

Commodity prices and real and financial processes in the Euro area: a Bayesian SVAR approach

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Abstract. This paper deals with links between global energy and non-energy commodity prices and real and financial processes in the euro area. We use monthly data spanning from 1997:1 to 2013:12 and the structural Bayesian VAR model with Sims-Zha prior specification. The analysis is performed for three sub-periods in order to capture potential changes in the reactions over time. Our main finding is that, commodity prices are indeed related to financial processes in the euro area macroeconomy: changes in the euro area interest rate have significant influence on commodity prices. There is no relation, however, between real processes in the euro area and commodity prices. Additionally, the relations between commodity prices have gradually become tighter over time.

Keywords: commodity prices, real economy, financial markets, BSVAR, Sims-Zha prior.

JEL Classification: C3, E37, E47, Q17, Q43

AMS Classification: 91G70

1 Introduction

Wide fluctuations in commodity prices have been observed in recent decades, which, according to Frankel and Rose [7], might have been caused by several factors including strong global growth (especially in China and India), easy monetary policy (low real interest rates or expected inflation), a speculative bubble and risk (geopolitical uncertainties). There are numerous articles devoted to the connections between commodity prices and financial or macroeconomic indicators. In most of them commodity prices are represented by oil prices and agricultural prices, while macroeconomic fundamentals include interest rates, inflation, exchange rates, industrial production or economic growth. Some researchers investigate the relationship between oil prices and economic growth (Cogni and Manera [5]), between oil prices and inflation rates (Chen [4]), and between oil prices and exchange rates (Śmiech and Papież [15]). Akram [1] finds evidence of a negative impact of interest rates on commodity prices, but Frankel and Rose [7] and Alquist et al. [2] do not find statistically significant relationships between real interest rates and oil prices. Tiwari [17] argues that the relationship between oil prices and German industrial production is ambiguous. Papież and Śmiech [11] analyse dependencies between the prices of crude oil and other commodities and financial investments.

The goal of this paper is to examine relations between global commodity prices and economic developments in the euro area macroeconomy in the period from 1997:1 to 2013:12. Since our intention is to include both real and financial processes we examine interdependencies between commodity prices on the one hand and economic activity (represented by industrial production) and financial conditions (represented by interest rate) in the euro area on the other hand. We also have the US dollar exchange rate in the analysis as it has been found in other studies to be related to global commodity prices (see e.g. Akram [1]). The Bayesian structural vector autoregression (BSVAR) model (see e.g. Sims and Zha [13]) is used to investigate these relations. It is quite likely that relations examined are not stable and that is why the sample is divided into three non-overlapping sub-periods.

2 Methodology

Let us consider m -dimensional structural VAR(p) model for a sample of size T (Sims and Zha [13]):

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$$y_t A_0 = y_{t-1} A_1 + \dots + y_{t-p} A_p + c + \varepsilon_t, t = 1, \dots, T, \quad (1)$$

where y_t is a vector of observations at time t , A_l is a coefficient matrix for the l^{th} lag, A_0 is a contemporaneous coefficient matrix, c is a vector of constants, ε_t is a vector of i.i.d. normal structural shocks. A distribution of ε_t conditional on y_{t-s} for $s > 0$ is normal and $E(\varepsilon_t | y_{t-s}, s > 0) = 0$, $E(\varepsilon_t' \varepsilon_t | y_{t-s}, s > 0) = I$. Let us rewrite (1) in the matrix form so that the columns of the coefficient matrices correspond to the equations $Y A_0 - X A_+ = E$. The matrix X contains the lagged Y 's and a column of 1's. The matrix A_+ is a matrix of the coefficients of the lagged variables and constants. The matrix E is a matrix of shocks. Define $a_0 = \text{vec}(A_0)$, $a_+ = \text{vec}(A_+)$, $A = \begin{bmatrix} A_0 \\ A_+ \end{bmatrix}$, $a = \text{vec}(A)$. The vec operator transforms a matrix into an vector by stacking the columns. We consider the identified model and assume that A_0 is an upper triangular nonsingular matrix.

Because of normally distributed shocks the conditional likelihood function for Y is a multivariate normal likelihood. The Sims-Zha prior is a certain modification of Litterman's structure (Sims and Zha [13], Litterman [10]). It is formed conditionally: $\pi(a) = \pi(a_0) \varphi(a_+; \mu(a_0), H(a_0))$, where $\mu(a_0)$ and $H(a_0)$ are a vector of prior mean parameters and the prior covariance for a_+ , respectively. The function φ is a multivariate normal density. The prior specification is first defined for the unrestricted VAR model and then mapped into the restricted prior parameter space (see Waggoner and Zha [18]). Under some additional assumption about symmetry the posterior distribution is tractable – it can be easily sampled. The prior conditional mean is given by $E(A_+ | A_0) = \begin{bmatrix} A_0 \\ 0 \end{bmatrix}$. Sims and Zha assumed the conditional prior independent across elements of A_+ with standard deviation $\lambda_0 \lambda_1 / \sigma_j t^{\lambda_3}$ for the element corresponding to the l^{th} lag of the variable j in the equation i . The parameter λ_0 controls the tightness of beliefs on A_0 , λ_1 controls tightness of the beliefs about the A_1 parameters and λ_3 controls the rate at which the prior variance shrinks with increasing lag length. The hyperparameters σ_j are introduced to correct scale. The prior standard deviation of constants equals a product of $\lambda_0 \lambda_4$, where λ_4 controls tightness around the intercept. Sims and Zha specified the prior distribution of each i^{th} column of A_0 as a normal $N(0, S_i)$, where S_i is some covariance matrix. They assumed a conditional independence of A_+ but unconditional prior dependence, corresponding to A_0 . The prior specification assumes the same correlation structure for the regression parameters and the structural residuals. More details about this prior specification and the hyperparameters λ_i might be found in Robertson and Tallman [12]. Posterior calculations are conducted by means of the Gibbs algorithm discussed in Waggoner and Zha [18].

3 Data

To examine relations between commodity prices on the one hand and real and financial processes in the euro area on the other hand, we use monthly data from 1997:1 to 2013:12. The data cover five series of variables. Industrial production index (IP) in the euro area describes real economy. Real 3-month interest rate in the euro area (IR) reflects financial conditional prevailing in the euro area. The data for both variables are taken from the Eurostat database. The real exchange rate of the US dollar (REX) is also included (its increase reflects the US dollar appreciation against the euro). The commodity price indices, i.e. energy price index (PENd) and non-energy commodity price index (PNEND), are based on nominal US dollars prices deflated with the US CPI. The data for these variables are taken from the databases of World Bank and Federal Reserve Bank of St. Louis. The energy price index (world trade-base weights) consists of crude oil (84.6%), natural gas (10.8%) and coal (4.6%). The non-energy commodity price index includes agriculture (64.8%), metals (31.6%), and fertilizers (3.6%). All series are expressed as indices equal 100 in 2010, seasonally adjusted and specified in natural logarithms.

The whole sample period is divided into three sub-periods: 1997:1-2002:12, 2003:1-2008:12, and 2009:1-2013:12. The division is motivated mainly by the disparate behaviour of commodity prices in these sub-periods. In the first sub-period the energy price index increased by 5.5%, while the non-energy commodity price index decreased by 29.2% (data are for indices deflated with the US CPI). In the middle sub-period, marked with large increases and decreases of commodity prices, the corresponding changes were 28.5% and 32.2%. In the last sub-period strong rebound effect was observed with 78.5% and 16.5% increases in commodity prices respectively. It

is interesting to observe that the average real interest rate was negative in the last sub-period (-0.8%) whereas it was positive in the other two sub-periods (2.0% and 0.7% on average respectively).

4 Empirical results

4.1 Time series properties of the data

A preliminary analysis of the series is carried out before estimating the main model. The standard augmented Dickey–Fuller (ADF) unit root tests for both the intercept and the trends specifications demonstrate that all variables have unit roots for each analysed sub-period. The number of lags in the test is established using AIC criterion⁵. Next, the presence of long-term relationship between integrated variables is investigated. The trace test statistic proposed by Johansen and Juselius [9] is used. If the variables are co-integrated, the VAR in first difference would not be correctly specified, and the long-term result would be very helpful in exploring the efficient parameters of short-term dynamics. According to trace test statistics and maximum eigenvalue test, there is no cointegration at 5 percent level in the first and third sub-periods. Test results demonstrate some evidence of the presence of cointegration only in the second sub-period. Trace test indicates one cointegrating equation at the 0.05 level. In contrast, maximum eigenvalue test indicates no cointegration at the 0.05 level⁶. Since the results of cointegration tests are at best ambiguous (if not suggesting the lack of cointegration), and the variables used are I(1), we use a VAR for the first differences in our five variables.

4.2 Structural impulse response analysis

We consider three BSVAR(p) models for $p=1, 2, 3$ under the same Sims-Zha prior specification with arbitrarily chosen hyperparameters $\lambda_0 = \lambda_1 = \lambda_3 = \lambda_4 = 1$. In other words, we choose a fairly diffuse prior distribution, which expresses the lack of prior knowledge, and shut off dummy observations (see Robertson and Tallman [12]). Selecting the most appropriate number p actually amounts to choosing the model ensuring the best data fit. The Bayesian factors point to the lag order $p=3$ for all sub-periods and therefore we limit our further considerations to the BSVAR(3). Results are obtained via the R package MSBVAR (Brandt and Davis [3]). The conclusions are based on 160,000 Gibbs draws, preceded by 60,000 burn-in cycles. Generating longer chains bears insignificant influence on the impulse response analysis. The impulse response analyses under the BSVAR(3) and the BSVAR(1) lead to almost the same conclusions. The main difference is that the posterior error bands are wider under the BSVAR(3). It means more uncertainty. The Gibbs sampler draws are normalized using the “DistanceMLA” method recommended in Waggoner and Zha [18]. It corresponds to the positive system shocks.

A five-variable VAR is estimated with industrial production (ΔIP), real interest rate (ΔIR), real exchange rate (ΔREX), energy price index ($\Delta PENd$) and non-energy commodity price index ($\Delta PNENd$). Industrial production index is set as the first variable in VAR because its adjustments to changes in other variables are assumed to rather sluggish. Since the aim of this study is to examine the response of commodity prices to real and financial processes, we place them after industrial production, real interest rate and real exchange rate. Here, we follow the ordering proposed by Akram [1] (more conventional ordering has been used by Hanson [8]). Figures 1-3 show the impulse responses to structural innovations of all our variables across sub-periods. For example, the first column of Figure 1 illustrates with a blue line the impulse response of each variable in the system to an innovation in industrial production index (output shock). Green and red lines determine 90% posterior intervals around the impulse responses based on the highest posterior density region.

In the sub-period 1997:1-2002:12 output shocks were neutral for all variables (Figure 1). The response of real interest rate, however, was positive and only marginally insignificant. Thus, the response was (almost) in line with an anti-inflationary orientation of the monetary authority: as output increases unexpectedly, the risk of inflation grows and monetary policy needs to be tightened.

Energy price index’s reaction to shocks in the real interest rate was consistent with Hotelling’s rule, which says that the benefit from storing a commodity should be equal to the interest rate. The benefit includes a revaluation gain and a convenience yield and is adjusted downwards by storage cost and risk premium (see Akram [1], Frankel and Rose [7] or Šmiech et al., [16]). Hotelling’s rule implies that an increase in the rate of interest, given expected future commodity prices, lowers the current commodity price. Such a reaction of energy price index to the interest rate makes them similar to asset prices (Svensson [14]).

⁵ Detailed test results are available from the authors upon request.

⁶ Since the length of the sample is not long, and there are four series in a vector of interests, a Monte Carlo experiment is performed and the empirical critical values of trace test are determined. We find that in such case null hypothesis of no cointegration is rejected too often. For detailed information contact the authors.

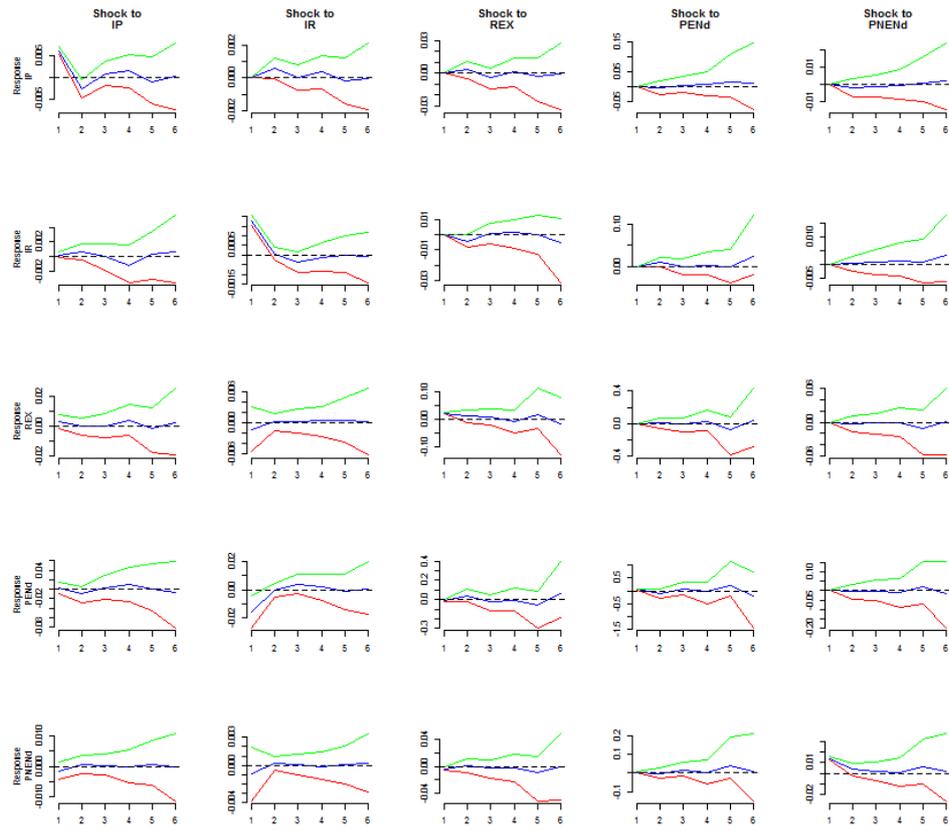


Figure 1 The impulse-responses results in sub-period 1997:1-2002:12

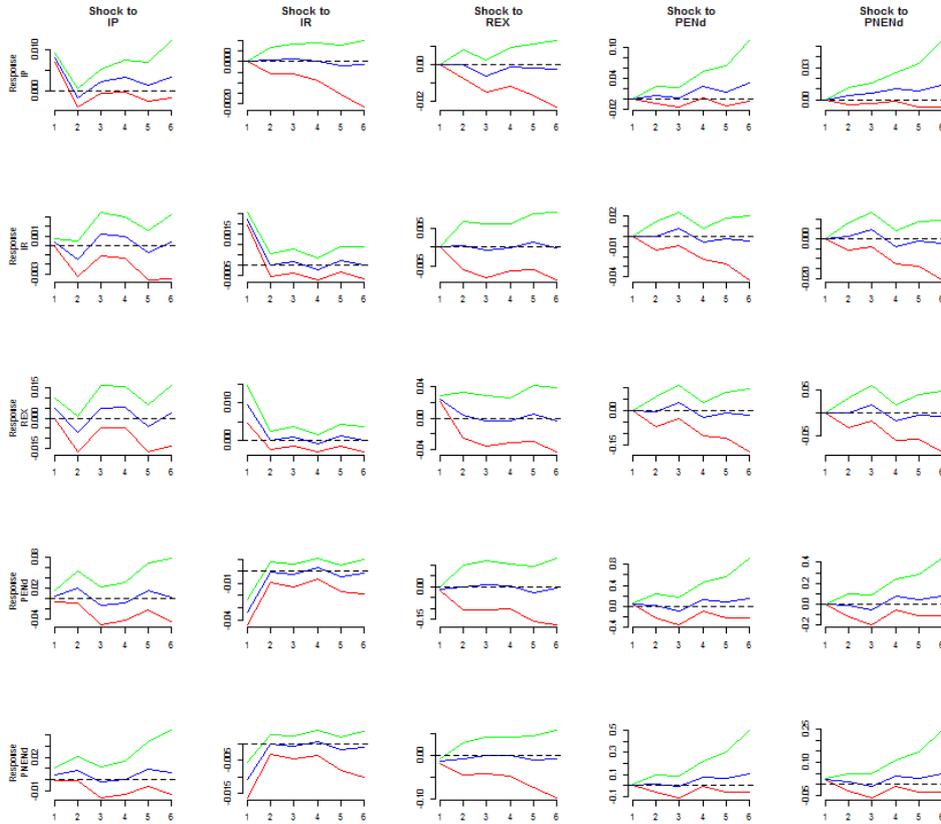


Figure 2 The impulse-responses results in sub-period 2003:1 – 2008:12

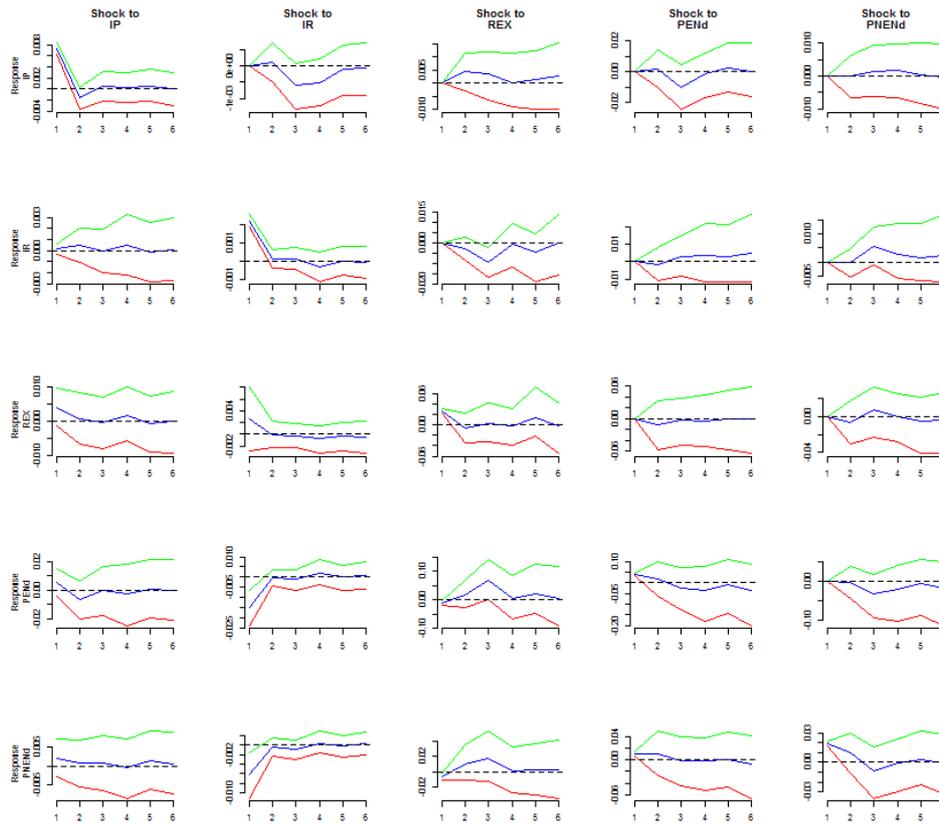


Figure 3 The impulse-responses results in sub-period 2009:1 - 2013:12

The US dollar real depreciation had a positive impact on both commodity price indices, though for energy price index at 10% level of significance only. Such reaction was found by other researchers as well, especially for the oil price (see e.g. Akram [1]). The conventional explanation is that commodity prices are quoted in US dollars, so when the dollar depreciates foreign currency commodity prices (i.e. converted into euros or pounds) are lower. Thus, the demand for commodities goes up, which results in an upward pressure on dollar price of commodities and some reversal of the initial depreciation of the US dollar.

Non-energy commodity prices respond positively to a shock in energy price index. Both commodities could be, therefore, seen as related to one another. In other words and less formally, non-energy commodity prices could not deviate too much from energy prices.

Responses in the third sub-period are almost the same as in the first sub-period (Figure 3). The only difference is that non-energy commodity price index responds to real interest rate shocks, which makes these commodities similar to energy prices.

There are more differences in the middle sub-period (Figure 2). Responses to output shocks of all variables are negligible. The importance of real shocks decreases in this sub-period and the links between both group of commodities on the one hand and the interest rate on the other hand are enhanced. It seems that commodity prices and interest rate detach from the real processes in the euro area, although not from one another. In fact, the responses of both energy and non-energy commodity price indices to interest rate shocks are twice as strong in this sub-period than in the two other sub-periods.

One more observation can be derived from Figures 1-3. The reaction of non-energy commodity price index to shocks in the energy price index has gradually increased from 0.004 to 0.009. This can be related to the developments in the biofuel market (e.g. at the share of biofuels in transport fuel consumption in the EU is estimated at 5-6%; see Demirbas [6]) and energy policy of the EU. The issue definitely requires further research.

5 Conclusion

This paper examines relations between global commodity prices and real and financial processes in the euro area macroeconomy in the period spanning from 1997:1 to 2013:12. The analysis is based on theoretical presumption that economic development in a large open economy that trades (mainly imports) commodities has non-

negligible effect on global commodity prices. We have used a Bayesian SVAR framework for three non-overlapping sub-period to allow for instability in the relations.

Our main findings are twofold. First, commodity prices are indeed related to financial processes in the euro area macroeconomy: changes in the euro area interest rate have significant influence on commodity prices. There is no relation, however, between real processes in the euro area and global commodity prices. This is especially visible in the middle sub-period, i.e. in the pre-crisis period, in which a (weak) link between interest rate and industrial production fades away as well. It seems that commodity prices and interest rate have detached from the real processes, although not from one another in this sub-period.

Second, the relation between non-energy commodity prices and energy prices has become tighter over time. This observation can be a symptom of a rising importance of biofuels in transport fuel consumption in the European Union and the EU energy policy. This, however, requires further research.

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