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 VAR and the index of Caldara and Iacoviello (2022), I find that a monetary contraction is less powerful when geopolitical risk is high, with significantly weaker effects on output and prices. The monetary-fiscal mix is central to this result: defense spending contracts after a monetary tightening in normal times but remains unresponsive to the shock during geopolitical tensions. A Lucas critique-robust counterfactual shows that this mechanism can quantitatively account for the documented state-dependent effects.

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

Keywords: monetary policy, geopolitical risk, nonlinear proxy VAR, counterfactuals, monetary-fiscal policy mix.

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1 Introduction

Global geopolitical tensions have soared since Russia's invasion of Ukraine in early 2022, with recent events spreading beyond Eastern Europe to 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. As former Federal Reserve Chair Jerome Powell remarked on February 4, 2024: *"Geopolitical risks, like the wars in Ukraine and the Middle East, are the greatest risk to stability. [...] It's a real risk and one we're aware of."* Moreover, the Federal Reserve Board highlighted in its November 2024 Financial Stability Report that an increase in geopolitical tensions could reduce economic activity and raise inflation. If these risks were to materialize, central banks would need to act to contain inflationary pressures. However, geopolitical tensions, which involve a convolution of supply, demand, and policy forces (Caldara et al., 2026), may alter the economy in ways that may have important implications for the conduct of monetary policy.

In this paper, I explore the interplay between geopolitical risk and monetary policy, a topic that has received little attention in the literature. Using the Geopolitical Risk (GPR) index of Caldara and Iacoviello (2022) and a nonlinear Proxy VAR for the U.S. economy, I show that monetary policy is less effective in periods of elevated geopolitical risk. A contractionary monetary shock does not reduce output and prices in the short run, with both variables declining only gradually thereafter while remaining statistically insignificant. Moreover, credit spreads and inflation expectations are less responsive to the shock. In contrast, a comparable monetary shock in normal times yields conventional effects.

I obtain these results by estimating the effect of a contractionary shock in normal times and in periods of high GPR using the nonlinear Proxy VAR approach developed by Debortoli et al. (2025) (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) until 2018m12. I perform extensive checks on the shock identification, sample period, alternative state variables, and estimation framework. Notably, these sensitivity checks yield no significant changes in the estimates.

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, the index combines episodes associated with potentially different macroeconomic dynamics. In particular, some geopolitical events—such as wars, military buildups, and threats to energy supply—are likely to operate through supply-side disruptions, while others, such as isolated terrorist attacks, are more closely associated with sharp declines in confidence and aggregate

demand. This distinction may be particularly relevant for the transmission of monetary policy, as different geopolitical episodes may be associated with different monetary-fiscal interactions. To this end, I use the subcomponents of the GPR index to construct an alternative measure that excludes the *Terror Acts* component, which is primarily driven by 9/11. The resulting index, which I label *GPR Threats and Wars*, isolates episodes associated with wars, military tensions, and geopolitical threats, which are more likely to be associated with a monetary-fiscal policy mix that is less favorable to the transmission of contractionary monetary policy.

Next, I shed light on the role of fiscal policy for the documented monetary state-dependent effects. The interaction between monetary and fiscal policy is important because monetary policy operates within a broader monetary-fiscal regime, so that its effects depend on fiscal backing and, more generally, on the prevailing policy mix (e.g., Bianchi, 2012; Bianchi and Melosi, 2017). More recently, the importance of monetary-fiscal interactions has also received growing empirical attention (Reichlin et al., 2023; Bouscasse and Hong, 2023; Breitenlechner et al., 2024).

Conventional wisdom in macroeconomics suggests that contractionary monetary policy constrains government budgets through two main channels: higher borrowing costs and lower tax revenues associated with a smaller tax base. These pressures should, in turn, lead to reductions in government spending. To investigate this mechanism empirically, I extend the baseline model to a quarterly VAR with fiscal variables. Consistent with conventional wisdom, a contractionary monetary shock lowers both tax revenues and total government spending in normal times. During periods of high GPR, however, the response of defense spending becomes substantially muted, even though tax revenues and nondefense spending continue to decline. This asymmetry is consistent with the idea that geopolitical tensions generate pressures to maintain defense expenditures, limiting the fiscal adjustment that would normally accompany a monetary tightening. As a result, periods of elevated geopolitical risk may be characterized by a less favorable monetary-fiscal policy mix, helping to explain the weaker transmission of monetary policy during periods of high GPR.

To assess whether fiscal policy can quantitatively account for the GPR-dependent effects of monetary policy, I estimate impulse responses to the monetary shock under an alternative fiscal policy rule. Specifically, I construct a counterfactual in which the response of defense spending during periods of high GPR is constrained to remain as close as possible to its response in normal times, while leaving all other variables unconstrained. In other words, this exercise isolates the contribution of defense spending to the state dependence of monetary policy transmission. To implement this counterfactual, I use the empirical methodology recently proposed by McKay and Wolf (2023). This approach has the advantage of being

robust to the Lucas critique, as it involves subjecting the economy to informationally distinct policy shocks at time 0, with no ex post surprises. The counterfactual scenario indicates that once defense spending in the high-GPR regime is set to respond similarly to that in normal times, monetary policy yields conventional effects. These findings suggest that the systematic response of defense spending can quantitatively explain the lower effectiveness of monetary policy in periods of geopolitical tensions.

In understanding how defense spending can significantly affect the transmission of monetary policy, two recent studies provide valuable insights. The first, Cox et al. (2024), documents that federal purchases – of which defense spending constitutes the largest component – explain approximately half of the variation in aggregate spending. This feature takes greater importance as federal purchases account for a small share in overall spending. Thus, even relatively small movements in defense spending can be important for fluctuations in overall spending. The second, Antolin-Diaz and Surico (2025), shows that defense spending has large and persistent real and nominal effects due to its impact on productivity. Consequently, this productivity-related mechanism may help explain the substantial role of defense spending in the transmission of monetary policy. When including a measure of utilization-adjusted TFP into the model, I find evidence in support of the productivity-related mechanism documented by Antolin-Diaz and Surico (2025).

Finally, I exploit the historical variability of the GPR index to construct policy experiments in which I simulate the effect of a contractionary monetary policy shock under different levels of GPR. These experiments are important because they can inform monetary policymakers about the effectiveness of policy actions at different *levels* of GPR. I find that higher levels of GPR are associated with a reduced responsiveness of defense spending to monetary shocks and a decline in the overall effectiveness of monetary policy, highlighting the critical role of the monetary-fiscal policy mix in shaping monetary transmission during periods of elevated geopolitical risk.

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.[1] 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) and Alpanda et al. (2021) for the

[1]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).

U.S. economy and a panel of advanced economies, respectively. More recently, Debortoli et al. (2025) develop a nonlinear Proxy VAR and, in an application to U.S. monetary policy, find evidence in support of countercyclical monetary policy as it is associated with relatively favorable inflation-unemployment trade-offs. Contrary to previous evidence, Bruns and Piffer (2023) find that monetary easing in the U.S. is more effective in recessions than monetary tightening in expansions.

While most of these papers have focused on monetary policy and the state of the business cycle, a growing body of literature explores alternative states of the economy. Within this literature, 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. Andreasen et al. (2023), using an Interacted-VAR, show that a more proactive monetary policy to stabilize output growth after an uncertainty shock is more effective in expansions than in recessions. Some other recent contributions include Falck et al. (2021) and Ascari and Haber (2022). The first paper uses nonlinear local projections and find that a U.S. monetary policy shock fails to stabilize real activity and inflation during periods of high disagreement in inflation expectation, while the shock exhibits conventional responses during periods of low disagreement. The second paper shows 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. Compared to this literature, my work differs in that I analyze a previously unexplored state of the economy, namely geopolitical risk.

The second strand of a recent and rapidly growing literature includes work that focuses on geopolitical risk. The seminal paper by Caldara and Iacoviello (2022) construct a news-based indicator of GPR and show that a shock to GPR has aggregate recessionary effects, with investment being affected more for industries that are positively exposed to GPR. Caldara et al. (2026) present a systematic analysis of the effects of GPR in a large panel of countries finding that on average supply-side disruptions dominate lower aggregate demand pressures during geopolitical episodes, leading to higher inflation and mixed effects for economic activity. A GPR shock is expansionary in the U.S. but recessionary in the other analyzed countries, with this divergence primarily attributed to higher defense spending. Bondarenko et al. (2024) extends the news-based analysis of Caldara and Iacoviello (2022) by focusing on country-specific newspapers to construct a GPR index that captures a local rather than a Western perspective on geopolitical tensions. Pinchetti (2024) classifies GPR shocks as being associated with disruptions on energy markets or with low aggregate de-

mand pressure. Brignone et al. (2024) find that geopolitical tensions primarily transmit to the economy through higher uncertainty when shocks are large. Kilian et al. (2024) develop a DSGE model of the global economy to study how geopolitical oil price risk is linked to economic fluctuations. The model predicts that GPR is mainly important for the dynamics of oil price volatility and that geopolitically driven oil price uncertainty shocks are not a major driver of fluctuations in macroeconomic aggregates.

The remainder of the paper is organized as follows. Section 2 presents an analysis of the geopolitical risk index. Section 3 describes the econometric methodology of the nonlinear Proxy VAR used. Section 4 presents the baseline empirical evidence on the GPR-dependent monetary effects and some robustness checks. Section 5 presents a decomposition of the GPR index based on its different macroeconomic implications. Section 6 first investigates the role of fiscal policy in monetary transmission, then presents policy counterfactuals and policy experiments. Section 7 concludes.

2 The geopolitical risk index

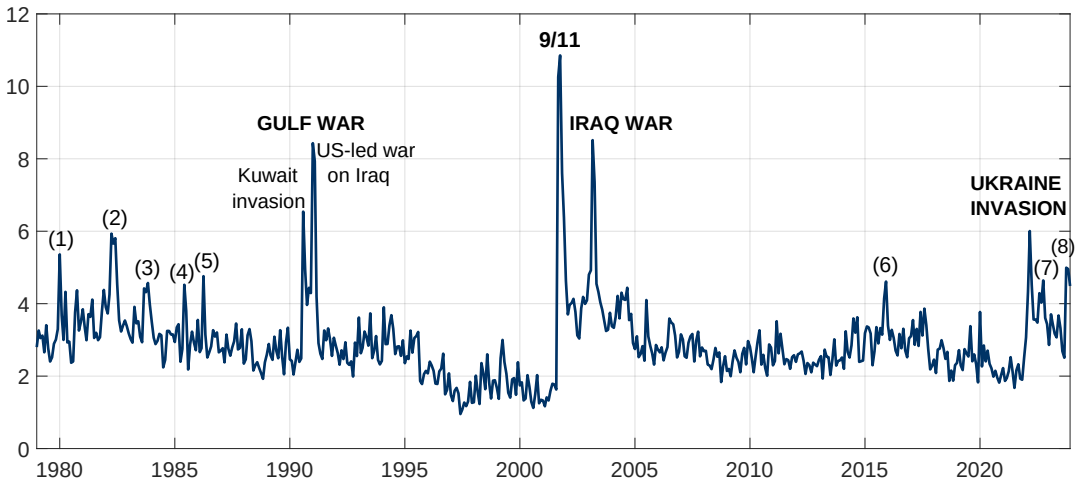
Understanding the interplay between geopolitical tensions and economic dynamics is crucial, as it can help us assess how GPR may affect the effectiveness of monetary policy. Therefore, I now examine the behavior of key economic variables during periods of high GPR, as measured by the index of Caldara and Iacoviello (2022). The authors define GPR 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 ten selected newspapers that mention a dictionary of words related to adverse geopolitical events.[2] Thus, this approach captures an index of how contemporaries in most Western countries perceived the risk associated with a particular geopolitical event, rather than an ex post perception.

Figure 1 plots the historical GPR index for the period 1979m7-2023m12. The major events in the sample are as follows. In August 1990, Iraq invaded the wealthy Gulf state of Kuwait, leading to the so-called Gulf War. In response, a coalition led by the U.S. attacked Iraq in January 1991. During these two events 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 more recent years, the increase in the GPR index reflects two major events. First, the Russian invasion of Ukraine in February 2022, which led to a significant energy

[2]The 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.

crisis in Europe. Second, the Hamas terrorist attack in October 2023 and the subsequent Israeli invasion of Gaza heightened tensions in the Middle East.

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). Major geopolitical events are labeled on the graph, while numbered markers indicate additional eight events: (1) January 1980: US-Soviet tensions as the Soviet Union invaded Afghanistan in December 1979; (2) April 1982: Falklands 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 ground incursion in Gaza.

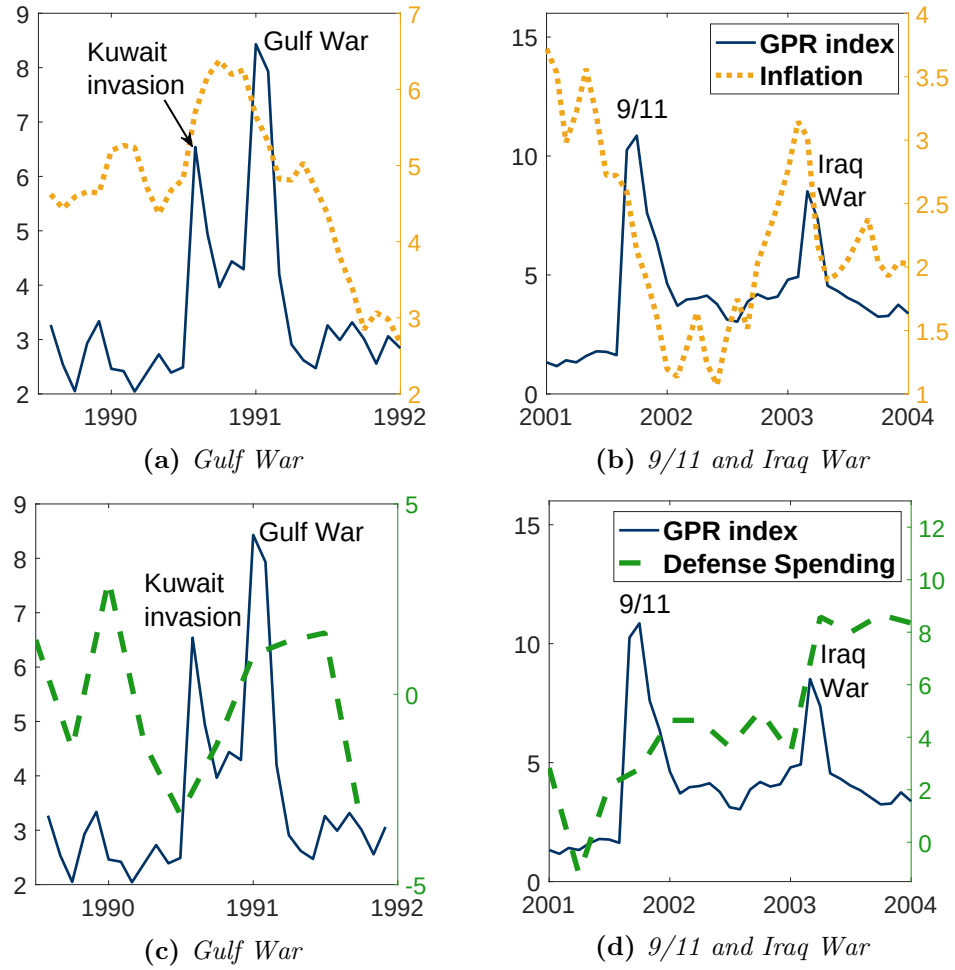
I now analyze the inflation dynamics during the major geopolitical events in the sample. The first row of Figure 2 reports the two series for the Gulf War in (2a), and 9/11 together with the Iraq War in (2b). The figure shows that periods of heightened geopolitical tension are associated with higher inflation rates in the majority of events. From a central bank perspective, these episodes may increase the risk of de-anchoring inflation expectations. By contrast, when geopolitical tensions are associated with weaker aggregate demand pressures, inflation declines significantly. In the sample used, this pattern is observed mainly around the 9/11 attacks. A similar pattern emerges for inflation expectations (see Figure C.1 in Online Appendix C).

As suggested by Caldara et al. (2026), geopolitical events also involve a policy response, particularly in the United States, where they are often associated with higher defense spending. Consequently, I now jointly analyze the dynamics of the GPR index and the annual growth rate of real federal defense spending. All plots in the second row of Figure 2 show that major geopolitical episodes in the sample are associated with increases in defense spending.

Overall, the GPR index of Caldara and Iacoviello (2022) appears to encompass two distinct macroeconomic dynamics: one associated with inflationary pressures, and another

characterized by weaker aggregate demand. At the same time, geopolitical episodes are generally accompanied by increases in defense spending, pointing to an important role for monetary-fiscal interactions during periods of heightened geopolitical tension.

Figure 2: GPR index: inflation (first row) and defense spending (second row)



(a) *Gulf War* (b) *9/11 and Iraq War*

(c) *Gulf War* (d) *9/11 and Iraq War*

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

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*. This assumption postulates partial invertibility, meaning 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). 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

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

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 the vector u_t^{-m}, and possibly nonlinear functions of u_t^{-m}. I include the level of d_t in the equation to isolate the nonlinear effect captured by the interaction $d_t u_t$ from the direct effect of the GPR index. Notice that estimating equation (1) using the lagged value of state variable d_{t-1} in both the level and the interaction terms does not materially alter the results.

I further assume that the vector $[u_t, u_t^{-m}]'$ is $i.i.d.$ zero mean with an identity covariance matrix.[3] The vector $\alpha(L) = \alpha_0 + \alpha_1 L + \alpha_2 L^2 ...$ 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, $\psi(L) = \psi_0 + \psi_1 L + \psi_2 L^2 + ...$ includes the impulse responses to the GPR

[3]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.

index and $\Gamma(L) = \Gamma_0 + \Gamma_1 L + \Gamma_2 L^2 + ...$ is a $n \times m$ matrix of impulse response functions to the remaining structural shocks and their nonlinear functions.

I combine the two terms $\alpha(L)$ and $\beta(L)$ to calculate the nonlinear impulse responses to the monetary policy shock. 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:

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

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

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

where $\Phi(L)e_t$ is the Wold representation of $\Gamma(L)\xi_t$.[4] Under the assumption of invertibility of the Wold representation, the structural model in (1) implies the existence of the following VARX representation:

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

where $\mu = A(1)\nu$, $A(L) = \Phi(L)^{-1}$, $\tilde{\alpha}(L) = A(L)\alpha(L)$, $\tilde{\beta}(L) = A(L)\beta(L)$ and $\tilde{\psi}(L) = A(L)\psi(L)$.[5] The above representation is equivalent to:

$$x_t = \mu + \tilde{A}(L)x_{t-1} + \tilde{\alpha}(L)u_t + \tilde{\beta}(L)d_t u_t + \tilde{\psi}(L)d_t + e_t \tag{4}$$

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 $\tilde{\alpha}(L)u_t + \tilde{\beta}(L)d_t u_t +$

[4] 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.

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

$\tilde{\psi}(L)d_t$ onto a constant and the past history of x_t gives

$$\tilde{\alpha}(L)u_t + \tilde{\beta}(L)d_t u_t + \tilde{\psi}(L)d_t = \theta + C(L)x_{t-1} + w_t. \tag{6}$$

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.

3.3 Identification

In the previous section, 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. When using a proxy variable to estimate the effect of a shock, the parameters of interest will be affected by an attenuation bias in proportion to the measurement error included in the proxy. In a linear framework, the estimated parameters are still affected by an attenuation bias, which, being proportional for all variables, can be eliminated by a standard normalization (Stock and Watson, 2018). Instead, in nonlinear models, measurement error introduces an attenuation bias that generally affects the linear and nonlinear terms differently, making it impossible to correct for the bias and thus combine the parameters from both components. Therefore, direct estimation of the VARX in equation (4) is not feasible (see Forni et al., 2025, for a formal discussion of this point).[6] Below, I discuss how to obtain a consistent estimate of the shock of interest that can be used to derive nonlinear impulse responses within this framework.

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

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

$$z_t = a + bu_t + \delta(L)'x_{t-1} + v_t, \tag{7}$$

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

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

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". This means that the variables in x_t are informationally sufficient to find the monetary policy shock u_t. In other words, the nonlinear terms are not needed to estimate the shock of interest. Note that the same assumption is also necessary in the linear case for the standard procedure to be valid. The assumption of partial invertibility is empirically testable, and the results of this test are presented in Section 4.

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 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 VAR 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

$$z_t = \lambda' \hat{\epsilon}_t + \eta_t. \tag{8}$$

Following Forni et al. (2023), an estimate of the 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 in the literature 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, relative to the nonlinear Proxy VAR 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 unique framework.

4 Baseline results

In this section, I present evidence on the effects of monetary policy shocks during periods of geopolitical tensions. I evaluate the impulse responses at a standardized GPR index value equal to 1.5 standard deviations above its sample mean in the baseline sample and compare them with the linear impulse responses, i.e., $\alpha(L)$ in equation (2). 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. Given that the excess bond premium 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 also provide evidence using a broader measure of credit spread, i.e. the Baa-Aaa corporate bond spread.

As a first step, I select the lag length of the model by standard information criteria and estimate the model to obtain the reduced-form residuals (point I in Section 3.4).[7] To estimate the first-stage regression in (8), I use the instrument of Miranda-Agrippino and Ricco (2021), extended by Degasperi and Ricco (2021) until 2018m12 (point II in Section 3.4), which satisfies the first-stage F-statistic criterion for instrument relevance. I then estimate the nonlinear model in (4) over the sample 1979m7 to 2019m12 by ordinary least squares using the standardized GPR index d_t (point III in Section 3.4).[8] I choose the starting

[7] I take the average of the AIC, BIC and HQC, which is six.

[8] 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.

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). Finally, I compute impulse responses according to equation (2) (point IV in Section 3.4).

To address concerns about the potential endogeneity of the GPR index, I follow Caldara and Iacoviello (2022), who show that macroeconomic, financial, and uncertainty developments do not Granger-cause the GPR index. I apply the same test to examine whether the GPR index is affected by the monetary policy shock identified in this paper. The results indicate that the GPR index is not Granger-caused by the monetary policy shock. Furthermore, the correlation between the two series is effectively zero (–0.007). Taken together, these findings confirm the view that the GPR index is exogenous to the monetary policy shock, both in terms of contemporaneous correlation and predictive content. To further address any residual concerns, I also find that the results are virtually unchanged when using the lagged, predetermined value of the GPR index (d_{t-1}) as state variable.

Before discussing the baseline results, I present the test concerning the validity of partial invertibility (Assumption A2). In other words, I ensure that the endogenous variables of the VAR are sufficient to identify the monetary policy shock.

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. (2023) by checking that the shock u_t is indeed invertible. The test is based on the theoretical result that, if the shock 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

$$z_t = \sum_{k=0}^{r} \lambda_k' \epsilon_{t+k} + \eta_t \tag{9}$$

The invertibility test proposed by Forni et al. (2023) 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 nonzero. I estimate the regression in equation (9) using up to the first five leads ($1 \leq r \leq 5$). The results, shown in Table (1), indicate that we cannot reject the null hypothesis of invertibility for all values of r at the conservative 10% confidence level.

Table 1: Invertibility test

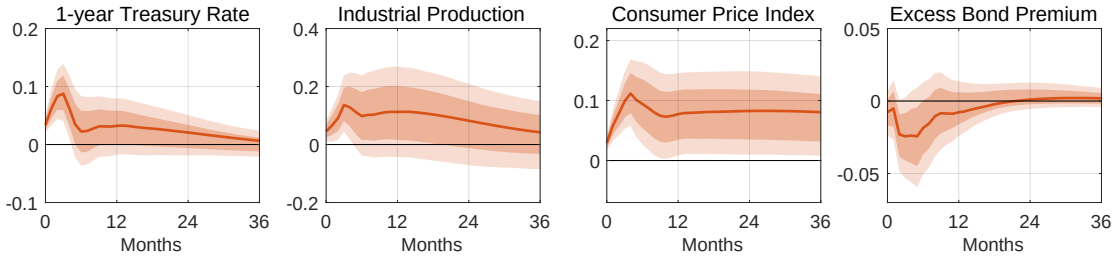
	Number of leads r				
	$r=1$	$r=2$	$r=3$	$r=4$	$r=5$
p-value	0.761	0.902	0.538	0.350	0.322

Notes: 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%) indicate that the shock is invertible.

4.2 The GPR-dependent effects of monetary policy

Before turning to state-dependent impulse responses, I first assess whether the data provide evidence in support of nonlinear effets of monetary policy when GPR is high. Figure 3 presents the results of the test, indicating that the null of linearity can be rejected, with the nonlinear component being most clearly visible in the response of consumer prices, and only marginally in industrial production, primarily over the first few months after the shock.

Figure 3: Monetary policy and GPR: Test of nonlinearity



Notes: Test of nonlinearity $H_0 : \beta(L) = 0$ in equation (2). The plot reports the impulse responses associated to the nonlinear term $\beta(L)$ in equation (2) estimated using a nonlinear Proxy VAR. The solid lines are the point estimates, and the shaded areas are 68% and 90% confidence bands of the extended.

Figure 4 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 in normal times (first row, blue lines), in periods of high GPR (second row, red lines with circles), and their difference (third row, black lines with cross markers).[9] The responses in periods of high GPR are obtained by summing the terms $\alpha(L)$ and $\beta(L)d_t$, with d_t being a value of 1.5 standard deviation, estimated from equation (2). The responses in normal times therefore correspond only to those of the linear term $\alpha(L)$. The shock is normalized to induce a 100 basis point increase in the 1-year Treasury rate. The impulse responses of the interest

[9]The responses for all variables included in the model are reported in Online Appendix C, Figure C.3.

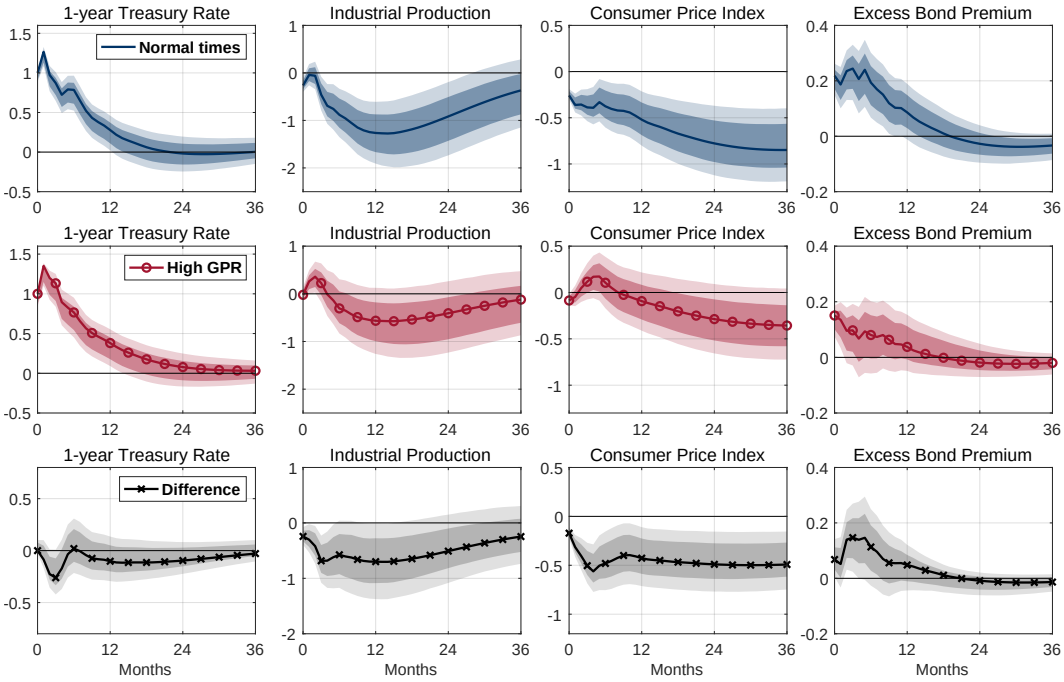
rate and the excess bond premium are measured in percentage point changes, and those of industrial production and prices are measured in percentage changes. The shaded areas are 68% and 90% confidence bands based on 10,000 bootstrap replications.[10] The maximum horizon of the impulse response functions is set at 36 months.

In normal times, a contractionary 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 the recent literature (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 contractionary shock of the same size leads to a temporary increase, although not statistically significant, in output and prices. These puzzling responses are followed by a non statistically significant decline. The excess bond premium rises less on impact than in normal times, before returning more quickly to its initial level.

The third row plots the difference in responses between normal times and the high-GPR scenario and provides information on the statistical relevance of the nonlinearity analyzed. This is equivalent to testing the null hypothesis $H_0 : \beta(L) = 0$ in equation (2), which is independent of the level chosen for defining periods of high GPR (see Figure 3). Comparatively, output responds by less to a monetary policy shock in periods of high GPR, at least for the first year 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 impulse response horizon. Finally, the difference in the excess bond premium response is also statistically significant in the first few months. Using the alternative specification with the Baa-Aaa corporate bond spread, I confirm the previous findings for prices, while I find larger differences in the output and credit spread response between the two regimes (see Figure C.2 in Online Appendix C). Overall, the results presented in this section provide a novel empirical evidence pointing to the lower effectiveness of monetary policy shocks when GPR is high.

[10]The confidence bands include the estimation uncertainty of the estimated monetary shock from the first-stage in (8), because both steps of the estimation are included in the bootstrapping procedure.

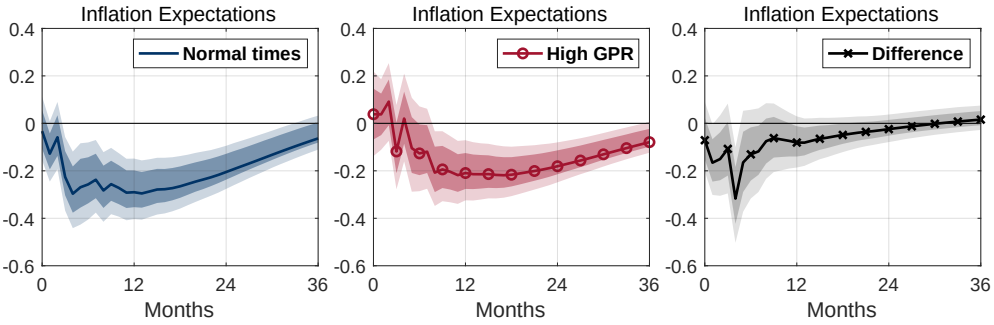
Figure 4: Monetary policy and geopolitical risk



Notes: Nonlinear impulse responses from a monetary policy shock in normal times (blue solid lines), high GPR (red lines with circles) and the difference between the two impulse responses (black lines with cross markers). 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.

Further evidence of the documented state-dependent effects comes from Figure 5, which plots the impulse responses of inflation expectations in the two regimes and their difference. The response of this variable is crucial to understanding the central bank's ability to achieve price stability. I obtain these responses by augmenting the baseline model with inflation expectations from the Michigan Survey of Consumers. Unsurprisingly, household inflation expectations respond sluggishly to monetary policy (D'Acunto et al., 2024), but they decline with a larger delay when GPR is high. The third column of the figure shows that this difference is also statistically significant during the first months. Overall, these results point to a less effective monetary policy in influencing inflation expectations in the short run during geopolitical tensions.

Figure 5: Monetary policy and geopolitical risk: Inflation expectations



Notes: Nonlinear impulse responses of inflation expectations from a monetary policy shock in normal times (blue solid lines), high GPR (red lines with circles) and the difference between the two impulse responses (black lines with cross markers). 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.

4.3 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 econometric model used to 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 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. Online 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 Online Appendix B plots the baseline impulse responses and confidence bands together with the point estimates obtained using the alternative instruments to identify a monetary policy shock. Despite minor differences in the estimates in the two regimes, which are due to the way these proxies are constructed, the state-dependent results obtained with the alternative instruments are fully coherent 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 binding and *(ii)* hold when

the sample is extended to include recent geopolitical events. First, Figure C.6 in the Online 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 (2008m12) 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.7 in Online 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.4 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.[11] 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, which is a more flexible approach than a threshold specification. Formally, I estimate the following VARX

$$x_t = \mu + \tilde{A}(L)x_{t-1} + \tilde{\alpha}(L)u_t + \tilde{\beta}(L)d_t u_t + \psi(L)d_t + \tilde{\Gamma}(L)r_t u_t + \tilde{\phi}(L)r_t + e_t \qquad (10)$$

where r_t takes the value of one during 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., $\tilde{\beta}(L)$. Figure C.5 in Online Appendix C presents a formal test of nonlinearity by plotting the coefficients $\tilde{\beta}(L)$ as orange solid lines together with their confidence bands in orange shaded areas. The figure also shows the point estimates and the 90% confidence bands as dashed lines for the baseline model in (2), i.e., $\beta(L)$. As the test shows, the role of GPR for the effect of monetary policy shocks is largely unaffected in the extended model. If anything, the test using the model in (10) indicates a larger nonlinearity for industrial production and the EBP. Instead, the parameters $\tilde{\Gamma}(L)$ confirm previous evidence on the presence of nonlinearities in the transmission of monetary policy across business cycle

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

regimes.[12] Overall, this sensitivity check suggests that the nonlinearities of monetary policy with respect to GPR and the business cycle should be seen as driven by separate regimes rather than mutually exclusive ones.

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:

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

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 VAR, 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.8 in Online Appendix C plots the baseline responses and confidence bands together with those estimated from local projections, reported as 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 forecast horizon, and display large variance. 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 obtained using the alternative (direct) approach raises questions about the relative advantages of local

[12] I find evidence in line with Bruns and Piffer (2023), in which monetary policy is more effective in recession. Results are not shown and are available upon request.

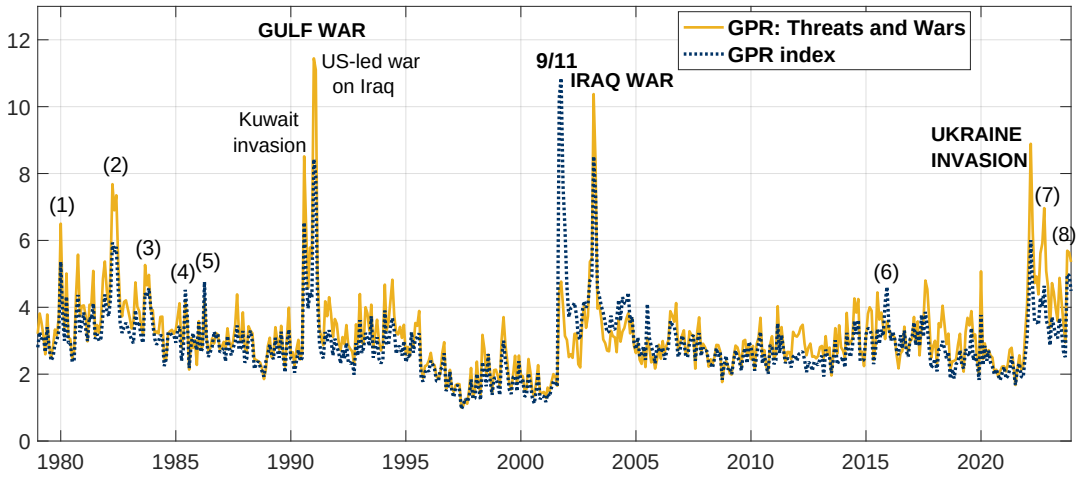
projections in nonlinear settings.

5 Disentangling the Geopolitical Risk index

As shown in Section 2, geopolitical episodes are heterogeneous in terms of their macroeconomic implications. While some events primarily generate uncertainty and declines in confidence, others are associated with military escalation and supply-side disruptions. This heterogeneity may be particularly relevant for the transmission of monetary policy, since the effects of monetary shocks depend importantly on the prevailing monetary-fiscal policy mix (e.g., Bianchi, 2012; Bianchi and Melosi, 2017). In this respect, different geopolitical episodes may give rise to markedly different monetary-fiscal interactions. To this end, I exploit the subcomponents of the GPR index constructed by Caldara and Iacoviello (2022). In particular, I construct an alternative measure that excludes the *Terror Acts* component, which is largely driven by events such as the 9/11 terrorist attacks. Relative to wars and military threats, episodes in this category appear to reflect more strongly the uncertainty and confidence channels associated with geopolitical risk. Historically, 9/11 has been characterized by expansionary fiscal measures, in terms of defense spending, alongside highly accommodative monetary conditions. The resulting index, which I label *GPR Threats and Wars*, isolates episodes associated with wars, military buildups, geopolitical threats, and related tensions.

Figure 6 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* index (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 United States specifically, but only for the global index. Nevertheless, the authors recognize that geopolitical events of global interest often imply the involvement of the United States. 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.9 in Online 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 for the U.S. economy.

Figure 6: Disentangling the geopolitical risk index: Threats and Wars

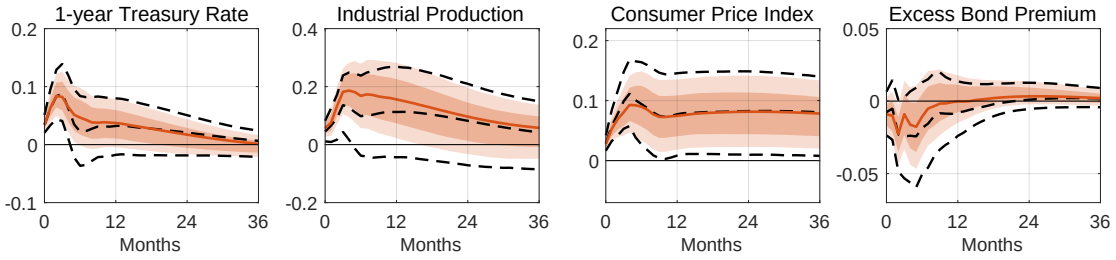


Notes: The dotted blue line is the monthly U.S. historical geopolitical risk index of Caldara and Iacoviello (2022) and the solid yellow line is the *Threats and Wars* index. The graph reports additional seven geopolitical events. (1) January 1980: US-Soviet tensions as the Soviet Union invaded Afghanistan in December 1979; (2) April 1982: Falklands 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.

I next assess whether using the refined GPR index has an impact on the degree of nonlinearity in the effects of monetary policy. Figure 7 compares the nonlinearity test obtained under the baseline GPR index (black dashed lines) with that based on the refined measure (orange solid lines and bands). The results indicate that the refinement matters quantitatively: compared with the baseline GPR measure, the refined index yields stronger evidence against linearity in the response of industrial production, especially at shorter horizons. Overall, the refined index yields more significant results in the test against $H_0 : \beta(L) = 0$ for consumer prices and marginally for the excess bond premium, although the nonlinear component for the latter remains statistically weak.

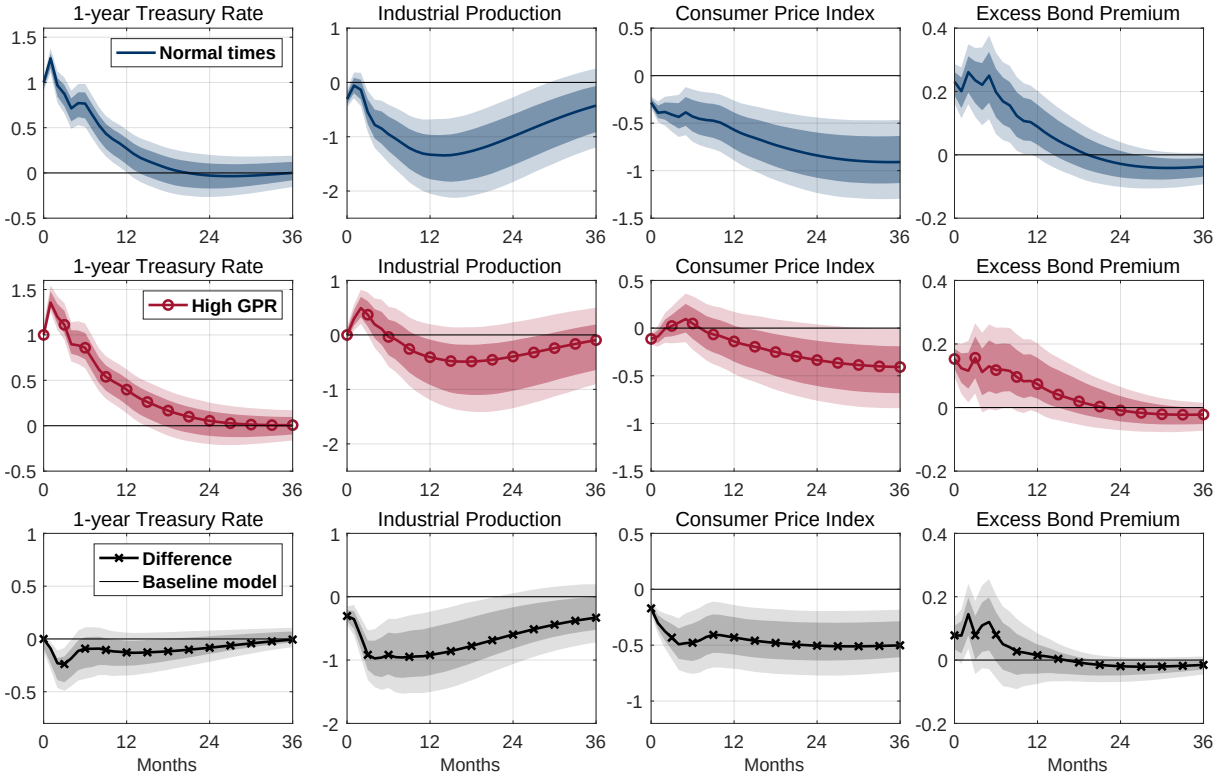
Figure 8 plots the impulse responses following a contractionary monetary shock using the GPR *Threats and Wars* index as the state variable in the model. The impulse responses are broadly consistent with the baseline GPR-dependent results and, as also shown in Figure 7, the third row indicates a more pronounced nonlinear effect of GPR on real activity than in the baseline findings.

Figure 7: Monetary policy and GPR Threats and Wars: Test of nonlinearity



Notes: Test of nonlinearity $H_0 : \beta(L) = 0$ in equation (2). The plot reports the impulse responses associated to the nonlinear term $\beta(L)$ in equation (2) estimated using a nonlinear Proxy VAR. The solid lines are the point estimates, and the shaded areas are 68% and 90% confidence bands of the GPR Threats and Wars index. The black dashed lines report the baseline point estimates and 90% bands of the test in Figure 3.

Figure 8: Monetary policy and GPR Threats and Wars



Notes: Six-variable VAR. Nonlinear impulse responses from a monetary policy shock in normal times (blue solid lines), high GPR (red lines with circles) and the difference between the two impulse responses (black lines with cross markers). 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.

5.1 The Role of Oil Market Stress

A natural question is whether the GPR-dependent monetary policy effects simply reflect oil-market stress episodes. This is a relevant concern as geopolitical events associated with threats of war and military conflict are arguably affecting the economy through oil-market disruptions. To explore this alternative interpretation, I re-estimate the nonlinear Proxy VAR using an indicator of future oil supply disruptions as state variable. Specifically, I use the WTI futures spread between the 1-month and 12-month contracts, so that higher values correspond to tighter oil-market conditions. Alquist and Kilian (2010) and Kilian and Murphy (2014) show that this spread captures the precautionary demand component of the real oil price driven by news of future oil supply disruptions.[13] Specifically, the model takes the form of equation (4), but with d_t defined as the WTI spread rather than the GPR index.

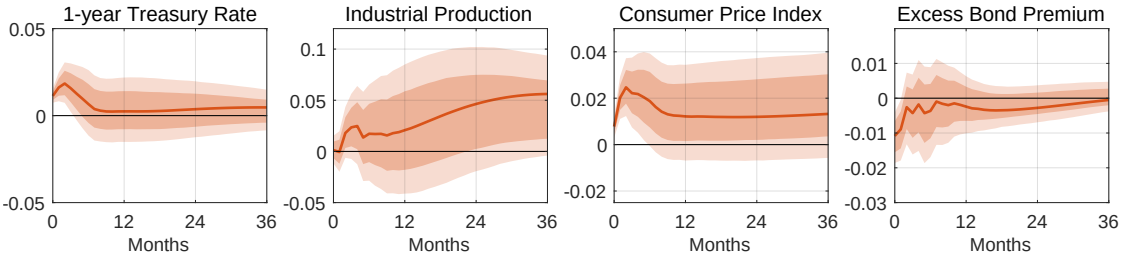
Figure 9 reports the corresponding test of nonlinearity by plotting the point estimates and bands for $\beta(L)$ in the corresponding equation (2). The responses are qualitatively aligned with those obtained under the refined GPR state: monetary policy appears less contractionary when oil-market stress is elevated. Quantitatively, however, the evidence is weaker. Relative to the refined GPR specification in Figure 7, the nonlinear component is mostly not statistically significant, with therefore only limited evidence against the null of linearity. This comparison suggests that oil-market disruptions are an important but incomplete component of the relevant state variation. Put differently, focusing on oil-market stress alone does not span the information contained in the (refined) GPR measure. The latter appears to capture geopolitical episodes in which oil-market conditions interact with broader dimensions of geopolitical tensions that are not summarized by oil-market variables alone. Hence, the results should not be interpreted as evidence that geopolitical risk matters independently of oil markets; rather, they indicate that oil-market stress alone is not sufficient to reproduce the GPR-dependent monetary transmission documented above.

This interpretation is further supported by the behavior of real oil prices following the monetary shock. Figure C.10 in Online Appendix C shows that the response of oil prices to the monetary policy shock is broadly similar across regimes, with the difference remaining statistically insignificant throughout the horizon. Hence, the weaker transmission of monetary policy during periods of high GPR does not appear to operate through a distinct endogenous response of oil prices to the monetary shock.[14]

[13] 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.

[14] The results are virtually identical when using the nominal oil price. Moreover, as shown in Figure C.3 (Online Appendix C), the same argument holds for a broader index of commodity prices.

Figure 9: Monetary policy and Oil Market Stress: Test of nonlinearity



Notes: Test of nonlinearity $H_0 : \beta(L) = 0$ in equation (2) when using as state variable the WTI futures spread between the 1-month and 12-month contracts. The plot reports the impulse responses associated to the nonlinear term $\beta(L)$ in equation (2) estimated using a nonlinear Proxy VAR as defined in Section 5.1. The solid lines are the point estimates, and the shaded areas are 68% and 90% confidence bands of the extended.

6 The Role of Fiscal Policy

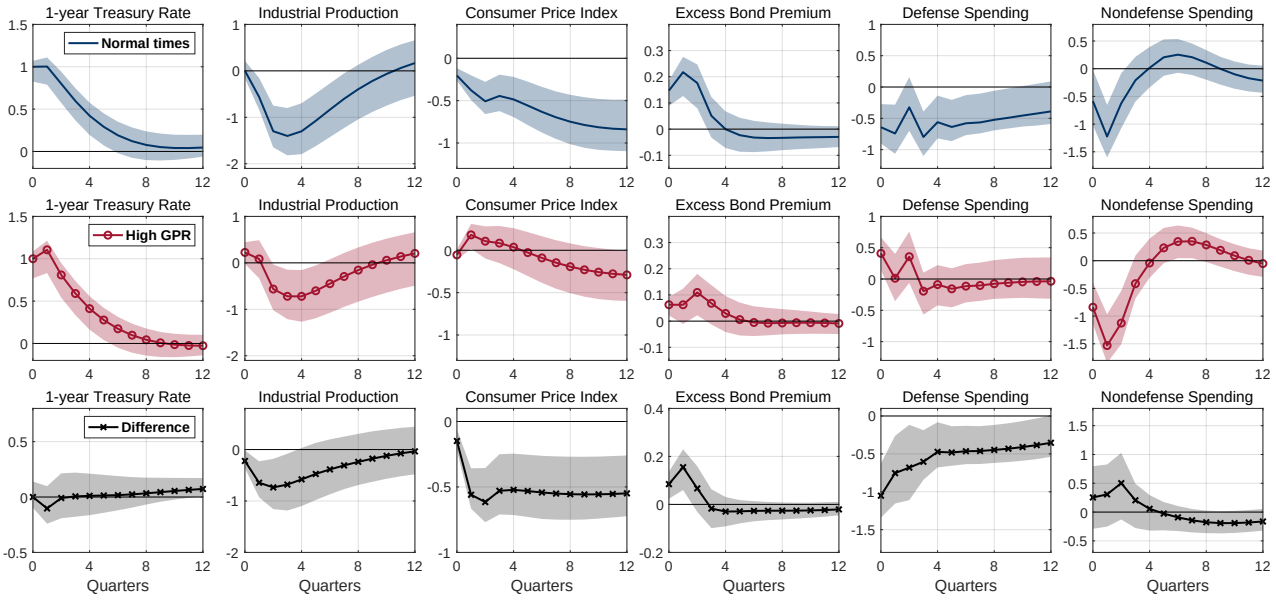
In this section, I study the response of fiscal variables to a monetary policy shock and explore the role that fiscal policy may play in its transmission mechanism. This is because monetary policy does not operate in isolation: its effects depend on the broader fiscal environment and on the prevailing monetary-fiscal policy mix (e.g., Bianchi, 2012; Bianchi and Melosi, 2017). In particular, fiscal backing can shape the aggregate effects of monetary policy by determining whether fiscal policy accommodates or offsets the adjustment induced by monetary tightening. More recently, the importance of the fiscal channel of monetary policy has been explicitly examined by Bouscasse and Hong (2023) and Breitenlechner et al. (2024). However, there is no definitive evidence on the fiscal response to monetary policy. Conventional wisdom in macroeconomics suggests that contractionary monetary policy constrains government budgets through two main channels: raising the cost of new borrowing and by reducing tax revenues through a lower tax base. These fiscal pressures should, in turn, lead to lower government spending.

Methodologically, I estimate a quarterly VAR extended to include defense spending, nondefense spending, tax receipts, and real public debt. Government spending is decomposed into defense and nondefense for the role that defense spending plays in periods of heightened geopolitical tension (Caldara and Iacoviello, 2022). All fiscal variables are at the federal level and are expressed in real per capita terms. The government budget variables enter the model as annual growth rates, and public debt enters in levels and is detrended according to Gordon and Krenn (2010) using an exponential trend for real GDP. The state variable used is the GPR *Threats and Wars* index.

Figure 10 plots the impulse responses as solid lines and 68% confidence bands as shaded

areas for the core variables plus defense and nondefense spending. The responses of tax revenues and public debt are plotted in Figure C.12 and commented in Online Appendix C.1.[15] The results confirm the findings of the monthly VAR for the core variables and reveal important dynamics for the response of government spending. In normal times, a monetary shock lowers both defense and nondefense spending. In periods of high GPR, however, the same shock generates a muted response in defense spending while still reducing nondefense spending. This differential response across regimes likely reflects the pressure on governments to increase defense spending during periods of heightened geopolitical tension, particularly in the United States, and highlights how the interplay between monetary and fiscal policy becomes central to this analysis.

Figure 10: Monetary policy and geopolitical risk: the role of fiscal policy



Notes: Nonlinear impulse responses from a monetary policy shock in normal times (blue solid lines), high GPR (red lines with circles) and the difference between the two impulse responses (black lines with cross markers). 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.

6.1 Policy counterfactuals

In this section, I quantitatively examine the role of fiscal policy in explaining the GPR-dependent effects of monetary policy. Specifically, I investigate whether the effects of monetary policy shocks in periods of high GPR can be explained by the differential response of

[15]The impulse responses include only 68% confidence bands because the quarterly model produces less precise estimates than its monthly counterpart.

defense spending in the two regimes. To conduct this analysis, I use the methodology of McKay and Wolf (2023) (MKW). Their method relies on the core assumption of instrument sufficiency, i.e., private agents do not care about the policy rule per se, but only about the movements in the policy instrument, which is defense spending in this application. This assumption holds in a broad class of macroeconomic models, including conventional New Keynesian theory.

6.1.1 McKay and Wolf's Methodology

Consider a linear, perfect foresight, infinite horizon economy with unobserved variables $\mathbf{w}$, observed non-policy variables $\mathbf{x}$, policy instruments $\mathbf{z}$, structural shocks $\boldsymbol{\varepsilon}$, and policy shocks $\boldsymbol{\nu}$.[16] The economy is summarized by the two-block system:

$$\mathcal{H}_w \mathbf{w} + \mathcal{H}_x \mathbf{x} + \mathcal{H}_z \mathbf{z} + \mathcal{H}_\varepsilon \boldsymbol{\varepsilon} = \mathbf{0} \tag{11}$$

$$\mathcal{A}_x \mathbf{x} + \mathcal{A}_z \mathbf{z} + \boldsymbol{\nu} = \mathbf{0} \tag{12}$$

where $\{\mathcal{H}_w, \mathcal{H}_x, \mathcal{H}_z, \mathcal{H}_\varepsilon\}$ represent the non-policy block, independent of the policy rule coefficients $\{\mathcal{A}_x, \mathcal{A}_z\}$. Assuming that (12) induces a unique equilibrium, the solution to (11)–(12) is

$$\begin{pmatrix} \mathbf{w} \\ \mathbf{x} \\ \mathbf{z} \end{pmatrix} = \underbrace{\begin{pmatrix} \boldsymbol{\Theta}_{w,\varepsilon,\mathcal{A}} & \boldsymbol{\Theta}_{w,\nu,\mathcal{A}} \\ \boldsymbol{\Theta}_{x,\varepsilon,\mathcal{A}} & \boldsymbol{\Theta}_{x,\nu,\mathcal{A}} \\ \boldsymbol{\Theta}_{z,\varepsilon,\mathcal{A}} & \boldsymbol{\Theta}_{z,\nu,\mathcal{A}} \end{pmatrix}}_{\equiv \boldsymbol{\Theta}_{\mathcal{A}}} \times \begin{pmatrix} \boldsymbol{\varepsilon} \\ \boldsymbol{\nu} \end{pmatrix} \tag{13}$$

where $\boldsymbol{\Theta}_{\mathcal{A}}$ contains impulse responses to structural and policy shocks under the prevailing rule. Denote by $\mathbf{x}_{\mathcal{A}}(\boldsymbol{\varepsilon})$ and $\mathbf{z}_{\mathcal{A}}(\boldsymbol{\varepsilon})$ the responses to $\boldsymbol{\varepsilon}$ when $\boldsymbol{\nu} = 0$.

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

$$\tilde{\mathcal{A}}_x \mathbf{x} + \tilde{\mathcal{A}}_z \mathbf{z} = \mathbf{0} \tag{14}$$

MKW's identification requires only $\{\boldsymbol{\Theta}_{x,\nu,\mathcal{A}}, \boldsymbol{\Theta}_{z,\nu,\mathcal{A}}\}$, i.e., the impulse responses of $\mathbf{x}$ and $\mathbf{z}$ to

[16] The vector $\boldsymbol{\nu}$ collects the full menu of contemporaneous and news shocks to the prevailing policy rule at all horizons.

policy shocks $\boldsymbol{\nu}$ under the prevailing rule.[17] The counterfactuals for rule $\{\tilde{\mathcal{A}}_x, \tilde{\mathcal{A}}_z\}$ are then:

$$\mathbf{x}_{\tilde{\mathcal{A}}}(\boldsymbol{\varepsilon}) = \mathbf{x}_{\mathcal{A}}(\varepsilon) + \boldsymbol{\Theta}_{x,\nu,\mathcal{A}} \times \tilde{\nu} \tag{15}$$

$$\mathbf{z}_{\tilde{\mathcal{A}}}(\boldsymbol{\varepsilon}) = \mathbf{z}_{\mathcal{A}}(\varepsilon) + \boldsymbol{\Theta}_{z,\nu,\mathcal{A}} \times \tilde{\nu} \tag{16}$$

where $\tilde{\nu}$ uniquely solve

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

Since in practice only n_s informationally distinct policy shocks are available, MKW suggest finding weights s on these n_s date-0 shocks to best approximate the desired counterfactual rule, by solving:

$$\min_{s} \|\tilde{\mathcal{A}}_x\big(\mathbf{x}_A(\boldsymbol{\varepsilon}) + \boldsymbol{\Omega}_{x,A} \times \mathbf{s}\big) + \tilde{\mathcal{A}}_z\big(\mathbf{z}_A(\boldsymbol{\varepsilon}) + \boldsymbol{\Omega}_{z,A} \times \mathbf{s}\big)\| \tag{17}$$

where columns of $\boldsymbol{\Omega}_{x,\mathcal{A}}$ and $\boldsymbol{\Omega}_{z,\mathcal{A}}$ store the impulse responses of $\mathbf{x}$ and $\mathbf{z}$ to each shock path, and s^* is the best approximation within the space of identified shock paths.

6.1.2 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 used for spending shocks.

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 of this approach.

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 (BP) 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.[18] The second is a *news* or *foresight* shock identified using the Survey

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

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

of Professional Forecasters (SPF) following the approach in Forni and Gambetti (2016), that reflects a shock to the longer end of government spending (see also Caggiano et al., 2015). This variable is defined as the difference between the forecast of government spending growth at time t for the following j quarters, $E_t\Delta g_{t+j}$, and the forecast made at time $t-1$ for the same quarters, $E_{t-1}\Delta g_{t+j}$. 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's variable.The longest common sample available for both shocks is 1981q4-2019q4, driven by the start date of the SPF, which largely overlaps with the sample of the baseline monetary model. Note that the identification of spending shocks focuses on total government spending rather than on defense spending alone, since the latter accounts for a substantial share of the variation in aggregate government expenditures (Ramey, 2011). Also, Cox et al. (2024) document that the BP shock series move mostly the defense rather than the nondefense component of federal purchases, a finding also supported by my estimates (Figure C.11).

6.1.3 Results

The desired counterfactual is obtained according to equations (15)-(16) by focusing on the nonlinear parameters $\beta(L)$ of the quarterly model used in Section 6. It is helpful to recall that the impulse responses in normal times correspond only to the linear coefficients $\alpha(L)$. Given this setup, I can construct the desired counterfactual by minimizing the nonlinear impulse response functions $\beta(L)$ associated with defense spending. I can then simply analyze the dynamics of the other (unrestricted) variables to assess the role that defense spending plays in the GPR-dependent effects of monetary policy.

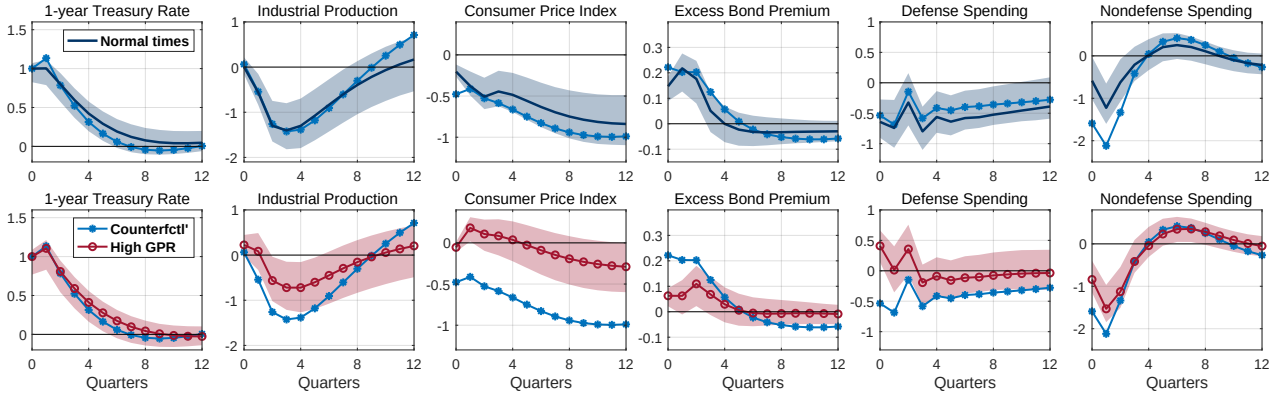
From equations (15)-(16), $\mathbf{x}_{\mathcal{A}}(\varepsilon)$ and $\mathbf{z}_{\mathcal{A}}(\varepsilon)$ correspond to the nonlinear responses of the macroeconomic observables $\mathbf{x}$ and policy instruments $\mathbf{z}$ included in $\beta(L)$ of equation (2). To be internally consistent, I 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 Section 6.1.2, as exogenous components.[19] The final ingredients in estimating policy counterfactuals are the weights $\tilde{\nu}$. These are obtained by solving the problem in (17) and correspond to finding the optimal weights s^* that can impose the desired counterfactual path for the target variable. To achieve this, I minimize the sum of squared deviations of the response of defense spending in the high-GPR regime from that in normal times. Notably, all other endogenous variables remain unconstrained. Overall, this counterfactual scenario allows me to quantitatively assess the extent to which the role of fiscal policy, in terms of defense spending, can account for the GPR-defendent effects of monetary policy.

[19]The impulse responses $\mathbf{\Theta}_{x,\nu,\mathcal{A}}$ and $\mathbf{\Theta}_{z,\nu,\mathcal{A}}$ are reported in Online Appendix C, Figure C.11.

Figure 11 compares the responses shown in Figure 10 (first two rows) with those from the counterfactual analysis for the high-GPR regime (light blue lines with asterisks). For ease of comparison, the counterfactual responses for the high-GPR regime are also reported in the first row. The distance in the defense spending response between normal times (blue solid lines) and the high GPR counterfactual response can be viewed as a measure of the accuracy of the counterfactual rule enforced using MKW's method. Note that this does not imply that the counterfactual rule is necessarily economically meaningful, but only that I am able to mimic the effects of a switch to a counterfactual policy rule. The minor discrepancy between the two responses indicates that the counterfactual rule is enforced with a high degree of accuracy.

Interestingly, the effects of a contractionary monetary shock in the high-GPR regime under the counterfactual rule mirror those estimated in normal times. The initial expansionary effects are now absent. Industrial production and prices fall on impact, closely following their responses in normal times. The same holds for the excess bond premium, which shows a larger response consistent with that in normal times. Finally, the interest rate and nondefense responses in the high-GPR regime remain similar to those under the prevailing rule, suggesting that the results are not driven by a more aggressive monetary contraction or a systematic reduction in nondefense spending. Overall, the evidence from this policy counterfactual supports the importance of the monetary-fiscal policy mix for the transmission of monetary policy. In particular, the differential response of defense spending in the high-GPR regime can quantitatively account for the GPR-dependent effects of monetary shocks.

Figure 11: Monetary policy and geopolitical risk: Policy counterfactual



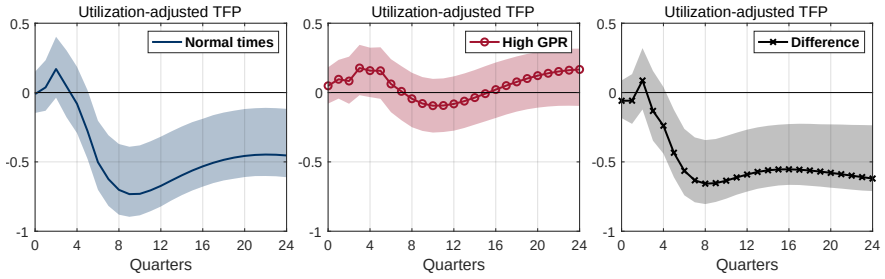
Notes: Nonlinear impulse responses from a monetary policy shock in normal times (blue solid lines) and high GPR (red lines with circles) under the prevailing rule and the counterfactual rule for the high-GPR regime (light blue lines with asterisks). Impulse responses under the counterfactual rule are also reported in the first row for comparison with those under the prevailing rule in normal times. 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.

6.2 Discussion

Two recent studies can help rationalize the important role of defense spending in the transmission of monetary policy. The first, Cox et al. (2024), documents that federal purchases – of which defense spending is the largest component – explain much of the variation in total government spending. This is particularly important as federal purchases represent a small share in overall spending. The second, Antolin-Diaz and Surico (2025), finds that an increase in defense spending has large and persistent real and nominal effects through its effect on productivity.

Two key insights emerge from these contributions. First, even relatively small changes in defense spending can significantly influence fluctuations in overall spending. Second, the mechanism identified by Antolin-Diaz and Surico (2025) can help explain why defense spending plays a major role in the transmission of monetary policy when GPR is high. Taken together, these contributions motivate the following conjecture: the novel productivity channel may operate differently during periods of high GPR. Specifically, if defense spending becomes less responsive to the monetary policy shock when GPR is high, then productivity should exhibit a more muted response to the monetary shock relative to normal times.

Figure 12: Monetary policy and geopolitical risk: the role of productivity



Notes: Nonlinear impulse responses from a monetary policy shock in normal times (blue solid lines), high GPR (red lines with circles) and the difference between the two impulse responses (black lines with cross markers). The model estimated is the quarterly VAR with fiscal variables including the utilization-adjusted TFP measure of Fernald (2014). 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.

To test this hypothesis, I estimate impulse responses of a productivity measure across the two regimes over a longer horizon of six years, consistent with the emphasis on long-run effects in Antolin-Diaz and Surico (2025).[20] Figure 12 reports the estimates. In normal times, productivity declines gradually after a contractionary monetary shock, which is in line with Jordà et al. (2024). In contrast, productivity shows no statistically significant response

[20]The productivity measure is the utilization-adjusted TFP series proposed by Fernald (2014). The model used is the quarterly VAR with fiscal variables as in Section 6 with three lags.

in times of high GPR, with the difference between the two scenarios being statistically significant. Overall, these results suggest that in times of high GPR, the different role of the productivity channel identified by Antolin-Diaz and Surico (2025) may explain the important role of defense spending in monetary policy transmission when GPR is high.

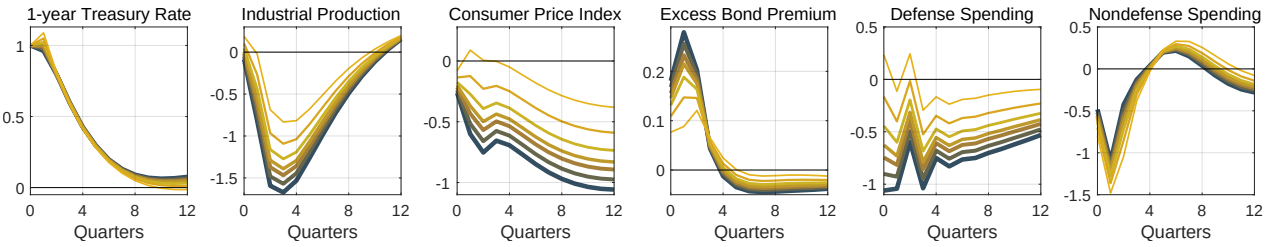
6.3 Policy experiments

I now exploit the historical variability of the GPR index to construct policy experiments around the linear effect of monetary policy by evaluating the nonlinear component of the model at different levels of GPR. These experiments are informative about how the effectiveness of monetary policy varies across different levels of GPR. The analysis uses the quarterly model described in Section 6.

Figure 13 reports the main result of this exercise and shows how the transmission of monetary policy varies across different levels of GPR. The point estimates are reported for a range of scenarios in which the level of geopolitical risk is evaluated at the 30th (thickest blue line), 40th, 50th, 60th, 70th, 80th, and 90th percentiles (thinnest yellow line), with thinner lines corresponding to higher percentiles of the index.[21]

The results suggest that contractionary monetary policy becomes progressively less effective as GPR increases. At higher levels of geopolitical tensions, output and prices respond less strongly to the monetary shock, while defense spending declines proportionally less. By contrast, the response of nondefense spending does not change systematically across GPR levels. Overall, these findings further support the importance of monetary-fiscal interactions for the transmission of monetary policy during periods of heightened geopolitical tensions.

Figure 13: Monetary Policy and GPR Threats and Wars



Notes: Nonlinear impulse responses (point estimates) from a contractionary monetary shock at the 30th (thickest blue line), 40th, 50th, 60th, 70th, 80th and 90th percentiles (thinnest yellow line) of the *Threats and Wars* index. The shock is normalized to induce a 100 basis point increase in the 1-year treasury rate.

[21] For reference, a value equal to 1.5 standard deviations above the mean of the standardized GPR index corresponds approximately to the 93rd–94th percentile of its empirical distribution.

7 Concluding Remarks

Geopolitical tensions involve a combination of supply, demand, and policy forces that can substantially alter the transmission of monetary policy. In this paper, I use a nonlinear Proxy VAR to show that contractionary monetary policy shocks are significantly less effective during periods of elevated geopolitical risk. In these episodes, output, prices, and credit spreads respond less to a monetary tightening than in normal times.

To rationalize these findings, I study the role of fiscal policy in the transmission mechanism of monetary policy. In normal times, contractionary monetary shocks reduce government spending and tax revenues. During periods of elevated geopolitical risk, however, the response of defense spending becomes substantially muted, even though nondefense spending and tax revenues continue to decline. This asymmetry is consistent with the idea that geopolitical tensions generate pressures to maintain defense expenditures, thereby limiting the fiscal adjustment that would otherwise accompany a monetary contraction.

A Lucas critique-robust counterfactual indicates that this mechanism can quantitatively account for the documented state dependence in monetary transmission. Overall, the results underscore the importance of monetary-fiscal interactions for the transmission of monetary policy during periods of heightened geopolitical tensions.

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Online Appendix

Central Banking in Times of High Geopolitical Risk

August 5, 2026

Alessandro Franconi

Appendix A Data Sources

The data used in the paper are the following:

- **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).
- **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), available at the URL https://www.federalreserve.gov/econres/notes/feds-notes/ebp_csv.csv.
- **The (Historical) Geopolitical Risk index, and subcomponents:** Caldara and Iacoviello (2022). Data are downloaded from https://www.matteoiacoviello.com/gpr.htm on January 25, 2024.
- **Baa-Aaa**: Moody's Seasoned Baa Corporate Bond Yield (FRED id: BAA) minus Moody's Seasoned Aaa Corporate Bond Yield (FRED id: AAA).
- **Inflation Expectations:** Median expected price change next 12 months from the Surveys of Consumers, University of Michigan (FRED id: MICH).
- **Defense Spending:** Federal defense spending (FRED id: A997RC1Q027SBEA) deflated using its price index (FRED id: A997RG3Q086SBEA) and divided by population (FRED id: POP). Transformed to percent change from quarter one year ago.
- **Nondefense Spending:** Federal current expenditure (FRED id: FGEXPND) minus Defense (FRED id: A997RC1Q027SBEA) deflated using the GDP deflator (FRED id: GDPDEF) and divided by population (FRED id: POP). Transformed to percent change from quarter one year ago.
- **Tax Revenues:** Federal government current tax receipts, Billions of Dollars, Quarterly, Seasonally Adjusted Annual Rate (FRED id: W006RC1Q027SBEA) deflated using the GDP deflator (FRED id: GDPDEF) and divided by population (FRED id: POP). Transformed to percent change from quarter one year ago.

- **Real Public Debt:** Total Public Debt, Millions of Dollars, Quarterly, Not Seasonally Adjusted (FRED id: GFDEBTN) deflated using the GDP deflator (FRED id: GDPDEF) and detrended with an exponential trend for real GDP as in Gordon and Krenn (2010) (see Section 6).

- **WTI spread:** calculated using WTI crude oil futures hh-month contract, settlement price (Datastream id: NCLC.hh).

- **Inflation:** Consumer Price Index for All Urban Consumers: All Items in U.S. City Average, Percent Change from Year Ago, Monthly, Seasonally Adjusted (FRED id: CPIAUCSL_PC1).

- **NBER Recession periods:** NBER based Recession Indicators for the United States from the Period following the Peak through the Trough, Monthly, Not Seasonally Adjusted (FRED id: USRECD).

- **Utilization-adjusted TFP:** Fernald (2014), available at the website: https://www.johnfernald.net/TFP.

Appendix B Alternative policy instruments

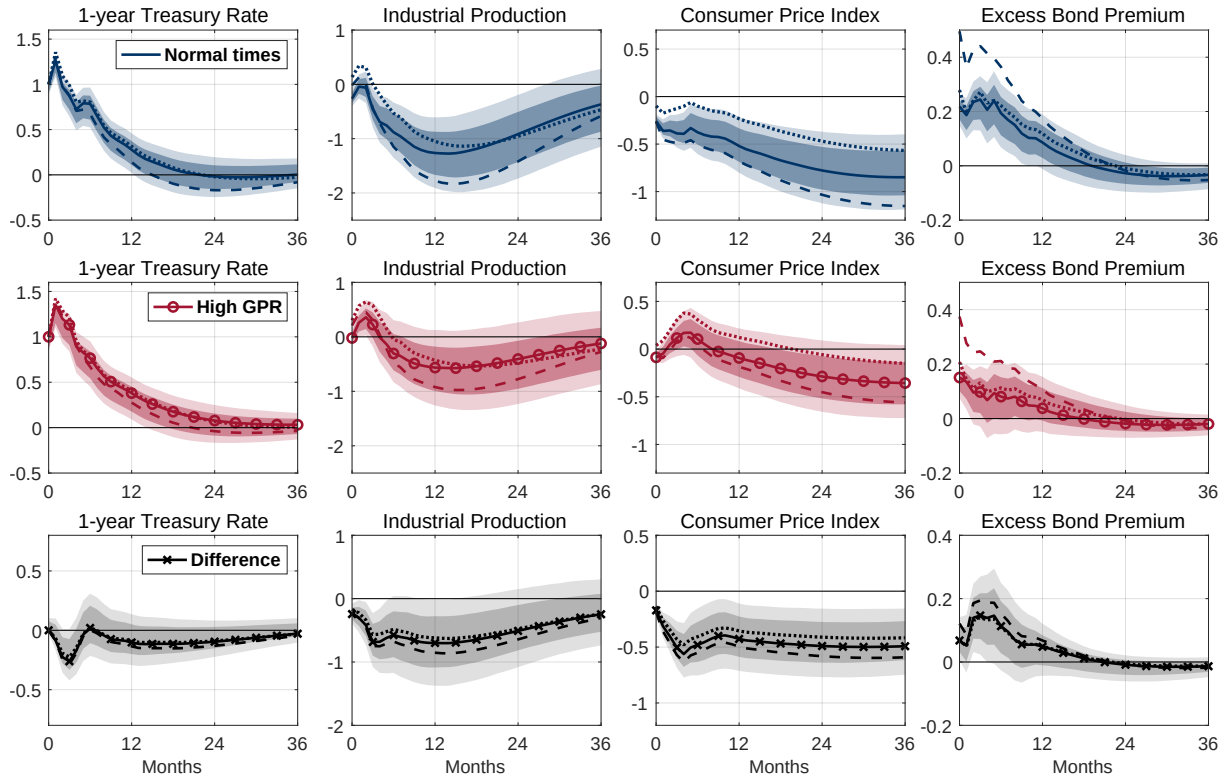
High-frequency instruments for monetary policy shocks are constructed using the financial market reaction around FOMC announcements. Instruments constructed in this way turn out to be predictable by publicly available macroeconomic or financial market information that precedes the FOMC announcement. 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 proxy do not affect the results, Figure B.1 plots the baseline responses along with those obtained identifying the shock using the proxies 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, these alternative identification strategies confirm the baseline state-dependent effects of monetary policy.

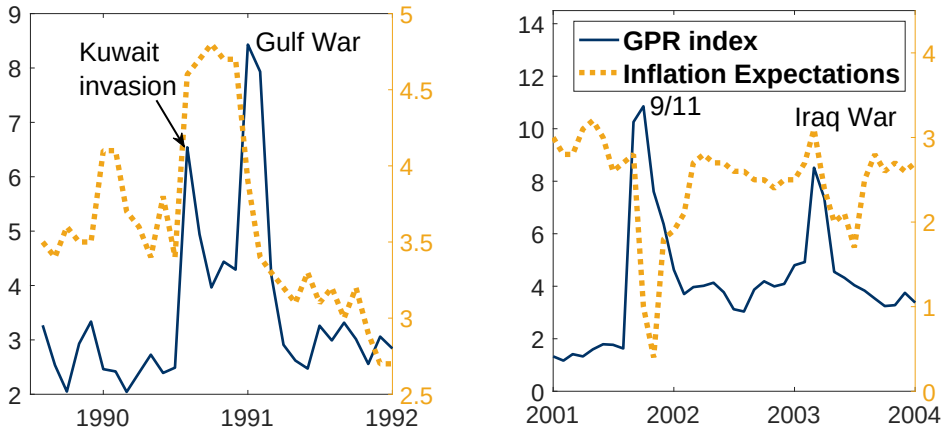
Figure B.1: Alternative shock identifications



Notes: Nonlinear impulse responses from a monetary policy shock in normal times (blue solid lines) and high GPR (red lines with circles) and the difference between the two impulse responses (black lines with cross markers). The monetary policy shock is identified using the baseline instrument of Miranda-Agrippino and Ricco (2021) 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

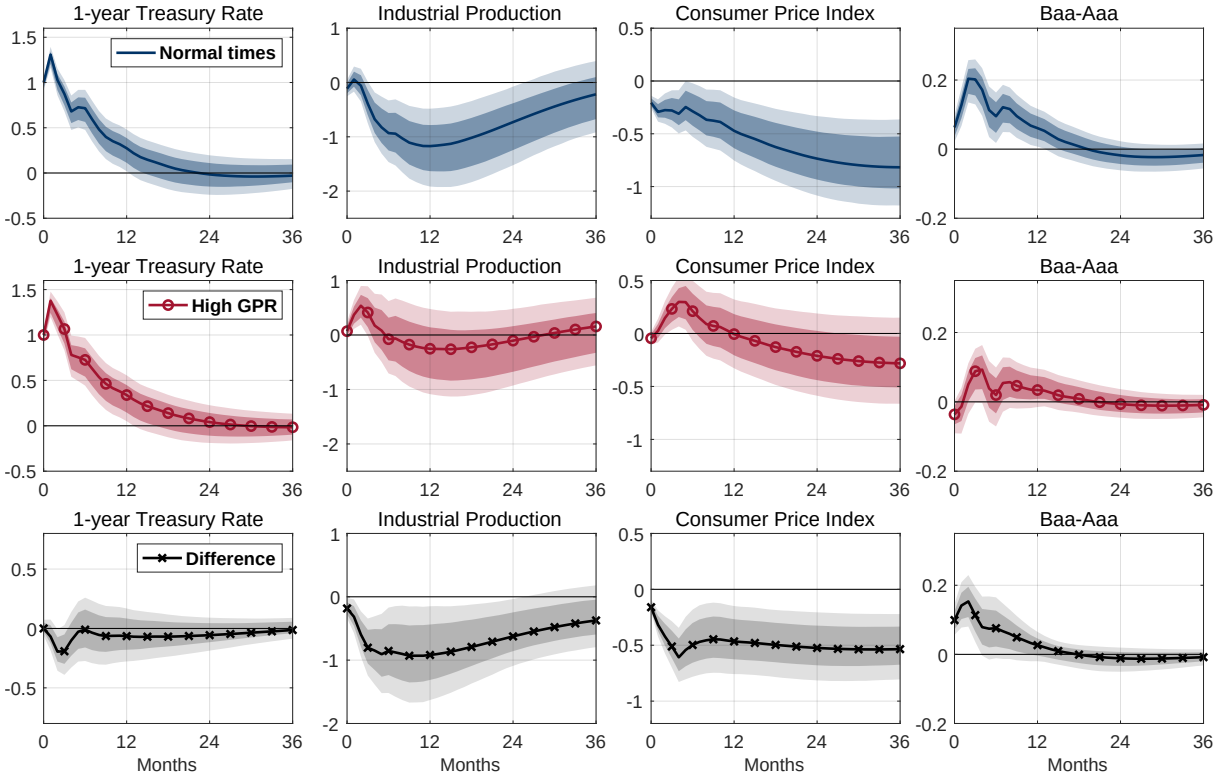
Figure C.1: Geopolitical risk and inflation expectations



(a) *Gulf War* **(b)** *9/11 and Iraq War*

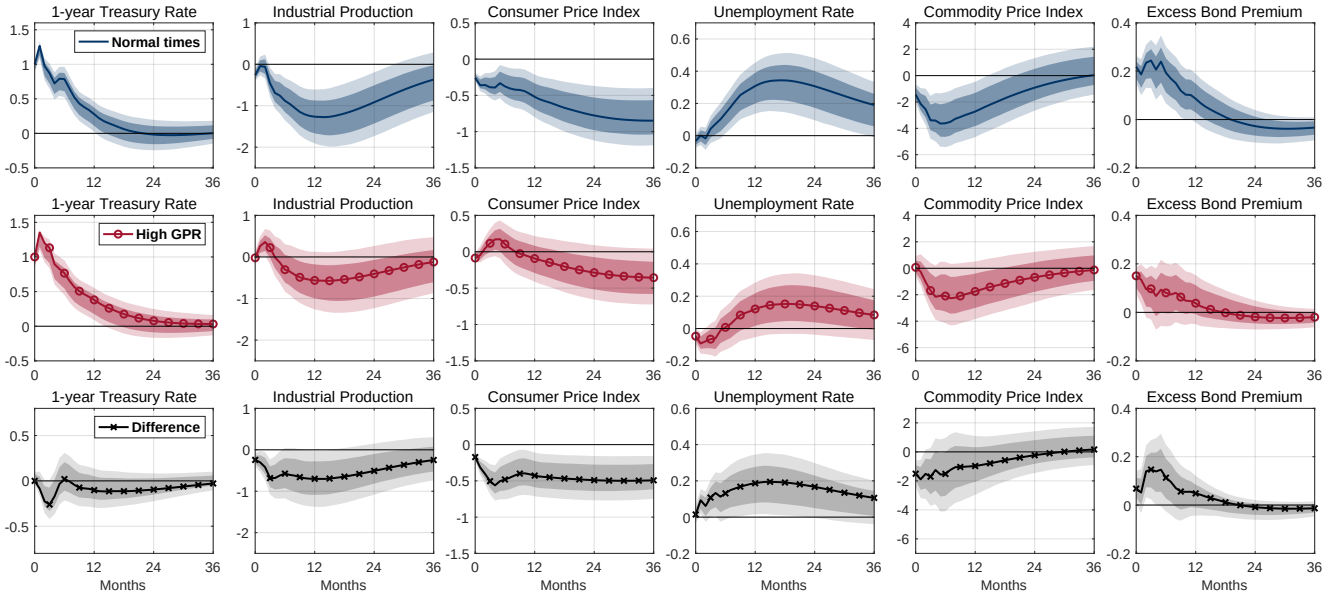
Notes: The solid blue line is the monthly U.S. historical geopolitical risk index of Caldara and Iacoviello (2022) and the yellow dotted line is the inflation forecasts of households from the University of Michigan Survey of Consumers.

Figure C.2: Monetary Policy and geopolitical risk: alternative model



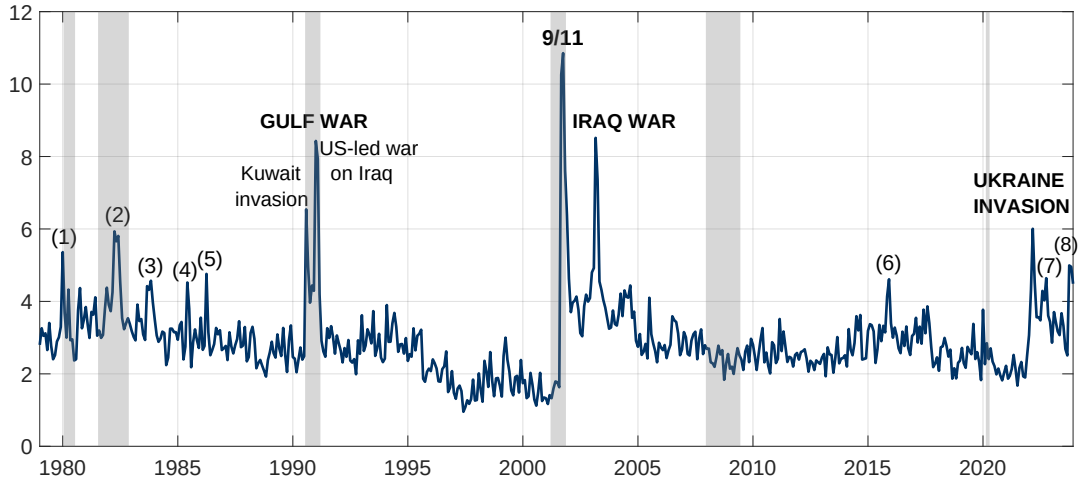
Notes: VAR with the Baa-Aaa spread as an alternative financial variable. Nonlinear impulse responses from a monetary policy shock for low (blue solid lines) and high (red lines with circles) levels of the U.S. geopolitical risk index, and the difference between the two impulse responses (black lines with cross markers). 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.3: Monetary Policy and geopolitical risk



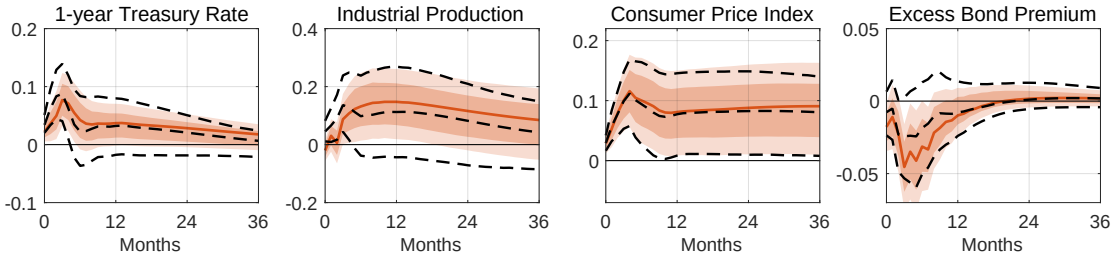
Notes: Nonlinear impulse responses from a monetary policy shock in normal times low (blue solid lines) and high (red lines with circles) levels of the U.S. geopolitical risk index, and the difference between the two impulse responses (black lines with cross markers). The shock is normalized to induce 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: The geopolitical risk index



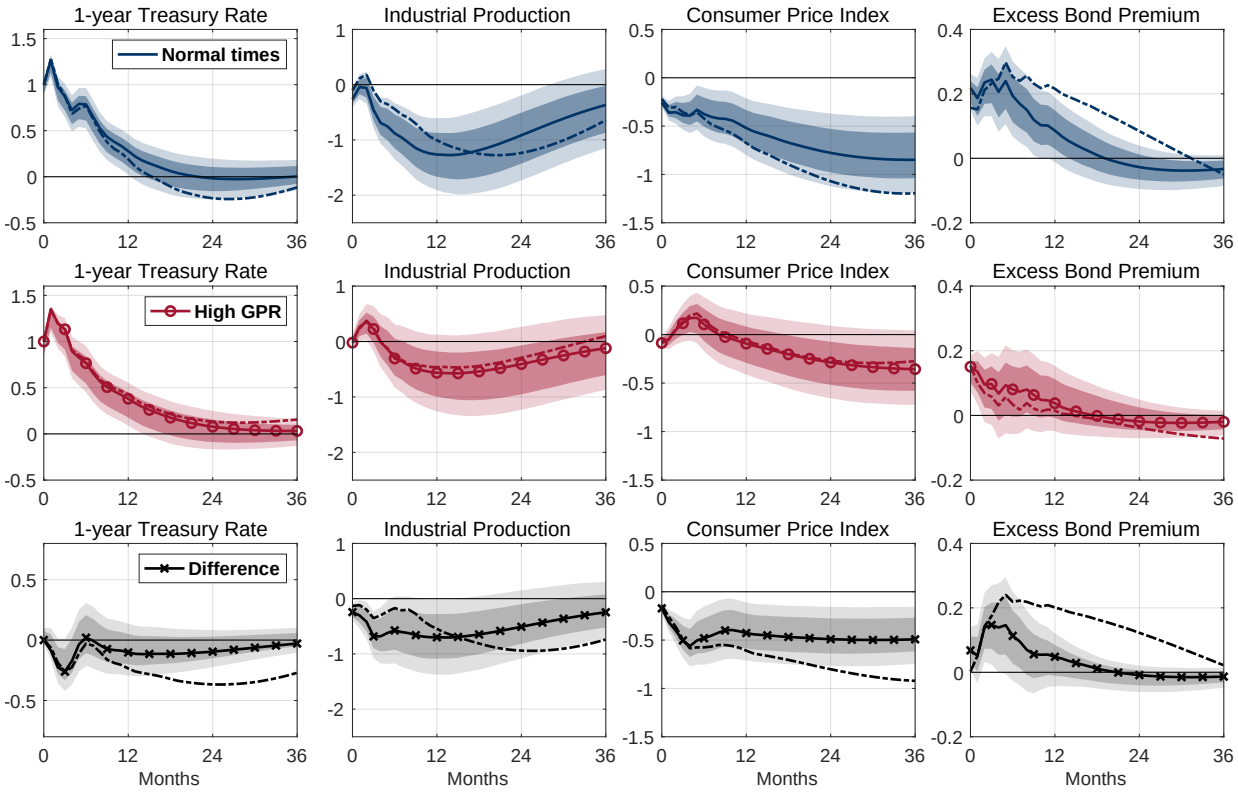
Notes: The solid blue line is the monthly U.S. historical geopolitical risk index of Caldara and Iacoviello (2022) and vertical shaded areas are NBER recession periods. Major geopolitical events are labeled on the graph, while numbered markers indicate additional eight events: (1) January 1980: US-Soviet tensions as the Soviet Union invaded Afghanistan in December 1979; (2) April 1982: Falklands 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.5: Monetary policy, GPR and the business cycle: Test of nonlinearity



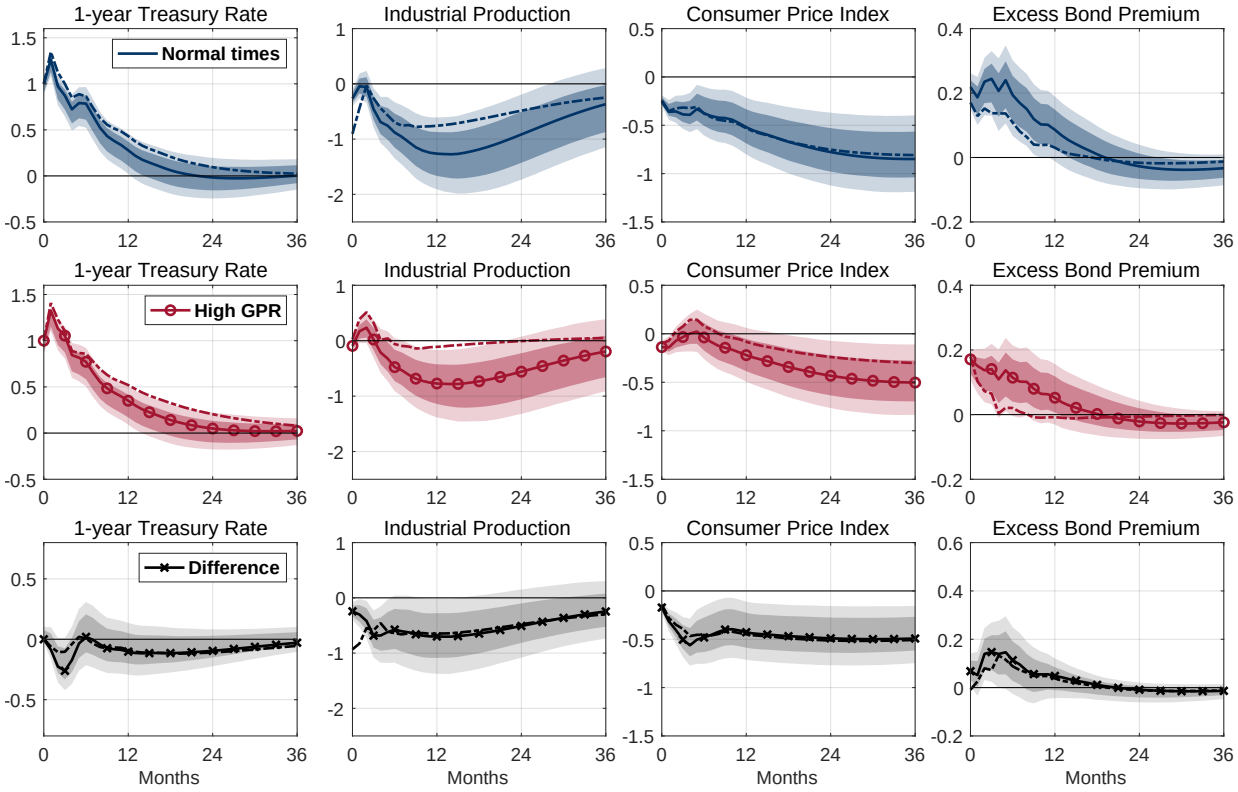
Notes: Test of nonlinearity $H_0 : \beta(L) = 0$ in equation (10). The plot reports the impulse responses associated to the nonlinear term $\beta(L)$ in equation (10) estimated using a nonlinear Proxy VAR. The solid lines are the point estimates, and the shaded areas are 68% and 90% confidence bands of the extended. The dashed black lines are the point estimates and 90% confidence bands for the baseline test of nonlinearity $H_0 : \beta(L) = 0$ in equation (2), shown in Figure 3.

Figure C.6: Monetary Policy and geopolitical risk: excluding the ZLB



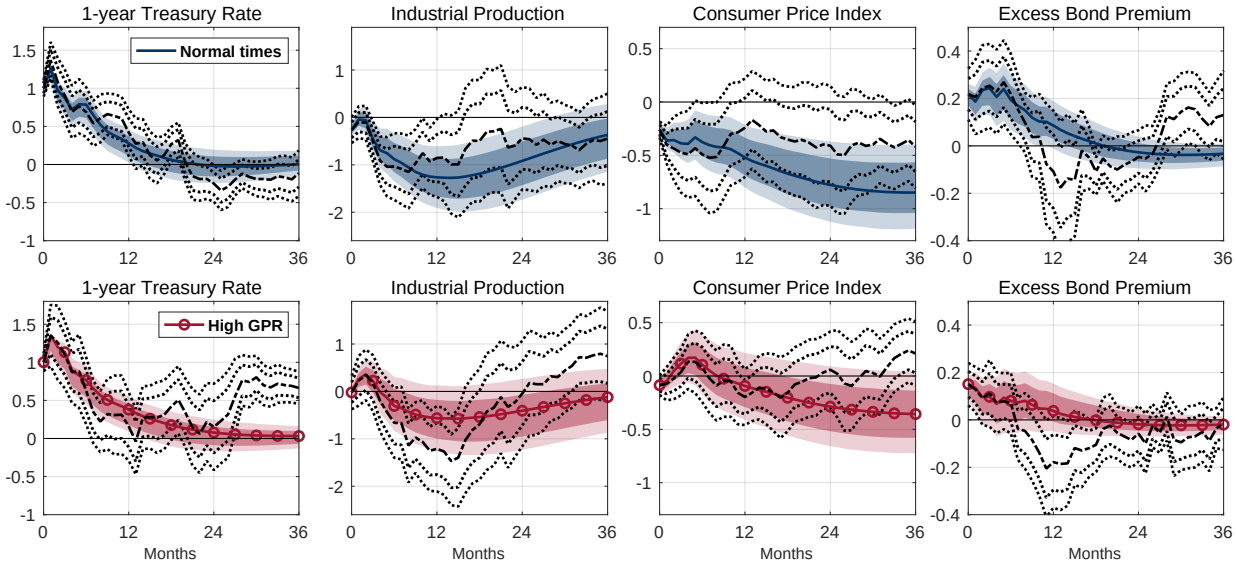
Notes: Nonlinear impulse responses from a monetary policy shock in normal times (blue solid lines), high GPR (red lines with circles) and the difference between the two impulse responses (black lines with cross markers). Dashed-dotted lines plot the alternative model excluding the Zero Lower Bound (ZLB). The shock is normalized to induce a 100 basis point increase in the 1-year treasury rate. Sample: 1979m7–2019m12 (Baseline), 1979m7–2008m12 (ZLB). The solid lines are the point estimates and the shaded areas are 68% and 90% confidence bands.

Figure C.7: Monetary Policy and geopolitical risk: including recent years



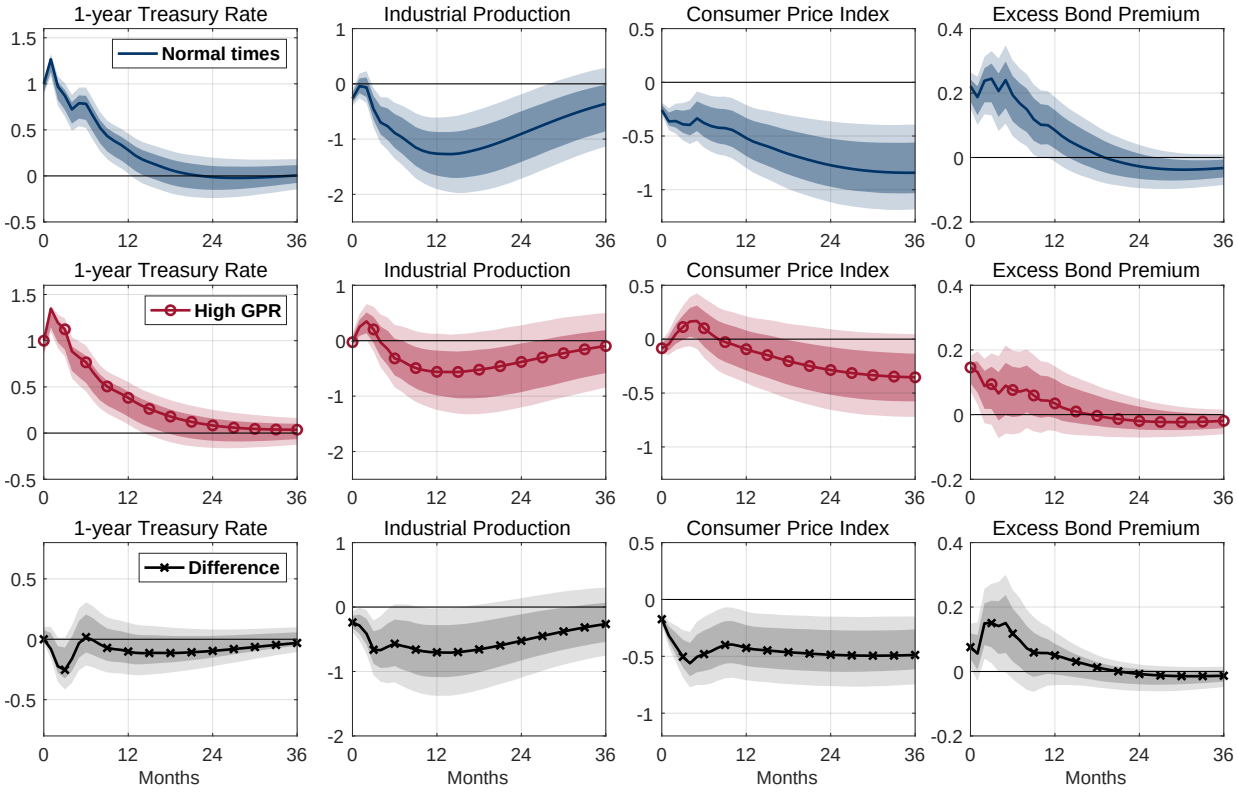
Notes: Nonlinear impulse responses from a monetary policy shock in normal times (blue solid lines), high GPR (red lines with circles) and the difference between the two impulse responses (black lines with cross markers). Dashed-dotted lines plot the alternative model for the recent sample. The shock 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.8: Monetary Policy and geopolitical risk: Local projections



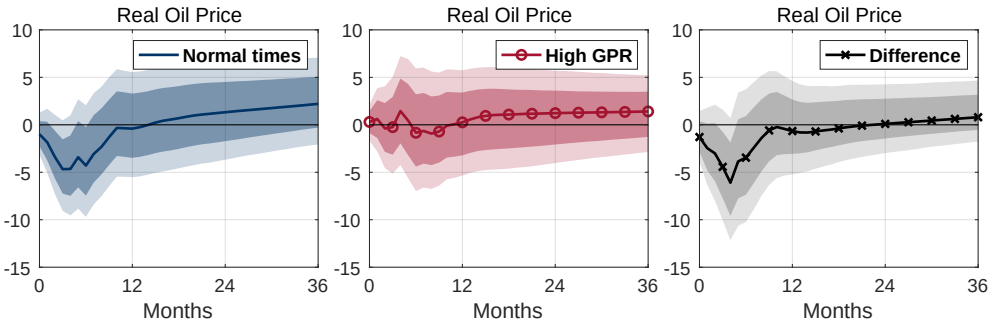
Notes: Nonlinear impulse responses from a monetary policy shock in normal times (blue solid lines) and high GPR (red lines with circles) estimated using the baseline nonlinear Proxy VAR. The solid lines are the point estimates and the shaded areas are 68% and 90% confidence bands. Dashed-dotted lines are point estimates obtained with local projections and dotted lines are 68% and 90% confidence bands. The shock is normalized to induce a 100 basis point increase in the 1-year treasury rate.

Figure C.9: Monetary Policy and global geopolitical risk



Notes: Six-variable VAR. Nonlinear impulse responses from a monetary policy shock in normal times (blue solid lines) high GPR (red lines with circles) and the difference between the two impulse responses (black lines with cross markers). 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 geopolitical risk: Real oil price

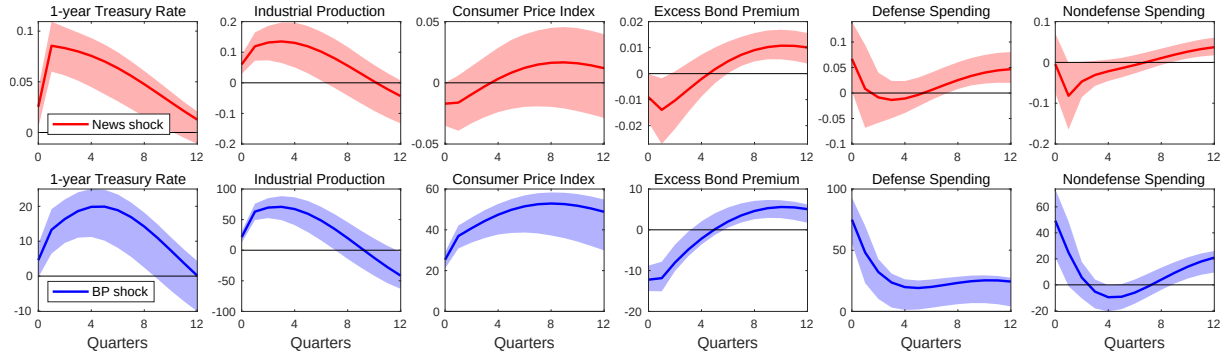


Notes: Nonlinear impulse responses of real oil prices from a monetary policy shock in normal times (blue solid lines), high GPR (red lines with circles) and the difference between the two impulse responses (black lines with cross markers). The shock is normalized to induce a 100 basis point increase in the 1-year treasury rate. The solid lines are the point estimates and the shaded areas are 68% and 90% confidence bands.

C.1 The role of fiscal policy: additional figures

Figure C.11 plots the impulse responses to the two spending shocks used in the counterfactual analysis presented in Section 6.1. Specifically, the impulse responses correspond to the parameters contained in $\mathbf{\Theta}_{x,\nu,\mathcal{A}}$ and $\mathbf{\Theta}_{z,\nu,\mathcal{A}}$. The first row reports the impulse responses to the news shock η^g_{13} (News shock) and the second row to the BP shock. Both shocks increase defense spending, although differently. Given the requirement of McKay and Wolf's methodology, this is an important feature of these two shocks. The BP shock, which is a shock to the shorter end of government spending, displays the largest increase in defense spending at impact with its effects dissipating after few quarters. The News shock, which is a shock to the longer end of government spending, leads to a sluggish increase in defense spending over time with its effects becoming significant not before the second year. Similar dynamics but with a smaller magnitude hold for nondefense spending. Finally, both spending shocks are expansionary for the 1-year treasury rate, output, prices, and the excess bond premium.

Figure C.11: Fiscal impulse response functions

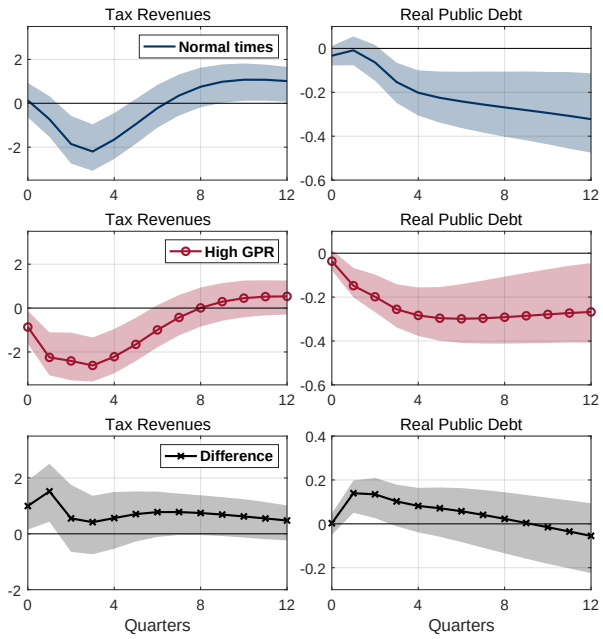


Notes: Impulse responses from the two fiscal spending shocks identified according to Section 6.1.2, corresponding to $\mathbf{\Theta}_{x,\nu,\mathcal{A}}$ and $\mathbf{\Theta}_{z,\nu,\mathcal{A}}$. The solid lines are the point estimates and the shaded areas are 90% confidence bands.

Figure C.12 indicates that tax revenues fall in both regimes, with the third row indicating that these dynamics are mostly statistically indifferent. Real debt exhibits a delayed response in normal times, possibly reflecting a higher cost of real debt due to the significant deflationary effects of monetary policy, before declining over the impulse response horizon. In contrast, when GPR is high, the real value of debt declines. This response is associated with an initial increase in prices following the monetary shock, which mechanically lowers the real value of nominal government liabilities through an inflation-induced revaluation of public debt, a standard mechanism in the fiscal theory of the price level (e.g., Woodford, 1995; Bianchi et al., 2023). As prices tend to fall, the dynamics of real debt become similar

to those in normal times.

Figure C.12: Monetary policy and geopolitical risk: additional fiscal variables



Notes: Nonlinear impulse responses from a monetary policy shock in normal times (blue solid lines), high GPR (red lines with circles) and the difference between the two impulse responses (black lines with cross markers). 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.