Natural Gas Prices, Inflation Expectations, and the Pass-Through to Euro Area Inflation*

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First Version: March 2023
This Version: November 2024

Abstract

This paper examines the recent increase in natural gas prices, the sensitivity of inflation expectations, and the pass-through to inflation. Using a semi-structural vector autoregression, we identify a natural gas price shock in the euro area with a combination of sign and zero restrictions. We rely on market-based measures of inflation expectations. The results show that shocks to the real price of natural gas affect both inflation and inflation expectations. To investigate the relative importance of the pass-through of inflation expectations to inflation, we conduct a structural scenario analysis in which inflation expectations are insensitive to movements in the real price of natural gas. The results indicate the presence of strong second-round effects via expectations. Our analysis provides guidance for policymakers to better understand the potential trade-offs of different policy responses to natural gas price shocks, particularly with respect to the potential for a de-anchoring of inflation expectations.

Keywords: Natural gas price shocks; Inflation expectations; Euro area; Counterfactual analysis.

JEL Codes: C32, E31, Q43.

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1. Introduction

After political tensions culminated in Russia's invasion of Ukraine in February 2022, European economies experienced dramatic hikes in energy prices, with natural gas being the most affected. However, the macroeconomic consequences of distortions in the natural gas market remain relatively unexplored. These price distortions affect inflation expectations, which facilitate the pass-through to inflation. The precise empirical strength of this pass-through has received much attention (Bernanke et al., 1997; Harris et al., 2009; Peersman and Van Robays, 2009; Coibion and Gorodnichenko, 2015b; Wong, 2015; Bjørnland, Larsen and Maih, 2018; Aastveit, Bjørnland and Cross, 2023). While the literature has focused on oil prices, this paper examines the pass-through arising from natural gas price distortions. As shown in Figure 1, pandemic-related factors led to an initial increase in short- and long-term market-based inflation expectations toward the 2% target in 2021, followed by a rapid acceleration in 2022. Real natural gas prices show similar dynamics, but a less pronounced co-movement compared to oil prices (see Figure 2). This has raised concerns among researchers and policymakers about the risk of a de-anchoring of inflation expectations (Schnabel, 2021; Blanchard, 2022; Reis, 2022; Schnabel, 2022b).

The goal of this paper is to investigate the recent surge in natural gas prices, its impact on inflation expectations, and the empirical strength of the pass-through to prices. We examine the relationship between key macroeconomic variables and natural gas prices using a semi-structural Bayesian vector autoregressive (VAR) model that covers a monthly sample from January 2004 to December 2022. We develop a combination of sign and zero restriction to identify a real natural gas price shock. The key identifying assumptions are a contemporaneous positive co-movement of the real natural gas price, inflation, and inflation expectations, while economic activity is constrained to not react on impact. Our results indicate that natural gas price shocks significantly affect both inflation and inflation expectations. A one standard deviation shock, equivalent to a 10% increase in the real price of natural gas, causes an immediate increase in inflation of 8 basis points (bps) and an increase in inflation expectations of 5 bps. After about one year, the maximum observed increases are 22 bps for inflation and 17 bps for inflation expectations. These considerably stronger effects at longer horizons suggest an important role for second-round effects via inflation expectations. In addition, we observe a protracted decline in economic activity and a tightening of the monetary policy stance.

Commodity prices, such as natural gas prices, feed into inflation through several channels (Peersman and Van Robays, 2009). A rise in commodity prices exerts a direct effect, particularly through the energy component of consumer prices, and an indirect effect that reflects higher production costs of non-energy goods and services. We refer to this as the *first-round effect*. We deem the effects working through inflation expectations as *second-round effects*. From a policymaker's perspective, monitoring *second-round effects* is crucial because they lead to a more persistent inflation response. Increases in costs, prices, or expectations affect price setting and wage bargaining, facilitating the pass-through to inflation (Wong, 2015; Coibion, Gorodnichenko and Kamdar, 2018; Aastveit, Bjørnland and Cross, 2023).[1] However, the empirical literature lacks consensus on the strength of this pass-through. Large increases in energy prices pose a risk to the anchoring of inflation expectations, necessitating appropriate policy intervention (Reis, 2023). This channel is thus of particular importance for the discussion on de-anchoring risks.[2] To differentiate between first- and second-round effects, we conduct a structural scenario analysis (Antolin-Diaz, Petrella and Rubio-Ramirez, 2021). For the scenario analysis, we additionally identify a shock to inflation expectations following Kilian and Zhou (2022a). The results highlight a pronounced *second-round* channel and a more muted *first-round* channel. After one year, inflation is 15 bps lower when inflation expectations are insensitive to real natural gas price shocks. The pass-through of inflation expectations to realized inflation is well below unity. Examining the entire term structure of inflation expectations, we find that short-term inflation expectations exhibit the strongest second-round effects before gradually declining and becoming insignificant for long-term expectations. The evidence suggests no de-anchoring risks of long-term expectations.

An important aspect of this analysis is the measurement of inflation expectations. We utilize inflation-linked swaps (ILS), which provide market-based expectations.[3] Since ILS data is available at a high frequency (i.e., monthly in our case), we can confidently estimate our semi-structural VAR over a relatively short sample period. However, it is important to note that the ILS also incorporate an inflation risk premium, which is part of the *expectational* component and is assumed to remain constant.[4][5] Nevertheless, ILS provide superior

[1] Another channel may counteract inflationary effects through a reduction in aggregate demand. This disinflationary effect occurs because commodity price shocks shift the aggregate supply curve upward along the downward-sloping aggregate demand curve.

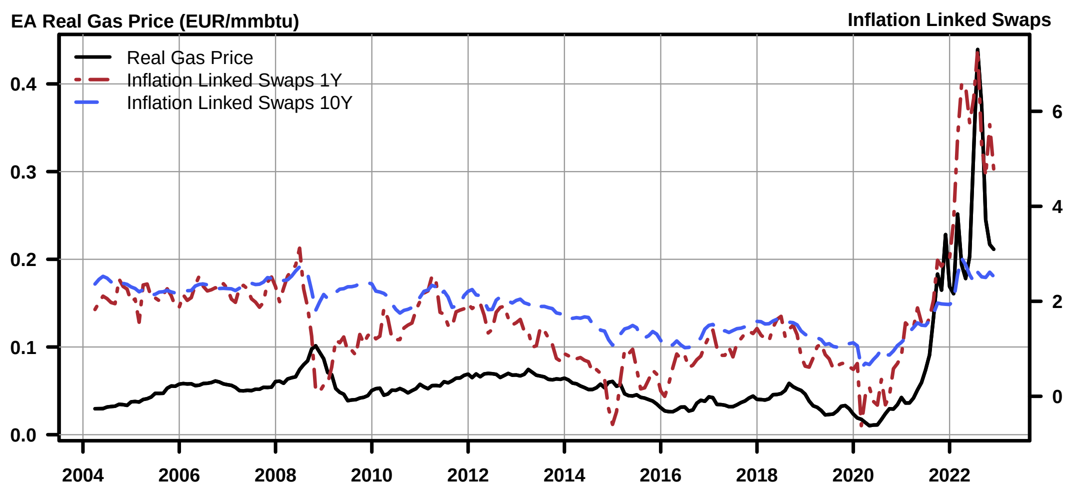
[2] This has been frequently pointed out in speeches by policymakers when discussing possible policy responses to the energy transition, to global energy price shocks, or fiscal policy responses to energy price increases (see, e.g., Schnabel, 2022a; Schnabel, 2022b; Lane, 2022a; Mester, 2022; Lane, 2022c). As highlighted by Ider et al. (2023) and partly by Castelnuovo, Mori and Peersman (2024), monetary policy can also affect energy price hikes through other channels as well.

[3] In principle, these swaps are derivative products that are linked to some sort of price index. By design, the swap is a forward contract between two parties, where the buyer party pays a (fixed) nominal rate and receives a real rate from the seller party. Consequently, the price of the swap is contingent upon the actual and anticipated inflation rates. This enables the instrument to be utilized for the purpose of hedging inflation

[4] For the US, the Federal Reserve of Cleveland provides an estimate of the inflation risk premium. This series fluctuates mildly around a long-run mean and shows no obvious correlations to business cycle fluctuations and/or historical episodes (e.g., the high inflation period of the 1970s).

[5] We provide robustness with the survey of professional forecasters, which is only available on a quarterly frequency.

Figure 1: Real Natural Gas Prices and Inflation Expectations in the Euro Area.



Notes: The figure shows the real natural gas price in Euro per mmbtu (transformed into Euro from US dollar and deflated with the harmonized index of consumer prices), and inflation linked swaps for 1Y and 10Y ahead. Vertical axis denotes real gas price (left axis) and inflation expectations in percent (right axis). Sample period ranges from 2004M1 to 2022M12. Horizontal axis denotes time in months.

information to other market-based measures (e.g., inflation-indexed treasury yields) as demonstrated by Haubrich, Pennacchi and Ritchken (2012). While market-based measures are similar to expectations of professional forecasters, they diverge from household inflation expectations, which tend to be substantially higher, as shown by D'Acunto, Malmendier and Weber (2023). Furthermore, Coibion and Gorodnichenko (2015a) document evidence of information rigidities in inflation expectations measured via ILS.

The paper contributes to several strands of the literature. The paper investigates the macroeconomic consequences of real natural gas price shocks. While the literature has traditionally focused on crude oil and other major commodities, natural gas markets have recently attracted more interest.[6] For instance, Nick and Thoenes (2014) and Rubaszek, Szafranek and Uddin (2021) develop structural models of the natural gas market for Germany and the US, respectively. Examining a rich set of structural drivers of euro area inflation, Bańbura, Bobeica and Martínez Hernández (2023) show the strong influence of gas prices on core inflation. Casoli, Manera and Valenti (2024) and Adolfsen et al. (2024) provide a structural model of the natural gas market that disentangles supply and demand forces. Using a narrative approach to identify natural gas supply shocks in the euro area, Alessandri and Gazzani (2023) show that such shocks materialize over longer horizons compared to oil supply shocks. A different approach by Kilian and Zhou (2023) jointly identifies

[6] The literature examining the macroeconomic effects of oil shocks is large and well established. See, inter alia, Barsky and Kilian (2002), Hamilton (2003), Kilian (2008), Kilian (2009), Bjørnland, Larsen and Maih (2018) and Baumeister and Hamilton (2019).

the impact of several energy prices, including natural gas and electricity prices, and finds that focusing on just one commodity is likely to underestimate the inflationary impact. To summarize, these studies highlight the importance of differentiating between different energy markets due to the different pass-through to inflation. Moreover, energy supply shocks play an important role in inflationary responses. Among energy shocks, natural gas supply shocks are highly relevant in the recent high inflation episode. This paper contributes to this literature by identifying a real natural gas price shock and investigating its pass-through to inflation.

The literature investigating the sensitivity of inflation expectations to energy price shocks and its pass-through on inflation focuses mostly on crude oil prices (Baumeister, Peersman and Van Robays, 2010; Clark and Terry, 2010; Wong, 2015). The literature finds mixed results regarding the empirical strength of the pass-through via inflation expectations and only attributes a limited role (or a declining role over time) to oil price shocks in driving inflationary responses. An exception is Aastveit, Bjørnland and Cross (2023) who disentangle oil supply and oil demand shocks and find that economic activity shocks have a significant pass-through to inflation. Investigating the impact of rising oil prices in 2020-23, Kilian and Zhou (2022b) also find limited evidence on the overall impact on inflation. Inflation expectations, however, play an important role for the impact and the transmission of energy price shocks. In this paper, we examine the second-round effects via inflation expectations to real natural gas price shocks for the euro area. We find similar effects for oil prices, which is in line with Peersman and Van Robays (2009) showing that second-round effects are much more pronounced in the euro area than in the US. To the best of our knowledge, we are thus the first to highlight the euro area's potential inflationary risks stemming from second-round effects of commodity price shocks, particularly of natural gas price shocks. In the presence of relevant second-round effects, the anchoring of inflation expectations is an important policy tool (Reis, 2021; Blanco, Ottonello and Ranosova, 2022).

The paper is organized as follows. Section 2 discusses the particularities of the natural gas market. Section 3 presents the econometric framework, the identification strategy, and how we construct the counterfactual experiment. Section 4 shows the baseline results and in Section 5 we offer some extensions. Finally, Section 6 concludes.

2. Some Structural Aspects of the Natural Gas Market

The natural gas market has several characteristics that make it an interesting subject to study. First, the demand for natural gas is increasing worldwide, but especially in Europe. Second, the natural gas market has evolved into an independent market, most likely decoupled from the oil market. Third, unlike oil, natural gas is traded much more locally due to the necessary infrastructure, resulting in different price dynamics around the world.

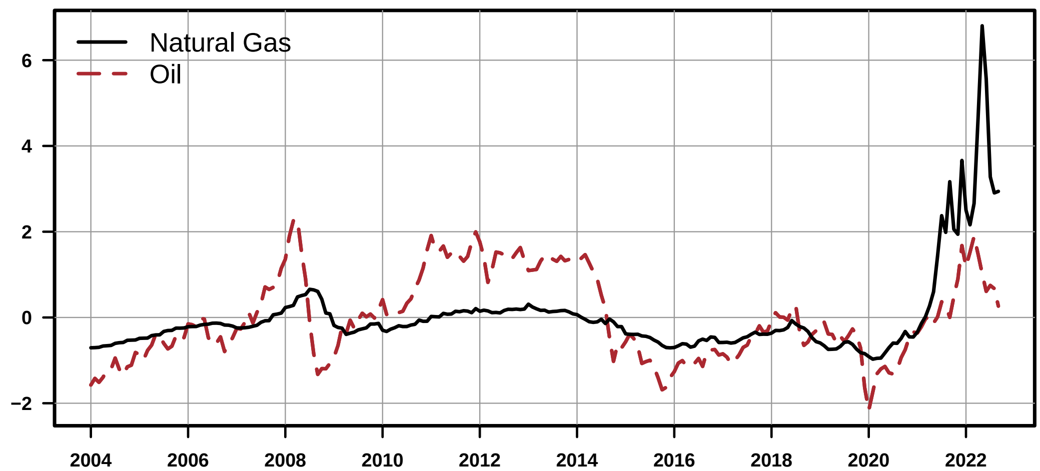
Natural gas is an important energy source for the European Union, where it represents 21.5% of the primary energy consumption and is the dominant source of energy for households with 32.1% (European Union Agency for the Cooperation of Energy Regulators, 2023). Moreover, it is considered a crucial energy resource during the transition to a green economy and therefore should gradually replace other fossil fuels.[7] While demand is growing, the EU27 has almost halved its domestic production over the last decade due to resource depletion and environmental concerns over a major European gas field, the Groningen field (U.S. Energy Information Administration, 2020; U.S. Energy Information Administration, 2022). Import-dependency has risen to about 80% in 2021, where the largest supplies come from Russia, Norway, and Algeria. Europe, particularly Germany, imports about 41% of the supply from Russia, which is not only rich in natural gas resources but also has the necessary infrastructure (European Union Agency for the Cooperation of Energy Regulators, 2023; Eurostat, 2023). Prior to the Russian invasion of Ukraine, this facilitated the flow of cheap energy reflected in the low volatility of the natural gas price, as shown in Figure 2.[8]

The natural gas market has undergone a process of liberalization over the last two decades. The transformation of natural gas markets through deregulation, including the development of mature gas trading hubs in Europe, has led to an overall increase in liquidity, transparency and thus competition (Heather, 2020). This has led to increasing market integration among European economies (Bastianin, Galeotti and Polo, 2019; Broadstock, Li and Wang, 2020). According to Szafranek and Rubaszek (2023), these changes have led to a gradual but significant shift from oil price indexation to gas-on-gas competition in opposite directions. Between 2005 and 2020, the latter increased from 15% to 80%, while the former decreased from 78% to

[7] The European Union has qualified natural gas as a transitional energy source, less carbon-intensive compared to coal and oil, to facilitate the green transformation through its Green Deal(European Commission, 2019). This strategy is also reflected in the European Commission's recent reclassification of natural gas as a green energy source during the green transition process (European Commission, 2021). During the expansion of renewable energy sources, natural gas, while still a fossil fuel, produces lower emissions and less air pollution compared to other hydrocarbons such as oil or coal.

[8] In a recent study, Chen et al. (2023) examine the structure of return volatility in the natural gas market. They find a clear structural break in the volatility pattern with the onset of Russia's invasion of Ukraine.

Figure 2: Real Natural Gas Prices and Real Oil Prices.



Notes: The figure shows the standardized real natural gas price in Euro per mmbtu (transformed into Euro from US dollar and deflated with the harmonized index of consumer prices), and the standardized real crude oil price (Brent). Vertical axis denotes standard deviations around long-run mean (zero). Sample period ranges from 2004M1 to 2022M12. Horizontal axis denotes time in months.

20%. This can be seen as a sign of the decoupling of gas and oil markets.[9] In addition, the European natural gas market has intensified its connectedness of market subsegments over the past decade (Zhang and Ji, 2018; Papież et al., 2022; Szafranek et al., 2023). Unlike the US, Europe has not developed large-scale shale gas production. A complete decoupling of natural gas from oil markets in Europe is therefore unlikely (Szafranek and Rubaszek, 2023).

These structural differences also explain the divergence of natural gas prices across geographic locations. For example, at the height of uncertainty immediately following the start of the Russian invasion, the US benchmark, Henry Hub, was quoted well below the European reference price, the Dutch TTF (Title Transfer Facility). On August 26, 2022, the peak price on the TTF spot market was just below 350 EUR/MWh, while the Henry Hub benchmark was quoted at 32.35 USD/MWh (or 32.47 EUR/MWh) on the same day. The US is meeting its own demand either through fracking or through the development of large standard gas fields. By now, the US has even become a net exporter of natural gas in recent years due to the shale gas revolution, which has also reduced its exposure to international supply disruptions (Geng, Ji and Fan, 2016;

[9] Along with the general increase in the efficiency of production processes and the use of alternative energy resources in line with the goals of the green transition, this observation is consistent with the finding that the sensitivity of real variables to oil price fluctuations has decreased over time (Blanchard and Gali, 2009; Baumeister and Peersman, 2013). This decreasing sensitivity is also evident in the responses of both expected and realized inflation to oil price shocks (Harris et al., 2009; Wong, 2015; Coibion and Gorodnichenko, 2015a; Conflitti and Luciani, 2019; Aastveit, Bjørnland and Cross, 2023).

Huang and Etienne, 2021). This suggests a *locality* of the market due to the more static and thus less flexible infrastructure (e.g. pipelines) required for transportation. This makes it difficult to find alternative suppliers, especially at short notice, leaving Europe particularly vulnerable to supply disruptions.[10]

Finally, a significant portion of natural gas in all countries is used not only as a direct energy resource, but also as an input to a wide range of production processes with potentially very limited substitutability.[11] Given the reliance on natural gas, severe disruptions or unexpected supply constraints can threaten not only price stability but also economic growth. Several studies have analyzed the impact of gas supply disruptions on the economy, mostly focusing on Germany. For example, Nick and Thoenes (2014) find that a natural gas supply disruption has a significant impact on the German economy. More recently, given the actual threat of a sudden supply disruption in light of the evolving geopolitical situation, a number of studies find that for Germany, a sudden Russian supply disruption would have large and persistent price effects, but output would react only moderately (see, e.g., Bachmann et al., 2024; Krebs, 2022; Güntner, Reif and Wolters, 2024). From a more global perspective, Albrizio et al. (2022) and Emiliozzi, Ferriani and Gazzani (2024) conducted similar analyses for Europe, explicitly including liquified natural gas (LNG). Their results show that greater EU integration into global LNG markets would create a price buffer, but other LNG-dependent countries would experience higher prices.

3. Empirical Methodology

To model the effects that real natural gas price shocks exert on expected and actual inflation, we develop a semi-structural model to disentangle the sources of variation in the price of natural gas, inflation expectations, inflation, and demand-side fluctuations. In a next step, we construct a structural scenario analysis in which inflation expectations are insensitive to movements in the real price of natural gas. This allows us to differentiate between first- and second-round effects. Our focus lies primarily on the identification of real natural gas price shocks transformed into the domestic currency. For the counterfactual analysis, we also identify an idiosyncratic inflation expectations shock. We extend the proposed structural models of Wong

[10] The same is true for the liquefied version of natural gas, LNG. While it is liquid and therefore easier to transport by ship, it also requires special infrastructure for regasification. For the time being, LNG simply cannot meet all of Europe's gas needs.

[11] According to the U.S. Energy Information Administration (2023), both electric power generation and the industrial sector account for over 70% of U.S. natural gas demand in 2021. In Europe, the residential sector accounts for the majority of natural gas demand, followed by power generation and the industrial sector. However, between 2000 and 2020, consumption by the industrial sector will decrease by 20%, with a 15% shift to power generation. Over time, the EU27 demand profile has changed significantly, reflecting the switch from coal to natural gas and the measures foreseen by the green transformation (European Union Agency for the Cooperation of Energy Regulators, 2023).

(2015) and Kilian and Zhou (2022a) and add monthly real GDP and the shadow rate (Wu and Xia, 2016) as a proxy of monetary policy to the model.

The econometric model is estimated on a monthly data frequency with a sample starting in January 2004 and ending in December 2022. The model features five variables $\boldsymbol{y}_t = [rgas_t, \Delta rgdp_t, sr_t, \pi_t, \pi_t^e]$, where $rgas_t$ denotes the log level of the real natural gas price (transformed into Euro from US dollar and deflated with the harmonized index of consumer prices, HICP), $\Delta rgdp_t$ the interpolated month-on-month growth rate of euro area real GDP, sr_t shadow short-term interest rate of the euro area by Wu and Xia (2016), π_t consumer price (year-on-year) headline inflation based on the HICP, and π_t^e inflation expectations measured through inflation swaps.[12] For the interpolation of real GDP, we use the Chow-Lin temporal disaggregation method using industrial production as input. Since our sample period covers the Covid-19 pandemic, various strategies have been proposed to address the outliers in this period. We follow the strategy of Cascaldi-Garcia (2022) and introduce dummy observations for the months of March to May 2020. An overview of the exact variable definitions, transformations, and sources is available in Appendix A.

3.1 The Semi-Structural VAR Model

We proceed with our structural model where the reduced-form VAR model representation is

$$\boldsymbol{y}_t = \boldsymbol{\Phi}_1 \boldsymbol{y}_{t-1} + \ldots + \boldsymbol{\Phi}_p \boldsymbol{y}_{t-p} + \boldsymbol{u}_t, \qquad \boldsymbol{u}_t \sim \mathcal{N}(\boldsymbol{0}, \boldsymbol{\Sigma}), \tag{3.1}$$

where $\boldsymbol{y}_t$ is an $n \times 1$ vector of macroeconomic variables, which are modeled as a function of their own past values (through the $n \times n$ coefficient matrices $\boldsymbol{\Phi}_j$, $j = 1, \ldots, p$) , and an $n \times 1$ vector $\boldsymbol{u}_t$ of forecast errors with an $n \times n$ covariance matrix $\boldsymbol{\Sigma}$. For the sake of brevity, Eq. (3.1) omits any possible deterministics such as the intercept and dummy variables (Cascaldi-Garcia, 2022). We allow up to $p = 12$ lags to enter the equation, accounting for the long and variable lags in the transmission of commodity price shocks (Hamilton and Herrera, 2004). We pursue a Bayesian approach to estimation as done in Chan (2022) but with a variant of global-local shrinkage priors. Specifically, we use the Normal-Gamma prior outlined in Huber and Feldkircher (2019). A detailed discussion of the estimation routine and the prior specification is provided in Appendix B. We sample 35,000 draws from the posterior distribution, from which we discard the first 10,000 as burn-ins. Finally, we use a thinning factor of 2, meaning we keep every second posterior draw.

[12] As noted above, these derivatives also contain an inflation risk premium. Moreover, the liquidity of the ILS market might pose additional challenges, potentially giving rise to a liquidity premium. Reis (2021), for instance, argues that the inflation swap market was only reasonably liquid starting in 2009. However, the later presented results are robust if we start the analysis in 2010. This further alleviates possible concerns that the global financial crisis is driving the effects.

The reduced-form shocks are a linear combination of n orthogonal structural disturbances $\boldsymbol{\varepsilon}_t$, which we write as $\boldsymbol{\varepsilon}_t = \boldsymbol{A}\boldsymbol{u}_t$. The matrix $\boldsymbol{A}$ governs the contemporaneous relationships between the endogenous variables. The structural VAR equation thus reads

$$\boldsymbol{A}\boldsymbol{y}_t = \boldsymbol{B}_1\boldsymbol{y}_{t-1} + \ldots + \boldsymbol{B}_p\boldsymbol{y}_{t-p} + \boldsymbol{\varepsilon}_t, \qquad \boldsymbol{\varepsilon}_t \sim \mathcal{N}(\mathbf{0}, \boldsymbol{I}_n), \tag{3.2}$$

where $\boldsymbol{\Phi}_j = \boldsymbol{A}^{-1}\boldsymbol{B}_j$ $(j = 1, \ldots, p)$ holds. We define $\boldsymbol{H} = \boldsymbol{A}^{-1}$ to be the structural impact matrix. By definition, structural shocks are mutually uncorrelated, i.e., $\mathrm{Var}(\boldsymbol{\varepsilon}_t) = \boldsymbol{I}_n$ being diagonal, and are thus identified up to a sign and scale convention. From the linear mapping of the shocks, $\boldsymbol{\Sigma} = (\boldsymbol{A}\boldsymbol{A}')^{-1}$ holds such that the identification problem is finding a suitable structural impact matrix $\boldsymbol{H}$.

We are interested in the identification of real natural gas price shocks, which are unpredictable surprises to the price of natural gas from the perspective of the European economy. We abstain from distinguishing between supply and demand forces on the natural gas market and thus the real natural gas price shock captures both natural gas supply and gas-specific demand/inventory shocks. As seen in Figure 1, the political turmoil ensuing from the Russian invasion of Ukraine in 2022 led to distortions in the natural gas market driving up its price tremendously. However, the natural gas market is a much more localized market than the global oil market and thus natural gas prices are not contemporaneously predetermined with respect to European variables. For this reason, we exploit a combination of sign and zero restrictions for the identification of a real natural gas price shock and avert timing assumptions. These restrictions are imposed on the structural impact matrix $\boldsymbol{H}$, as shown in Eq. (3.3). We use the algorithm outlined in Arias, Rubio-Ramírez and Waggoner (2018). Specifically, we search for an orthonormal matrix $\boldsymbol{Q} = (\boldsymbol{q}_1, \ldots, \boldsymbol{q}_n)$, such that $\boldsymbol{Q}\boldsymbol{Q}' = \boldsymbol{I}$ holds. The algorithm searches for each column vector of the matrix $\boldsymbol{Q}$ recursively, conditional on the zero restrictions.[13] This yields the structural impact matrix $\boldsymbol{H} = \boldsymbol{L}\boldsymbol{Q}$, where $\boldsymbol{L}$ is the lower-triangular Cholesky factor of the variance-covariance matrix.

The main goal is the identification of the real natural gas price shock. For this shock, we assume that an increase in the real price of natural gas leads on impact to a surge of inflation. Rising energy prices increase the costs of production, which firms pass-through to final goods. On real GDP growth we impose a zero restriction to not allow for any demand-driven effects (giving rise to an upshot in real GDP) and to account for the long transmission of real natural gas price shocks to the real economy (Alessandri and Gazzani, 2023). We

[13] In more detail, we use the Algorithm 2 outlined in Arias, Rubio-Ramírez and Waggoner (2018) and not their proposed importance sampler in Algorithm 3, which extends Algorithm 2. We abstain from doing so because we depart from their normal-generalized-normal distribution. As they note, this is permissible with the drawback that the distribution is not invariant to a reordering of the shocks.

assume an immediate and positive response of the central bank through the shadow rate on impact. This can be justified by the shadow rate also capturing a broader perspective of monetary policy (e.g., unconventional policies) and by recent evidence that the central bank is capable of actively fighting inflation through rising energy prices (Ider et al., 2023). Lastly, we also assume an increase in inflation expectations to a real natural gas price shock given the evidence in Binder (2018).

We are also interested in the identification of the idiosyncratic inflation expectations shock. The importance of such a shock is highlighted in Madeira and Zafar (2015). It is assumed not to affect the real price of natural gas, real GDP, the shadow rate, and inflation on impact. Hence, all movements in expectations that impact actual consumer prices are then already captured by the remaining shocks. These restrictions are also consistent with sign restriction approaches on other energy-related studies, like on gasoline prices as provided in Kilian and Zhou (2022a).

While a fully identified system is not necessary for our research purpose here, it can improve inference even if some shocks are not essential for the analysis (Canova and Paustian, 2011). We thus identify European a demand, residual supply, and monetary policy shock. The imposed sign restrictions are relatively standard. The residual supply shock is assumed to raise inflation and inflation expectations and to lower industrial production on impact. However, we do not restrict the impact on the interest rate as the central bank may see through possible non-energy related, residual supply shocks (e.g., a markup shock or other production cost-related shocks). The demand shock is assumed to raise real natural gas prices, inflation, inflation expectations, and economic activity. This shock accounts for demand-side effects on the real natural gas price shock (e.g., higher demand of natural gas due to heating purposes). For the monetary policy shock, we assume that the central bank is following a Taylor rule: on impact, a monetary policy shock decreases the real natural gas price, real GDP, and inflation. In addition and consistent with recent developments to take expectations explicitly into account (e.g., mentioned in central banker's speeches, like Mester, 2022; Lane, 2022c), inflation expectations are assumed to decrease after such a shock.

Jointly, these restrictions imply

$$\begin{pmatrix} u_t^{rgas} \\ u_t^{\Delta rgdp} \\ u_t^{sr} \\ u_t^{\pi} \\ u_t^{\pi^{exp}} \end{pmatrix} = \begin{bmatrix} + & + & - & - & 0 \\ 0 & + & - & - & 0 \\ + & + & + & * & 0 \\ + & + & - & + & 0 \\ + & + & - & + & + \end{bmatrix} \begin{pmatrix} \varepsilon_t^{\text{real gas price shock}} \\ \varepsilon_t^{\text{demand-side shock}} \\ \varepsilon_t^{\text{monetary policy shock}} \\ \varepsilon_t^{\text{residual supply-side shock}} \\ \varepsilon_t^{\text{idiosyncratic inflation expectation shock}} \end{pmatrix}, \tag{3.3}$$

where + denotes a positive, − a negative reaction, and 0 a zero restriction. A star $*$ indicates that we impose no sign restriction on impact.

3.2 Structural Scenario Analysis Counterfactuals

If the real natural gas price shock causes movements in inflation expectations that subsequently pass-through to inflation, we define this as the second-round effect. Note, that even if inflation expectations rise in response to a real gas price shock, this does not automatically imply that second-round effects are at work. To investigate the second-round effects, we construct a counterfactual in which inflation expectations are insensitive, thereby isolating first-round effects. We feed a sequence of inflation expectations shocks into the model to mute the inflation expectations response. We use the framework of Antolin-Diaz, Petrella and Rubio-Ramirez (2021) adapted to the case of impulse response analysis by Breitenlechner, Georgiadis and Schumann (2022) to construct the structural scenario analysis counterfactuals.

The unconditional forecast of the observed variables in the VAR, denoted with the $nh \times 1$ vector $\boldsymbol{y}_{T+1,T+h} = (\boldsymbol{y}'_{T+1}, \boldsymbol{y}'_{T+2}, \ldots, \boldsymbol{y}'_{T+h})'$, can be written as

$$\boldsymbol{y}_{T+1,T+h} = \boldsymbol{b}_{T+1,T+h} + \boldsymbol{M}' \boldsymbol{\varepsilon}_{T+1,T+h}, \tag{3.4}$$

where the vector $\boldsymbol{b}_{T+1,T+h}$ is predetermined and depends on the full history of the observables and the reduced-form parameters. In the absence of any future shocks, $\boldsymbol{b}_{T+1,T+h}$ denotes the dynamic forecast of the system. The $nh \times 1$ vector $\boldsymbol{\varepsilon}_{T+1,T+h} = (\boldsymbol{\varepsilon}'_{T+1}, \boldsymbol{\varepsilon}'_{T+2}, \ldots, \boldsymbol{\varepsilon}'_{T+h})'$ thus denotes all future values of the structural shocks. Lastly, the $nh \times nh$ matrix $\boldsymbol{M}$ constitutes the dynamic propagation of future structural shocks and is a function of the structural VAR parameters. Note that if the VAR is stationary, in steady state at T, $\boldsymbol{b}_{T+1,T+h} = \boldsymbol{0}$, and if there is only a single future shock $\boldsymbol{\varepsilon}_{T+1,T+h} = (\boldsymbol{e}'_1, \boldsymbol{0}_{n(h-1)\times 1})$, then $\boldsymbol{M}$ reflects the usual impulse response functions to a unit shock. $\boldsymbol{e}_i$ denotes the unit vector with unity on the i-th position. For instance, for the impulse responses to a real gas price shock, we have $\varepsilon_{1,T+1} = 1$, $\varepsilon_{1,T+s} = 0$ for $s > 1$ and

$\varepsilon_{j,T+s} = 0$ for $s > 0$ and $j \neq 1$. We denote this in the following as the *unconditional* impulse response function.[14]

In the framework of Antolin-Diaz, Petrella and Rubio-Ramirez (2021), the structural VAR parameters captured in $\boldsymbol{M}$ remain *unchanged* in the counterfactual. Hence, the analysis does not risk falling into the criticism put forward by Lucas (1976) as long as the structural shocks used to construct the counterfactuals are not *too unusual*. We use the modesty statistic proposed by Leeper and Zha (2003) and the q-divergence distribution proposed in Antolin-Diaz, Petrella and Rubio-Ramirez (2021) to safeguard us against these concerns. In Appendix C, we provide the details on how to implement these tests. In order to satisfy the imposed constraints on the impulse response $\tilde{\boldsymbol{y}}_{T+1,T+h}$, additional shocks are allowed in $\tilde{\boldsymbol{\varepsilon}}_{T+1,T+h}$ to materialize over the impulse response horizon. We choose those values such that we offset the effects of inflation expectations to a real natural gas price shock.

We implement the constraints on the paths of one endogenous variable (i.e., inflation expectations) in $\tilde{\boldsymbol{y}}_{T+1,T+h}$ as follows

$$\overline{\boldsymbol{C}}\tilde{\boldsymbol{y}}_{T+1,T+h} = \overline{\boldsymbol{C}}\boldsymbol{M}'\tilde{\boldsymbol{\varepsilon}}_{T+1,T+h} \sim \mathcal{N}\left(\overline{\boldsymbol{f}}_{T+1,T+h}, \overline{\boldsymbol{\Omega}}_f\right), \tag{3.5}$$

where $\overline{\boldsymbol{C}}$ is a $k_o \times nh$ selection matrix, $\overline{\boldsymbol{f}}_{T+1,T+h}$ is a $k_o \times 1$ vector, and $\overline{\boldsymbol{\Omega}}_f$ a $k_o \times k_o$ matrix. $\overline{\boldsymbol{f}}_{T+1,T+h}$ and $\overline{\boldsymbol{\Omega}}_f$ are the mean and covariance matrix restrictions. This formulation also accommodates the special case $\boldsymbol{\Omega}_f = \boldsymbol{0}$, which we will adopt. This resembles the classic "hard" conditional forecasting exercise as defined in Waggoner and Zha (1999). Here, we impose the restriction that the inflation expectations are insensitive to real natural gas price shocks. The constraints on the structural shocks are given by

$$\boldsymbol{\Xi}\tilde{\boldsymbol{\varepsilon}}_{T+1,T+h} \sim \mathcal{N}\left(\boldsymbol{g}_{T+1,T+h}, \boldsymbol{\Omega}_g\right), \tag{3.6}$$

where $\boldsymbol{\Xi}$ is a $k_s \times nh$ selection matrix. $\boldsymbol{g}_{T+1,T+h}$ is a $k_s \times 1$ vector and $\boldsymbol{\Omega}_g$ is a $k_s \times k_s$ matrix and denote the mean and covariance matrix restrictions. Again, we implement exact restrictions and fix $\boldsymbol{\Omega}_g = \boldsymbol{0}$. Here, we use the structural idiosyncratic inflation expectation shock as the offsetting force such that the impulse response to inflation expectation to real natural gas price shocks is zero. To do so, we impose that all structural shocks are zero over the whole impulse response horizon except the structural shock to natural gas prices in the first period and the structural shocks to inflation expectation along the impulse response horizon.

[14] Technically, the impulse response function is *conditional* on a shock in the first period. Nevertheless, we deem the term appropriate since both – the baseline impulse response and the counterfactual impulse response – are *conditional* on a shock in the first period. Hence, we distinguish between *conditional* counterfactual impulse responses and *unconditional* impulse responses to a shock in the first period.

Antolin-Diaz, Petrella and Rubio-Ramirez (2021) show how to obtain the solution in terms of $\tilde{\boldsymbol{\varepsilon}}_{T+1,T+h}$, which satisfies the constraints in Equation (3.5) and Equation (3.6). The counterfactual impulse response is then given by $\tilde{\boldsymbol{y}}_{T+1,T+h} = \boldsymbol{M}'\tilde{\boldsymbol{\varepsilon}}_{T+1,T+h}$. We refer to Appendix C for further technical details.

4. Results

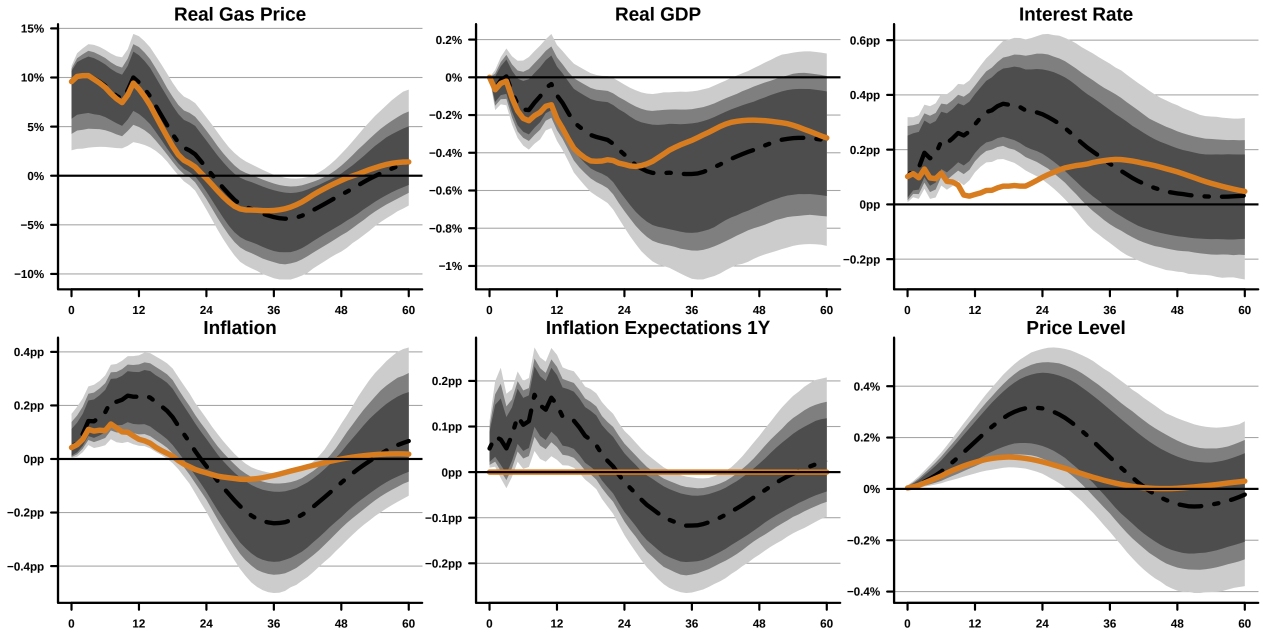
In this section, we discuss the effects of a real natural gas price shock and investigate the second-round effects via inflation expectations and the role of their respective horizon. In terms of shock sizes, we standardize it to a one standard deviation increase in the real price of natural gas. We accumulate the responses of real GDP growth, which allows us to interpret the effects in percent deviations from the pre-shock level. From the impulse response of annualized inflation, we back out price level deviations by accumulating the effects. In all specifications, we report the median impulse response functions (IRFs) along with their 68/80/90 percent confidence bands. The black, dashed lines always denote the median IRF, while the orange, solid lines report the counterfactual IRF.

4.1 The Effects of Real Natural Gas Price Shocks

Figure 3 presents the main results. The real natural gas price shock raises real gas prices by about 10% on impact. The reaction is persistent for about one year, before gradually returning to its pre-shock level. Real GDP reacts only with a strong lag. The reaction is insignificant up to twelve months, before a delayed economic downturn is visible. Inflation and inflation expectations increase and show strong mean-reverting behavior with an undershooting after two years. Monetary policy reacts contractionary and turns more restrictive, reaching a peak increase of the interest rate after eighteen months.

Our interest lies on the transmission of real natural gas price shocks to inflation. In terms of magnitudes, we observe an 8 bps increase in inflation to a one standard deviation real natural gas price shock on impact. At maximum, inflation rises by about 22 bps after one year. This aligns well with the corresponding weight in the consumption basket of the HICP with about 1.29% for natural gas, translating a 10% increase in the natural gas price to an increase in the HICP of about 13 bps. The hump-shaped IRF shows considerable stronger effects at longer horizons, pointing towards effects beyond the first-round channel. The results are quite consistent with evidence from microdata provided by Lafrogne Joussier, Martin and Mejean (2023), who examine the cost pass-through to inflation to energy price shocks in the French manufacturing sector.

Figure 3: Impulse Response Functions to a Real Natural Gas Price Shock.



Notes: Real natural gas price shock is identified with sign and zero restrictions and standardized to a one standard deviation increase in the real price of natural gas. Real GDP is the cumulative response of real GDP growth. The price level is computed afterwards as cumulative sum of the inflation response. Black dashed lines denote the posterior median responses while gray shaded areas depict the 68/80/90 percent confidence intervals. The orange solid line denotes the counterfactual response. The vertical axis denotes the effect sizes of the real gas price, real GDP and the price level in percent, while the interest rate, inflation and inflation expectations are scaled to annualized percentage points. The horizontal axis denotes the impulse response horizon in months.

These effects remain robust in shape and magnitude to a number of specification choices as discussed in the robustness section further below.[15]

Inflation expectations also react positively to a real natural gas price shock over a sustained time period. This is consistent with the empirical evidence that inflation expectations are sensitive to commodity price shocks (Harris et al., 2009; Coibion and Gorodnichenko, 2015b; Aastveit, Bjørnland and Cross, 2023). In terms of magnitudes, inflation expectations react less pronounced than the inflation series. This implies a persistent positive forecast error of inflation for about two years.[16] This is consistent with prior studies that document an underreaction of inflation expectations (Coibion and Gorodnichenko, 2012; Coibion and Gorodnichenko, 2015a).

[15] In the appendix, Figure D2 reports the comparison to the baseline model. The specification is robust to various perturbations.
[16] We back out the implied forecast error impulse response function of inflation (constructed as the difference between realized inflation and the previous year's 1-year expected inflation) to validate this claim, as shown in Figure D1.

4.2 Second-Round Effects of Inflation Expectations

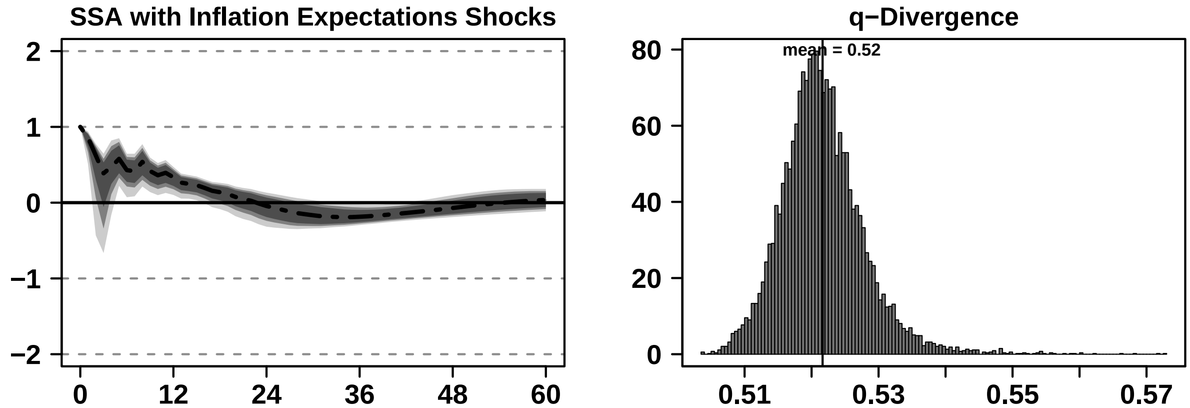
We construct a structural scenario analysis in which inflation expectations are insensitive to shocks of the real natural gas price. This allows us to investigate whether movements in inflation expectations exert a pass-through to inflation. To differentiate between this first- and second-round effects, we back out the differential response of inflation. To visualize the results of this exercise, the solid orange lines in Figure 3 depict the counterfactual impulse response, where the impulse response of inflation expectations is zero at all horizons. Note that we only use the structural shocks of inflation expectations to construct a sequence of shocks to offset this response. Put differently, only the idiosyncratic inflation expectations shock is allowed to deviate from its unconditional impulse response, eventually changing the dynamics of the whole system while maintaining the estimated structural relationships.[17]

When inflation expectations are insensitive to movements in the real price of natural gas, we observe that the response of inflation (and the implied response of the price level) is much more muted, pointing to strong adjustments via second-round effects. While the conditional and unconditional response of inflation do not deviate from each other on impact, they start to diverge more strongly after about six months. After one year, inflation is 15 bps lower in the counterfactual. The pass-through is close to but clearly below unity. Interest rates show an attenuated pattern as well, pointing to an attentive monetary authority towards second-round effects. This highlights that policymakers are monitoring inflation expectations closely in the face of global energy shocks (see, for instance, Schnabel, 2022b). Lastly, both the real natural gas price and the real GDP response do not deviate strongly from their unconditional response. This is reassuring because this means that neither the nature of the shock is strongly changed in the counterfactual nor demand-side effects seem to play a strong role.

Overall, our results stand in stark contrast to the findings of the literature for other commodity price shocks. Wong (2015), for instance, uses a smaller model and conducts an analysis for the US and for oil price shocks. Nevertheless, he finds only limited evidence for second-round effects of inflation expectations and concludes that the US offers an environment where inflation expectations are well anchored. Further evidence for that is provided by Kilian and Zhou (2022b) who investigate the increase in oil and gasoline prices since mid-2020. They provide evidence that these kinds of shocks have not moved long-run household inflation

[17] We relax this assumption in a robustness exercise. Instead of using one particular shock (*the idiosyncratic inflation expectations shock*) as the offsetting force, we use the combination of all other shocks to offset the response of the inflation expectations series. For more details, we refer to the robustness section and the presentation of the robustness exercise in Figure D2. Outcomes are robust to this choice.

Figure 4: Plausibility Statistics of Counterfactual Scenario.

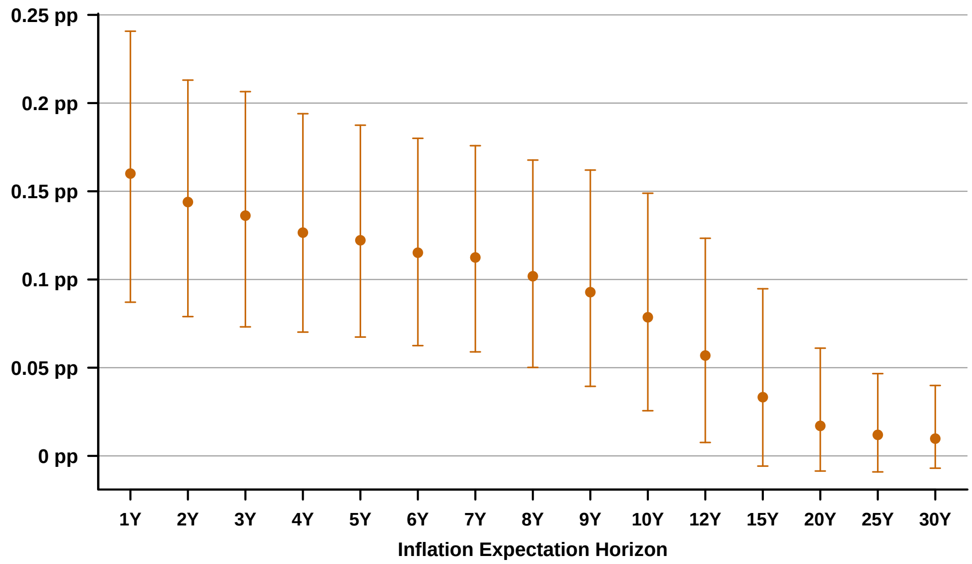


Notes: The left figure shows the modesty statistic of Leeper and Zha (2003) and the right figure shows the distribution of the q-divergence proposed by Antolin-Diaz, Petrella and Rubio-Ramirez (2021). The modesty statistic reports the implied shocks that impose the counterfactual constraint for inflation expectations. The black dashed line denotes the posterior median responses while gray shaded areas depict the 68/80/90 percent confidence intervals.

expectations. We will return to these points when comparing the results to the US in several extensions below. Lastly, we assume that the inflation risk premium is not time-varying in this analysis and thus not a major driver of our findings. To alleviate possible concerns, we will return to this point later on when re-doing the analysis with inflation expectations originating from the survey of professional forecasters.

The plausibility of the counterfactuals obtained by the structural scenario analysis depends on the offsetting structural shocks, i.e., the idiosyncratic inflation expectation shock. Specifically, we risk falling into the criticism by Lucas (1976) if the required shocks are unusually large or persistent. Under such a situation, agents may update their beliefs about the policy regime and the structure of the economy more substantially. Against this backdrop, we implement the modesty statistic of Leeper and Zha (2003) and the q-divergence proposed by Antolin-Diaz, Petrella and Rubio-Ramirez (2021), which are presented in Figure 4. The left figure shows the modesty statistic, which denotes the implied offsetting shocks that impose the counterfactual constraint for inflation expectations. The offsetting shocks are *modest* if the statistic is smaller than two in absolute values. This is confirmed and thus the materialization is unlikely to induce agents to adjust their expectation formation and beliefs about the structure of the economy leaving no room for the Lucas critique. In the right graph, the q-divergence indicates how strongly the distribution of offsetting shocks in the counterfactual deviates from their unconditional distribution translated into a comparison of the binomial distribution of a fair and a biased coin. Again, the test does not indicate that the distribution of offsetting shocks in the counterfactual is notably different from the unconditional distribution.

Figure 5: Effect Sizes for Inflation with Varying Horizon of Inflation Expectations.



Notes: Maximum difference between the unconditional and the counterfactual impulse response function of inflation in percentage points (pp) in the model identified with sign and zero restrictions. Dots refer to the median of the maximum response, while the whiskers denote the 68 percent confidence region. The horizontal axis denotes the inflation expectations horizon in years.

4.3 Effects Along the Term Structure of Inflation Expectations

We exploit the availability of ILS inflation expectations along the term structure (up to thirty years ahead). We re-estimate the model identified with sign and zero restrictions exchanging the series of inflation expectations along the maturity horizon. Note that we exchange the measure of inflation expectations once at a time and do not pursue estimating a model including all the horizons.

We start with short-run expectations of one year ahead (used in the baseline model) and move along until we reach long-run inflation expectations (30 years ahead). Then, for each estimated model we pick the maximum difference between the unconditional and the counterfactual impulse response of inflation. For example, we can directly compare the outcome of the maximum difference of 1Y inflation expectations to the difference in Figure 3. Here, the maximum difference between the impulse responses occurs after 24 months with about 15 bps. Put differently, without second-round effects via the expectation channel, inflation is 15 bps (annualized) lower on average. Additionally, we also report confidence sets for the differences such that the whiskers in Figure 5 are the full 68 % confidence interval of the differences' posterior distribution.

Figure 5 reveals several interesting results. First, the maximum differences between the unconditional and counterfactual impulse responses are statistically and significantly different from zero until 12Y ahead.

Second, the median difference response using short-term (i.e., one year) inflation expectations exhibits the highest difference. Although the effects gradually decrease from short- to long-term horizons, as expected, differences are not statistically significant. The differences in the medium-term expectations, from four to eight years, remain rather constant, while the differences start to decline more strongly after 10Y ahead. The overall picture points to a downward shift of the short- and medium-term inflation expectations, while long-term expectations do not move. Starting with eight years ahead, the maximum effect of the counterfactual exercise is already below 10 bps. Furthermore, but not visible in the plot, the maximum responses are usually reached after about one and half to two years, typically quicker for shorter-term expectations. Overall, the picture shows that the channel via inflation expectations accounts for 10-15 bps of inflation. While our approach does not distinguish between movements of inflation risk premia in market-based inflation expectations, we conclude that the second-round effects of real natural gas price shocks are non-negligible over the near and medium term structure of expectations.

4.4 Discussion of the Results

Summing up the results so far, we show that real gas price shocks have inflationary tendencies both via first- and second-round effects. The counterfactual analysis reveals strong second-round effects through inflation expectations. However, the time frame of the sample is crucial for the results, which is already indicated in Figure 1. If we exclude the period of the recent natural gas price surge, we do not obtain these pronounced reactions.[18] Furthermore, our findings indicate that second-round effects are strong and relevant for short-term inflation expectations, before turning near zero and insignificant for long-term expectations. This is consistent with evidence that short-term expectations are more important compared to longer ones in determining inflation (Fuhrer, 2012; Fuhrer, Olivei and Tootell, 2012).

Still, a few questions about the interpretation of the results remain. For instance, the results point to the fact that inflation expectations in the euro area are rather sensitive to natural gas price shocks or, at least, turned so in the last year. Why are these reactions comparatively strong and why do they actually drive inflation? Three possible – however, not mutually exclusive – interpretations offer an explanation. The first concerns issues around the anchoring of inflation expectations in the euro area. A second interpretation points towards

[18] For this exercise, we split the sample before the onset of the pandemic (end of December 2019) and before the recent gas price surge (end of June 2021). In both cases, natural gas price shocks do not reveal strong effects on inflation and inflation expectations. This holds specifically for natural gas prices and, to a lesser extent, for oil prices. Results are available from the authors upon request.

the particularities of expectation formation processes. Finally, there could also be demand-side forces outside of our framework at work that affect inflation expectations.

With respect to the first question, we expect inflation expectations to not react strongly in an environment in which they are well-anchored (Reis, 2021; Carvalho et al., 2023). Monetary policy authorities put an emphasis on managing inflation expectations to ultimately stabilize inflation through various factors. As a result, inflation should thus not respond beyond the cost channel, or first-round effects. Our results, however, allow the interpretation that expectations are susceptible to real natural gas price shocks. We find substantial second-round effects, indicating a sizable pass-through to inflation. Effects are strongest for short-term expectations and relatively small for longer-term expectations, pointing to minor concerns regarding de-anchoring risks. Nevertheless, policymakers at central banks are closely monitoring potential de-anchoring risks (Lane, 2022b; Schnabel, 2022b).

Second, the expectation formation process of inflation expectations may be distorted in a way that it does not resemble rational expectation. A wide array of papers has shown that agents, may it be firms or households, are informationally constrained when forming inflation expectations, which holds true independently of how inflation expectations are measured (Coibion and Gorodnichenko, 2012; Coibion and Gorodnichenko, 2015a). We confirm this in our analysis with market-based expectations. Specifically, D'Acunto, Malmendier and Weber (2023) and Weber, Gorodnichenko and Coibion (2023) point out that information provided by policymakers is often ignored or wrongly interpreted by economic agents and that personal experience, human cognition or gender play a larger role for households in forming inflation expectations. In terms of monetary policy, the pervasiveness of information rigidities in the economy has led to the conclusion that an optimal policy should respond aggressively to fluctuations in inflation (Reis, 2009). All of the above-mentioned studies point to the fact that information frictions do not differ strongly between the US and the euro area. This empirical fact motivates our consecutive analysis, where we re-do the analysis for the United States.

Third, inflation expectations are also affected by additional demand-side forces outside of our model. While natural gas prices have strongly gained momentum after the Russian invasion of Ukraine in February 2022, they started to rise already in mid-2021. In this time period, economies around the world were recovering economically from the Covid-19 pandemic. Part of this recovery process were generous fiscal transfers and support to households and firms, which we do not take explicitly into account with our model stance. Thus, these fiscally induced demand forces are not considered by our identification, unless they are captured by real GDP. According to Coibion, Gorodnichenko and Weber (2021) inflation expectations are sensitive to fiscal considerations, such as taxes and government spending. Specifically, news about future debt leads to anticipatory inflation expectation reactions, both in the short and long run. For example, using a new consumer survey in the euro area, Georgarakos and Kenny (2022) show in a randomized control trial that a more positive assessment of fiscal interventions improves household expectations about income prospects or future access to credit and financial sentiment. Needless to say, this serves only as indirect evidence for an effect of inflation expectations. Still, we acknowledge that our approach can only partially filter out the effects of the various fiscal interventions during the Covid 19 pandemic, which we will thus leave for further research.

4.5 Robustness

We provide robustness to the baseline model in terms of the specification and the identification with results reported in the appendix, see Figure D2.

Regarding the specification, the outcomes are robust to varying the degree of lags (six instead of twelve lags), the choice of target inflation measure (core instead of headline inflation), the choice of the economic activity variable (industrial production instead of real GDP), and the transformation of the real natural gas prices (log-differences instead of logarithm). Overall, the conclusions from the benchmark model remain robust to these perturbations.

We relax several assumptions in the identification of the model. First, we also impose a negative sign on industrial production instead of a zero restriction. Second, we relax the assumption that monetary policy reacts immediately to the real natural gas price, residual supply, and the demand shock. Third, we only identify the real natural gas price shock and the idiosyncratic inflation expectations shock and abstain from identifying the remaining shocks. Fourth, we only identify the real natural gas price shock. In this case, we have to construct the structural scenario analysis differently. Instead of using one particular shock (*the idiosyncratic inflation expectations shock*) as offsetting force to construct the counterfactual, we use the combination of all other shocks to offset the response of the inflation expectations series. Here, the results suggest a stronger case for second-round effects, which can be traced back to using a variety of (non-identified) shocks. The details on the exact implementation are provided in Appendix D. The baseline model is robust to all these choices.

5. Extensions

We provide extensions along three lines. First, we re-do the analysis with the survey of professional forecasters (SPF) of the European Central Bank (ECB) on a quarterly frequency. Second, crude oil prices show a strong comovement (see the discussion in section 2). Hence, the second extension is based on the real price of oil instead of natural gas. The third extension tackles the aforementioned question if there is a different reaction to these shocks in the US.

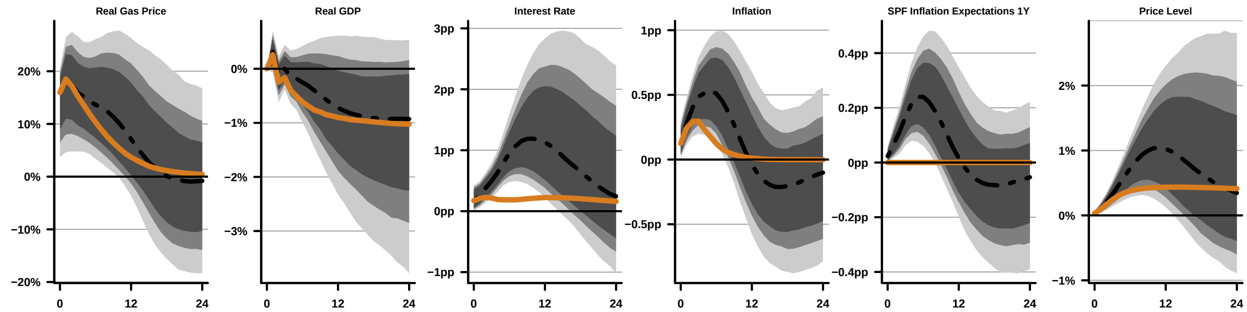
5.1 Inflation Expectations based on the Survey of Professional Forecasters

We re-do the analysis with the inflation expectations from the SPF of the ECB to provide robustness with a survey-based measure of inflation expectations, which compared to their marked-based counterpart does not feature liquidity risks or a latent risk premium component. However, the SPF data is only available on a quarterly frequency, which leaves us with only 72 observations from 2004Q1 to 2022Q4. We use the 12-month ahead and longer-term forecast (5 years) of HICP for the analysis.

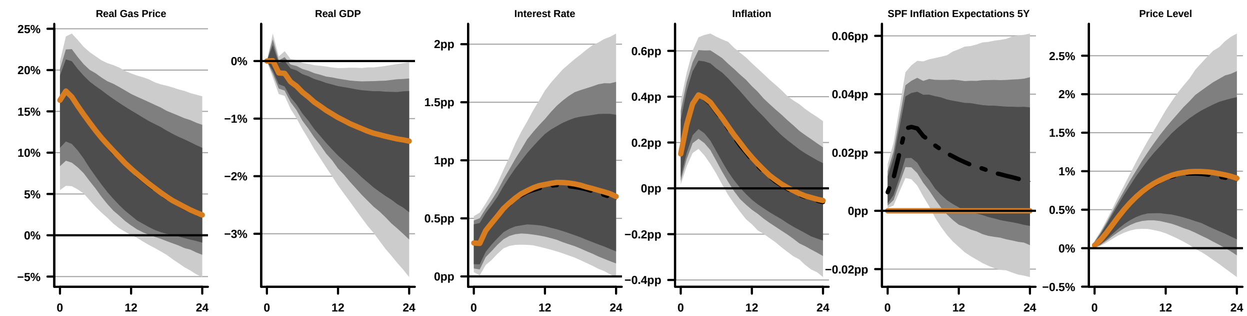
The results are presented in Figure 6. Overall, they confirm the picture presented so far. A real natural gas price shocks elicits a jump in the real price of natural gas of about 15%. Real GDP shows a protracted decline while the monetary authority raises interest rates. Inflation and its expectations increase. Notably, longer-term expectations increase much less pronounced. While 1Y ahead expectations peak at 25 bps after two years, 5Y ahead expectations only amount to a 3 bps reaction after one year at maximum. The counterfactual exercise (solid orange lines) reveals that the second-round effects are quite sizable in the model with short-term expectations but vanish for longer-term expectations. The pass-through in short-term expectations is now only about 0.4-0.5, while long-term expectations do not alter the transmission channel visibly. Again, in both models (with 1- and 5-year expectations) expectations clearly underreact to new information. We corroborate our earlier findings that the duration of an underreaction to information is about two years (as seen in Figure D1) which is consistent with earlier findings (Coibion and Gorodnichenko, 2012; Coibion and Gorodnichenko, 2015a). Furthermore, the uncertainty bounds are generally more sizable, which is due to the limited number of observations in the quarterly sample.

The SPF data confirms our initial findings, although second-round effects are already not visible anymore for 5Y ahead expectations. Interestingly, in the model with survey-based expectations, inflation and 1Y ahead inflation expectations show more sensitivity towards real natural gas price shocks, which yields a

Figure 6: Impulse Response Functions to a Real Gas Price Shock (SPF).



(a) SPF: 1Y Ahead Inflation Expectations.



(b) SPF: 5Y Ahead Inflation Expectations.

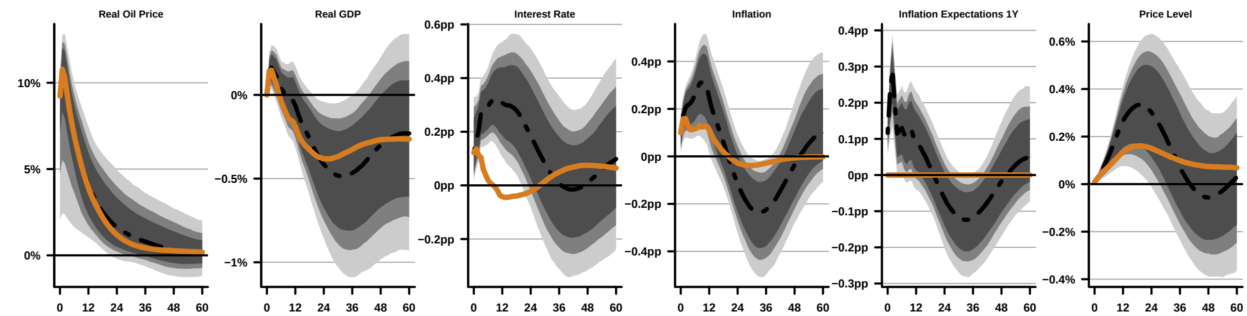
Notes: The model features inflation expectations from the survey of professional forecasters (SPF) on quarterly frequency. The upper panel uses 1-year ahead and the lower panel 5-year ahead inflation expectations. The shock is identified with sign and zero restrictions and standardized to a one standard deviation increase in the real price of natural gas. Real GDP is the cumulative response of real GDP growth. The price level is computed afterwards as cumulative sum of the inflation response. Black dashed lines denote the posterior median responses while gray shaded areas depict the 68/80/90 percent confidence intervals. The orange solid line denotes the counterfactual response. The vertical axis denotes the effect sizes of the real gas price, real GDP and the price level in percent, while the interest rate, inflation and inflation expectations are scaled to annualized percentage points. The horizontal axis denotes the impulse response horizon in quarters.

comparatively stronger effect on the price level. These findings confirm that short-term inflation expectations matter most for the pass-through of real natural gas price shocks. Long-term professional expectations, however, are more stable than market-based ones.

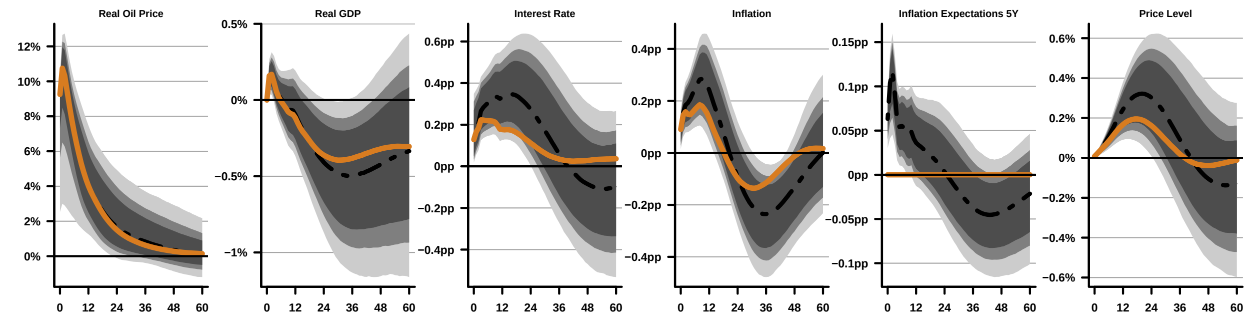
5.2 Real Oil Prices as External Shocks

We look into an alternative fossil fuel used either as energy or as a direct production input, thus also serving as a source of external shocks to energy prices in the euro area: crude oil. Oil prices are historically an interesting case and still – while with a vanishing effect – important for economic activity. Crude oil exhibits a strong comovement with natural gas (see Figure 2) with a correlation coefficient of 0.45. If we end the sample before mid-2021, the correlation coefficient amounts to 0.76. While natural gas prices were mostly determined by oil prices in the past, this is further evidence for decoupling tendencies of these markets

Figure 7: Impulse Response Functions to the Real Oil Price Shock.



(a) Alternative Commodity Price: Oil with 1Y Ahead Inflation Expectations.



(b) Alternative Commodity Price: Oil with 5Y Ahead Inflation Expectations.

Notes: The model for the real price of Brent oil features five variables, where the shock is identified with sign and zero restrictions and standardized to a one standard deviation increase in the real price of crude oil. The upper panel uses 1-year ahead and the lower panel 5-year ahead inflation expectations (ILS). Real GDP is the cumulative response of real GDP growth. The price level is computed afterwards as cumulative sum of the inflation response. Black dashed lines denote the posterior median responses while gray shaded areas depict the 68/80/90 percent confidence intervals. The orange solid line denotes the counterfactual response. The vertical axis denotes the effect sizes of the real oil price, real GDP and the price level in percent, while the interest rate, inflation and inflation expectations are scaled to annualized percentage points. The horizontal axis denotes the impulse response horizon in months.

(Szafranek and Rubaszek, 2023; Szafranek et al., 2023). We substitute the real natural gas price with the real price of crude oil (Brent). The identification strategy to isolate a real oil price shock resembles the same sign and zero restrictions as before.

The findings are presented in Figure 7. The oil price shock shows similar dynamics as in the baseline model. A one standard deviation shock elicits a jump in oil prices of about 10% on impact. Inflation and inflation expectations increase to the real oil price shock. We observe an increase of 10 bps for inflation and inflation expectations on impact in the model with short-run expectations. Inflation expectations react less pronounced for 5Y ahead expectations. This is comparable to the baseline model, pointing to similar direct effects. However, the dynamic reaction of inflation expectations are different, showing no hump-shaped response. Industrial production shows a protracted decline and interest rates a contractionary monetary policy stance.

Turning to the counterfactual exercise, i.e., the orange lines in Figure 7, the outcomes are again qualitatively similar to the baseline model. Second-round effects are visible for inflation and the implied price level for both short- and long-term expectations. Compared to the baseline model, second-round effects are smaller in magnitude and not significant for long-term expectations. The implied price level is about 40-50% lower than in the unconditional response (moving from a 0.23% increase to less than a 0.18% increase).

This shows that also real oil price shocks can account for sizable second-round effects, mostly transmitted through short-term expectations. Effects are a magnitude smaller than in the model with real natural gas prices. This corroborates the findings of Aastveit, Bjørnland and Cross (2023) that inflation expectations are sensitive to oil price shocks. For the euro area, Peersman and Van Robays (2009) show that inflationary responses via second-round effects are sizable. These results also align well with the findings of Bańbura, Bobeica and Martínez Hernández (2023), which point to a larger contribution of gas supply shocks compared to other energy price shocks.

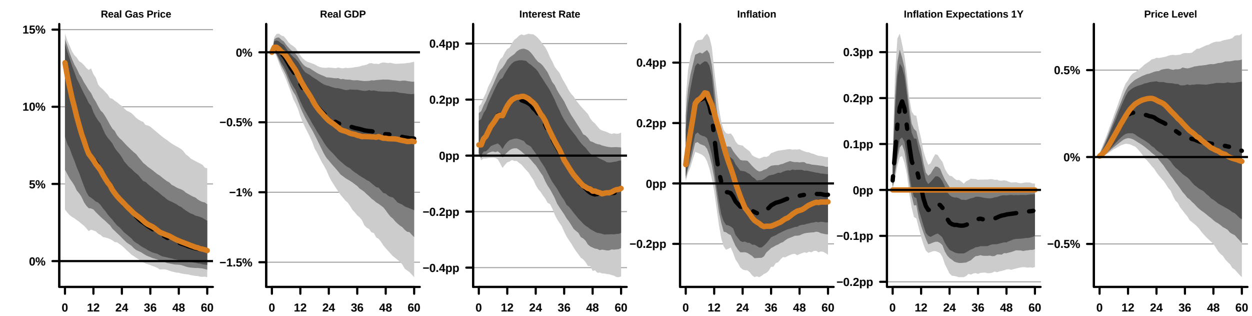
5.3 Real Gas Price Shocks in the US

In this extension, we re-do the analysis for the US. We exchange the real natural gas price to the price benchmark of the US (Henry Hub) and use US-specific macroeconomic quantities. The specification of the VAR, identification strategy, and sample is the same as before. We refer to Appendix A for the exact data details.

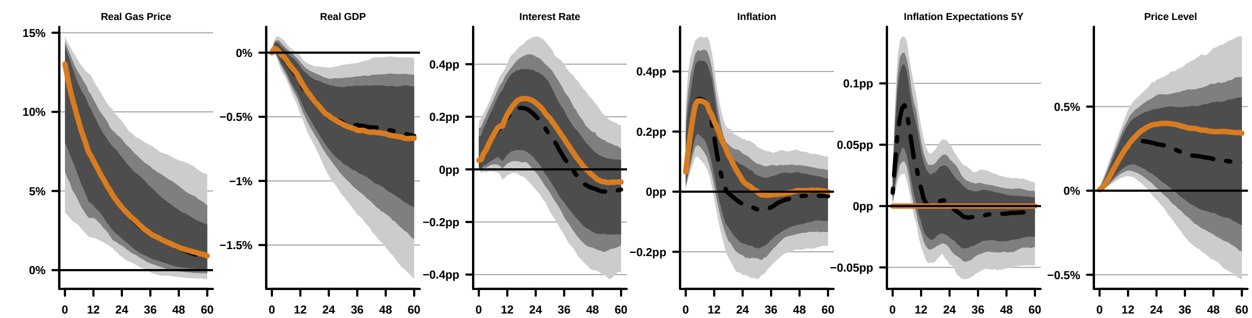
Results are presented in Figure 8. The real natural gas price shock elicits a jump in the price of about 13%. Inflation is elevated for about one year, while inflation expectations increase only temporarily with a quick mean reversion. Real GDP shows again a sluggish decline and interest rates show a contractionary monetary policy stance. The counterfactual exercise, however, is remarkably different from the euro area results. Shutting down second-round effects via inflation expectations does not alter the unconditional responses substantially. This holds for both short- and long-term expectations. Inflation does not show a strong difference from its unconditional response and thus we do not see a strong impact on the price level. This corroborates the findings of Wong (2015), who finds only mild evidence (in a sample dating back to the early 80s) that inflation expectations pass-through to inflation in response to a real oil price shock.

These results help to understand the differences between the euro area and the US. We have discussed the anchoring of inflation expectations, information rigidities in expectation formation process, and demand-side forces due to the recovery from the Covid-19 pandemic. Coibion and Gorodnichenko (2015a) show that

Figure 8: Impulse Response Functions to a Real Gas Price Shock for the US.



(a) Real Gas Price Shock for the US with 1Y Ahead Inflation Expectations



(b) Real Gas Price Shock for the US with 5Y Ahead Inflation Expectations

Notes: The extended model for the US features five variables, where the shock is identified with sign and zero restrictions and standardized to a one standard deviation increase in the real price of natural gas. The upper panel (a) uses 1Y inflation expectations while the lower panel (b) 5Y expectations. Real GDP is the cumulative response of real GDP growth. The price level is computed afterwards as cumulative sum of the inflation response. Black dashed lines denote the posterior median responses while gray shaded areas depict the 68/80/90 percent confidence intervals. The orange solid line denotes the counterfactual response. The vertical axis denotes the effect sizes of the real gas price, real GDP and the price level in percent, while the interest rate, inflation and inflation expectations are scaled to annualized percentage points. The horizontal axis denotes the impulse response horizon in months.

information rigidities are relatively similar in a set of countries, including the US and the euro area. Given the findings for the US, we discard this explanation. Similarly we discard the fiscal story since both economies experienced strong fiscal interventions as response to the pandemic, potentially outside of our model. Hence, the results point toward second-round effects. In line with Wong (2015), expectations are tightly anchored in the US. In contrast, Peersman and Van Robays (2009) point out the presence of sizable second-round effects of rising wages in the euro area for oil price shocks. They argue that wage-price spirals can lead to persistent inflationary effects. They also point out strong heterogeneities in the euro area due to different labor market dynamics. Clearly, these heterogeneities lead to monetary policy actions that does not fit all member countries. We corroborate these findings for real natural gas price shocks.

6. Concluding Remarks

This paper investigates the recent natural gas price surge and its implications for inflation and inflation expectations in the euro area. We are particularly interested in the sensitivity of inflation expectations to the real natural gas price shock and the empirical strength of the pass-through to inflation. To investigate this issue, we develop a semi-structural VAR model and use a combination of sign and zero restrictions to identify a real natural gas price shock. Then, we construct a counterfactual exercise in which the responses of inflation expectations are insensitive to the real natural gas price shock. We measure inflation expectations via market-based expectations. This allows us to inspect the pass-through along the term structure of inflation expectations. We also provide several extensions, in which we re-do the analysis with survey-based expectations, oil prices, and US data.

We find that both inflation and inflation expectations react positively to real natural gas price shocks. The counterfactual exercise reveals that second-round effects via inflation expectations are present and sizable, which indicates a limited role of the direct cost channel. Second-round effects are strongest for short-term expectations and vanish for long-term expectations. This points towards a relatively stable inflation expectations anchor. These findings are robust to a number of specification choices. The extensions reveal that the findings hold when using inflation expectations from the survey of professional forecasters. Inflation and inflation expectations are also sensitive to real oil price shocks but the pass-through is attenuated when compared to real natural gas price shocks. The findings are sensitive to the inclusion of the period starting in mid-2021 and cannot be, in general, replicated for the US.

We discuss possible reasons for these findings. Central banks have to ensure that expectations remain anchored in periods of sudden commodity price shocks, such as real natural gas price shocks. Monetary authorities take inflation expectations generally into account to achieve their objective of price stability. Our results show that second-round effects via inflation expectations are present in the euro area. The effect of a real natural gas price shock is sizable for short-term expectations, while long-term expectations remain well anchored. We leave for further research to investigate potential heterogeneities due to different labor market dynamics in the euro area. Hence, monetary policy actions taken by the ECB cannot fit all individual member countries. This suggests that the de-anchoring risks of expectations were not an immediate threat in the previous episode of energy market disruption. However, central bank policymakers are, and should be, closely monitoring the potential risk of de-anchoring. As we have argued, the EU is vulnerable to supply-side disruptions in energy markets, with additional threats to price stability via second-round effects. This logic broadly applies to other supply-side disruptions as well. Monitoring expectations is therefore an essential task for the central bank.

Declaration of Interest

The authors declare to have no conflict of interest.

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A. Data Appendix

All series were gathered from the sources listed below, including the FRED database (McCracken and Ng, 2016), the World Bank Commodity Price Data (*Pink Sheet*) (The World Bank, 2023), the statistical data warehouse of the European Central Bank, or Macrobond. If necessary, series are seasonally adjusted with the X-13ARIMA-SEATS model (Sax and Eddelbuettel, 2018). All series are approximately stationary.

Table A1: Variable Definitions.

Variable	Transformation	Details	Source
Euro area			
$\mathbf{rgas}_t$	$\ln\left(\frac{\mathtt{PGAS}_{it}}{\mathtt{E}_t^{US/EUR}\times\mathtt{HICP}_t}\right)$	real natural gas price	constructed
$\mathbf{roil}_t$	$\ln\left(\frac{\mathtt{POIL}_{it}}{\mathtt{E}_t^{US/EUR}\times\mathtt{HICP}_t}\right)$	real crude oil price	constructed
$\mathbf{rgdp}_t$	$100\times\ln\mathtt{RGDP}_t^m$	logarithm of real GDP	constructed
$\mathbf{sr}_t$	$\mathtt{SR}$	Shadow rate for euro area by Wu and Xia (2016)	website of Jing Cynthia Wu
π_t	$100\times\ln\left(\frac{\mathtt{HICP}_t}{\mathtt{HICP}_{t-12}}\right)$	year-on-year growth rate of harmonized index of consumer prices	constructed
π_t^e	$\mathtt{ILS}^{xY}$	inflation-linked swaps with x year ahead	Macrobond
π_t^e	$\mathtt{SPF}^{xY}$	survey of professional forecasters with $x=\{1,5\}$ years ahead	SPF ECB
$\mathbf{PGAS}_t$	$\mathtt{PGAS}_t$	price of natural gas (TTF) in $/mmBTU from Pink Sheet	World Bank
$\mathbf{POIL}_t$	$\mathtt{POIL}_t$	crude oil prices: Brent - Europe	FRED
$\mathbf{E}_t^{US/EUR}$	$\mathtt{E}_t^{US/EUR}$	US dollars to Euro spot exchange rate	FRED
$\mathbf{HICP}_t$	$\mathtt{HICP}_t$	harmonized index of consumer prices	FRED
$\mathbf{HICP}_t^{core}$	$\mathtt{HICP}_t^{core}$	harmonized index of consumer prices excluding food, energy, alcohol, and tobacco	FRED
$\mathbf{IP}_t$	$\mathtt{IP}_t$	industrial production index	SDW ECB
$\mathbf{RGDP}_t^m$	$\mathtt{RGDP}_t^m$	monthly real GDP: interpolated using Chow-Lin temporal disaggregation method with industrial produduction as input	SDW ECB, constructed
United States			
$\mathbf{rgas}_t$	$\ln\left(\frac{\mathtt{PGAS}_{it}}{\mathtt{CPI}_t}\right)$	real natural gas price	constructed
$\mathbf{roil}_t$	$\ln\left(\frac{\mathtt{POIL}_{it}}{\mathtt{CPI}_t}\right)$	real crude oil price	constructed
$\mathbf{rgdp}_t$	$100\times\ln\mathtt{RGDP}_t^m$	logarithm of real GDP	constructed
$\mathbf{sr}_t$	$\mathtt{SR}$	Shadow rate for the US by Wu and Xia (2016)	website of Jing Cynthia Wu
π_t	$100\times\ln\left(\frac{\mathtt{CPI}_t}{\mathtt{CPI}_{t-12}}\right)$	year-on-year growth rate of the consumer prices index	constructed
π_t^e	$\mathtt{ILS}^{xY}$	inflation-linked swaps with x year ahead	Macrobond
$\mathbf{PGAS}_t$	$\mathtt{PGAS}_t$	price of natural gas (Henry Hub) in $/mmBTU from Pink Sheet	World Bank
$\mathbf{POIL}_t$	$\mathtt{POIL}_t$	crude oil prices: West Texas Intermediate (WTI), US dollars per Barrel	FRED
$\mathbf{CPI}_t$	$\mathtt{CPI}_t$	consumer prices index for all urban consumers	FRED
$\mathbf{CPI}_t^{core}$	$\mathtt{CPI}_t^{core}$	consumer price index for all urban consumers excluding food and energy	FRED
$\mathbf{IP}_t$	$\mathtt{IP}_t$	industrial production index	FRED
$\mathbf{RGDP}_t^m$	$\mathtt{RGDP}_t^m$	monthly real GDP: interpolated using Chow-Lin temporal disaggregation method with industrial production as input	FRED, constructed

B. Econometric Details

In this section, we briefly describe the estimation strategy of the macroeconomic model. The estimation of the VAR is based on a Bayesian framework with the Normal-Gamma shrinkage prior, a variant of a global-local shrinkage prior (Griffin, Brown et al., 2010; Huber and Feldkircher, 2019). Hence, following Equation (3.1), the reduced-form VAR(p) model for the time series process $\boldsymbol{y}_t$ reads

$$\boldsymbol{y}_t = \boldsymbol{\mu} + \boldsymbol{\Phi}_1 \boldsymbol{y}_{t-1} + \ldots + \boldsymbol{\Phi}_p \boldsymbol{y}_{t-p} + \boldsymbol{\Gamma} \boldsymbol{d}_t + \boldsymbol{u}_t, \quad \boldsymbol{u}_t \sim \mathcal{N}_n(\boldsymbol{0}, \boldsymbol{\Sigma}), \tag{B.1}$$

where p is the lag order, $\boldsymbol{c}$ is an $n \times 1$ vector of constants, $\boldsymbol{\Phi}_1, \ldots, \boldsymbol{\Phi}_p$ are $n \times n$ coefficient matrices, and $\boldsymbol{u}_t$ denotes an $n \times 1$ vector of reduced-form Gaussian distributed innovations with covariance matrix $\boldsymbol{\Sigma}$, factorized as follows $\boldsymbol{\Sigma} = \boldsymbol{S}^{-1} \boldsymbol{\Lambda} \boldsymbol{S}^{-1'}$. Additionally, the model may feature exogenous variables in the data matrix $\boldsymbol{d}_t$ of size $n_d \times 1$ and the corresponding coefficient matrix $\boldsymbol{\Gamma}$ of size $n \times n_d$. This allows us to introduce dummy variables for the pandemic months (Cascaldi-Garcia, 2022). We collect all VAR coefficients in $\boldsymbol{\alpha} = (\boldsymbol{\mu}', \boldsymbol{\Phi}_1', \ldots, \boldsymbol{\Phi}_p', \boldsymbol{\Gamma}')'$. $\boldsymbol{\Lambda}$ is a diagonal matrix with generic jth element λ_j. These coefficients are gathered in $\boldsymbol{\lambda} = (\lambda_1, \ldots, \lambda_n)'$. $\boldsymbol{S}^{-1}$ is a lower-triangular matrix with ones on its main diagonal.

Estimation. For the estimation, we pursue the approach by Chan and Eisenstat (2018) and Chan (2022). For that, we re-write the system in its triangularized form:

$$\boldsymbol{S} \boldsymbol{y}_t = \tilde{\boldsymbol{x}}_t \tilde{\boldsymbol{\alpha}} + \tilde{\boldsymbol{\varepsilon}}_t, \qquad \tilde{\boldsymbol{\varepsilon}}_t \sim \mathcal{N}\left(\boldsymbol{0}, \boldsymbol{\Lambda}\right), \tag{B.2}$$

where $\tilde{\boldsymbol{x}}_t = (1, \boldsymbol{y}_{t-1}', \ldots, \boldsymbol{y}_{t-p}', \boldsymbol{d}_t)$. We can easily recover the reduced-form parameters by $\boldsymbol{\alpha} = \boldsymbol{S}^{-1} \tilde{\boldsymbol{\alpha}}$, the reduced-form covariance matrix $\boldsymbol{\Sigma} = \boldsymbol{S}^{-1} \boldsymbol{\Lambda} \boldsymbol{S}^{-1'}$, and reduced-form shocks by $\boldsymbol{u}_t = \boldsymbol{S}^{-1} \tilde{\boldsymbol{\varepsilon}}_t$. Note that $\tilde{\boldsymbol{\varepsilon}}$ are not equal to the structural shocks $\boldsymbol{\varepsilon}$ of Eq. (3.2). Consequently, we re-write the ith equation of the system as

$$y_{i,t} = \tilde{\boldsymbol{w}}_{i,t} \boldsymbol{s}_i + \tilde{\boldsymbol{x}}_t \tilde{\boldsymbol{\alpha}}_i + \varepsilon_{i,t}, \qquad \tilde{\varepsilon}_{i,t} \sim \mathcal{N}(0, \lambda_i^2), \tag{B.3}$$

where $\tilde{\boldsymbol{w}}_{i,t} = (-y_{1,t}, \ldots, -y_{i-1,t})$ and $\boldsymbol{s}_i$ are the elements first $i-1$ elements in the ith row of $\boldsymbol{S}$. Note that $y_{i,t}$ depends on the contemporaneous variables $y_{1,t}, \ldots, y_{i-1,t}$. We estimate the system in its triangular form and let $\boldsymbol{x}_{i,t} = (\tilde{\boldsymbol{w}}_{i,t}, \tilde{\boldsymbol{x}}_t)$, then it simplifies to

$$y_{i,t} = \boldsymbol{x}_{i,t} \boldsymbol{\theta}_i + \varepsilon_{i,t}, \qquad \tilde{\varepsilon}_{i,t} \sim \mathcal{N}(0, \lambda_i^2), \tag{B.4}$$

where $\boldsymbol{\theta}_i = (\boldsymbol{s}_i', \tilde{\alpha}_i')$ is of dimension $k_i = np+i+n_d$. This allows us to estimate the VAR equation-by-equation and to impose asymmetries in the amount of shrinkage per variable and equation. Important to note here is that we specify priors directly on the *triangularized* coefficients and not the *reduced-form* coefficients. This variant of VAR estimation has no order invariance issues as in Carriero, Clark and Marcellino (2019) and Carriero et al. (2022). Then we can back out the reduced-form coefficients.

Prior Specification. We have to elicit a prior distribution on $(\boldsymbol{\theta}, \boldsymbol{\lambda})$. We assume that the parameters are a priori independent across equations, such that $p(\boldsymbol{\theta}, \boldsymbol{\lambda}) = \prod_{i=1}^{n} p((\boldsymbol{\theta}_i, \lambda_i)$.

Specifically, for $i = 1, \ldots, n$, we assume:

$$\boldsymbol{\theta}_i \sim \mathcal{N}(\boldsymbol{m}_i, \boldsymbol{V}_i). \tag{B.5}$$

Following Huber and Feldkircher (2019), we consider a Normal-Gamma (NG) shrinkage prior setup for the VAR coefficients, which is given by

$$V_{ij} \mid \kappa_i^2, \theta_{ij} \sim \mathcal{N}(\underline{V}_{ij}, 2\kappa_i^{-2}\theta_{ij}), \qquad \theta_{ij} \sim G(\tau_\theta, \tau_\theta), \quad \kappa_i^2 \sim \mathcal{G}(d_\kappa, e_\kappa), \tag{B.6}$$

where V_{ij} denotes the j-th diagonal element of the matrix $\boldsymbol{V}_i$. τ_{ij} denotes the *local* shrinkage parameter that is coefficient specific and λ_i is a *global* shrinkage term that pulls all elements in $\boldsymbol{V}_i$ towards zero. This can be viewed as a common equation-specific scaling factor with the θ_{ij} allowing for coefficient-specific deviations in light of a large value of κ_i^2. On both the global and local parameters, we impose Gamma distributed priors with hyperparameters τ_θ, d_κ, and e_κ. τ_θ controls the tail behavior of the prior with small values placing more prior mass on zero and leading to heavier tails. The remaining two hyperparameters d_κ and e_κ control the amount of global shrinkage with small values (i.e. of order 0.01) leading to heavy shrinkage towards the origin.

Finally, for the volatilities, we specify Inverse-Gamma prior distributions, which reads for $i = 1, \ldots, n$:

$$\lambda_i \sim IG(c_0, d_0), \tag{B.7}$$

where $c_0 = 3$ and $d_0 = 0.03$.

C. Details on Structural Scenario Analysis Counterfactuals

Building on the work of Waggoner and Zha (1999), the structural scenario analysis framework of Antolin-Diaz, Petrella and Rubio-Ramirez (2021) provides a general framework on how to impose specific paths on observed variables in a VAR model as conditional forecasts with and without constraints on the set of offsetting – or *driving* – shocks. Breitenlechner, Georgiadis and Schumann (2022) adapt this to the case of impulse response analysis with structural scenario analysis (SSA). Again, iterate the VAR model in Equation (3.1) forward and re-write it as

$$\boldsymbol{y}_{T+1,T+h} = \boldsymbol{b}_{T+1,T+h} + \boldsymbol{M}'\boldsymbol{\varepsilon}_{T+1,T+h}, \tag{C.1}$$

where the $nh\times 1$ vector $\boldsymbol{y}_{T+1,T+h} = (\boldsymbol{y}'_{T+1}, \boldsymbol{y}'_{T+2}, \ldots, \boldsymbol{y}'_{T+h})'$ denotes future values of the endogenous variables, $\boldsymbol{b}_{T+1,T+h}$ an autoregressive component that is due to initial conditions as of period T, and the $nh \times 1$ vector $\boldsymbol{\varepsilon}_{T+1,T+h} = (\boldsymbol{\varepsilon}'_{T+1}, \boldsymbol{\varepsilon}'_{T+2}, \ldots, \boldsymbol{\varepsilon}'_{T+h})'$ future values of the structural shocks. The $nh \times nh$ matrix $\boldsymbol{M}$ reflects the impulse responses and is a function of the structural VAR parameters. The definition of $\boldsymbol{M}$ is as follows

$$\boldsymbol{M} = \begin{bmatrix} \boldsymbol{M}_0 & \boldsymbol{M}_1 & \ldots & \boldsymbol{M}_{h-1} \\ \boldsymbol{0} & \boldsymbol{M}_0 & \ldots & \boldsymbol{M}_{h-2} \\ \vdots & \vdots & \ddots & \vdots \\ \boldsymbol{0} & \boldsymbol{0} & \ldots & \boldsymbol{M}_0 \end{bmatrix}, \tag{C.2}$$

where $\boldsymbol{M}_0 = \boldsymbol{S}^{-1}$ and $\boldsymbol{M}_i = \sum_{j=1}^{i} \boldsymbol{M}_{i-j}\boldsymbol{B}_j$ with $\boldsymbol{B}_j = \boldsymbol{0}$ if $j > p$. From this representation, it is clear that the matrix $\boldsymbol{M}$ only depends on the structural parameters. Furthermore, note that $\boldsymbol{M}'\boldsymbol{M}$ only depends on the reduced-form parameters. Thus, one only needs the history of observables and the reduced-form parameters to characterize the distribution of the unconditional forecast.

Then, the unconditional forecast is distributed

$$\boldsymbol{y}_{T+1,T+h} = \mathcal{N}\left(\boldsymbol{b}_{T+1,T+h}, \boldsymbol{M}'\boldsymbol{M}\right). \tag{C.3}$$

In the framework of Antolin-Diaz, Petrella and Rubio-Ramirez (2021), structural scenarios involve

i) *Conditional-on-observables* forecasting, i.e., specifying paths for a subset of observables in $\boldsymbol{y}_{T+1,T+h}$ that depart from their unconditional forecast, and/or

ii) *Conditional-on-shocks* forecasting, i.e., specifying the subset of structural shocks $\boldsymbol{\varepsilon}_{T+1,T+h}$ that are allowed to deviate from their unconditional distribution to produce the specified path of the observables in (i).

In the following, we will discuss how to implement both options. Therefore, one should note that

$$\tilde{\boldsymbol{y}}_{T+1,T+h} \sim \mathcal{N}\left(\boldsymbol{\mu}_y, \boldsymbol{\Sigma}_y\right), \tag{C.4}$$

denotes the distribution of the future values of the *constrained* observables. The goal is to determine $\boldsymbol{\mu}_y$ and $\boldsymbol{\Sigma}_y$ such that the constraints in (i) and (ii) are satisfied simultaneously.

Under (i), *conditional-on-observables* forecasting can be implemented as follows. Let $\overline{\boldsymbol{C}}$ be a $k_o \times nh$ selection matrix, with k_o denoting the number of restrictions. Then, *conditional-on-observables* restrictions can be written as

$$\overline{\boldsymbol{C}}\tilde{\boldsymbol{y}}_{T+1,T+h} \sim \mathcal{N}\left(\overline{\boldsymbol{f}}_{T+1,T+h}, \overline{\boldsymbol{\Omega}}_f\right), \tag{C.5}$$

where the $k_o \times 1$ vector $\overline{\boldsymbol{f}}_{T+1,T+h}$ is the mean of the distribution of the observables constrained under the conditional forecast, and the $k_o \times k_o$ matrix $\overline{\boldsymbol{\Omega}}_f$ is the associated variance-covariance matrix.

Under (ii), *conditional-on-shocks* forecasting can be implemented as follows. Let Ξ be a $k_s \times nh$ selection matrix, with k_s denoting the number of restrictions. Then, *conditional-on-shocks* restrictions can be written as

$$\Xi\tilde{\boldsymbol{\varepsilon}}_{T+1,T+h} \sim \mathcal{N}\left(\boldsymbol{g}_{T+1,T+h}, \boldsymbol{\Omega}_g\right), \tag{C.6}$$

where the $k_s \times 1$ vector $\boldsymbol{g}_{T+1,T+h}$ is the mean of the distribution of the shocks constrained under the conditional forecast and the $k_s \times k_s$ matrix $\boldsymbol{\Omega}_g$ is the associated variance-covariance matrix. Under invertibility, the shocks can always be expressed as a function of observed variables and allow us to re-write the restrictions:

$$\begin{aligned}\Xi\boldsymbol{M}'^{-1}\tilde{\boldsymbol{y}}_{T+1,T+h} &= \Xi\boldsymbol{M}'^{-1}\boldsymbol{b}_{T+1,T+h} + \Xi\tilde{\boldsymbol{\varepsilon}}_{T+1,T+h} \\ \underline{\boldsymbol{C}}\tilde{\boldsymbol{y}}_{T+1,T+h} &= \underline{\boldsymbol{C}}\boldsymbol{b}_{T+1,T+h} + \Xi\tilde{\boldsymbol{\varepsilon}}_{T+1,T+h},\end{aligned} \tag{C.7}$$

and thus

$$\underline{\boldsymbol{C}}\tilde{\boldsymbol{y}}_{T+1,T+h} = \underline{\boldsymbol{C}}\boldsymbol{b}_{T+1,T+h} + \Xi\tilde{\boldsymbol{\varepsilon}}_{T+1,T+h} \sim \mathcal{N}\left(\underline{\boldsymbol{f}}_{T+1,T+h}, \underline{\boldsymbol{\Omega}}_f\right), \tag{C.8}$$

where $\underline{\boldsymbol{\Omega}}_f = \boldsymbol{\Omega}_g$.

Now we can combine the k_o restrictions on the observables under *conditional-on-observables* forecasting and the k_s restrictions on the structural shocks under *conditional-on-shocks* forecasting. This amounts to $k = k_o + k_s$ total restrictions. We define the $k \times nh$ matrices $\boldsymbol{C} = [\overline{\boldsymbol{C}}', \underline{\boldsymbol{C}}']'$ and $\boldsymbol{D} = [\boldsymbol{M}\overline{\boldsymbol{C}}', \boldsymbol{\Xi}']'$, which allows us to write

$$\boldsymbol{C}\tilde{\boldsymbol{y}}_{T+1,T+h} = \boldsymbol{C}\boldsymbol{b}_{T+1,T+h} + \boldsymbol{D}\tilde{\boldsymbol{\varepsilon}}_{T+1,T+h} \sim \mathcal{N}\left(\boldsymbol{f}_{T+1,T+h}, \boldsymbol{\Omega}_f\right), \tag{C.9}$$

where the $k \times 1$ vector $\boldsymbol{f}_{T+1,T+h} = [\overline{\boldsymbol{f}}'_{T+1,T+h}, \underline{\boldsymbol{f}}'_{T+1,T+h}]'$ stacks the means of the distribution and the $k \times k$ matrix $\boldsymbol{\Omega}_f = \mathtt{diag}(\overline{\boldsymbol{\Omega}}_f, \underline{\boldsymbol{\Omega}}_f)$ denotes the associated variance-covariance matrix.

Following the framework in Antolin-Diaz, Petrella and Rubio-Ramirez (2021) and given the restrictions specified above, we can derive solutions for $\boldsymbol{\mu}_y$ and $\boldsymbol{\Sigma}_y$. Define the restricted future shocks

$$\tilde{\boldsymbol{\varepsilon}}_{T+1,T+h} \sim \mathcal{N}(\boldsymbol{\mu}_\varepsilon, \boldsymbol{\Sigma}_\varepsilon), \tag{C.10}$$

where $\boldsymbol{\Sigma}_\varepsilon = \boldsymbol{I}_n h + \boldsymbol{\Psi}_\varepsilon$, such that $\boldsymbol{\mu}_\varepsilon$ and $\boldsymbol{\Psi}_\varepsilon$ denote the deviation of the mean and covariance matrix from their unconditional counterparts. Using Eq. (C.9), we match the first and second moments to get

$$\boldsymbol{f}_{T+1,T+h} = \boldsymbol{C}\boldsymbol{b}_{T+1,T+h} + \boldsymbol{D}\boldsymbol{\mu}_\varepsilon, \tag{C.11}$$

$$\boldsymbol{\Omega}_f = \boldsymbol{D}(\boldsymbol{I}_n h + \boldsymbol{\Psi}_\varepsilon)\boldsymbol{D}'. \tag{C.12}$$

Depending on k, the number of restrictions, and nh, the length of $\tilde{\boldsymbol{y}}_{T+1,T+h}$, the systems of Eq. (C.11) and Eq. (C.12) may have multiple solutions ($k < nh$), one solution ($k = nh$), or no solution ($k > nh$). Since $k < nh$ is the most interesting case, the solution is given by

$$\boldsymbol{\mu}_\varepsilon = \boldsymbol{D}^*\left(\boldsymbol{f}_{T+1,T+h} - \boldsymbol{C}\boldsymbol{b}_{T+1,T+h}\right), \tag{C.13}$$

$$\boldsymbol{\Psi}_\varepsilon = \boldsymbol{D}^*\boldsymbol{\Omega}_f\boldsymbol{D}^{*'} - \boldsymbol{D}^*\boldsymbol{D}\boldsymbol{D}'\boldsymbol{D}^{*'}, \tag{C.14}$$

where $\boldsymbol{D}^*$ is the Moore-Penrose inverse of $\boldsymbol{D}$. Eq. (C.13) shows that the path of the implied structural shocks under the conditional forecast depends on its deviation from the unconditional forecast. Furthermore, Eq. (C.14) shows that the variance of the implied future structural shocks depends on the uncertainty the researcher attaches to the conditional forecast. If the uncertainty is zero ($\boldsymbol{\Omega}_f = \boldsymbol{0}$), then $\boldsymbol{\Sigma}_\varepsilon = \boldsymbol{0}$. This means that a unique path for $\boldsymbol{\mu}_\varepsilon$ can be found.

Combining Eq. (C.3), Eq. (C.13), and Eq. (C.14), we get

$$\boldsymbol{\mu}_y = \boldsymbol{b}_{T+1,T+h} + \boldsymbol{M}'\boldsymbol{D}^*\left(\boldsymbol{f}_{T+1,T+h} - \boldsymbol{C}\boldsymbol{b}_{T+1,T+h}\right), \tag{C.15}$$

$$\boldsymbol{\Sigma}_y = \boldsymbol{M}'\boldsymbol{M} - \boldsymbol{M}'\boldsymbol{D}^* \left(\boldsymbol{\Omega}_f - \boldsymbol{D}\boldsymbol{D}'\right) \boldsymbol{D}^{*'} \boldsymbol{M}. \tag{C.16}$$

As before, if $\boldsymbol{\Omega}_f = \mathbf{0}$, then $\boldsymbol{\Sigma}_y = \mathbf{0}$ and thus there is no uncertainty about the path of the observables under the imposed restrictions.

C1. Restrictions in the VAR

In our VAR, we have $\boldsymbol{y}_t = [rgas_t, \Delta rgdp_t, sr_t, \pi_t, \pi_t^e]$ and want to constrain the effect of a real gas price shock on inflation expectations π_t^e to be zero. Denote with $\boldsymbol{e}_i$ an $n \times 1$ vector of zeros with unity at the i-th position.

Under (i), *conditional-on-observable* forecasting, we impose

$$\overline{\boldsymbol{C}} = \boldsymbol{I}_h \otimes \boldsymbol{e}_5', \tag{C.17}$$

$$\overline{\boldsymbol{f}}_{T+1,T+h} = \mathbf{0}_{h\times 1}, \tag{C.18}$$

$$\overline{\boldsymbol{\Omega}}_f = \mathbf{0}_{h\times h}. \tag{C.19}$$

These equations impose that the conditional forecast that underlies the impulse response of inflation expectations (which is ordered fifth in the VAR) is constrained to be zero over all horizons $T+1, \ldots, T+h$. Furthermore, we do not allow for any uncertainty.

Under (ii), *conditional-on-shocks* forecasting, we impose

$$\boldsymbol{\Xi} = \begin{bmatrix} \boldsymbol{e}_1' & \mathbf{0}_{1\times n(h-1)} \\ (\mathbf{0}_{n-2\times 1}, \boldsymbol{I}_{n-2}) & \mathbf{0}_{n-2\times n(h-1)} \\ \mathbf{0}_{(h-1)(n-1)\times n} & \boldsymbol{I}_{h-1} \otimes (\boldsymbol{I}_{n-2}, \mathbf{0}_{n-2\times 1}) \end{bmatrix}_{h(n-1)\times nh} \tag{C.20}$$

$$\underline{\boldsymbol{f}}_{T+1,T+h} = \boldsymbol{g}_{T+1,T+h} = [1, \mathbf{0}_{1\times n-2}, \mathbf{0}_{1\times (n-1)(h-1)}]', \tag{C.21}$$

$$\underline{\boldsymbol{\Omega}}_f = \boldsymbol{\Omega}_g = \mathbf{0}_{h(n-1)\times h(n-1)} \tag{C.22}$$

The first row in Eq. (C.20) selects the real gas price shock ordered first in $\boldsymbol{\varepsilon}_t$ and the first row in Eq. (C.21) constrains it to be unity in the impact period $T+1$. In the second row in Eq. (C.20) we select the structural shock to industrial production, short-term interest rate, and inflation (ordered from the second to second-last position in the VAR) and the second entry of Eq. (C.21) constrains these structural shocks to be zero in period $T+1$. Hence, in $T+1$ the only structural shock which is allowed to vary is the one of inflation expectations. Similarly, the third row selects the first $n-1$ structural shocks over the remaining impulse response horizon

$T+2, T+3, \ldots, T+h$ and constrains them to zero in Eq. (C.21). Hence, in $T+2, T+3, \ldots, T+h$ the only structural shock which is allowed to vary is again the one of inflation expectations. Lastly, Eq. (C.22) specifies that we allow for no uncertainty.

We also consider the case, in which we use all all shocks as offsetting force to construct the scenario analysis. In this case, we have slightly different equations

$$\boldsymbol{\Xi} = \boldsymbol{I}_h \otimes \boldsymbol{e}_1' \tag{C.20a}$$

$$\underline{\boldsymbol{f}}_{T+1,T+h} = \boldsymbol{g}_{T+1,T+h} = [1, \boldsymbol{0}_{1\times n-2}, \boldsymbol{0}_{1\times(n-1)(h-1)}]', \tag{C.21a}$$

$$\underline{\boldsymbol{\Omega}}_f = \boldsymbol{\Omega}_g = \boldsymbol{0}_{h(n-1)\times h(n-1)}, \tag{C.22a}$$

where only the first equation differs markedly from before. The first shocks is not allowed to deviate from its unconditional distribution, as before. However, all the other shocks are allowed to move freely to create an offsetting force.

It is also interesting to consider the stacked matrices $\boldsymbol{C}$ and $\boldsymbol{D}$ which look as follows

$$\boldsymbol{C} = \begin{pmatrix} \overline{\boldsymbol{C}}_{h\times nh} \\ \underline{\boldsymbol{C}}_{h(n-1)\times nh} \end{pmatrix}_{hn\times nh}, \qquad \boldsymbol{D} = \begin{pmatrix} \overline{\boldsymbol{C}}_{h\times nh}\boldsymbol{M}'_{nh\times nh} \\ \boldsymbol{\Xi}_{h(n-1)\times nh} \end{pmatrix}_{hn\times nh}, \tag{C.23}$$

where $\underline{\boldsymbol{C}} = \boldsymbol{\Xi}\boldsymbol{M}'^{-1}$.

C2. How plausible is the counterfactual?

Generally, structural scenario analysis counterfactuals based on SVARs are not prone to the Lucas critique (Lucas, 1976). However, if the implied shocks are so *unusual* the analysis might become subject to the Lucas critique anyway. Hence, measures of the plausibility of the created counterfactual scenario are a remedy. We use two measures: the q-divergence proposed in Antolin-Diaz, Petrella and Rubio-Ramirez (2021) and adapted to the case of impulse response functions by Breitenlechner, Georgiadis and Schumann (2022) and the modesty statistic proposed by Leeper and Zha (2003). These measures intend to measure how much the structural scenario deviates from its unconditional counterpart. When this deviation becomes too large, the scenario might be implausible.

Antolin-Diaz, Petrella and Rubio-Ramirez (2021) propose to use the Kullback-Leibler (KL) divergence as a measure of how plausible a scenario is. Denote with $\mathcal{D}(\mathcal{N}_{SS}||\mathcal{N}_{UF})$ the KL divergence between the distributions of the structural scenario analysis $\mathcal{N}_{SS}$ and the unconditional distribution $\mathcal{N}_{UF}$. While it is straightforward to compute $\mathcal{D}(\mathcal{N}_{SS}||\mathcal{N}_{UF})$, it is difficult to grasp whether any value for the KL divergence is large or small. In other words, the KL divergence can be easily used to rank scenarios, but it is hard to understand how far away they are from the unconditional forecast. Therefore, Antolin-Diaz, Petrella and Rubio-Ramirez (2021) propose to compare the KL divergence with the divergence between two binomial distributions, one with probability q and the other with probability $p = 0.5$. The idea is to compare the implied counterfactual distribution with their unconditional distribution, which translates into a comparison of the binomial distributions of a fair and a biased coin. If the probability q is near p, then this suggests that the distribution of the offsetting shocks is not at all far from the unconditional distribution. Antolin-Diaz, Petrella and Rubio-Ramirez (2021) suggest calibrating the KL divergence from $\mathcal{N}_{UF}$ to $\mathcal{N}_{SS}$ to a parameter q that would solve the following equation $\mathcal{D}\left(\mathcal{B}(nh, 0.5)||\mathcal{B}(nh, q)\right) = \mathcal{D}\left(\mathcal{N}_{SS}||\mathcal{N}_{UF}\right)$. The solution to the equation is

$$q = 0.5 * \left(1 + \sqrt{1 - \exp\left(-\frac{2z}{nh}\right)}\right) \quad \text{with} \quad z = \mathcal{D}\left(\mathcal{N}_{SS}||\mathcal{N}_{UF}\right). \tag{C.24}$$

As Breitenlechner, Georgiadis and Schumann (2022) point out, in the context of impulse responses the KL divergence has to be slightly adjusted because Antolin-Diaz, Petrella and Rubio-Ramirez (2021) propose their measure in the context of conditional forecasts relative to an unconditional forecast. As before, the unconditional scenario is the case with only a single shock of unity size, which occurs in $T+1$ with certainty. More formally, $\boldsymbol{\varepsilon}_{T+1,T+h} = (\boldsymbol{e}_1', \boldsymbol{0}_{n(h-1)\times 1})'$ denotes the *unconditional* impulse response of a natural gas price shock. $\boldsymbol{e}_i$ denotes the unit vector with unity on the i-th position. For the structural scenario analysis counterfactual, we impose the restrictions specified above (i.e., inflation expectations do not react to a natural gas price shock). Hence, we set

$$\texttt{UF:} \qquad \boldsymbol{\mu}_{UF} = \boldsymbol{M}'(\boldsymbol{e}_1', \boldsymbol{0}_{n(h-1)\times 1})' \tag{C.25}$$

$$\texttt{SS:} \qquad \boldsymbol{\mu}_{SS} = \boldsymbol{\mu}_y, \tag{C.26}$$

where $\boldsymbol{\mu}_y$ is given by Equation (C.15). Since we impose this with certainty, $\boldsymbol{\Psi} = \boldsymbol{0}$ such that the shocks have their unconditional variance. Hence, $\boldsymbol{\Sigma}_{UF} = \boldsymbol{\Sigma}_{SS} = \boldsymbol{\Sigma}_\varepsilon = \boldsymbol{I}$. The KL divergence between the distribution of the shocks under the unconditional and conditional scenario is then given by

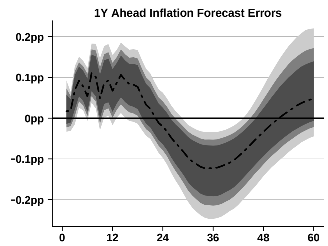
$$\mathcal{D}(\mathcal{N}_{SS}||\mathcal{N}_{UF}) = \frac{1}{2}\left(\mathrm{tr}\left(\boldsymbol{\Sigma}_{SS}^{-1}\boldsymbol{\Sigma}_{UF}\right) + (\boldsymbol{\mu}_{SS} - \boldsymbol{\mu}_{UF})'\boldsymbol{\Sigma}_{SS}^{-1}(\boldsymbol{\mu}_{SS} - \boldsymbol{\mu}_{UF}) - nh + \ln\left(\frac{\det \boldsymbol{\Sigma}_{SS}}{\det \boldsymbol{\Sigma}_{UF}}\right)\right), \tag{C.27}$$

where $\boldsymbol{\mu}_{\varepsilon}$ and $\boldsymbol{\Sigma}_{\varepsilon}$ are given by Equation (C.13) and Equation (C.14). Furthermore, we discard any SSA counterfactuals when the offsetting shocks are particularly unlikely. We set this to be above $q > 0.9$.

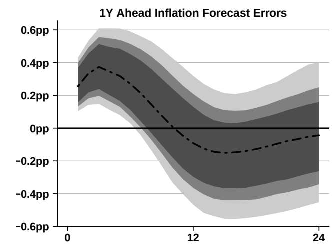
The second plausibility measure is the one of *modest intervention* or *modesty statistic* used in Leeper and Zha (2003). The measure reports how unusual the path for policy shocks is relative to the typical size of these shocks, which are needed to impose the counterfactual restriction. For instance, if the counterfactual implies a sequence of shocks close to their unconditional mean, the policy intervention is considered *modest*, in the sense that the shocks are unlikely to induce agents to revise their beliefs about policy rules and the structure of the economy. Instead, if the counterfactual involves an unlikely sequence of shocks, the analysis is likely to be prone to the critique by Lucas (1976). The offsetting shocks are considered to be modest if the statistic is smaller than two in absolute value.

D. Additional Results

Figure D1: Implied Impulse Response Functions to a Real Gas Price Shock.



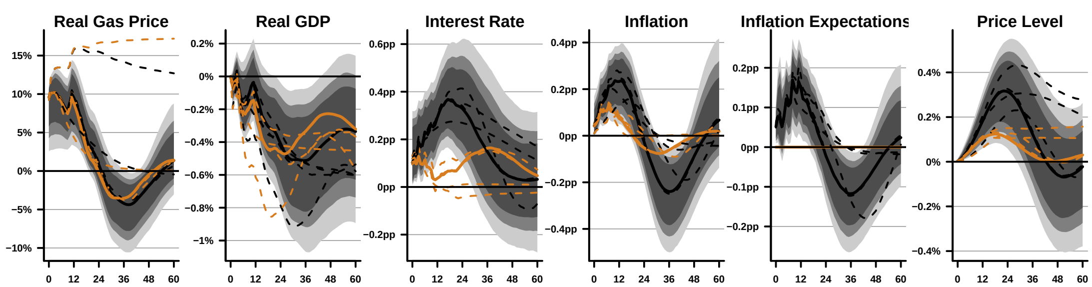
(a) ILS: Implied Forecast Error.



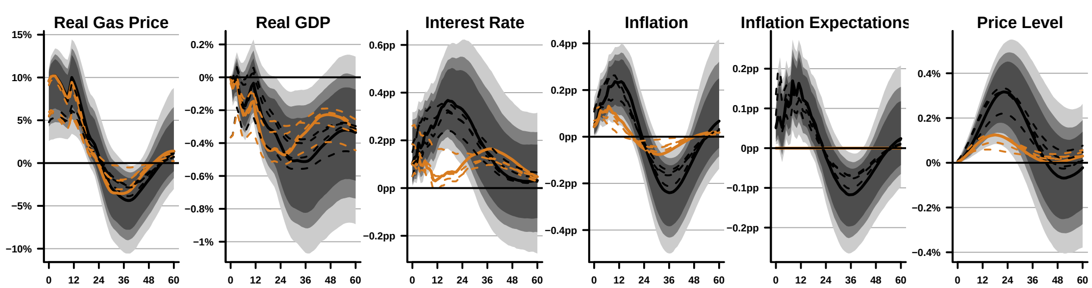
(b) SPF: Implied Forecast Error.

Notes: Implied impulse response function of forecast errors (constructed as the difference between realized inflation and the previous year's 1-year average expected inflation). The underlying model features five variables and is identified with sign and zero restrictions. Black dashed lines denote median responses, while gray shaded areas denote the 68/80/90 percent confidence intervals. The horizontal axis denotes the impulse response horizon in months (a) and quarters (b).

Figure D2: Robustness to the Baseline Model.



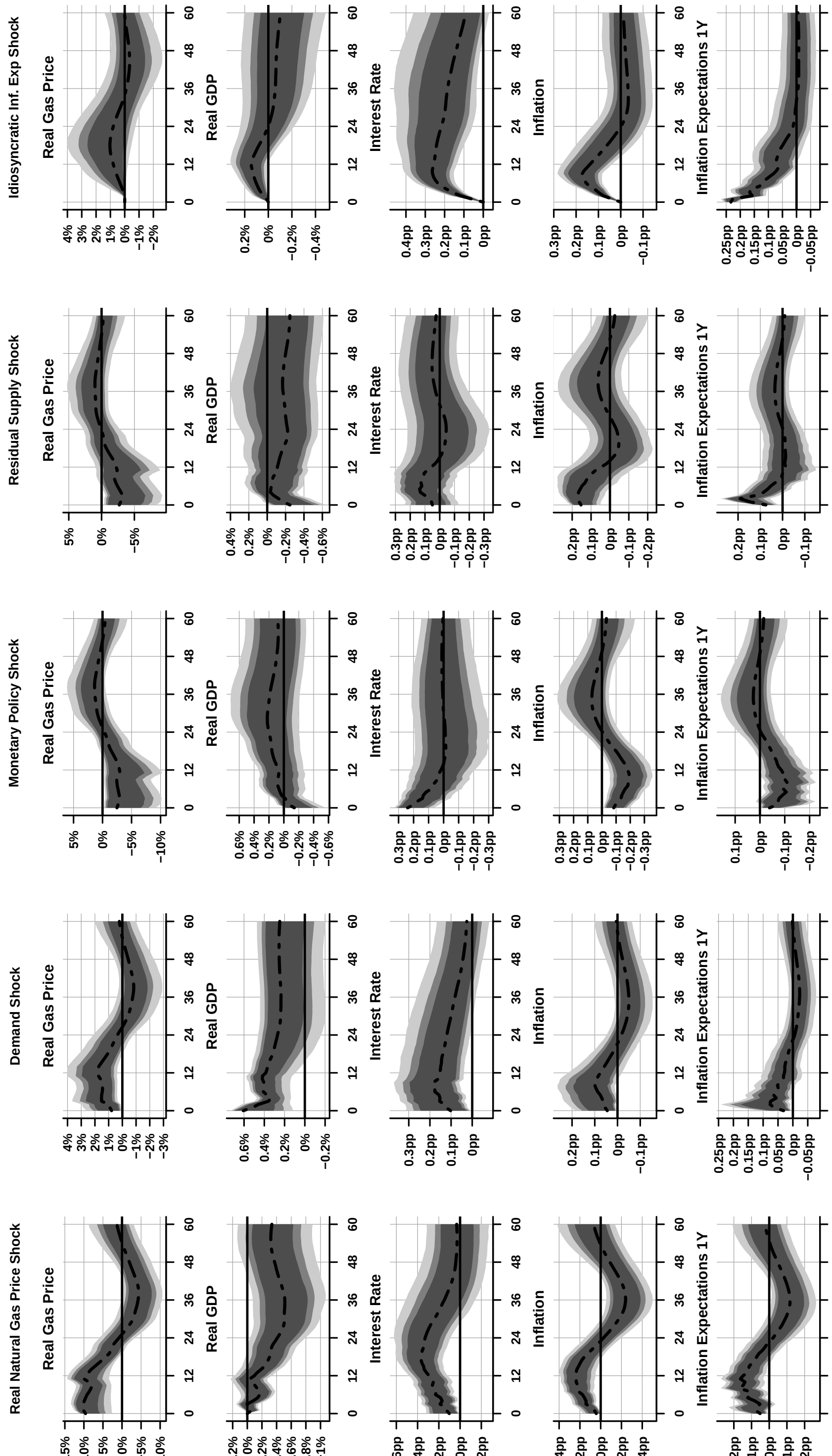
(a) Robustness: Specification Choices.



(b) Robustness: Identification Choices.

Notes: Robustness to the baseline model. Real GDP is the cumulative response of real GDP growth. The price level is computed afterwards as cumulative sum of the inflation response. Black solid line denotes median response, while gray shaded areas denote the 68/80/90 percent confidence intervals. The orange solid line denotes the counterfactual response. Dashed black and orange lines denote different specification choices. The vertical axis denotes the effect sizes of the real gas price, real GDP and the price level in percent, while the interest rate, inflation and inflation expectations are scaled to annualized percentage points. The horizontal axis denotes the impulse response horizon in months.

Figure D3: Full Set of Responses with 1Y Ahead Inflation Expectations.



Notes: Full set of impulse responses of the baseline model identified with sign and zero restrictions. Real GDP is the cumulative response of real GDP growth. Black dashed lines denote median responses, while gray shaded areas denote the 68/80/90 percent confidence intervals. The vertical axis denotes the effect sizes of the real gas price, real GDP and the price level in percent, while the interest rate, inflation and inflation expectations are scaled to annualized percentage points. The horizontal axis denotes the impulse response horizon in months.