Central Banking in Times of High Geopolitical Risk*

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Abstract

This paper investigates how the response of the U.S. economy to monetary policy shocks depends on the level of geopolitical risk. Using a nonlinear Proxy-SVAR model and the geopolitical risk index from Caldara and Iacoviello (2022), I find that the effects of monetary policy are significantly weakened in periods of high geopolitical risk. A tightening shock produces short-run expansionary effects and significantly smaller effects than in normal times. These findings can be rationalized through the monetary-fiscal interaction: defense spending increases following a monetary tightening shock only when geopolitical risk is high. A Lucas critique-robust policy counterfactual reveals that this monetary-fiscal interaction alone accounts for most of the documented state-dependent effects. Further policy experiments show that both the nature and the level of geopolitical risk are crucial in predicting the outcomes of monetary tightening and easing.

JEL codes: C32, C54, D81, E52, E62.

Keywords: monetary policy, geopolitical risk, nonlinear proxy-SVAR, state-dependence, monetary-fiscal interaction.

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1 Introduction

"Geopolitical risks, like the wars in Ukraine and the Middle East, are the greatest risk to stability. [...] It could be the price of oil. It could just the be the spreading conflict and the blow that would strike to public confidence. But we don't see that yet. It's a risk. It's a real risk and one we're aware of." – Jerome Powell (February 4, 2024)

Global geopolitical tensions have soared since Russia's invasion of Ukraine in early 2022. More recently, geopolitical risks have spread beyond Eastern Europe, with tensions rising in the Middle East. Geopolitical episodes, such as these recent events, pose significant risks to the world economic outlook and require vigilant monitoring by central banks and international organizations. The Federal Reserve Board, in its April 2024 Financial Stability Report, highlights that an increase in geopolitical tensions could reduce economic activity and boost inflation. Should these risks materialize, central banks must be prepared to adopt a contractionary stance to contain inflationary pressures.

In this paper, I examine whether the effects of monetary policy shocks depend on the level of geopolitical risk (hereafter GPR). Using U.S. data from mid-1979, I provide novel evidence that the effectiveness of monetary policy shocks is substantially weakened in periods of high GPR: a tightening shock leads to expansionary short-run real effects, followed by a statistically insignificant decline in output. The price response is similar, but not statistically significant. Inflation expectations fall with a larger lag in periods of high GPR than in normal times. The excess bond premium shows a qualitative smaller response compared to normal times, while a broader credit spread (Baa-Aaa corporate bond spread) shows a significant delayed response during periods of high GPR. In contrast, a monetary policy shock in normal times produces conventional responses. I obtain these results estimating the effect of an interest rate hike shock in normal times and during periods of high geopolitical tensions using the GPR index constructed by Caldara and Iacoviello (2022) and the nonlinear Proxy-SVAR approach developed by Debortoli et al. (2023) (hereafter DFGS). The nonlinearity in the model is introduced by interacting the monetary shock with the GPR index. The estimation sample runs from 1979m7 to 2019m12, and I identify the monetary policy shock using the instrument of Miranda-Agrippino and Ricco (2021) extended by Degasperi and Ricco (2021). I perform extensive checks on the shock identification, sample period, state variable and estimation framework. I find no significant changes in the estimates.

I rationalize these findings through the monetary-fiscal interaction. In periods of high GPR, unlike in normal times, real defense spending increases immediately following a contractionary shock. This is likely due to the risk of large military buildups following elevated GPR (see Caldara and Iacoviello, 2022).[1] These results suggest that an unfavorable policy mix could potentially explain the ineffectiveness of a monetary tightening shock during periods of high geopolitical tensions: the effect of a monetary shock is less effective in periods of high GPR if there is an increase in the systematic response of federal defense spending. To test how much this mechanism can account for the observed GPR-dependent monetary effects, I estimate the same responses to the monetary shock under an alternative counterfactual rule for fiscal policy. Specifically, I isolate the contribution of defense spending in affecting the GPR-dependent effects of monetary policy shocks. I implement this counterfactual using the empirical methodology recently proposed by McKay and Wolf (2023). The advantage of their method is that it is robust to the Lucas critique. This is because it involves subjecting the economy to informationally distinct policy shocks at time 0, with no ex-post surprises. The results of this counterfactual analysis suggest that once defense spending in the high GPR regime is set to respond similarly to normal times, the impulse responses during periods of high GPR resemble those in normal times: the expansionary short-run real effects are now absent, and output falls. Prices do not rise after the shock, but exhibit a lagged decline that closely follows the response of prices in normal times. The same is true for the excess bond premium. Overall, I conclude that the systematic fiscal (defense spending) policy response can account for most of the difference in the GPR-dependent effects of monetary policy.

So far, I have assumed that the GPR index of Caldara and Iacoviello (2022) reflects a homogeneous set of events in terms of their macroeconomic implications. However, this index incorporates two distinct macroeconomic developments: news of future supply-side disruptions in the energy market and lower aggregate demand pressures. This is an important distinction, particularly for monetary policymakers. While the first type of geopolitical event is associated with inflationary pressures that prompt the central bank to tighten interest rates, the second type of event is associated with deflationary

[1] To include real defense spending, I estimate a quarterly VAR and, alternatively, I maintain the monthly VAR and apply a time disaggregation method to obtain a monthly index of real defense spending. Both models convey the same message.

pressures that prompt the central bank to loosen interest rates. Leveraging on the subcomponents of the GPR index, I construct a new index that can better isolate geopolitical episodes linked with news of future supply-side disruptions in the energy market. Specifically, I remove the *Terror Acts* subcomponent, which peaks on 9/11, reflecting geopolitical events associated with lower aggregate demand pressures.[2] I show that when using this new index as state variable, which I call GPR *Threats and Wars*, I find larger statistically significant differences in output and the excess bond premium between periods of high GPR and normal times. These results suggest that the refined index is a more appropriate state-dependent variable in this empirical application.

Finally, I exploit the historical variability of the GPR index to construct policy experiments around the average effect by arbitrarily varying the level of GPR in the indirect effect. Using this feature of the data, I show that the effectiveness of monetary policy varies with the level of GPR. This is an important contribution, especially for monetary policymakers, as it provides a prescription for their actions under a given *level* and *nature*, i.e., its underlying macroeconomic implications, of GPR. I consider both monetary tightening and easing shocks with shock-specific state variables: I select the GPR *Threats and Wars* index for a tightening shock and the GPR *Terror Acts* index for an easing shock. Specifically, the *Terror Acts* index reflects geopolitical episodes associated with lower aggregate demand pressures and thus is a potentially more appropriate state variable when simulating an expansionary monetary shock in periods of GPR. The results suggest that a monetary tightening shock becomes less effective as geopolitical tensions rise, with clear output and price puzzles and a negative response of the excess bond premium when the index is at the 99th percentile.[3] These effects are accompanied by rising defense spending for higher GPR levels. A monetary easing shock, which is followed by a similar increase in defense spending, becomes more effective during periods of high GPR. These results confirm that the monetary-fiscal interaction mechanism is relevant for both types of shocks and nature of GPR.

[2]This is a clear example of an event that increases the real-option value of waiting for information (e.g., Bernanke, 1983; Bloom, 2009)

[3]While the 99th percentile of the GPR *Threats and Wars* index is 8.3, events such as the Kuwait invasion, the Gulf War and the Iraq War peak at 8.5, 11.4, and 10.3, respectively.

Related literature. This paper is closely related to two strands of literature. The first one includes the large literature studying the effects of monetary policy shocks.[4] While most of the studies have focused on linear models, recent papers have examined the nonlinear effects of monetary policy shocks. Among them, Tenreyro and Thwaites (2016) use nonlinear local projections and find that the effects of monetary policy in the U.S. are less powerful in recessions and that tightening shocks have stronger effects than easing shocks. The latter result is confirmed by Barnichon and Matthes (2018) using a functional approximation of impulse responses, which find that a monetary tightening shock in the U.S. raises unemployment, while an expansionary shock has little effect on unemployment. Using nonlinear local projections and data from 18 advanced economies, Alpanda et al. (2021) confirm that monetary policy is less effective in recessions and find similar results during periods of high interest rates. They also find that monetary policy shocks have larger real effects when household debt is high, with its effect being driven primarily by countries with adjustable-rate mortgages. More recently, Debortoli et al. (2023) develop a nonlinear Proxy-SVAR, which is used in this paper, and in an application to U.S. monetary policy find that: (i) a monetary easing shock can be an effective tool for stimulating the economy in recessions because it leads to a large reduction in unemployment and a small cost in terms of inflation, and (ii) a monetary tightening shock has large effects on prices and smaller effects on unemployment in expansions than in recessions. Overall, their results support the use of countercyclical monetary policy as it is associated with relatively favorable inflation-unemployment trade-offs. Contrary to previous evidence, Bruns and Piffer (2023), who develop a tractable method for estimating the parameters of the smooth transition function, find that monetary easing in the U.S. is more effective in recessions than monetary tightening in expansions.

While most of the research on nonlinear effects of monetary policy has focused on the sign or the business cycle, some others explore alternative state dependencies. Particularly relevant contributions are the ones analyzing the role of uncertainty for the monetary policy transmission. Using a medium-scale Threshold VAR, Castelnuovo and Pellegrino (2018) find that monetary easing has milder real effects and stronger nominal effects during periods of high macroeconomic uncertainty than in normal times.

[4] Recent advances in terms of shock identification come from Gertler and Karadi (2015), Arias et al. (2019), Caldara and Herbst (2019), Miranda-Agrippino and Ricco (2021), Jarociński and Karadi (2020) and Bauer and Swanson (2023b).

Aastveit et al. (2017) and Pellegrino (2021) use an Interacted-VAR to provide evidence on the lower real effects of monetary policy in the U.S. during (financial) uncertain times. A similar result is obtained by Pellegrino (2018) for the Euro area. Andreasen et al. (2023), using a similar empirical model, show that a more proactive monetary policy to stabilize output growth after an uncertainty shock is more effective in expansions than in recessions. Finally, other recent contributions are Falck et al. (2021) and Ascari and Haber (2022). The first paper uses nonlinear local projections to study the role of disagreement in inflation expectation for the transmission of U.S. monetary policy. They find that a monetary policy shock fails to stabilize real activity and inflation during periods of high disagreement, while the shock exhibits conventional responses during periods of low disagreement. The second paper documents that large monetary policy shocks lead proportionally larger initial responses of the price level, and that a monetary policy shock has large nominal effects and small real effects in a high trend inflation regime.

The second strand of the literature includes papers that focus on geopolitical risk. The main reference papers are Caldara and Iacoviello (2022) and Caldara et al. (2024). In the former, the authors develop a news-based index of GPR, and in the latter they present a systematic analysis of the effects of geopolitical risk in a large panel of countries. They find that on average supply-side disruptions dominate lower aggregate demand pressures during GPR episodes, leading to a simultaneous decline in economic activity and higher inflation. In an application to shipping freight rates, Drobetz et al. (2021) find that a positive shock to global GPR has an immediate positive but gradually decreasing effect on shipping freight rates, while a positive shock to economic policy uncertainty has a negative effect on shipping freight rates. Francis et al. (2022) examine the comovement of international business cycles and find that higher geopolitical risk is one of the leading indicators of international recessions. Jalloul and Miescu (2023) examine the comovement of equity returns across G7 countries and find that these returns are more correlated when GPR is at higher levels, with the results driven by geopolitical threats rather than their actual realization.

The rest of the paper is organized as follows. Section 2 presents an analysis of the geopolitical risk index and Section 3 describes the econometric methodology of the nonlinear Proxy-SVAR used. Section 4 presents the baseline empirical evidence on the GPR-dependent monetary effects and some robustness checks. Section 5 rationalizes these results through the monetary-fiscal interaction and then presents policy counterfactuals. Section 6 extends the baseline results by disentangling the GPR index and by providing policy experiments. Section 7 concludes the paper.

2 The geopolitical risk index

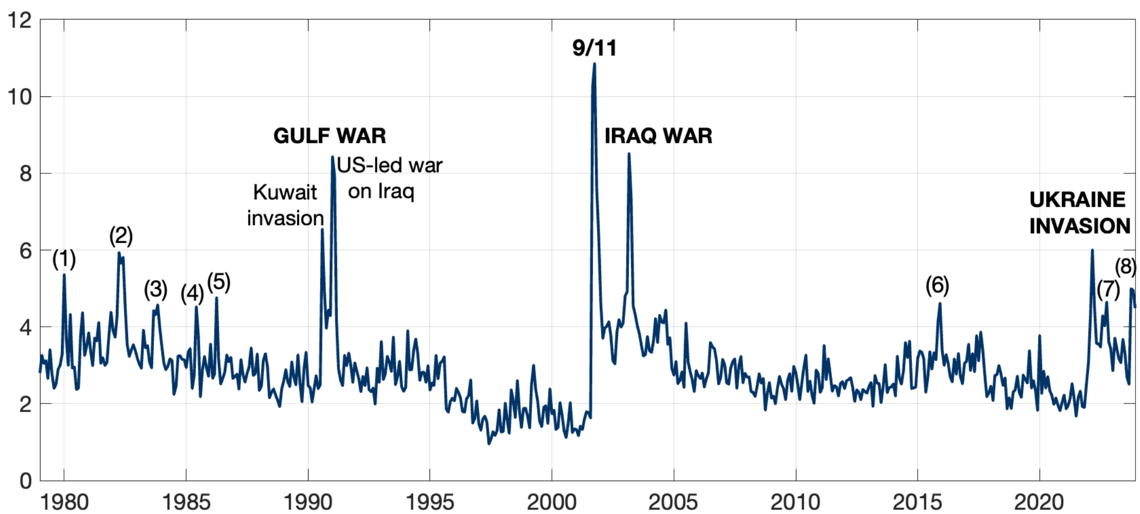
In this section, I explore the evolution of macroeconomic variables during periods of high GPR, as measured by the index of Caldara and Iacoviello (2022). This is important to understand the relevance of these periods for the effectiveness of monetary policy shocks. The authors define this risk as the threat, occurrence and escalation of adverse events related to war, terrorism and geopolitical tensions. To construct the index, they calculate the share of articles in leading newspapers that mention a dictionary of words related to adverse geopolitical events. Thus, this approach captures an index of how contemporaries perceived the risk associated with a specific geopolitical event, not an ex post perception. After constructing the index, Caldara and Iacoviello (2022) present evidence on the macroeconomic effects of a shock to the GPR index and find persistent contractionary effects on investment, employment and stock prices, which are stronger for more exposed industries. However, the authors do not examine the specific sources of geopolitical risk, which is left to future research. In a more recent paper, Caldara et al. (2024) show that the transmission of GPR to inflation involves supply disruptions, demand forces and a policy response. I now proceed narratively to examine the major events associated with increases in GPR.

Figure 1 plots the historical GPR index for the period 1979m7-2023m12. The major events in the sample are the following. In August 1990, Iraq invaded the wealthy Gulf state of Kuwait in an attempt to gain greater control over the lucrative Middle East oil supply. This event gave rise to the so-called Gulf War. In response, coalition forces led by the U.S. attacked Iraq in January 1991. In correspondence with these two events, in which oil was the most tangible interest, we observe two spikes in the GPR index. In 2001, we observe that the 9/11 terrorist attack corresponds to the largest spike in the index. The second largest spike in the GPR Index occurred in March 2003, when U.S. forces invaded Iraq in an oil-related war.[5] Looking at more recent years, the increase in the GPR index reflects (*i*) the Russian invasion of Ukraine in February 2022, which led

[5]This was also recognized by leading U.S. figures such as Alan Greenspan in his memoirs.

to a significant energy crisis that was particularly felt in Europe, and (*ii*) the Hamas terrorist attack in October 2023 and the subsequent large-scale Israeli invasion of Gaza, which led to heightened tensions in the Middle East, especially with Iran, one of the world's largest oil producers. Overall, it is clear that most of these episodes are related to the global oil market.

Figure 1: The geopolitical risk index



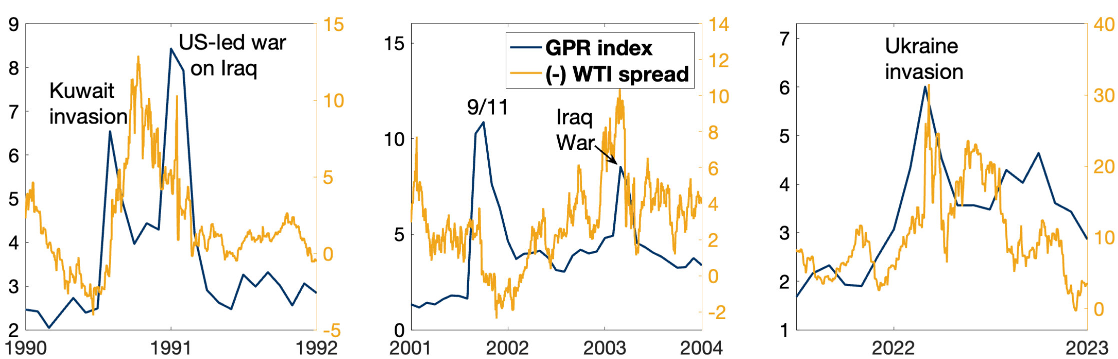
Notes: The solid blue line is the monthly U.S. historical geopolitical risk index of Caldara and Iacoviello (2022). The graph reports additional seven geopolitical events. (1) January 1980: US-Soviet tensions as the Soviet Union invaded Afghanistan in December 1979; (2) April 1982: Falkland war; (3) September 1983: Nuclear war scare; (4) June 1985: Hijacking of TWA Flight 847; (5) April 1986: US bombing of Libya; (6) November 2015: Paris terrorist attack; (7) October 2022: Crimea bridge attack and Russia strikes on Ukraine's energy infrastructure; (8) October 2023: Hamas Terrorist attack and Israeli invasion of Gaza.

Figure 2 provides additional information on the role of the oil market in geopolitical episodes by plotting the GPR index together with the negative of the daily WTI oil futures spread between 12-month and 1-month contracts (yellow line) around the major geopolitical events in the sample. As suggested by Alquist and Kilian (2010) and Kilian and Murphy (2014), this spread can be viewed as an indicator of fluctuations in the price of oil driven by a precautionary demand component of the real oil price due to news of future oil supply disruptions. This spread is multiplied by minus one to interpret positive spikes in this indicator as increases in this precautionary demand component. The rationale is as follows. If agents anticipate an increase in uncertainty about future oil supply, they will demand above-ground oil inventories or buy oil futures as insurance against future higher oil prices. This implies an immediate increase in the spot price or larger increase in short-term futures prices compared to longer-term futures prices.

Figure 2a shows that the increase in this indicator coincides with the start of the Gulf War in August 1990 with Iraq's invasion of Kuwait. After two months of steady increase, the spread reversed in October as the U.S. began planning an offensive to expel Iraqi forces from Kuwait. The spread continued to fall in the following months as the UN Security Council issued Resolution 678, calling for the use of "all necessary means" to force Iraq to withdraw from Kuwait by January 15, 1991. The failure of Iraqi forces to comply with UN Resolution 678 led to a spike in the spread on January 16 as uncertainty over oil supply production grew. However, this increase was quickly reversed when a U.S.-led coalition launched a campaign of air and missile strikes against targets in Iraq and Kuwait in the early morning hours of January 17, significantly reducing uncertainty about future oil supply conditions. This is an example of a GPR event related to the oil market, but involving news of favorable future supply conditions.

In Figure 2b, the 9/11 terrorist attack, which led to an 8% increase in the price of oil over the following five days and a subsequent 27% decline, is associated with a decrease in precautionary demand as world demand for crude oil fell, making a shortage less likely (see also Kilian, 2009; Alquist and Kilian, 2010). The Iraq war, on the contrary, is clearly related to news of future supply disruptions, as we observe an increase in the

Figure 2: Geopolitical risk and oil futures spread



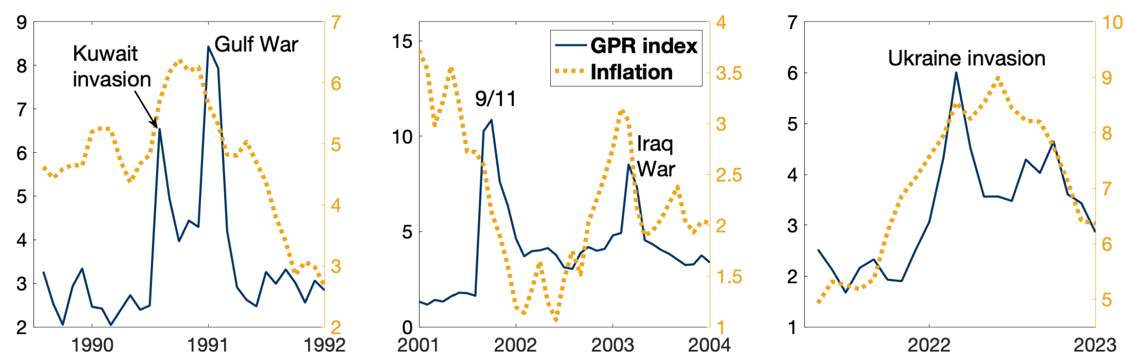
(a) *Gulf War* **(b)** *9/11 and Iraq War* **(c)** *Russia's war in Ukraine*

Notes: The solid blue line is the monthly U.S. historical geopolitical risk index of Caldara and Iacoviello (2022) and the yellow solid line is the (negative) of the daily oil futures spread (12-month to 1-month).

precautionary demand component of oil prices. Finally, Figure 2c shows that this was also the case during the Russian invasion of Ukraine. These conclusions are broadly in line with Pinchetti (2024), who distinguish shocks to the GPR index that involve a channel related to disruptions in energy markets from a second channel that can be attributed to lower aggregate demand pressures. A feature not covered in Pinchetti (2024) is that the oil-related channel can also include positive news about future supply production and thus, as will be shown in the next figure, be associated with deflationary pressures. This is the case for the second phase of the Gulf War.

I now analyze the inflation dynamics during the major geopolitical events in the sample. Figure 3 plots the two series for the Gulf War (3a), 9/11 and the Iraq War (3b) and Russia's war in Ukraine (3c). The figure shows that periods of heightened geopolitical tension are associated with higher inflation rates when these events are related with increased uncertainty about future supply disruptions in the oil market as discussed earlier. From a central bank perspective, these episodes represent a high risk of losing the inflation anchor, as (expected) inflation is well above the central bank's target. On the contrary, when higher geopolitical tensions are associated with lower aggregate demand pressures or news of favorable future oil supply conditions, the inflation rate falls significantly. The same argument holds for inflation expectations (see Figure C.1 in Appendix C).

Figure 3: Geopolitical risk and inflation



(a) *Gulf War* **(b)** *9/11 and Iraq War* **(c)** *Russia's war in Ukraine*

Notes: The solid blue line is the monthly U.S. historical geopolitical risk index of Caldara and Iacoviello (2022) and the yellow dotted line is the inflation rate.

As suggested by Caldara et al. (2024), geopolitical events also involve a policy response. Specifically, it leads to an increase in defense spending, especially in the United States. Therefore, I now jointly analyze the dynamics of the GPR index and the annual growth rate of real federal defense spending. All the plots in Figure 4 show that these geopolitical episodes are associated with increases in defense spending. This is an important association because while fiscal policy tends to be expansionary during all geopolitical events, central bank actions must be conditional on the inflation dynamics associated with the specific geopolitical episode.

Overall, the GPR index of Caldara and Iacoviello (2022) appears to operate broadly through two main mechanisms, one of which is associated with negative news about future supply disruptions in the oil market and the other with a lower aggregate demand component.[6] While the first type of geopolitical event is associated with inflationary pressures that cause central banks to raise interest rates, the second type of event is associated with deflationary forces that cause central banks to adopt an expansionary stance. From a fiscal policy perspective, all geopolitical episodes imply an increase in defense spending. The result is a policy mix that is not always coordinated.

Figure 4: Geopolitical risk and defense spending

(a) *Gulf War* **(b)** *9/11 and Iraq War* **(c)** *Russia's war in Ukraine*

Notes: The solid blue line is the monthly U.S. historical geopolitical risk index of Caldara and Iacoviello (2022) and the yellow dotted line is the real federal national defense consumption expenditures and gross investment (annual rate).

[6]News of favorable future oil supply conditions, such as the U.S.-led war against Iraq during the Gulf War, can be included in this second type of geopolitical event.

3 Econometric methodology

In this section, I present the econometric methodology used in this paper. I begin with a brief overview of the DFGS framework, followed by a detailed description.

3.1 Overview

In this econometric framework, the underlying economy is represented by a structural Vector Moving Average including potentially many nonlinear terms of the shock of interest. Under appropriate conditions, the variables of interest have a VARX representation, where the shock and its nonlinear functions are the exogenous variables which, however, are unobserved. To overcome this limitation, I leverage on the key result of DFGS: I can estimate the shock of interest as the standardized fitted value of the regression of a valid instrument on the residuals of a standard linear VAR. The requirement of this methodology is that the endogenous variables in the model are *informationally sufficient*, which postulates partial invertibility.[7] After testing the validity of the invertibility assumption, I can obtain a consistent estimate of the shock which allows the estimation of the VARX and obtain nonlinear impulse responses.

3.2 Nonlinear representation

Let x_t be a n-dimensional vector of stationary macroeconomic variables with the following structural representation

$$x_t \;=\; \nu + \alpha(L)u_t + \beta(L)d_t u_t + \Gamma(L)\xi_t, \tag{1}$$

where ν is a vector of constants, u_t is the monetary policy shock, d_t is the geopolitical risk index of Caldara and Iacoviello (2022), and ξ_t is a m-dimensional vector of structural shocks other than the monetary policy shock, collected in vector u_t^-, and possibly nonlinear functions of these shocks. I further assume that the vector $[u_t, u_t^-]'$ is *i.i.d.* zero mean with an identity covariance matrix.[8] The vector $\alpha(L) = \alpha_0 + \alpha_1 L + \alpha_2 L^2 ...$

[7] Partial invertibility means that a subset of structural economic shocks can be estimated as linear combinations of the residuals of the linear projection of a vector of variables onto their past values (see also: Forni and Gambetti, 2014; Forni et al., 2019).

[8] Notice that shock serial and mutual independence implies that all structural shocks are uncorrelated with the lags of $d_t u_t$ and x_t.

includes the impulse response functions to the monetary policy shock; the vector $\beta(L) = \beta_0 + \beta_1 L + \beta_2 L^2 + ...$ includes the impulse response functions to the interaction term between the monetary shock and the geopolitical risk index. Finally, $\Gamma(L) = \Gamma_0 + \Gamma_1 L + \Gamma_2 L^2 + ...$ is a $n \times m$ matrix of impulse response functions to the remaining structural shocks and their non-linear functions.

The nonlinear impulse response functions to the monetary policy shock are derived by combining the two terms $\alpha(L)$ and $\beta(L)$. Formally, the total effect of a monetary policy shock $u_t = \bar{u}$ is then given by the sum of the linear and nonlinear term:

$$IRF(u_t = \bar{u}) = \alpha(L)\bar{u} + \beta(L)d_t\bar{u}. \tag{2}$$

The total responses defined in equation (2) simply correspond, in this nonlinear context, to the Generalized Impulse Response Functions defined as $E(x_{t+h}|u_t = \bar{u}) - E(x_{t+h}|u_t = 0)$, $h = 0, 1, ...$ I discuss below how to estimate the model and the implied impulse response functions.

From (2), it is obvious that the impulse response is nonlinear given the interaction term $d_t u_t$. By choosing different values of d_t, we can characterize the effects of a monetary policy shock in periods of, say, high geopolitical risk.

Stationarity of $\Gamma(L)\xi_t$ ensures the existence of the following representation:

$$x_t = \nu + \alpha(L)u_t + \beta(L)d_t u_t + \Phi(L)e_t, \tag{3}$$

where $\Phi(L)e_t$ is the Wold representation of $\Gamma(L)\xi_t$.[9] Under the assumption of invertibility of the Wold representation, the structural model in (1) implies the existence of the following VARX representation:

$$A(L)x_t = \mu + \tilde{\alpha}(L)u_t + \tilde{\beta}(L)d_t u_t + e_t \tag{4}$$

where $\mu = A(1)\nu$, $A(L) = \Phi(L)^{-1}$, $\tilde{\alpha}(L) = \Phi(L)^{-1}\alpha(L)$ and $\tilde{\beta}(L) = \Phi(L)^{-1}\beta(L)$.[10]

[9] If the structural representation $\Gamma(L)\xi_t$ is invertible, then $\Phi(L) = \Gamma(L)\Gamma_0^{-1}$ and $e_t = \Gamma_0\xi_t$, Γ_0^{-1} being either the inverse of Γ_0, if $m = n$, or a left inverse of Γ_0, if $m < n$. If the structural representation $\Gamma(L)\xi_t$ is not invertible (e.g. when $m > n$ or $m = n$ but $\Gamma(L)$ vanishes on the unit circle), then e_t is a linear combination of the present *and past* values of ξ_t and $\Phi(L)e_t$ is just a statistical represention, devoid of economic meaning.

[10] I make the standard assumption in the VAR literature that all the matrices of polynomials in L can be approximated by finite order matrices.

The above representation is equivalent to:

$$x_t = \mu + \tilde{A}(L)x_{t-1} + \tilde{\alpha}(L)u_t + \tilde{\beta}(L)d_t u_t + e_t$$

where $\tilde{A}(L) \equiv [A_1 + A_2 L + \cdots + A_p L^{p-1}] = -[A(L) - I]/L$.

Notice also that stationarity of x_t implies the existence of the Wold representation. Under the assumption that such representation is invertible, the vector x_t admits the following VAR representation:

$$x_t = \vartheta + B(L)x_{t-1} + \epsilon_t = \vartheta + \sum_{j=1}^{\infty} B_j x_{t-j} + \epsilon_t \tag{5}$$

where ϵ_t is orthogonal to x_{t-j}, $j = 1, \ldots, \infty$.

Next, we derive the relation between the VARX representation in (4) and the VAR representation in (5). Starting from equation (4), the linear projection of $\tilde{\alpha}(L)u_t + \tilde{\beta}(L)d_t u_t$ onto a constant and the past history of x_t gives

$$\tilde{\alpha}(L)u_t + \tilde{\beta}(L)d_t u_t = \theta + C(L)x_{t-1} + w_t.$$

From the above equation, we get that $\vartheta = \mu + \theta$, $B(L) = \tilde{A}(L) + C(L)$ and $\epsilon_t = e_t + w_t$.

Notice that, if $\tilde{\beta}(L) = 0$, the structural representation in (1) reduces to a linear model and standard SVAR analysis can be conducted using the model in (5). Hence, the linear model is nested in this more general framework. It is straightforward to test for nonlinearity, and it can be done in two ways. One is to test for the null $\tilde{\beta}(L) = 0$ in equation (4), or alternatively for $\beta(L) = 0$ in the impulse-response functions in (2).

3.3 Identification

In the previous subsection, I presented the conditions under which the nonlinear economy admits the VARX representation in (4). At this point, one might be tempted to proceed with direct estimation of the model in (4) or an equivalent local projection. Unfortunately, the exogenous variables in the model are unobservable. The econometrician rarely observes the shock of interest and, in most cases, has only a proxy variable that is a noisy measure of the actual shock. In a linear framework, the estimated parameter is still affected by an attenuation bias, which, being proportional for all variables, can be eliminated by a standard normalization (Stock and Watson, 2018). In a nonlinear model, instead, the estimated parameters for the nonlinear term would generally be biased in a way that is not proportional across variables and cannot be corrected in any way. Therefore, direct estimation of the VARX in equation (4) is not feasible (see Forni et al., 2023a, for a formal discussion of this point).[11] Below I discuss how to obtain a consistent estimate of the exogenous shock that can be used to estimate impulse responses within this econometric framework.

The identification procedure relies on two standard assumptions in the proxy-SVAR literature. The first requires the existence of a valid instrument, as specified below.

Assumption A1 (*Proxy*). The proxy z_t is given by

$$z_t = a + bu_t + \delta(L)' x_{t-1} + v_t, \tag{6}$$

where $\delta(L)$ is a vector of polynomials of degree p in the lag operator L, $b \neq 0$ and v_t is an error independent of the structural shocks at all leads and lags. Notice that under Assumption A1, the standard conditions for a valid instrument, i.e. $\text{cov}(z_t, u_t) = b \neq 0$ (relevance) and $\text{cov}(z_t, \xi_t) = 0$, are satisfied (see Mertens and Ravn, 2013; Stock and Watson, 2018).

The second assumption ensures that the monetary policy shock can be estimated as a combination of current and past data and is formalized below.

Assumption A2 (*Informational sufficiency*). The monetary policy shock is a linear combination of the current and past values of the endogenous variables x_t.

Assumption A2 postulates "partial invertibility" of u_t, i.e. that the variables in x_t are informationally sufficient to find the monetary policy shock. In other words, the nonlinear terms are not needed to estimate the shock of interest. Notice that the same assumption has to hold also in the linear case in order for the standard procedure to be valid. Fortunately this is a testable assumption. In the empirical section I will assess whether it holds.

Under assumptions A1 and A2, DFGS show that the shock of interest can be obtained as the standardized fitted value of the linear projection of z_t onto the VAR

[11] If the monetary policy shock were perfectly observable, then equation (4) or a local projection version of it could be estimated by OLS.

innovations ϵ_t. This is the basic result underlying the proposed procedure. First, we get an estimate of the monetary policy shock by using a standard proxy SVAR method; having an estimate of the shock, we can use it to estimate the VARX representation in (4).

3.4 Estimation

More in detail, the estimation procedure is the following.

I. Estimate (5) with OLS to obtain consistent estimates of the residuals ϵ_t, call them $\hat{\epsilon}_t$.

II. Estimate the first-stage regression

$$z_t = \lambda' \hat{\epsilon}_t + \eta_t. \tag{7}$$

Following Forni et al. (2023b), an estimate of the normalized shock is obtained by standardizing the fitted value of the above regression, i.e. $\hat{u}_t = \hat{\lambda}' \hat{\epsilon}_t / \text{std}(\hat{\lambda}' \hat{\epsilon}_t)$.

III. Estimate equation (4) using as regressors the current value and the lags of the estimated shocks $\hat{u}_t$ and its nonlinear function $d_t \hat{u}_t$. This gives the estimates of $A(L)$, $\tilde{\beta}(L)$ and $\tilde{\alpha}(L)$. From these parameters, one can obtain estimates of $\alpha(L) = A(L)^{-1}\tilde{\alpha}(L)$ and $\beta(L) = A(L)^{-1}\tilde{\beta}(L)$.

IV. Finally, compute the impulse responses according to equation (2).

3.5 Comparison with alternative nonlinear models

I find the model used in this paper particularly appealing compared to alternative nonlinear models for a number of reasons. First, unlike the FAIR approach proposed by Barnichon and Matthes (2018), it does not require the assumption of a specific shock distribution. Second, the alternative nonlinear Proxy-SVAR implemented in Pellegrino et al. (2023) it has greater flexibility to handle nonlinearities directly related to the shock. Finally, compared to nonlinear local projection methods (e.g., Tenreyro and Thwaites, 2016), it has a more parsimonious parameterization, reducing estimation uncertainty, and it is internally consistent, integrating the identification of the shock and the estimation of nonlinear impulse responses in a single framework.

4 Baseline results

In this section, I present novel evidence on the effects of monetary policy shocks in periods of high GPR in comparison to normal times. I define the economy to be in a high GPR regime when the Caldara and Iacoviello (2022)'s index is at the 96th percentile ($d_{t,P_{96}} = 4.57$), and normal times as when the index is at its median value ($d_{t,P_{50}} = 2.73$) in the baseline (pre-Covid) sample.[12] The choice of the 96th percentile to define a period of high GPR is motivated by the positive skewed distribution of the index, with a skeweness of 2.7. This is clearly visible from the plot in Figure 1. In Section 6.1, I perform some policy experiments by studying the effects of monetary policy shocks at different historical levels of GPR.

The endogenous variables of the VAR is set accordingly to Miranda-Agrippino and Ricco (2021) and includes the 1-year Treasury rate, the log of industrial production, the log of the consumer price index, the unemployment rate, the log of a commodity price index and the Gilchrist and Zakrajšek (2012) excess bond premium spread. As a first step, I select the VAR lag length by standard information criteria.[13] I then estimate the VAR over the sample 1979m7 to 2019m12 by ordinary least squares.[14] To estimate the first-stage regression in (7), I use the instrument of Miranda-Agrippino and Ricco (2021) extended by Degasperi and Ricco (2021) up to 2017m12. Before discussing the baseline results, I present two tests. The first concerns the validity of Assumption A2. Specifically, I make sure that the endogenous variables of the VAR are sufficient to identify the monetary policy shock. The second relates to testing whether the proposed GPR-dependent effects of monetary policy are statistically relevant.

4.1 Testing for invertibility

To test whether the assumption of *informational sufficiency* (Assumption A2) holds, I adopt the testing procedure developed in Forni et al. (2023b) by checking that the shock u_t is indeed invertible. The test is based on the theoretical result that, if the shock

[12]In the baseline VAR estimation the sample ends before the outbreak of the Covid pandemic (2019m12), while Figure 1 plots the GPR index up to 2023m12. Over the longer sample, the values of $d_{t,P_{96}}$ and $d_{t,P_{50}}$ are identical up to the first decimal. As a robustness check, I provide evidence for the VAR estimated up to 2023m12.

[13]I take the average of the AIC, BIC and HQC, which is 6.

[14]I choose the starting point to coincide with the beginning of Volcker's tenure as Federal Reserve Chairman, as it marked a change in the conduct of monetary policy (see, e.g., Bianchi, 2013).

is non-invertible, then it is a function of current and future VAR residuals, instead of a combination of current residuals only. More specifically, the test requires regressing the instrument z_t on the first r leads of the Wold residuals $\hat{\epsilon}_t$. Formally

$$z_t = \sum_{k=0}^{r} \lambda_k' \epsilon_{t+k} + \eta_t \tag{8}$$

The invertibility test proposed by Forni et al. (2023b) is an F-test for the significance of the r leads, where the null hypothesis is $H_0 : \lambda_1, \lambda_2, \ldots, \lambda_k = 0$ against the alternative that at least one of the coefficients is non-zero. I estimate the regression in equation (8) using different numbers of leads ($6 \leq r \leq 12$). The results, shown in Table (1), indicate that we cannot reject the null hypothesis of invertibility for all values of r at the 10% confidence level (at 5% for $r = 9$), which is a very conservative level.

Table 1: Invertibility test

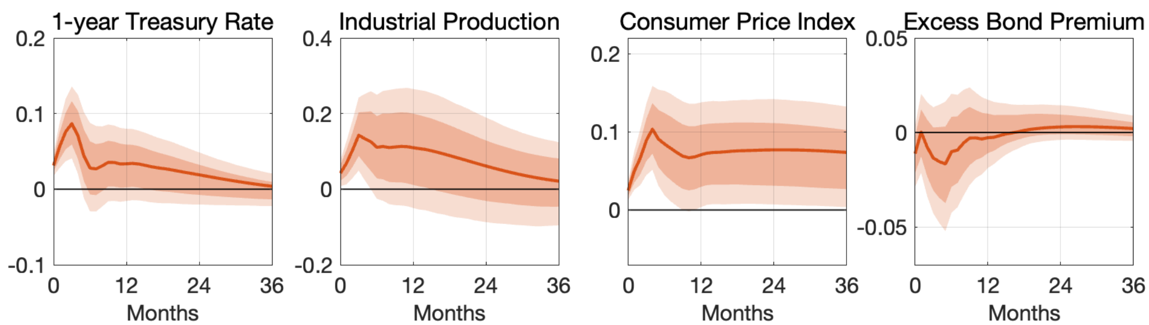
	Number of leads r						
	$r=6$	$r=7$	$r=8$	$r=9$	$r=10$	$r=11$	$r=12$
p-value	0.195	0.211	0.160	0.094	0.148	0.136	0.178

The table shows the p-values for each regression including the current value and up to r leads of the Wold residuals. Values above the confidence level (1%, 5%, 10%) indicates that the shock is invertible.

4.2 Are nonlinearities important?

In this section, I formally test for nonlinearity by testing the null hypothesis H_0 : $\beta(L) = 0$ in equation (2). Specifically, Figure 5 informs about the statistical relevance of the nonlinear term $d_t u_t$ in equation (2) by analyzing the corresponding impulse responses $\beta(L)$ and the 68% and 90% confidence intervals. Industrial production and prices both respond positively to the interaction term of the monetary policy shock with the GPR index. However, the response of industrial production is statistically different from zero at the 90% confidence level only for the first four months, while the response of prices is positive and statistically different from zero for the entire forecast horizon. Lastly, for the excess bond premium, which measures a residual component of the spread between an index of corporate bond yields and the yield on a government bond of similar maturity after removing the default risk component (e.g., risk aversion), I find no statistical support of nonlinearitiy. To provide evidence on a broader measure of credit spread, Figure C.2 in Appendix reports the test when using an alternative model specification using the Baa-Aaa corporate bond spread. Using this alternative specification, the nonlinear term has a statistically negative impact on this broader measure of credit spreads in the first two months after the shock.

Figure 5: Monetary policy and geopolitical risk: Test of nonlinearity



Notes: Six-variable VAR. Impulse responses to the nonlinear term in equation (2) estimated using a nonlinear Proxy-SVAR. The solid lines are the point estimates and the shaded areas are 68% and 90% confidence bands.

4.3 The GPR-dependent effects of monetary policy

Figure 6 plots the impulse responses of the 1-year Treasury rate, the industrial production, the consumer price index and the excess bond premium to a monetary policy shock conditional on periods of normal times (first row, blue line) and periods of high GPR (second row, red line).[15] These responses are obtained by summing the terms $\alpha(L)$ and $\beta(L)d_{t,P_k}$, where $k = (50, 96)$, estimated from equation (2).[16] The shock is normalized to induce a 100 basis point increase in the 1-year Treasury rate. The impulse responses of the interest rate and the excess bond premium are measured in percentage point changes, and those of industrial production and prices are measured in percentage changes. The shaded areas are 68% and 90% confidence bands based on 10,000 bootstrap replications.[17] The maximum horizon of the impulse response

[15]The responses for all variables included in the model are reported in Appendix Figure C.4.

[16]The VARX includes five lags for both endogenous and exogenous variables. As explained in subsection 3.2, the number of lags in the VARX must be lower than the number of lags in the first-step VAR used to identify the shock to avoid collinearity problems.

[17]The confidence bands include the estimation uncertainty of the estimated monetary shock from the first-stage in (7), because both steps of the estimation are included in the bootstrapping procedure.

functions is set at 36 months.

In normal times, a tightening shock leads to impulse responses consistent with standard macroeconomic theory: output and prices contract sluggishly while the excess bond premium rises at impact. The magnitude of these effects are also in line with recent contributions (e.g., Miranda-Agrippino and Ricco, 2021; Bauer and Swanson, 2023b). Some interesting facts emerge when these responses are compared with those in periods of high GPR. A tightening shock of the same size leads to a temporary increase in output and prices, although only the former is statistically significant. These puzzling responses are then followed by a decline, which is not statistically different from zero at all forecast horizons. The excess bond premium increases at impact by less than in normal times before returning to its initial level a few months after the shock.

The third row plots the difference in responses between normal times and the high GPR regime. Comparatively, output is significantly less responsive to a monetary policy shock in periods of high GPR, at least for the first six months at the 90% confidence interval. Instead, the difference in price responses emerges as the most important GPR-dependent effect of monetary policy, with a statistically significant and persistent difference over the entire forecast horizon. Lastly, as expected, the excess bond premium responds similarly across the two regimes. The results from the alternative model specification with the Baa-Aaa spread, reported in Appendix C Figure C.3, confirm the previous findings for output and prices, but also indicate that while the Baa-Aaa spread rises on impact in normal times, it rises with a delay when GPR is high, resulting in a significant difference between the two regimes. Overall, the evidence presented in this section points to the ineffectiveness of monetary policy shocks when GPR is high.

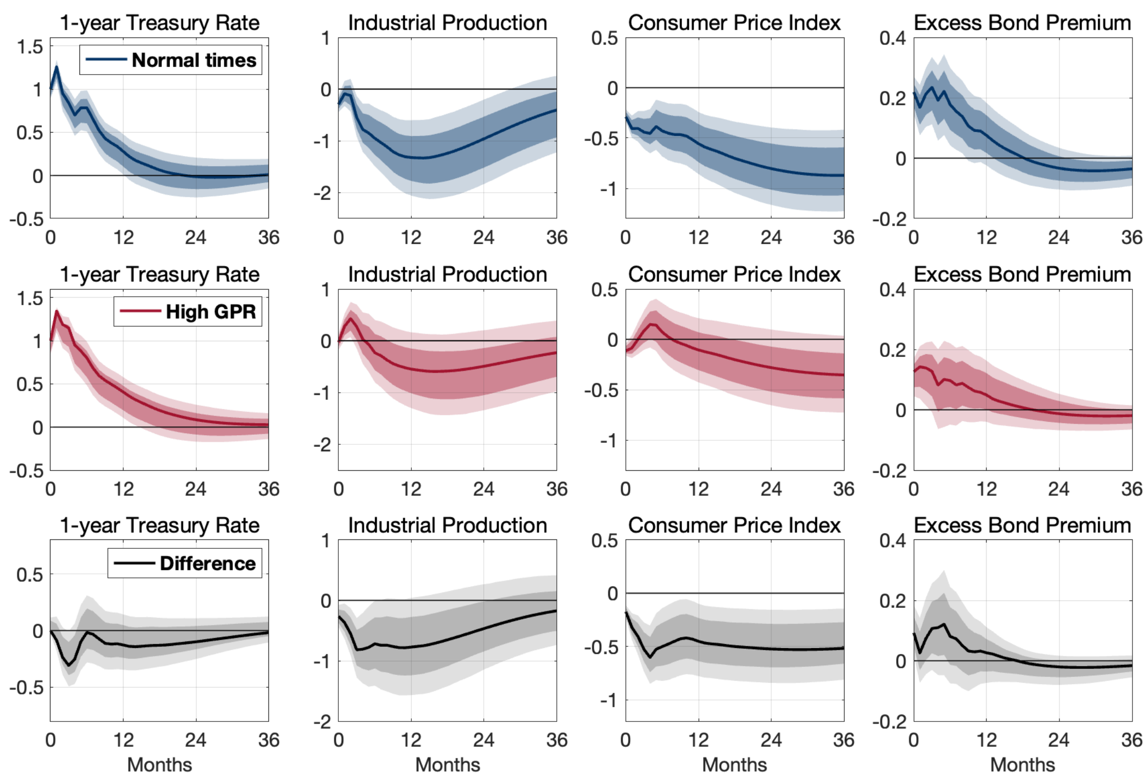
From an economic standpoint, the results presented in this section provide a novel empirical evidence pointing to the ineffectiveness of monetary policy shocks when GPR is high.

Further evidence of the documented state-dependent effects of monetary policy comes from Figure 7, which plots the impulse responses of expected inflation in the two regimes and their difference. The dynamic response of this variable is crucial to understanding the central bank's ability to achieve and maintain price stability. To obtain these responses, I augment the baseline specification with inflation expectations from the Michigan Survey of Consumers. Not suprisingly, households' inflation expectations respond sluggishly to monetary policy (D'Acunto et al., 2024), but they decline with a larger lag during periods of high GPR relative to normal times. Overall, the figure shows that monetary policy is also less powerful in influencing expectations in periods of high GPR. This is true in the short run.[18]

4.4 Sensitivity checks

In this Section, I present some sensitivity checks concerning the identification of monetary policy shocks, the estimation sample, the state variable and the model used to

Figure 6: Monetary policy and geopolitical risk



Notes: Six-variable VAR. Nonlinear impulse responses from a monetary policy shock in normal times (50th percentile of the U.S. GPR index, blue lines), high GPR (96th percentile of the U.S. GPR index, red lines) and the difference between the two impulse responses (black lines). The shock is normalized to induce a 100 basis point increase in the 1-year treasury rate. The solid lines are the point estimates and the shaded areas are 68% and 90% confidence bands.

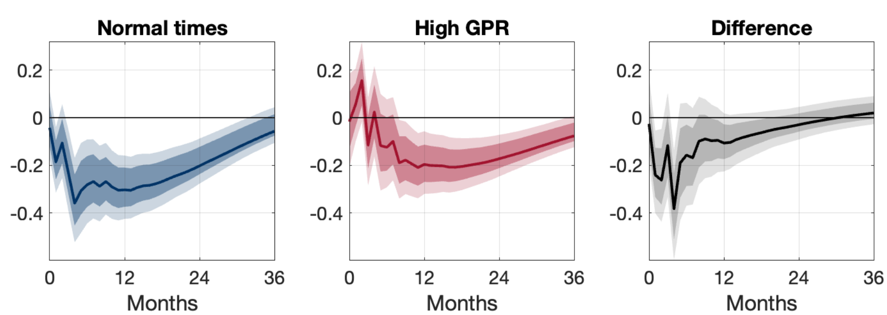
[18]For a formal test of nonlinearity for expected inflation, see Figure C.5 in Appendix.

estimate nonlinear impulse responses.

Shock identification. The results presented are obtained using the high-frequency monetary policy surprise series of Miranda-Agrippino and Ricco (2021), which accounts for the presence of information frictions using the Greenbook forecasts. This methodology has proven to be one of the many alternatives to *adjust* the identified series of high-frequency changes in interest rates around FOMC announcements. Appendix B presents an overview of this issue and assesses the robustness of the documented state-dependent effects of monetary policy shocks using the alternative instruments of Jarociński and Karadi (2020) and Bauer and Swanson (2023b). Figure B.1 in the Appendix B plots the baseline impulse responses and confidence bands together with the point estimates, in dashed and dotted lines, obtained using the alternative instruments to identify a monetary policy shock. Despite some differences in the estimates in the two regimes, which are due to the way the proxies are constructed, the state-dependent results obtained with the alternative instruments are fully consistent with the baseline results.

Estimation sample. The estimation sample includes a sustained period of time in which the zero lower bound (ZLB) on nominal interest rates was binding but also excludes recent geopolitical tensions as it ends in the pre-Covid period. Thus, I now assess whether the results are *(i)* driven by periods when the ZLB constraint was bind-

Figure 7: Monetary policy and geopolitical risk: Inflation expectations



Notes: Seven-variable VAR. Nonlinear impulse responses from a monetary policy shock in normal times (50th percentile of the U.S. GPR index, blue lines), high GPR (96th percentile of the U.S. GPR index, red lines) and the difference between the two impulse responses (black lines). The shock is normalized to induce a 100 basis point increase in the 1-year treasury rate. The solid lines are the point estimates and the shaded areas are 68% and 90% confidence bands.

ing and *(ii)* hold when the sample is extended to include recent geopolitical events. First, Figure C.13 in the Appendix C plots the baseline impulse responses and confidence bands together with the point estimates when the sample ends at the start of the ZLB (2008m11) in dashed-dotted lines. Second, I estimate nonlinear impulse responses over the sample ending in 2023m12, controlling for the lockdown imposed in response to the outbreak of the Covid pandemic using a dummy variable. Figure C.14 in Appendix C plots the baseline impulse responses and confidence bands together with the point estimates when the sample ends at 2023m12 in dashed-dotted line. Overall, the baseline findings turn out to be coherent with respect to both alternative sample specifications.

The role of the business cycle. Figure C.7 plots the GPR index along with the official NBER recession periods. Given that geopolitical episodes in the first half of the sample occur at the beginning of or during recessions, one might be concerned that the state-dependent results documented in this paper are not due to geopolitical episodes but rather to periods of recession. First, I find that the correlation of the GPR index with the unemployment rate, as an indicator of labor market slack, is quite low at 0.18.[19] Then, to formally address these concerns, I run a race against this potential alternative explanation for the results obtained in this paper. To do so, I extend the baseline model to include an additional interaction term of the shock with a dummy indicating periods of recession. Formally, I estimate the following VARX

$$x_t = \mu + \tilde{A}(L)x_{t-1} + \tilde{\alpha}(L)u_t + \tilde{\beta}(L)d_t u_t + \tilde{\Gamma}(L)r_t u_t + e_t \tag{9}$$

where r_t denotes official NBER recession periods. To test the relevance of this potential alternative source of state-dependent effects of monetary policy, I can look at the coefficients associated with the interaction of the monetary shock with the GPR index, i.e., $\beta(L)$. Figure C.8 in Appendix C presents the nonlinearity test by plotting the coefficients $\beta(L)$ as orange solid lines together with their confidence bands in orange shaded areas. The figure also shows the baseline estimates for the point estimates and the 90% confidence bands as dashed lines. As the test shows, the role of GPR for the effect of monetary policy shocks is largely unaffected in the extended model.

[19] Caldara and Iacoviello (2022) find that macroeconomic, financial, and uncertainty variables do not granger cause the GPR index.

If anything, the test using the model in (9) indicates a larger role of GPR for prices and the EBP, which is still negative but now also statistically significant. Overall, this sensitivity check indicates that the role of the GPR in the effectiveness of monetary policy shocks is not driven by periods of recession.

Local projections. An important assumption behind the VAR approach is that the model adequately captures the true generating process of the data. Jordà (2005) proposed a direct approach to estimating impulse responses that does not require specification and estimation of the underlying multivariate dynamic system. In a linear setting, this advantage over VARs translates into lower bias, but at the cost of higher variance (Li et al., 2024).

Local projections in this nonlinear setting involve running the following set of regressions:

$$y_{i,t+h} = c_h + \alpha_h u_t + \beta_h d_t u_t + \Gamma_h(L) w_{t-1} + \xi_{t+h} \qquad \text{for} \qquad h = 0, 1, 2, \dots,$$

where c_h is a vector of constants, $y_{i,t+h}$ is the outcome variable of interest, u_t is the monetary policy shock identified from the Proxy-SVAR, d_t is the geopolitical risk index of Caldara and Iacoviello (2022), $\Gamma(L)$ is a polynominal in the lag operator of order five, w_{t-1} is a vector of lagged controls in line with the VAR and $\xi_{t,h}$ is a potentially serially correlated error term. Note that the direct projection using the proxy z_t of Miranda-Agrippino and Ricco (2021) as the *true* shock u_t in a nonlinear setting would yield unbiased OLS estimates only if the proxy is a noisy-free measure of the shock. In other words, z_t should not be contaminated by measurement error, a condition that does not hold in general (see also Section 3.3).

The total effect of monetary policy shocks at horizon h is obtained by summing up the linear and nonlinear term as follows:

$$IRF_h(u_t = \bar{u}) = \alpha_h \bar{u} + \beta_h d_t \bar{u}$$

Figure C.15 in Appendix C plots the baseline responses and confidence bands together with those from local projections in dashed-dotted lines. As expected, the impulse responses estimated using local projections are more erratic as they are not bounded by any dynamic restriction over their horizon. At short horizons, the impulse responses estimated using local projections are equivalent. This is not surprising given the results in Plagborg-Møller and Wolf (2021). At longer horizons, some responses show erratic jumps. Nevertheless, the GPR-dependent monetary effects are broadly confirmed using local projections. However, the erratic behavior of the estimates using the alternative approach suggests that the benefits of local projections in this nonlinear setting are quite small compared to their costs.

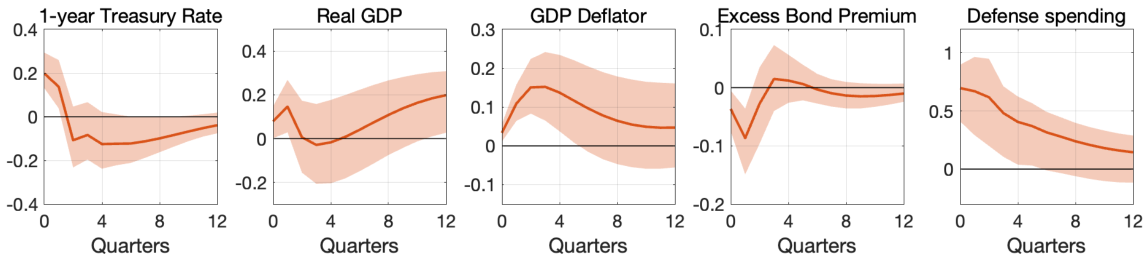
5 The monetary-fiscal interaction

The baseline results presented in the previous section indicate that during periods of high GPR, a tightening shock has expansionary or non-significant real and nominal effects. What could be driving these results? In this section, I shed light on the mechanism behind these results, drawing on the monetary-fiscal interaction that takes place during geopolitical tensions. To this end, I extend the baseline model to include the growth rate of real federal defense spending as the seventh endogenous variable of the model.[20] Since national accounts are only available at a quarterly frequency, I am left with two options. The first is to estimate a lower frequency model, which would necessarily imply less precise estimates. The second is to use a temporal disaggregation method (e.g., Chow-Lin) to obtain a monthly index of real federal defense spending. I use the quarterly VAR as the baseline model, but to account for the shortcomings of each strategy, I also include the same exercise using the monthly model. It turns out that they produce comparable results and convey the same message.

Before presenting the impulse responses in both regimes, I conduct a formal test of nonlinearity in the quarterly model. Figure 8 plots the coefficients and confidence bands associated with the interaction term $d_t u_t$ in (2). The impulse responses include only 68% confidence bands because the quarterly model produces less precise estimates than its monthly counterpart. The test indicates that real defense spending responds positively to the interaction term, which is statistically significant for the first year and a half.

[20]Notice that while this variable is correlated with the GPR index, it is uncorrelated with the interaction term between the monetary shock and the GPR index, which is an exogenous component in my model.

Figure 8: Policy mix: Test of nonlinearity



Notes: Seven-variable quarterly VAR. Impulse responses to the nonlinear term in equation (2) estimated using a nonlinear Proxy-SVAR. The solid lines are the point estimates and the shaded areas are 68% confidence bands.

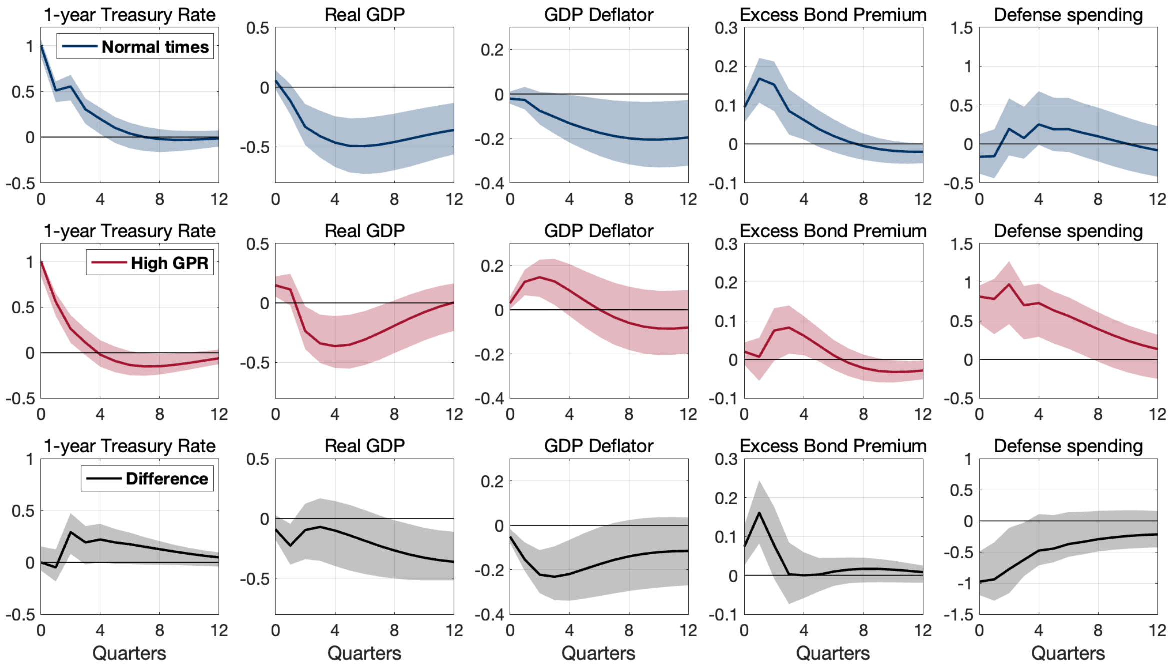
Figure 9 plots the impulse responses of the seven-variable quarterly VAR, which augment the baseline specification with the fiscal variable. In this model, real GDP and the GDP deflator substitute the industrial production index and the consumer price index. This is done to avoid additional bias from aggregating monthly indicators to a quarterly frequency. The results indicate that while in normal times defense spending does not respond in a statistically significant way to a monetary tightening shock, it increases at impact when GPR is high. This may be the case because geopolitical events are characterized by large U.S. military buildups. These results point to the unfavorable policy mix in place when GPR is high, where the monetary shedding light on the mechanisms underlying the ineffectiveness of monetary policy shocks in these periods.[21]

Figure C.9 in Appendix plots the impulse responses for the monthly model.[22] The responses are broadly in line with the lower frequency model, confirming the unfavorable policy mix following a monetary tightening shock. The major difference concerns the response of federal defense spending when GPR is high. In the quarterly model, this variable reacts at impact with a larger magnitude before converging to the monthly counterpart response within a few months following the shock.

[21]I consider defense spending a good indicator of the fiscal stance as nondefense spending is a minor part of federal spending over the (pre-covid) sample used.

[22]As the relevant monthly indicators for Chow-Lin, I include Industrial Production Equipment: Defense and Space Equipment and the GPR index. The choice of the latter variable is dictated by the fact that geopolitical events in the U.S. have been accompanied by large military buildups.

Figure 9: Monetary policy and geopolitical risk: Policy-mix



Notes: Seven-variable quarterly VAR. Nonlinear impulse responses from a monetary policy shock in normal times (50th percentile of the U.S. GPR index, blue lines), high GPR (96th percentile of the U.S. GPR index, red lines) and the difference between the two impulse responses (black lines). The shock is normalized to induce a 100 basis point increase in the 1-year treasury rate. The solid lines are the point estimates and the shaded areas are 68% confidence bands.

5.1 Policy counterfactuals

In this subsection, I examine the role of the monetary-fiscal interaction in explaining the GPR-dependent effects of monetary policy shocks. Specifically, I investigate whether this interaction mechanism alone can account for the observed pattern of initially expansionary followed by weakened effects of a contractionary monetary shock in periods of high GPR. To assess this, it would be ideal to obtain a counterfactual scenario in which the impulse response of defense spending does not differ from that in normal times. From there, a simple comparison of the effects of monetary policy shocks in the two regimes can provide information on the relevance of the monetary-fiscal interaction for the documented GPR-dependent nonlinearity.

A recent paper by McKay and Wolf (2023) (hereafter MKW) provides the empirical tools to estimate such a counterfactual. The novelty of their method is to subject the economy to multiple informationally distinct policy shocks at time 0 in order to enforce the desired counterfactual policy rule both ex ante in private sector expectations and ex post along the equilibrium path. Informationally distinct policy shocks mean that each shock should reflect a rather different policy path. This requirement is easily met given the progress made in the recent literature in identifying shocks with quite different effects on future policy paths and is addressed in the next subsection. The basic idea of this methodology is that we can mimic the effects of changes in a counterfactual policy rule by combining the policy paths implied by these informationally distinct shocks. Since this approach does not involve ex post surprises, it is robust to the Lucas critique.

5.1.1 McKay and Wolf's Methodology

I will now present a summary of MKW's methodology. Consider a linear, perfect foresight, infinite horizon economy. Boldface denotes time paths for $t = 0, 1, 2, \ldots,$ and all variables are expressed in deviations from the deterministic steady state. The model includes unobserved variables collected in an n_w-dimensional vector $\mathbf{w}$, observed non-policy variables in an n_x-dimensional vector $\mathbf{x}$, and policy instruments in an n_z-dimensional vector $\mathbf{z}$. The model also includes a n_ε-dimensional vector of structural shocks $\boldsymbol{\varepsilon}$ and an n_z-dimensional vector $\boldsymbol{\nu}$ of policy shocks.[23] The economy is summarized by the following two-block system:

$$\mathcal{H}_w \mathbf{w} + \mathcal{H}_x \mathbf{x} + \mathcal{H}_z \mathbf{z} + \mathcal{H}_\varepsilon \boldsymbol{\varepsilon} = \mathbf{0} \tag{10}$$

$$\mathcal{A}_x \mathbf{x} + \mathcal{A}_z \mathbf{z} + \boldsymbol{\nu} = \mathbf{0} \tag{11}$$

where the linear maps $\{\mathcal{H}_w, \mathcal{H}_x, \mathcal{H}_z, \mathcal{H}_\varepsilon\}$ represent the non-policy block, which does not depend on the coefficients of the policy rule $\{\mathcal{A}_x, \mathcal{A}_z\}$. A central assumption of this methodology is the separation between the private sector (non-policy) and policy blocks, which is a feature in common with the linear representation obtained within the workhorses of modern macroeconomic models. Assuming that the policy rule in (11) induces a unique equilibrium, we can write the solution to the system (10)-(11) as

[23] The vector $\boldsymbol{\nu}$ collects the full menu of contemporaneous and news shocks to the prevailing policy rule at all horizons. News shocks should be interpreted as the entries of this vector for $t > 0$.

$$
\begin{pmatrix} \mathbf{w} \\ \mathbf{x} \\ \mathbf{z} \end{pmatrix} = \underbrace{\begin{pmatrix} \boldsymbol{\Theta}_{w,\varepsilon,\mathcal{A}} & \boldsymbol{\Theta}_{w,\nu,\mathcal{A}} \\ \boldsymbol{\Theta}_{x,\varepsilon,\mathcal{A}} & \boldsymbol{\Theta}_{x,\nu,\mathcal{A}} \\ \boldsymbol{\Theta}_{z,\varepsilon,\mathcal{A}} & \boldsymbol{\Theta}_{z,\nu,\mathcal{A}} \end{pmatrix}}_{\equiv \boldsymbol{\Theta}_{\mathcal{A}}} \times \begin{pmatrix} \boldsymbol{\varepsilon} \\ \boldsymbol{\nu} \end{pmatrix} \tag{12}
$$

where $\boldsymbol{\Theta}_{\mathcal{A}}$ contains the impulse responses of the endogenous variables $\mathbf{w}$ and $\mathbf{x}$ and the policy instruments $\mathbf{z}$ to the structural $\boldsymbol{\varepsilon}$ and policy $\boldsymbol{\nu}$ shocks under the prevailing rule. Thus, the impulse responses of the macroeconomic observables $\mathbf{x}$ and policy instruments $\mathbf{z}$ to the structural shock $\boldsymbol{\varepsilon}$, when the policy rule is followed perfectly ($\boldsymbol{\nu} = 0$) are, $\mathbf{x}_{\mathcal{A}}(\boldsymbol{\varepsilon})$ and $\mathbf{z}_{\mathcal{A}}(\boldsymbol{\varepsilon})$, respectively.

Now suppose we want to study the effect of a structural shock $\boldsymbol{\varepsilon}$ under a counterfactual policy rule of the form

$$
\tilde{\mathcal{A}}_x \mathbf{x} + \tilde{\mathcal{A}}_z \mathbf{z} = \mathbf{0} \tag{13}
$$

To construct policy counterfactuals, MKW's identification results require only knowledge of $\{\boldsymbol{\Theta}_{x,\nu,\mathcal{A}}, \boldsymbol{\Theta}_{z,\nu,\mathcal{A}}\}$.[24] In other words, an econometrician interested in constructing counterfactual paths needs the impulse response functions of policy instruments $\mathbf{z}$ and macroeconomic observables $\mathbf{x}$ to contemporaneous as well as all possible future policy shocks $\boldsymbol{\nu}$ to the prevailing rule. Under the outlined assumption, the policy counterfactuals $\mathbf{x}_{\tilde{\mathcal{A}}}(\boldsymbol{\varepsilon})$ and $\mathbf{z}_{\tilde{\mathcal{A}}}(\boldsymbol{\varepsilon})$ for a counterfactual rule $\{\mathcal{A}_x, \mathcal{A}_z\}$ can be recovered as follows:

$$
\begin{aligned}
\mathbf{x}_{\tilde{\mathcal{A}}}(\boldsymbol{\varepsilon}) &= \mathbf{x}_{\mathcal{A}}(\varepsilon) + \boldsymbol{\Theta}_{x,\nu,\mathcal{A}} \times \tilde{\nu} & (14) \\
\mathbf{z}_{\tilde{\mathcal{A}}}(\boldsymbol{\varepsilon}) &= \mathbf{z}_{\mathcal{A}}(\varepsilon) + \boldsymbol{\Theta}_{z,\nu,\mathcal{A}} \times \tilde{\nu} & (15)
\end{aligned}
$$

where $\tilde{\nu}$ uniquely solve

$$
\tilde{\mathcal{A}}_x[\mathbf{x}_{\mathcal{A}}(\boldsymbol{\varepsilon}) + \boldsymbol{\Theta}_{x,\nu,\mathcal{A}} \times \tilde{\nu}] + \tilde{\mathcal{A}}_z[\mathbf{z}_{\mathcal{A}}(\boldsymbol{\varepsilon}) + \boldsymbol{\Theta}_{z,\nu,\mathcal{A}} \times \tilde{\nu}] = \mathbf{0}
$$

The basic idea of this methodology is that we can mimic the effects of changes in a counterfactual policy rule by combining the policy paths implied by informationally distinct policy shocks. However, this is not an empirically feasible approach because we would need to know every element of $\boldsymbol{\Theta}_{x,\nu,\mathcal{A}}$ and $\boldsymbol{\Theta}_{z,\nu,\mathcal{A}}$, i.e., the impulse responses

[24] The underlying assumption is that the counterfactual policy rule induces a unique equilibrium.

of $\mathbf{x}$ and $\mathbf{z}$ to each possible shock path. In fact, in the real world we have only available a limited number of informationally distinct policy shocks. Suppose, we have access to estimates of n_s distinct policy shocks associated with n_s alternative response paths of the policy instruments z. In this setup, MKW suggest finding the weights s on the n_s date-0 shocks to obtain the desired counterfactual rule. Formally, we need to solve the following optimization problem:

$$\min_{s} \|\tilde{\mathcal{A}}_x\big(\mathbf{x}_A(\boldsymbol{\varepsilon}) + \boldsymbol{\Omega}_{x,A} \times \mathbf{s}\big) + \tilde{\mathcal{A}}_z\big(\mathbf{z}_A(\boldsymbol{\varepsilon}) + \boldsymbol{\Omega}_{z,A} \times \mathbf{s}\big)\| \tag{16}$$

where the impulse responses of $\mathbf{x}$ and $\mathbf{z}$ to a shock path are stored in each column of $\boldsymbol{\Omega}_{x,\mathcal{A}}$ and $\boldsymbol{\Omega}_{z,\mathcal{A}}$. The optimal weights s^* obtained in (16) determine the best approximation to the desired policy counterfactual within the space of empirically identified policy shock paths.

5.1.2 Identification of spending shocks

MKW stated that their *"empirical method can be applied in the empirically relevant case of a researcher with access to only a couple of distinct shocks"* (p. 1696). Therefore, I will now present the identification strategy for spending shocks adopted in this paper.

The macroeconomic literature on fiscal policy has approached the identification of spending shocks in a variety of ways, ranging from structural identification to the use of narrative records or professional forecasts. Although all of these strategies target shocks to government spending, the resulting shock series reflect different types of fiscal news and imply very distinct policy paths. This is particularly appealing given the methodological requirements outlined in the previous subsection. I estimate counterfactual impulse responses by subjecting the model in (4), with the same baseline endogenous variables, to two distinct types of spending shocks at date 0 simultaneously. The first is the shock identified by Caldara and Kamps (2017) using the Blanchard-Perotti approach under general rules that, given that this strategy relies on contemporaneous innovations in spending, can be viewed as a shock to the shorter end of government spending.[25] The second is a *news* or *foresight* shock identified using the Survey of Professional Forecasters (hereafter SPF) following the approach in Forni and Gambetti

[25]The estimation uncertainty in this shock is directly accounted for by having estimated the full posterior distribution.

(2016), that reflects a shock to the longer end of government spending. This variable is defined as the difference between the forecast of government spending growth at time t for the following quarters, as reported by the SPF, and the forecast for the same quarters at time $t-1$. The reason I choose this variable instead of the well-known defense news series of Ramey (2011) is dictated by the use of a post-Korean sample, which implies a large loss of information in the Ramey variable.[26] The longest common sample available for both shocks is 1981q4-2019q4, driven by the start date of the SPF, which is almost entirely coincident with the baseline sample.

5.1.3 Results

The desired counterfactual is obtained according to equations (14)-(15) by focusing on the nonlinear responses of the quarterly monetary model introduced in Section 5. Intuitively, the goal of this strategy is to switch off the monetary-fiscal interaction in the high GPR regime and then compare the monetary effects on real GDP, prices and the credit spread in the two regimes.[27] From equations (14)-(15), $\mathbf{x}_{\mathcal{A}}(\varepsilon)$ and $\mathbf{z}_{\mathcal{A}}(\varepsilon)$ correspond to the nonlinear responses of the macroeconomic observables $\mathbf{x}$ and policy instruments $\mathbf{z}$ included in $\beta(L)$ of equation (2). To be internally consistent, I estimate $\mathbf{\Theta}_{x,\nu,\mathcal{A}}$ and $\mathbf{\Theta}_{z,\nu,\mathcal{A}}$ using the same VARX from the baseline analysis with the two spending shocks, identified according to the previous subsection, as exogenous components. The final ingredient in estimating policy counterfactuals are the weights $\tilde{\nu}$. These are obtained by solving (16) and correspond to the optimal weights s^*. In doing so, I minimize the distance of the response of defense spending in the high GPR regime from that in normal times. All other endogenous variables are left unconstrained. Overall, this counterfactual scenario allows me to quantitatively assess the extent to which the monetary-fiscal interaction can account for the different responses to a monetary shock in the two regimes.

Figure 10 reports the responses and confidence bands shown in Figure 9 (first two rows) and adds those obtained from the counterfactual analysis in the high GPR regime as black dotted lines. For ease of comparison, the counterfactual responses are also shown in the first row, along with the baseline responses of the high GPR regime.

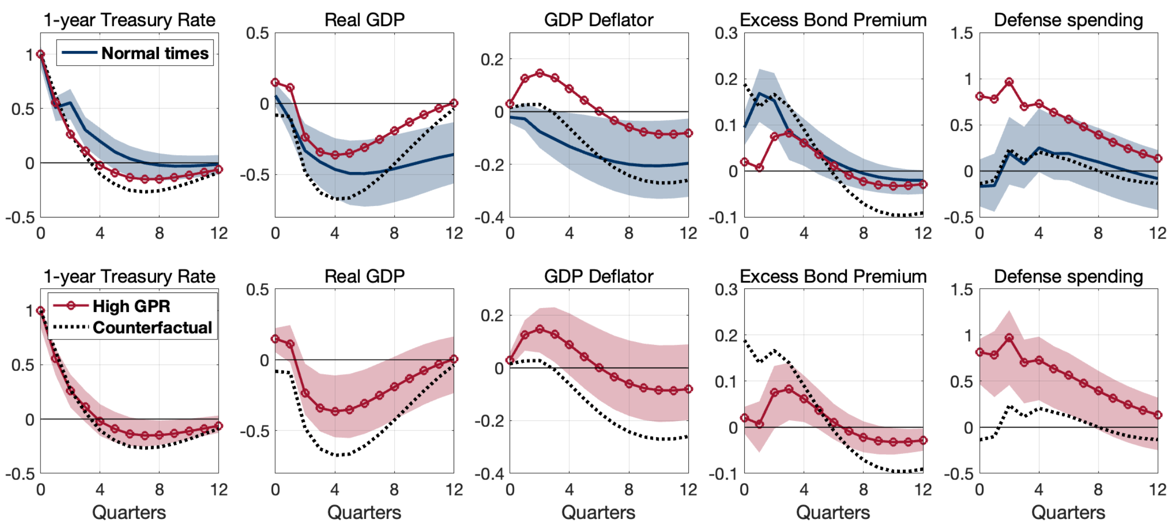
[26] In addition, Forni and Gambetti (2016) find that *"Ramey's* [military spending] *variable seems unsuited to correctly capture the foresight shock"* (p. 14).

[27] Remember that the responses in normal times correspond to only the linear term $\alpha(L)$ in (2). This is because the GPR index d_t is normalized around the median.

The distance in the defense spending response between normal times (blue solid line) and the high GPR counterfactual response (black dotted line) can be regarded as a measure of the accuracy of the counterfactual rule enforced using MKW's method. The minor discrepancy between the two responses indicates that the counterfactual rule is enforced with a significantly high degree of accuracy.[28]

Interestingly, the effects of a contractionary monetary shock in the high GPR regime are similar to those in normal times once the monetary-fiscal interaction mechanism is switched off. The "puzzling" expansionary short-run real effects are now absent, and real activity reacts negatively with a hump-shaped response. Prices do not rise after the shock, but exhibit a lagged decline that closely follows the response in normal times. The same is true for the credit spread, as the excess bond premium shows a larger response, closer to that in normal times. This is true at least for the first six quarters. Finally, the interest rate response in the high GPR regime is not significantly different from that under the prevailing rule, suggesting that the results are not driven

Figure 10: Monetary policy and geopolitical risk: Policy counterfactual



Notes: Seven-variable quarterly VAR. Nonlinear impulse responses from a monetary policy shock in normal times (50th percentile of the U.S. GPR index, blue solid lines) and high GPR (96th percentile of the U.S. GPR index, red lines with circles) under the prevailing baseline rule and the counterfactual for high GPR (black dotted lines). The shock is normalized to induce a 100 basis point increase in the 1-year treasury rate. The solid lines are the point estimates and the shaded areas are 68% confidence bands.

[28]Note that this does not imply that the counterfactual rule is necessarily economically meaningful, only that I am able to mimic the effects of a switch to a counterfactual policy rule.

by a more aggressive monetary contraction. Overall, the main takeaway from this policy counterfactual is that the monetary-fiscal interaction mechanism is *quantitatively* important in explaining the GPR-dependent effects of monetary policy shocks.

6 Disentangling the Geopolitical Risk index

As shown in Section 2, geopolitical episodes are associated with two distinct macroeconomic developments. The first is linked to news of future supply disruptions in the oil market, and is therefore associated with inflationary pressures. The second, by contrast, with a lower aggregate demand component and is therefore associated with deflationary pressures. Given these distinct macroeconomic developments linked with different GPR, it becomes crucial to isolate the relevant geopolitical episodes for the contractionary monetary policy shock considered.

In this section, I disentangle episodes in the GPR index associated with the first mechanism from those associated with the second. The rationale is to remove episodes from the GPR index that are associated with lower aggregate demand pressures, since during these events, most notably 9/11, monetary contractions are not present according to historical records and the identified shock series. This approach improves the analysis of how GPR affects the effectiveness of a monetary tightening shock by providing a more appropriate state variable.

Caldara and Iacoviello (2022) calculate the geopolitical risk index by counting the number of articles related to adverse geopolitical events in each selected newspaper for each month and take the percentage over the total number of news articles.[29] Then, they characterize geopolitical events based on the following subcategories: War Threats, Peace Threats, Military Buildups, Nuclear Threats, Terror Threats, Beginning of War, Escalation of War, Terror Acts.

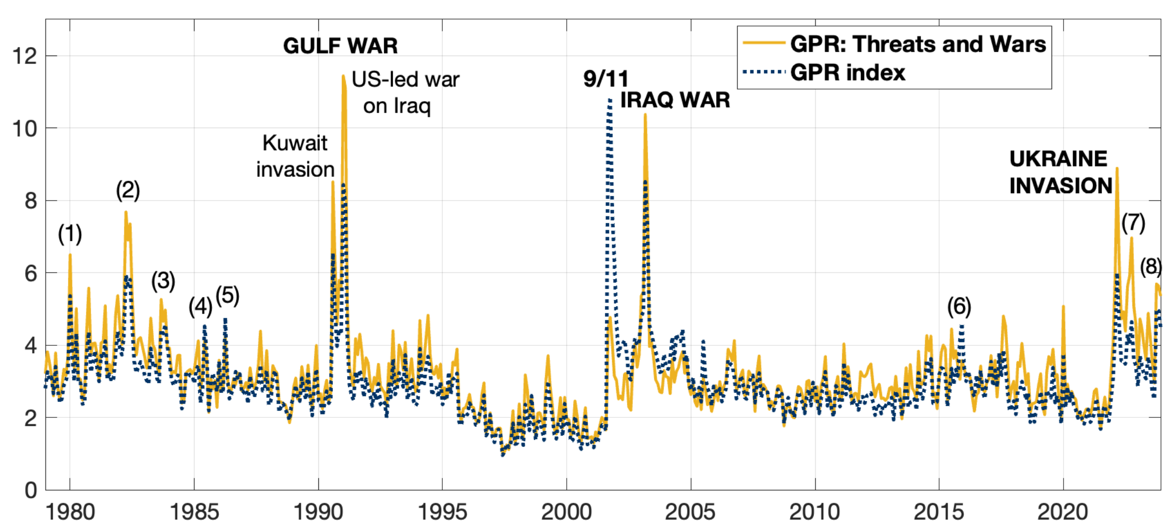
I noticed that geopolitical episodes in the subcategory *Terror Acts* are a clear example of events that increase the real-option value of waiting for information (e.g., Bernanke, 1983; Bloom, 2009). A notable example is 9/11, which led to a slowdown in consumption and investment through its effect on consumer and business confidence and a temporary decline in stock prices. Such geopolitical events require an accom-

[29]The selected newspapers are: Chicago Tribune, the Daily Telegraph, Financial Times, The Globe and Mail, The Guardian, the Los Angeles Times, The New York Times, USA Today, The Wall Street Journal, and The Washington Post.

modative stance by the Federal Reserve. On the other hand, geopolitical tensions associated with news of future oil supply disruptions, such as the Gulf War or the Iraq War (see Figure 2), force the Federal Reserve to act in defense of its price stability mandate by adopting a contractionary stance. On the fiscal side, as shown in Section 2, both situations imply expansionary fiscal (military) interventions. Therefore, when studying the GPR-dependent effects of a monetary tightening shock it is very important to exclude the subcomponent *Terror Acts*, which highlights geopolitical episodes associated with deflationary forces, and thus inconsistent in the data with a tightening stance by the Federal Reserve.

Figure 11 plots the baseline GPR index (blue dotted line) along with the alternative index, the GPR index net of the *Terror Acts* component, which I call *Threats and Wars* (yellow solid line). While it may be tempting to proceed directly to using this index as a state variable, it should be noted that this decomposition is not available for the U.S. specifically, but only for the global index. Nevertheless, the authors recognize

Figure 11: Disentangling the geopolitical risk index: Threats and Wars

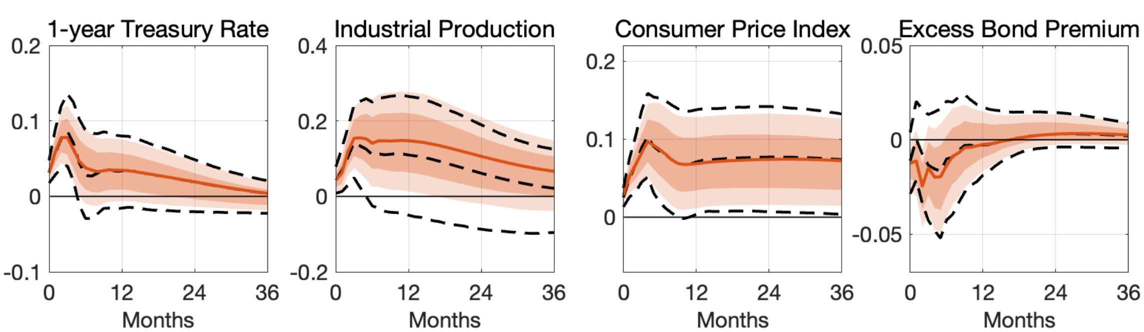


Notes: The dotted blue line is the monthly U.S. historical geopolitical risk index of Caldara and Iacoviello (2022) and the solid yellow line is the *Threats and Wars* index. The graph reports additional seven geopolitical events. (1) January 1980: US-Soviet tensions as the Soviet Union invaded Afghanistan in December 1979; (2) April 1982: Falkland war; (3) September 1983: Nuclear war scare; (4) June 1985: Hijacking of TWA Flight 847; (5) April 1986: US bombing of Libya; (6) November 2015: Paris terrorist attack; (7) October 2022: Crimea bridge attack and Russia strikes on Ukraine's energy infrastructure; (8) October 2023: Hamas Terrorist attack and Israeli invasion of Gaza.

that geopolitical events of global interest often imply U.S. involvement. To ensure that the baseline results hold when the global GPR index is used instead of the U.S.-specific index, I re-estimate the model in (4) using the global GPR index. Figure C.10 in Appendix C plots the impulse responses when the global GPR index is used as the state variable and confirms the findings documented in the baseline analysis, thus validating the use of the global subcomponents of the GPR index as state variables for the U.S. economy.

Figure 12 presents a formal test for nonlinearity when choosing the GPR *Threats and Wars* index as the state variable. Specifically, it plots the impulse responses to the nonlinear term, i.e., $\beta(L)$, and the confidence bands from equation (2). The figure also includes the point estimates and the 90% confidence bands for the test estimated using the baseline model, reported in Figure 5, as dashed black lines. The test shows that the role of the GPR *Threats and Wars* index is statistically significant and contributes positively to industrial production for the first 16 months. On the other hand, it is positive and statistically significant only for the first 4 months when the baseline GPR index is used. Regarding prices, the difference with the baseline model is the smaller estimation uncertainty. Another important difference is the role of the nonlinear term for the excess bond premium. The interaction is now statistically significant and contributes negatively to the dynamic response of this credit spread following the shock. By comparison, the test in the baseline analysis does not indicate the presence of GPR-dependent effects on the excess bond premium.

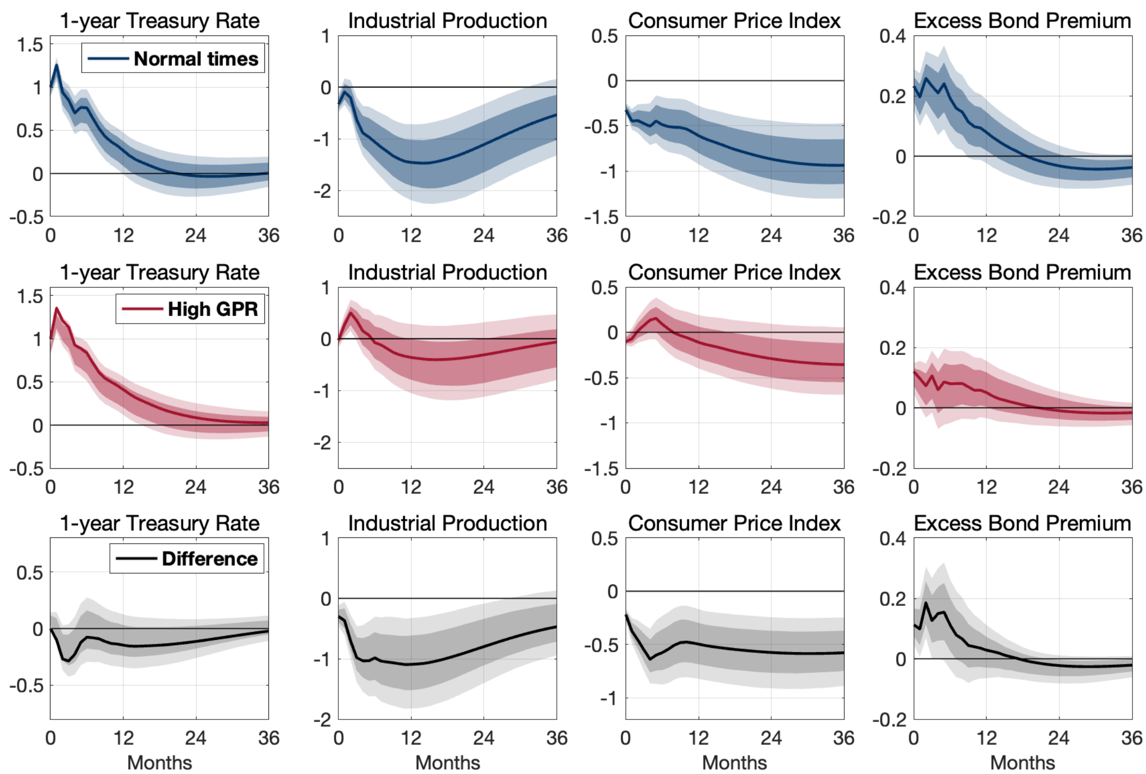
Figure 12: Monetary policy and GPR Threats and Wars: Test of nonlinearity



Notes: Six-variable VAR. Impulse responses to the nonlinear term in equation (2) estimated using a nonlinear Proxy-SVAR and the GPR *Threats and Wars* index. The solid lines are the point estimates and the shaded areas are 68% and 90% confidence bands. Dashed black lines are the point estimates and the 90% confidence bands for the baseline nonlinearity test with the GPR index as state variable.

Figure 13 plots the impulse responses following a tightening shock with the GPR *Threats and Wars* index as the state variable. As indicated by the nonlinearity test, there are now larger differences in the responses of industrial production and the excess bond premium compared to the baseline model. In view of the GPR-dependent evidence documented in this paper, the findings suggest that the GPR *Threats and Wars* index provides a better state-dependent variable for a monetary tightening shock. As discussed in Section 5, this likely reflects the more unfavorable policy mix associated with these events.[30]

Figure 13: Monetary policy and GPR Threats and Wars



Notes: Six-variable VAR. Nonlinear impulse responses from a monetary policy shock in normal times (50th percentile of the U.S. GPR index *Threats and Wars*, blue lines), high GPR (96th percentile of the U.S. GPR index *Threats and Wars*, red lines) and the difference between the two impulse responses (black lines). The shock is normalized to induce a 100 basis point increase in the 1-year treasury rate. The solid lines are the point estimates and the shaded areas are 68% and 90% confidence bands.

[30]One might be concerned that, as shown in Figure 2, the second phase of the Gulf War, i.e. the U.S.-led war against Iraq, is mostly associated with news of stronger future oil supply production. To address this concern, I reduce the weight of this event. The results turn out to be largely unaffected. If anything, the test for industrial production is not statistically significant for the first 19 months.

6.1 Policy experiments

I now conduct some policy experiments by studying the effects of monetary policy shocks under different levels of geopolitical risk. The main advantage of these experiments is that they provide a historical perspective on the findings that are documented in the paper. This is important because it can inform monetary policymakers about the effectiveness of their actions under a given *nature* and *level* of geopolitical risk. The model used is the core six-variable monthly VAR augmented with real federal defense spending.[31]

6.1.1 Monetary tightening: GPR *Threats and Wars* index

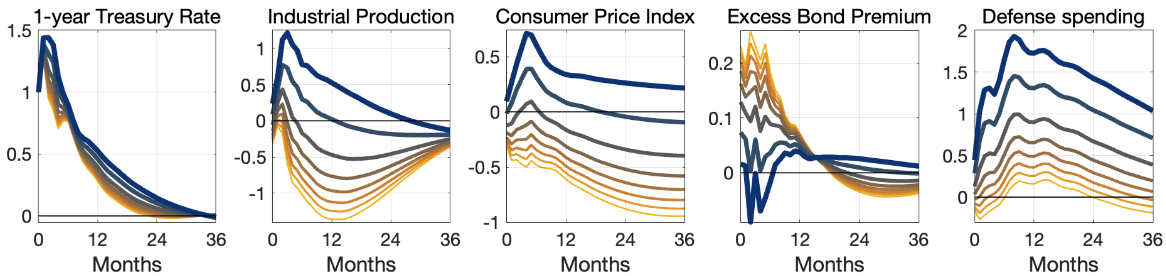
Figure 14 reports the main result of this exercise, and presents evidence on how the effectiveness of monetary tightening varies as I select different levels of GPR risk. The state variable adopted is the GPR *Threats and Wars* index, as I have shown in the previous section to be a more relevant state variable for a tightening shock than the aggregate GPR index. The point estimates are reported for a range of scenarios in which I select the level of geopolitical risk at the 50th (yellow thinnest line), 60th, 70th, 80th, 90th, 95th, 98th, and 99th percentiles (blue thickest line). Thicker lines correspond to higher percentiles of the index. To put things in perspective, events such as the Kuwait invasion, the Gulf War, and the Iraq War peaked at 8.5, 11.4, and 10.3 on the GPR *Threats and Wars* index, respectively, and the 98th and 99th percentiles correspond to 6.4 and 8.3, respectively. These values are lower than the peaks of these major geopolitical events, which are above the 99th percentile, but in the real world central banks do not always react in the same month of the geopolitical episode. This motivates the use of smaller values of the index in the paper. Choosing larger values of the index would only strengthen the following arguments.

The results of this experiment suggest that a monetary tightening shock becomes less effective as geopolitical tensions rise. When the *Threats and Wars* index is at the 98th or 99th percentile, a tightening shock has clear expansionary effects. As documented in Section 5, these results must be viewed through the lens of the monetary-fiscal interaction. Real defense spending increases with higher geopolitical tensions,

[31] As shown earlier, this model is consistent with its quarterly counterpart. The reason for choosing the higher frequency model is to obtain more precise estimates, which is critical for the type of exercise developed here.

peaking at almost a 2% increase in the growth rate within the first year following the shock when the *Threats and Wars* index is at the 99th percentile. This adverse policy mix can explain the temporary expansionary effects of monetary tightening when geopolitical risk is significantly high.

Figure 14: Monetary tightening and GPR Threats and Wars



Notes: Seven-variable VAR. Nonlinear impulse responses (point estimates) from a monetary tightening shock at the 50th (thinnest yellow line), 60th, 70th, 80th, 90th, 95th, 98th and 99th percentiles (thickest blue line) of the *Threats and Wars* index. The shock is normalized to induce a 100 basis point increase in the 1-year treasury rate.

6.1.2 Monetary easing: GPR *Terror Acts* index

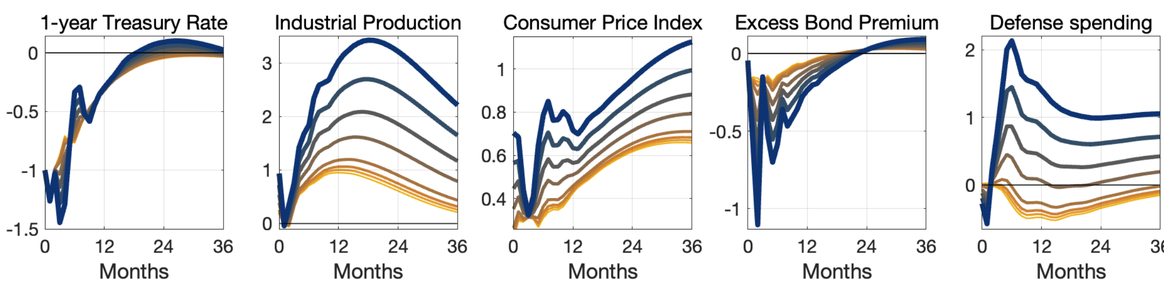
In section 6, I have isolated the components of the GPR index related to news of future supply disruptions, i.e., *Threats* and *Wars*. However, the residual index, which includes the *Terror Acts* component provides valuable information for geopolitical episodes associated with lower aggregate demand pressure. During geopolitical episodes of this nature, the central bank will be induced to adopt an expansionary stance. Following the monetary-fiscal interaction mechanism presented in Section 5, such geopolitical episodes would lead to a favorable policy mix, resulting in larger effects of an expansionary monetary policy shock.

To test this hypothesis, I conduct a second policy experiment along the lines of the previous one, but now using the GPR *Terror Acts* index and focusing on a monetary easing shock.[32] Figure 15 reports the main result of this exercise and, as in the previous experiment, presents evidence on how the effectiveness of a monetary easing shock varies across different levels of the state variable. The point estimates are reported for

[32]By simply changing the state variable to the *Terror Acts* index for monetary easing, the model naturally generates the appropriate impulse responses without the need to explicitly model sign asymmetries. This approach allows the model to accurately capture the different effects of monetary policy shocks depending on the prevailing geopolitical episode.

a range of scenarios in which I select the level of geopolitical risk at the 50th (yellow thinnest line), 60th, 70th, 80th, 90th, 95th, 98th, and 99th percentiles (blue thickest line). Thicker lines correspond to higher values of the index. Figure C.12 in Appendix plots the baseline GPR index together with the GPR *Terror Acts* index. The results of this experiment confirm the importance of the monetary-fiscal interaction mechanism also in the context of these alternative geopolitical events and monetary action. In this case, however, the effects of monetary easing are amplified by the coordinated fiscal response. Indeed, real defense spending increases with higher geopolitical tensions, peaking at an increase in the growth rate of about 2% within the first year following the shock when the *Terror Acts* index is at the 99th percentile. This amplification is increasing in the level of the *Terror Acts* index and is transmitted to both real, nominal and financial variables.

Figure 15: Monetary easing and GPR Terror Acts



Notes: Seven-variable VAR. Nonlinear impulse responses (point estimates) from a monetary easing shock at the 50th (thinnest yellow line), 60th, 70th, 80th, 90th, 95th, 98th and 99th percentiles (thickest blue line) of the *Terror Acts* index. The shock is normalized to induce a 100 basis point increase in the 1-year treasury rate.

7 Concluding Remarks

Geopolitical tensions are a major risk to stability and central banks are closely monitoring the risks they pose to the domestic and global economic outlook. This is important as geopolitical events can be of different nature. A first type is related to supply disruptions in the energy market. A second type relates to lower aggregate demand pressures. The two types of events imply opposite central bank responses, while both trigger expansionary fiscal interventions in the form of military buildups. In this paper, I document that a monetary tightening shock is less powerful during periods of geopolitical tensions related to supply disruptions in the energy market than in normal times for real, nominal and financial variables. This is because the contractionary stance of the central bank is countered by expansionary fiscal interventions, leading to an unfavorable policy mix. On the contrary, a monetary easing shock during geopolitical tensions related to lower demand pressures is more powerful than in normal times for real, nominal and financial variables because both the central bank and the fiscal authority adopt an expansionary stance, leading to a favorable policy mix. Overall, both the nature and level of geopolitical risk are crucial in predicting the outcomes of monetary tightening and easing shocks.

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Appendix A Data Sources

The data used in the VAR are described below:

- Industrial production: Total Index, Index 2017=100, Monthly, Seasonally Adjusted (FRED id: INDPRO)
- Consumer price index: All Items in U.S. City Average, Index 1982-1984=100, Monthly, Seasonally Adjusted (FRED id: CPIAUCSL)
- 1-year Treasury rate: Market Yield on U.S. Treasury Securities at 1-Year Constant Maturity, Quoted on an Investment Basis, Percent, Monthly, Not Seasonally Adjusted (FRED id: GS1)
- Read GDP: Real Gross Domestic Product per capita, Chained 2017 Dollars, Quarterly, Seasonally Adjusted Annual Rate (FRED id: A939RX0Q048SBEA)
- GDP Deflator: Gross Domestic Product: Implicit Price Deflator, Index 2017=100, Quarterly, Seasonally Adjusted (FRED id: GDPDEF)
- Unemployment rate: Percent, Monthly, Seasonally Adjusted (FRED id: UNRATE)
- Commodity price index: CRB Commodity Price Index (Bloomberg ticker: CRB CMDT Index)
- Excess bond premium: Gilchrist and Zakrajšek (2012)
- Baa-Aaa corporate bond spread: Moody's Seasoned Baa Corporate Bond Yield minus Moody's Seasoned Aaa Corporate Bond Yield (FRED-MD id: BAA, AAA)
- Expected inflation: Median expected price change next 12 months from the Surveys of Consumers, University of Michigan (FRED id: MICH)
- Federal defense spending: Real Government Consumption Expenditures and Gross Investment: Federal: National Defense, Percent Change from Quarter One Year Ago, Quarterly, Seasonally Adjusted (FRED id: A824RO1Q156NBEA)
- The Geopolitical risk index (and subcomponents): Caldara and Iacoviello (2022). Data are downloaded from https://www.matteoiacoviello.com/gpr.htm on January 25, 2024
- WTI crude oil futures hh-month contract, settlement price (Datastream id: NCLC.hh)
- NBER based Recession Indicators for the United States from the Period following the Peak through the Trough, Monthly, Not Seasonally Adjusted (FRED id: USRECD)

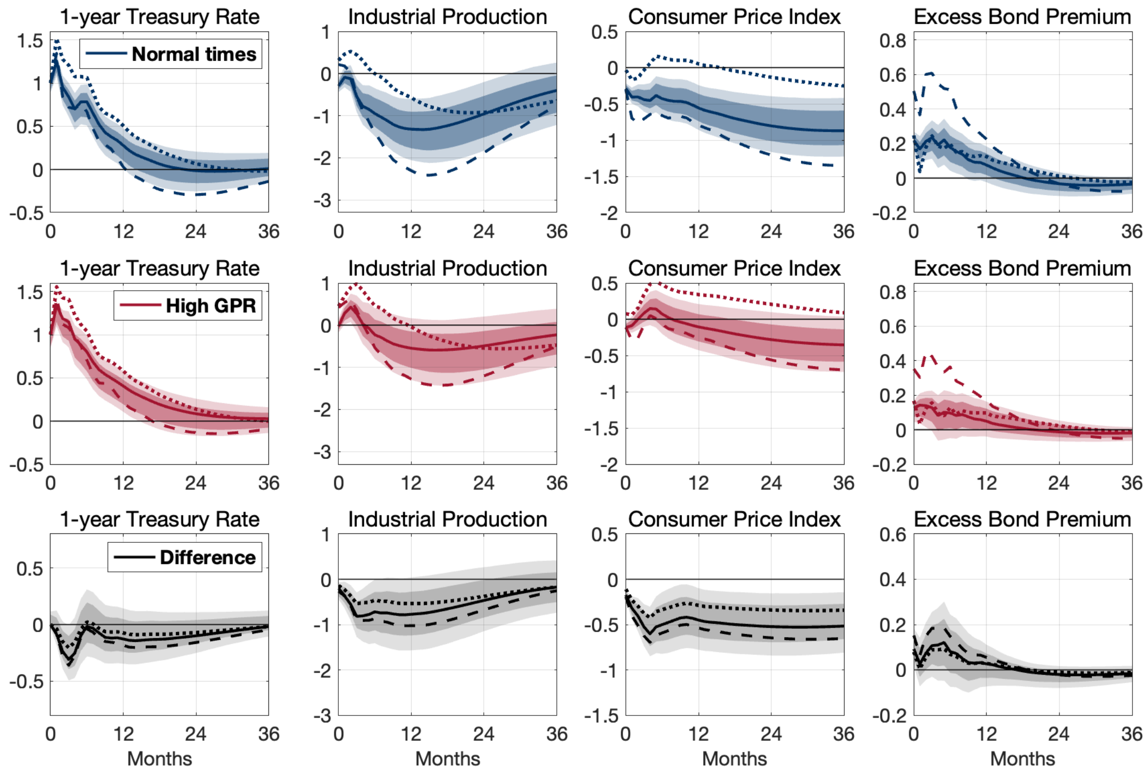
Appendix B Alternative policy instruments

High-frequency instruments for monetary policy shocks are constructed using the financial market reaction around FOMC announcements. Instruments constructed in this way turn out to be predictable by publicly available macroeconomic or financial market information that precedes the FOMC announcement. This is a problem because monetary policy surprises should capture only unanticipated movements in interest rates. To address this problem, Miranda-Agrippino and Ricco (2021) use Greenbook forecasts, and forecast revisions, to purge the high-frequency instrument of the central bank's information set at the time of the announcement.

This approach is not the only one available in the literature, and other authors have reached a similar conclusion using alternative strategies. Jarociński and Karadi (2020) look at the direction of stock market reactions over a 30-minute window following the FOMC announcement to distinguish whether the shock can be classified as a *pure* monetary policy shock. This is defined as an unanticipated increase in interest rates followed by a decline in stock prices. Alternatively, they define it as a central bank information shock, indicating that the central bank reveals private information about current and future demand conditions. An alternative approach comes from Bauer and Swanson (2023a,b), where they question the idea that the Fed has superior information relative to the private sector. They show that there is nothing special about the Greenbook forecasts by obtaining the same results as Miranda-Agrippino and Ricco (2021) using publicly available information, such as the Blue Chip forecasts. Alternatively, they also propose a parsimonious set of predictors that has an intuitive connection to the Fed's monetary policy rule.

To ensure that the choices made by Miranda-Agrippino and Ricco (2021) on the treatment of the instrument do not affect my results, Figure B.1 plots the baseline impulse responses along with those obtained identifying the monetary policy shock using the instruments of Jarociński and Karadi (2020) and Bauer and Swanson (2023a,b). The 68% and 90% confidence bands are reported only for the baseline results. Overall, the alternative identification strategies point to the same state-dependent effects of monetary policy shocks documented in the baseline specification.

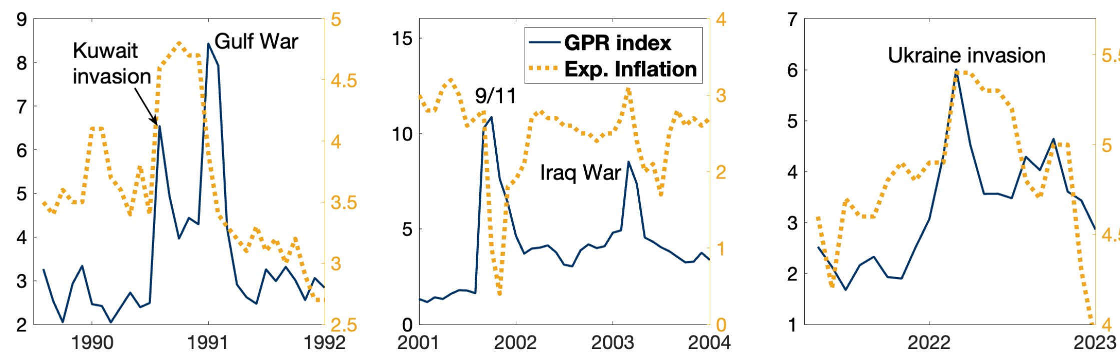
Figure B.1: Alternative shock identifications



Notes: Six-variable VAR. Nonlinear impulse responses from a monetary policy shock in normal times (50th percentile of the geopolitical risk index, blue lines) and high (96th percentile of the geopolitical risk index, red lines) levels of the U.S. geopolitical risk index, and the difference between the two impulse responses (black lines). The monetary policy shock is identified using the baseline instrument of Miranda-Agrippino and Ricco (2021) (solid lines) and two alternative instruments: Jarociński and Karadi (2020) (dotted lines) and Bauer and Swanson (2023b) (dashed lines). The shock is normalized to induce a 100 basis point increase in the 1-year treasury rate. The shaded areas are 68% and 90% confidence bands and are reported only for the baseline model.

Appendix C Additional figures

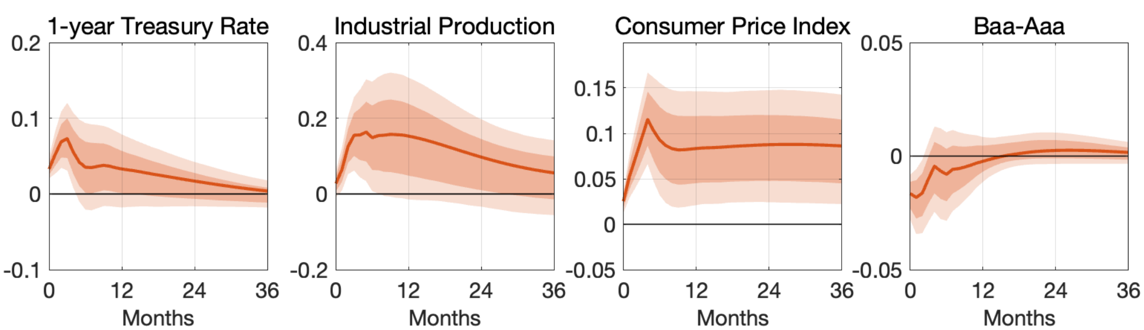
Figure C.1: Geopolitical risk and inflation expectations



(a) *Gulf War* **(b)** *9/11 and Iraq War* **(c)** *Russia invades Ukraine*

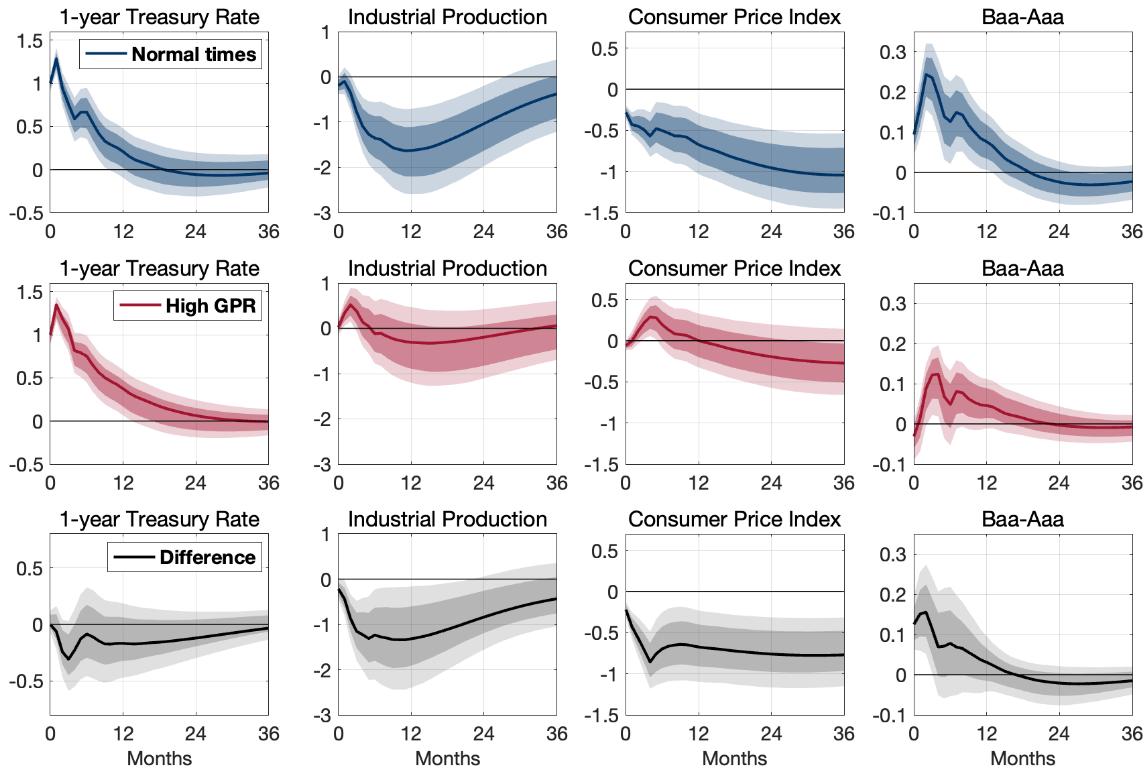
Notes: The solid blue line is the monthly U.S. historical geopolitical risk index of Caldara and Iacoviello (2022) and the yellow dotted line is the inflation forecasts of households from the University of Michigan Survey of Consumers.

Figure C.2: test of nonlinearity - Alternative model



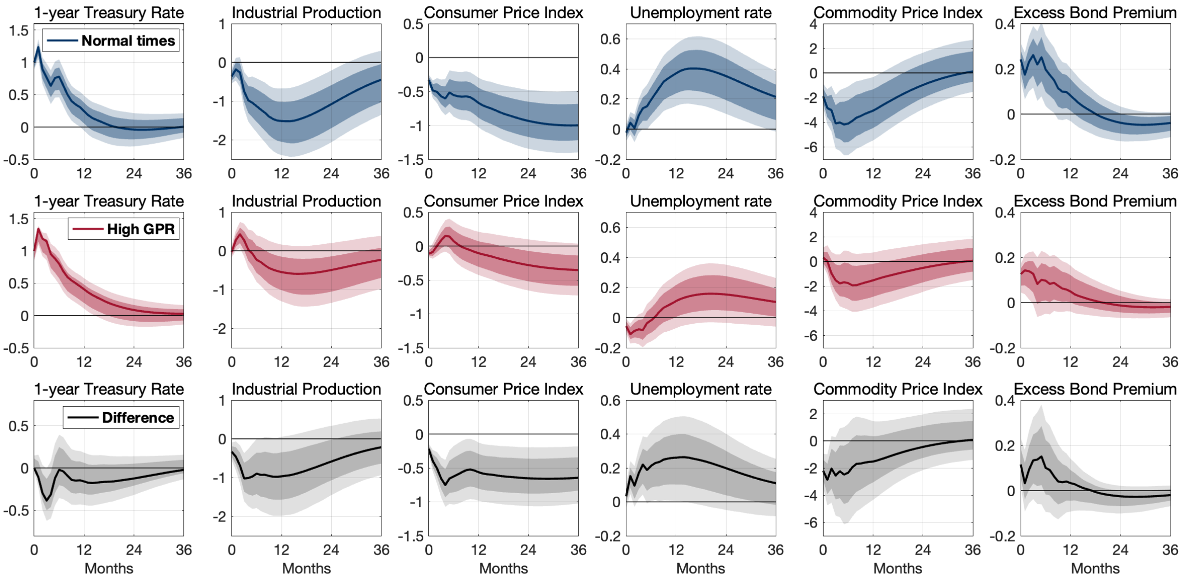
Notes: Six-variable VAR with the Baa-Aaa spread as an alternative financial variable. Impulse responses to the nonlinear term in equation (2) estimated using a nonlinear Proxy-SVAR. The solid lines are the point estimates and the shaded areas are 68% and 90% confidence bands.

Figure C.3: Monetary Policy and geopolitical risk: alternative model



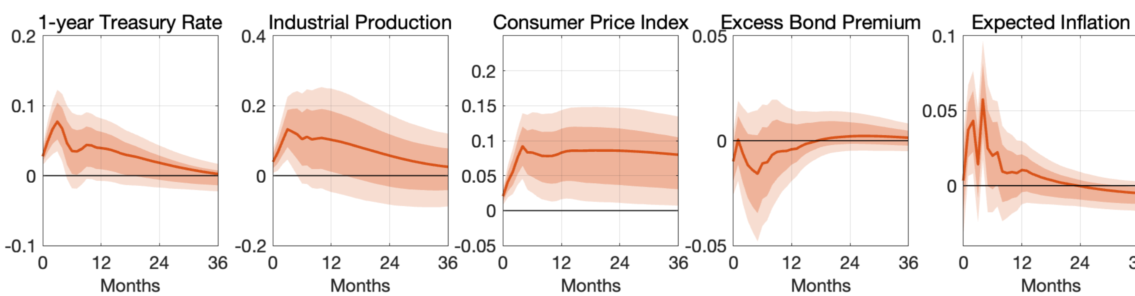
Notes: Six-variable VAR with the Baa-Aaa spread as an alternative financial variable. Nonlinear impulse responses from a monetary policy shock for low (10th percentile of the geopolitical risk index in the sample, blue lines) and high (90th percentile of the geopolitical risk index in the sample, red lines) levels of the U.S. geopolitical risk index, and the difference between the two impulse responses (grey line). The shock is normalized to induce a 100 basis point increase in the 1-year treasury rate. The solid lines are the point estimates and the shaded areas are 68% and 90% confidence bands.

Figure C.4: Monetary Policy and geopolitical risk



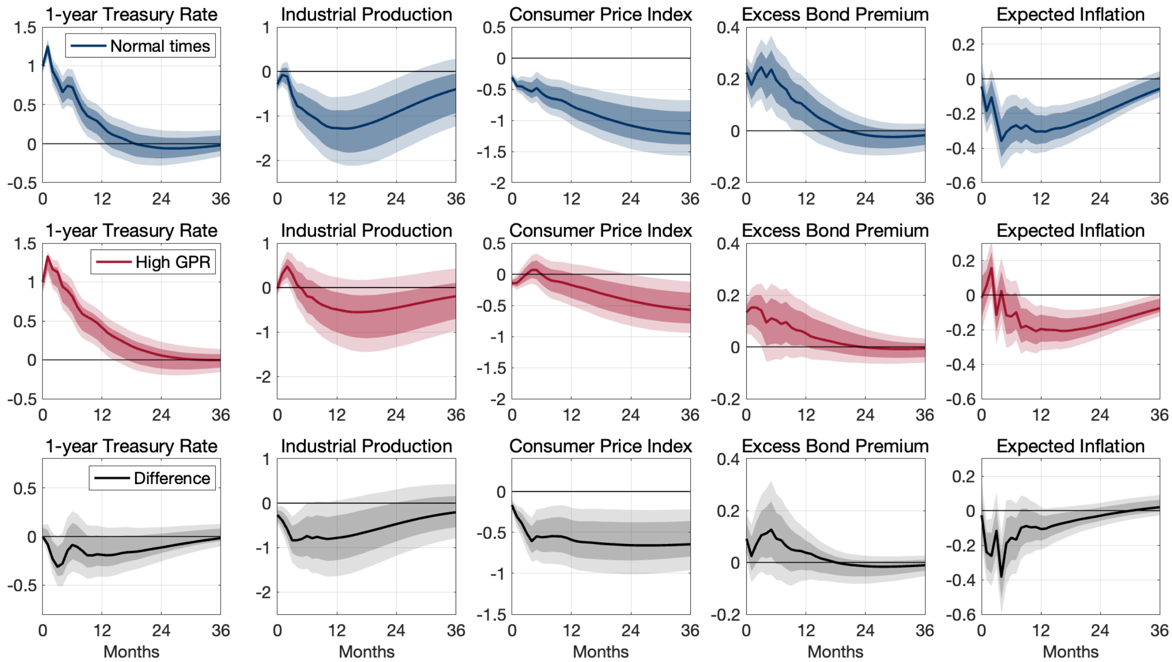
Notes: Six-variable VAR. Nonlinear impulse responses from a monetary policy shock in normal times low (50th percentile of the geopolitical risk index, blue lines) and high (90th percentile of the geopolitical risk index, red lines) levels of the U.S. geopolitical risk index, and the difference between the two impulse responses (black lines). The shock is normalized to induce a 100 basis point increase in the 1-year treasury rate. The solid lines are the point estimates and the shaded areas are 68% and 90% confidence bands.

Figure C.5: test of nonlinearity - Extended model with expected inflation



Notes: Seven-variable VAR, adding expected inflation. Impulse responses to the nonlinear term in equation (2) estimated using a nonlinear Proxy-SVAR. The solid lines are the point estimates and the shaded areas are 68% and 90% confidence bands.

Figure C.6: Monetary policy and geopolitical risk: Inflation expectations

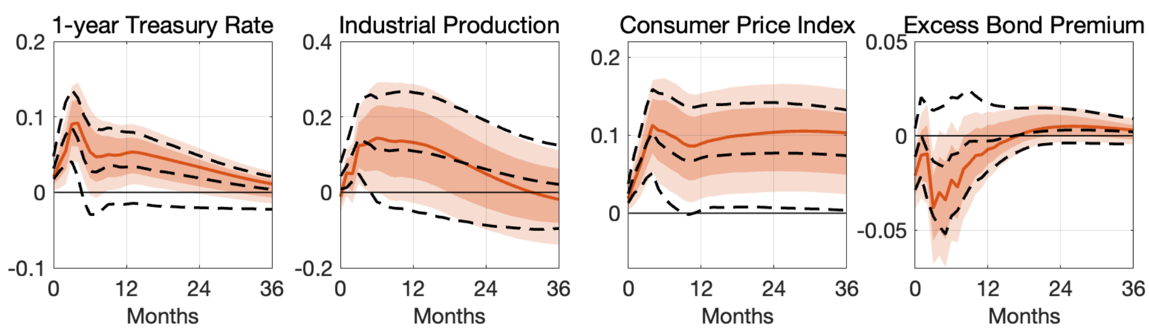


Notes: Seven-variable VAR. Nonlinear impulse responses from a monetary policy shock in normal times (50th percentile of the geopolitical risk index, blue lines), high (96th percentile of the geopolitical risk index, red lines) levels of the U.S. geopolitical risk index and the difference between the two impulse responses (black lines). The shock is normalized to induce a 100 basis point increase in the 1-year treasury rate. The solid lines are the point estimates and the shaded areas are 68% and 90% confidence bands.

Figure C.7: The geopolitical risk index

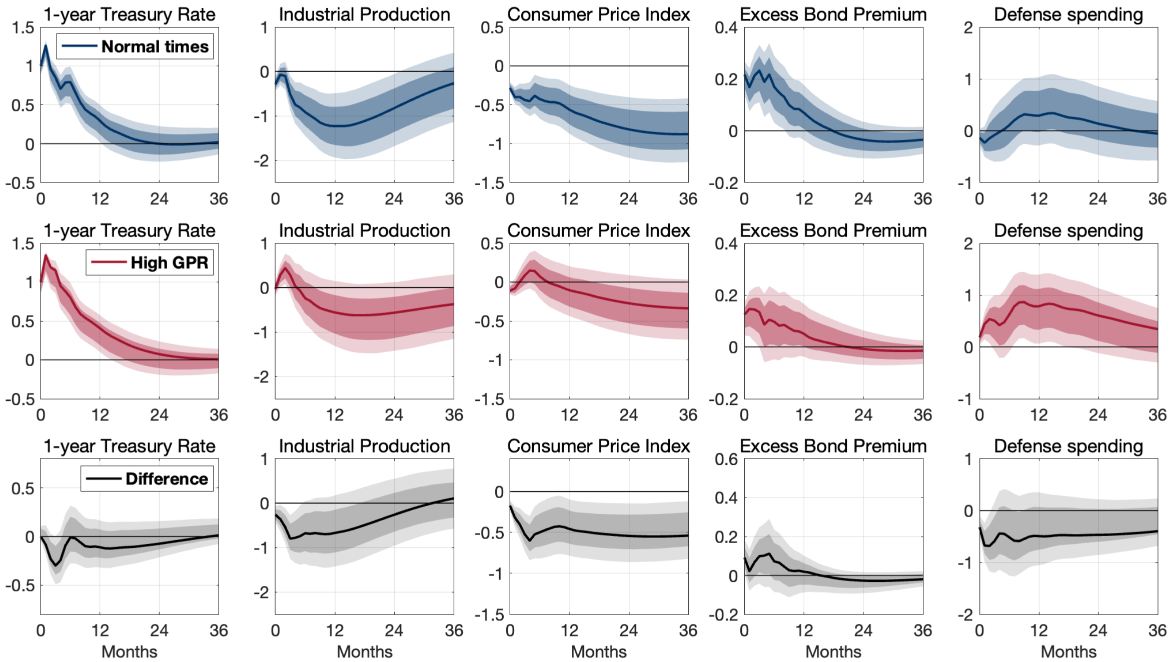
Notes: The solid blue line is the monthly U.S. historical geopolitical risk index of Caldara and Iacoviello (2022). The graph reports additional seven geopolitical events. (1) January 1980: US-Soviet tensions as the Soviet Union invaded Afghanistan in December 1979; (2) April 1982: Falkland war; (3) September 1983: Nuclear war scare; (4) June 1985: Hijacking of TWA Flight 847; (5) April 1986: US bombing of Libya; (6) November 2015: Paris terrorist attack; (7) October 2022: Crimea bridge attack and Russia strikes on Ukraine's energy infrastructure; (8) October 2023: Hamas Terrorist attack and Israeli invasion of Gaza. Shaded bands are NBER recession periods.

Figure C.8: Monetary policy, GPR and the business cycle: Test of nonlinearity



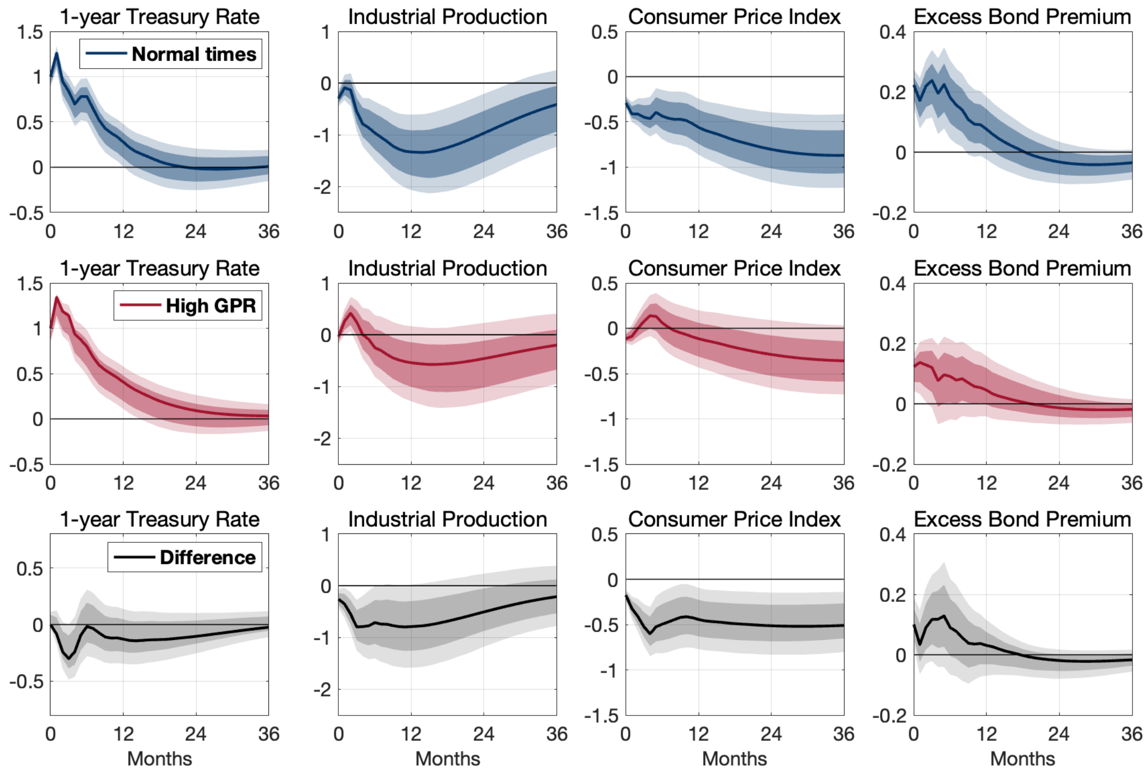
Notes: Six-variable VAR. Impulse responses to the coefficients associated with the nonlinear term $\beta(L)$ in equation (9) estimated using a nonlinear Proxy-SVAR. The solid lines are the point estimates, and the shaded areas are 68% and 90% confidence bands of the extended. The dashed black lines are the point estimates and 90% confidence bands for the baseline nonlinearity test shown in Figure 5.

Figure C.9: Monetary policy and geopolitical risk: Policy-mix



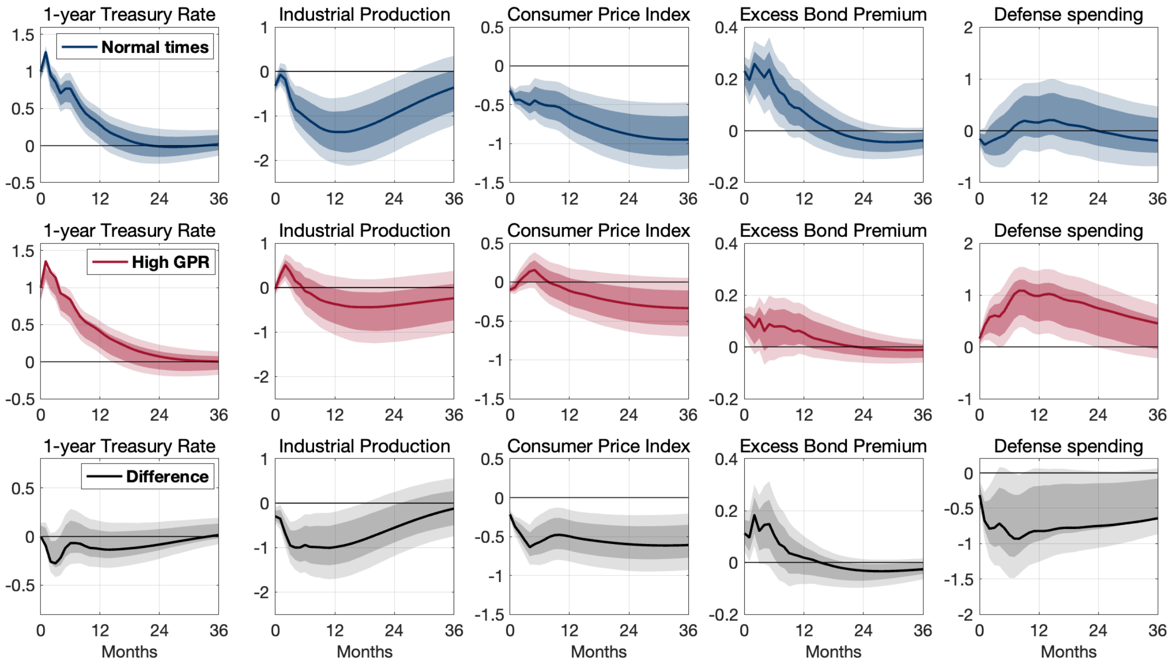
Notes: Seven-variable VAR. Nonlinear impulse responses from a monetary policy shock in normal times (50th percentile of the U.S. GPR index, blue lines), high GPR (96th percentile of the U.S. GPR index, red lines) and the difference between the two impulse responses (black lines). The shock is normalized to induce a 100 basis point increase in the 1-year treasury rate. The solid lines are the point estimates and the shaded areas are 68% and 90% confidence bands.

Figure C.10: Monetary Policy and global geopolitical risk



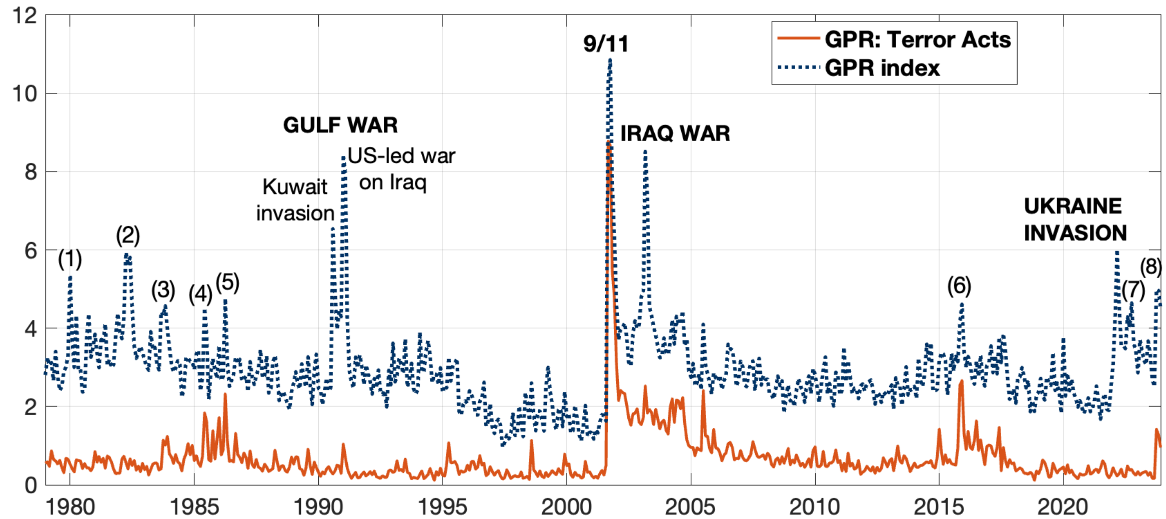
Notes: Six-variable VAR. Nonlinear impulse responses from a monetary policy shock in normal times (50th percentile of the global geopolitical risk index, blue lines) and high (96th percentile of the global geopolitical risk index, red lines) levels of the global geopolitical risk, and the difference between the two impulse responses (black lines). The shock is normalized to induce a 100 basis point increase in the 1-year treasury rate. The solid lines are the point estimates and the shaded areas are 68% and 90% confidence bands.

Figure C.11: Threats and Wars: Policy-mix



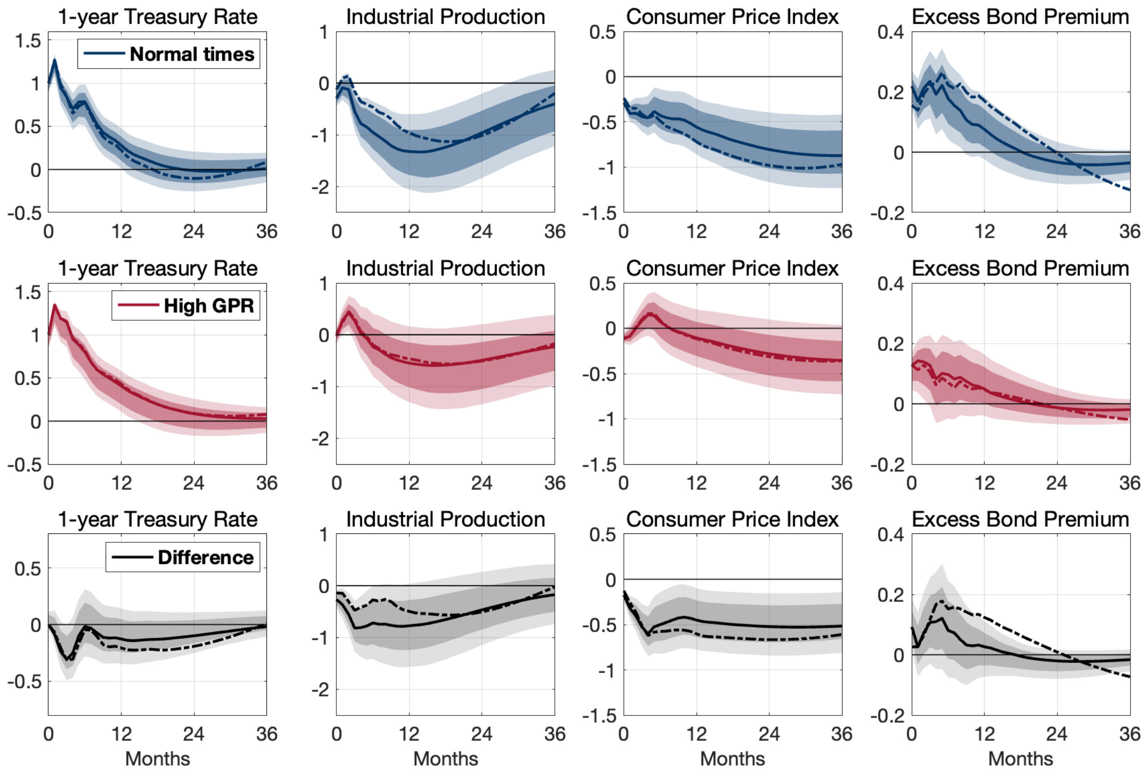
Notes: Seven-variable VAR. Nonlinear impulse responses from a monetary policy shock in normal times (50th percentile of the geopolitical risk index, blue lines), high (96th percentile of the geopolitical risk index, red lines) levels of the geopolitical risk index *Threats and Wars* and the difference between the two impulse responses (black lines). The shock is normalized to induce a 100 basis point increase in the 1-year treasury rate. The solid lines are the point estimates and the shaded areas are 68% and 90% confidence bands.

Figure C.12: Disentangling the geopolitical risk index: Terror Acts



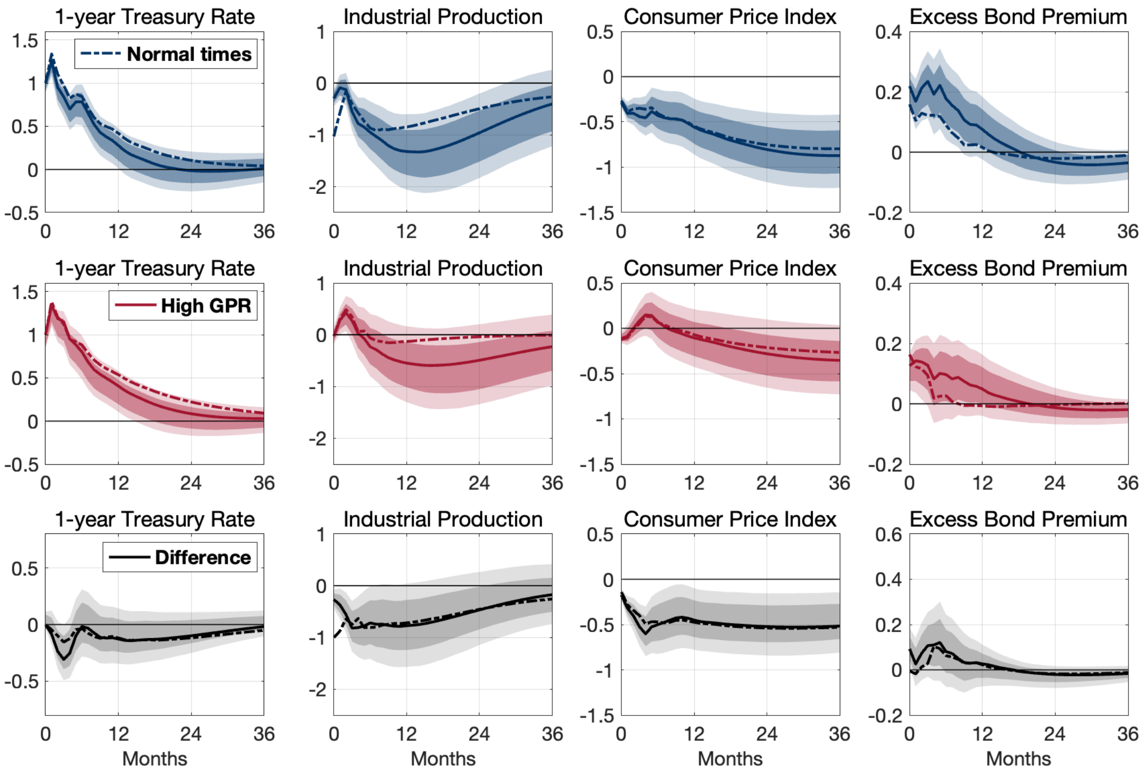
Notes: The dotted blue line is the monthly U.S. historical geopolitical risk index of Caldara and Iacoviello (2022) and the solid orange line is the *Terror Acts* index. The graph reports additional seven geopolitical events. (1) January 1980: US-Soviet tensions as the Soviet Union invaded Afghanistan in December 1979; (2) April 1982: Falkland war; (3) September 1983: Nuclear war scare; (4) June 1985: Hijacking of TWA Flight 847; (5) April 1986: US bombing of Libya; (6) November 2015: Paris terrorist attack; (7) October 2022: Crimea bridge attack and Russia strikes on Ukraine's energy infrastructure; (8) October 2023: Hamas Terrorist attack and Israeli invasion of Gaza.

Figure C.13: Monetary Policy and geopolitical risk: excluding the ZLB



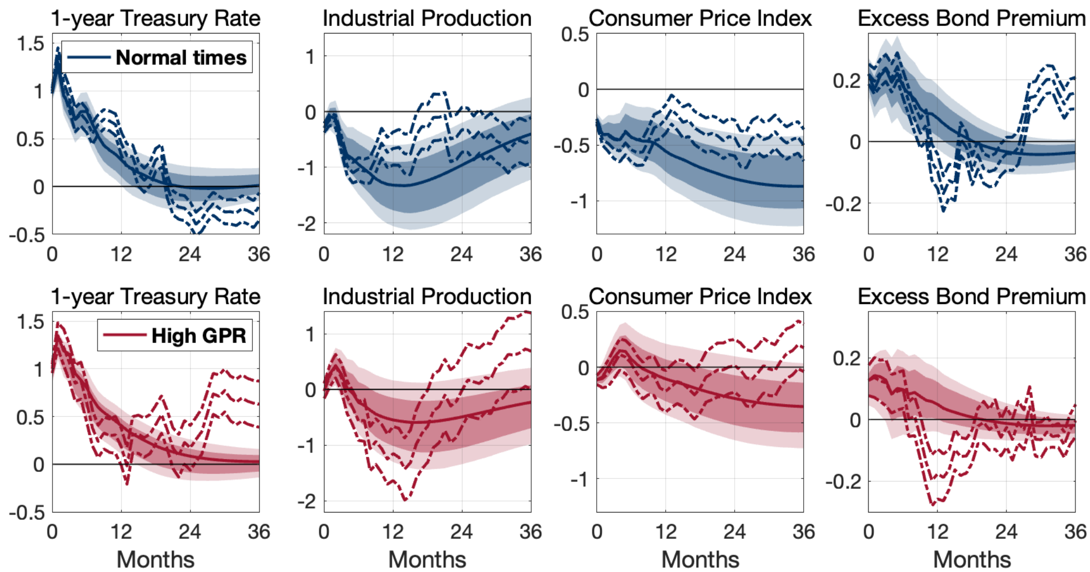
Notes: Six-variable VAR. Nonlinear impulse responses from a monetary policy shock in normal times (50th percentile of the geopolitical risk index, blue lines), high (96th percentile of the geopolitical risk index, red lines) levels of the U.S. geopolitical risk index and the difference between the two impulse responses (black lines). Dashed-dotted lines plot the alternative model excluding the Zero Lower Bound (ZLB). The shock is normalized to induce a 100 basis point increase in the 1-year treasury rate. Sample: 1979m7–2019m12 (Baseline), 1979m7–2008m11 (ZLB). The solid lines are the point estimates and the shaded areas are 68% and 90% confidence bands.

Figure C.14: Monetary Policy and geopolitical risk: including recent years



Notes: Six-variable VAR. Nonlinear impulse responses from a monetary policy shock in normal times (50th percentile of the geopolitical risk index, blue lines), high (96th percentile of the geopolitical risk index, red lines) levels of the U.S. geopolitical risk index and the difference between the two impulse responses (black lines). Dashed-dotted lines plot the alternative model for the recent sample. The shock is normalized to induce a 100 basis point increase in the 1-year treasury rate. Sample: 1979m7–2019m12 (Baseline), 1979m7–2023m12 (recent sample). The solid lines are the point estimates and the shaded areas are 68% and 90% confidence bands.

Figure C.15: Monetary Policy and geopolitical risk: Local projections



Notes: Nonlinear impulse responses from a monetary policy shock in normal times (50th percentile of the geopolitical risk index, blue lines) and high (96th percentile of the geopolitical risk index, red lines) levels of the U.S. geopolitical risk index estimated using the baseline nonlinear Proxy-SVAR. The solid lines are the point estimates and the shaded areas are 68% and 90% confidence bands. Dotted-dashed lines are the nonlinear point estimates using local projections together with 90% confidence bands based on heteroscedasticity-robust standard errors. The shock is normalized to induce a 100 basis point increase in the 1-year treasury rate.