Review of Economic Perspectives – Národohospodářský obzor
Vol. 21, Issue 3, 2021, pp. 309–346, DOI: 10.2478/revecp-2021-0014

Oil price shock in the US and the euro area – evidence from the shadow rate and the term premium

Martin Pažický[1]

Abstract: The aim of this article is to investigate the consequences of oil price changes for the economy of the US and the euro area. Oil price transmission channel is assessed using Granger causalities and structural vector autoregressive (*VAR*) specifications (applying the Cholesky factorization and the restrictions following the method of Blanchard and Quah). The conventional oil price transmission channel is extended by a shadow policy rate and term premium, as the importance of both indicators has been growing rapidly in recent years. The results confirm that the oil price shock is not negligible in the aftermath of the Global Financial Crisis and in the subsequent period of monetary policy normalization. The findings are confirmed by the outcomes of the Bayesian VAR specification with sign restrictions. The consequences of changes in oil prices have significantly grown since the introduction of unconventional monetary instruments. The magnitude of the response of industrial production, price level and shadow interest rate to the oil price shock is strongest in the period corresponding to the unconventional monetary policy. In many cases, however, the reaction is short-lived. The conventional instrument (policy rate) in the euro area has still not been sufficient to stabilize the economy in the recent period of monetary policy normalization in the US.

Keywords: oil price, unconventional monetary policy, VAR, shadow rate, term premium

JEL Classification: B23, E42, E52, E58, Q43

Received: 1 December 2020 / Accepted: 9 June 2021 / Sent for Publication: 9 September 2021

Introduction

Monetary policy has had a good reputation since the Great Moderation era in the mid 1980s. Especially after the Global Financial Crisis (GFC) in 2007 and the European sovereign debt crisis in 2014, the policymakers in the US and the euro area have adopted unconventional monetary instruments designed to ease and stimulate financial markets and the real economy. This set of unconventional monetary policy measures has resulted in the unprecedented expansion in the balance sheets of both central banks[2] and in the

[1] Comenius University in Bratislava, Faculty of Management, Odbojárov 10, 820 05 Bratislava, Slovak Republic; e-mail: Martin.Pazicky@fm.uniba.sk

[2] The balance sheet of the Federal Reserve increased from 0.9 trillion USD in August 2008 to more than 4.5 trillion USD in December 2014. The ECB's balance sheet expanded from 1.5 billion EUR in 2007 to 4.5 billion EUR in December 2018.

compression of interest rates to historical lows. In particular, the term premia have been squeezed below zero due to unconventional measures.

Although the monetary policy in the US gradually started to normalize and the interest rate slowly increased from 0 to 2.75 percent in March 2019, the ECB's key interest rate remains at zero. The economic fundamentals in the euro area (particularly in German car industry) deteriorated in 2019, which caused that the ECB's Governing Council responded by a new package of monetary easing in September 2019.[3] At the end of 2019, the ECB did not have many options to suppress the upcoming recession. This was also clear to the ECB's representatives, who at a press conference in September 2019 called on the member governments to reconsider the fiscal stimulus. A remarkable commentary on the effectiveness of central banks and their interdependence on fiscal policy is available in Hartwell's study (2018).

Given that the instruments that central banks have at their disposal after a long period of unconventional monetary policies are limited, the ability of central banks to mitigate shocks was questionable. Their ability to withstand shock was tested in early 2020, when, as a result of the corona crisis, the world's economies fell into the deepest recession in modern history. Although the response of monetary authorities both in the US and in the euro area has been unprecedented in its promptness and magnitude, maintaining the labour market stability would not have been possible without effective government interventions, resulting in a sharp rise in government debt in individual countries[4]. One of the side effects of the corona shock was a sudden drop in global demand, which manifested itself through a heavy drop in oil prices. The aim of our work is to evaluate the oil price transmission channel in the US and the euro area in the environment of low interest rates shortly before the coronavirus outbreak. We verify whether the economies of these countries were more sensitive to economic shocks than before the adoption of unconventional monetary policies. Global economic shocks in particular, such as the oil price shock, could paralyze even large economies such as the US and the euro area. Due to the lack of monetary policy ammunition (i.e., low interest rates), the oil price shock could potentially have a much more significant impact on the economies of these countries than it used to have in the past.

Multiple studies have documented that since the introduction of unconventional monetary policies, traditional instruments have been rendered useless for the assessment of monetary policy – for example Borio and Zabai (2016), Coenen et al. (2017), Glick and Leduc (2018), Mouabbi and Sahuc (2019) or Pažický (2018, 2019). The apparent lack of sensitivity of longer-term interest rates to changes in short-term rates has already been

[3] A renewed QE package of the ECB released on 12 September 2019 consisted of: (1) an open-ended QE program for public and private assets that will run at a pace of 20bn EUR per month until "shortly before" key interest rates start rising; (2) a 10bp cut in the deposit rate along with strengthened forward guidance; (3) two-tier system for reserve remuneration; (4) easier TLTRO-III terms for the banks.

[4] The balance sheet of the Federal Reserve almost doubled to 7.1 trillion USD in May 2020, while the ECB's balance sheet exceeded 6.4 billion EUR in August 2020. The Coronavirus Aid, Relief and Economy Security Act ("CARES Act") in the US is estimated at 2.3 trillion USD (around 11% of GDP). The announced joint debt of the European Union is expected to amount to 750 billion EUR.

emphasized by Rudebusch, Sack and Swanson (2007). Greenspan (2005) noted that the broadly unanticipated behaviour of global bond market remains a conundrum. Economists have therefore focused on alternative indicators of monetary policy. *The shadow policy rate* (see Krippner, 2012 or Lombardi and Zhu, 2018 or Wu and Xi, 2016) and term premium have become particularly popular – Rudebusch and Wu (2004), Kim and Wright (2005), Bańbura et al. (2008), Carboni (2014) or Cohen, Hördahl and Xia (2018).

Gagnon et al. (2011) examine the changes in long-term bond yields that are driven by the changes in expectations of the future path of short-term interest rates. Bond yields could be decomposed into a part reflecting the change in the rationally anticipated average level of short rates and a part reflecting a change in the size of the *term premium*. With this approach, the term premium is a candidate for an alternative monetary policy indicator at the time of unconventional monetary instruments (see discussion in Woodford, 2012 or Bundick, Herriford and Smith, 2017 or Hubert and Labondance, 2018). Given that the importance of the term premium grows only in the period of unconventional monetary policy, in our work, we consider the shadow policy rate as an indicator of monetary policy. First, therefore, we verify the response of the shadow policy rate to the oil price shock. We then also verify the response of the term premium as an alternative indicator of monetary policy.

Another stream of literature focuses on the interactions between monetary policy and commodity prices and their effects on the real economy. Wanke (2017) shows that while the relationship between the price of crude oil and yields of US Treasury bonds with longer maturities was still negative in the autumn of 2016, it has since reversed and become increasingly stronger. Vannelli (2018) claims that the term premium was historically related to oil prices, as energy shocks were largely a function of the US dependence on foreign oil. Segal (2007), on the other hand, claims that oil prices have never been as important as is usually thought. Sekine and Tsuruga (2018) estimate a cross-country panel to assess the effects of commodity price shocks on headline inflation. They find that the effects of commodity price shocks on inflation disappear within about one year after the shock. For further evidence from the US, see Fornero and Kirchner (2018), Choi et al. (2017) or Elder and Serletis (2010). For observations from Europe, see Akram et al. (2015) or Sussman (2015). Different approach was used by De Vijlder (2018), who applied the *NiGEM* model[5] to assess the oil price shock.

Our study contributes to the existing literature by assessing the effectiveness of the Federal Reserve's and the ECB's monetary policy since the introduction of unconventional monetary instruments. We focus on comparing the impacts of oil price shocks before and after the adoption of unconventional monetary policy measures, and in the recent period of monetary policy normalization. First, we verify the transmission channel through which oil price changes affect the real economy. Negative oil price

[5] *NiGEM* (National Institute Global Econometric Model) is a peer reviewed global econometric model that has evolved over 30 years. It is a detailed country level model including over 60 countries and regions, which provides access to forecasts, scenarios and stochastic output for over 5,000 macro variables.

shock[6] is assumed to decelerate industrial production, which motivates investors to reallocate their resources into safer government bonds, squeezing bond yields and their term premia. A rise in oil prices is likely to result in an increase in overall prices. On the other hand, a more expensive commodity may cause the inflation expectations to fall due to the expected economic slowdown, leading to a decline in overall prices. To verify the transmission channel, we test the following transmission channel:

H1: A negative oil price shock lowers production and increases price level, which motivates central banks to decrease policy rate and hence term premium.

We further assume that since the introduction of unconventional monetary instruments, the economy is more sensitive to external shocks. Given the key policy zero lower bound, negative term premium and inflation below the target of 2 percent (at least in the euro area), the effectiveness of monetary policy in the recent period can be questioned. As a result, the economy is more vulnerable to external shocks than in the past. The consequences of the oil price shock are examined in individual periods testing the following hypothesis:

H2: The oil price shock has greater consequences for the economy since the introduction of unconventional monetary instruments.

Our analyses confirm that a negative oil price shock leads to a decrease in production and to an increase in overall price level. However, we have found some deviations from the transmission channel in the individual sub-periods. Our findings suggest that the importance of the term premium grew substantially in times of unconventional monetary policy, and its impact on treasury yield may have even outweighed the signal from the expectations component. We further document that the consequences of the changes in oil prices have significantly grown since the introduction of unconventional monetary instruments. In many cases, however, the reaction is short-lived. Imbalances and vulnerabilities may accumulate during the use of unconventional monetary instruments as conventional instruments have become ineffective.

The rest of our study is structured as follows: section 2 describes our dataset, while a detailed methodology and empirical results are available in sections 3 and 4, respectively. The conclusion follows in section 5.

1. Data

Our work draws from a dataset that consists of monthly oil prices and macroeconomic variables for the US and for the euro area. The oil prices (OP) are approximated by the WTI crude oil price index and the same data are used for both the US and the euro area economy.

For the US, we collected monthly data ranging from May 1983 to December 2019 (392 observations) on industrial production (IND) as an indicator of economic activity,

[6] We refer to an increase in oil prices as a negative oil price shock, which leads to an increase in inflation and a fall in production.

Consumer Price Index (CPI)[7] as an approximation of core inflation, shadow policy rate (SR) as an indicator of monetary policy and 10-year Treasury term premium (TP). All the data for the US were obtained from the Federal Reserve economic database, except for the shadow policy rate that we predicted using Wu and Xi (2016) dataset (see Appendix C). In respect of the unconventional monetary policy of the US Federal Reserve, we identify two important events, which result in three sub-periods: (1) November 2007, just before the US Federal Reserve began coordinating the unprecedented action of five leading central banks around the world on 13 December 2007 to offer billions of dollars in loans to banks; (2) 12 December 2012, when an increase in open-ended asset purchases from 40 billion USD to 85 billion USD per month was announced by the FOMC. The announcement is known as QE3. Although this was not the first QE (i.e., quantitative easing) measure, it was an unprecedented policy announcement due to its large scale and infinite nature. As a result, our first sub-period ranging from May 1983 to November 2007 represents a period when *conventional monetary policy instruments* were used. The second sub-period ranging from December 2007 to November 2012 covers the *time of unconventional monetary policy*. The last sub-period spans from December 2012 to December 2019 and captures the recent period of *monetary policy normalization*. The data overview for each sub-period is available in Appendix A.

We use exactly the same variables for the euro area, but we only managed to collect data from July 2005[8] to December 2019 (174 monthly observations). For industrial production (*IND*) and *CPI*[9], we obtained the data from Eurostat. The data on the shadow policy rate are estimated using Wu and Xi (2016) dataset (see Appendix C). No time series with estimated term premium (*TP*) for the euro area is available. For this reason, we produce our own estimate of term premia for the euro area using affine term structure model developed by Adrian, Crump and Moench (2013, 2014). To estimate the 10-year term premium for the euro area, we use the Eonia curve as an approximation of the euro area risk-free curve. The estimation is described in Appendix B. We ignore the fact that the time frame of our euro area dataset is compressed compared to the US dataset and use exactly the same sub-periods as for the US. The data description is available in Appendix A. The visual overview of the original data is captured in Figure A.1, Appendix A.

2. Methodology

The models presented in this paper belong to the class of vector autoregressive (*VAR*) models. We first estimate the *VAR* model with short-run restrictions, then, we estimate the *VAR* model with long-run restrictions in the fashion of Blanchard and Quah (1989), and finally, we impose sign restrictions. All three *VAR* specifications are estimated separately for each sub-period in the US and in the euro area.

Before turning to estimating the *VAR* models, we first perform the Augmented Dickey-Fuller (*ADF*) test for unit root to verify the stationarity of included variables and their

[7] As a measure of core inflation in the US, we use CPI index for all items except for food & energy from the US Bureau of Labor Statistics.

[8] Data on the OIS curve, used to calculate the term premium in the euro area, have been available since July 2005.

[9] As a measure of core inflation in the euro area, we use HICP overall index excluding energy, food, alcohol and tobacco from Eurostat.

order of integration. The generally accepted method for tackling non-stationarity is data transformation in first difference – see Enders (2010). We found our data to be non-stationary, integrated at *I(1)* – see Table A.2 in Appendix A. As further applied time series analyses require the data to be stationary, we transfer the data in differenced form instead of level. We further check Pearson correlations and use the Granger causality test to determine the relationships between the variables. The Granger causality test verifies a unidirectional causality between two variables. The disadvantage of the Granger causality test is that it shows the unidirectional causality while taking into account only the information between the two examined variables. We use the Granger causality test as a supplementary source of information on the oil price transmission channel.

We proceed with a general *VAR* model specification. Each variable in the *VAR* has an equation explaining contemporaneous effects based on its own lagged values, the lagged values of the other variables in the equation system, and an error term. In our models, we inspect the oil price shock, which is assumed to be an exogenous shock to the system of equations. The oil prices (OP) are expected to drive the changes in the industrial production (IND) and the price level in the economy (CPI). A central bank reacts to the changes in inflation by adjusting its key policy rate (SR), which affects the components of treasury yields – the expectations component and the term premium component (TP).

A complete technical procedure for estimating the baseline *VAR* model extended by short-run restrictions using the Cholesky decomposition and by long-run restrictions using the identification of Blanchard and Quah (1989) is available in Appendix D. The order of variables in the vector autoregressive process is as follows: oil price, industrial production, CPI inflation, shadow policy rate and term premium. The proposed order is in line with the standard view on how the changes in oil prices affect the variables of the real economy and therefore assumes a usual channel for the transmission of oil prices (see for example Lippi and Nobili, 2008 or Melolinna, 2012).

Finally, the sign restrictions are imposed by the means of the Bayesian framework using the maximal likelihood estimator (*MLE*) instead of *OLS*. The process of constructing the Bayesian *VAR* model and imposing sing restrictions is described in Appendix D. As we are concerned with the oil price shock, we show the proposed sign restrictions capturing the expected response of variables in the system (*OP*, *IND*, *CPI, SR, TP*) to the negative oil price shock (*OP*) in Table 1.

Table 1. Sign restrictions

Variables	**Oil price shock** ↑
Oil price	n/a
Industrial production	−
CPI	+
Shadow rate	+
Term premium	n/a

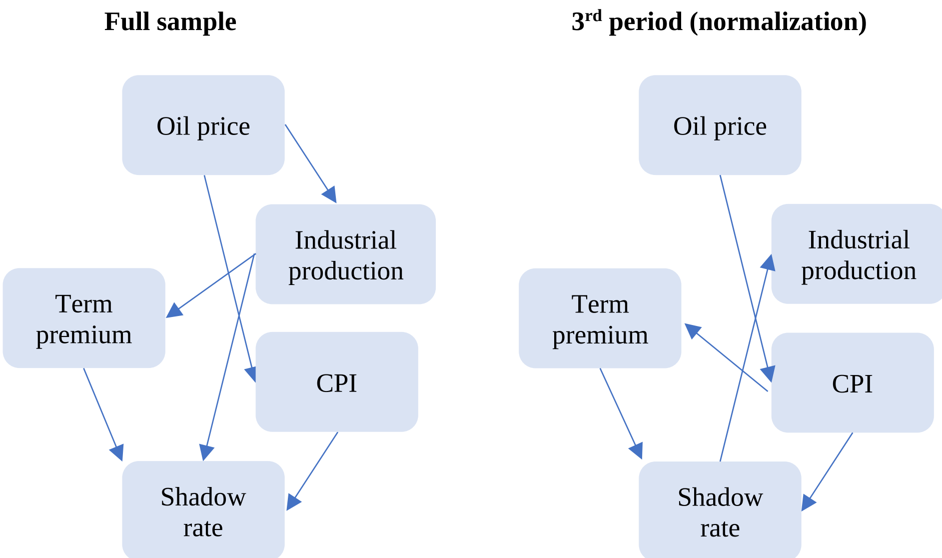
Source: own prepared

The sign restrictions proposed in Table 1 reflect the oil price transmission channel. A negative oil price shock is expected to increase the production costs, which then reduces the industrial production. A supply shortage causes a higher price level. The central bank responds to higher inflation by a policy rate hike, which should translate into a rise in the expectations component. Because we do not know whether the expectations component increases at a slower or faster pace than the policy rate, we are unable to decide on the final response of the term premium component. Finally, we do not impose any restrictions on the response of oil prices to the oil price shock. We expect that the oil price shock will translate into the oil prices in a same direction. Therefore, we aim to avoid imposing additional restrictions, which ultimately verifies the reliability of our model.

3. Empirical results

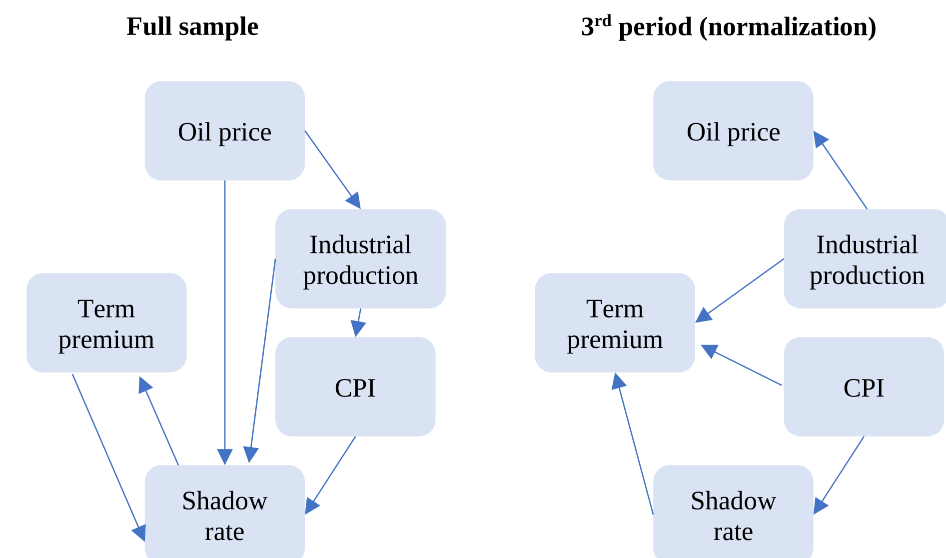
To identify the unidirectional causalities between variables, we used the Granger causality test to evaluate the oil price transmission channel. The results of the Granger causality test for the **US** are captured in Figure 1, where we observe that, in the full sample period, oil price Granger causes industrial production and price level. We further observe that the changes in industrial production and prices lead to the changes in shadow policy rate, which is consistent with the economic theory. Changes in industrial production cause a change in the term premium and the changes in the term premium ultimately affect the level of shadow policy rate. Based on the findings of the Granger causality test for the full sample period, we can confirm the oil price transmission channel for the US. At the same time, however, we observe that the transmission channel was disrupted in the period of unconventional monetary policy and was only partially renewed in the last period of monetary policy normalization – see Figure 1. The complete results of the Granger causality test for the US are captured in Table E.1, Appendix E. For the euro area, we confirm the oil price transmission channel in the full sample period. In general, the oil price affects the industrial production, which further affects the price level. The central bank responds to the changes in price level by adjusting the interest rate. The test does not confirm a unidirectional causality between the shadow policy rate and the term premium, but instead indicates that the variables interact. As in the US, the oil price transmission channel was disrupted in the period corresponding to the unconventional monetary policy. As shown in Figure 2, the oil price transmission channel in the euro area, unlike in the US, has not recovered since GFC (see Figure D.2, Appendix E for the complete Granger causality results).

Figure 1. Oil price transmission channel based on the Granger test in the US



Note: The diagrams are created on the basis of the results of the Granger causality test captured in Table E.1, Appendix E. The arrows indicate unidirectional causality between variables.
Source: prepared by the author

Figure 2. Oil price transmission channel based on the Granger test in the euro area



Note: The diagrams are created on the basis of the results of the Granger causality test captured in Table E.2, Appendix E. The arrows indicate unidirectional causality between variables.
Source: prepared by the author

4.1 Short-run and long-run restrictions

In this sub-section, we present our results related to short-run restrictions imposed by Cholesky factorization and long-run restrictions proposed by Blanchard and Quah (1989). We focus only on the oil price shock.

The results for the **US** are shown in Figure 3. For a better comparison of the short-run and long-run effects, we display the cumulative impulse response functions and their 68% confidence bands of Cholesky factorization and Blanchard and Quah factorization always within the same graph.

It is clear from Figure 3 that the impulse response of ***industrial production*** to a negative oil price shock is negative in the full sample period, which is in accordance with the general expectations. In the full sample period, the oil price shock has a relatively small immediate effect on industrial production. About 6 months after the shock, however, the response increases by −0.3 standardized units[10] and remains persistent until the end of the time horizon. The long-run response is much stronger at the beginning (about −0.15 standardized units) and further decreases to −0.4 standardized units. The profile of the impulse response function under long-run restrictions is consistent with the profile under short-run restrictions. In the first sub-period, the short-run response of industrial production is non-significant. The profile is almost identical under long-run restrictions, yet the immediate response is positive and significant, and disappears around 3 months after the shock. The results in the second and third sub-period are implausible, as the responses are not significant under both short- and long-run restrictions.

Turning to the response of the price level, we achieved plausible results, given that the responses to the oil price shock are positive and statistically significant under both short- and long-run restrictions in all sub-periods. Interestingly, the magnitude of price level response to the oil price shock in the second period (representing unconventional monetary policy) increased compared to the period of conventional monetary policy. Such an observation can be interpreted as reinforcing the effect of the oil price transmission channel on inflation in times when conventional monetary instruments are less effective. In the following period of monetary policy normalization, the magnitude of the impulse response function scaled back again.

The presented results are also consistent for the ***shadow policy rate***. Figure 3 confirms that a central bank seeks to mitigate a negative oil price shock by raising its policy rate. In the full sample period, we observe a positive response of shadow policy rate, which does not disappear until the end of the entire horizon. Although the immediate response is significant under long-run restrictions, its significance disappears 3 periods after the shock. In the period of conventional monetary policy, the shadow rate was, as a proxy for monetary policy rate, an effective tool of a central bank for managing inflation, as the responses are significant. By contrast, in times corresponding to the unconventional monetary policy, the interest rate does not have a sufficient power to stabilize the changes in oil prices. We know that in that period the level of inflation was very low in spite of the zero key policy rates – in other words, the central bank was not able to stabilize the inflation using only conventional measures. In the period of monetary policy

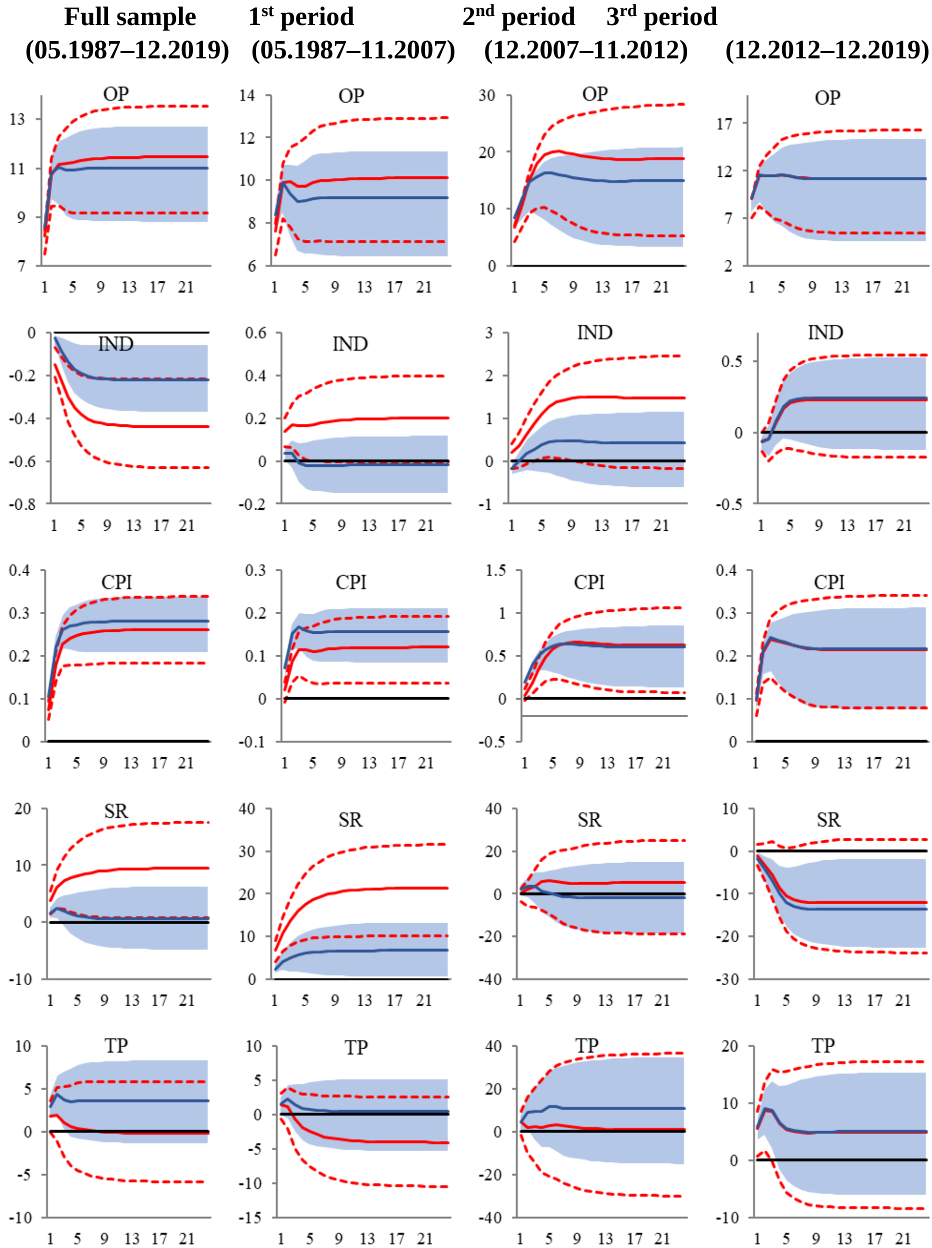
[10] A 'standardized unit' corresponds to the 1st difference of the data transformed in natural logarithm.

normalization, the long-run response of the shadow policy rate is still non-significant, while the short-run response turns to be significant but negative. Such a counter-intuitive response could have been caused by a relatively stable oil price level in the period under examination. In addition to the positive effect on prices, there is also a negative effect on the industrial production, to which the monetary authority could respond by lowering its policy rate. Therefore, the monetary policy trade-off between stabilizing inflation and stabilizing output can also be an explanation for the non-intuitive response, as purely stabilizing inflation might be costly.

Finally, with regard to the ***term premium***, we confirm that the immediate response of the term premium to the oil price shock is positive and significant. The significance vanishes around 3 periods after the shock. Such a result confirms that the term premium moves together with the interest rate, as it is its component. However, we document that the response of the term premium was barely significant in the first and the second sub-period. Its importance increases in the third period, which captures the normalization of monetary policy. Interestingly, the direction of the response is opposite to the response of the shadow policy rate in a corresponding period. Such a finding indicates that the term premium component outweighs the expectations component – indicating that the importance of term premium has grown in the recent period.

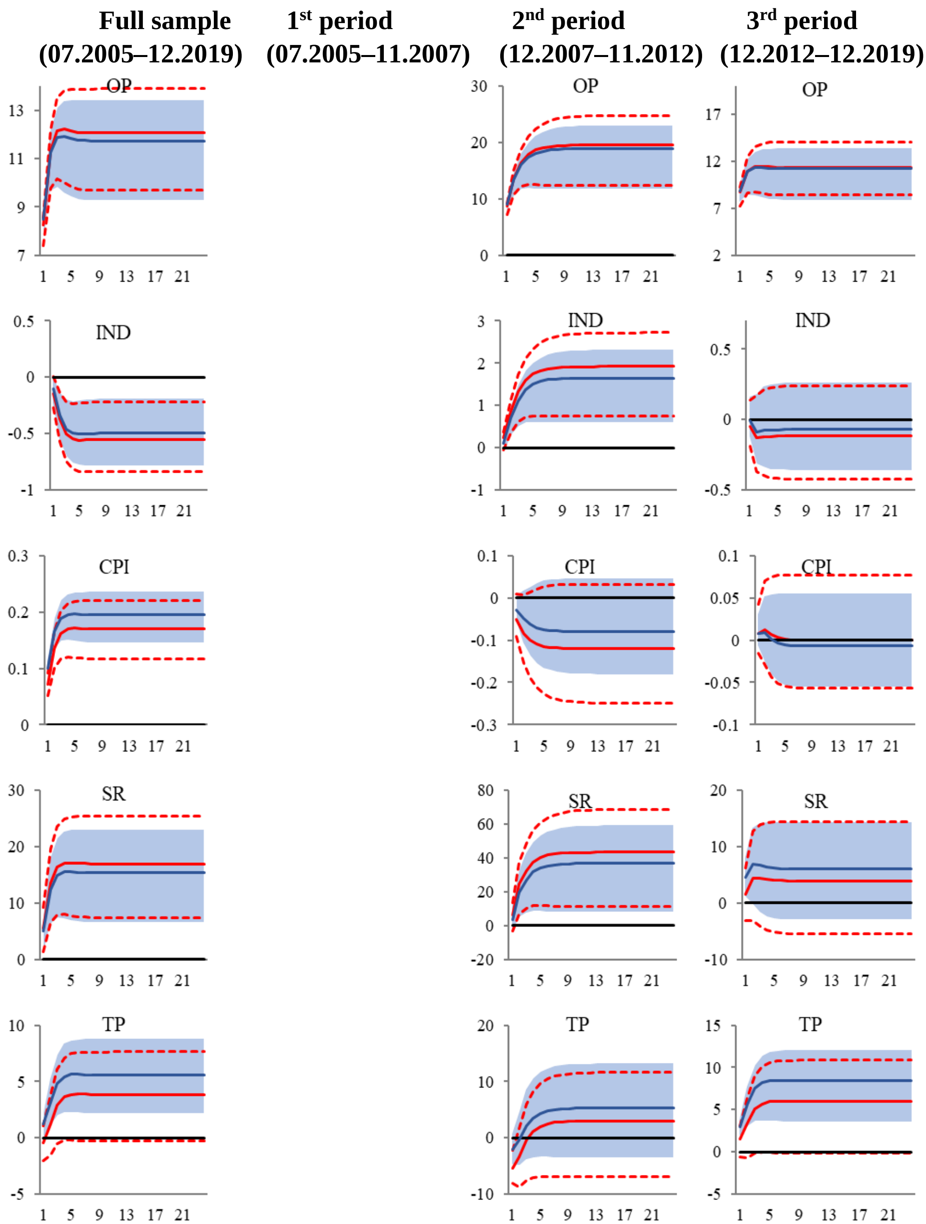
The results for the euro area are available in Figure 4, which shows the impulse responses to the oil price shock for all variables in each sub-period except the first one due to the lack of data (only 29 observations).

Figure 3. Short-run and long-run responses to the oil price shock in the US



Note: blue line – cumulative impulse response under Cholesky factorization; red line – cumulative impulse response under Blanchard-Quah long-run restriction; dashed lines and shaded area represent 68 percent confidence bands
Source: own calculations based on the Federal Reserve economic database

Figure 4. Short-run and long-run responses to the oil price shock in the euro area



Note: blue line – cumulative impulse response under Cholesky factorization; red line – cumulative impulse response under Blanchard-Quah long-run restriction; dashed lines and shaded area represent 68 percent confidence bands
Source: own calculations based on the Federal Reserve economic database and Eurostat

The response of *industrial production* in the euro area is generally consistent with that in the US. Compared to the results for the US, the response of industrial production to the oil price shock is positive and significant in the second period, which represents a period of unconventional monetary policy. Such a response is at odds with the economic theory and may reflect a deep distortion between oil price movements and industrial production. We admit that, in the given sub-period, other factors than oil prices (for example monetary or fiscal stimuli) could improve the response of industrial production and revert the effects of oil prices.

As for *inflation*, we document that the expected spill-overs from oil price changes to the overall prices apply in the euro area as well. However, in the period of unconventional monetary policy, we can clearly observe that the changes in oil prices are irrelevant for the CPI index. The same applies for the US. What is different, however, is the response in the last period of policy normalization. We have already mentioned that the third period does not approximate the normalization of monetary policy for the euro area. The response of CPI in the euro area during the third period confirms that the economy has not fully stabilized yet and as a consequence, the response of the price level is still not usual.

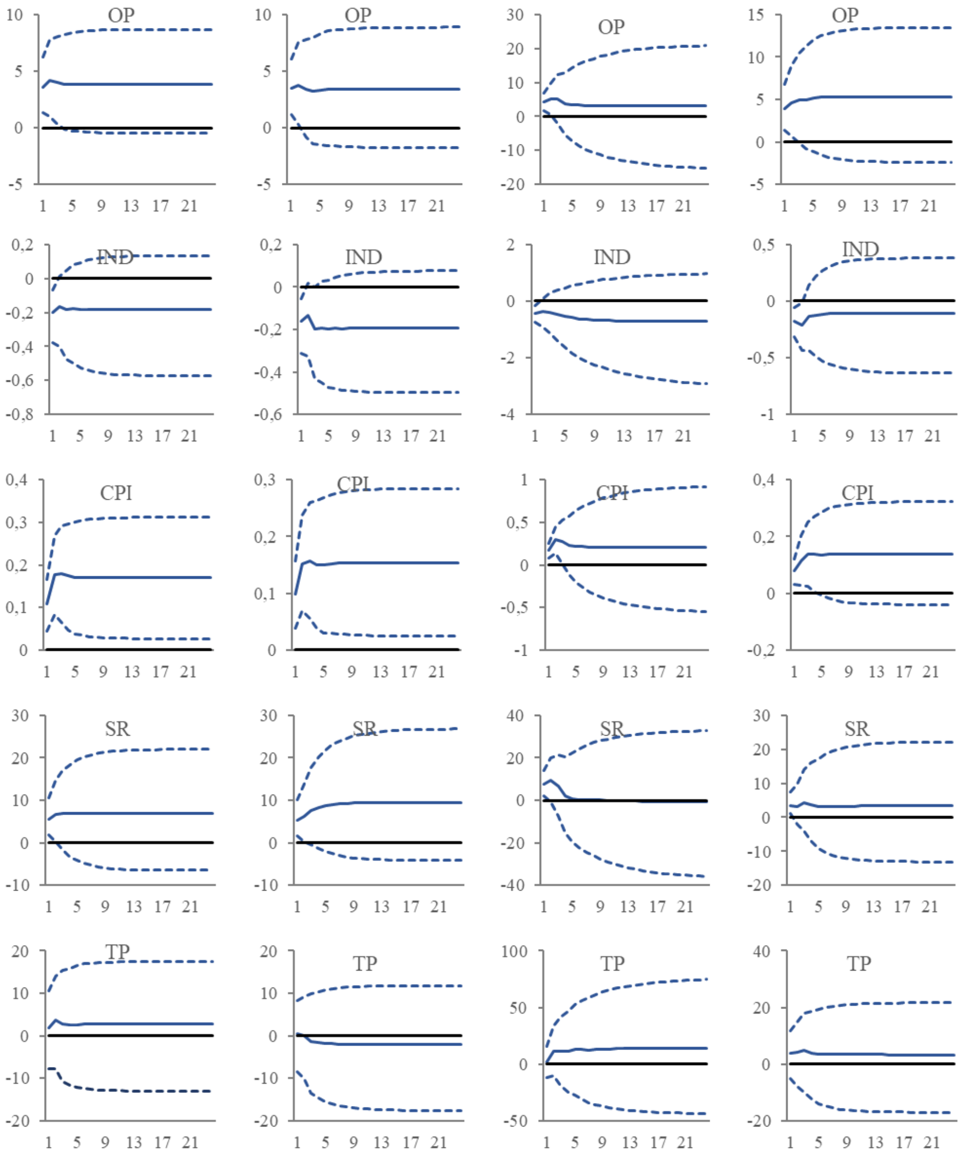
Additional evidence to support this assumption comes from the response of the *shadow policy rate*. While its response is in the full sample period positive and significant, it remains non-significant in the period of monetary normalization. This implies that the central bank does not need to intervene by changing its policy rate, as it is ineffective in the respective period. Finally, our results are, with respect to the *term premium* in the euro area, consistent with those in the US. The importance of the term premium has obviously increased over the last years.

4.2 Sign restrictions

In this section, we present the results of the Bayesian *VAR* model. We impose sign restrictions on the oil price shock as indicated in Table 1. The results of Bayesian *VAR* model cannot be used for assessing the oil price transmission channel, because the direction of the responses is already given by the restrictions. We can, however, make use of the results of the Bayesian *VAR* model in order to evaluate the changes in the magnitude of the responses between individual sub-periods. We present the results of our Bayesian *VAR* model defined in sub-section 3.3 in Figure 5 and Figure 6, respectively. Figure 5 shows the cumulative impulse response functions and their 68% confidence bands for the US.

Figure 5. Sign restricted responses to the oil price shock in the US

Full sample (05.1987–12.2019)	1st period (05.1987–11.2007)	2nd period (12.2007–11.2012)	3rd period (12.2012–12.2019)



Note: blue line – cumulative impulse response under sign restrictions; dashed lines – 68 percent confidence bands
Source: own calculations based on the Federal Reserve economic database

A comparison of the impulse responses of industrial production suggests that the immediate response of industrial production was the strongest (−0.44 standardized unit) in the second period, when the US Fed actively pursued unconventional monetary measures. We further observe that in the subsequent period of monetary policy normalization, the magnitude of initial response returned to −0.15 standardized unit. The results thus confirm that in the period of unconventional monetary policy in the aftermath of the Global Financial Crisis, the industrial production was more sensitive to the oil price turbulences and as a consequence, the economy was more vulnerable than in the other sub-periods. These results are consistent with our previous findings presented in Figure 3.

A similar pattern can be observed in the response of the *price level*. The magnitude of the initial responses in the third and fourth sub-periods are much higher compared to the first sub-period. Moreover, the response is relatively short-lived – it disappears approximately 3 months after the shock. These results are fully in accordance with the outcomes of Cholesky and Blanchard and Quah *VAR* model specifications.

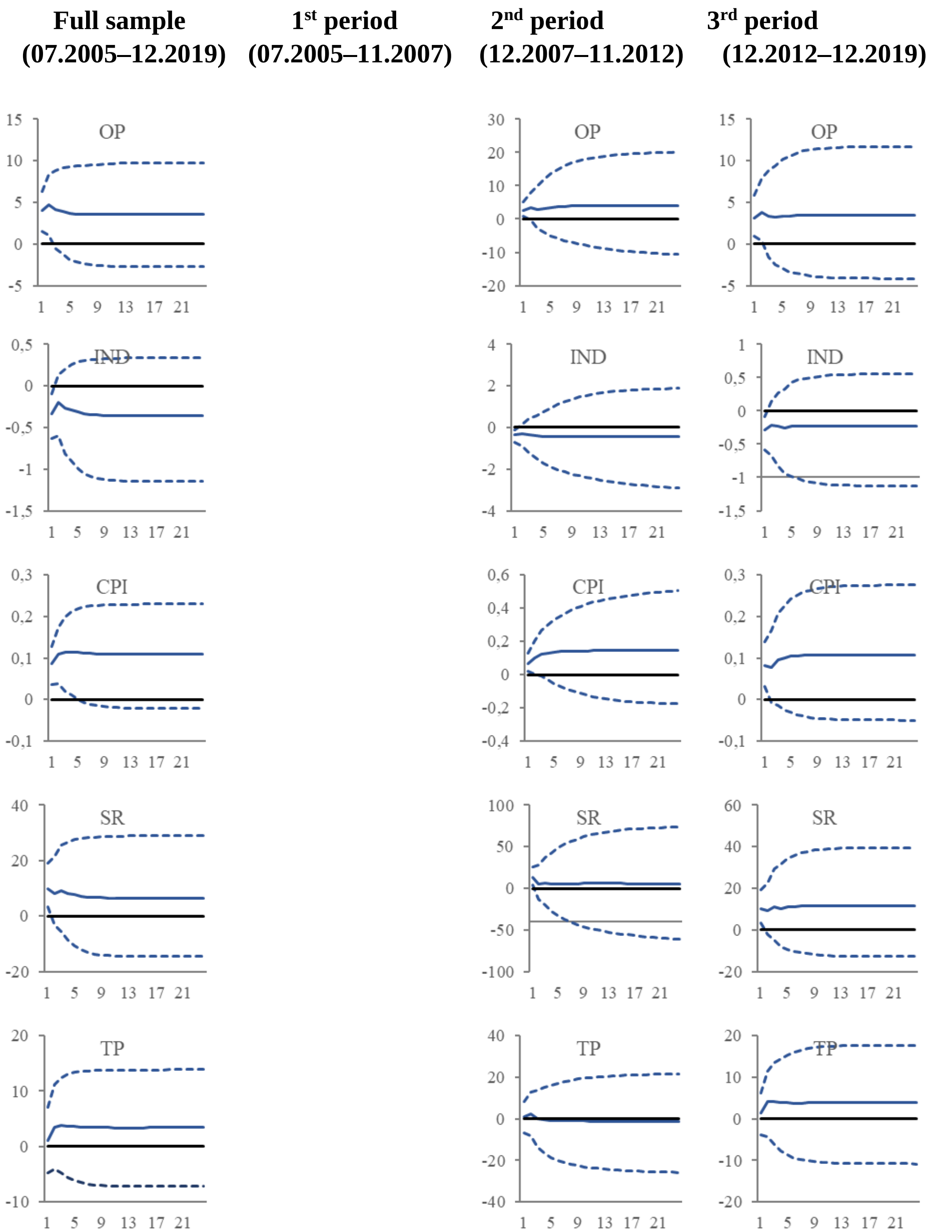
The Bayesian *VAR* model further confirms that in the period of conventional monetary policy, the *shadow policy rate* was an effective instrument of the central bank to stabilize inflation. In the period corresponding to the unconventional monetary policy, however, the initial response of shadow rate is somewhat smaller and returns to zero 5 months after the shock. Moreover, the significance of the response disappears soon, which implicates that the monetary policy (in particular conventional) was insufficient in the aftermath of the Global Financial Crisis.

Finally, the response of the *term premium* is non-significant for all the examined sub-sample periods. Even though that the initial response of the term premium in the second sub-period is somewhat higher, it fails to be significant. The findings of the Bayesian *VAR* model therefore do not confirm the hypothesis of growing importance of the term premium in the US economy in the period of unconventional monetary policy.

Turning to the euro area, the results of Bayesian *VAR* model are shown in Figure 5. It was not possible to estimate the *VAR* model for the first sub-period due to the lack of data, as we have only 29 observations. The results for industrial production, price level and shadow interest rate confirmed the previous results. The magnitude of the response is strongest in the period of unconventional monetary policy. In the recent period, the initial response is somewhat weaker, but still stronger than average in the full sample period. The results thus confirm that in the euro area, the conventional tools are in the third sub-period still not powerful enough.

For a better comparison of our results from the three different *VAR* specifications (i.e., with short-term, long-term and sign restrictions), in Appendix F, we present the response functions (IRFs) from all three specifications together in one graph without confidence bounds both for the US and the euro area economy.

Figure 6. Sign restricted responses to the oil price shock in the euro area



Note: blue line – cumulative impulse response under sign restrictions; dashed lines – 68 percent confidence bands
Source: own calculations based on the Federal Reserve economic database and Eurostat

4. Conclusion

Using monthly data, we examined the consequences of oil price changes for the economy of the US and the euro area. To verify the transmission channel through which the oil price shock affects the economy, we used the unidirectional Granger causality test supplemented with the interpretation of the impulse response functions of several *VAR* specifications. We conclude that none of the applied methods and the model specifications contradict each other. On the contrary, they complement each other and create a complete mosaic, which can be viewed as a proof of the robustness of our results.

Overall, we confirmed the assumption given by our first hypothesis that a negative oil price shock leads to a decrease in production and to an increase in the overall price level, which motivates the central banks in the US and in the euro area to decrease their policy rates. Such result is in accordance with the findings of other authors who performed similar analyses (see for example Peersman, 2011 or Bundick, Herriford and Smith, 2017). However, we have identified some deviations from the transmission channel in the individual sub-periods. For example, we have confirmed that it became much more difficult to stabilize the price level in the aftermath of the Global Financial Crisis – both in the US and in the euro area. A new aspect in our study is the fact that we have extended the traditional oil price transmission channel by the effects on the treasury term premium in the US and in the euro area. Most of the papers dealing with the term premium examines its proper calculation and factors that influence its dynamics. We rather examined the position of the term premium in the oil price transmission channel and its change over time. Our findings suggest that the importance of the term premium grew substantially in times of unconventional monetary policy, and its impact on treasury yield may have even outweighed the signal from the expectations component.

With regard to our second hypothesis, we conclude that the consequences of changes in the oil prices have significantly grown since the introduction of unconventional monetary instruments. This is clear from the results of our Bayesian *VAR* models and confirmed by the results of the *VAR* model under short-run and long-run restrictions. The magnitude of the response of industrial production, the price level and shadow interest rate to the oil price shock is strongest in the period of unconventional monetary policy. In many cases, however, the reaction is short-lived. Not surprisingly, while the response magnitude for most variables in the US returned to previous levels during the third sub-period of monetary normalization, in the case of the euro area, we observe that the response of most variables remained overstated. The same conclusions come from the responses of term premia. The results thus confirm that the conventional instruments in the euro area are still not sufficient to stabilize the economy in the third sub-period. Based on our result, we agree with Wanke (2017), De Vijlder (2018) or Vannelli (2018) that the oil price shock is not negligible for several economic variables even in times of monetary policy normalization.

We conclude that imbalances and vulnerabilities may accumulate during the use of unconventional monetary instruments, as conventional instruments have become ineffective. While the US central bank managed to raise its key interest rates before the corona crisis outbreak, the ECB still had zero interest rates and was forced to restart quantitative easing. Given the very low level of term premia, it would be more difficult to stabilize the economy in the event of a future economic recession. In the light of these conclusions, we see the use of unconventional monetary policy instruments in both the US and the euro area as effective for stabilizing the economy, although we acknowledge that they may also cause some imbalances. In an event of a large-scale negative shock, we suggest considering the adoption of unconventional monetary instruments again. Depending on the extent of the shock, it is also possible to consider extending the measures used so far in a combination with an adequate use of fiscal stimulus.

Acknowledgement: I thank the participants of the SIdE Summer School: Principles, ideas and theory in econometric time series, which took place from 16 to 22 June 2019 in Bertinoro, Italy, and was organized by Società Italiana di Econometria (SIdE). I also benefited from attending the Barcelona Graduate School of Economics (BGSE) - Summer School 2019. I also thank two anonymous reviewers for their comments and feedback. The article presents author's calculations and views. All errors remain author's.

Disclosure statement: No potential conflict of interest was reported by the author.

References

ABBRITTI, M., DELL´ERBA, S., MORENO, A., (2018). Global Factors in the Term Structure of Interest Rates. *International Journal of Central Banking*, Vol. 14, No. 2, March 2018.

ADRIAN, T., CRUMP, R.K., MOENCH, E., (2013). Pricing the Term Structure with Linear Regressions. *Federal Reserve Bank of New York*, Working paper, DOI: http://dx.doi.org/10.1016/j.jfineco.2013.04.009

ADRIAN, T., CRUMP, R.K., MOENCH. E., (2014). Regression-based Estimation of Dynamic Asset Pricing Model." *Journal of Financial Economics*, Vol. 110, Issue 1, pp. 110-138, DOI: http://dx.doi.org/10.1016/j.jfineco.2015.07.004

AKRAM, Q.F., MUMTAZ, H., (2016). The Role of Oil Prices and Monetary Policy in the Norwegian Economy since the 1980s. *Norges Bank Research*, Working paper 1/2016.

BAŃBURA, M., GIANNONE, D., REICHLIN, L., (2008). Large Bayesian VARs. *European Central Bank Working Paper Series*, No. 966/November 2008.

BLANCHARD, O.J., QUAH, D., (1989). The Dynamic Effects of Aggregate Demand and Supply Disturbances. *The American Economic Review*, Vol. 79, No. 4, (Sep., 1989), pp. 655-673.

BORIO, C., ZABAI, A., (2016). Unconventional Monetary Policies: a Re-appraisal. *Bank for International Settlements BIS Working Papers*, No. 570, July 2016.

BUNDICK, B., HERRIFORD, T., SMITH, A.L., (2017). Forward Guidance, Monetary Policy Uncertainty, and the Term Premium. *Federal Reserve Bank of Kansas City,* working paper No. 17-07, July, DOI: http://doi.org/10.18651/RWP2017-07

CARBONI, G. (2014). Term Premia Implications of Macroeconomic Regime Changes. *European Central Bank Working Paper Series*, No. 1694/July 2014, DOI: http://dx.doi.org/ 10.1017/S1365100514000352

COENEN, G., EHRMANN, M., GABALLO, G., HOFFMANN, P., NAKOV, A., NARDELLI, S., PERSSON, E., STRASSER, G., (2017). Communication of Monetary Policy in Unconventional Times. *European Central Bank Discussion Papers*, No. 2080/June 2017, DOI: http://dx.doi.org/10.2139/ssrn.3043098

COHEN, B.H., HÖRDAHL, P., XIA, D., (2018). Term Premia: Models and Some Stylised Facts. *Bank for International Settlements BIS Quarterly Review*, September 2018.

DE VIJLDER, W., (2018). The Decline of Commodity Prices: A Matter of Concern? *BNP Paribas EcoWeek*, No. 18-45, 7 December 2018.

ELDER, J., SERLETIS, A., (2010). Oil Price Uncertainty. *Journal of Money, Credit and Banking*, Vol. 42, Issue 6, 1137-1159, DOI: https://doi.org/10.1111/j.1538-4616.2010.00323.x

ENDERS, W. (2010). Applied Econometric Time Series. 3rd Edition., *John Wiley & Sons*, Hoboken.

FORNERO, J., KIRCHNER, M., (2018). Learning about Commodity Cycles and Saving-Investment Dynamics in a Commodity-Exporting Economy. *International Journal of Central Banking*, Vol. 14, No. 2, March 2018.

GAGNON, J., RASKIN, M., REMACHE J., SACK, B., (2011). The Financial Market Effects of the Federal Reserve's Large-Scale Asset Purchases, *International Journal of Central Banking*, Volume 7, Issue 1, pp.3-43.

GAMBACORTA, L., HOFMANN, B., PEERSMAN, G., (2012). The Effectiveness of Unconventional Monetary Policy at the Zero Lower Bound: A Cross-Country Analysis. *Bank for International Settlements BIS Working Papers*, No. 384, August 2012, available at http://dx.doi.org/10.1111/jmcb.12119

GLICK, R., LEDUC, S., (2018). Unconventional Monetary Policy and the Dollar: Conventional Signs, Unconventional Magnitudes. *International Journal of Central Banking*, Vol. 14, No.5.

GREENSPAN, A., (2005). Testimony of Chairman Alan Greenspan. *Federal Reserve Board's semiannual Monetary Policy Report to the Congress*, February 16, 2005.

HARTWELL, Ch. A., (2018). On the impossibility of central bank independence: four decades of time- (and intellectual) inconsistency. *Cambridge Journal of Economics*. Vol. 43, Issue 1, pp. 61-84, DOI: https://doi.org/10.1093/cje/bex083

HUBERT, P., LABONDANCE, F., (2018). The Effect of ECB Forward Guidance on the Term Structure of Interest Rates. *International Journal of Central Banking*, Vol. 14, No. 5, December 2018.

CHOI, S., FURCERI, D., LOUNGANI, P., MISHRA, S., POPLAWSKI-RIBEIRO, M., (2017). Oil Prices and Inflation Dynamics: Evidence from Advanced and Developing Economies. *IMF Working Paper*,WP/17/196, DOI: http://dx.doi.org/10.1016/j.jimonfin.2017.12.004

KIM, D.H., WRIGHT, H., (2005). An Arbitrage-Free Three-Factor Term Structure Model and the Recent Behavior of Long-Term Yields and Distant-Horizon Forward

Rates. *Finance and Economics Discussion Series*, 2005-33, DOI: http://dx.doi.org/10.2139/ssrn.813267

KRIPPNER, L., (2012). Modifying Gaussian Term Structure Models when Interest Rates are Near the Zero Lower Bound. GAMA Woking Paper No. 2012-05.

LANNE, M., LUOTO, J., (2019). Identification of Economic Shocks by InequalityConstraints in Bayesian Structural Vector Autoregression. *Oxford Bulletin of Economics and Statistics*, 0305–9049, DOI: https://doi.org/10.1111/obes.12338

LIPPI, F., NOBILI, A. (2008). Oil and the Macroeconomy: A Structural VAR Analysis with Sign Restrictions. *CEPR Discussion Paper* No. DP6830, May 2008, available at SSRN: https://ssrn.com/abstract=1143193

LOMBARDI, M.J., ZHU, F., (2018). A Shadow Policy Rate to Calibrate U.S. Monetary Policy at the Zero Lower Bound. *International Journal of Central Banking*. Vol. 14, No. 5, December 2018.

MELOLINNA, M., (2012). Macroeconomic Shocks in an Oil Market VAR. *ECB Working Paper* No. 1432, May 3, 2012, available at SSRN: https://ssrn.com/abstract=2050326

MOUABBI, S., SAHUC, J. G., (2019). Evaluating the Macroeconomic Effects of the ECB's Unconventional Monetary Policies. *Journal of Money, Credit and Banking*, Vol. 51, Issue 4, 831-858, DOI: https://doi.org/10.1111/jmcb.12628

PAŽICKÝ, M., (2018), The consequences of unconventional monetary policy in euro area in times of monetary easing. *Oeconomia Copernicana*, Vol.9, Issue 4, DOI: https://doi.org/10.24136/oc.2018.029

PAŽICKÝ, M., (2019). Interest rate effectiveness in era of quantitative easing in euro market. *International Journal of Accounting and Finance*, Vol.8, Issue 4. DOI: https://doi.org/10.1504/IJAF.2018.098341

PEERSMAN, G., (2011). Macroeconomic Effects of Unconventional Monetary Policy in the Euro Area. *European Central Bank Working Paper Series, No. 1397/November 2011*.

RONAYNE, D., (2011). Which Impulse Response Function? *The University of Warwick. Warwick Economic Research Papers*, No. 971, October 2011.

RUBIO-RAMIREZ, J.F., WAGGONER, D., ZHA, T., (2010). Structural Vector Autoregressions: Theory of Identification and Algorithms for Inference. *Review of Economic Studies*, 2010, Vol. 77, issue 2, 665-696, DOI: http://dx.doi.org/10.1111/j.1467-937X.2009.00578.x

RUDEBUSCH, G.D., SACK, B. P., SWANSON, E.T., (2007). Macroeconomic Implications of Changes in the Term Premium. *Federal Reserve Bank of St. Louis Review*, July/August 2007, 89(4), pp. 241-69, DOI: http://dx.doi.org/10.20955/r.89.241-270

RUDEBUSCH, G.D., WU. T., (2004). A Macro-Finance Model of the Term Structure, Monetary Policy, and the Economy. *Federal Reserve Bank of San Francisco Working Paper Series*, Working paper 2003-17, DOI: http://dx.doi.org/10.1111/j.1468-0297.2008.02155.x

SEKINE, A., TSURUGA T., (2018). Effects of commodity price shocks on inflation: a cross-country analysis. *Oxford Economic Papers*, Volume 70, Issue 4, October 2018, pp. 1108–1135, DOI: https://doi.org/10.1093/oep/gpy015

SEGAL, P., (2007). Why Do Oil Price Shocks No Longer Shock? *Oxford Institute for Energy Studies*, WPM 35, October 2007.

SUSSMAN, N., ZOHAR, O., (2015). Oil Prices, Inflation Expectations, and Monetary Policy. *Bank of Israel*, 16 September 2015.

VANNELLI, S., (2018). Thoughts on the Term Premium. Knowledge Leaders Capital, blog contribution from September 7, 2018.

WANKE, S., (2017). Oil Prices and Bond Yields – Hand-in-Hand Again. *KfW Research Economics in Brief*, No. 140, 27 July 2017.

WOODFORD, M., (2012). Methods of Policy Accommodation at the Interest-Rate Lower Bound. Columbia University, September 16, 2012.

WU, J.C., XIA, F.D., (2016). Measuring the Macroeconomic Impact of Monetary Policy at the Zero Lower Bound, *Journal of Money, Credit, and Banking*, 48(2-3), 253-291, 2016.

Appendix A. Data characteristics

Table A.1: Statistical characteristics

		US						Euro area				
		TP	*IND*	*CPI*	*OP*	*SR*		*TP*	*IND*	*CPI*	*OP*	*SR*
Num.	**Full sample (05.1987 - 12.2019)**	392	392	392	392	392	**Full sample (07.2005 - 12.2019)**	174	174	174	174	174
Avg.		1,3	88,9	79,1	46,5	3,2		1,3	100,6	94,9	77,3	-0,7
Std.		1,0	15,5	17,5	32,7	3,0		1,0	4,9	6,1	24,9	3,0
Max		3,5	110,6	109,0	132,7	9,9		3,1	110,1	104,3	132,7	4,3
Min		-0,9	59,3	47,7	9,8	-3,0		-0,7	86,6	82,9	30,7	-5,5
Skew		0,1	-0,6	0,0	0,9	-0,1		-0,1	-0,4	-0,4	0,4	0,0
Kurt		-0,9	-1,1	-1,2	-0,5	-0,8		-1,3	0,0	-1,1	-1,1	-1,1
Num.	**1.period (conventional) (05.1987 - 11.2007)**	247	247	247	247	247	**1.period (conventional) (07.2005 - 11.2007)**	29	29	29	29	29
Avg.		1,5	81,6	67,8	27,1	4,9		2,3	104,1	85,1	66,1	3,1
Std.		0,9	14,7	11,0	16,4	2,1		0,3	3,0	1,3	8,9	0,7
Max		3,5	105,3	89,0	92,4	9,9		2,8	108,7	87,0	92,4	4,0
Min		-0,1	59,3	47,7	9,8	1,0		1,8	98,4	82,9	53,7	2,0
Skew		0,1	0,0	0,0	1,8	0,0		0,5	-0,2	0,0	1,1	-0,3
Kurt		-1,0	-1,6	-1,0	2,4	-0,4		-0,7	-1,1	-1,4	1,4	-1,6
Num.	**2.period (unconventional) (12.2007 - 11.2012)**	60	60	60	60	60	**2.period (unconventional) (12.2007 - 11.2012)**	60	60	60	60	60
Avg.		1,7	96,3	92,9	91,8	-0,1		2,1	97,8	91,9	91,8	0,9
Std.		0,8	4,8	2,6	24,3	1,4		0,6	5,8	2,5	24,3	1,6
Max		3,0	105,3	97,7	132,7	4,2		3,1	110,1	96,7	132,7	4,3
Min		0,2	87,1	89,2	40,0	-1,5		1,0	86,6	87,2	40,0	-1,1
Skew		-0,4	-0,1	0,4	-0,4	1,4		-0,2	0,3	-0,1	-0,4	1,1
Kurt		-0,9	-0,7	-1,2	-0,8	1,3		-0,7	-0,2	-0,5	-0,8	0,0
Num.	**3.period (normalization) (12.2012 - 12.2019)**	85	85	85	85	85	**3.period (normalization) (12.2012 - 12.2019)**	85	85	85	85	85
Avg.		0,5	105,1	102,3	70,9	0,6		0,3	101,3	100,4	70,9	-3,2
Std.		0,7	2,9	3,3	24,4	2,3		0,5	3,5	1,7	24,4	1,8
Max		2,1	110,6	109,0	116,1	3,8		1,5	108,1	104,3	116,1	0,5
Min		-0,9	100,8	97,6	30,7	-3,0		-0,7	95,1	96,8	30,7	-5,5
Skew		0,3	0,5	0,5	0,6	-0,2		0,4	-0,1	0,5	0,6	0,5
Kurt		-0,6	-1,1	-1,1	-1,0	-1,6		-0,3	-1,2	0,0	-1,0	-1,3

Source: own calculations based on Federal Reserve economic database and Eurostat

Table A.2. Results of Augmented Dickey-Fuller test for unit root

US

Variable	*TP*		*ΔTP*	*IND*		*ΔIND*	*CPI*		*ΔCPI*	*OP*		*ΔOP*	*SR*		*ΔSR*
	no trend	**trend**		**no trend**	**trend**		**no trend**	**trend**		**no trend**	**trend**		**no trend**	**trend**	
Num. obs.	388	388	387	387	387	387	390	390	389	390	390	389	388	388	388
Stat	-2.25	-2.74	-8.68	-1.31	-2.22	-5.37	-2.87	-3.38	-13.19	-2.02	-2.97	-11.22	-1.85	-1.89	-6.34
CV	-3.45	-3.98	-3.45	-3.45	-3.99	-3.45	-3.45	-3.98	-3.45	-3.45	-3.98	-3.45	-3.45	-3.98	-3.45
p value	*(0.19)*	*(0.22)*	*(0.00)*	*(0.62)*	*(0.48)*	*(0.00)*	*(0.05)*	*(0.05)*	*(0.00)*	*(0.28)*	*(0.14)*	*(0.00)*	*(0.35)*	*(0.66)*	*(0.00)*
Num. lags	3	3	3	4	4	3	1	1	1	1	1	1	3	3	3
Integration	**I(1)**			**I(1)**			**I(1)**			**I(1)**			**I(1)**		
Conclusion	**non-stationary**			**non-stationary**			**non-stationary**			**non-stationary**			**non-stationary**		

Euro area

Variable	*TP*		*ΔTP*	*IND*		*ΔIND*	*CPI*		*ΔCPI*	*OP*		*ΔOP*	*SR*		*ΔSR*
	no trend	**trend**		**no trend**	**trend**		**no trend**	**trend**		**no trend**	**trend**		**no trend**	**trend**	
Num. obs.	172	172	171	170	170	170	172	172	171	172	172	171	171	171	171
Stat	-0.52	-2.91	-8.58	-2.63	-2.64	-4.65	-2.04	-2.19	-6.67	-2.56	-2.65	-7.21	-0.48	-2.92	-7.96
CV	-3.49	-4.02	-3.49	-3.49	-4.02	-3.49	-3.49	-4.02	-3.49	-3.49	-4.02	-3.49	-3.49	-4.02	-3.49
p value	*(0.89)*	*(0.16)*	*(0.00)*	*(0.09)*	*(0.26)*	*(0.00)*	*(0.27)*	*(0.50)*	*(0.00)*	*(0.10)*	*(0.26)*	*(0.00)*	*(0.90)*	*(0.16)*	*(0.00)*
Num. lags	1	1	1	3	3	2	1	1	1	1	1	1	2	2	1
Integration	**I(1)**			**I(1)**			**I(1)**			**I(1)**			**I(1)**		
Conclusion	**non-stationary**			**non-stationary**			**non-stationary**			**non-stationary**			**non-stationary**		

Note: Δ represents 1st difference transformation; p values are in parenthesis; CV stands for critical value; Augmented Dickey–Fuller procedure tests whether a variable follows the unit-root process; the null hypothesis states that the variable contains a unit root against alternative that the variable is generated by a stationary process; lag selection was performed based on SIC (Schwarz information criterion); conclusions are based on 0.01 level of significance; testing procedure including drift is not presented.

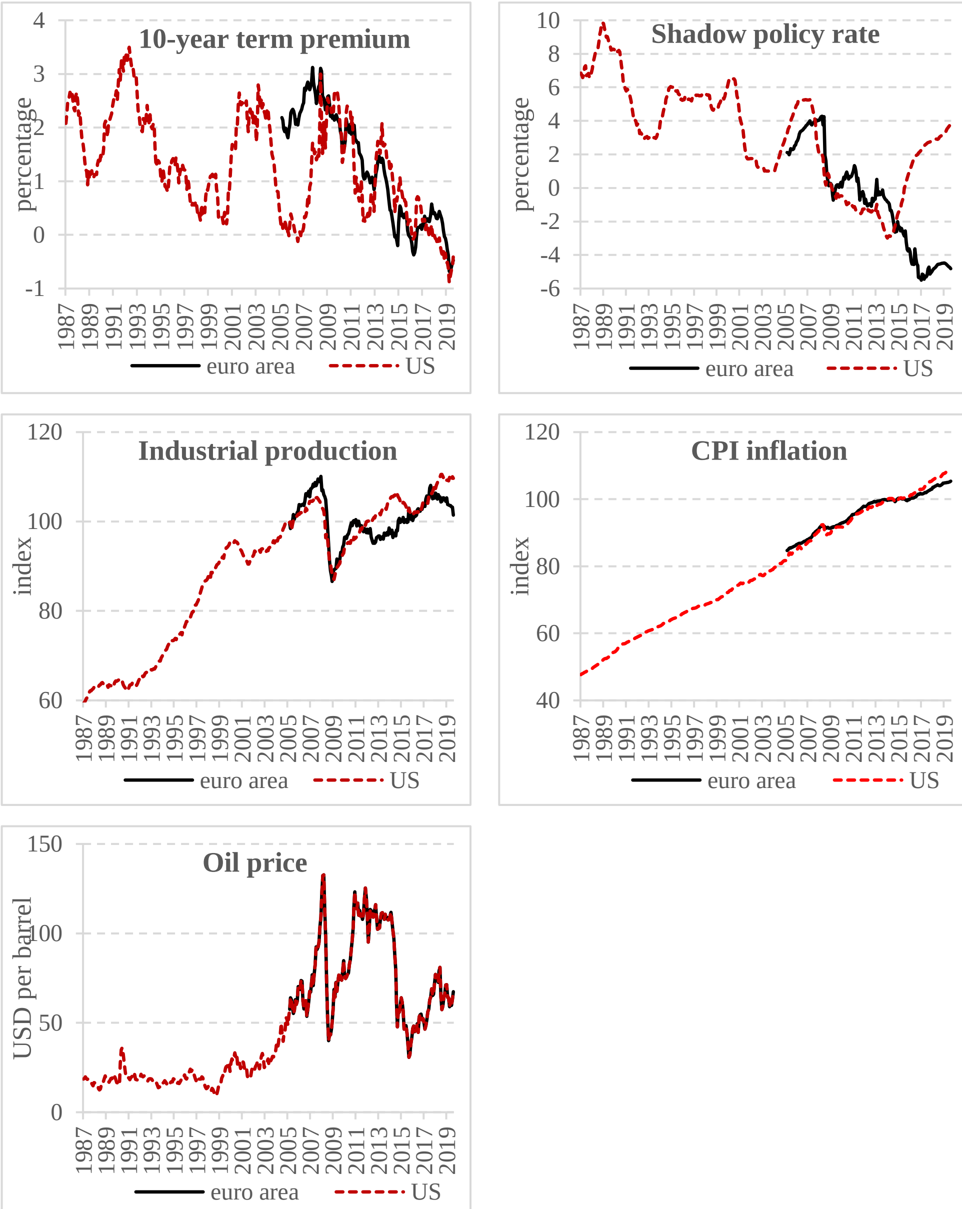
Source: own calculations based on Federal Reserve economic database and Eurostat

Table A.3. Correlations

		US data					Euro area data				
		TP	*IND*	*CPI*	*OP*	*SR*	*TP*	*IND*	*CPI*	*OP*	*SR*
FULL SAMPLE	*TP*	1					1				
	IND	-0,61	1				-0,12	1			
	CPI	-0,52	0,92	1			-0,89	-0,01	1		
	OP	-0,26	0,67	0,77	1		0,33	-0,12	-0,10	1	
	SR	-0,05	-0,62	-0,75	-0,70	1	0,91	-0,01	-0,91	0,37	1
1st PERIOD	*TP*	1					1				
	IND	-0,58	1				0,83	1			
	CPI	-0,49	0,96	1			0,83	0,97	1		
	OP	-0,47	0,68	0,75	1		0,75	0,56	0,56	1	
	SR	-0,30	-0,50	-0,60	-0,21	1	0,77	0,96	0,98	0,47	1
2nd PERIOD	*TP*	1					1				
	IND	-0,64	1				0,23	1			
	CPI	-0,66	0,36	1			-0,91	-0,40	1		
	OP	-0,52	0,74	0,69	1		-0,38	0,64	0,29	1	
	SR	0,17	0,32	-0,70	-0,13	1	0,75	0,75	-0,86	0,15	1
3rd PERIOD	*TP*	1					1				
	IND	-0,59	1				-0,57	1			
	CPI	-0,83	0,84	1			-0,70	0,75	1		
	OP	0,60	-0,10	-0,37	1		0,78	-0,60	-0,42	1	
	SR	-0,88	0,54	0,89	-0,56	1	0,74	-0,91	-0,72	0,74	1

Note: data were transformed in 1st difference, since they are I(1) type
Source: own calculations based on Federal Reserve economic database and Eurostat

Figure A.1. 10-year Eonia decomposition



Source: own prepared based on Federal Reserve economic database and Eurostat

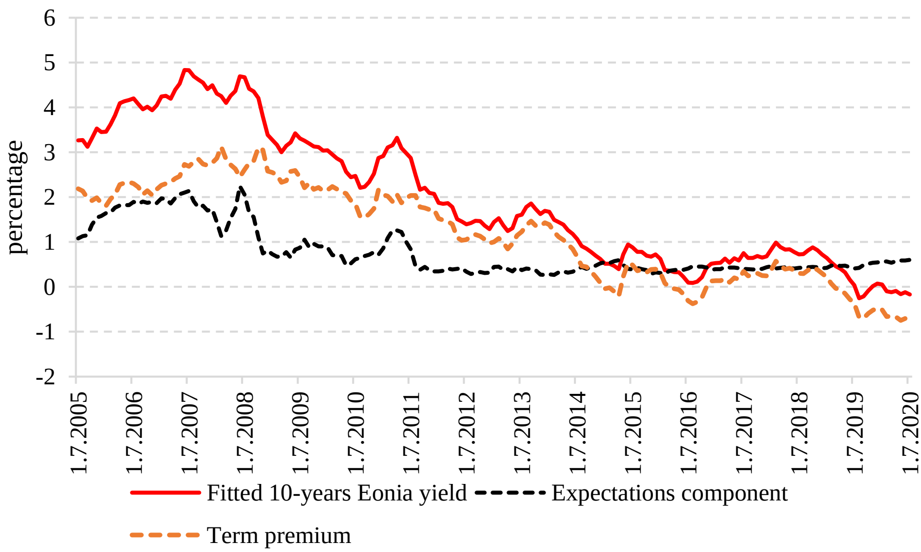
Appendix B. Term premium estimation in the euro area

We made use of the Overnight Index Swap (OIS) yields to estimate term premia of the euro area. We work with monthly yields of Eonia curve from July 2005 to October 2018 with maturities from one to ten years and one-month maturity as an approximation of the risk-free rate. We estimate the model developed by Adrian, Crump and Moench (2013, 2014), the so-called ACM model. The model parameters are estimated by a three-step regression method and application of the OLS estimators.

First, we compute the pricing factors by the principal components of the yields. However, we apply only the first three principal components, as they satisfactorily explain the variation in the yields. In their original work, Adrian, Crump and Moench (2013) applied the first five principal components in the decomposition of the US Treasury yields. We use the Principal Component Analysis (PCA) applied to the data, adjusted by the means and volatility to extract the principal components from the term structure of OIS. Cumulatively the first three principal components explain around 99.97% of the overall variance in the term structure of OIS. For example, Abbritti et al. (2018) used *FAVAR* model to show that global factors account for more than 80% of term premia in advanced economies. For the rest of the estimation, we replicate the ACM model developed by Adrian, Crump and Moench (2013, 2014) using our dataset.

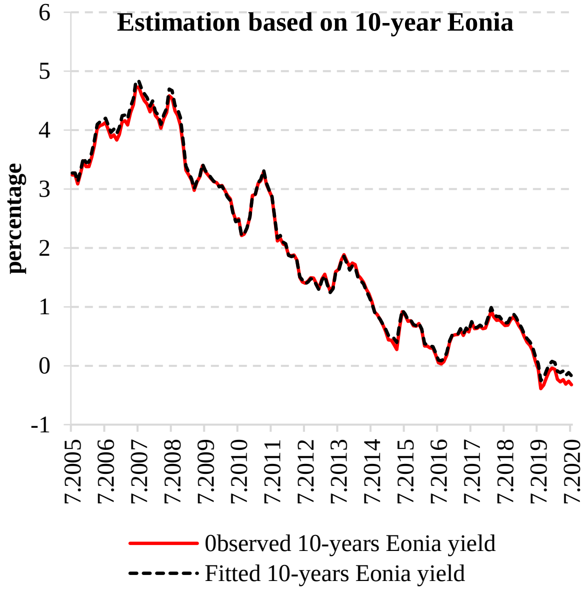
For the purpose of our study, we take into consideration the term premium only for 10-year maturity because it captures the dynamics of all shorter maturity term premia. The decomposition of the yields on 10-year OIS into average short rate expectations and 10-year term premium is shown in Figure (B.1). The performance of our model can be assessed based on Figure (B.2), which compares the observed 10-year OIS curve with the 10-year OIS curve fitted by our model.

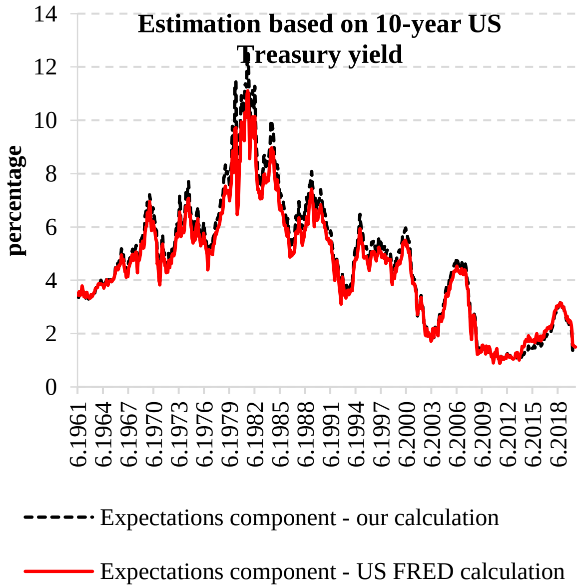
Figure B.2. 10-year Eonia decomposition



Source: own calculations based on ECB data

Figure B.3. ACM model performance





Source: own calculations based on ECB data and Federal Reserve economic database

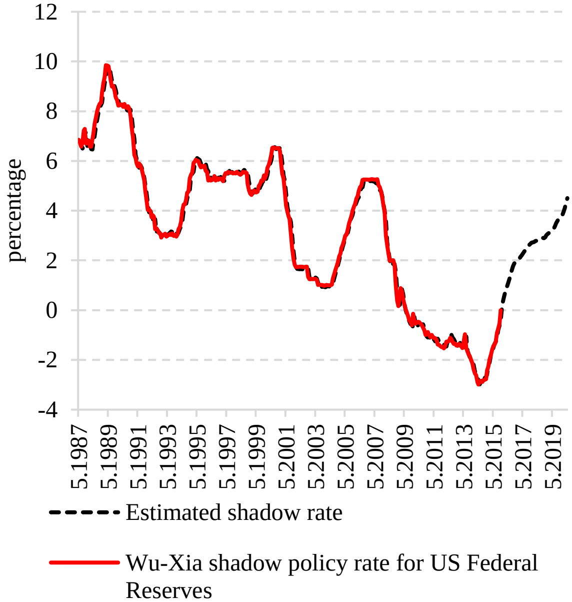
Appendix C. Shadow policy rate prediction

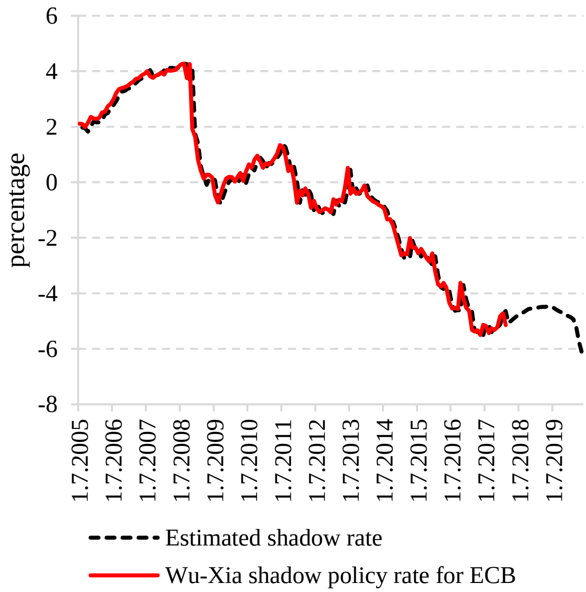
The data of shadow policy rate for both euro area and the US are already available in Wu and Xia (2016). Unfortunately, their dataset does not cover the entire period of our interest. We overcome this lack of data by estimating our own shadow policy rates for both economies using Wu and Xia (2016) dataset as dependent variables, while the data applied in our VAR models are used as explanatory variables. Estimated coefficients are then used for prediction of missing data of shadow rates. The regression model for both economies has a following form:

$$SR_t = \beta_0 + \beta_1 SR_{t-1} + \beta_2 TP_t + \beta_3 IND_t + \beta_4 CPI_t + \beta_5 OP_t + \varepsilon_t \quad \text{(C.1)}$$

where Wu and Xia (2016) shadow rate SR is estimated using its own lagged values (reflecting the persistence in the data and hence ensuring the smooth monetary policy over time), term premium TP, industrial production IND, price index CPI and oil prices OP. An error term is captured by ε_t and parameters β_0, β_1, β_2, β_3, β_4, β_5 are regression coefficients used to estimate missing shadow rate values. Goodness of fit is decent given that the R^2 is higher than 0.9 in both economies. Regression results are available upon request. Our predicted shadow policy rates compared with the original shadow rates of Wu and Xia (2016) are shown in Figure C.1.

Figure C.1: Estimated shadow rate for the US and Euro area





Source: Wu and Xia (2016) data and own calculations based on Federal Reserve economic database and Eurostat

Appendix D. Derivation of used VAR models

In our specification we include 2 lags based on the Bayesian information criterion (*BIC*), meaning that we estimate the following *VAR(3)* process:

$$Y_t = A_1 Y_{t-1} + A_2 Y_{t-2} + A_3 Y_{t-3} + \varepsilon_t \qquad (1)$$

Parameter ε_t represents an error term $\varepsilon_t \sim WN(0, \Omega)$, matrices A_1, A_2 and A_3 are the matrices whose elements we attempt to estimate and term Y_t denotes a following vector of variables:

$$Y_t = (\Delta \log OP, \Delta \log IND, \Delta \log CPI, \Delta SR, \Delta TP) \qquad (2)$$

where OP represents the oil prices, IND stands for industrial production, CPI represents inflation SR is a shadow policy rate and TP is a 10-year term premium. The log operator indicates a natural logarithm and Δ represents the transformation of the time series in first differences. The proposed order reflects a usual channel for the transmission of oil prices (see for example Lippi and Nobili, 2008 or Melolinna, 2012).

The process in Equation (1) can be expressed in a more compact form, abstracting from the subscript t. This will be useful for the *OLS* estimation of the matrices A_1, A_2 and A_3. We can rewrite the process in Equation (1) as *SUR* representation as follows:

$$\mathrm{Y} = \mathrm{X}\Pi + u \qquad (3)$$

where

$$\mathrm{Y} = (\mathrm{Y}_{p+1}, \dots, \mathrm{Y}_T) \qquad (4)$$

and

$$\mathrm{X} = (\mathrm{Y}_p, \dots, \mathrm{Y}_{T-p}) \qquad (5)$$

and

$$u = \left(\varepsilon_{p+1}, \dots, \varepsilon_T\right) \qquad (6)$$

and

$$\Pi = (C, A_1, A_2, A_3)' \qquad (7)$$

where subscript p denotes the number of lags ($p = 3$); subscript T denotes the number of observations ($T = 392$ for the US and $T = 174$ for the euro area) and parameter C is a vector of constants. The estimation of a *VAR* model by *OLS* reduces to the standard formula:

$$\widehat{\Pi} = (\mathrm{X}'\mathrm{X})^{-1}\mathrm{X}\mathrm{Y} \qquad (8)$$

If the conditions for stationarity of included variables are satisfied, then the series $\sum_{j=0}^{\infty} A^J$ converges and parameter Y_t has a *VMA*(∞) representation in the terms of the Wold shock ε_t, given by:

$$Y_t = \sum_{j=0}^{\infty} A^J \varepsilon_{t-J} = C(L)\varepsilon_{t-J} \qquad (9)$$

Element $C(L) = (C_0 + C_1 L + C_2 L^2 + \cdots)$ and $C_0 = 1$, $C_1 = A^1$, …, $C_j = A^J$.

Using the sequence, we can compute the *impulse responses* by *OLS* and construct the matrix A and the matrices C_j for $j = 0,1,2,\dots,J$, where j is a time horizon of our responses. We construct impulse responses over the two-year horizon, which equals 24 months.

The confidence bands are constructed applying the bootstrap for the same specification as is outlined above. For the bootstrap, we run 1,000 iterations, from which we create upper and lower bands using a 68 percent confidence level – see Gambacorta, Hofmann and Peersman (2012) or Peersman (2011).

Short-run restrictions

VAR specification outlined in the previous sub-section allows for the correlations between two and more external shocks coming from different variables. It is therefore important to orthogonalize the shocks by imposing short-run and long-run restrictions.

The Cholesky factor is defined as the lower triangular matrix S such that $S' = \Omega$, where Ω is the variance-covariance matrix. This allows to rewrite a *VAR(p)* process in terms of orthogonal shocks $\eta_t = S^{-1}\varepsilon_t$ with identity variance-covariance matrix:

$$A(L)y_t = S\eta_t \qquad (10)$$

where $var(\eta_t) = S^{-1}SS'(S^{-1})' = I$. The *impulse responses* to orthogonalized shocks η_t are then calculated from the *VMA*(∞) representation as follows:

$$y_t = A(L)^{-1}S\eta_t = C(L)S\eta_t = \sum_{j=0}^{\infty} C_j S\eta_{t-j} \qquad (11)$$

Using the Cholesky factorization, we are imposing the restriction that the shock of the second variable, for instance, has no contemporaneous impact on the first variable. As is clear from the following matrix, the element s_{12} of matrix S is simply zero.

$$S = \begin{pmatrix} s_{11} & 0 & 0 & 0 \\ s_{21} & s_{22} & 0 & 0 \\ s_{31} & s_{32} & s_{33} & 0 \\ s_{41} & s_{42} & s_{43} & s_{44} \end{pmatrix} \qquad (12)$$

The *confidence bands* are estimated in the fashion outlined in the previous sub-section. Finally, we compute *cumulative responses* for a horizon of 24 months as $(C_0 + C_1 + \cdots C_{24})$. Ronayne (2011) claims that cumulating the responses has the effect of smoothing the spikiness of the differenced variables, making inference more comfortable.

Long-run restrictions

We use long-run restrictions in the fashion of Blanchard and Quah (1989) to identify oil price shocks. Blanchard and Quah (1989) assume that the long-run effect is a lower triangular matrix. Thus, the restriction can be implemented, again, using a Cholesky factor as follows:

$$S = chol(C(1)\Omega C(1)') \qquad (13)$$

and

$$K = C(1)^{-1}S \qquad (14)$$

Applying principles defined by Equations (13) and (14), we can obtain *impulse responses* by rewriting Equation (11) as follows:

$$y_t = C(L)K\tau_t = H(L)\tau_t \qquad (15)$$

where $\tau_t = K^{-1}\varepsilon_t$. The long-run restriction is then implemented as:

$$H(1) = C(1)K = C(1)C(1)^{-1}S = S \qquad (16)$$

As S is a lower triangular matrix, the element $H(1)_{12} = 0$, similar to the case of Cholesky factorization. It means in our case that, for example, the shock in industrial production has no long-run effect on the level of oil prices. As in the Cholesky factorization, the errors are orthogonal since $var(\tau_t) = S^{-1}SS'(S^{-1})' = I$.

We estimate *confidence bands*, again using a bootstrap method. *Cumulative responses* are calculated for a horizon of 24 months as $(H_0 + H_1 + \cdots H_{24})$.

Sign restrictions

Finally, we impose sign restrictions on the *VAR* model. Unlike in the structural *VAR* models, we impose sign restrictions by means of the Bayesian framework using maximal likelihood estimator (*MLE*) instead of *OLS*. Bayesian structural *VAR* is widely used for identification of various economic shocks - see recent study of Lanne and Luoto (2019). We begin, again, with the *SUR* representation of the reduced *VAR(3)* specification:

$$\mathrm{Y} = \mathrm{X}B + \mathrm{U} \qquad (17)$$

where $\mathrm{Y} = (\mathrm{Y}_{p+1}, \dots, \mathrm{Y}_T)$, $\mathrm{X} = (1, \mathrm{Y}_{-1}, \dots, \mathrm{Y}_{-p})$, $\mathrm{U} = \left(u_{p+1}, \dots, u_T\right)$ and $B = (C, A_1, A_2, A_3)'$.

Equation (17) can be expressed in the vector form as follows:

$$\mathrm{y} = (I_n \otimes \mathrm{X})\beta + \mathrm{u} \qquad (18)$$

where $y = vec(\mathrm{Y})$, $\beta = vec(B)$, $\mathrm{u} = \mathrm{vec}(\mathrm{U})$ and $\mathrm{u} \sim N(0, \Sigma \otimes I_{T-p})$. Again, subscript T denotes the number of monthly observations and subscript p represents the number of lags.

Assuming the normality of errors, it is possible to express the likelihood of the sample, conditional on model parameters and a set of regressors X, as follows:

$$L(y|\mathrm{X}, \beta, \Sigma) \propto \left|\Sigma \otimes I_{T-p}\right|^{-\frac{T-p}{2}} exp\left\{\frac{1}{2}(y - I_n \otimes \mathrm{X}\beta)'(\Sigma \otimes I_{T-p})^{-1}(y - \ I_n \otimes \mathrm{X}\beta)\right\} \qquad (19)$$

where n represents, again, the number of included variables (four in our case). Assume that $\hat{B} = (\mathrm{X}'\mathrm{X})^{-1}\mathrm{X}'\mathrm{Y}$ is the $MLE = OLS$ estimate, and $S = \left(\mathrm{Y} - \mathrm{X}\hat{B}\right)'(\mathrm{Y} - \mathrm{X}\hat{B})$ is the sum of squared errors. Denoting $\hat{\beta} = vec(\hat{B})$ allows us to rewrite Equation (19) as follows:

$$L(y|\mathrm{X}, \beta, \Sigma) \propto \left|\Sigma \otimes I_{T-p}\right|^{-\frac{T-p}{2}} exp\left\{\frac{1}{2}\left(\beta - \hat{\beta}\right)'(\Sigma^{-1} \otimes \mathrm{X}'\mathrm{X})(\beta - \hat{\beta})\right\} exp\left\{-\frac{1}{2}tr(\Sigma^{-1}\mathrm{S})\right\} \qquad (20)$$

We choose a non-informative prior for B and Σ that is proportional to $|\Sigma|^{-\frac{n+1}{2}}$. We can then calculate the posterior of the parameters given the data using Bayes rule as follows:

$$P(B,\Sigma|y,\mathrm{X}) = L(y|\mathrm{X},\beta,\Sigma)p(B|\Sigma)p(\Sigma)$$
$$= |\Sigma|^{-\frac{T-p+n+1}{2}} exp\left\{\frac{1}{2}(\beta-\hat{\beta})'(\Sigma^{-1}\otimes \mathrm{X}'\mathrm{X})(\beta - \hat{\beta})\right\} exp\left\{-\frac{1}{2}tr(\Sigma^{-1}\mathrm{S})\right\} \tag{21}$$

Hence $\beta|\Sigma,y,\mathrm{X}\sim N(\hat{\beta},\Sigma\otimes(\mathrm{X}'\mathrm{X})^{-1})$ and $\Sigma|y,\mathrm{X}\sim IW(S,v)$ where $v = T - p - (np+1) - n - 1$.

The *impulse responses* of the reduced form *VAR*, are obtained through the *Gibbs sampler*, which is generating Σ by drawing from $IW(S,v)$. Parameter S is obtained by estimating Equation (17) using *MLE*. We then compute $\Sigma^0\otimes(\mathrm{X}'\mathrm{X})^{-1}$ and draw β^0 from $N(\hat{\beta},\Sigma^0\otimes(\mathrm{X}'\mathrm{X})^{-1})$, where $\hat{\beta} = vec(\hat{B})$ and $\hat{B} = (\mathrm{X}'\mathrm{X})^{-1}\,\mathrm{X}'\mathrm{Y}$. We repeat this procedure 3,000 times and then discard 1,000 initial draws and finally, we pick 1,000 of the draws.

Equation (17) can be rewritten as follows:

$$Y_t = C + AY_{t-1} + u_t \tag{22}$$

which can be expressed in terms of *VMA(∞)* as follows:

$$Y_t = \sum_{j=0}^{\infty} A^j u_{t-j} = C(L)u_t \tag{23}$$

where $C_0 = 1$, $C_1 = A^1$, …, $C_j = A^J$.

We construct 68 percent *confidence bands* taking the 16th and 84th percentile of C_j.

We have just outlined the procedure of computing impulse responses and confidence bands. In what follows, we need to impose sign restrictions. For this purpose, let $u_t = A\varepsilon_t$, where $\varepsilon_t\sim N(0,I_n)$ and A has the characteristics as $AA' = \Sigma$. We assume that A is a Cholesky factorization of Σ. In order to identify all the shocks in the system we need additional $\frac{n(n-1)}{2}$ conditions, where n represents the number of included variables.

The additional conditions correspond to the sign restrictions and are imposed using the *QR* decomposition algorithm proposed by Rubio-Ramirez, Waggoner and Zha (2010). As we consider the oil price shock, we show the proposed sign restrictions capturing the expected response of variables in the system (*OP*, *IND*, *CPI*, *SR*, *TP*) to the negative oil price shock (*OP*) in Table 1.

Appendix E. Results of Granger causality test

The Granger causality test is applied to identify a unidirectional causality between two variables. The null hypothesis states that *Y* variable does not Granger cause *X* variable. The results of Granger causality test for the USA and for the euro are area shown in Table A and in Table B, respectively. The Granger causality test was applied to the estimated baseline *VAR(3)* model outlined in Section 3. The lag length selection was based on the SIC information criteria. Recall that the VAR model estimate was preceded by testing the stationarity of the data using the ADF test, the results of which are shown in Table A.2, Appendix A. All data used were therefore transferred to their natural logarithm (except for the shadow policy rate and the term premium variables) and subsequently stationarised using the first differences – see Equation 2 in Methodology.

Table E.1. Results of Granger causality test for the US

X	Y	Full sample (03.1983-09.2018)	1st period (03.1983-11.2007)	2nd period (12.2007-11.2012)	3rd period (12.2012-12.2019)
OP	*IND*	70.605 *(0.070)*	0.698 *(0.952)*	10.942* *(0.027)*	0.792 *(0.940)*
OP	*CPI*	70.239 *(0.071)*	11.638* *(0.020)*	49.988 *(0.287)*	30.724 *(0.546)*
OP	*SR*	28.482 *(0.416)*	88.551 *(0.065)*	23.846 *(0.665)*	16.834 *(0.794)*
OP	*TP*	39.804 *(0.264)*	88.994 *(0.064)*	39.601 *(0.411)*	17.098 *(0.789)*
IND	*OP*	16.841* *(0.007)*	15.338* *(0.004)*	11.822* *(0.019)*	48.757 *(0.300)*
IND	*CPI*	46.036 *(0.203)*	17.226 *(0.787)*	15.073* *(0.005)*	18.474 *(0.764)*
IND	*SR*	3.240 *(0.356)*	16.380* *(0.003)*	17.329 *(0.785)*	13.895* *(0.008)*
IND	*TP*	39.076 *(0.272)*	1.735 *(0.784)*	92.064 *(0.056)*	24.151 *(0.660)*
CPI	*OP*	51.366* *(0.000)*	25.833* *(0.000)*	30.862* *(0.000)*	28.298* *(0.000)*
CPI	*IND*	76.372 *(0.054)*	15.045 *(0.826)*	55.646 *(0.234)*	28.787 *(0.578)*
CPI	*SR*	70.855 *(0.069)*	2.815 *(0.589)*	12.926* *(0.012)*	39.679 *(0.410)*
CPI	*TP*	70.385 *(0.071)*	22.674 *(0.687)*	10.273* *(0.036)*	63.274 *(0.176)*
SR	*OP*	44.348 *(0.218)*	86.387 *(0.071)*	25.554 *(0.635)*	50.721 *(0.280)*
SR	*IND*	16.868* *(0.001)*	12.583* *(0.014)*	14.352* *(0.006)*	92.924 *(0.054)*
SR	*CPI*	22.062* *(0.015)*	10.879* *(0.028)*	3.801 *(0.434)*	22.874* *(0.000)*
SR	*TP*	18.957* *(0.000)*	3.029 *(0.553)*	75.898 *(0.108)*	32.450* *(0.000)*
TP	*OP*	24.538 *(0.484)*	12.456 *(0.871)*	10.574 *(0.901)*	5.219 *(0.266)*
TP	*IND*	23.063* *(0.000)*	56.847 *(0.224)*	31.987* *(0.000)*	75.938 *(0.108)*
TP	*CPI*	17.429 *(0.627)*	0.812 *(0.937)*	0.774 *(0.942)*	39.211 *(0.417)*
TP	*SR*	53.089 *(0.151)*	8.731 *(0.068)*	14.537 *(0.835)*	39.533 *(0.412)*

Notes: Table shows the Chi-square statistics and respective p-values are in the parenthesis
** indicates that we do not reject null hypothesis at the level of 0.05 (that Y Granger causes X)*
3 lags were included in all VAR specifications.
Source: own calculations based on the Federal Reserve economic database

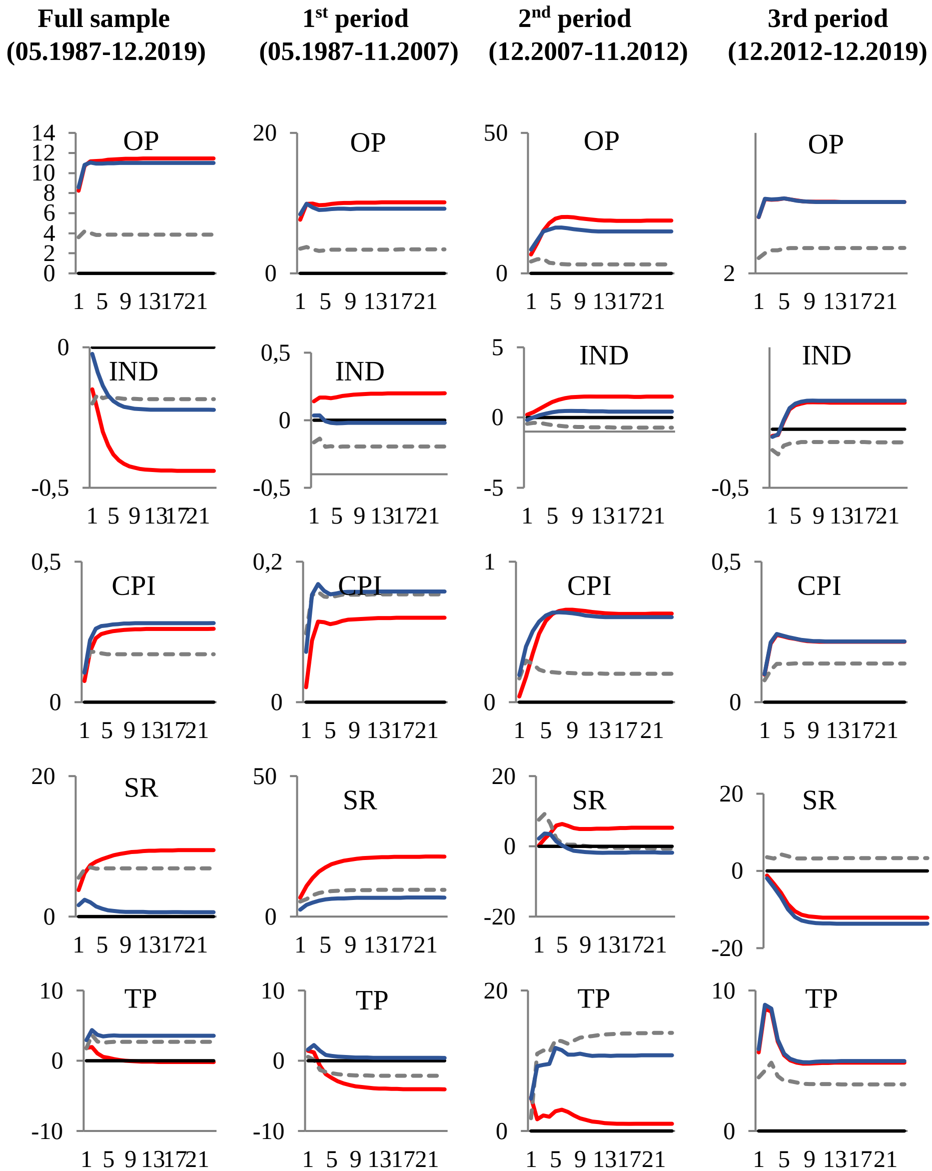
Table E.2. Results of Granger causality test for the euro area

X	Y	Full sample (07.2005-09.2018)	1st period (07.2005-11.2007)	2nd period (12.2007-11.2012)	3rd period (12.2012-12.2019)
OP	*IND*	65.643 *(0.087)*	n.a. n.a.	26.907 *(0.442)*	86.287* *(0.035)*
OP	*CPI*	22.792 *(0.517)*	n.a. n.a.	14.721 *(0.689)*	11.671 *(0.761)*
OP	*SR*	13.524 *(0.717)*	n.a. n.a.	96.381* *(0.022)*	13.068 *(0.728)*
OP	*TP*	14.148 *(0.702)*	n.a. n.a.	31.111 *(0.375)*	0.574 *(0.902)*
IND	*OP*	11.626* *(0.009)*	n.a. n.a.	13.292* *(0.004)*	63.706 *(0.095)*
IND	*CPI*	63.748 *(0.095)*	n.a. n.a.	94.591* *(0.024)*	40.773 *(0.253)*
IND	*SR*	28.771 *(0.411)*	n.a. n.a.	23.264 *(0.507)*	38.376 *(0.280)*
IND	*TP*	48.771 *(0.181)*	n.a. n.a.	33.888 *(0.335)*	1.304 *(0.728)*
CPI	*OP*	32.648 *(0.353)*	n.a. n.a.	13.882 *(0.708)*	17.034 *(0.636)*
CPI	*IND*	11.622* *(0.009)*	n.a. n.a.	6.475 *(0.091)*	49.447 *(0.176)*
CPI	*SR*	17.778 *(0.620)*	n.a. n.a.	4.001 *(0.261)*	62.166 *(0.102)*
CPI	*TP*	33.777 *(0.337)*	n.a. n.a.	55.187 *(0.138)*	66.555 *(0.084)*
SR	*OP*	14.170* *(0.003)*	n.a. n.a.	15.089* *(0.002)*	27.762 *(0.427)*
SR	*IND*	11.500* *(0.009)*	n.a. n.a.	0.591 *(0.899)*	46.938 *(0.196)*
SR	*CPI*	11.541* *(0.010)*	n.a. n.a.	64.158 *(0.093)*	8.239* *(0.041)*
SR	*TP*	99.222* *(0.019)*	n.a. n.a.	48.419 *(0.184)*	35.318 *(0.317)*
TP	*OP*	72.443 *(0.065)*	n.a. n.a.	12.053* *(0.007)*	25.375 *(0.469)*
TP	*IND*	48.325 *(0.184)*	n.a. n.a.	10.905* *(0.012)*	12.111* *(0.007)*
TP	*CPI*	46.635 *(0.198)*	n.a. n.a.	4.341 *(0.227)*	85.331* *(0.036)*
TP	*SR*	16.588* *(0.001)*	n.a. n.a.	10.776* *(0.013)*	15.084* *(0.002)*

Notes: Table shows the Chi-square statistics and respective p-values are in the parenthesis
** indicates that we do not reject null hypothesis at the level of 0.05 (that Y Granger causes X)*
1st period covers short period of time yielding only 29 observations, which compromised the results of the Granger causality test
3 lags were included in all VAR specifications.
Source: own calculations based on the Federal Reserve economic database and Eurostat

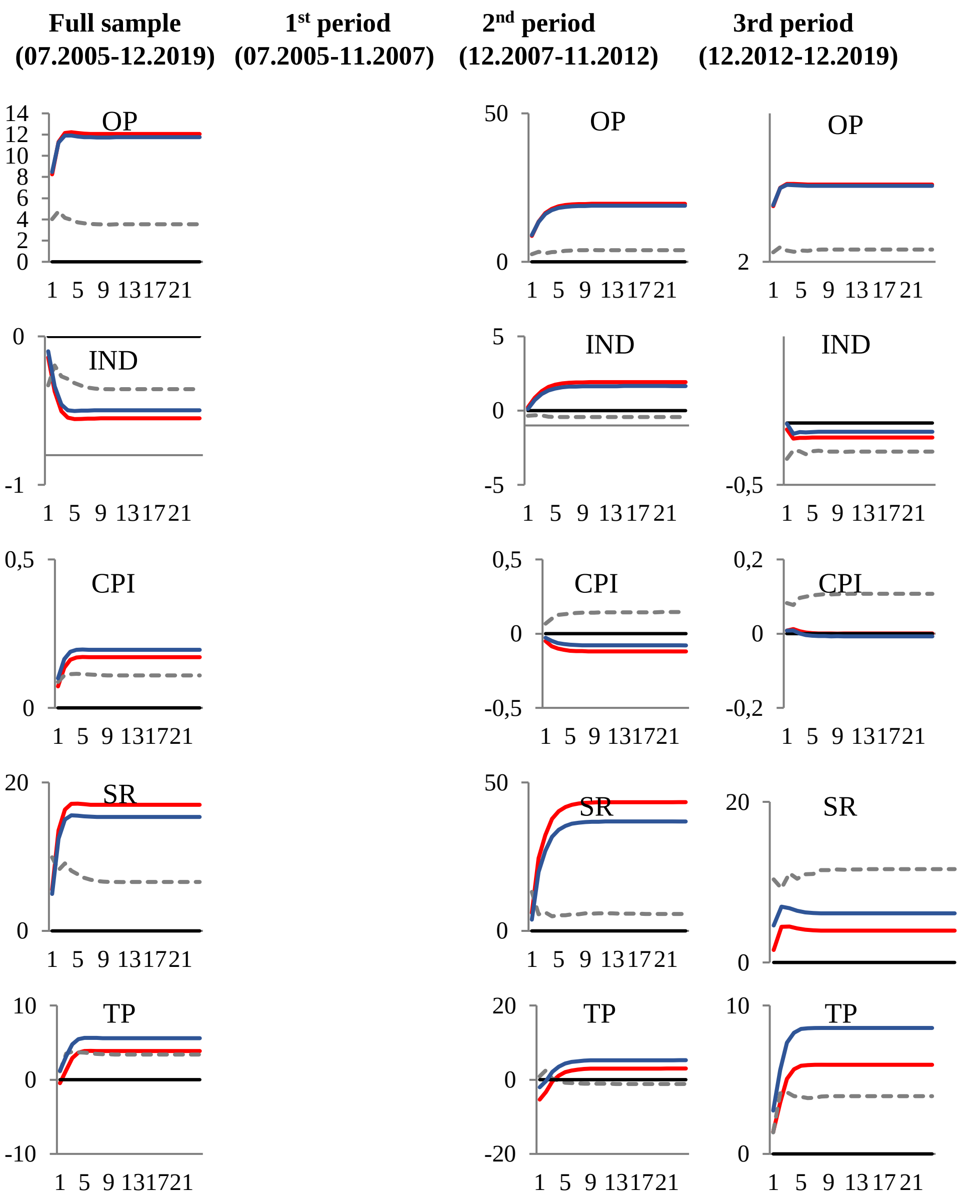
Appendix F. Comparison of functions from all VAR specifications

Figure F.1. Comparison of functions from all VAR specifications for the US



Note: blue line – cumulative impulse response under Cholesky factorization; red line – cumulative impulse response under Blanchard-Quah long-run restriction; dashed shadow line – cumulative impulse response under sign restrictions
Source: own calculations based on the Federal Reserve economic database

Figure F.1. Comparison of functions from all VAR specifications for the euro area



Note: blue line – cumulative impulse response under Cholesky factorization; red line – cumulative impulse response under Blanchard-Quah long-run restriction; dashed shadow line – cumulative impulse response under sign restrictions
Source: own calculations based on the Federal Reserve economic database