Abstract

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SECTION 1. INTRODUCTION

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The goal of pioneering works in exchange rate pass through estimation area was mainly in determining industry-specific effects in specific economies: among others, (Schembri, 1985) examines Canadian exports, (Menon, 1992, 1993) — Australian exports and Imports of Motor Vehicles, (Athukorala et al., 1994; Khosla, 1991) — Japanese exports, (Cowling et al., 1989 — UK and West German car market, Athukorala, 1991 — Korean exports, Baldwin, 1988; Feenstra, 1989; Hooper et al., 1989) — US imports. These papers show that there is a heterogeneity in pass-through across industries as well as countries though challenging data measurement errors and model misspecifications. A huge contribution to review these attempts is made in (Menon, 1993, 1995).

Looking for exchange rate pass-through for whole economies, (Khosla et al., 1989) estimate shock-independent ERPT to export prices for 23 countries using calculated quarterly nominal effective exchange rate for each economy and fitting OLS regressions. They find that pass-through effect varies drastically across countries: for developed economies this value is high, meanwhile developing ones experience low pass-through.

A more advanced methods are used in (Y. Kim, 1990) — author examines pass-through to US import prices and influence of exchange rate to mark-up using a model with time-varying parameters. It is shown that a mark-up negatively correlates with US dollar exchange rate, though a direct effect of the latter to prices fell from 1980s.

In (Deravi et al., 1995) a vector autoregression (VAR) is applied to fit US broad money aggregate, dollar exchange rate and consumer price index (CPI) with a main emphasis on monetary supply shock. Via causality test It is underlined that there is a significant causality effect of broad money to other macrovariables. Variance analysis suggests the effects to CPI from innovations to other two variables are nearly equal after four years.

(K.-H. Kim, 1998) employs vector error-correction model (VECM) in order to study pass-through to US import prices. This paper reveals a significant negative effect of US exchange rate appreciation to producer price index (PPI) and conducts causality test for this dependency, which confirms an influence of exchange rate. Moreover, author argues that previous works were using inefficient methods to examine ERPT.

In his renown paper, Taylor (2000) provides strong theoretical framework for understanding exchange rate pass-through nature. The author simulates three-equation model of individual and aggregate prices and output and shows that when the inflation is low, pricing power of firms declines as well leading to lower pass-through. Hence, if a producer wants to raise or lower their individual price due to change in costs or, equivalently, exchange rate, he or she would expect other firms stay on the remaining price level due to low inflation.

Another approach of examining exchange rate pass-through is contained in literature based on general equilibrium models, although there are few ones specially structured for studies in this particular field. Mainly based on purely statistical approach, this particular paper refers only to several works of this kind, leaving the rest to the reader.

One of the works is Adolfson, 2001, where author examines optimal policy of monetary authority under different completeness of pass-through. The main consequence of this study is that the lower pass-through is, the less important nominal economy is, as interest rate response to shocks from outside is lower and exchange rate fluctuations are higher.

The seminal paper in this field is Obstfeld et al., 2002. It does not directly touch the pass-through problem, however, it is a starting point for many papers in this field. In the paper, a cooperation of monetary authorities in a two-country model is examined. The main result of this paper is that even if monetary authorities do not coordinate with each other, benefits from macroeconomic stabilization can outweigh lack of coordination, and coordination under fixed exchange rate is more preferred than one under the floating rate.

Looking for effects of exchange rate volatility, (Devereux et al., 2002) develop a multi-economy new-Keynesian general equilibrium model based on the model from aforementioned paper. Authors show that fluctuations in nominal exchange rate appear to compensate pass-through to prices nominated in local currencies. It is argued that even if there is a little volatility in fundamental macroeconomic variables, fluctuations of exchange rate may be quite high. This model lacks empirical research though, constrained only by simulations with different parametrisation.

Basing on the same foundations, an attempt to make an empirical research based on DSGE model is done in (Smets et al., 2002), where Euro area data is used to calibrate a model and estimate exchange rate pass-through in an economy with optimal monetary policy. As a result, authors claim that under an assumption of presence of import price stickiness in the economy, its effect is similar as stickiness of domestic prices.

Gagnon et al. (2004) use Monte-Carlo approach with multi-equation model to show that there was a decline in exchange rate pass-through since 80s due to inflation stabilisation policy conducted by many central banks across the world. To find more evidence, authors fit an OLS regression with lags of exchange rate summed with foreign CPI for two subsamples individually chosen for 20 countries. Additionally, they estimate interest rate rule coefficients in order to find changes in monetary policy. Finally, authors argue that the hypothesis is confirmed.

A new wave in studying exchange rate pass-through — use of structural vector autoregressions (SVAR) — starts from (Hahn, 2003) for Euro area macro data from 1971 to 2002. In this remarkable work, a recursive (also known as *Cholesky*) identification scheme is used in order to recover macroeconomic shocks to PPI and HICP from other different macroeconomic variables (oil price, interest rate, output gap and non-oil import prices). To address statement about pass-through decline in (Gagnon et al., 2004), author conducts a robustness test and finds out that there was no significant change in pass-through effect for the Euro area.

The same conclusion about decline, among other ones, is made in (Campa et al., 2005). Searching for the pass-through effect to import prices, authors examine data for 23 countries and assert that the pass-through effect is incomplete for all countries in the short run and for overwhelming majority of them in the long run.

(Ca' Zorzi et al., 2007) and (McCarthy, 2007) papers resemble previously cited (Hahn, 2003). The first work studies data for 12 developing economies and employs recursive SVAR to estimate shock-dependent ERPT; authors find that pass-through effect fades down to the distribution chain and argue that when inflation in a developing economy is low, ERPT is comparable to one of developed countries.

On the opposite side is (McCarthy, 2007) work, where data for nine developed countries are examined applying Cholesky identification scheme to VAR. Author states that pass-through in developed economies is quite low, and inflation in the US is mainly driven by oil shocks, producer price shocks and internal CPI shocks.

In (Shambaugh, 2008) paper author uses long run restrictions for SVAR in order to identify link between exchange rate and CPI together with import prices. Author uses data for 16 countries for the time frame from 1973 to 1999 and obtains supportive evidence that low inflation declines pass-through — for some countries, CPI growth rate does not respond to exchange rate shocks in the same magnitude as producer price index growth rate.

Data's granularity higher than quarterly is not usually found in the studies, although (Amstad et al., 2010) observe monthly Swiss CPI and NEER from 1993 to 2008. This work employs event study approach to estimate an effect of monthly import price time series release to ERPT. Author underlines that this method is more suitable for policymakers due to possibility of using the most current data and does not rely on VAR restrictions, which may be controversial. The criticism of SVAR is quite questionable in this light, since the monthly data does not impair a possibility of proper identification of shocks, while the benefits of shock-dependent ERPT are higher for monetary and macroprudential authorities.

An innovative identification method is introduced by An et al. (2012) — author employs sign identification scheme in order to obtain price-to-exchange rate ratio (*PERR*, shock-dependent exchange rate pass-through), which will be described in the following Section. Author fits the model for eight developed economies and claims that for the most cases pass-through is incomplete. Another conclusion is that pass-through is higher for small-sized economies with more volatile monetary policy.

The work of Delatte et al. (2012) is devoted to determination of pass-through asymmetry for four countries (Germany, Japan, UK, US) from 1980 to 2009. An ARDL with nominal exchange rate changes divided into two variables (with positive and negative increments) is estimated to determine both short-run and long-run asymmetric ERPT. Author argues that pass-through is smaller during local currency appreciations.

(Brun-Aguerre et al., 2012) paper's aim is to find what drives ERPT to import prices. Authors use both ECM and panel fixed effects (FE) model to catch time- and country-specific effects for 37 countries on 1980–2009 period; again, pass-through asymmetry is considered. The conclusion is that there is no evidence of pass-through declining for both developed and emerging economies, although domestic tariffs and import-to-export ratio matter.

Monthly data of Taiwanese economy under deflation are examined in (Lin et al., 2012). In this work, a two- and three-regime threshold autoregression (TAR) models are fit to find non-linearities in pass-through relation. It is argued that pass-through declines only when inflation is close to zero, and the link of ERPT and inflation is V-shaped. With this non-trivial result, high rates of deflation are unpleasant for an economy additionally from the side of exchange rate pass-through.

Another work observing Asian economy is (Jiang et al., 2013). Authors estimate SVARs with custom shock matrix in order to find PERR for China. This method is more flexible than recursive identification scheme as the shock matrix does not necessarily need to be triangular, although application of such scheme is quite situational. Authors conclude that PERRs are incomplete, which is usual for the literature in this field.

(Yamada, 2013) paper is devoted to study exchange rate regime effect to inflation among developing and emerging economies. Author fit treatment effects model with propensity score matching based on GDP and geographical characteristics in order to calculate between inflation targeting regime and other ones. The conclusion is that inflation targeting exchange rate regime performs at least not worse than fixed regime in terms of inflation lowering.

Multi-currency study for 17 countries of Euro area is done by de Bandt et al. (2014) to estimate effect of exchange rate fluctuations to import prices for multiple trade partners. Currencies chosen are US dollar, UK pound-sterling and Chinese Renminbi. Authors estimate FE model in order to calculate ERPT and find out that in the short run pass-through is incomplete, but its completeness is confirmed fore the long run.

In order to look for the changes in pass-through after 2008 financial crisis, Jasova et al. (2016) estimate 6-year rolling ERPT for both developing (11) and advanced (22 countries) economies completing their study by fitting two-way FE model. Authors assert that pass-through declined during financial crisis for developing economies, meanwhile ERPT of developed countries remained on a relatively stable level.

In (D. Comunale, 2017) paper data for four Euro area countries — France, Germany, Italy and Spain — are studied to find both ERPT and PERR under the zero lower bound (ZLB) environment. Instead of short-term interest rate, authors make use of calculated *shadow interest rates* and estimate Bayesian VAR with sign and zero restrictions. The results of the study are that pass-through is high and volatile to import prices and, in general, is dependent on shocks evolving. Moreover, authors state that the process of choosing identification scheme is quite sophisticated, and the identified shocks are true only conditional on the scheme involved.

Both FE model and sign and zero restricted SVAR are estimated in (K. J. Forbes et al., 2017), where authors try to analyse time- and country-specific differences in pass-through on the sample of 26 countries. It is argued that structural variables, like the first two statistical moments of inflation and exchange rate are important for time and country effect explanation, while structural shocks are crucial for explanation of macro-variable variation in time.

A quite remarkable paper of the same collective is (K. Forbes et al., 2018). Authors study UK economy pass-through before and during Brexit using SVAR model with sign and zero restrictions. The study shows that pound-sterling's depreciation periods during 2008 financial crisis and Brexit have different ground and discrepancy in inflation rates are caused by different shocks affecting the economy. Authors admit that set up in this fashion, a model cannot capture all the complexity of pass-through nature, although identification of shocks can help to improve relevant policy by monetary authority.

Another work employing the same method is (An et al., 2020), though *narrative* sign restrictions (simply put, signs dictated by historical events) are added. The main drivers of Japanese pass-through are examined in this study. Authors argue that narrative sign approach is more promising in terms of shock identification procedure.

Time-varying ERPT is examined in (Leiva-Leon et al., 2019), where authors estimate time-varying parameters (TVP) dynamic factor model and SVAR with sign restrictions for Euro area. A TVP approach is quite innovative in pass-through literature, as it is highly likely to solve problems of non-linear ERPT estimation. The paper's conclusion is that inflation is mostly driven by exogenous exchange rate shocks, though core inflation are less exposed to these fluctuations.

Colavecchio et al. (2019) use local projection method in order to capture non-linear pass-through effects for the 19 countries in Euro area. Plainly speaking, local projections are h-ahead forecasts on the basis of current data. The results show that there is no complete pass-through for all the countries neither after a one year nor two years. Authors also find there is a sign and exchange rate shock size non-linearity for some countries.

The recent work of M. Comunale (2020) is devoted to a comparison of Bayesian SVARs and DSGEs for the purposes of PERR estimation. This particular work is important in the sense that SVAR and DSGE models can give controversial results; hence, a policymaker needs to distinguish an appropriate aims for both setups. Author finds out that just after a shock PERR's are identical for both models, although in the long-run estimates from DSGE are higher due to endogenous response of macrovariables.

DSGE is also employed in (Garcıa-Cicco et al., 2020), where comparison of shock-independent ERPT's and PERR's is done on Chilean data. It is argued throughout the work that pass-through conditional on shocks gives a full picture of macroeconomic variables' relations and that DSGE models are helpful to generate prudent monetary policy.

In their latest work, K. Forbes et al. (2020) estimate both SVAR and FE model for a set of 26 countries in order to review *"shocks vs. structure"* dilemma. Authors claim that both structural characteristics and shocks are important for better understanding pass-through. Also they find an evidence that monetary shocks are associated with large PERR, which made a big contribution to price fluctuations in advanced economies that are not close to the lower bound.

All in all, there has been a shift in the literature from industry-specific studies to understanding of shock importance during exchange rate pass-through estimation. This drift is dictated not only by an evolution of methodology (this point is prevalent though), but data availability and, what is the most important, a switch to macroprudential policy. The intervention of vector autoregressions and Bayesian estimation techniques, especially sign restrictions, have given, in some sense, the second breath to the research. On the other hand, further development of DSGE models and acceptance of them by central banks globally gave an idea of how shock-dependent ERPT should look like for each country. The idea of SVAR being guided by DSGE (at least for a signs) is given in (Ortega et al., 2020), which is a brilliant review on the topic of exchange rate pass-through estimation with a focus on Euro area and the US.

Russia and CIS

One of the first works exploring the Commonwealth of Independent States (CIS) is (Korhonen et al., 2006), where authors estimate VAR models for each country-member using the data from 1999 to 2004. This work is quite disputable as, for instance, Russia and Kazakhstan have *negative* ERPT, which is highly unlikely even considering the policy in 2000-2005. On the example of Russia, the most obvious omission in this study is that the model is fitted on the data after 1998 default, which led to huge surge in exchange rate due to risk premium shock. After this event, it seems that the ruble was underappreciated, and its exchange rate was lowering from 2000 till 2008 financial crisis together with rise in oil price. Being fitted with this data, VAR may generate biased results, as effects from several shocks are not taken into account: the influence of risk premium shock was declining together with rising oil price shock. Due to this issue, the cited work is a good example why shock-dependent ERPT is important for proper policy implications.

Oomes et al. (2005) study the relation of inflation and money demand on the example of Russia in 1996–2004. The main focus of this work is economy dollarisation — the influence of foreign currency holdings in the country. As a side effect, authors estimate ERPT to control for influence of money aggregate to inflation.

DSGE model for Russia and China is presented in (Sosunov et al., 2006), where a response of exchange rate to foreign currency accumulation policy by central bank is studied.

Analysis shows that low level of money in the economy for Russia in 2001–2005 is a pro-inflammatory factor. Moreover, it is underlined that the management of real exchange rate by means of currency accumulation has a little effect to it.

The model for Russia in Dobrynskaya et al., 2008 is fitted on the sample from 1995 to 2002. This work employs two single-equation regressions (simple and extended ones) in order to estimate ERPT for the country. The choice of time-frame and model (OLS) there is questionable, since the surge of CPI and exchange rate after 1998 crisis probably lead to unreliable estimates, although results in the paper do not contradict economic intuition. Authors argue that exchange rate pass-through in Russia is fast, as a huge part of it comes into inflation right in one month after exchange rate shock.

In (Kataranova, 2010), a more recent (2000–2008) data are used. Author fits different specifications of OLS models accounting for asymmetry in order to evaluate ERPT. The results are that the presence of pass-through asymmetry is confirmed for Russia, and the effect of ruble decline caused by 2008 financial crisis, indeed, was only partial, since the following ruble appreciation strengthen credibility of the local currency.

Beckmann et al. (2013) estimate exchange rate pass-through for CIS countries on the data from 1999 to 2010. Authors obtain negative ERPT estimate from VAR model for USD-RUB pair and the same estimate close to zero for EUR-RUB pair, which contradicts reality. Authors address this issue by pointing out that this pass-through estimate is shock-independent and there is an uncontrolled effect of oil shocks.

In (O. Faryna, 2016a) paper author estimate bilateral panel VAR for both Russia and Ukraine to study cross-country spillovers. Author claims that depreciation of ruble causes increase in Ukrainian inflation, which follows intuition as Russian and Ukrainian economies have been quite integrated (even after 2014 political crisis this degree of integration remains high). Moreover, it is observed in this paper that inflation in Ukraine responds to USD-RUB changes higher than to USD-UAH (Ukrainian hryvna — the Ukrainian local currency).

The other work of the same author is (O. Faryna, 2016b) examines non-linear ERPT for Ukraine exclusively on the 2007–2016 data. As it is occurred in the literature for Russia, local currency depreciation causes more effect to inflation than its appreciation, although pass-through is higher for Ukraine.

The latest work of H. Faryna (2018) studies some CIS countries, including Ukraine and, partially, Russia. This paper employs Global VAR model (VAR with equations for each country) fitted on the 2001–2016 data. This model shows a close relationship of Euro area output and output of CIS countries. An oil shocks is definitely positive for Russia, and what is more interesting, it is positive for CIS countries due to spillover effect.

A VECM model with ERPT asymmetry is used in (Ponomarev et al., 2016) for 2000–2012 data. Authors break time series into two sub-periods: before and after 2008 financial crisis, to achieve robustness of results. Their findings are that pass-through effect reveals fully in 6-12 month period, and this effect is incomplete.

(M. Comunale et al. (2018)) estimate fixed effects panel data model with ERPT asymmetry for CIS countries on a 1999–2014 time-frame. Authors claim that pass-through is high for CIS countries, reaching its maximal level in one year. Moreover, they argue that there is a little evidence of asymmetric pass-through, although it doesn't apply to Russia, as this country is just partially included in the study.

In Sinyakov et al., 2019 authors calibrate a simplified multi-equation static model of small open economy in order to evaluate industry-specific asymmetry of ERPT. Their findings are that electrical appliance manufacturers and paper producers are less likely to transfer costs from exchange rate fluctuations directly to the final good prices keeping ERPT in this area low. On the opposite, textile and wholesale trading industries have very high pass-through (around 0.5–0.6). The final conclusion is that when produces are aware of competitors' actions in a specific industry, the pass-through is high.

A comprehensive research is done by Khotulev (2020). In this paper, author evaluate both ERPT by means of OLS model and PERR using the Bank of Russia's DSGE model. There are five exogenous shocks in the latter model: oil price, monetary policy, country risk premium, government expenditures and reserves. A quite disputable result is that there is a huge negative PERR due to government spendings shock (-1.596), which may be a result of purely technical restrictions (first-order approximation of DSGE equations).

The evolution of pass-through literature concerning Russia and CIS is fairly limited, since the last work calculates shock-dependent ERPT. Some works include negative shock-independent ERPT for Russia, which is rather dubious for Russian structure of external trade. Table 1 summarizes shock-independent pass-through estimates across the literature.

Paper	Currency	Data	Infl. aggr.	Length	ERPT
Oomes et al., 2005	NEER	1996–2004	CPI	Short-run	0.4–0.5
Korhonen et al., 2006	USD	1999–2004	ULC[1]	Long-run	-0.42
Dobrynskaya et al., 2008	NEER	1995–2002	CPI	Long-run	0.35
Kataranova, 2010	USD	2000–2008	CPI	Short-run	0.6–0.20
Beckmann et al., 2013	USD	1999–2010	CPI	Long-run	-0.17
Ponomarev et al., 2016	NEER	2000–2012	CPI	Short-run	0.046
O. Faryna, 2016a	USD	2000–2015	CPI Core	Long-run	0.1
Sinyakov et al., 2019	NEER	(2016–2017)	CPI	Long-run	0.35
Khotulev, 2020	NEER	2005–2019	CPI	Long-run	0.16

Table 1: Earlier shock-independent ERPT estimates for Russia.

SECTION 3. METHODS

Theoretical Framework

Before going into technical details, I explain why pass-through estimation is important, and why there may not be complete pass-through (a one-to-one correspondence of exchange rate fluctuations and inflation movements). In this step, I follow (D. Comunale, 2017) and (K. Forbes et al., 2018), which use the most popular approach in the pass-through literature.

First, consider the following aggregate pricing equation of an imported good:

$$P_t^{imp} = E_t \cdot P_t^{exp,f}, \tag{1}$$

where P_t^{imp} is price of imported good, E_t is exchange rate and P_t^{exp} is foreign export price.

Then, taking logarithm on both sides and using $P_t^{exp,f} = Markup_t + MC_t$, we get:

$$p_t^{imp} = e_t + markup_t + mc_t, \tag{2}$$

where p_t^{imp} is a logarithm of import price level, e_t is a logarithm of exchange rate, $markup_t$ is producer mark-up and mc_t is marginal costs. From this relation, any exchange rate fluctuation can be either fully incorporated into import price, or markup and marginal costs can be adjusted. Assuming that firms-exporters are forward-looking, and taking into account price rigidities, one expects mark-up to be adjusted over marginal costs for any macroeconomic shock.

The same logic, indeed, can be applied to firms-importers, such as electronics and clothing stores for Russia. These firms has marginal costs mainly consisting of foreign wholesale prices of goods imported. Due to price stickiness and oligopolistic nature of the market in these industries, the firms adjust not only the final price of good, nominated in a local currency, but their mark-up. Hence, depending on the shock emerged in the economy, the pass-through can be incomplete.

Pass-through Definitions

A simple technique that has been widely used in the literature determining shock-independent ERPT (*exchange rate pass-through*) is estimation of the following OLS model:

$$p_t = \beta_0 + \beta_1^e e_t + \beta_2^e e_{t-1} + \beta_3^e e_{t-2} + \beta_4^e e_{t-3} + \ldots + \varepsilon_t, \tag{3}$$

where ellipsis denotes other control variables. Hence, ERPT is calculated as the sum of exchange rate lags coefficients:

$$ERPT_T = \sum_{j=0}^{T} \beta_t^e - j, \tag{4}$$

where T is an ERPT horizon. An ERPT can be characterised as *incomplete* for a horizon T, if $ERPT_T < 1$.

The most important benefit of this method is its flexibility: a researcher can specify different forms of the price equation, e.g. panel regressions for spatial data, allowing asymmetry of exchange rate effect or time/regime-dependent effects. Non-linear effects are important for studying pass-through dependence on shifts in monetary policy or structural breaks.

In the same time, ERPT measure lacks identification of economy shocks, which is a huge shortcoming of this approach. This issue is misleading in the sense that it can result into wrong policy conclusions made by authorities: if shock-independent pass-through is measured involving the historical data with a particular set of shocks, the result cannot be applied to the further periods when the economy is exposed to other set of shocks, since ERPT will be different.

In this work I calculate shock-dependent pass-through known in the literature as *price-to-exchange rate* coefficient (PERR). This measure involves estimation of structural vector autoregression, either purely data-driven or derived from DSGE equations. PERR is defined as follows:

$$PERR_T^{\text{shock}} = \frac{\sum_{t=0}^{T} IRF_t^{\text{shock},e}}{\sum_{t=0}^{T} IRF_t^{\text{shock},p}}, \tag{5}$$

where $IRF_t^{\text{shock,response}}$ is a value of impulse-response function calculated at time t for *response* variable caused by exogenous *shock*. e and p stand for exchange rate and price variables.

Econometric Model and Identification Strategies

I estimate vector autoregression (VAR) with six variables: oil price in US dollars, world import price index, local short-term interest rate, local nominal effective exchange rate, local GDP per capita index and local consumer price index (CPI). All variables, except interest rate, are used in differences of logarithms.

The choice of variables here is dictated mainly by (Hahn, 2003;D. Comunale, 2017; K. Forbes et al., 2018; Leiva-Leon et al., 2019). Throughout the shock-dependent pass-through literature widely used variables are import/export price index, short-term interest rate, exchange rate, output and inflation aggregate. Moreover, quarterly data are used instead of monthly series due to primarily absence of monthly output data. In addition to this, I include oil price series, as this variable is crucial for determining exchange rate fluctuations and movement of other macrovariables in Russia. More information about data is in the following section.

After VAR estimation a researcher gets its reduced form, which is given as follows:

$$Y_t = A_1 Y_{t-1} + A_2 Y_{t-2} + \ldots + u_t, \tag{6}$$

where Y_{t-j} is vector of variables' values of lag j, A_i is matrix of coefficients at lag j, u_t is residual term.

Since variance-covariance matrix Ξ_{u_t} is not diagonal and normalized, there is no economic interpretation of residuals u_t. Then, a researcher has to refer to the generalized form of VAR:

$$B_0 Y_t = B_1 Y_{t-1} + B_2 Y_{t-2} + \ldots + \epsilon_t, \tag{7}$$

$$Y_t = B_0^{-1} B_1 Y_{t-1} + B_0^{-1} B_2 Y_{t-2} + \ldots + B_0^{-1} \epsilon_t, \tag{8}$$

$$Y_t = B_0^{-1} B_1 Y_{t-1} + B_0^{-1} B_2 Y_{t-2} + \ldots + u_t, \tag{9}$$

From this result and equivalence $B_0^{-1}\epsilon_t = u_t$, an idea is to specify such matrix B_0 that a variance-covariance matrix $\Xi_{\epsilon_t} = V(B_0 u_t)$ is diagonal and normalized ($\Xi_{\epsilon_t} = I$). Again, using $B_0^{-1}\epsilon_t = u_t$ a researcher gets:

$$\Xi_{u_t} = V(B_0^{-1}\epsilon_t) = B_0^{-1} V(\epsilon_t \epsilon_t') B_0^{-1\prime} = B_0^{-1} I B_0^{-1\prime} = B_0^{-1} B_0^{-1\prime}. \tag{10}$$

This result is noticeable, as it appears that $B_0^{-1} B_0^{-1\prime}$ is a Cholesky decomposition of Ξ_{u_t}. Then, $\dfrac{n(n-1)}{2}$ restrictions are needed to exactly identify shocks.

There are several methods to do this. The first method is the simplest one — Cholesky decomposition of variance-covariance matrix Ξ_{u_t}, as it is given above. From it, on the example of the model estimated in this work, shocks are identified as follows:

$$\begin{pmatrix} \epsilon_t^{\text{Global supply shock}} \\ \epsilon_t^{\text{Global demand shock}} \\ \epsilon_t^{\text{Mon. shock}} \\ \epsilon_t^{\text{Ex. rate shock}} \\ \epsilon_t^{\text{Output shock}} \\ \epsilon_t^{\text{Inflation shock}} \end{pmatrix} = \begin{pmatrix} S_{11} & 0 & 0 & 0 & 0 & 0 \\ S_{21} & S_{22} & 0 & 0 & 0 & 0 \\ S_{31} & S_{32} & S_{33} & 0 & 0 & 0 \\ S_{41} & S_{42} & S_{43} & S_{44} & 0 & 0 \\ S_{51} & S_{52} & S_{53} & S_{54} & S_{55} & 0 \\ S_{61} & S_{62} & S_{63} & S_{64} & S_{65} & S_{66} \end{pmatrix} \begin{pmatrix} u_t^{\text{Oil}} \\ u_t^{\text{Imp. infl}} \\ u_t^{\text{Int. rate}} \\ u_t^{\text{NEER}} \\ u_t^{\text{Real GDP}} \\ u_t^{\text{CPI}} \end{pmatrix} \tag{11}$$

Then, the first variable in the vector is not directly affected by any shock but one that is associated with this variable, while the last variable is affected by fluctuations of all the shocks in the model. The main drawback of this identification scheme is that it implies very stiff restrictions to the shocks, and there may not be desired order of variables that produces sensible result. For example, one can reasonably assume that ruble's exchange rate and inflation have effect to the interest rate in short-run (monetary policy rule reaction).

This method is used in order to provide robustness test for the results, obtained from the main model described below. Under this specific variable order, a reader shouldn't expect precise results from identification of supply (output) shock, as it is restricted to have effect only to inflation, which is a primitive assumption. Hence, I expect PERR coefficients to match at least for the first four structural shocks.

In the recent literature estimating shock-dependent pass-through, a sign and zero restrictions method is widely used. A reader can refer to (Arias et al., 2014) for rigorous explanations of this technique, although, in a nutshell, this is a *guess-and-verify* technique: one has to randomly generate matrix Q such that QQ' is diagonal and its product with matrix from Cholesky decomposition produces such IRF's that satisfy both sign and zero restrictions. As the model is underidentified, there are multiple matrices Q satisfying there restrictions. After n times generating the matrices and obtaining random IRF's, their median is calculated as a result.

The main drawback of this approach is that researcher has to intuitively arrange signs of response functions, which is usually non-trivial. (Ortega et al., 2020) advice to use DSGE models for this purpose, although data are not necessarily to match IRF produced by theoretical justifications. Another shortcoming is that sign and zero approach is hard to apply to data that presume heterogeneous covariance matrix, and it mostly comes from the problem of sign setting. Moreover, this approach is quite sensitive to data atomicity: for example, quarterly data require less effort to determine signs than monthly series, since data with less periodicity is averaged, and there are less fluctuations, which harmfully affect the process of model fitting and further identification.

Talking about Russian data, one can recall the exchange rate policy shift in November 2014, when banded exchange rate was replaced by floating one. Hence, it is not expected that Russian macroeconomic time series exhibit homogeneity, while economic agents' adjustment to the new policy has been continuous. Then, sign and zero approach seems fraught with difficulties; moreover, there can be no plausible identification found if sign restrictions are set thoroughly from the theory.

In this work, I use smooth covariance transition technique proposed in (Lütkepohl et al., 2017), which seems to produce more credible results than Cholesky decomposition, on the one side, and is more data-driven than sign and zero restrictions. The idea of this method is that if the residual term is heteroscedastic, its variance-covariance matrix can be defined as a sum of two non-linearly weighted variance-covariance matrices, which represents gradual and time-dependent transition from one volatility pattern to another one. Below I provide the widely used transition function, which is employed in this work:

$$\Xi_{u_t,t} = f(t)\Xi_a + (1 - f(t))\Xi_b, \tag{12}$$

$$\Xi_a = B_0^{-1} B_0^{-1\prime},\ \Xi_b = B_0^{-1} \Lambda B_0^{-1\prime}, \tag{13}$$

$$f(t) = \frac{1}{\left(1 + e^{e^{\gamma} \cdot (t-c)}\right)}, \tag{14}$$

where c is period of transition and γ is smoothness parameter. The search for optimal c and γ is done via grid search, although these values can be specified. I specify c to be equal to the period when the Central Bank of Russia has switched to the floating exchange rate regime (November 2014) and do search grid for optimal γ on the space $(-5, 5)$ with step 0.02. In order to estimate the model, I use implementation from `svars` package in R (Lange et al., 2021).

The model's parsimony is predominantly caused by identification issues, while DSGE models may include dozens of variables.[2] A problem of lack of identification grows with the number of variables due to requirements of identification schemes and irrelevance of odd

[2] On the other hand, the process of DSGE model estimation is not the same as the same process for SVAR, as it is either calibration or frequentist/Bayesian estimation of only few parameters.

variables.

Post-Estimation Analysis

After structural shocks are identified, a researcher is usually interested in response of variables after a unit structural shock. For this purposes, an impulse response function (IRF) is calculated, which represents deviation from a steady state of a variable for different periods after an (orthogonal) shock.

In order to calculate IRFs, the model should be written in terms of vector moving average (VMA) model:

$$Y_t = \sum_{j=0}^{\text{inf}} \Phi_j \epsilon_{t-j}, \tag{15}$$

where Φ_j is a matrix of MA coefficients obtained by VAR inversion. Finally, IRFs are defined as:

$$IRF_j^{\text{imp,resp}} = \Phi_j^{\text{imp,resp}}, \tag{16}$$

where ”imp” is a row index of matrix Φ_j, ”resp” is a column index of the same matrix.

SECTION 4. DATA

I use monthly time series of six variables from 2003M3 to 2020M8: oil price in USD, world import prices, short-term interest rate, nominal effective exchange rate, real GDP per capita and CPI. Monthly data are more interesting for policy conclusions in terms that it allows to make policy responds more rapidly, although quarterly data are more popular in the literature. It comes mostly from absence of monthly data in open access for the most countries (primarily, GDP) and purely academical purpose of studies.

The data time-frame is bounded by the first observation of the import price index and by the last observation of real GDP per capita. There are 186 observations in total. World import price index growth rate and CPI growth rate are seasonally adjusted according to X13-ARIMA-SEATS; GDP per capita index growth rate is available already seasonally adjusted. NEER is taken with opposite sign, as it is more convenient to work with a direct quote. See Table 2 for data description and sources.

I associate oil price fluctuations with global persistent shock (as in K. Forbes et al., 2018), import price fluctuations with global supply shocks, interest rate fluctuations with monetary policy shocks, exchange rate fluctuations with exogenous exchange rate shocks, domestic product fluctuations with aggregate supply shocks and CPI fluctuations with aggregate demand shocks.

A quite intuitive fact that oil price is negatively correlated with exchange rate follows from the fact that oil producers convert their yields from convertible currencies (mostly,

Variable	Description	Source
Oil price	Brent oil price nominated in US dollars, monthly average	Bloomberg
Import prices	World import price index, manufactured goods only (2005m1 = 100), seasonally adjusted	WTO
Interest rate	MIACR 31–180 days, monthly average. Absent values were imputed by last available rate	CBR
Exchange rate	Nominal effective exchange rate (NEER) of ruble, Ruble appreciation = NEER decline, MoM, Central Bank of Russia methodology	CBR
Output	Russian real GDP per capita (1995m1 = 100), OECD leading indicator, seasonally adjusted	FRED OECD
Inflation	Consumer price index, MoM, seasonally adjusted	Rosstat

Table 2: Description of Data and Sources

US dollar and Euro) to Russian ruble, which creates excessive supply of foreign currency. In the same time, ruble appreciation declines consumer prices due to the import-dependent structure of consumer basket (huge share of import in manufactured goods). These relations form a pass-through channel.

In general, the Russian economy can be described by the following stylized facts:

1. There is a strong negative correlation of oil price and ruble exchange rate (especially, USD and EUR);
2. There is a strong negative correlation of interest rate in Russia and Russian GDP;
3. Oil price and Russian GDP are positively correlated;
4. Exchange rate and CPI are positively correlated;
5. Exchange rate is more volatile than CPI;
6. Russian GDP and CPI are negatively correlated.

Appendix contains additional information about characteristics of time series. Table a displays correlation coefficients of variables, Table b provides unit root tests for them. MIACR exhibits unit root presence, although there is no visual evidence of it, and short-term interest rates are used in the literature without differencing.

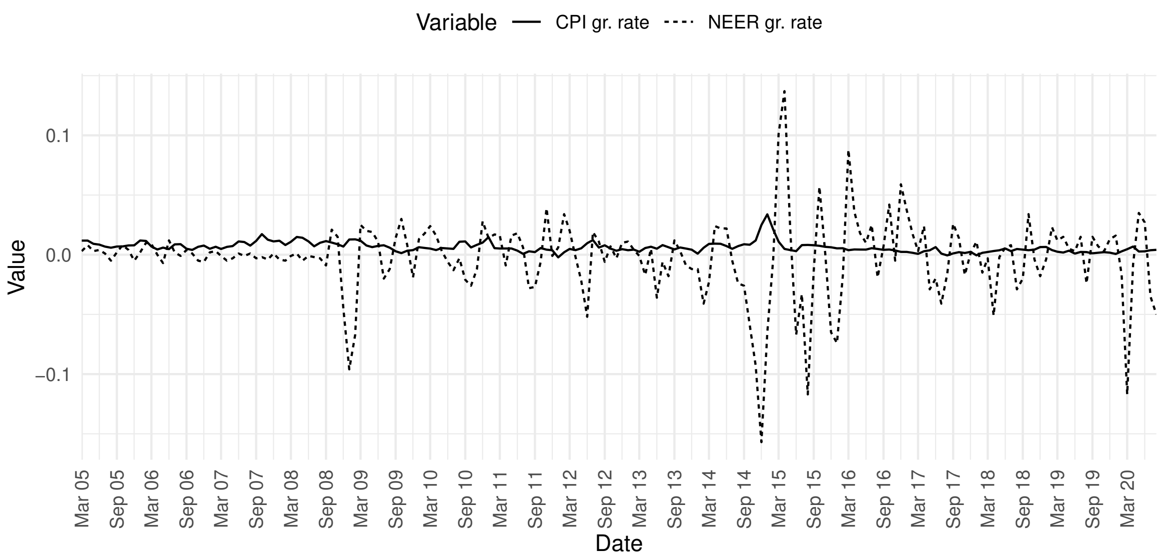

Figure 1: Time series of NEER and CPI (seasonally adjusted), growth rates.

SECTION 5. RESULTS

To obtain the results, I estimate VAR of order 4. The order of model is picked based on AIC (Akaike information criterion) calculated for models up to 10 lags. I use the following order of variables: oil price growth rate, import price growth rate, short-term interest rate, NEER growth rate, real GDP per capita growth rate, CPI growth rate. Although variable order is not important for smooth covariance transition, it is crucial for robustness test purposes.

After VAR estimation, smooth covariance transition algorithm is applied in order to recover orthogonal impulse-response functions. Then, PERR coefficients are calculated for two time periods: the short-run (4 months) and the long-run one (12 months). For purposes of positive analysis, historical decomposition for CPI is calculated.

As a robustness test, I apply recursive (Cholesky) identification scheme. This approach is more rigid; hence, a reader shouldn't expect this method to be more precise than the first scheme.

Tables 3 and 4 provide PERR estimations. Shock-dependent pass-through coefficients calculated from smooth transition scheme are incomplete for all the variables, except global demand and output shocks. The latter shock is substantial for the economy, as even small output fluctuations generate huge deviations of other macroeconomic variables. Hence, this result might be the case for Russian economy. On the same time, there is a huge difference of results from the recursive scheme. I don't expect PERR for output from Cholesky decomposition to be robust as output shocks have influence only to CPI (due to variable order), while there may be more effects to other variables.

The result for global demand shock is more controversial, although there is little difference in coefficients from two models. It might be a result of weak identification as well as natural result due to behaviour of retail sellers, which hedge import price volatility risks by rising local prices.

IRFs for smooth covariance transition identification scheme are provided in Appendix (Figure d and e).

Figure 2 contains historical decomposition for CPI growth rate for observations from 2017. See Figure f in Appendix for full historical decomposition graph.

Shock	Smooth cov.	Cholesky dec.
Global persistent shock	0.1219	0.1216
Global demand shock	0.4844	0.2354
Monetary shock	0.4703	0.4798
Exog. exchange rate shock	0.1509	0.1514
Output shock	3.7154	0.1591

Table 3: PERR calculated from different models, short-run (4 months).

Shock	Smooth cov.	Cholesky dec.
Global persistent shock	0.1361	0.1354
Global demand shock	1.4662	1.5997
Monetary shock	0.4111	0.4229
Exog. exchange rate shock	0.2858	0.2871
Output shock	3.0022	0.0664

Table 4: PERR calculated from different models, long-run (12 months).

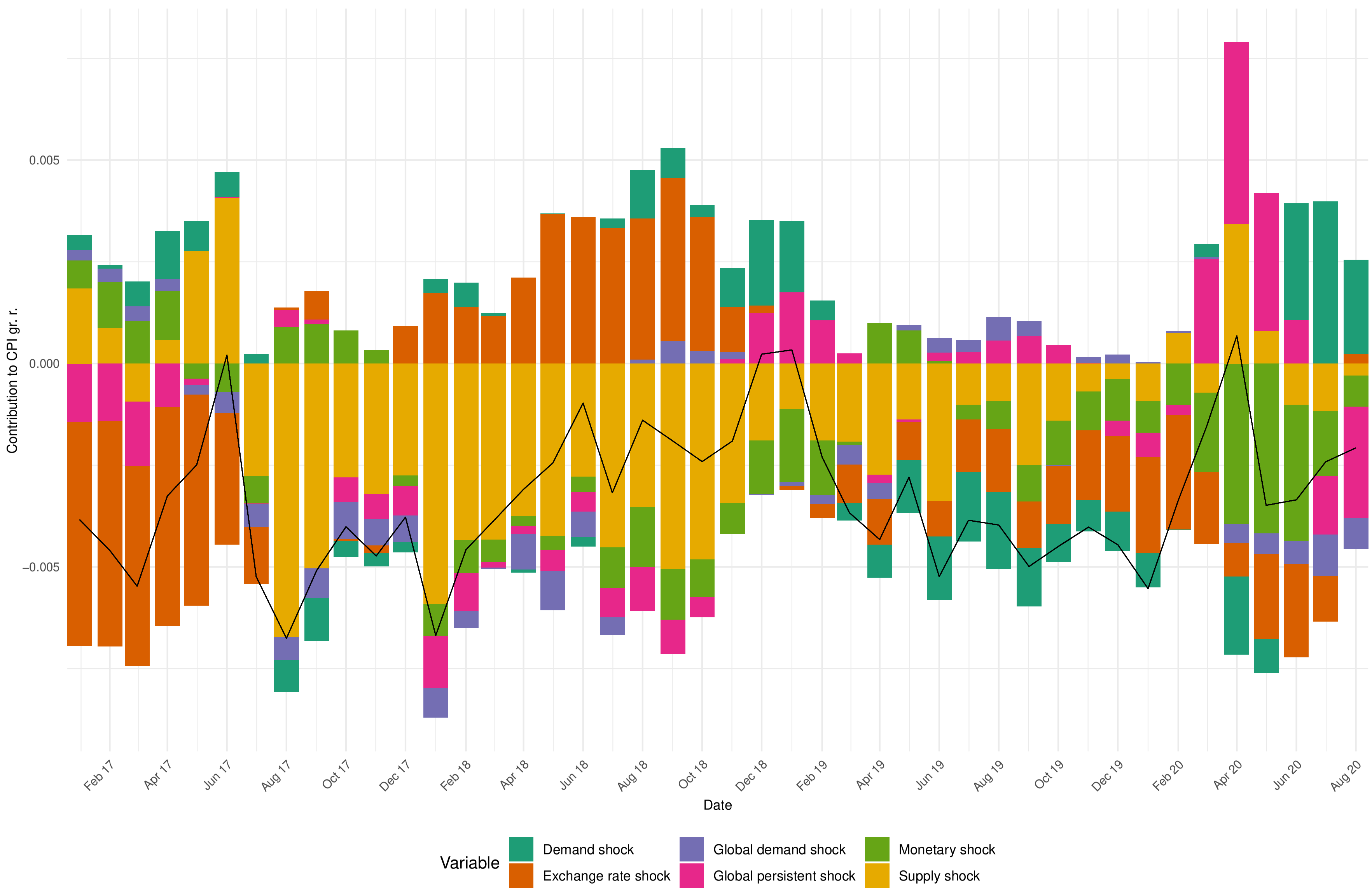

Figure 2: Historical decomposition of demeaned CPI gr. r. time series (seasonally adjusted), observations from 2017. Black line denotes demeaned observed CPI gr. r.

DISCUSSION

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Appendix

	Oil price	Import price	MIACR	NEER	GDP	CPI
Oil price	1					
Import price	0.12	1				
MIACR	-0.13	-0.30***	1			
NEER	-0.35***	-0.04	0.16*	1		
GDP	0.19**	0.32***	-0.57***	-0.13	1	
CPI	-0.05	-0.10	0.29***	0.27***	-0.19**	1

Table a: Correlation table (Pearson) of variables, growth rates where applicable.
$^{*}p < 0.05,\ ^{**}p < 0.01,\ ^{***}p < 0.001.$
CBR = Central Bank of Russia.
FRED = Federal Reserve Economic Data.
OECD = Organization of Economic Co-operation and Development.
Rosstat = Federal State Statistics Service of Russia.
WTO = World Trade Organization.

Variable	ADF	PP
Oil price, gr. r	-6.2036, p<0.01	-10.868, p<0.01
Import price, gr. r.	-4.7868, p<0.01	-10.880, p<0.01
MIACR	-2.8510, p≈0.22	-2.9313, p≈0.19
NEER, gr. r.	-6.0899, p<0.01	-7.7579, p<0.01
GDP per cap., gr. r	-5.3102, p<0.01	-3.2687, p≈0.08
CPI, gr. r.	-3.8145, p≈0.02	-5.3466, p<0.01

Table b: Unit root tests for time series used in the model. Null hypothesis for both models: unit root, alternative hypothesis: stationarity.

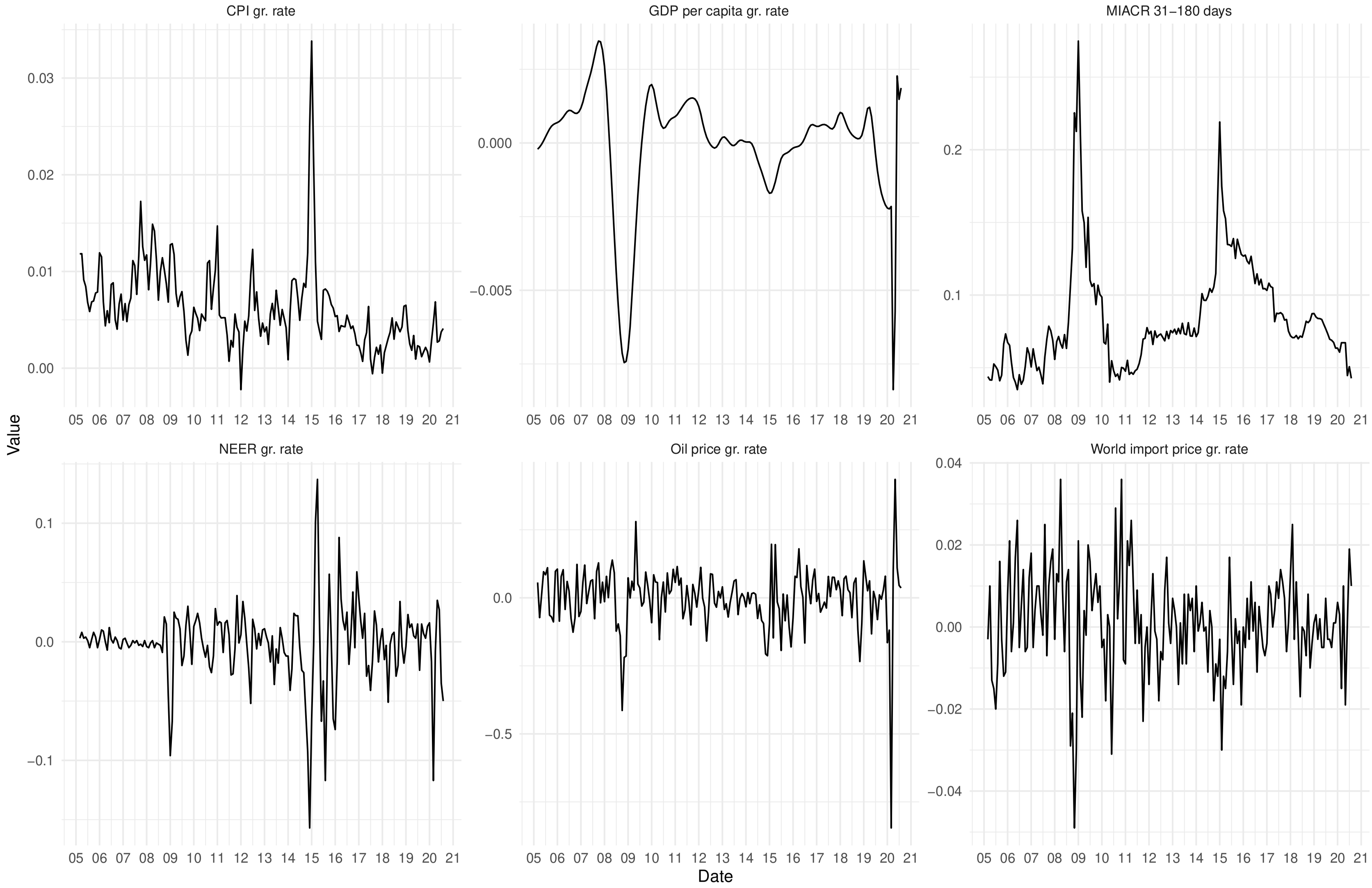

Figure c: Time series of variables used in the model, seasonally adjusted. Growth rates, where applicable.

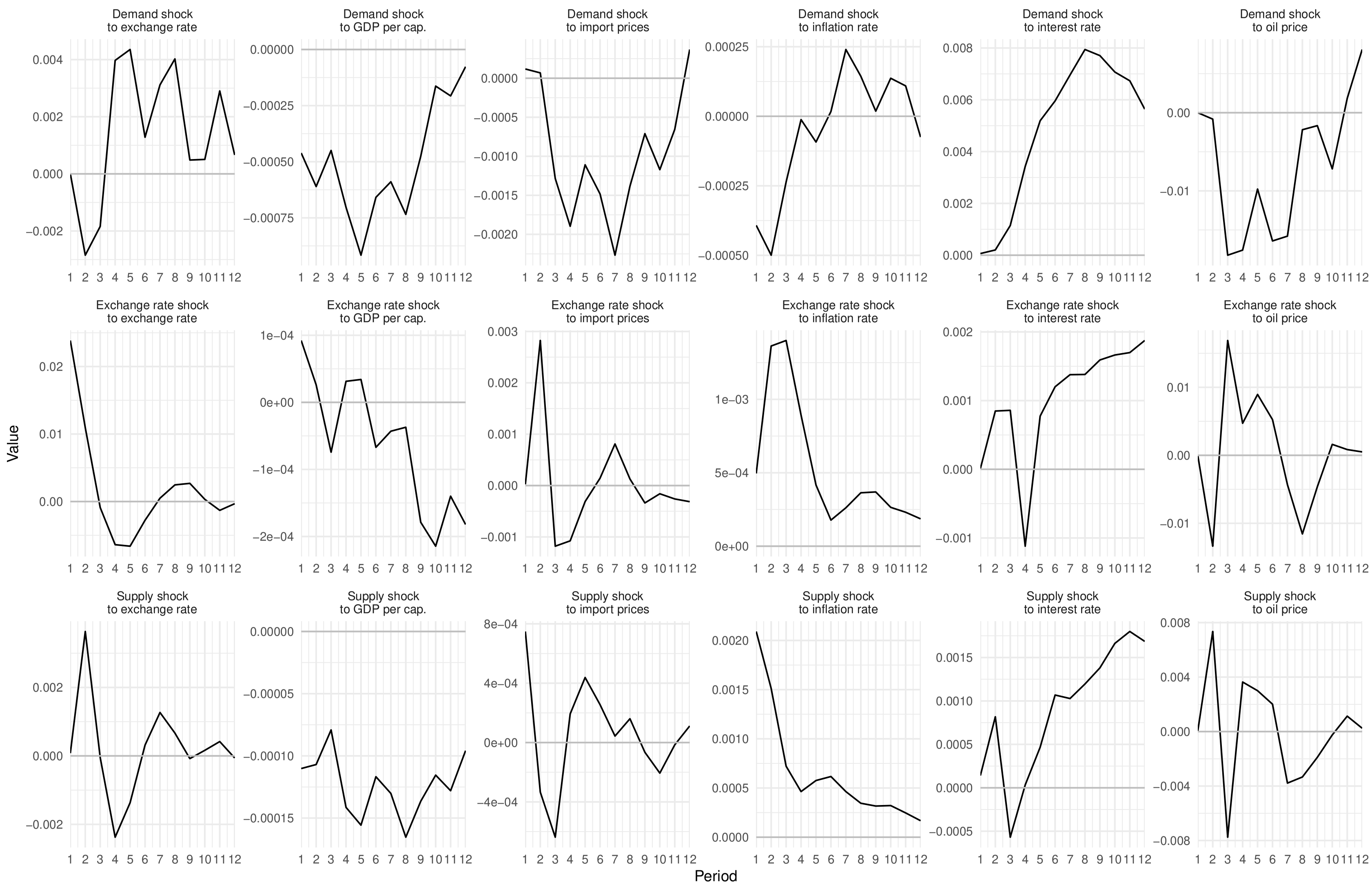

Figure d: Impulse response functions for smooth covariance transition identification scheme (1).

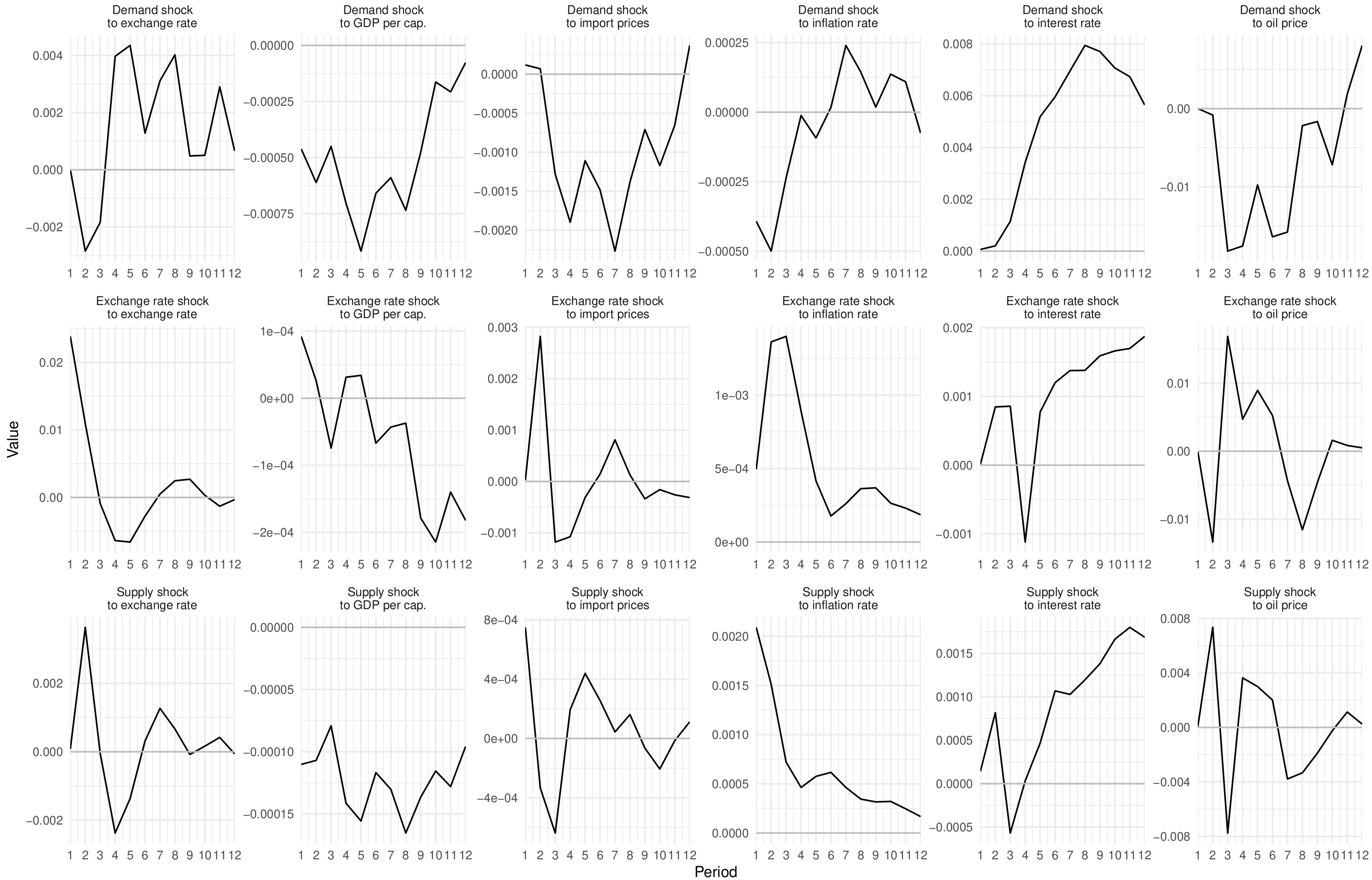

Figure e: Impulse response functions for smooth covariance transition identification scheme (2).

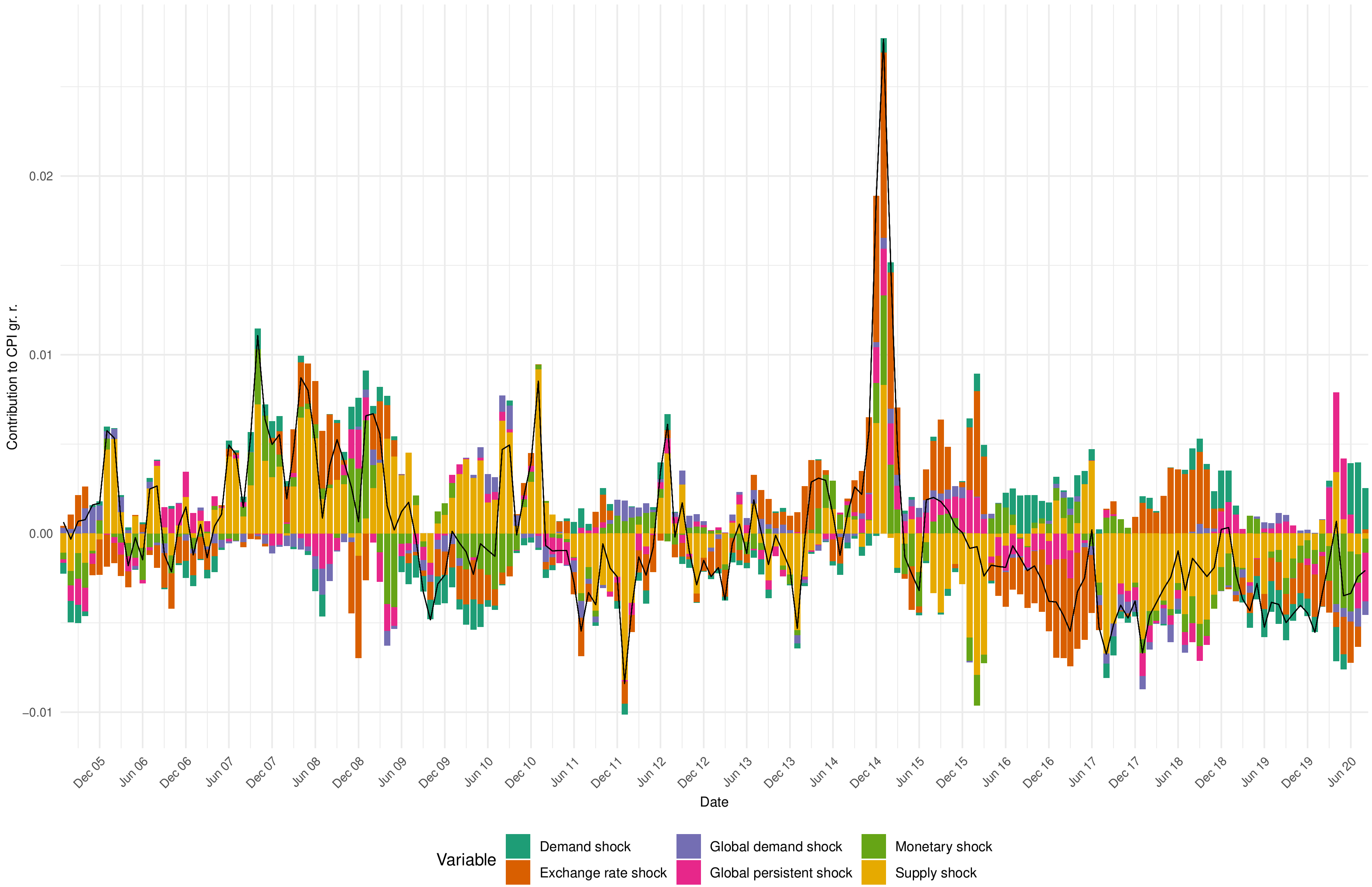

Figure f: Historical decomposition of demeaned CPI gr. r. time series (seasonally adjusted). Black line denotes demeaned observed CPI gr. r.