungary (black line))





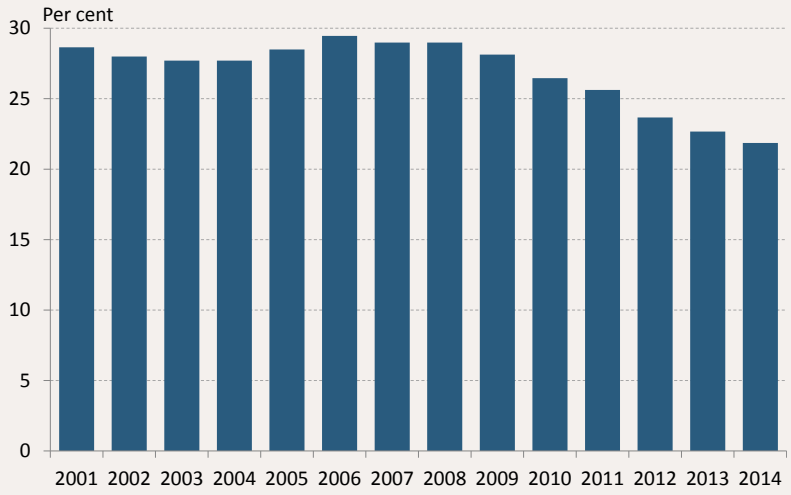
by applying the band-pass filter to isolate short term fluctuations in the exchange rate from the trend. Using monthly data we specify the cyclical component of the time series as cycles with duration shorter than one quarter. Putting the extracted components into a simple OLS regression leads to the result that while the trend component has a significant effect on prices the cyclical component does not. Economic agents only seem to react to persistent changes in the exchange rate. With the increased variance in the exchange rate, agents may currently find it more difficult to disentangle persistent changes from cyclical volatility. Therefore the pass-through of exchange rate shocks may be slower than in the past.

The decline in ERPT may partly result from changes in the composition of household consumption. The weight of industrial goods within the CPI basket has decreased from around 30 per cent in 2001 to just above 20 per cent in 2014 (Figure 4). As industrial goods are highly tradable, the exchange rate plays a particularly important role in their price dynamics. Therefore, the diminishing share of industrial goods may have contributed to the decline in ERPT. We do not however control for this change when evaluating the degree of pass-through. [1]

[1] The only exception is the error correction framework used in the near-term inflation forecasting process. We estimate the exchange rate pass-through at the aggregate level by weighting together the results for the subcomponents. Therefore we can control for the changes in the consumption basket. The results suggest that part of the decline is indeed due to the diminishing share of industrial goods.

Figure 4
CPI weight of industrial goods

(Source: HCSO, MNB)



4 Empirical models of exchange rate pass-through

In this section the exchange rate pass-through to prices is analysed using several empirical approaches. Prices are measured by core inflation adjusted for indirect taxes. Indirect tax changes are excluded because they are not informative about underlying inflation developments. The identification of the pass-through would be more difficult for non-core items. The price regulation mechanism of administered products has changed substantially in recent years, which loosened the link between the exchange rate and retail energy prices. Unprocessed food prices are characterised by high volatility that typically arises from exogenous events such as weather conditions. Fuel prices are an exception: there is essentially full and immediate pass-through in this subgroup. Our sample starts in June 2001 and ends in June 2014. The period before the introduction of inflation targeting is disregarded as the gradual, pre-announced depreciation of the currency during the crawling peg exchange rate regime may imply substantially different macroeconomic relationships. Seasonal adjustment is applied if appropriate. The first approach is a rolling regression estimated by OLS. The second approach proposes an error correction framework. Finally a vector autoregressive model is applied to the data. Depending on the particular analysis, estimation was carried out for monthly or quarterly data. As a caveat, the relatively short sample period and high volatility in the data generate wide confidence bands and small sample bias may distort the results.

4.1 Rolling window analysis

A common underlying assumption for time series models is that the parameters are constant over time. We start our analysis of the exchange rate pass-through by estimating a single equation model of inflation on a rolling window using ordinary least squares. The methodology is a simple and convenient approach to assess the stability of parameters.

Our calculations are based on quarterly data, the sample is from 2001Q3 to 2014Q2. We use a window size of seven years. [2] The first window covers the time period from 2001Q3 to 2008Q2. Subsequently, the window is moved one step ahead, such that we can estimate the model for 25 windows, with the last one covering the sample period from 2007Q3 to 2014Q2.

We run the rolling regression using the following specification:

$$\Delta logpc_t = \alpha + \beta_1 \Delta logpc_{t-1} + \beta_2 \Delta loge_{t-1} + \beta_3 GAP_{t-1} + \beta_4 logcore\_ea_t + \beta_5 \Delta logcpb_t + u_t \quad (1)$$

where, pc_t is core inflation excluding indirect taxes, e_t is the forint/euro exchange rate, *GAP* is the output gap, *core_ea* is core inflation in the euro area, and *CPB* is the world price of processed industry goods. [3] The output gap is included to control for the inflation pressure coming from the domestic economy, while import prices in foreign currency are captured by core inflation in the eurozone (in euros) and the world price of processed goods (in U.S. dollars).

Based on our results (7 year windows)[4], a 1 per cent depreciation of forint/euro exchange rate increased core inflation excluding indirect taxes by approximately 0.05-0.06 per cent in short term (from one quarter to another) in the period leading up to the

[2] Choosing the size of the rolling window is related to the response time between the variables. Longer windows can be appropriate for systems with slower response time. Longer rolling window sizes tend to yield smoother rolling window estimates than shorter ones. Shorter rolling window size produces a better picture of short term dynamics, however the estimates of the parameters are more uncertain, which is reflected in the width of the confidence bands.

[3] Source: CPB Netherlands Bureau for Economic Policy Analysis

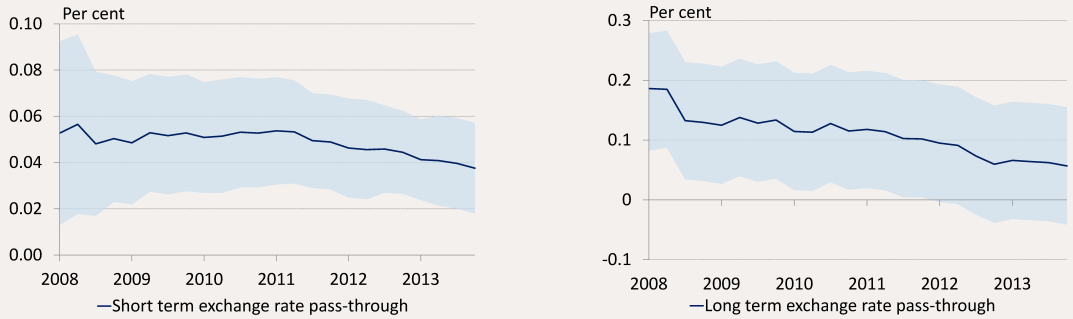
[4] Detailed estimatation results are available in the appendix.

crisis, whereas this value decreased to 0.04 per cent following 2008 (Figure 5, left panel). The long term effect of the forint/euro exchange rate on core inflation, which can be calculated using the persistence parameter of core inflation is as follows:

$$Long\ term\ pass-through = \frac{\beta_{2,t}}{1-\beta_{1,t}} \tag{2}$$

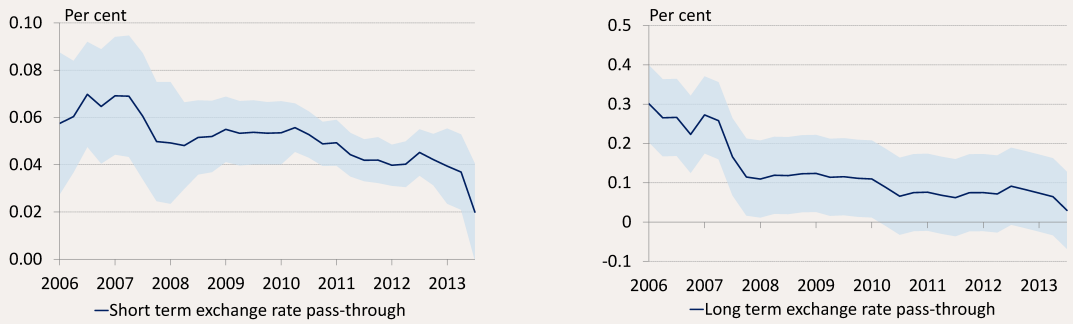
The calculated long term impact has also weakened. While a 1 per cent depreciation of forint/euro exchange rate increased core inflation excluding indirect taxes by 0.2 per cent before 2009, this value decreased to 0.1 per cent in the period following the financial crisis (Figure 5, right panel).

Figure 5
Rolling-window estimation with 7-year window
(Blue bands represent 95 per cent confidence interval.)



As a robustness check we carried out the same exercise with a window size of 5-year, and the results were qualitatively similar (Figure 6). Although these estimates are not as smooth, they nevertheless provide useful information. The shorter interval gives a better picture of the short term dynamics, and therefore the latest developments in the parameter values. At the end of the sample there is a sharp decline in estimates of the coefficients, however it also illustrates the fact that there were periods when the coefficients for exchange rate pass-through were increasing over time.

Figure 6
Rolling-window estimation with 5-year window
(Blue bands represent 95 per cent confidence interval.)



4.2 An error correction framework for forecasting inflation subcomponents

To investigate the exchange rate pass-through to consumer prices, we reestimated the disaggregate models applied in the near-term inflation forecasting process. The empirical methodology is based on the error correction model of Várpalotai (2003). The models follow an error correction framework, where prices are determined by cost factors in the long run, but they can temporarily diverge from the long term equilibrium, for example due to changes in demand. The error correction framework may give a more accurate picture of pass-through at longer horizons because it explicitly takes into account long-run relationships between the levels of endogenous variables. The model is employed by MNB staff to forecast the subcomponents of core inflation excluding indirect taxes: industrial goods, market services, and processed food prices.[5] We estimate the exchange rate pass-through at the aggregate level by weighting together the results for the sub-components. In addition, the model also contains a separate equation for unprocessed food prices. The main advantage of the disaggregated approach is that it can accommodate the heterogeneity among inflation subcomponents.

We used logarithmized, seasonally adjusted data at a monthly frequency. The estimation period runs from June 2001 to June 2014. The use of monthly data enables us to investigate the price dynamics at the high frequency required in near term forecasting and ensures sufficient number of observations. The main explanatory variables are producer prices for agricultural products, unit labour costs[6] and energy prices in the services and manufacturing sector. As a proxy for imported inflation, we consider euro area inflation for processed and unprocessed food products, durable and nondurable industrial goods. The imported inflation proxy is expressed in local currency. Table 1 summarizes the explanatory variables used for the different inflation subcomponents.

Table 1
Variables in long term equations

	industrial goods		market services	food	
	durable	nondurable		processed	nonprocessed
euro area durable prices in HUF	x				
euro area nondurable prices in HUF		x			
euro are processed food prices in HUF				x	
euro are nonprocessed food prices in HUF					x
ULC in manufacturing	x	x		x	x
ULC in services	x	x	x	x	x
energy for companies in manufacturing sector				x	x
energy for companies in services sector		x	x	x	x
agricultural producer prices				x	x
CPI	x	x	x	x	x

The long run cointegration relationship for the subcomponents is the following:

$$p_{i,t} = a_i + \alpha_i t + \sum_n \beta_n c_{n,t} + u_{i,t} \tag{3}$$

where $p_{i,t}$ is the price level for subcomponents, $c_{n,t}$ the cost factors, t is a linear trend, and $u_{i,t}$ is the error term. We assume that $\sum_n \beta = 1$, so that price elasticity is 1 in the long run: if all cost factors increase by 1 per cent, the prices also grow by 1 per

[5] There is no separate equation for alcohol and tobacco which is part of core inflation, because price development in the case of these products are determined predominantly by changes in excise duties, while the exchange rate does not significantly affect price developments.

[6] In case of unit labour costs the data are not available at monthly frequency. Therefore, the quarterly data are interpolated.

cent. The model may not capture all relevant cost factors; we assume that the prices of omitted cost factors grow at the same rate as the consumer price index.

We apply a Bayesian approach to estimate the model. Expert knowledge and previous parameter estimates are summarized in the prior. This leads to the following optimization problem:

$$min_{\beta}((1-\omega)(y-X\beta)'(y-X\beta)+\omega\sum(\beta-\beta^{pr})^2) \tag{4}$$

where β^{pr} is the prior, while ω is the weight on the prior in the estimation. It is assumed that $\beta_j > 0, \forall\, j$.

In the short run, fluctuations in demand can push prices away from the long term equilibrium. The variables in short term equations are the same as in the long term equations expect that they are differenced. They also contain the lagged dependent variable ($\Delta p_{i,t-j}$) and the error term from the long run relationship ($u_{i,t-1}$) to capture the short term disequilibrium in prices. Demand is captured by the moving average of changes in consumption expenditure[7] (Δce_t). The short term equations have the following form:

$$\Delta p_{i,t} = b_i + \delta_i \Delta ce_t + \sum_n \gamma_n B_n(L)\Delta c_{n,t} + \sum_j \Delta p_{i,t-j} + \phi u_{i,t-1} + \varepsilon_{i,t} \tag{5}$$

where ϕ_i is the error correction parameter. It captures how prices react to the short run deviations from the equilibrium. The error correction parameters represent the speed of adjustment in absolute value. If the error correction parameter is negative, the prices revert to the long term equilibrium. Based on the whole sample, the error correction parameters are found to be negative (Table 2).

Table 2
Error correction parameters in the short term equation

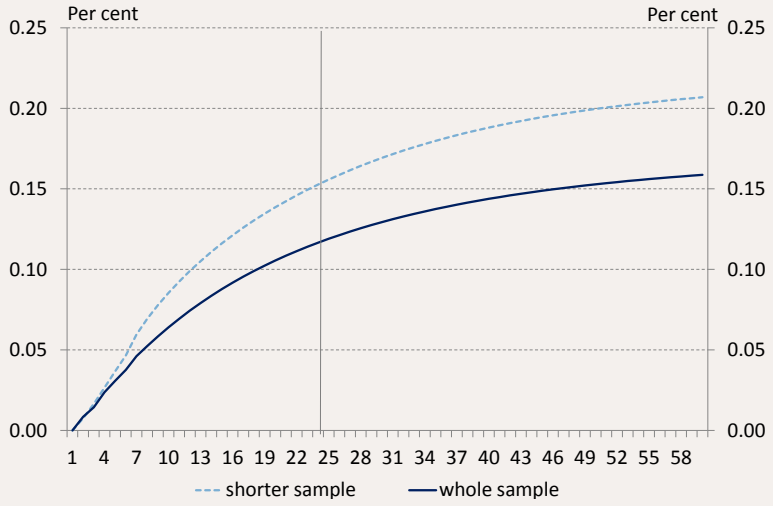
Subcomponent	Parameter value
Nonprocessed food	-0.18
Processed food	-0.08
Market services	-0.01
Durable industrial goods	-0.02
Nondurable industrial goods	-0.04

The previous version of the model was based on estimation carried out in 2010. Our aim in this exercise was to reestimate the parameters and to update the framework to include new data up to mid-2014, while preserving the basic structure of the model. However, the number of lags and explanatory variables was slightly reduced to obtain a more parsimonious model.

Exchange rate pass-through to prices is analysed by means of impulse response functions. Figure 7 shows the response of prices to import prices expressed in HUF. The shock is standardized to 1 per cent permanent depreciation. We estimated the model on 2 different samples: a shorter sub period leading up to the financial crisis (June 2001 - December 2008), and the entire sample (June 2001 – June 2014). Looking at the full sample, a 1 per cent depreciation of the exchange rate increases the consumer prices by roughly 0.1 per cent after two years. Comparing this value to the result based on the period leading up to the crisis confirms that the exchange rate pass-through has decreased over time mostly due to the decline of weight of industrial goods.

[7] In case of consumption the data are not available at monthly frequency. Therefore, the quarterly data are interpolated.

Figure 7
The effect of 1 per cent increase in import prices on core inflation



4.3 Vector autoregression on quarterly data

This section applies standard reduced form methods (vector autoregressions, VAR) to investigate the pass-through of exchange rate shocks. The basic framework follows McCarthy (2000), capturing the macroeconomic dynamics in a model of the distribution chain of pricing. The underlying structural shocks are recovered by a recursive identification scheme, imposing short-term restrictions on the covariance matrix. The size and speed of exchange rate pass-through is derived from impulse response functions. This method has been applied to study the transmission of external shocks in the euro area and emerging market countries (Hahn, 2003, Ca'Zorzi et al., 2007).

The underlying assumptions when applying standard VAR is that (i) international spillovers can be captured by exogenous variables, (ii) nonlinearities are negligible, and (iii) parameters are constant through time. The setup provides a useful benchmark. The assumptions could be relaxed by applying more elaborate methods: for example time-varying VAR, time-varying panel VAR analysis or other non-linear models.

The benchmark VAR has the representation:

$$Y_t = A(L)Y_{t-1} + B(L)X_t + \mu_t \tag{6}$$

where Y_t is a vector of endogenous variables and X_t is a vector of exogenous variables. Real GDP is used as the measure of output (y_t). Prices are measured by core inflation adjusted for indirect taxes (p_t). Within industrial producer prices we consider the prices of the sectors producing consumer goods (ppi_t). $Neer_t$ denotes the nominal effective exchange rate against 14 trade partners. In the studies of Hahn (2003) and Ca'Zorzi et al. (2007) the model also includes non-oil import prices and a nominal interest rate. We leave out these variables, and estimate a more parsimonious model due to the short sample size.

$$Y_t' = [y_t, neer_t, ppi_t, p_t] \tag{7}$$

The estimated VAR contains two exogenous variables: euro area real GDP growth (ea_t) and a world commodity price index (cp_t).

$$X_t' = [ea_t, cp_t] \tag{8}$$

These variables are included to control for changes in foreign demand and inflation. This is implicitly assuming that there is no feedback from the country specific endogenous variables to the exogenous variables.

The analysis considers quarterly data on the period between 2001Q3-2014Q2. Seasonal adjustment is applied if appropriate, after which variables are expressed as quarterly growth rates. Standard tests are used to determine the lag-order of the VARs. Considering the information criteria the VAR are estimated with one lag of the endogenous variables.

In terms of identification assumptions, we limit ourselves to the standard Cholesky decomposition of the covariance matrix (short-term restrictions). This is equivalent to assuming a recursive exogeneity structure, where we place GDP first in the ordering of the variables, followed by the exchange rate, and then prices. As production only tends to react to persistent changes in the exchange rate, and can be subject to adjustment costs, we assume that shocks to the exchange rate have no contemporaneous impact on output growth, but may affect prices immediately. The ordering among the price variables follows the distribution chain of pricing. Robustness was tested using other orderings of the variables, and the results were robust as regard other plausible orderings in the Cholesky decomposition.

The transmission of an exchange rate shock to prices is analyzed by means of impulse response functions in Figure 8. The accumulated impulse responses are displayed over a horizon of twelve quarters. The exchange rate shock is standardized to a one-percent to illustrate the transmission of depreciation. Inflation pass-through is estimated to be around 0.15 per cent assuming the exchange rate permanently weakens by 1 per cent. The change in exchange rate finds its way to prices and therefore inflation in approximately 2 years. This analysis considers a relatively short period, and therefore produces wide confidence bands for the impulse responses.

Figure 8
Response of prices to 1 per cent permanent exchange rate shock
(Point estimate with one and two-standard error band)

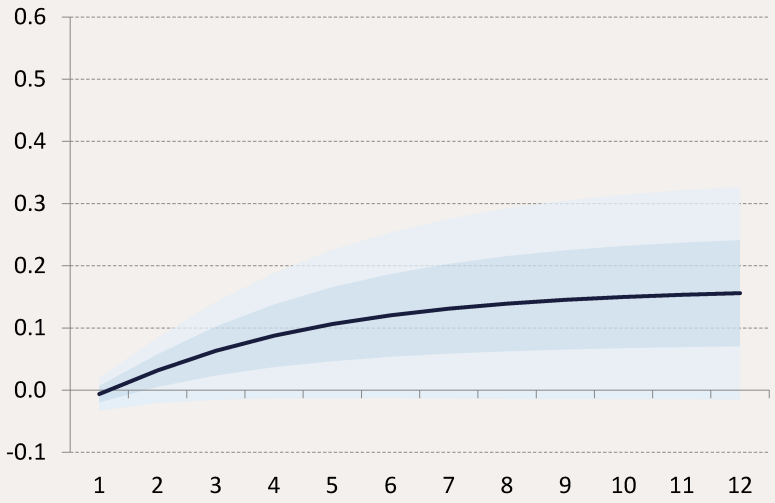
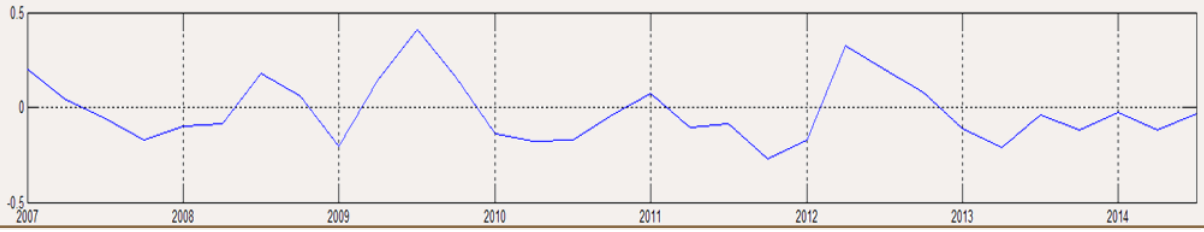


Figure 9 shows the historical contribution of the exchange rate shocks to core inflation adjusted for indirect taxes. The depreciation following the outbreak of the 2008-2009 financial crisis and the beginning of 2012 significantly contributed to the rise in inflation. However, the episodes of depreciation since the beginning of 2013 have not had any inflationary impact.

Using the recursive Chow breakpoint test the null hypothesis that the benchmark VAR is stable over the estimation period can be rejected. A sensitivity analysis is carried out by estimating the model over a recursive sample[8], and investigating the stability

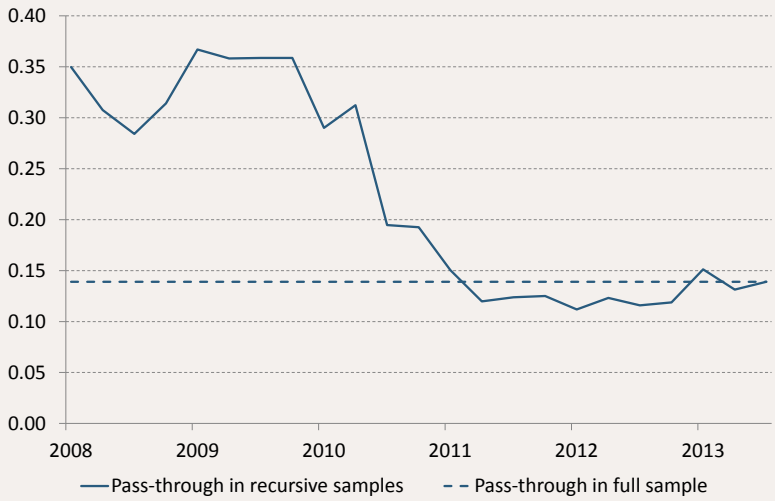
Figure 9
Cumulative effect of exchange rate shock on inflation



of the impulse responses. This allows us to infer whether there is any evidence that the transmission of exchange rate shocks has changed over time. This is a computationally simpler alternative to estimating a time varying parameter VAR. The necessity of having to estimate a huge number of parameters is avoided, however this method does not provide detailed information on the exact timing of changes. The shortest subsample looks at the period leading up to the financial crisis (2001Q3-2008Q4). We report the cumulative impulse response to a 1 per cent depreciation after 2 years (Figure 10).

Inflation pass-through to the level of the core adjusted for indirect taxes is estimated to be around 0.3-0.35 in samples ending in the previous decade. This result is in line with estimates conducted before the crisis. The impact on prices is about double in size than that for the full sample. However the results for the subperiods ending after 2011 are close to those found in the baseline model. In conclusion, the sensitivity analysis of the impulse response functions suggests that only in recent years has the exchange rate had less pronounced effects on prices.

Figure 10
Response of prices after 2 years in recursive samples



The timing of the decline in ERPT may be linked to the depreciation of the HUF against the CHF in 2010-2011. The large share of foreign currency liabilities in households' balance sheets is a crucial feature of the Hungarian economy. The depreciation increased foreign currency debt and the debt service of households in domestic currency terms. As a result, consumption declined significantly, which may have contributed both to the change in ERPT and the increased cyclical sensitivity of Hungarian inflation.

[8] An alternative approach would be to use rolling samples. Due to the short sample period, and the high number of parameteres this estimation can only be done with large uncertainty.

5 Disentangling the factors behind the decline in pass-through

5.1 Threshold VAR model

The crisis may have triggered changes in economic agent's behaviour which may have altered economic relationships. The investigation in this section is aimed at answering two central questions, notably: is the transmission of exchange rate shocks different when aggregate demand is subdued compared to normal times? Has the decline in inflation and the adjustment of inflation expectations contributed to lower ERPT? To investigate these hypotheses we estimate a threshold-VAR model which is particularly appropriate to model abrupt, discrete changes in economic dynamics. The basic setup of the model in Franses and van Dijk (2000) is as follows:

$$Y_t = \begin{cases} \alpha_1 + A_1(L)Y_{t-1} + \mu_t, \ if \ q_{t-d} \leq \gamma_1 \\ \alpha_2 + A_2(L)Y_{t-1} + \mu_t, \ if \ \gamma_1 \leq q_{t-d} \leq \gamma_2 \\ \alpha_3 + A_3(L)Y_{t-1} + \mu_t, \ if \ q_{t-d} > \gamma_2 \end{cases} \quad (9)$$

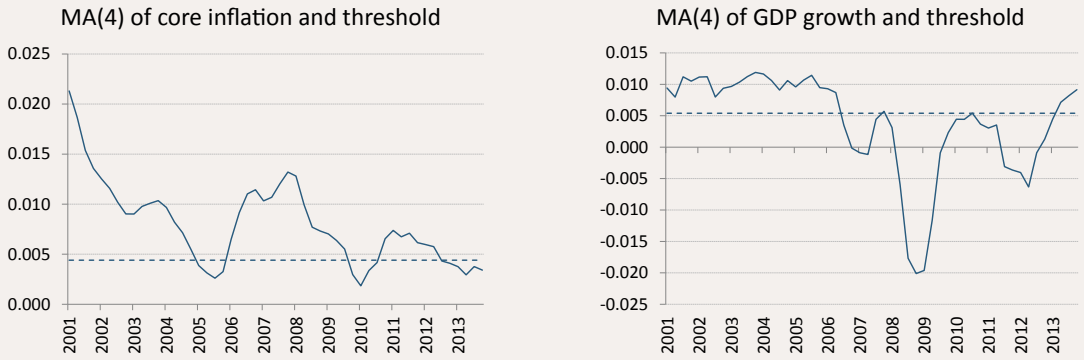
It is assumed that several (minimum 2) regimes exist, and that a threshold variable (q_t) determines the regime shifts of the dynamic process Y_t. The framework has several advantages. It provides a formal framework to test for nonlinearities. Furthermore the framework also provides a comparison of how exchange rate shocks affect prices in the different regimes.

We use seasonally adjusted data at a quarterly frequency. We estimate a two-regime Threshold VAR with one lag where Y_t is a multi-dimensional vector of endogenous variables (GDP growth, nominal effective exchange rate, and core inflation excluding indirect taxes). First, we examine statistical evidence for nonlinearities in ERPT corresponding to different levels of inflation in the economy. We estimate various specifications with 2 regimes, allowing for different coefficients during episodes of high inflation and low inflation. Because the quarter on quarter growth rates are very volatile, we use as threshold variables 4 quarter moving averages of the corresponding variables. Possible threshold values are restricted so that at least 15 per cent of the observations are in each regime. The multivariate extension of the linearity test of Hansen (1999) proposed in Lo and Zivot (2001) takes the ratio of the likelihoods to compare the covariance matrices of the two models. For our preferred specification the null hypothesis of linearity is rejected with a p-value of 0.000.

With the estimated threshold of 0.004 (annualized inflation of 1.8 per cent) there are three low inflation episodes. The first period follows the EU accession, while the other 2 follow the financial crisis of 2007-2008 and the recent slowdown and decline of inflation in the euro area countries. Figure 13 in the appendix displays the impulse responses. Depreciation has a substantially larger effect on inflation when the system is in the high inflation regime. This is particularly true for the larger shock (2 standard deviations). When the exchange rate shock is standardized to a one-percent permanent depreciation, inflation pass-through is estimated to be around 0.6 per cent in the high inflation regime, while in the low inflation regime ERPT declines to 0.1 per cent (Figure 12). This result is in line with the hypothesis put forward by Taylor (2000), in particular that ERPT depends positively on inflation. The results also suggests the existence of asymmetric responses: depreciation has a larger effect on prices than an appreciation of the exchange rate.

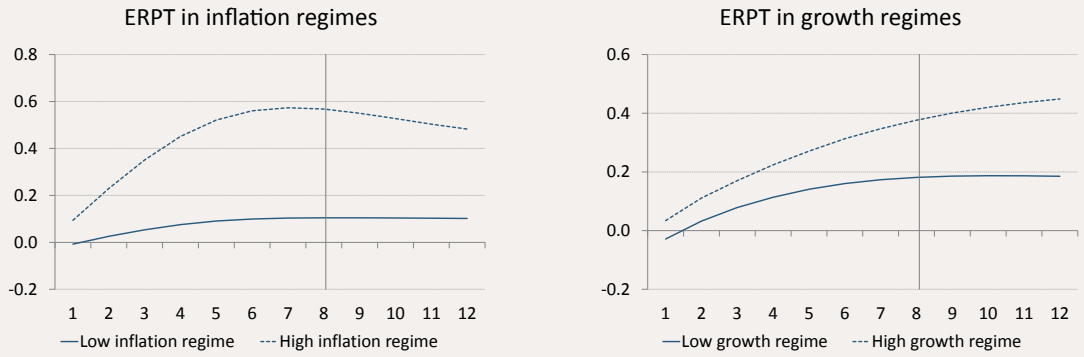
Next we choose GDP growth as the threshold variable, to examine whether cyclical factors may have also contributed to the decline. The results suggest that ERPT was significantly stronger in the period leading up to the middle of 2006 when GDP

Figure 11
Alternative threshold variables and estimated threshold values



growth was above an annualized rate of 2 per cent. The decline in pass-through is significant, but less sizable in this case. It should also be noted that according to the estimated model and threshold value, output growth has recently picked up to a point where pass-through is projected to increase. However, the findings come with the caveat that each estimation conditions only on one potential explanation for the changes in ERPT. In fact the interaction among all these factors determines the degree of pass-through.

Figure 12
ERPT, conditional on regime and threshold variable



6 Conclusion

In this paper we provide empirical evidence on pass-through of exchange rate movements. The analysis suggests that ERPT has decreased since the crisis. This decline can be attributed to several factors. Subdued demand (a cyclical change) has contributed to the weakening of the relationship, as companies have been less able to price in exchange rate movements. In addition, the decline in inflation may also have played a role (a structural change). Before the crisis, inflation had remained persistently above MNB's target of 3 per cent. In recent years however, inflation has moderated and expectations adapted to lower inflation. This anchoring of expectations may have affected the behaviour of economic agents, reducing the impact of cost-push shocks.

Before the crisis a 1 per cent change in the exchange rate resulted in a 0.3 per cent change in price level after two years. Currently, the pass-through is estimated to be in the range of 10–20 per cent over a two-year horizon. Since the decline can be partly attributed to cyclical reasons – which are considered to be temporary –, in the medium term the pass-through could revert to the upper end of the range. When interpreting the results, it is important to keep in mind that the degree of the pass-through may continue to change over time.

The results of our estimates are consistent with the values estimated for other economies in the region. The Czech National Bank (Kucharčuková et al., 2013) and the National Bank of Poland (Arratibel and Michaelis, 2013) both identified a value around 0.1 per cent in time-varying parameter models. In comparison, the values estimated for the euro area range between 0.02 per cent (Faruqee, 2006) and 0.08 per cent (Hahn, 2003).

References

[1] Aleem, A. - Lahiani, A. (2014): "Monetary policy credibility and exchange rate pass-thorugh: Same evidence from emerging countries", *Economic Modelling*, Vol. 43, December 2014, pp. 21-29.

[2] Arratibel, O. – Michaelis, H. (2013): "The impact of monetary policy and exchange rate shocks in Poland: Evidance from a Time-Vaying VAR ", *Munich Discussion Paper*, No. 2013-36, December 2013, pp. 1-36.

[3] Ca'Zorzi, M. – Hanh, E. – Sanchez, M. (2007): "Exchange rate pass-throuugh in emering markets ", *ECB Working paper*, No. 739, March 2007, pp. 1-33.

[4] Campa, J. M. – Goldberg, L. S. (2005): "Exchange rate pass-through into import prices ", *The review of economics and statistics*, 87(4), November 2005, pp. 679-690.

[5] Correa, A. S. – Minella, A. (2006): "Nonlinear mechanisms of the exchange rate pass-through: A Phillips curve model with threshold for Brazil ", *Banco Central Do Brasil Working Paper*, No. 122, November 2006, pp. 1-30.

[6] Darvas, Zs. (2001): "Exchange rate pass-through and real exchange rate in EU candidate countries ", *Discussion Paper of Economic Research Centre of the Deutsche Bundesbank*, 10/01, May 2001, pp. 1-76.

[7] Faruqee, H. (2006): "Exchange rate pass-through in the Euro Area ", *IMF Staff Papers*, Vol. 53, No. 1, pp. 63-88.

[8] Franses, P. H. and van Dijk, D. (2000): "Nonlinear Time Series Models in Empirical Finance ", *Cambridge University Press*

[9] Giannone, D. – Lenza, M. – Primiceri, G. (2012): "Prior selection for vector autoregressions ", *ECB Working paper*, No. 1494, November 2012, pp. 1-41.

[10] Goldfajn, I. – Werlang, S. R. C. (2000): "The pass-through from depreciation to inflation: A panel study ", *Banco Central Do Brasil*, No. 5., September 2000, pp. 1-47.

[11] Hanh, E. (2003): "Pass-through of external shocks to euro area inflation ", *ECB Working paper*, No. 243, July 2003, pp. 1-61.

[12] Hansen, B. E. (1999): "Testing for linearity ", *Journal of Economic Surveys*, Vol. 13, No. 5, December 1999, pp. 551-576.

[13] Kucharcukova, O. B. – Franta, M. – Hajkova, D. – Kral, P. – Kubicova, I. – Podpiera, A. – Saxa, B. (2013): "What we know about monetary policy transmission in the Czech Republic: Collection of empirical results ", *Research And Policy Notes*, No.1, October 2013, pp. 1-58.

[14] Lo, M. and Zivot, E. (2001): "Threshold cointegration and nonlinear adjustment to the law of one price ", *Macroeconomic Dynamics, Cambridge University Press*, Vol. 5(4), September 2001, pp. 533-576.

[15] Taylor, J. B. (2000): "Low inflation, pass-through, and the pricing power of firms ", *European Economic Review*, Vol. 44, Issue 7, June 2000, pp. 1389-1408.

[16] Várpalotai, V. (2003): "Dezaggregált költségbegyűrűzés-alapú ökonometriai infláció-előrejelző modell ", *MNB Munkafüzet*, 2003/4, April 2003, pp. 1-53., *only in Hungarian*

[17] Vonnák, B. (2010): "Risk premium shocks, monetary policy and exchange rate pass-through in the Czech Republic, Hungary and Poland ", *MNB Working Papers*, 2010/1, January 2010, pp. 1-33.

7 Appendix

Figure 13
Response of prices and NEER to shocks, conditional on inflation regime

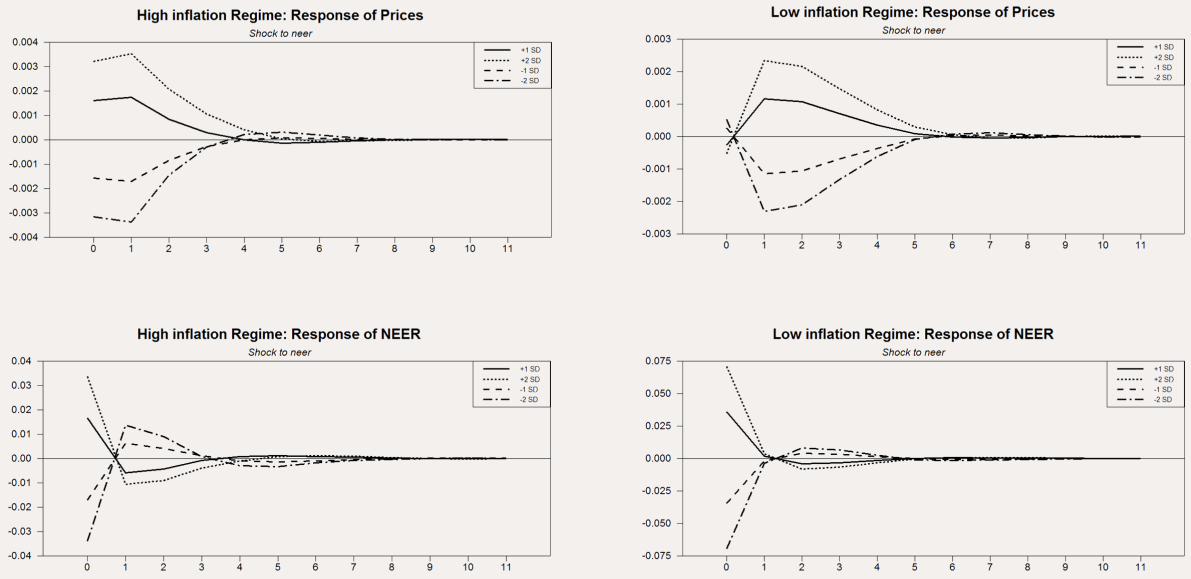


Figure 14
Response of prices and NEER to shocks, conditional on GDP growth regime

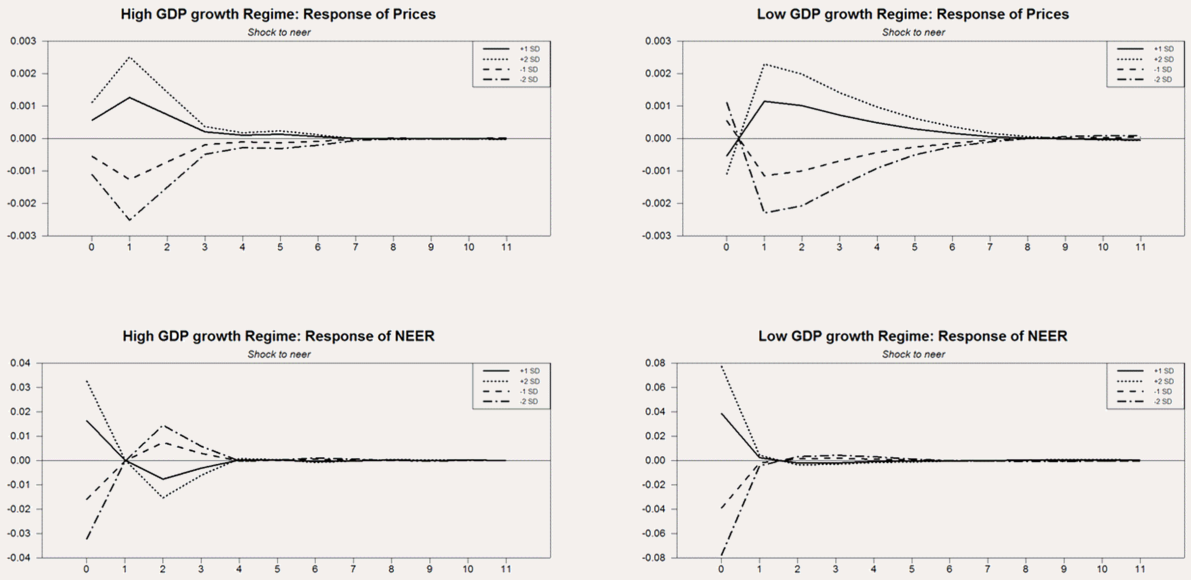


Table 3
Coefficients estimates of the rolling window with 7-year windows

Variable	OLS estimates
constant	0.0035 [0.0012]
$dlogPC_t$	0.3393 [0.1576]
$dlogE_t$	0.0376 [0.0098]
GAP	0.0007 [0.0002]
$dlogCPB_t$	0.0874 [0.0216]
$dlogCORE\_EA_t$	0.0884 [0.0468]

Standard errors are in bracket, below coefficients

Table 4
Coefficients estimates of the rolling window with 5-year windows

Variable	OLS estimates
constant	0.0026 [0.0019]
$dlogPC_t$	0.4658 [0.2090]
$dlogE_t$	0.0394 [0.0159]
GAP	0.0006 [0.0006]
$dlogCPB_t$	0.0723 [0.0318]
$dlogCORE\_EA_t$	0.1130 [0.0594]
Standard errors are in bracket, below coefficients	

Table 5
Coefficients in vector error correction model: the long term equation

	industrial goods		market services	food	
	durable	nondurable		processed	nonprocessed
constant	-3.56	-2.97	-2.72	-2.21	-0.91
CPI	0.25	0.31	0.33	0.13	0.06
EA durable prices in HUF	0.48				
EA nondurable prices in HUF		0.35			
EA processed food prices in HUF				0.10	
EA unprocessed food prices in HUF					0.02
ULC in manufacturing	0.11	0.05		0.20	0.02
ULC in services	0.16	0.22	0.57	0.13	0.07
energy for companies in manufacturing sector				0.05	0.06
energy for companies in services sector		0.06	0.10	0.08	0.25
agricultural producer prices				0.32	0.53
trend	0.00	0.00	0.00	0.00	0.00

Table 6
Coefficients in vector error correction model: the short term equation

	industrial goods		market services	food	
	durable	nondurable		processed	nonprocessed
constant	0.00	0.00	0.00	0.00	0.00
CPI	-0.04	-0.02	-0.01	-0.08	-0.18
EA durable prices in HUF(−3)	0.02				
EA durable prices in HUF(−6)	0.02				
EA nondurable prices in HUF(−1)		0.01			
energy for companies in services sector(−9)			0.02		
ULC in manufacturing(−2)	0.03				
consumption(−2)				-0.01	
consumption(−12)		0.06	0.14		
agricultural producer prices					0.27
agricultural producer prices(−1)					0.12
agricultural producer prices(−3)				0.04	
own lag_1	0.01	0.12	0.13	0.21	0.35
own lag_2		0.31	0.17	0.22	
own lag_3		0.13	0.15		

MNB Occasional Papers 121

Exchange rate pass-through after the crisis: the Hungarian experience

Budapest, September 2015