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VECTOR AUTOREGRESSIVE ANALYSIS OF WESTERN SANCTIONS, ECONOMIC GROWTH, AND THE REAL EXCHANGE RATE OF ROUBLE

Master's thesis

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ABSTRACT

This study aims to examine the effect that current anti-Russian economic sanctions have had on the evolution of the real exchange rate of rouble per euro. Theory suggests that the ban on foreign trade and the outflow of capital accompanying economic sanctions should cause the real exchange rate to depreciate. On the other hand, an increase in inflation caused by a growing share of non-tradable goods in the consumption basket of the target country should lead to the real appreciation of the currency. Thus, the net effect of sanctions on the real exchange rate is ambiguous. To resolve this ambiguity, a structural VAR(4) model is set up incorporating quarterly data from 1999Q1 to 2017Q3 on five variables: the sanctions index measuring the intensiveness of sanctions, the real GDP growth differential between Russia and the euro area, the short-term interest rate differential between the two regions, the price of oil, and the real bilateral Rouble/Euro exchange rate to perform impulse-response and variance decomposition analysis. The estimated impulse response functions suggest that the sanctions shock has a net negative impact on the real exchange rate, contributing to the real appreciation of the rouble, but at the same time, the introduction of sanctions might have made the real exchange rate more responsive to the shock of falling oil prices which cause the Russian currency to depreciate.

Keywords: economic sanctions, Harrod-Balassa-Samuelson effect, structural vector autoregression, impulse response functions, Russia, euro area

INTRODUCTION

The conflict between Russia and Ukraine that started in March 2014 called for the imposition of economic sanctions by the EU and its partners, against Russia. The sanctions imposed have been frequently referred to as “smart” sanctions, meaning that they target only specific sectors of the Russian economy, as well as certain individuals and entities closely related to the Russian political elite. The end goal of these sanctions, as we can see today, has not been reached – Crimea is still under the Russian control and Eastern Ukraine continues to be racked by armed conflict. It does not mean, though, that the sanctions in place have not had *any* effect on the Russian economy.

This paper is based on the assumption that if current economic sanctions target the overall competitiveness of the Russian economy, then their effect should be visible in the development of the real exchange rate of the Russian rouble over the period when the sanctions have been in place and aims to investigate the following questions:

- what was the impact of the sanctions shock on the bilateral real exchange rate of rouble per euro in terms of sign and magnitude between 2014Q1 and 2017Q3?
- what effect did conventional macroeconomic variables – such as GDP, interest rates and oil prices – have on the development of the real exchange rate of rouble per euro since the introduction of euro in 1999?
- compared with the pre-sanctions period, did the principal relationships between macroeconomic variables and the real exchange rate experience significant changes after the introduction of sanctions?

For the purposes of this thesis, five time series have been obtained: a sanctions index measuring the intensiveness of Western sanctions introduced in Dreger *et al.* (2016), real GDP growth differentials between Russia and the euro area, the short-term interest rate differentials between the two regions, the price of oil, and the real bilateral Rouble/Euro exchange rate. GDP data were obtained in real terms from the Russian Federal Statistics Service and Eurostat, respectively. Interest rates were obtained from the OECD database and combined into the real interest rate differential using inflation expectations published by the European Central Bank and the Russian Central Bank. Oil price and currecy exchange data were taken in nominal terms and converted to real ones using OECD's CPI tables. Oil spot prices were obtained from the Thomson Reuters Eikon terminal and the RUB/EUR exchange rates from the European Central Bank's database.

Overall, the following hypotheses are tested:

- the impact of sanctions on the real exchange rate of rouble per euro was positive – i.e. the imposition of sanctions contributed to the real depreciation of the value of rouble,
- during the period of sanctions, the oil price growth has been much more relevant in explaining the development of rouble's exchange rate than sanctions,
- from 1999 through 2017, the real exchange rate of rouble per euro has been on average *positively* correlated with the real interest rate differential between the Russian and euro zone regions,
- from 1999 through 2017, the real exchange rate of rouble per euro has been on average *negatively* correlated with the productivity differential between the Russian and euro zone economies.

To test the above, an econometric analysis is carried out using a structural VAR(4) model incorporating quarterly data from 1999Q1 to 2017Q3. Identification is achieved through recursive short-run restrictions. Following the estimation step, an impulse-response analysis and variance decomposition are performed to investigate the relationship of the real exchange rate with the other four endogenous variables. To evaluate the effect that sanctions might have had on the relationship between the real exchange rate and macroeconomic variables, a second estimation is conducted that excludes the sanctions period from the sample. Finally, to check the robustness of the obtained results, a third estimation is carried out using an alternative ordering of the variables in the VAR system. All estimations and plotting are performed in the Gretl software.

The rest of the paper is organized as follows. Chapter 1 presents some of the most popular theories of exchange rate formation and gives an overview of the previous empirical research on the economic impact of sanctions. Chapter 2 introduces the specificities of the VAR approach in econometrics. Chapter 3 concludes with the presentation and discussion of empirical results and suggestions for further research.

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1. THEORETICAL BACKGROUND

After the abandonment of the gold standard in 1976, the mechanism of currency exchange fluctuations became more or less of a mystery. Exchange rates were fluctuating but no one has so far managed to explain their movement with a single, mathematically precise model that would fit the real world data. As Lucio Sarno laconically puts it in his 2005 article: "The international finance profession has not yet been able to produce theories and, therefore, empirical models that allow us to explain the behavior of exchange rates with a reasonable degree of accuracy" (Sarno 2005). The current paper fits into the enourmous strand of currency exchange literature that attempts to empirically explain, if only to the slightest degree, the exchange rate behavior. Particularly, it examines the impact of the current Western economic sanctions on the Russian rouble's fluctuation over 2014-2017 alongside a combination of economic fundamentals relative to uncovered interest parity and purchasing power parity theories. Following previous research, a multivariate VAR model comprising conventional macroeconomic variables and an aggregate index measuring the intensity of economic sanctions is tested. Next follows an overview of several currency exchange theories that enter the model through macroeconomic variables.

1.1. Uncovered interest parity and the rejection of such

According to the efficient markets hypothesis formulated by Eugene Fama (1970), in a speculative market where asset prices fully reflect all information available to market participants and the market participants are risk-neutral and rational, it is impossible to earn excess returns to speculation because the expected foreign exchange gain from holding one currency rather than another must be offset by the opportunity cost of choosing to hold funds in this currency rather than the other. This condition is termed the uncovered interest rate parity (UIP):

$$\Delta_k s_{t+k}^e = i_{t,k} - i_{t,k}^* \qquad (1.1)$$

where s_t denotes the logarithm of the spot exchange rate (domestic price of the foreign currency) at time t, $i_{t,k}$ and $i_{t,k}^*$ are the nominal interest rates available on comparable domestic and foreign assets, respectively, for k holding periods, $\Delta_k s_{t+k} \equiv s_{t+k} - s_t$, and the superscript e denotes the market expectation based on information at time t. Under the assumption that the forward market correctly reflects future spot exchange rates, we can substitute s_{t+k} with f_t^k, the logarithm of the k-period forward rate, to arrive at the covered interest parity (CIP):

$$f_t^k - s_t = i_{t,k} - i_{t,k}^* \tag{1.2}$$

Replacing the interest rate differential $i_{t,k} - i_{t,k}^*$ with the forward premium (or discount) $f_t^k - s_t$, numerous scholars have tested UIP by estimating a regression of the form

$$\Delta s_{t+1} = \alpha + \beta(f_t^1 - s_t) + u_{t+1} \tag{1.3}$$

where u_{t+1} denotes an error term that is uncorrelated with the information available at time t. Under *UIP*, the intercept α must equal zero and the slope parameter β must equal unity. (Fama 1984). In practice, however, successive empirical research has routinely rejected UIP. For instance, estimations of equation (1.3) for various currencies and time periods by Hodrick (1987), Froot and Thaler (1990) and Engel (1996) find that the slope parameter β is significantly less than one, and negative.

The negativity of the estimated UIP slope implies that the larger is the forward premium, the less the home currency is predicted to depreciate. Equivalently, the more domestic interest rates exceed foreign interest rates, the more the domestic currency tends to appreciate over the holding period. In both cases, a trader may realize excess returns to speculation by borrowing funds in the foreign currency, converting them to the domestic currency, investing in the domestic deposit and converting it back to the foreign currency at the end of the holding period.

A rejection of the UIP condition can therefore be attributed to the failure of one or both joint hypothesis assumptions – non-rationality of market participants and/or risk aversion of agents who require higher returns for investing in foreign relative to domestic assets (Sarno and Taylor 2003). Briefly, the failure of these assumptions modifies the UIP condition in the following manner:

(i) The failure of rational expectations: market participants do not use available information efficiently enough to form correct expectations. In this case, the UIP condition is extended with the forecast error term η_{t+1} , which is correlated with the information available at time *t*:

$$\Delta s_{t+1} = \beta_0 + \beta_1(f_t^1 - s_t) + \eta_{t+1} + u_{t+1} \tag{1.4}$$

where u_{t+1} is a white-noise disturbance term. The literature identifies at least four possible causes of irrational expectations: learning about regime shifts (Lewis 1989a); rational bubbles (Lewis 1989b); the "peso problem" (Rogoff 1979), or inefficient information processing (Bilson 1981).

(ii) Departure from risk neutrality: risk-averse market participants demand a higher rate of return than just the interest differential to compensate for the risk of holding foreign currency. Their expected gain will hence consist of the expected rate of depreciation of the domestic currency $f_t^1 - s_t$ and the risk premium p_t:

$$\Delta^e s_{t+1} = \beta_1(f_t^1 - s_t) - p_t + u_{t+1} \tag{1.5}$$

where u_{t+1} is a white-noise error term. An earlier literature attempted to understand foreign exchange risk with the help of the static capital asset pricing model (CAPM). Most of these studies provide evidence that the risk aversion parameter is very large but oftentimes not significantly different from zero (e.g., Adler and Dumas 1983; Frankel 1982; Giovannini and Jorion 1989; Engel 1992). Subsequent empirical works have leant on this research and analyzed the significance of the risk premium in a dynamic general equilibrium context, primarily in the dynamic, two-country, two-good general equilibrium model of Lucas (1982), whose main inference is the constant relative risk aversion utility function:

$$E_t s_{t+1} - f_t^1 = \phi cov_t(s_{t+1}, c_{t+1}) - \frac{1}{2} var_t(s_{t+1}) + cov_t(s_{t+1}, p_{t+1}) \tag{1.6}$$

where $E_t s_{t+1}$ is the expected future spot exchange rate, f_t^1 is the current domestic one-period forward exchange rate, c_{t+1} is the total consumption of the domestic and foreign good by domestic agents in period $t+1$, ϕ is the relative risk aversion coefficient, p_{t+1} is the next period price of the domestic good and var_t and cov_t denote the conditional variance and covariance based on information at time *t*. These works have generally shown that it is hard to explain deviations from UIP by an appeal to risk premia alone: either ϕ, the relative risk aversion coefficient, must be incredibly large, or the conditional covariance between the spot rate and consumption must be incredibly high (Lewis 1995).

Nevertheless, using more advanced econometric apparatus, recent research has found evidence for the ability of forward rates to forecast future spot rates. Clarida and Taylor (1997) apply a linear VECM representation to the spot dollar exchange rate during the recent floating exchange rate regime to show that sufficient information may be extracted from the term structure of forward rates in order to predict future spot exchange rates. Clarida *et al.* (2003) generalized the linear

VECM to a multivariate Markov-switching model and reported evidence of the non-linearities in the term structure of forward rates and demonstrated that the Markov-switching forecasts are strongly superior to the random walk forecasts at out-of-sample forecasts for up to 52 weeks ahead.

1.2. Purchasing power parity and the real exchange rate

1.2.1. Definition of the purchasing power parity

Usage of the term “purchasing power parity” (PPP) is relatively recent — it was introduced into the economic lexicon by Gustav Cassel 100 years ago (Cassel 1918). The concept itself and discussions relating to the relationship between the exchange rate and prices more generally have, however, a very much longer history in economics, and date back as far as the writings of scholars of the University of Salamanca in the fifteenth and sixteenth centuries (Sarno and Taylor 2003). The purchasing power parity (PPP) hypothesis states that national price levels should be equal when expressed in a common currency. The validity of this simple proposition over the long run has been examined empirically either by testing whether nominal exchange rates and relative prices move together or by testing for the existence of a stable long-run equilibrium level of the real exchange rate (RER). The latter approach is motivated by the fact that the real exchange rate may be defined as the nominal exchange rate adjusted for relative national price levels. More formally, the real exchange rate, q_t, may be expressed in logarithmic form as

$$q_t \equiv s_t - p_t + p_t^* \tag{1.7}$$

where p_t and p_t^* denote the logarithms of the domestic and foreign price levels, respectively. The real exchange rate, q_t thus constitutes a measure of the deviation from PPP and must be stationary for long-run PPP to hold (*Ibid.*).

Long-run PPP has several significant economic implications. In particular, the degree of persistence in the real exchange rate can reveal the principal impulses behind the exchange rate movements. For example, if the real exchange rate is highly persistent or resembles a random walk process, then the shocks are likely to be real-side, principally technology shocks, whereas if it is not very persistent, then the shocks must originate from aggregate demand, for example, through innovations to monetary policy (Rogoff 1996).

Regardless of the great interest in this area of research, the validity of long-run PPP and the properties of PPP deviations remain the subject of an ongoing controversy. Specifically, for developed-country data, earlier cointegration studies generally reported the absence of considerable mean reversion of the real exchange rate for the post-Bretton Woods era (Mark 1990) but found some evidence of reversion toward PPP during the gold standard period (McCloskey and Zecher 1984; Diebold, Husted, and Rush 1991), as well as for the interwar float (Taylor and McMahon 1988), and for the U.S.-Canadian float in the 1950s (McNown and Wallace 1989). Exchange rates of high-inflation countries were found to revert to a long-run equilibrium even after the abolition of the gold standard (Choudhry, McNown, and Wallace 1991).

It should also be noted that, in testing for mean reversion in real exchange rates, most studies in the literature have examined a cointegrating relationship between real exchange rates and official price indices. If, however, one assumes that the real exchange rate adjustment towards the PPP equilibrium is founded on the tenet of arbitrage in international goods markets, the choice of the appropriate price index to be used in implementing PPP becomes increasingly important. Summers and Heston (1991) attempted to make a breakthrough in the literature by constructing internationally comparable price indices taking into account traded goods only, although their datasets have little practical use for time series econometricians, since they are constructed at infrequent and long time intervals. Due to this, economists normally use consumer price indices (CPI) published by official sources when constructing the real exchange rate, despite their limitations for the purpose of testing for the validity of the long-run PPP hypothesis. However, some work on PPP looks at the cost of a basket of goods from the producers' point of view - producer price index - rather than the cost of a basket of goods to consumers. For instance, Lafrance, Osakwe, and Normandin (1998) and Sarno and Chowdhury (2003) provide evidence that PPP is more likely to hold if it is based on the costs of production – primarily unit labor costs – or on indices covering exclusively tradable goods, rather than on consumer price indices from official sources.

1.2.2. Harrod-Balassa-Samuelson effect

One conceptual problem that economists face when studying PPP is that it makes an implicit assumption that the long-run real exchange rate is constant. One potential explanation as to why the equilibrium real exchange rate, at least as measured using relative consumer price indices, may vary is based on the so-called Harrod–Balassa–Samuelson (HBS) effect (Harrod 1933; Balassa 1964; Samuelson 1964). For the sake of illustration, consider a situation where prices are equalized internationally among traded goods through international arbitrage, so that the “law of one price” (LOOP) holds and there are two countries, one of which is growing faster than the other. In the fast-growing economy, productivity growth will tend to be concentrated in the traded goods sector, since the nontraded goods sector will have a higher proportion of services, which are less responsive to productivity innovations. This will result in wage rises for the traded goods sector while the prices of traded goods stay unaffected. Hence the LOOP continues to hold and the nominal exchange rate remains unchanged. However, workers in the nontraded goods sector will demand comparable pay rises for themselves, which will cause an increase in the prices of non-tradables and an overall rise in the consumer price index. Since the LOOP holds among traded goods and, by assumption, the nominal exchange rate has remained unchanged, the rise in the domestic CPI will not be matched by an equivalent upward movement in the nominal exchange rate and the domestic currency will now appear undervalued compared to the real exchange rate estimated using relative CPIs. The crucial assumption of the HBS effect is that faster growing economies will experience consistently higher price levels of domestically produced goods, as measured by consumer price indices - when converted at the prevailing market exchange rates – and hence a fall in the level of the real exchange rate, i.e. a real appreciation of the domestic currency.

For the long-run PPP to hold, a real appreciation of the external value of the currency – i.e. a loss in competitiveness – must have a net long-run deflationary impact on the economy and facilitate mean reversion to the equilibrium real exchange rate. Overall, the empirical evidence on the HBS effect is quite mixed (Sarno and Taylor 2003). Japan is often referred to as a good example of the Harrod-Balassa-Samuelson effect in operation. Taylor *et al.* (2001), for example, studied real exchange rates of four developed industrialized countries against the US dollar from 1973 to 1996 and found that, although there is evidence of nonlinear mean reversion in the US dollar–Japanese yen real exchange rate, the speed of mean reversion for this exchange rate is much higher than for the other countries examined. They suggest that this may be due to HBS effects, given that Japan has been the fastest-growing economy for much of the post-World War II period.

Lothian and Taylor (2008) examine the dollar–sterling real exchange rate in a non-linear setting by allowing for the HBS effects through relative national output per capita differentials. They find significant evidence of an HBS effect and corroborate the argument that including a proxy of the HBS effects into the econometric procedure does indeed improve the speed of mean reversion.

Coakley *et al.* (2005) take a slightly different approach by testing for a *relative* rather than *absolute* long-run PPP. They argue that, even if there were some important real shocks disturbing the equilibrium real exchange rate, relative PPP may still hold in the sense that a difference in aggregate price inflation between two countries will, in the long-run lead to a proportionate rate of depreciation of their nominal exchange rate. Having studied a panel data for a large group of countries in this manner, they are unable to reject long-run relative PPP.

Overall, many economists would argue that HBS effects should be present to a greater or lesser degree in the exchange rate relationship of any two currencies, however, there may be other important real factors explaining a positive correlation between productivity growth and rising prices. Lipsey and Swedenborg (1996), for instance, find strong evidence of the inflationary impact of protectionist measures in the agricultural sector of developed countries (e.g., with tariffs on agricultural imports). They show that if wealthy countries support their native producers by imposing barriers to free trade in protected industries, prices for protected goods will rise making them more expensive in the domestic market than in the international market. The result of this move will be similar to that of the HBS effect - relative price level of the protectionist countries will rise and, *ceteris paribus*, the real value of the national currencies will appreciate.

1.3. Link between exchange rates and oil prices

Commodity prices are generally found to explain much of the real exchange rate fluctuations in commodity-exporting countries (Chen and Rogoff 2003; Cashin *et al.* 2004). Oil prices in particular are considered to affect real exchange rates both in oil-exporting countries, as well as in advanced economies (Ricci *et al.* 2008). Therefore, econometric models of real exchange rate equilibria often include commodity prices among their explanatory variables. Overall, commodity prices are thought to have a positive impact on the RER: a rise in commodity terms of trade[1] causes the exporting country's real exchange rate to appreciate. The currencies of exporting countries that follow this pattern are usually termed "commodity currencies".

Chen and Rogoff (2003) focus on three "commodity currencies" issued by Australia, Canada and New Zealand. They find that commodity prices have a stron impact on their real exchange rates,

[1] Calculated as the ratio of commodity export price to commodity import price, both terms of the ratio being a weighted average price of the commodities specifically exported and imported by each country.

particularly for Australia and New Zealand; the result for Canada due to its more diversified export structure. Cashin *et al.* (2004) examine 58 commodity-exporting countries and find the effect of commodity terms of trade on the real exchange rate in about a third of them. Ricci *et al.* (2008) perform a panel analysis of cointegration relationship between economic fundamentals and the real effective exchange rate (REER)[2] in a sample of 48 countries, developed and developing. Among their explanatory variables, they include commodity terms of trade and determine that the long-run elasticity of the REER towards commodity terms of trade is generally 0.6, which means that a 10% rise in the commodity terms of trade implies a 6% appreciation of the REER in the long run.

Being a key natural resource, oil provides the bulk of the wealth of oil-exporting countries and is normally expected to have a more powerful impact than other commodities on the exchange rates of such countries. Habib and Kalamova (2007), for instance, focus on three oil-exporting countries – Russia, Norway and Saudi Arabia – and study cointegration between the real oil price and the real exchange rate on a single-country basis. Their findings evidence a long-run relationship between the two series only for Russia.

Korhonen and Juurikkala (2007) take a different approach and perform a cross-country study. They find that the price of oil has had a significant, positive effect on the real exchange rates of OPEC and three Commonwealth of Independent States (CIS) countries from 1975 to 2005. Alongside the oil price, they include real GDP per capita as their explanatory variable but fail to evidence a clear effect of GDP on the exchange rate. According to the authors, it may be due to the difficulty to disentangle the separate effects of productivity and oil price, since an increase in the oil price is analagous to the Harrod-Balassa-Samuelson effect discussed above: a country's relative productivity at new relative prices rise in the tradable energy sector, which pushes up wages and prices in other sectors of the economy.

The real appreciation of the currency that follows an increase in the oil price may cause a special case of the HBS phenomenon referred to as the "Dutch disease" (Corden 1984). "Dutch disease" progresses as follows. A rise in the oil price or output generates higher revenues for the oil sector, which in turn leads to higher wages in the sector and rising aggregate demand. As part of the

[2] The real effective exchange rate (REER) is the weighted average value of a country's currency relative to a basket of its major trading partners, adjusted for the effects of inflation. The weights are determined by the ratio of each trade partner in the country's total export and import.

demand is aimed at domestically produced services, service prices rise as well, whereas the prices of oil and manufacturing goods, being determined in the international market, are not affected. This triggers a real exchange rate appreciation through the rise of the country's relative price level. Next, assuming that the labor force is mobile, all other sectors of the economy are forced to raise their wages as well. Since the tradable sectors, whose prices are determined abroad, cannot offset the pay rise by raising their prices, their profits drop. This eventually induces a decline in manufacturing output and employment.

1.4. Empirical research of sanctions' economic effect

Sanctions as an instrument of political coersion have been rising in popularity among political leaders and economists alike since World War II. Empirical research of economic sanctions have generally been focused on the cost of sanctions to the sender(s), i.e. the imposing country(s), the target, or the target's neighbors (the so-called "bystanders"). By design, economic sanctions restrict the ability of the targeted country to engage in international commerce reducing its exposure to foreign trade or finance (Lektzian and Patterson 2015). The resulting effects on the welfare of the target country have been traced through various channels of macroeconomic and fiscal indicators.

The empirical evidence on the effects of economic sanctions is mixed. Trade restrictions, for example, can harm both the target and the sanctioning country by cutting established trading routes and increasing transportation costs. Non-sanctioning countries with strong economic ties with the target can either be hit through reduced perspectives for growth or favored through increased trade with the rest of the world (Caruso 2003). Using a gravity regression approach, Caruso (*Ibid.*) reports that, over the period 1960-2000, various extensive unilateral sanctions imposed by the Unites States had a large negative impact on the bilateral trade between 49 target countries and the other G-7 countries, while limited and moderate US sanctions, in contrast, induced a minor positive growth in other G-7 countries' bilateral trade with the targets.

The impact of economic sanctions on the target states' GDP growth is another popular line of research. Neuenkirch and Neumeier (2015), for instance, assess the impact of economic sanctions imposed by the UN and the USA between 1976-2012. Using panel data estimation techniques and a sample that contains 68 countries, they find that UN sanctions have a relatively large and economically significant effect in reducing target state's real per capita GDP growth rate by 2.3-3.5 percentage points over a period of 10 years, with the most severe sanctions (that is, embargoes affecting almost all economic activity) yielding a more than 5 percentage point decrease in GDP. In contrast, the effect of the US sanctions is much smaller, accounting for an average of 0.5-0.9 percentage point decrease in the GDP growth rate over a period of 7 years.

Yahia and Saleh (2008) analyze the link between economic sanctions, oil price fluctuations, and the labor market in Libya. Using a multiple regression model and the Johansen cointegration technique, authors find that the periods of sharp decline in oil prices (1983-1998) and economic sanctions (1990-2003) had a negative effect on the movement of the skilled non-Libyan labor.

Hoffmann and Neuenkirch (2015) investigate the impact of sanctions on Russian stock returns during the period from November 21, 2013 to September 29, 2014. They utilize a newly constructed news-based indicator of the development of conflict in the Ukraine and find that the degree of (de-)escalation accounts for a total variation of 6.5 percentage points in the Russian stock market and that the intensification of the conflict reduces Russian stock returns. The intuition behind this is simple: the risk of war increases the risk of assets related to the parties involved in the conflict and induces sales of risky assets by investors.

Gurvich and Prilepskiy (2016) execute a scenario analysis of the impact of financial sanctions on the Russian economy. Based on the actual data available at the second quarter of 2015, they simulate time series of the capital inflow of Russian companies and overall macroeconomic and fiscal indicators up until the end of 2017. In their setup, financial sanctions are threefold and include limits on extensions of credit to energy and defence technology sector companies, a US ban on foreign exchange transactions with entities whose owners were subjected to individual sanctions, and the so-called “soft” sanctions, i.e. tightened regulatory control over bank payments that slows down their execution and increases transaction costs.

Authors estimate the overall negative effect of sanctions on gross capital inflow over 2014-2017 at $280bn and note that the net effect will be 40% lower ($160-170bn) due to Russian companies’ active self-adjustment, evidenced by their debt repayments using accumulated foreign assets and an overall decrease in gross capital outflow (*Ibid.*). Authors conclude that the drop in oil prices has a much larger effect on the Russian economy: falling oil prices lead to GDP losses of 8.5 p.p. cumulatively from 2014 through 2017, while financial sanctions account for only 2.4% losses of pre-crisis GDP, explaining that with a comparatively slow reaction of the economy to the cheapening oil via decreased production of non-tradable goods resulting from falling domestic demand.

Dreger *et al.* (2016) utilize a cointegrated vector autoregression (VAR) technique to examine the effects of the sanctions related to Russia and Ukraine and the fall in the oil price on the daily nominal exchange rate of the rouble throughout 2014. They supply their VAR model with data on interest rates, actual and unanticipated santions and find that the exchange rate is affected more by oil prices than by the economic sanctions. Next, they extend the model by a multivariate Generalized Autoregressive Conditional Heteroscedasticity (GARCH) process and establish that unanticipated sanctions matter for the conditional volatility of the variables envolved. Kholodilin and Netsunajev (2017) employ the sanctions index developed in Dreger *et al.* (2016) and investigate bilateral effects of sanctions on the real GDP growth of Russia and the euro area as well as on the real effective exchange rates of the Russian rouble and the euro using structural VAR. They show that the effect of sanctions is assymetric: a negative effect on the euro is estimated to be larger compared to the rouble; at the same time, the effect on the respective euro area GDP growth is negligible, while the Russian GDP growth decreased by almost 2%.

2. METHODOLOGY

The current paper utilizes a structural Vector Autoregression (SVAR) model, a multivariate version of univariate autoregressive (AR) time series modeling, to assess the effects of economic sanctions against Russia for the Russian currency, compared to a number of macroeconomic variables. The estimable VAR system includes five dependent variables: a sanctions index measuring the intensiveness of Western sanctions introduced in Dreger *et al.* (2016), real GDP growth differential between Russia and the euro area, the short-term interest rate differential between the two regions, the price of oil, and the bilateral Rouble/Euro exchange rate.

In their basic form, VARs represent summaries of the correlation structure embedded in observational data, but if some sort of structure is induced into the interdependancies of individual regressors, it becomes feasible to make inferences about the causality between the VAR components. The following chapter introduces the specificities of how a structural VAR is set up and the ways to estimate and interpret it.

2.1. Data description

The following data have been used in the vector autoregression analysis. All of the variables have been checked for seasonality using Gretl's *TRAMO Analysis* add-on and seasonally adjusted, where necessary. The data are quarterly and run from 1999Q1 through to 2017Q3 (75 observations).

- $\boldsymbol{s_t}$ – sanctions index, adopted from the Dreger *et al.* (2016), updated with new data and slightly overhauled. Before 2014Q1, the index takes the value of zero and thereafter represents a gradual intensification of the economic sanctions against Russia between 2014 and 2017 with a single drop in the composite index taking place in 2017Q2, when some of the sanctions expired. Following Dreger *et al.'s* example (*Ibid.*), the index assigns integer power weights from 1 to 3 to the measures targeting: 1) individuals, 2) businesses, and 3) entire industries, respectively, as well as trade weights equal to the share of imposing countries in the total external trade (imports plus exports) of the Russian Federation between 2009 and 2013. The maximum trade weight assigned was 0.443 and belonged to the European Union, the largest trading partner of Russia before and after the onset of sanctions. The author's amendments to the original index design consisted of adding sanctions data and/or corresponding power and trade weights that were not reflected in the initial dataset, extending the series with new data available after 2015Q1 and, finally, assigning time weights - the longer a sanction was in action throughout a certain quarter, the larger time weight between 0 and 1 it is assigned. Appendix 1 provides a side-by-side comparison of the original and modified sanctions data. The ultimate composite sanction index representing sanctions against Russia is defined here as a cumulative sum of individual sanction entries, S_t:

$$S_t = \sum_{\tau}^{t} \sum_{i=1}^{I} \sum_{j=1}^{J} w_i w_j w_\tau s_{ij}$$

where,

w_i– is the power weight of sanction *i*,
w_j – is the trade weight of country *j*,
w_τ – is the time weight of sanction *i* in quarter *τ*,
s_{ij} – is an indicator function of individual sanction i taking values 1 or 0, depending on whether the sanction *i* is in action in period *τ* or not.

- $\boldsymbol{\Delta p_t^{oil}}$ – log difference of the real Urals oil price, the Russian reference oil brand. The average quarterly spot prices of oil have been downloaded from Eikon's Datastream terminal and adjusted by the Russian CPI (2010 = 100) taken from the OECD database. Given the high share of oil and gas in the total exports of Russia, this variable accounts for a downward pressure on Russia's foreign trade following a drastic drop in oil prices that began in summer 2014,
- $\boldsymbol{\Delta y_t^{RU} - \Delta y_t^{EA}}$ – differential of real GDP growth rates between Russia and euro area. Russian GDP data are obtained from the Russian Federation Federal State Statistics Service (Rosstat), data for the euro zone are obtained from Eurostat and cover 19 euro area countries. This variable enters the model as a proxy for productivity differential between the two regions to test for the presence of HBS effects,
- $\boldsymbol{i_t^{RU} - i_t^{EA}}$ – real short-term interest rate differential, calculated as a difference between the real Russian 3-month interbank rate and the real 3-month EURIBOR. Interest rate data are taken from the OECD database and converted to real values using the Fisher equation:

$$real\ interest\ rate = \frac{(1 + nominal\ interest\ rate_t)}{(1 + expected\ inflation\ rate_{t+1})} - 1$$

whereas inflation expectations were obtained from the European Central Bank and Russian Central Bank archives. This variable represents the UIP condition discussed above, which states that the more domestic interest rates exceed foreign interest rates, the more domestic currency should depreciate, provided that market agents are risk-neutral and endowed with rational expectations,

- $\boldsymbol{e_t^{RUB}}$ – log of the real bilateral RUB/EUR exchange rate, calculated as:

  $RER = Nominal\ Exchange\ Rate \times EU19\ 2010\ CPI/Russian\ 2010\ CPI$

  Nominal exchange rates are obtained from the European Central Bank and represent average quarterly RUB/EUR reference prices. CPI data are taken from OECD. An increase in the level of $\boldsymbol{e_t^{RUB}}$ represents a real depreciation of the rouble against the euro.

Figure 1 presents the individual time series plots and Table 1 summarizes the descriptive statistics of the variables under inspection.

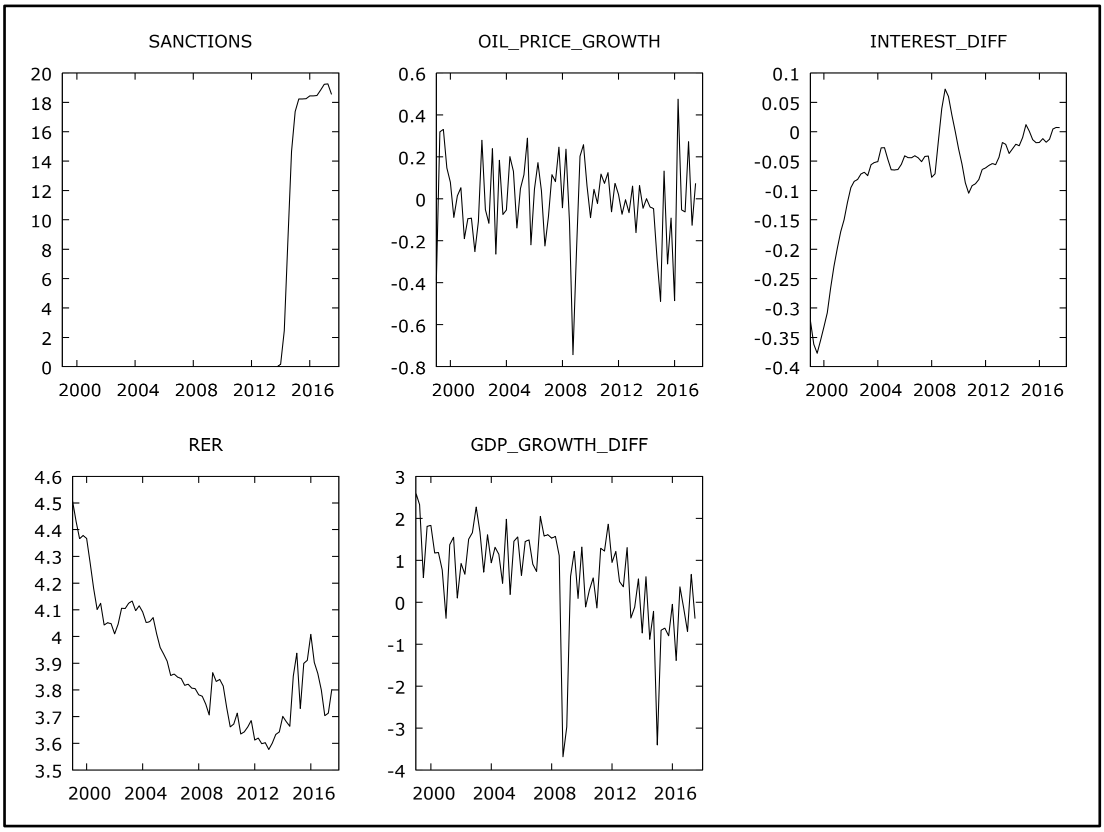

Figure 1. Time-series plots of variables under inspection
Source: OECD, Eurostat, ECB, RCB, Rosstat, Thomson Reuters Eikon

Table 1. Descriptive statistics of the variables under inspection

Variable	Mean	Median	Minimum	Maximum	St.Dev.
Sanctions	3.056	0	0	19.253	6.733
oil_price_growth	-0.009	-0.004	-0.743	0.475	0.204
GDP_differential	0.628	0.767	-3.684	2.606	1.188
interest_differential	-0.074	-0.051	-0.377	0.073	0.097
RER	3.896	3.852	3.577	4.507	0.222

Source: author's calculations

2.2. Structural vector autoregressive model

Vector autoregressive models were popularized in econometrics by Sims (1980) as a more flexible form of univariate AR models. Unlike the latter, VARs allow a variable to depend on more than just its own lags or combinations of white noise terms. The structural VAR takes the form of

$$\mathrm{A}y_t = \Phi_0 + \Phi_1 y_{t-1} + \Phi_2 y_{t-2} + \cdots + \Phi_p y_{t-p} + e_t \tag{2.1}$$

where,

$y_t -$ is a vector of observable variables, equal to $(s_t, p_t^{oil}, \Delta y_t^{RU} - \Delta y_t^{EU}, i_t^{RU} - i_t^{EU}, e_t^{RUB})'$,
$p -$ is equal to 4, based on the lag length selection test results presented below,
$\mathrm{A} -$ is an invertible $(K \times K)$ matrix containing the instantaneous relations between the LHS variables,
$\Phi_0 -$ is a $(K \times 1)$ constant term,
$(\Phi_1 - \Phi_p) -$ is a $(K \times K)$ coefficient matrix,
$e_t -$ is a K-dimensional serially uncorrelated vector of structural residuals with mean zero and identity covariance matrix Σ_e.

Premultiplying (2.1) with A^{-1}, the testable reduced-form model appears as

$$y_t = \mathrm{A}_0 + \mathrm{A}_1 y_{t-1} + \mathrm{A}_2 y_{t-2} + \cdots + \mathrm{A}_p y_{t-p} + u_t \tag{2.2}$$

where,

$u_t -$ is a K-dimensional serially uncorrelated vector of zero mean reduced-form residuals with a non-singular covariance matrix Σ_u,
$\mathrm{A}_i = \Phi_i \times \mathrm{A}^{-1}$
$u_t = \mathrm{e}_t \times \mathrm{A}^{-1}$

The estimation of (2.2) is done according to these steps:

1) A_i's and Σ_u are estimated using OLS,

2) Φ_i's and e_t are obtained using the Cholesky decompostion of Σ_u into $A^{-1}\,A^{-1\prime}$ and assumption that A^{-1} is lower triangular: $\begin{pmatrix} a_{11} & \cdots & 0 \\ \vdots & \ddots & \vdots \\ a_{K1} & \cdots & a_{Kp} \end{pmatrix}$. Since the recursive method of identification implies that equations in the system are ordered in terms of their relative exogeneity, a restricting assumption must be made as to which variables can instantaneously impact the other variables in the system and which cannot.

It should be emphasized that there are other methods used to identify SVARs apart from the recursive scheme suggested by Sims: sign restrictions, zero long-run restricions, theory- and heteroskedasticity based restrictions among many others (Kilian 2011). However, none of them escapes some sort of criticism. Sims (1980) championed the idea that unrestricted vector autoregressions altogether provide a better understanding of macroeconomic relationships than structural models because structural models use "incredible" identifying restrictions. Cooley and Dwyer (1998) argue that structural VARs (SVARs) "are certainly not invariant to the identifying assumptions and may not be reliable as vehicles for identifying the relative importance of shocks." Furthermore, it has been shown in a number of articles that SVARs generated less accurate out-of-sample forecasts than their unrestricted counterparts (Sims 1980; McNees 1986). This may arise from the unfortunate fact that identifying constraints are usually based on subjective (non data-based) considerations. Recent studies, however, have attempted to find a more objective way to identify simultaneous vector autoregressive equations. Lanne *et al.* (2017), for instance, have shown that the SVAR model can be uniquely identified by statistical properties of the data. The complexity of such approach, though, would warrant a more extensive research than the current one.

Often, financial theory will have little to say on what is an appropriate lag length *p* for a VAR, or, put differently, how long changes in the variables should take to work through the system. In such instances, there are broadly two approaches that could be taken to arrive at the optimal lag length: cross-equation restrictions and information criteria (Brooks 2008). Cross-equation restrictions imply that variables in each equation may have the same or different number of lagged values. The former is more in the spirit of VAR estimation proposed by Sims (1980), since the latter, a VAR with different lag lengths for each equation, could be viewed as a restricted VAR where coefficients on certain lags of some of the variables, in specific equations, are restricted to zero.

Nevertheless, a more straightforward and preferable approach to lag length restricion is to specify the same number of lags in each equation and to determine the model order using a likelihood ratio test or information criteria (Brooks 2008). A likelihood ratio test performs a pairwise comparison between the unrestricted model, which contains p lagged values of observable variables (e.g., 4 lags for quarterly data), and the restricted model containing $p-1$ lags. An alternative approach to selecting the optimal VAR lag length would be to use multivariate information criteria, which do not assume the normality of error distributions (*Ibid.*) Similar to the likelihood ratio test, the estimation of the information criteria for the unrestricted and restricted models is conducted for 0, 1, ..., p lags (up to some pre-defined maximum p). The appropriate number of lags would be that number minimising the value of the given criterion.

Table 2 presents the values of the Akaike (AIC), Schwarz (BIC) and Hannan–Quinn (HQC) information criteria, as well as the likelihood ratio test statistic, computed in Gretl from VARs of order 1 to the chosen maximum of 4 based on the quarterly data compiled for the current paper. The output of the optimal lag length test contains mixed results: Akaike and Hannan-Quinn criteria suggest that the optimal lag length is 4, the likelihood ratio test selects a VAR of order 1, while the Schwarz criterion chooses a VAR of order 2. Since VAR(4) is consistent with the previous studies by Dreger *et al.* (2016) and Kholodilin and Netsunajev (2017), the choice of the model order is made in favor of lag length 4.

Table 2. VAR lag length selection results

Lag	**loglik**	**p(LR)**	**AIC**	**BIC**	**HQC**
1	144.103		-3.214	-2.258	-2.834
2	206.624	0.000	-4.271	-2.518*	-3.574
3	245.083	0.000	-4.650	-2.101	-3.636
4	294.808	0.000	-5.347*	-2.001	-4.016*

Note: maximum lag order 4, the asterisks indicate the best values of the information criteria
Source: author’s calculations

2.3. Interpretation of the estimated model

One fundamental weakness of the VAR approach to modeling is that it's a-theoretical nature (akin to ARMA models) and the large number of parameters involved make the estimated models difficult to interpet. In particular, some lagged variables may have coefficients which change sign across the lags and this, complicated with the interconnectivity of the equations, could render it hard to see what effect a given change in a variable would have upon the future values of the variables in the system. Two sets of statistics are usually constructed for an estimated SVAR model: impulse responses and variance decompositions. This sub-section summarizes the technicalities of each of them.

2.3.1. Impulse response functions

Impulse response functions (IRFs) are needed to identify the sign and duration of the response of one variable to an impulse in another variable in a system that contains a number of further variables as well. In particular, once the estimation step is completed and the A^{-1} matrix has been obtained, econometric software constructs a structural vector moving average representation of the model:

$$y_t = \mathrm{A}^{-1}e_t + \Theta_1\mathrm{A}^{-1}e_{t-1} + \cdots \tag{2.3}$$

where $\{\Theta_n\}_{i,j} = \frac{\partial y_{i,t+n}}{\partial e_{j,t}}$, is the impulse response function – the rate of change in variable y_i at time $t + n$ to a one-time unit shock in variable y_j at time t, with all other variables dated t held constant. Since the variance of structural shocks e_t is one ($E[e_t {e'}_t] = I$), a unit shock is just a shock of one standard deviation. Since the exchange rate variable is the logarithm of the actual exchange rate of roubles per euro, the unit of measurement of the IRFs plotting the response of the real exchange rate is equal to

$$\Theta = \frac{\partial y}{\partial e} = \frac{\partial \ln(Y)}{\partial e} = \frac{\partial Y}{Y\partial e} \tag{2.4}$$

Since the shock size is equal to 1, $\partial e = 1$. This simplifies the above to

$$\Theta = \frac{\partial Y}{Y} = \frac{(Y_{t+1} - Y_t)}{Y_t} \tag{2.5}$$

IRFs for the variable e_t^{RUB}, therefore, represent percentage changes of the real exchange rate between time t and $t + 1$, divided by 100.

It is important to note that the unit shock is applied exclusively to the first equation while keeping the error terms of all other equations unchanged. This, however, is unrealistic since the reduced-form residuals u_{it} can contain the information about the instantenous relationships between the variables in the system and thus be contemporaneously correlated. The more highly correlated are the error terms from an estimated model, the more the variable ordering becomes important.

2.3.2. Variance decompositions

Forecast error variance decompositions (FEVDs) assess the proportion of the movements in the dependent variables that are due to their own pure shocks versus shocks to the other variables. They show how much of the h-step-ahead forecast error variance of a given variable is explained by innovations to each explanatory variable for $h = 1, 2, \ldots$ In practice it is often observed that own series shocks explain the largest portion of the forecast error variance of the series in a VAR (Brooks 2008).

Once the residual vector u_t from the reduced-form equation (2.2) is decomposed into simultaneously uncorrelated innovations e_t, the h-step forecast error variance can be written as

$$\Sigma_h = \sum_{i=0}^{h-1} \Theta_i \Theta_i' \tag{2.6}$$

where $\Theta_0 = \mathrm{A}^{-1}$ and $\Theta_i = \Phi_i \times \mathrm{A}^{-1}$ (Lütkepohl 2010). The forecast error variance of the kth element of the forecast error vector is then seen to be

$$\sigma_k^2(h) = \sum_{j=0}^{h-1} (\theta_{k1,j}^2 + \cdots + \theta_{kK,j}^2) \tag{2.7}$$

where the term $(\theta_{k1,j}^2 + \cdots + \theta_{kK,j}^2)$ represents the contribution of the jth innovation to the h-step forecast error variance of variable k (*Ibid.*).

3. EMPIRICAL RESULTS

Before proceeding to the OLS estimation of the structural VAR in the reduced form of (2.2), stationarity has been checked using an augmented Dickey Fuller test – all variables appear to be stationary at the 5-percent significance level, except for the real exchange rate, which is stationary at the 10-percent significance level in a model that includes a constant (see Appendix 2).

The initial ordering of the variables in vector y_t is chosen so that to match the economic theory and empirical literature on exchange rate fluctuations presented in the first and second chapter of this paper: the first variable in the order is the sanctions index as the most exogenous and atheoretical one, whose structural shock on the real exchange rate this study seeks to identify, next follows the oil price, which is also quite exogenous in the current setup, since it should abide to the world's demand and supply of energy resources, followed by the productivity differential, which in case of Russia can be particularly sensible towards fluctuating oil prices; the last two variables are the interest rate differential and the real exchange rate, whereas, according to the UIP theory, interest rates take presedence of exchange rates. The latter are, based on theory and empirical evidence, most endogenous in the given setting and, accordingly, can be contemporaneously affected by all of the preceding variables. Consequently, the initial ordering of the endogenous variables will be: sanctions, oil price growth, GDP growth differential, interest rate differential and the real exchange rate.

As has been noted before, a disadvantage of using the Cholesky decomposition is that the results of IRFs and variance decompositions can be sensitive to changes in the ordering. Hence, to check the robustness of the proposed model, a sensitivity analysis is performed in the last sub-section of this chapter, where an alternative ordering is employed.

3.1. Impulse-response analysis and variance decomposition of the real exchange rate – the entire sample

In order to answer the first and second research questions, two objects of interest are obtained from the estimated SVAR model: impulse response functions and variance decompositions. First, impacts on the real exchange rate of unit shocks in the other four endogenous variables – sanctions, oil price growth, GDP growth differential and interest rate differential – are tested. This is done by calculating the impulse response functions that trace the reaction of the target variable to a shock on another variable in the system, and how it evolves after the shock. The sign, magnitude and persistence of shocks is evaluated using visual analysis. The results of the impulse response functions for the four shock variables are presented in Figure 2. The IRFs depict the rate of change in the level of the real exchange rate at time $t + h$, for $h = (0 - 20)$, caused by a one standard deviation increase in the shock variable at time t, with 68% confidence intervals built from 2000 bootstrap replications.

In response to the sanctions shock, the real exchange rate shows a statistically significant immediate decrease of about 2.2%, signalling the real appreciation of the rouble. This finding is supported by the recent evidence provided in Tyll *et al.* (2018), who found that since the introduction of sanctions, the exchange rate between rouble and USD has been actively managed by the Russian Central Bank in efforts to tie the exchange rate of rouble to the oil prices. Considering that the escalation of sanctions has been accompanied by the decline in oil prices, it is natural to assume that an unexpected increase in sanctions has met with the Russian Central Bank's immediate purchases of the domestic currency in the FX market.

The sign of the effect, however, is quickly shifted into the positive area in the next period following the shock, when the rouble depreciates 0.5%. Statistically significant depreciation lasts for almost 11 periods after the innovation in sanctions, with the peak of 1.9% in the seventh quarter after the shock. The depreciation of the rouble in this timeframe is supported by the economic theory since the sanctions represent foreign trade distortions (Caruso 2003) and a drop in the gross capital inflow (Gurvich and Prilepskiy 2016).

The IRF then continues its fall in the confidence interval covering both the positive and negative sides of the value range, crossing the zero line again in the 13th period after the shock, indicating a shift to the real appreciation of the rouble. Statistically significant appreciation lasts from the 18th through the 20th period following the sanctions shock. Overall, these results are consistent with the findings of Kholodilin and Netsunajev (2017), who demonstrate an appreciation of the real effective exchange rate of rouble at the very beginning of the response horizon of almost 7% in response to a two unit shock in the sanctions index, which then gives way to the real depreciation

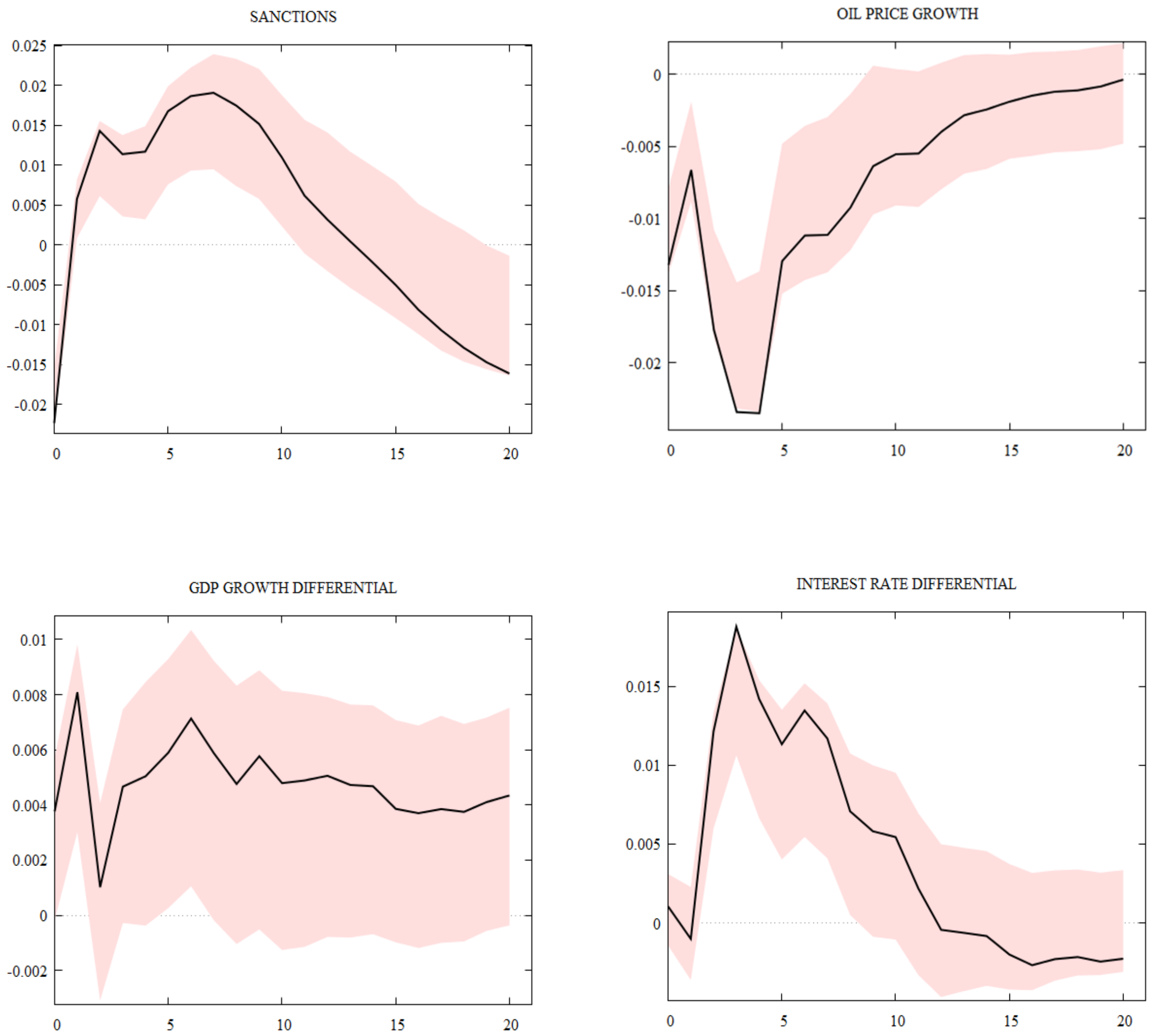

Figure 2. Response of the real exchange rate to a unit shock in the endogenous variables, 1999Q1-2017Q3
Source: author’s calculations using Gretl
Note: X-axes denote quarters following the shock. The pink filled curves indicate 68% confidence intervals based on 2000 bootstrap replications.

of the rouble for five consecutive periods, followed by 15 periods of real appreciation. However, while the sanctions shock in Kholodilin and Netsunajev (*Ibid.*) completely dies out after 20 quarters, the IRF of the real exchange rate from the sanctions shock calculated in this paper does not converge to zero at the end of the horizon period. This discrepancy could be caused by the differences in the identification approaches: the current paper utilizes a simple short-run recursive identification scheme, while Kholodilin and Netsunajev (*Ibid.*) resort to a set of narrative sign restrictions which allow them to mechanically disentangle the sanctions shock from the shock in oil prices and limit the effect that other endogenous variables can have on the evolution of the sanctions index. In this light, the descending negative IRF of the real exchange rate at the end of the 20-period horizon towards an unexpected increase in sanctions, as shown in Figure 2, could be interpreted in several ways. It can indicate an unstable and hardly predictable relationship between the sanctions shock and the long-run equilibrium of the real exchange rate or it could also mean that the effect of the sanctions shock on the real exchange rate is extremely long-lived. To come to conclusions about the longer-term effect of sanctions on the real exchange rate, further research is warranted, possibly in a vector error correction setting.

Next, the effect of unit shocks in the other macroeconomic variables on the evolution of the real exchange rate is studied. The effect of the oil price growth is statistically significant and negative from the beginning of the response horizon (−1.3%) up until the ninth quarter after the shock, with the fastest appreciation of the rouble taking place in the fourth quarter after the shock (−2.3%). From there on, the confidence intervals include zero, but still with a much larger probability mass on the negative side. At the end of the 20th period following the unexpected increase in the oil prices, the shock subsides and the real exchange rate reverts back to its previous mean level. This behaviour is in line with the empirical literature discussed above (Chen and Rogoff 2003; Habib and Kalamova 2007; Korhonen and Juurikkala 2007): rising oil prices lead to the real appreciation of the currency of an oil exporting country.

The response of the real exchange rate to a one standard deviation shock in the GDP growth differential is positive throughout the entire horizon, however, taking into account confidence intervals, it is significant only in the first, sixth and ninth quarters following the shock. What is noteworthy here is that the sign of the IRF is in drastic contrast with the Harrod-Balassa-Samuelson theory presented above: a sudden widening of the GDP growth differential in favor of the domestic country leads to a persistent real depreciation of its currency. This may either be the result of a misspecification or present some new evidence on the HBS effects in Russia. On the one hand, this is reminiscent of the results attained by Korhonen and Juurikkala (2007). In their study, while confirming the significance of the oil price in the exchange rate formation, they failed to ascertain a statistically significant link between the real exchange rates and the GDP per capita. According to the authors, it may be due to the difficulty to disentangle the separate effects of productivity and oil price, since an increase in the oil price is analagous to the Harrod-Balassa-Samuelson effect: a country's relative productivity at new relative prices rises in the tradable energy sector, which pushes up wages and prices in other sectors of the economy. On the other hand, empirical literature does contain examples of the reverse HBS effects. Benigno and Thoenissen (2002), for instance, use a dynamic general equilibrium model that includes home bias for the UK economy to show that overall domestic productivity improvement lowers the price of home goods, which translates into the real depreciation of the real exchange rate when home goods account for a bigger share of the domestic consumption basket.

Lastly, the IRF for the interest rate differential shock exhibits some interesting results. After the initial two insignificant responses of 0.1% in both directions, the real exchange rate depreciates for nine consecutive quarters, with the fastest depreciation of 1.9% taking place in the third period after the shock. This is consistent with the UIP condition discussed above: a widening interest rate differential causes the real exchange rate to depreciate, which diminishes excess returns in the interest rate arbitrage. During the second half on the response horizon, the confidence intervals are distributed almost equally between the positive and negative sides, implying that the value of the IRF is likely to be zero. However, ignoring the confidence intervals, it is notable that the real exchange rate appears to stabilize at a new equilibrium level that is 0.2% lower than the previous one.

The impulse-response analysis is complemented with the variance decomposition of the real exchange rate's forecast errors. The estimated variance decompositions at the 20-period horizon are reported in Figure 3. Most of the forecast error variance can be attributed to the own shocks of the real exchange rate.

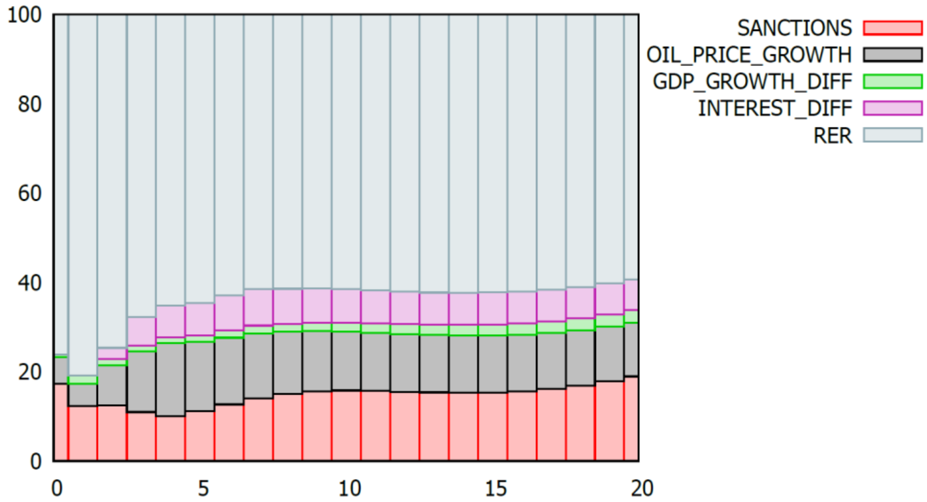

Figure 3. Decomposition of errors for the forecast of the RUB/EUR real exchange rate, 1999Q1-2017Q3
Source: author's calculations using Gretl

The sanctions and oil price growth shocks seem to explain equal portions of the total real exchange rate volatility – around 17% quarterly. Based on this finding and on the IRFs constructed before, the second hypothesis of the current thesis can be denied at this point: the oil price is not more relevant in explaining the variance of the real exchange rate than sanctions.

The validity of two other hypotheses can also be decided upon at this stage. Firstly, on impact, sanctions have had a negative effect on the level of the real exchange rate, contributing to the real appreciation of the currency, which denies the first hypothesis. Secondly, the real exchange rate has indeed been on average positively correlated with the real interest rate differential, which supports the third hypothesis. Regarding the fourth and last hypothesis, considering the IRF of the GDP growth differential shock with a barely detectable statistical significance and its relatively small contribution to the total volatility of the real exchange rate, a sensitivity analysis needs to be performed before jumping to conclusins about the validity of the HBS effects.

3.2. Impulse-response analysis and variance decomposition of the real exchange rate – subsample estimation

In order to answer the third research question about the possibility that the introduction of economic sanctions has changed the role of the macroeconomic variables under inspection in the development of the rouble's bilateral real exchange rate, subsample estimation is performed. The data covering the sanctions period after 2014Q1 is excluded from the sample, leaving 60 observations and four variables in the dataset, keeping the original ordering intact. The results of the re-calculated impulse response functions and variance decompositions are presented in Figures 4a and 4b.

Three observations can be made after comparing Figure 2 and Figure 3 with Figures 4a-4b. First, the IRF of the real exchange rate in response to the oil price growth shock is practically zero at the beginning of the pre-sanctions response horizon and seems to be trailing around zero starting from the 10th quarter after the shock onwards, while the analogous shock evaluated using the entire sample takes twice as long to subside. The fact that the real exchange rate has been more responsive to oil price shocks in Figure 2 than in Figure 4a once again corroborates the assumption made by Tyll *et al.* (2018) that since the introduction of sanctions, the nominal exchange rate of rouble has been actively managed by the Russian Central Bank in efforts to tie the exchange rate to the oil prices.

Second, the magnitude and significance of the GDP growth differential shock is higher in the pre-sanction period (a 0.4% response on impact with the highest consequtive response of 0.8% during the entire sample compared to a 0.75% response on impact and a maximum response of 2.2% thereafter in the subsample).

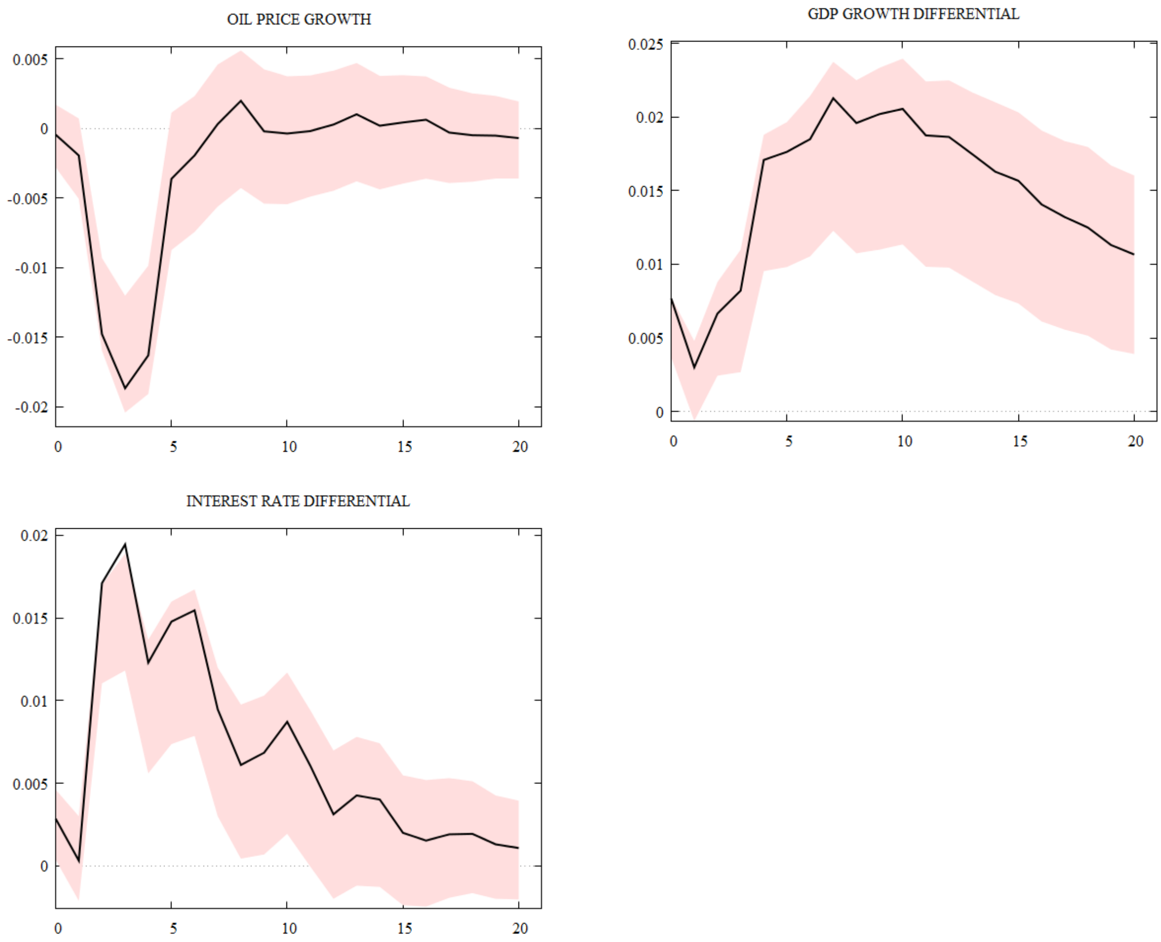

Figure 4a. Response of the real exchange rate to a unit shock in the endogenous variables, 1999Q1-2013Q4
Source: author's calculations using Gretl
Note: X-axes denote quarters following the shock. The pink filled curves indicate 68% confidence intervals based on 2000 bootstrap replications.

Moreover, while the IRF calculated on the pre-sanction subsample appears to be in a downward trend towards zero at the end of the horizon, the analogous IRF from the entire sample does not show signs of dying out. This may indicate that the introduction of sanctions has distorted the relationship between the GDP growth differential and the real exchange rate. This assumption is supported by Figure 4b: the GDP growth differential seems to be the second largest source of the total volatility for the real exchange rate during the pre-sanction period, compared to its small contribution to the fluctuation of the real exchange rate in the entire sample covering the period when the sanctions have been in place (see Figure 3). The fact that the sign of the reaction is positive in both settings implies that moving the GDP growth differential ahead of the sanctions in the variables' ordering will not yield differently signed results, which is important for the following sensitivity analysis.

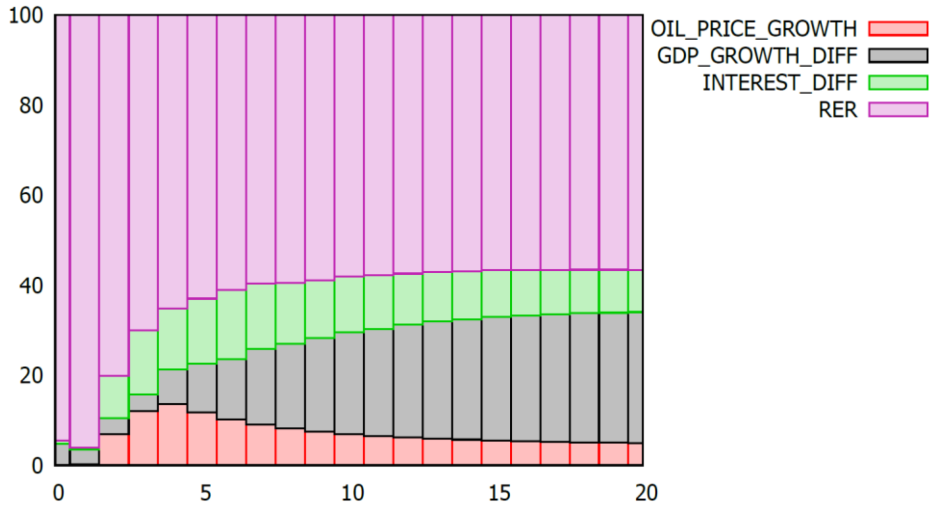

Figure 4b. Decomposition of errors for the forecast of the RUB/EUR real exchange rate, 1999Q1-2013Q4
Source: author's calculations using Gretl

Third, the relationship between the real interest rate differential and the real exchange rate has not significantly changed with the imposition of sanctions. The IRFs are statistically significant a positively signed between the second and ninth period after the shock (in the subsample estimation – between the second and 11th period), signalling that earning excess returns in the interest rate arbitrage has been difficult both before and after the breakpoint in 2014Q1. This result is quite in line with Russia-oriented UIP studies that generally provide evidence supporting the validity of the UIP condition in modern Russia. Recently, Vasilyev *et al.* (2017), for example, in a panel regression analysis of 2001-2014 data, found that the UIP in Russia holds better than in other emerging market economies.

3.3. Sensitivity analysis – alternative ordering of the variables

In order to check the validity of the reverse HBS effect found in sub-sections 3.2 and 3.3., a sensitivity analysis is performed, where the oil price growth and the GDP growth differential variables switch places. The alternative ordering of the endogenous variables will therefore be:

sanctions, GDP growth differential, oil price growth, interest rate differential and the real exchange rate. The reasoning behind the new ordering is supported by the evidence of reverse causality going from macroeconomic variables, such as real GDP, to oil prices, provided by Barsky and Kilian (2004).

Re-estimated IRFs and the variance decomposition of the real exchange rate are reported in Figures 5a and 5b.

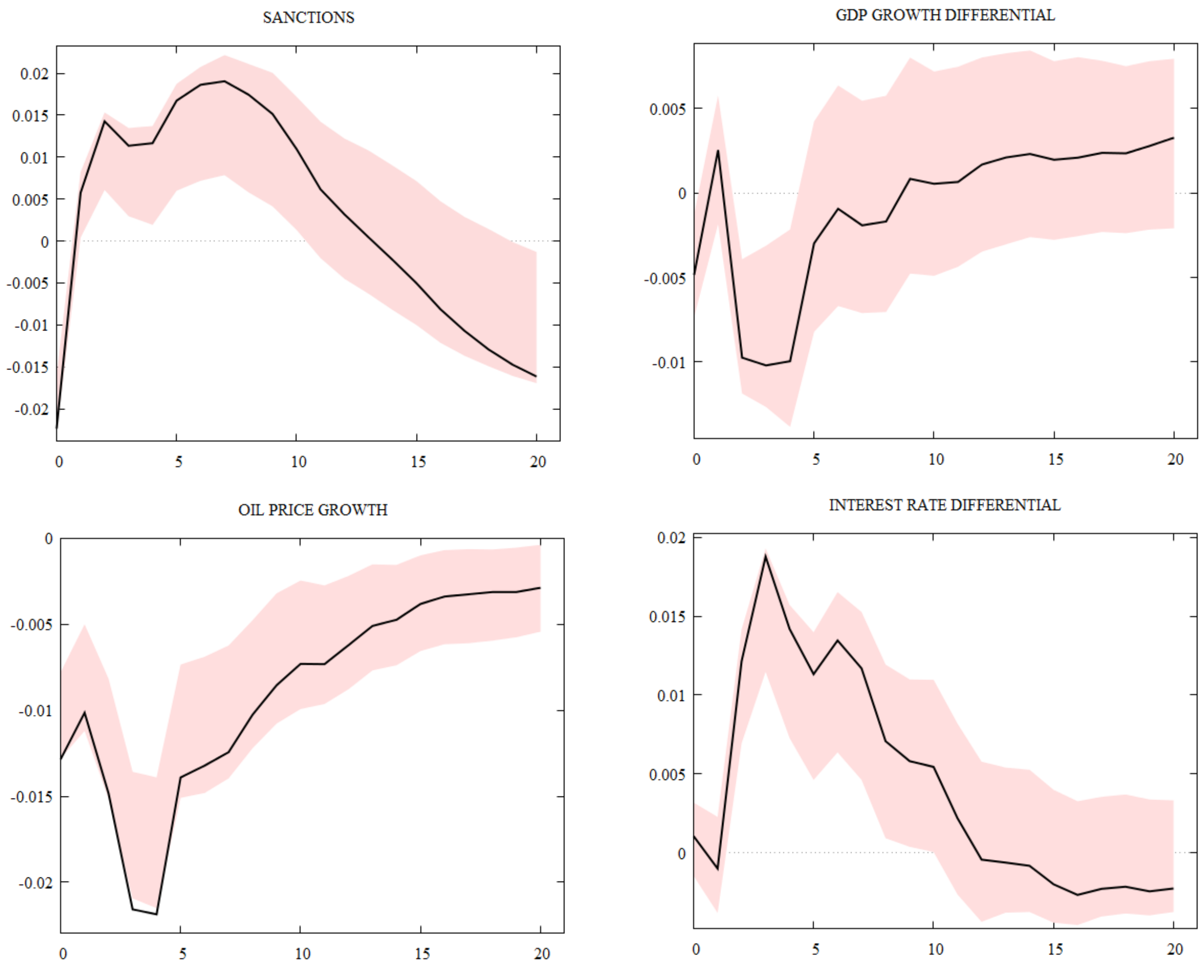

Figure 5a. Response of the real exchange rate to a unit shock in the endogenous variables, 1999Q1-2017Q3
Source: author's calculations using Gretl
Notes:

1) X-axes denote quarters following the shock. The pink filled curves indicate 68% confidence intervals based on 2000 bootstrap replications.
2) The ordering has been changed to: sanctions, GDP growth differential, oil price growth, interest rate differential, real exchange rate.

No significant changes have occured in the variance decomposition (Figure 5b compared to Figure 3). The IRFs in Figure 5a plotting the response of the real exchange rate to unit shocks in sanctions and the interest rate differential are also identical to the ones in Figure 2, which is expected. The oil price growth shock seems to have a very similar effect on the real exchange rate as in the original ordering, albeit not quite subsided in the 20th period after the innovation.

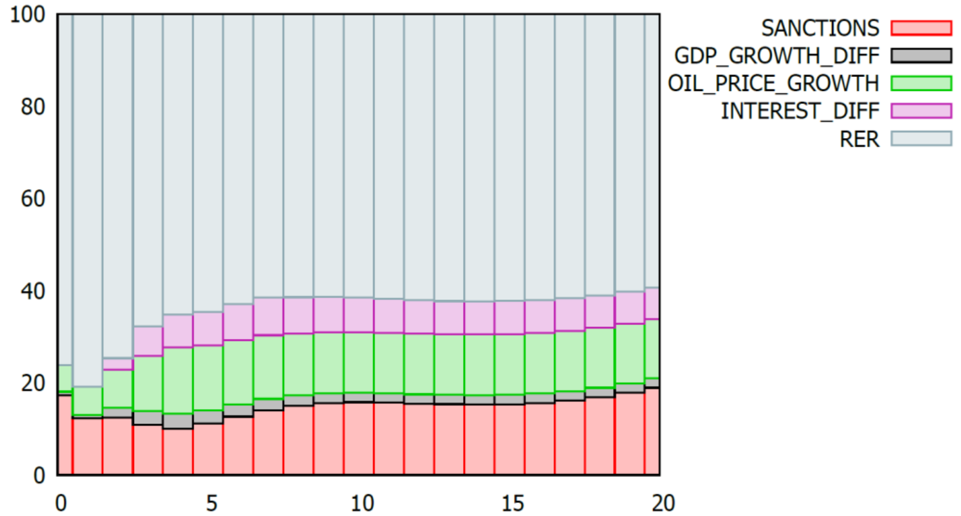

Figure 5b. Decomposition of errors for the forecast of the RUB/EUR real exchange rate, 1999Q1-2017Q3
Source: author's calculations using Gretl
Note: the ordering has been changed to: sanctions, GDP growth differential, oil price growth, interest rate differential, real exchange rate

The IRF of the real exchange rate from the GDP growth differential shock, however, shows some noteworthy variations: a positive 0.4% impact at the beginning of the response horizon in the initial ordering has now been reversed to a negative 0.5% impact. The following movement of the new IRF also seems to be inclined towards the negative side of the value range, which is more in line with the HBS theory. Nevertheless, its statistical significance remains low: statistically significant are only the initial response and responses in the second, third and fourth periods after the shock. Taking into account this new evidence, no definite conclusion can be made using the current dataset and model specification, about whether or not the real exchange rate has experienced impact from the Harrod-Balassa-Samuelson effects. This problem could further be explored by proxying productivity differentials to other macroeconomic variables, such as labor productivity.

CONCLUSIONS

The aim of this master's thesis was to analyze the effect that current anti-Russian economic sanctions have had on the evolution of the real exchange rate of rouble per euro. For this purpose, an econometric analysis was performed using a structural VAR(4) model incorporating quarterly data from 1999Q1 to 2017Q3 on five variables: the sanctions index measuring the intensiveness of sanctions, the real GDP growth differential between Russia and the euro area, the short-term interest rate differential between the two regions, the price of oil, and the real bilateral Rouble/Euro exchange rate. The reasoning behind choosing these specific variables was based on the theories of the uncovered interest rate parity, Harrod-Balassa-Samuelson effect, and previous empirical studies on the link between real exchange rates and the price of oil.

Identification is achieved through recursive short-run restrictions. The ordering of the variables is chosen so that the first variable is the most exogenous in the sense that it exerts a unilateral contemporaneous effect on the other variables and the last one is the most endogenous meaning that it experiences a unilateral contemporaneous impact from all the previous variables. Following the estimation step, 20-period impulse-response analyses and variance decompositions are performed to investigate the relationship of the real exchange rate with the other four endogenous variables. To evaluate the effect that sanctions might have had on the relationship between the real exchange rate and macroeconomic variables, a second estimation is conducted that excludes the sanctions period from the sample. Finally, to check the robustness of the obtained results, a third estimation is carried out using an alternative ordering of the variables in the SVAR system.

The first and second round of estimations covering the entire sample between 1999Q1 and 2017Q3 and the subsample spanning from 1999Q1 till 2013Q4, respectively, yielded the following results.

- The sanctions shock has an immediate negative effect of 2.2% on the real exchange rate, signalling the real appreciation of the rouble. Interestingly, the oil price growth shock has a similar negative impact of 1.3% on the RER based on the estimation of the entire sample including the period of sanctions and an almost non-existent impact at the beginning of the response horizon calculated using pre-sanctions data. This corroborates the recent finding in the empirical literature that since the introduction of sanctions, the nominal exchange rate of rouble has been actively managed by the Russian Central Bank in efforts to tie the exchange rate to the oil prices.

- The sign of the sanctions' effect on the real exchange rate between the first and 11th period of the response horizon is, in contrast, positive and varies between 0.5% and 1.9%, which is expected since the studied economic sanctions represent foreign trade distortions and a drop in the gross capital inflow.
- The real exchange rate does not show a mean reversion behaviour in response to the sanctions shock over a 20-period horizon.
- The oil price has not been more relevant in explaining the volatility of the real exchange rate than sanctions.
- The interest rate differential has consistently exerted a positively signed effect on the real exchange rate up to a maximum of 1.9% per quarter in both estimation rounds. This implies that earning excess returns in the interest rate arbitrage between the Russian and European currencies has been difficult both before and after the breakpoint in 2014Q1. This result is quite in line with Russia-oriented UIP studies that generally provide evidence supporting the validity of the UIP condition in modern Russia.

In the third round of the estimations, the order of the GDP growth differential and the oil price growth was switched to further check the validity of the Harrod-Balassa-Samuelson effect for the real exchange rate of rouble. Overall, due to the low statistical significance and opposite signs of the two comparable IRFs, no definite conclusion could be made using the current dataset and model specification, about whether or not the real exchange rate has experienced impact from the Harrod-Balassa-Samuelson effects.

The quality of the results in the current research could further be improved by

1) employing a vector error correction model instead of a structural VAR to explore the long-run relationship between sanctions and the real exchange rate;
2) proxying the productivity differential to other macroeconomic variables, such as labor productivity, instead of the GDP growth differential, to (in)validate the presence of the HBS effects.

On a related note, further research is warranted to investigate the impact that current economic sanctions have had on the currencies of the non-sanctioning neighbors and trading partners of Russia – the members of the Commonwealth of Independent States.

KOKKUVÕTE

LÄÄNE SANKTSIOONIDE, MAJANDUSKASVU JA RUBLA REAALSE VAHETUSKURSI STRUKTUURNE VEKTORAUTOREGRESSIIVNE ANALÜÜS

Katerina Savchenko

Käesoleva magistritöö eesmärgiks on analüüsida mõju, kuidas praegused Venemaa vastased majanduslikud sanktsioonid on mõjutanud rubla reaalse vahetuskursi arengut euro kohta. Selleks viidi läbi ökonomeetriline analüüs, kasutades struktuurilist VAR (4) mudelit, mis sisaldas kvartaliandmeid ajavahemikus 1999Q1 kuni 2017Q3 viie muutuja kohta: sanktsioonide intensiivsust mõõtev sanktsioonide indeks, Venemaa ja euroala SKP reaalkasvu erinevus, lühiajaline intressimäärade erinevus kahe piirkonna vahel, nafta hind ja reaalne kahepoolne rubla/euro vahetuskurss. Nende spetsiifiliste muutujate valimise põhjendused põhinesid katmata intressimäära pariteedi teooriatel, Harrod-Balassa-Samuelsoni efektil ja varasematel empiirilistel uuringutel, mis käsitlesid reaalseid vahetuskursse ja nafta hinda.

Identifitseerimine toimub rekursiivsete lühiajaliste piirangute abil. Muutujate järjestus valitakse nii, et esimene muutuja on kõige eksogeensem selles mõttes, et see avaldab ühepoolselt samaaegset mõju teistele muutujatele ja viimane on kõige endogeensem, mis tähendab seda, et sellel on eelnevate muutujate ühepoolne samaaegne mõju. Hindamisetapi järel viiakse läbi 20-perioodiline impulss-vastuse analüüs ja dispersioonide lahutus, et uurida reaalse vahetuskursi suhet teise nelja endogeense muutujaga. Hindamaks mõju, mida sanktsioonid võisid avaldada reaalse vahetuskursi ja makromajanduslike muutujate vahelistele suhetele, viiakse läbi teine hinnang, mis välistab valimisse võetud sanktsioonide perioodi. Lõpuks, saadud tulemuste usaldusväärsuse kontrollimiseks viiakse läbi kolmas hinnang, kasutades SVAR-süsteemi muutujate alternatiivset järjestust.

Esimese ja teise vooru hinnangute ring, mis hõlmas kogu valimit ajavahemikus 1999Q1 kuni 2017Q3 ja osavalimit vastavalt ajavahemikus 1999Q1 kuni 2013Q4, andsid järgmised tulemused.

- Sanktsioonide šokil on otsene negatiivne mõju 2,2% reaalsele vahetuskursile, mis annab märku rubla reaalsest kallinemisest. Huvitaval kombel on nafta hinna tõusu šokil sarnane 1,3-protsendiline negatiivne mõju, tuginedes kogu valimi hindamisele, sealhulgas sanktsioonide perioodile ja peaaegu olematu mõju reaktsiooniaja alguses, mis arvutatakse sanktsioonidele eelnevate andmete alusel. See kinnitab hiljutisi empiirilise kirjanduse järeldusi, et pärast sanktsioonide kehtestamist on Venemaa keskpank aktiivselt juhtinud rubla nominaalset vahetuskurssi, et jõuda naftahindade ja vahetuskursi sidumiseni.
- Märk sanktsioonide mõjust reaalsele vahetuskursile esimese ja üheteistkümnenda perioodi vahel on vastupidi positiivne ja varieerub vahemikus 0,5-1,9%, mida on oodata, kuna uuritud majanduslikud sanktsioonid kujutavad endast väliskaubanduse moonutusi ja kapitali sissevoolu vähenemist.
- Reaalne vahetuskurss ei näita sanktsioonide šoki järgse 20 perioodi jooksul taandumist keskmisele tasemele.
- Naftahind ei ole olnud asjakohasem selgitades reaalse vahetuskursi volatiilsust kui sanktsioonid.
- Intressimäärade erinevus on avaldanud reaalse vahetuskursile järjekindlalt positiivset mõju, mis on kvartalis mõlemas hindamisringis maksimaalselt 1,9%. See tähendab, et Venemaa ja Euroopa valuutade intressimäära arbitraaži ülejäägi teenimine on olnud raske nii enne kui ka pärast 2014. aasta 1. kvartali murdepunkti. See tulemus on küllaltki kooskõlas Venemaale orienteeritud UIP uuringutega, mis üldiselt annavad tõendeid UIP tingimuste kehtivuse kohta tänapäeva Venemaal.

Hinnangu kolmandas voorus muudeti SKP kasvutendentsi ja naftahinna kasvu järjekorda, et veelgi kontrollida Harrod-Balassa-Samuelsoni mõju kehtivust rubla reaalse vahetuskursi suhtes. Üldiselt ei saa praeguse andmekogumi ja mudeli spetsifikatsiooni abil kahe võrreldava IRFi ja vastandlike märkide tõttu lõplikku järeldust teha, kas Harrod-Balassa-Samuelsoni mõju on reaalset vahetuskurssi mõjutanud või mitte.

Praeguste teadusuuringute tulemuste kvaliteeti võiks veelgi parandada

3) rakendades vektori veaparandusmudelit struktuurilise variandi asemel, et uurida pikaajalist suhet sanktsioonide ja reaalse vahetuskursi vahel;
4) määrates tootlikkuse erinevust teiste makromajanduslike muutujate abil, näiteks tööjõu tootlikkuse näol, et kinnitada (mitte kinnitada) HBS-i mõju olemasolu.

LIST OF REFERENCES

Adler, M. and Dumas, B. (1983). International Portfolio Choice and Corporate Finance: a Synthesis. *The Journal of Finance, 38,* 925-984.

Balassa, B. (1964). The Purchasing Power Parity Doctrine: A Reappraisal. *Journal of Political Economy, 72*, 584-96.

Barsky, R. B. and Kilian, L. (2004). Oil and the Macroeconomy Since the 1970s. *Journal of Economic Perspectives, 18*, 115-34.

Benigno, G. and Thoenissen, C. (2002). *Equilibrium Exchange Rates and Supply-Side Performance.* Bank of England Working Paper No. 156.

Bilson, J. F. (1981). The "Speculative Efficiency" Hypothesis. *Journal of Business*, *54,* 435-51.

Brooks, C. (2008). *Introductory Econometrics for Finance* (2nd ed.). Cambridge: Cambridge University Press.

Caruso, R. (2003). The Impact of International Economic Sanctions on Trade: an Empirical Analysis. *Peace Economics, Peace Science, and Public Policy, 9,* 1-36.

Cashin, P., Céspedes, L. F., Sahay, R. (2004). Commodity Currencies and the Real Exchange Rate. *Journal of Development Economics*, 239-68.

Cassel, G. (1918). Abnormal Deviations in International Exchanges. *Economic Journal*, *28*, 413-15.

Chen, Y. and Rogoff, K. (2003). Commodity Currencies and Empirical Exchange Rate Puzzles. *DNB Staff Reports (discontinued)*(76).

Choudhry, T., McNown, R., and Wallace, M. (1991). Purchasing Power Parity and the Canadian Float in the 1950s. *The Review of Economics and Statistics*, *73*, 558-63.

Clarida, R. H. and Taylor, M. P. (1997). The Term Structure of Forward Exchange Premia and the Forecastibility of Spot Exchange Rates: Correcting the Errors. *Review of Economics and Statistics, 79*, 353-61.

Clarida, R., Sarno, L., Taylor, M., Valente, G. (2003). The Out-of-Sample Success of Term Structure Models as Exchange Rate Predictors: A Step Beyond. *Journal of International Economics*, *60,* 61-83.

Coakley, J., Flood, R. F., Fuertes, A.-M., and Taylor M. P. (2005). Long-run Purchasing Power Parity and the Theory of General Relativity: the First Tests. *Journal of International Money and Finance, 24*, 293-316.

Cooley, T., Dwyer, M. (1998). Business Cycle Analysis Without Much Theory: a Look at Structural VARs. *Journal of Econometrics, 83*, 57-88.

Corden, W. M. (1984). Booming Sector and Dutch Disease Economics: Survey and Consolidation. *Oxford Economic Papers, New Series, 36*, 359-80.

Diebold, F., Husted, S., and Rush, M. (1991). Real Exchange Rates under the Gold Standard. *Journal of Political Economy, 99*, 1252-71.

Dreger, C., Kholodilin, K., Ulbricht, D., and Fidrmuc, J. (2016). Between the Hammer and the Anvil: the Impact of Economic Sanctions and Oil Prices on Russia's Ruble. *Journal of Comparative Economics, 44*, 295-308.

Engel, C. (1992). On the Foreign Exchange Risk Premium in a General Equilibrium Model. *Journal of International Economics, 32*, 305-19.

Engel, C. (1996). The Forward Discount Anomaly and the Risk Premium: a Survey of Recent Evidence. *Journal of Empirical Finance, 3*, 123–92.

Fama, E. (1970). Efficient Capital Markets: A Review of Theory and Empirical Work. *Journal of Finance*, 25, 383-423.

Fama, E. F. (1984). Forward and Spot Exchange Rates. *Journal of Monetary Economics, 14*, 319–38.

Frankel, J. A. (1982). In Search of the Risk Premium: A Six-Country Model Using Mean-Variance Optimization. *Journal of International Money and Finance, 1*, 255-74.

Froot, K. A. and Thaler, R. H. (1990). Foreign exchange. *Journal of Economic Perspectives, 4*, 179–92.

Giovannini, A. and Jorion, P. (1989). The Time Variation of Risk and Return in the Foreign Exchange and Stock Markets. *The Journal of Finance, 44*, 307-25.

Gurvich, E. and Prilepskiy, I. (2016). The Impact of Financial Sanctions on the Russian Economy. *Voprosy Economiki, 1*, 36-62.

Habib, M. M. and Kalamova, M. (2007). *Are There Oil Currencies? The Real Exchange Rate of Oil Exporting Countries.* European Central Bank Working Paper Series No. 839.

Harrod, R. (1933). *International Economics.* London: James Nisbet.

Hodrick, R. J. (1987). *The Empirical Evidence on the Efficiency of Forward and Futures Foreign Exchange Markets.* London: Harwood.

Hoffmann, M. and Neuenkirch, M. (2015). *The Pro-Russian Conflict and its Impact on Stock Returns in Russia and the Ukraine*. University of Trier Research Papers in Economics No. 1/15.

Kholodilin, K. A. and Netsunajev, A. (2017). *Crimea and Punishment: the Impact of Sanctions on Russian and European Economies.* Bank of Estonia Working Papers wp2017-5.

Kilian, L. (2011). *Structural Vector Autoregressions.* CERP Discussion Papers 8515.

Korhonen, I. and Juurikkala, T. (2007). *Equilibrium Exchange Rates in Oil-Dependent Countries.* BOFIT Discussion Paper No. 8/2007.

Lafrance, R., Osakwe, P., and Normandin, M. (1998). Evaluating Alternative Measures of the Real Effective Exchange Rate. *Bank of Canada, Working Paper*, 98-20.

Lanne, M., Meitz, M., and Saikkonen, P. (2017). Identification and Estimation of Non-Gaussian Structural Vector Autoregressions. *Journal of Econometrics*, 288-304.

Lektzian, D. and Patterson, D. (2015). Political Cleavages and Economic Sanctions: The Economic and Political Winners and Losers of Sanctions. *International Studies Quarterly, 59*(1).

Lewis, K. K. (1989a). Can Learning Affect Exchange-Rate Behaviour? The Case of the Dollar in the Early 1980s'. *Journal of Monetary Economics, 23*, 79-100.

Lewis, K. K. (1989b). Changing Beliefs and Systematic Rational Forecast Errors with Evidence from Foreign Exchange. *American Economic Review, 79*, 621-36.

Lewis, K. K. (1995). Puzzles in International Financial Markets. In G. M. Grossman, & K. Rogoff, *Handbook of International Economics* (Vol. III, pp. 1913-71). Amsterdam: North-Holland.

Lipsey, R. E. and Swedenborg, B. (1996). The High Cost of Eating: Causes of International Differences in Food Prices. *Review of Income and Wealth, 42*(2).

Lothian, J. and Taylor, M. (2008). Real Exchange Rates Over the Past Two Centuries: How Important is the Harrod-Balassa-Samuelson Effect? *Economic Journal, 118,* 1742-63.

Lucas, R. E. (1982). Interest Rates and Currency Prices in a Two-Country World. *Journal of Monetary Economics, 10*, 335-59.

Lütkepohl, H. (2010). Variance Decomposition. rmt: S. N. Durlauf, *Macroeconometrics and Time Series Analysis* (lk 369-71). London: Palgrave Macmillan.

Mark, N. (1990). Real and Nominal Exchange Rates in the Long Run: an Empirical Investigation. *Journal of International Economics, 28*, 115-36.

McCloskey, D. N. and Zecher, J. R. (1984). The Success of Purchasing-Power Parity: Historical Evidence and Its Implications for Macroeconomics. In M. D. Bordo, *A Retrospective on the Classical Gold Standard, 1821-1931* (pp. 121-72). Chicago: University of Chicago Press.

McNees, S. K. (1986). Modeling the Fed: A Forward-Looking Monetary Policy Reaction Function. *New England Economic Review*, 3-8.

McNown, R. and Wallace, M. S. (1989). National Price Levels, Purchasing Power Parity, and Cointegration: A Test of Four High Inflation Economies. *Journal of International Money and Finance, 8*, 533-45.

Neuenkirch M. and Neumeier, F. (2015). The Impact of UN and US Economic Sanctions on GDP Growth. *European Journal of Political Economy, 40*, 110-125.

Ricci, L. A., Milesi-Ferretti G., and Lee, J. (2008). *Real Exchange Rates and Fundamentals: A Cross-Country Perspective.* IMF Working Paper 08/13.

Rogoff, K. (1979). *Essays on Expectations and Exchange Rate Volatility. Ph.D. Dissertation.* Massachusetts Institute of Technology.

Rogoff, K. (1996). The Purchasing Power Parity Puzzle. *Journal of Economic Literature, 34*, 647–68.

Samuelson, P. A. (1964). Theoretical Notes on Trade Problems. *Review of Economics and Statistics*, 145-54.

Sarno, L. (2005). Viewpoint: Towards a Solution to the Puzzles in Exchange Rate Economics: Where Do We Stand? *Canadian Journal of Economics, 38*, 673–708.

Sarno, L. and Chowdhury, I. (2003). The Behaviour of the Real Exchange Rate: Evidence from an Alternative Price Index. *Economic Notes, 32*, 295-333.

Sarno, L. and Taylor, M. P. (2003). *The Economics of Exchange Rates.* Cambridge: Cambridge University Press.

Sims, C. (1980). Macroeconomics and Reality. *Econometrica, 48*, 1-48.

Summers, R. and Heston, A. (1991). The Penn World Table (Mark 5): An Expanded Set of International Comparisons, 1950–1988. *The Quarterly Journal of Economics, 106*, 327-68.

Taylor, M. and McMahon, P. C. (1988). Long-Run Purchasing Power Parity in the 1920s. *European Economic Review*, 179-97.

Taylor, M. P., Peel, D. A., and Sarno, L. (2001). Nonlinear Mean-Reversion in Real Exchange Rates: Toward a Solution to the Purchasing Power Parity Puzzles. *International Economic Review, 42*, 1015-42.

Tyll, L. P. (2018). The Impact of Economic Sanctions on Russian Economy and the RUB/USD Exchange Rate. *Journal of International Studies*, 21-33.

Vasilyev, D., Busygin, V., and Busygin, S. (2017). Testing and Interpreting Uncovered Interest Parity in Russia. *Russian Journal of Economics, 3*, 158-73.

Yahia, A. F. and Saleh, A. S. (2008). Economic Sanctions, Oil Price Fluctuations and Employment: New Empirical Evidence from Libya. *American Journal of Applied Sciences, 5*, 1713-19.

APPENDICES

Appendix 1. Composition of the sanctions index

Year	Month	Day	Country and sanction type – Dreger *et al.* (2016)		Country and sanction type – author's compilation	
			1 – directed against individuals, 2 – against certain companies, 3 – against entire industries			
2014	3	6	USA	1	USA	1
2014	3	17	USA	1	USA	1
2014	3	17	EU	1	EU	1
2014	3	17	Canada	1	Canada	1
2014	3	17	Japan	3	Japan	3
2014	3	19	Canada	1	Canada	1
2014	3	19	Australia	1	Australia	1
2014	3	20	USA	2	USA	2
2014	3	21	Canada	2	Canada	2
2014	3	21	EU	1	EU	1
2014	3	28	Canada	2	Canada	2
2014	4	11	Albania, Iceland, Montenegro, Ukraine	1	Albania, Iceland, Montenegro, Ukraine	1
2014	4	11	USA	2	USA	2
2014	4	28	USA	3	USA	3
2014	4	29	Japan	1	Japan	1
2014	4	29	EU	2	EU	2
2014	5	4	Canada	2	Canada	2
2014	5	12	Canada	2	Canada	2
2014	5	12	EU	2	EU	2
2014	5	21	Australia	2	Australia	2
2014	6	21	Canada	2	Canada	2
2014	6	24	Canada	2	Canada	2
2014	7	12	EU	1	EU	1
2014	7	16	USA	2	USA	2
2014	7	25	EU	2	EU	2
2014	7	29	USA	3	USA	3
2014	7	30	EU	2	EU	2
2014	7	31	EU	3	EU	3
2014	7	31	–	–	EU	3
2014	7	31	–	–	EU	3
2014	8	6	Canada	2	Canada	2
2014	8	14	Ukraine	2	Ukraine	2
2014	9	8	EU	2	EU	1

Appendix 1 continued

2014	9	12	USA	2	USA	1
2014	9	12	–	–	EU	3
2014	9	12	–	–	EU	3
2014	9	12	–	–	EU	3
2014	9	16	Canada	2	Canada	2
2014	11	28	–	–	EU	2
2014	12	18	–	–	EU	3
2014	12	19	Canada	3	Canada	2
2014	12	19	–	–	Canada	3
2014	12	19	–	–	USA	2
2015	2	9	EU	2	EU	2
2015	2	16	–	–	EU	2
2015	2	17	Canada	2	Canada	2
2015	3	31	Australia	3	Australia	3
2015	12	22	–	–	USA	1
2015	12	22	–	–	USA	2
2015	12	22	–	–	USA	3
2016	9	1	–	–	USA	3
2016	11	9	–	–	EU	1
2016	11	14	–	–	USA	1
2016	12	20	–	–	USA	1
2016	12	20	–	–	USA	2
2016	12	20	–	–	USA	3
2017	6	20	–	–	USA	1
2017	6	20	–	–	USA	2
2017	6	20	–	–	USA	3
2017	8	4	–	–	EU	2
2017	9	29	–	–	USA	3
2017	9	29	–	–	USA	3

Source: Dreger *et al.* (2016, 301), European Union External Action, US Department of the Treasury

Appendix 2. Unit-root test results from Gretl

```
Augmented Dickey-Fuller test for SANCTIONS
testing down from 11 lags, criterion AIC
sample size 63
unit-root null hypothesis: a = 1

  test without constant
  including 11 lags of (1-L)SANCTIONS
  model: (1-L)y = (a-1)*y(-1) + ... + e
  estimated value of (a - 1): -2.73468
  test statistic: tau_nc(1) = -4.44762
  asymptotic p-value 9.572e-006
  1st-order autocorrelation coeff. for e: 0.003
  lagged differences: F(11, 51) = 48.549 [0.0000]
```

Appendix 2 continued

```
  test with constant
  including 11 lags of (1-L)SANCTIONS
  model: (1-L)y = b0 + (a-1)*y(-1) + ... + e
  estimated value of (a - 1): -2.71311
  test statistic: tau_c(1) = -4.41081
  asymptotic p-value 0.0002789
  1st-order autocorrelation coeff. for e: 0.003
  lagged differences: F(11, 50) = 47.084 [0.0000]

  with constant and trend
  including 11 lags of (1-L)SANCTIONS
  model: (1-L)y = b0 + b1*t + (a-1)*y(-1) + ... + e
  estimated value of (a - 1): -2.6542
  test statistic: tau_ct(1) = -4.39644
  asymptotic p-value 0.00216
  1st-order autocorrelation coeff. for e: 0.001
  lagged differences: F(11, 49) = 45.704 [0.0000]

Augmented Dickey-Fuller test for RER
testing down from 11 lags, criterion AIC
sample size 74
unit-root null hypothesis: a = 1

  test without constant
  including 0 lags of (1-L)RER
  model: (1-L)y = (a-1)*y(-1) + e
  estimated value of (a - 1): -0.00272181
  test statistic: tau_nc(1) = -1.48675
  p-value 0.1274
  1st-order autocorrelation coeff. for e: -0.105

  test with constant
  including 0 lags of (1-L)RER
  model: (1-L)y = b0 + (a-1)*y(-1) + e
  estimated value of (a - 1): -0.0891809
  test statistic: tau_c(1) = -2.89226
  p-value 0.05106
  1st-order autocorrelation coeff. for e: -0.112

  with constant and trend
  including 0 lags of (1-L)RER
  model: (1-L)y = b0 + b1*t + (a-1)*y(-1) + e
  estimated value of (a - 1): -0.120289
  test statistic: tau_ct(1) = -2.38842
  p-value 0.3825
  1st-order autocorrelation coeff. for e: -0.092

Augmented Dickey-Fuller test for OIL_PRICE_GROWTH
testing down from 11 lags, criterion AIC
sample size 74
unit-root null hypothesis: a = 1

  test without constant
  including 0 lags of (1-L)OIL_PRICE_GROWTH
  model: (1-L)y = (a-1)*y(-1) + e
  estimated value of (a - 1): -1.02355
  test statistic: tau_nc(1) = -8.96012
  p-value 3.131e-030
  1st-order autocorrelation coeff. for e: 0.039

  test with constant
  including 0 lags of (1-L)OIL_PRICE_GROWTH
  model: (1-L)y = b0 + (a-1)*y(-1) + e
  estimated value of (a - 1): -1.02446
```

Appendix 2 continued

```
  test statistic: tau_c(1) = -8.89772
  p-value 3.921e-007
  1st-order autocorrelation coeff. for e: 0.039

  with constant and trend
  including 0 lags of (1-L)OIL_PRICE_GROWTH
  model: (1-L)y = b0 + b1*t + (a-1)*y(-1) + e
  estimated value of (a - 1): -1.04265
  test statistic: tau_ct(1) = -9.04578
  p-value 3.058e-008
  1st-order autocorrelation coeff. for e: 0.035

Augmented Dickey-Fuller test for GDP_GROWTH_DIFF
testing down from 11 lags, criterion AIC
sample size 73
unit-root null hypothesis: a = 1

  test without constant
  including one lag of (1-L)GDP_GROWTH_DIFF
  model: (1-L)y = (a-1)*y(-1) + ... + e
  estimated value of (a - 1): -0.37706
  test statistic: tau_nc(1) = -3.57859
  asymptotic p-value 0.0003413
  1st-order autocorrelation coeff. for e: -0.026

  test with constant
  including 0 lags of (1-L)GDP_GROWTH_DIFF
  model: (1-L)y = b0 + (a-1)*y(-1) + e
  estimated value of (a - 1): -0.55016
  test statistic: tau_c(1) = -5.32202
  p-value 2.56e-005
  1st-order autocorrelation coeff. for e: -0.081

  with constant and trend
  including 0 lags of (1-L)GDP_GROWTH_DIFF
  model: (1-L)y = b0 + b1*t + (a-1)*y(-1) + e
  estimated value of (a - 1): -0.73178
  test statistic: tau_ct(1) = -6.44245
  p-value 3.788e-006
  1st-order autocorrelation coeff. for e: -0.009

Augmented Dickey-Fuller test for INTEREST_DIFF
testing down from 11 lags, criterion AIC
sample size 63
unit-root null hypothesis: a = 1

  test without constant
  including 11 lags of (1-L)INTEREST_DIFF
  model: (1-L)y = (a-1)*y(-1) + ... + e
  estimated value of (a - 1): -0.0690916
  test statistic: tau_nc(1) = -2.22348
  asymptotic p-value 0.02525
  1st-order autocorrelation coeff. for e: 0.024
  lagged differences: F(11, 51) = 7.259 [0.0000]

  test with constant
  including 11 lags of (1-L)INTEREST_DIFF
  model: (1-L)y = b0 + (a-1)*y(-1) + ... + e
  estimated value of (a - 1): -0.117403
  test statistic: tau_c(1) = -2.65174
  asymptotic p-value 0.08268
  1st-order autocorrelation coeff. for e: 0.025
  lagged differences: F(11, 50) = 7.423 [0.0000]
```

Appendix 2 continued

```
with constant and trend
including 11 lags of (1-L)INTEREST_DIFF
model: (1-L)y = b0 + b1*t + (a-1)*y(-1) + ... + e
estimated value of (a - 1): -0.160272
test statistic: tau_ct(1) = -3.01411
asymptotic p-value 0.1282
1st-order autocorrelation coeff. for e: 0.018
lagged differences: F(11, 49) = 7.659 [0.0000]
```