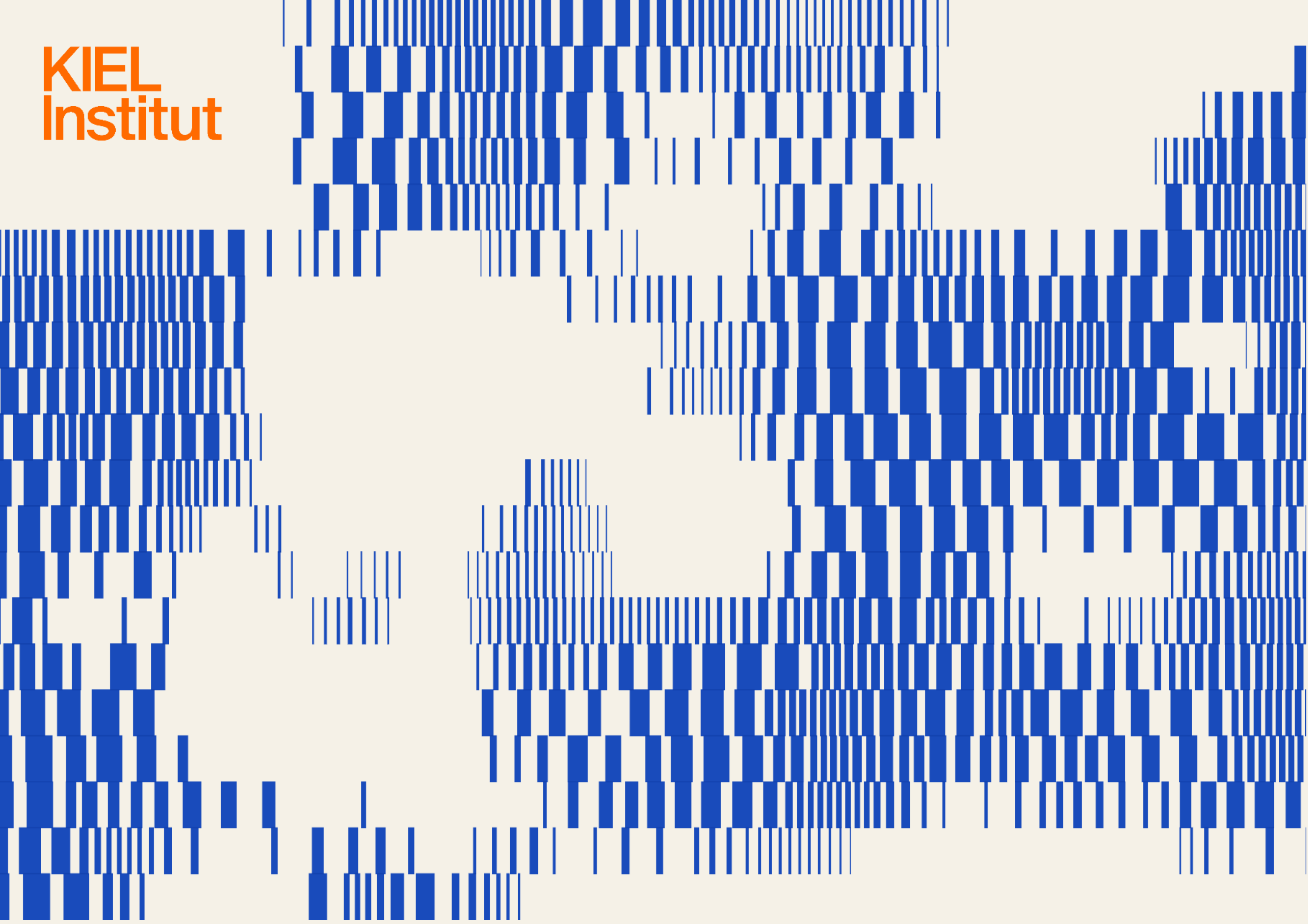

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June 2026

Wars, Threats, and the Sovereign Bond Market

Kiel Working Paper

No. 2323, ISSN 1862-1155
Kiel Institute for the World Economy

Abstract

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We study how wars and military threats affect financial markets. Leveraging more than 300,000 monthly price observations since 1822, we create an external currency bond index for more than 90 countries – the EXBI. Using the EXBI, we document large effects of wars on returns and borrowing costs. In a global external bond portfolio, a one-standard-deviation war shock lowers returns by five percentage points. At the country level, wars at home generate sharp losses and increase default risk. Military threats depress bond prices in threatened states, but not in threatening ones, highlighting their role as a channel for eoeconomic coercion.

Keywords: sovereign bond returns, war risk, geopolitical risk, militarized disputes, sovereign default, global financial history, conflict and finance, international capital markets
JELs: F34, G15, H56, N20, E44

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*We thank Gernot J. Müller, Tarun Ramadorai, as well as our discussants, Barry Eichengreen and Hillary Stein for helpful feedback. Christoph Trebesch acknowledges financial support from the European Union (ERC, Great.Power.Finance, 101087838). Views and opinions are those of the authors and do not necessarily reflect those of the European Union or the European Research Council Executive Agency. Neither the European Union nor the granting authority can be held responsible for them.

1 Introduction

Military conflicts are among the largest shocks to sovereign creditworthiness and financial markets. While wars are infrequent events for any individual country—rare disasters in the sense of Barro (2006)—they are a persistent feature of the international system: since 1870, there has not been a single year without an ongoing war somewhere in the world, and in more than 90 percent of years at least one new conflict has begun. Despite the importance and persistence of geopolitical conflict, we still know relatively little about how war risks are priced in sovereign bond markets.

This paper uses two centuries of bond prices and returns to study how investors price explicit threats and outright war. We show that military conflicts have large and persistent effects on sovereign bond valuations, both globally and at the country level. Wars generate large investor losses, raise sovereign borrowing costs, and increase default risk. A second main insight is that war *threats*, measured as verbal or physical shows of force short of actual violence, depress the bond prices of targeted states. Explicit military threats move financial markets and constitute a distinct channel of geoeconomic coercion akin to Clayton, Maggiori and Schreger (2025*a*).

Our analysis focuses on external sovereign bonds traded in London and New York, which provide a continuous and globally comparable record across countries. Unlike equity markets, which often close or malfunction during major wars, sovereign bonds in financial centers continue to trade.[1] Government bond markets may even expand during wars, as governments continue borrowing to finance military operations. The uninterrupted trading history of external sovereign bonds in US and UK markets therefore allows us to study how wars and military threats are reflected in global capital markets.

We construct a new 200-year external sovereign bond index, the EXBI, at both the global and country level. The EXBI is the first long-run global index of external sovereign bonds, providing monthly bond returns and yields for more than 90 emerging and advanced economies. Existing long-run asset pricing work has focused almost exclusively on *domestic* sovereign bonds, typically covering around 20 countries and relying on annual data.[2] The EXBI fills this gap by providing a monthly global series on hard-currency sovereign bond pricing. It builds on, but condenses and extends, the granular dataset of Meyer, Reinhart

[1] For instance, many stock markets shuttered for extended periods during World War I; see Jorion and Goetzmann (1999) or Silber (2005).

[2] See, for example, Jorda et al. (2019) and Dimson, Marsh and Staunton (2002).

and Trebesch (2022), which contains more than 250,000 monthly bond price observations. Our contribution is to construct standardized, value-weighted index series that are easy to use and extend the data through 2022. In the modern period, the EXBI is highly correlated with JP Morgan's EMBI. However, the EMBI begins only in the 1990s, while our index starts 150 years earlier.

For studying war risk in asset markets, the EXBI has three advantages. First, the series extend back to the nineteenth century, allowing us to analyze periods in which wars were large, frequent, and involved major advanced and emerging economies.[3] Second, the data cover more than 90 countries, enabling us to examine both the global spillovers of major geopolitical shocks—such as World War I and II—and the local effects of more idiosyncratic wars across diverse historical and geographic settings. Third, the monthly frequency of the data is critical for studying how markets price threats of coercion and war. By combining monthly bond price data with monthly records of military threats, we can identify the timing and magnitude of market reactions to threats, both historically and today.

We merge these financial data with newly geolocated information on wars since 1870, drawn from the Price of War dataset of Federle et al. (2026), the Correlates of War Project, and the Uppsala Conflict Data Program. These sources provide data on interstate wars between sovereign states, as well as intrastate wars (e.g., civil wars) and extrastate wars (e.g., wars of independence). Our baseline specification uses the combined sample of wars from these sources, though we show that the results are robust to alternative definitions and subsets of conflicts.

To measure threats directed toward other countries, we use data on militarized interstate disputes (MIDs) from the Correlates of War Project, which records explicit threats and displays of force at monthly frequency (Maoz et al., 2019). The dataset codes "explicit threats," in which one government announces that it may respond with military force if another government fails to act, or does not refrain from acting, in a specified matter (Jones, Bremer and Singer, 1996; Palmer et al., 2022). The data also record coercive displays of force, such as troop mobilizations near borders or naval deployments. These threats are frequently followed by actual war. Together, these sources allow us to study how the onset and intensity of wars, as well as the anticipation of conflict through military threats, are reflected in sovereign bond returns.

[3] In recent decades, wars have mostly been internal rather than external and typically occurred in smaller, poorer, and economically peripheral countries.

Our results can be summarized in three findings. First, using the global EXBI time series, we find that wars exert a major influence on the value-weighted global bond market. A one-standard-deviation global war shock—corresponding to casualties of roughly two basis points of the world population—reduces global bond returns by about five percentage points. The magnitude is comparable to that of major global recessions. We find that the global incidence of wars is strongly associated with the prevalence of economic crises—including currency, inflation, and banking crises, as well as domestic and external sovereign bond crises and stock market crashes.

Second, in a country-year panel, *domestic wars*, defined as wars with military actions on a country's own territory, are rare but disastrous for sovereign bonds. Despite an annual probability of only about 3%, domestic wars account for roughly one quarter of all bond collapses in the historical sample. The average domestic war lowers returns by almost 10 percentage points on impact and raises the probability of default by about 7 percentage points.

Third, coercive military threats have significant effects on bond markets. When one government threatens another, the probability of war rises between the two countries. However, the increase in domestic war risk is concentrated in the targeted country rather than the threatening country. Bond markets price this asymmetry immediately: bonds of targeted countries experience losses of roughly one percentage point at the time threats are issued, whereas bonds of threatening countries show little response.

Related literature Our findings relate closely to research on the financial and economic consequences of wars and war risks, a literature that goes back at least to Clark (1916) and is currently seeing a revival. Pflueger and Yared (2024) examine the joint determination of military spending, geopolitical risk, and government bond prices in a dynamic model of global power rivalry. Federle et al. (2026) study the macroeconomic effects of wars—particularly on GDP and inflation—in a panel extending to 1870. They address the empirical challenge that war onsets may be influenced by domestic economic conditions by focusing on a subsample of interstate wars whose origins are narratively identified as unrelated to short-run economic considerations. We use a similarly broad sample of countries and years and build on their identification strategy and data on war sites. However, we leverage both monthly and annual data and focus on sovereign debt markets, analyzing bond returns, yields, and default risk, and we additionally consider military threats short of war. Hirshleifer, Mai and

Pukthuanthong (2023, 2025) study the impact of war discourse, captured by millions of New York Times news articles, on the pricing of equities in the United States across 160 years. The papers find significant effects of perceived war risks both on the time series and cross-section of US stock returns. Our paper also uses a historical perspective, but studies pricing effects for bonds rather than stocks, considers both perceived and realized war risk, and includes more than 90 countries worldwide. Also closely related is Ferguson (2006) who studies how nineteenth-century political crises affected European sovereign bond markets before World War I. Ferguson argues investors appeared surprisingly insensitive to geopolitical tensions in the decades before 1914, suggesting that World War I came as a "bolt from the blue" for financial markets. Additional papers on the consequences of war on the economy and trade include, among others, Ferguson (2008), Glick and Taylor (2010), Acemoglu et al. (2011), Rohner, Thoenig and Zilibotti (2013), Couttenier et al. (2019), Caldara and Iacoviello (2022), or Korovkin and Makarin (2023).[4]

We also contribute to the fast-growing literature on geoeconomics (for surveys see Clayton, Maggiori and Schreger, 2025*b*; Mohr and Trebesch, 2025). Clayton, Maggiori and Schreger (2025*a*, 2024) examine how governments use economic and financial threats as instruments of statecraft to achieve economic or geopolitical objectives. We show that war threats affect investor returns and sovereign borrowing costs in targeted states, but not in threatening ones. This finding speaks directly to a core question in the geoeconomics literature: how geoeconomic threats or acts of "geoeconomic pressure" shape real and financial outcomes (Clayton et al., 2025). Notably, the influential book by Blackwill and Harris (2016) explicitly lists a country's ability to influence foreign bond markets as one of the most effective tools of geoeconomic power. We provide the first systematic evidence supporting this mechanism: by issuing military threats, countries can depress the bond prices of their adversaries.

More broadly, this paper is among the first to connect the emerging literature on geoeconomics and interstate coercion with the longstanding literature on sovereign debt and default (Panizza, Sturzenegger and Zettelmeyer, 2009; Mitchener and Trebesch, 2023).[5]

Our paper also relates to the literature on financial crises and macroeconomic disasters, e.g. Barro (2006), Barro and Ursúa (2008), Reinhart and Rogoff (2009), Sufi and Taylor (2016). We document that, from a macro-financial perspective,

[4]On the broader economic literature on wars and conflict see Dube et al. (2024).

[5]Horn, Reinhart and Trebesch (2024) show that global wars have substantial effects on sovereign bond issuance in international markets.

domestic wars are "multicrises." Rather than isolated fiscal or geopolitical shocks, we show that major wars trigger large capital destruction and concurrent disruptions in output, inflation, public finances, sovereign risk, and the stock market and currency system. These multifactor dynamics echo the literature on "twin and triplet crises," which shows that banking, currency, and sovereign default crises often occur together and amplify each other, and that defaults, banking crises, currency crashes, and inflationary episodes tend to cluster historically (Kaminsky and Reinhart, 1999; Reinhart and Rogoff, 2009). In line with this literature, we view the bond market response to domestic wars through the same lens, since wars, as we show below, often trigger all of these crises at once.

The remainder of this paper is structured as follows. In section 2, we detail the data sources and construction of the EXBI. Section 3 presents the empirical framework and results for sovereign bond responses to domestic war. In Section 4, we turn to the pricing of military threats. Lastly, Section 5 offers a brief conclusion.

2 Bond Prices and Returns: The EXBI Index

In Meyer, Reinhart and Trebesch (2022), the authors compile a unique 200-year dataset on sovereign bond prices and restructurings to study long-run investor returns in the external bond market. Their data cover 255,000 monthly bond price observations for about 1,550 foreign-currency sovereign bonds traded in London and New York between 1815 and 2016, spanning 91 countries. The database includes British-pound and US-dollar denominated government bonds with fixed coupons and maturities of at least one year. It draws from archival sources such as The Economist, Money Market Review, Investor's Monthly Manual, Commercial and Financial Chronicle, and later JP Morgan's EMBI Global data. The dataset is complemented by a newly constructed archive of over 300 sovereign debt restructurings since 1815, documenting missed payments, restructuring terms, and estimated investor haircuts. We extend these data until 2022 using financial data from JP Morgan.

The resulting dataset allows for a consistent computation of returns on sovereign bonds across centuries, including during wars, defaults, and crises. It captures both price dynamics and credit events, linking monthly prices to detailed information on defaults and recoveries. Notably, the median haircut in defaults is below 50%, and outright repudiations are rare.

We use these data beginning in 1870 for our analysis on wars and from 1822 onward for threats, to conform with the conflict datasets, which we describe in more detail below. We report full sample summary statistics on bonds, recognizing that the analyses of wars and threats in Sections 3 and 4 use a truncated sample due to data availability.

Our goal is to construct two types of return and yield series: a global one, which we dub EXBI, and a country-specific one, which we term EXBI-C. In the following, lower-cased variables refer to the EXBI-C, and upper-cased variables refer to the EXBI.

Monthly total returns on each bond incorporate both price changes and coupon payments. Specifically, for bond i in month t, the nominal return is defined as

$$r_{i,t} = \frac{p_{i,t} + c_{i,t} - p_{i,t-1}}{p_{i,t-1}},$$

where $p_{i,t}$ denotes the end-of-month bond price, $p_{i,t-1}$ is the price in the previous month, and $c_{i,t}$ is the coupon (or partial payment) received during month t. Nominal returns are converted to real returns using the relevant consumer price index for the United Kingdom or the United States. Coupons and prices are adjusted for missed or partial payments during defaults and restructurings, using the authors' detailed database on sovereign debt crises. This results in realized ex-post total returns that reflect actual investor experience, inclusive of price changes, coupon income, and default-related losses or recoveries.

We compute global monthly value-weighted yields of all external sovereign bonds i traded in month t as

$$Y_t = \sum_{i=1}^{N} y_{i,t} * \frac{w_{i,t}}{\sum_{i=1}^{N} w_{i,t}}, \tag{2.1}$$

where $w_{i,t}$ is the value-weight of bond i. We use as weight the market capitalization quantified by the product of amount issued$_{i,t}$ and $p_{i,t-1}$.

Similarly, returns are defined as

$$R_t = \sum_{i=1}^{N} r_{i,t} * \frac{w_{i,t-1}}{\sum_{i=1}^{N} w_{i,t-1}}. \tag{2.2}$$

For returns, weights are based on the market capitalization at the end of the previous month. Country returns and yields are defined analogously to 2.1 and 2.2. We deflate yields and returns of the bonds, which are originally quoted in nominal USD, using the US CPI.

Table 1: Summary statistics (1822–2022)

	Mean	SD	Percentiles				
			5th	25th	50th	75th	95th
Panel A: Global bond return index, EXBI (yearly)							
Real log returns	4.38	10.19	-16.50	-0.27	5.28	9.04	20.79
Real yields	6.11	5.40	-1.12	3.27	5.65	8.79	14.91
Panel B: Country-level bond return index, EXBI-C (yearly)							
Real log returns	6.28	27.02	-31.70	-2.55	4.54	12.68	51.50
Real yields	14.39	40.57	-4.53	2.46	5.33	11.25	56.70
Panel C: Country-level bond return index, EXBI-C (monthly)							
Real log returns	0.56	6.93	-7.89	-0.91	0.41	2.08	9.39
Real yields	14.33	39.57	-4.83	2.53	5.29	11.30	56.41

Note: Descriptive statistics for real log returns and yields in annualized percent, except for Panel C returns, which are in monthly percent.

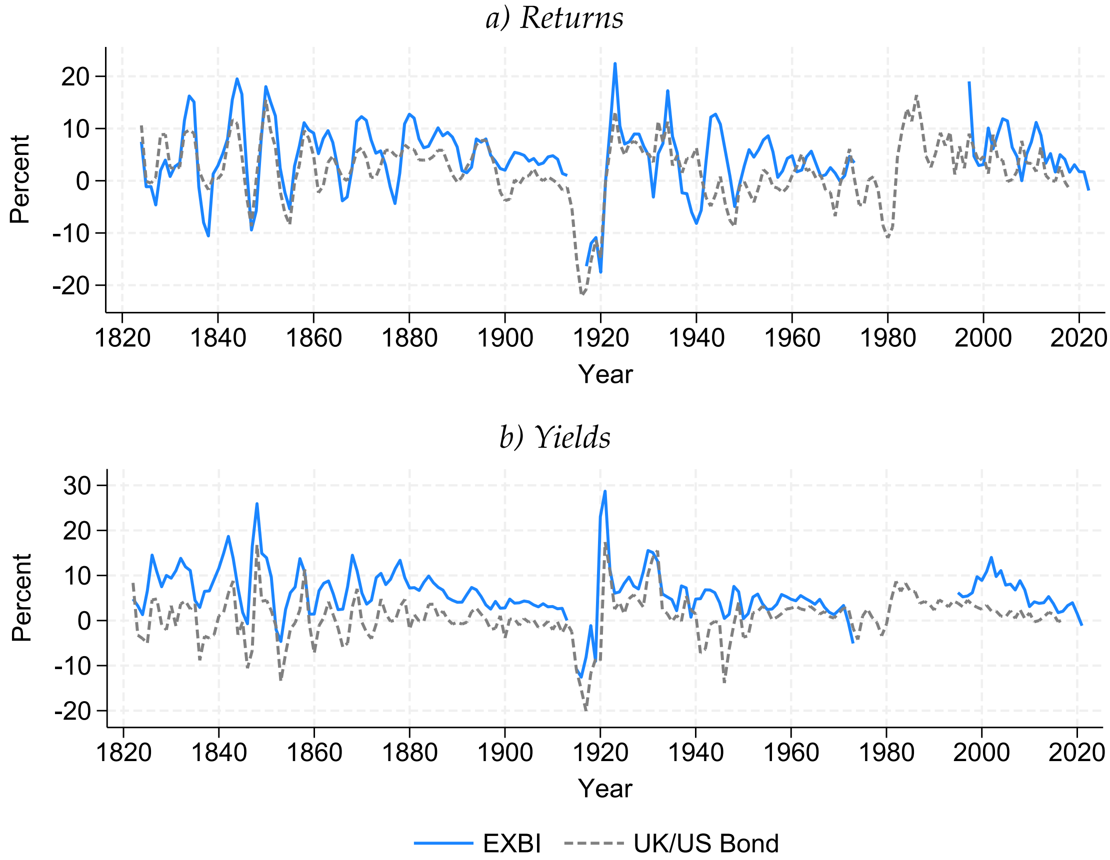
Table 1 summarizes the statistical properties of the EXBI yields and portfolio returns for the period 1822–2022. Panels A and B show summary statistics for annual returns and yields of the EXBI and the EXBI-C, respectively. Panel C shows statistics for monthly returns and yields, which we use when we turn to our analysis of threats in Section 4.

As can be seen in Panel A, annual returns average 4.4 percent. The standard deviation of returns is 10 percent per year. Panel B shows that the standard deviation of returns and yields is significantly higher for country-level returns, at 27 percent and 41 percent per year, respectively. Lastly, Panel C shows monthly returns and annualized monthly yields at the country level. The monthly log returns stand at about 0.6 percent and have a standard deviation of roughly 7 percent. The annualized monthly yields are about 14 percent. These numbers are in line with the annual observations outlined in Panel B.

Figure 1 plots the annual value-weighted mean real log return (Panel a) and real yield (Panel b) of the EXBI, together with long-term UK/US government bond counterparts.[6] The series highlights how external sovereign bond markets have been shaped by major historical events. Returns exhibit relatively low volatility during the first age of globalization, just before 1914. Volatility then

[6]Until 1917, the grey dashed line depicts UK bond returns and yields. From 1918 onwards, it shows US bond returns and yields.

Figure 1: EXBI Returns and Yields (real)

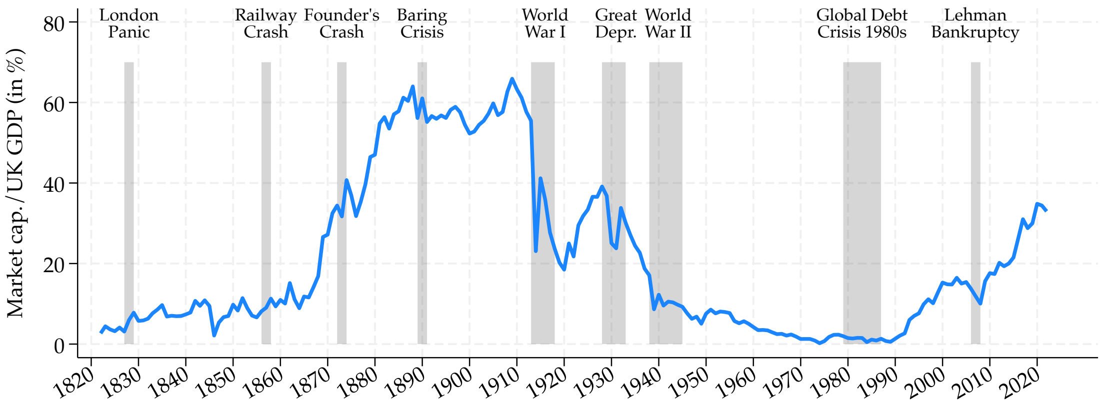


Note: Figure shows returns (Panel a) and yields (Panel b) of external sovereign bond index. Inflation/deflation dynamics can change the correlation between real yields and real returns, e.g. after WW1.

rises sharply in the interwar period, with large drawdowns around World War I and the Great Depression. In contrast, the post-2010 period looks more like the pre-1914 era, with relatively muted swings in both returns and yields.

One advantage of the EXBI relative to stock market data is that the underlying bond market operates almost continuously: while sovereign issuance goes through booms and busts, trading in external bonds persists even during major wars, when many equity markets close or malfunction. This continuity allows us to track how investors priced risk around episodes such as the World Wars, the 1930s wave of sovereign defaults, and the Latin American debt crisis, periods for which reliable and liquid equity data are often difficult to obtain. Annual real log returns on the EXBI are 37% correlated with the GFD Emerging Markets Stock Index in the overlapping 1921–2022 period, and similarly correlated with the GFD US stock-market index; correlations in nominal terms are slightly higher.

Figure 2: External Sovereign Bonds: Market Size



Note: Figure shows monthly aggregate market capitalization for more than 300,000 monthly bond price observations for about 2,150 foreign-currency sovereign bonds traded in London and New York between 1822 and 2022, spanning more than 90 countries. The database includes British-pound and US-dollar denominated government bonds with fixed coupons.

Table 2: Countries in EXBI over time

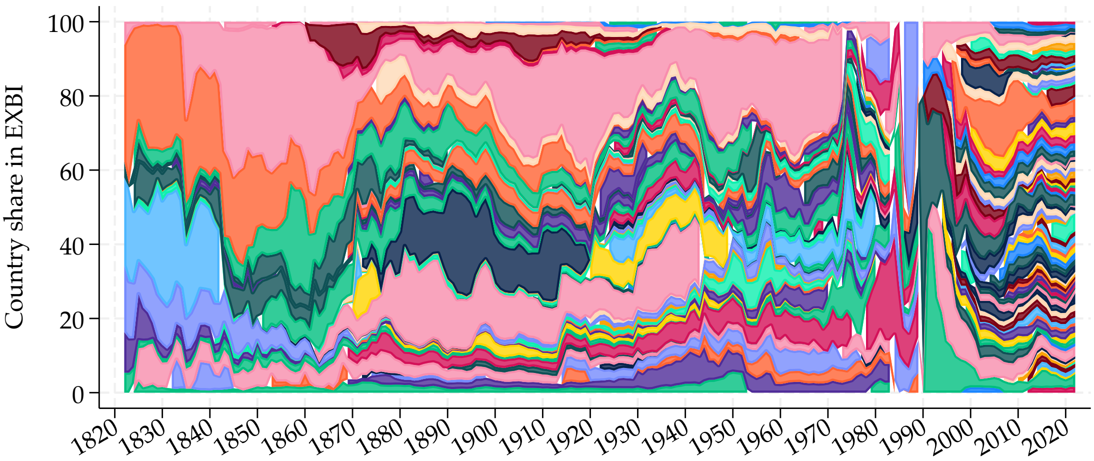
	Number of countries					
Region	1830	1870	1900	1940	1980	2020
East Asia & Pacific	0	3	4	5	6	7
Europe & Central Asia	6	11	11	18	12	18
Latin America & Caribbean	7	12	14	13	11	22
Middle East & North Africa	0	2	1	1	0	11
North America	0	1	1	1	1	0
South Asia	0	0	0	0	0	2
Sub-Saharan Africa	0	0	1	1	2	14
Total	13	29	32	39	32	74

Note: Table shows number of countries in EXBI by World Bank region at selected years.

Figure 2 shows the total market size for external bonds, expressed in percent of UK GDP. Between 1880 and 1910, the market is steady at approximately 60 percent of UK GDP. The market collapses during World War I, and continues to decline in size until the 1970s. Beginning in the 1990s, the market recovers with significant issuance. Despite being a relatively small market in the middle decades of the 20th century, there are still over 50 bonds trading every year.

Table 2 reports the number of countries represented in the EXBI by region at selected benchmark years. In the early period, the index is small and geographi-

Figure 3: EXBI Composition



Note: Figure shows value-weighted share of countries in EXBI over time.

cally concentrated: in 1830 it comprises only 13 countries, located exclusively in Europe and Latin America. Coverage then broadens over the nineteenth century, reaching 29 countries by 1870 and 32 by 1900, as East Asian, Middle Eastern, North American, and a first Sub-Saharan African issuer enter the market. The count peaks at 39 countries in 1940, driven largely by an expansion of European issuers. Reflecting the contraction of the external bond market in the mid-twentieth century—and the shift toward syndicated bank lending in the 1970s and 1980s—coverage falls back to 32 countries by 1980. The modern era brings a pronounced recovery and diversification: by 2020 the index spans 74 countries, with the largest gains in regions that were previously thinly represented, including Sub-Saharan Africa, the Middle East, North Africa, and South Asia, alongside a renewed and broad presence in Latin America. North America, by contrast, drops out of the sample entirely by 2020, as Canada no longer issues the foreign-currency external bonds that define the index.

Figure 3 shows the evolution of the EXBI's country composition by year. Composition refers to monthly country weights analogous to those used for weighting yields and returns. In the early period, before 1870, the index consists mainly of bonds from Spain, Russia, and Prussia. Spanish bonds account for a particularly large share and remain prominent until the 1970s. From the interwar period through around 1970, the EXBI becomes more dispersed across countries. In the modern period, the largest weights are held by Saudi Arabia, Mexico, Indonesia, and Turkey.[7]

There is also significant variation over time in the average remaining maturity. It is longer in the pre-war years, but falls below 20 years on average in the modern era; see also Figure A.1 in the Online Appendix.

3 Wars and bond returns

3.1 Data on wars

Our analysis of wars is based on war sites, defined as countries experiencing military destruction on their own territory, including all *interstate* and *other* wars documented by the Correlates of War Project (Sarkees and Wayman, 2010; Dixon and Sarkees, 2015) from 1870 to 2007, and by the Uppsala Conflict Data Program (UCDP) from 2008 onward (Gleditsch et al., 2002; Davies, Pettersson and Öberg, 2022). Interstate wars refer to conflicts between sovereign states, whereas other wars comprise intrastate wars (e.g., civil wars) and extrastate wars (including wars of independence). We source data on war sites from Federle et al. (2026), who provide geolocated, battle-level data on casualties for all wars listed in the above datasets. Casualties are defined as the total number of dead, missing, wounded, or prisoners of war captured on a given country's territory.[8]

Table 3 summarizes the war sites in our sample. In the combined sample of interstate and other wars, we observe 694 war site onsets at the country-year level; for 188 of these, contemporaneous bond return data are available. On average, war sites incur casualties of roughly 1 percent of the local population and experience fighting for about three years. Approximately one-third of the wars in our sample are interstate wars, with the remainder classified as other wars.

3.2 Time-series analysis: Global wars and global returns

Our first analysis is to study the impact of war incidence on the overall sovereign bond market through a time-series analysis.

[7] Table A.1 in the Appendix provides a detailed overview of the EXBI weights and market capitalization by country in 2022.

[8] Throughout, the ratio of casualties to the local population is winsorized at the 95% level; see A.3 in the Online Appendix.

Table 3: War sites in sample

	Wars	Casualties / population (in %)		Length	
	Total	Mean	Median	Mean	Median
Panel A: All wars					
Full sample	694	1.65	0.15	3	2
Bond sample	188	1.08	0.11	3	2
Panel B: Interstate wars only					
Full sample	225	3.50	0.26	4	3
Bond sample	61	2.20	0.26	4	3
Panel C: Other wars only					
Full sample	469	0.74	0.13	3	2
Bond sample	127	0.55	0.09	2	2

Note: Table shows properties for all war sites. "Total" refers to total number of war sites; "Bond sample" refers to total number of war sites for which bond returns are available in year of war onset; mean and median casualties (number of dead, missing, wounded, or prisoners of war) denoted in percent of pre-war population; mean and median length of wars denoted in years. Numbers of casualties and length refer to restricted sample where bond return data is available. Ratio of casualties to the local population is winsorized at the 95% level; see A.3 in the Online Appendix.

3.2.1 Variable construction and empirical framework

We follow the methodology of Federle et al. (2026) in defining the severity of wars. If $Wars_t$ is the set of all countries becoming war sites in t, the global impact of war can be characterized by three terms. First, the weight of all countries becoming war sites in t, which we define as the pre-war ratio of their population relative to the world population:

$$Weight_t = \sum_{i \in Wars_t} \frac{Pop_{i,t-1}}{WorldPop_{t-1}} . \tag{3.1}$$

Second, the severity of all wars starting in year t, defined, for each war, as the casualties incurred over the course of the war starting on the soil of i in year t, relative to i's local pre-war population:

$$Severity_t = \sum_{i \in Wars_t} \frac{Casualties_{i,t}}{Pop_{i,t-1}} . \tag{3.2}$$

And, lastly, the interaction of both terms, which forms a Global War Index (GWI) capturing the global impact of all war onsets in a given year t

$$GWI_t = \sum_{i \in Wars_t} \frac{Casualties_{i,t}}{WorldPop_{t-1}} . \tag{3.3}$$

Equipped with these variables, we can trace out the dynamic effects of global wars over the k years following the global war shock on the EXBI using a simple set of local projections of the following form:

$$\begin{aligned} x_{t+k} - x_{t-1} = \alpha_k + \beta_k \times Severity_t + \gamma_k \times Weight_t + \delta_k \times GWI_t \\ + \zeta_k' Z_t + \varepsilon_{t+k}, \end{aligned} \tag{3.4}$$

where x_t is either the cumulative log return series of the EXBI or the global economic crises tally in t. The crises tally counts the number of ongoing currency, inflation, banking, external and domestic sovereign debt crises as well as stock market crashes and is sourced from Reinhart and Rogoff (2009). $Severity_t$, $Weight_t$, and GWI_t are, as defined above, the aggregate severity of all wars, the weight of the countries that are host to the conflicts, and the interaction of these terms forming the GWI. As we do at the country level, we winsorize these variables at the 95% level to alleviate the influence of particularly severe episodes. [9] The controls, Z_t, comprise three lags of x_t in first differences as well as three lags of the explanatory variables. Lastly, ε_t is the error term. Standard errors are robust to heteroscedasticity and serial correlation up to k lags, with k at least 3 (Newey and West, 1986).

Our war-shock measure is constructed using realized casualties over the full duration of wars that begin in year t. The resulting estimates should therefore not be interpreted as implying predictable excess returns or implementable trading strategies. At the time of war onset, investors do not observe the eventual scale or duration of the conflict. Rather, the impulse responses trace the ex post dynamics associated with wars that ultimately turn out to be more severe.

In the analyses that follow, we report responses to a one-standard-deviation global war shock, which corresponds to casualties of roughly two basis points of the world population and is allocated to weight and severity in equal parts. To put this magnitude in perspective, it is comparable to the casualties—again measured relative to world population—incurred on Vietnamese soil over the

[9] Absent winsorization, responses of returns are robust and about twice as large as our baseline estimates. We flag significant influence of the World Wars for the unwinsorized sample, see Figure B.1 in the Appendix.

course of the Vietnam War, and somewhat below those of the Korean War, which on its own accounted for about three basis points of the world population.

3.2.2 Descriptive statistics and results

Figure 4 provides some descriptive statistics on the GWI. The top panel of Figure 4 plots the GWI—casualties relative to world population in basis points—as a 3-year moving average on the red solid line (left y-axis), alongside the global economic crises tally on the blue dashed line (right y-axis) that counts, at each point in time, the number of ongoing currency, inflation, and banking crises, external and domestic sovereign debt crises, and stock market crashes. A clear pattern emerges as the two lines move closely together: spikes in the war series tend to be followed by run-ups in the crises tally, which often remains elevated for some time even as war intensity recedes. These dynamics are particularly pronounced during the two World Wars and the 1960s. Taken at face value, the evidence is consistent with war triggering macro-financial distress at the global level.

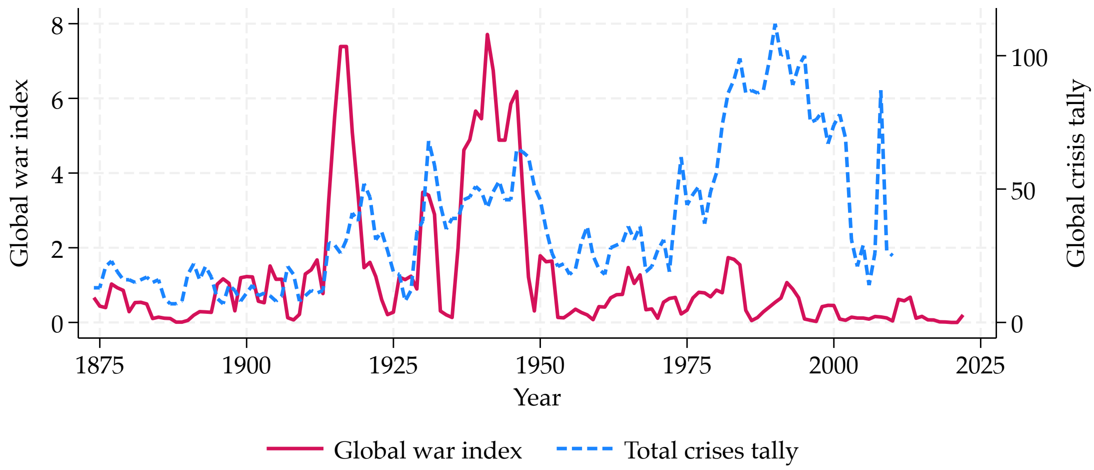
The bottom panel compares the index against three widely used newspaper-based measures of geopolitical risk. Each subpanel plots a binned scatter of our GWI (y-axis) against, respectively, the threat and acts indices of Caldara and Iacoviello (2022) and the war discourse index of Hirshleifer, Mai and Pukthuanthong (2023), all in z-scores. The figures suggest a strong positive relationship between the GWI and other measurements of geopolitical risk. At the same time, the series are far from perfectly aligned, a feature we examine in more detail below.

Motivated by the strong co-movement between the GWI and the global crises tally in the top panel of Figure 4, we now turn to our empirical analysis of how global wars trigger economic crises: The left panel of Figure 5 plots the impulse responses of the global crises tally to a one-standard-deviation global war shock (casualties of 2 basis points of the world population), as yielded by estimating regression (3.4) above. The responses show that global wars have sizable and persistent effects on the number of crises worldwide, with a one-standard-deviation war shock associated with roughly three additional crises at the global scale.

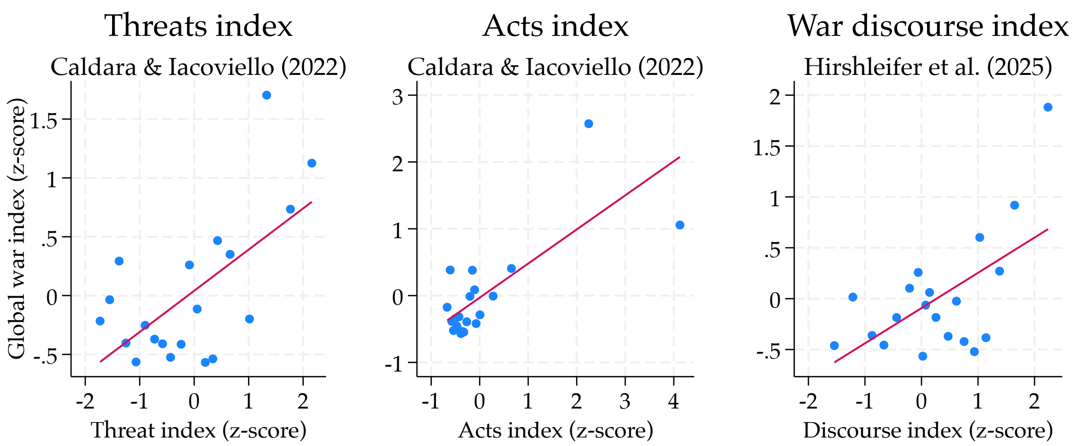
The right panel of Figure 5 provides a complementary perspective, showing cumulative global bond returns in response to war onsets. A one-standard-deviation war shock is associated with a decline in global bond returns of about 2 percentage points on impact and more than 5 percentage points at the trough

Figure 4: Global war index (GWI)

a) GWI and the global economic crises tally

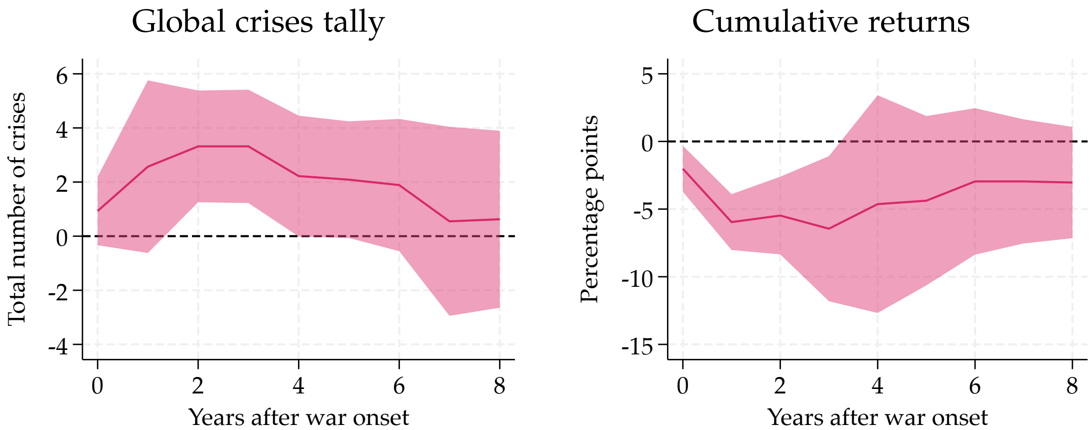


b) Comparing GWI to other indices of geopolitical risk



Note: Top panel a) plots 3-year moving average of GWI (casualties of all wars starting in a given year relative to world population) in basis points on left y-axis (solid red line) and total number of ongoing crises (currency, inflation, and banking crises, external and domestic sovereign-debt crises, and stock-market crashes; see Reinhart and Rogoff (2009)) on right y-axis (dashed blue line) against years on x-axis. Bottom panel b) shows binned scatter plots comparing GWI derived from Equation (3.3) to newspaper-based risk measures of Caldara and Iacoviello (2022) and Hirshleifer, Mai and Pukthuanthong (2023).

Figure 5: Response of global returns and economic crises to a 1 SD war shock (Casualties/World population ≈ 2bps)



Note: Figure depicts results of estimating Equation (3.4), showing impulse responses to a one-standard-deviation global war shock (casualties/world population ≈ 2bps; allocation to weight and severity in equal parts). Left panel shows response of global crises tally (total of ongoing currency, inflation, and banking crises, external and domestic sovereign-debt crises, and stock-market crashes; see Reinhart and Rogoff (2009)); Right panel shows response of cumulative log return series in percentage points. Shaded area denotes 90% confidence interval. Standard errors robust to heteroscedasticity and serial correlation (Newey and West, 1986).

in the following years.[10] As in the case of global crises, bond prices remain depressed for several years after war onset, although statistically significant effects persist only up to the third year.

It follows that wars have a strong negative effect on global bond returns. This is particularly important because, from a global perspective, wars are frequent events: In every single year since 1870, wars have been fought, and in more than 90% of years, a war *started* somewhere. Thus, while it is fair to say that wars are rare disasters from the perspective of any given country, at the global level, they are a regular driver of variation in returns.

The high explanatory power of wars for bond returns becomes apparent when we compare war shocks to other forms of crises in Table 4. The table reports the response of cumulative returns in the year of the shock ($k = 0$, first six columns) and over the three years following the shock ($k = 3$, last six columns). Within each time horizon, the first column shows the effects of a one-standard-deviation GWI shock; the second those of a recession onset in the financial center; the third those of a stock-market crash onset in the financial

[10]Our global war measure is forward-looking: it aggregates all casualties incurred over the full duration of wars *starting* in year t. Its explanatory power for future returns does not imply return predictability.

Table 4: Global bond returns, wars, and crises

	Cumulative EXBI log return ($R_{t+k} - R_{t-1}$)											
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
	$k = 0$						$k = 3$					
GWI (1 SD)	-3.075**					-2.551*	-8.858**					-7.941**
	(1.191)					(1.377)	(3.904)					(3.905)
Center recession		-4.198*						-2.129				
		(2.355)						(4.124)				
Center stock crash			-7.202***						-10.08			
			(2.364)						(7.258)			
ΔGlobal crises tally				-0.397***		-0.335***				-0.435**		-0.460***
				(0.0905)		(0.0840)				(0.177)		(0.144)
Commodity returns					0.219**	0.102					-0.0490	-0.234
					(0.0973)	(0.0976)					(0.121)	(0.142)
Constant	4.832***	4.963***	4.949***	4.174***	4.309***	4.542***	20.07***	16.78***	17.58***	16.63***	15.84***	19.83***
	(0.794)	(0.954)	(0.939)	(0.815)	(0.911)	(0.699)	(2.369)	(2.871)	(3.088)	(3.022)	(3.130)	(2.539)
R^2	0.050	0.026	0.041	0.241	0.096	0.296	0.130	0.001	0.016	0.062	0.001	0.195
Observations	128	126	116	116	112	104	122	122	113	113	110	102

Note: Table shows results of estimating regression (3.4). No lagged controls included to facilitate interpretation of share of variation explained (R^2). Standard errors are reported in parentheses and robust to heteroscedasticity and serial correlation (Newey and West, 1986). ***$p < 0.01$, **$p < 0.05$, *$p < 0.1$.

center; the fourth those of the year-on-year difference in the global economic crises tally; the fifth those of returns on the *Economist* commodity price index; and the sixth includes global wars, the year-on-year difference in the crises tally, and commodity returns simultaneously.[11] On impact ($k = 0$), crises and wars have a significantly negative effect on global bond returns: a global war shock lowers returns by about 3 percentage points, while recession onsets and stock-market crash onsets in the financial center reduce returns by roughly 4 and 7 percentage points, respectively. In turn, each additional crisis in the global crises tally is associated with a decline in bond returns of about 0.4 percentage points.

Commodity returns are positively associated with EXBI returns. We attribute this to the large number of emerging market economies in the index, for whom rising commodity prices may represent beneficial terms-of-trade shocks (see also Federle, 2026). Notably, including war and commodity returns jointly (column 6) leaves the war coefficient largely unchanged, indicating that the effect of war is not materially driven by commodity price movements.

In terms of fit, global wars alone explain about 5% of the variation in global bond returns, more than is explained by recessions in the financial center and similar to the share explained by center stock-market crashes.[12] The global crises tally accounts for the largest share of variation, explaining about 24% of global bond returns.

Turning to cumulative three-year returns, global wars explain more than 10% of the variation in returns. This far exceeds the variation explained by center recessions, stock-market crashes, or the change in the global crises tally over the same horizon. Our reading is that wars significantly shape short- to medium-run returns at the global level, accounting not only for severe collapses but also for a substantial share of overall variation.

Table 5 compares the GWI to several widely used measures of global (geopolitical) risk. The table reports the contemporaneous response of cumulative EXBI log returns to a one-standard-deviation war shock. Alongside the GWI, we sequentially add the yearly averages of the newspaper-based "acts" and "threats" indices of Caldara and Iacoviello (2022) and the war-discourse index of Hirshleifer, Mai and Pukthuanthong (2023), as well as a set of historical VIX proxies for the US and UK, and a spliced version of them.[13] In the specifications where

[11] We define the financial center as the UK until 1917, and the US from 1918 onwards.

[12] In contrast to the estimations above, the table includes no lagged controls to facilitate the interpretation of the R^2.

[13] As proxies for the volatility index, we build realized volatility measures by computing, for each period, the standard deviation of the UK and US stock market indices. These data are

Table 5: Global bond returns, war, and other geopolitical risk indices ($k = 0$)

	Cumulative EXBI log return ($R_{t+k} - R_{t-1}$)							
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
GWI (1 SD)	-3.325**						-3.305***	-2.826***
	(1.283)						(1.202)	(1.068)
Acts (Caldara & Iacoviello, 2022)		0.244						0.465
		(1.580)						(1.520)
Threats (Caldara & Iacoviello, 2022)		-2.283**						-1.983
		(1.018)						(2.148)
War discourse (Hirshleifer et al., 2025)			-2.029*					-0.296
			(1.193)					(2.487)
ΔVIX proxy, USA				-0.403***				
				(0.123)				
ΔVIX proxy, UK					-0.930***			
					(0.293)			
ΔVIX proxy, spliced						-0.398***	-0.421***	-0.444***
						(0.138)	(0.133)	(0.134)
Constant	4.469***	3.651***	4.473***	3.953***	4.006***	3.941***	4.308***	3.852***
	(0.946)	(0.979)	(0.918)	(1.068)	(1.051)	(1.068)	(0.888)	(1.384)
R^2	0.058	0.052	0.036	0.071	0.177	0.068	0.131	0.171
Observations	92	92	92	92	92	92	92	92

Note: Table shows results of estimating regression (3.4) for $k = 0$. No lagged controls included to facilitate interpretation of share of variation explained (R^2). Standard errors are reported in parentheses and robust to heteroscedasticity and serial correlation (Newey and West, 1986). ***$p < 0.01$, **$p < 0.05$, *$p < 0.1$.

they enter, changes in implied stock-market volatility are strongly negatively related to EXBI returns. However, the GWI retains substantial explanatory power even after we add these volatility measures: war intensity helps explain variation in global sovereign bond returns above and beyond what is captured by swings in global risk. Overall, the R^2 ranges from roughly 0.04 to 0.18, with the highest explanatory power in specifications that combine the GWI with VIX-type measures, reinforcing the view that wars perform well as an indicator of global bond returns when benchmarked against both text-based geopolitical risk indicators and standard market-based risk measures.

Across all specifications that include it, the coefficient on the GWI remains negative, economically large, and statistically significant: a one-standard-deviation global war shock is associated with an immediate decline in EXBI returns of roughly 3.5 percentage points, even once we control for alternative risk indices.[14] By contrast, the newspaper-based indices have a more mixed performance. The threats and war-discourse measures enter with negative and statistically significant coefficients, indicating that spikes in geopolitical tensions and war-related news are indeed associated with lower bond returns. However, their inclusion leaves the war coefficient largely unchanged, suggesting that the GWI captures a distinct component captured only in part by the newspaper-based measurements.

Robustness tests show that these findings are not driven by singular episodes, as we arrive at similar results when dropping the 5% of most intense years in terms of the GWI. We also arrive at consistent responses of returns even when restricting our sample to the subset of interstate wars that are narratively identified to have not been started for short-run economic developments. Conversely, the responses are not driven by these interstate wars alone, as they are robust to restricting the sample to intrastate and extrastate wars. We defer the robustness tests to Figures B.2–B.4 in the Online Appendix. We also show that our GWI performs similarly against other measures of risk in the medium run in Table B.1 in the Online Appendix.

A key identification concern is that war onsets need not be uniformly exogenous to bond returns—or, more broadly, to short-run developments of the global and domestic economies.[15] While our baseline analysis uses the full

sourced from Schwert (1990), Thomas and Dimsdale (2017), and Thomson Reuters.

[14] Note that the regressions are estimated on a restricted subsample from 1901–2016 where data on all indices is available.

[15] For example, rising fuel prices contributed to the ongoing Haitian crisis, see also "Haiti crisis deepens amid fuel protests and gang warfare", Financial Times, 29 September 2022.

sample (interstate and other wars) to document the general patterns, we show in our robustness that our findings on how wars affect returns persist when we restrict our analysis to a subsample of interstate wars that have been narratively identified as initiated for reasons orthogonal to short-run domestic economic conditions (Federle et al., 2026); see Figure B.3 in the Online Appendix.

3.3 Domestic wars and country-level returns

At the country level, wars are rare but catastrophic events. In our sample, the probability that a country becomes a war site is only about 3% per year, yet wars account for roughly one quarter of all bond collapses. This section studies these rare disasters using a panel local-projection framework.

3.3.1 Variable construction and empirical framework

To estimate the effects of wars that take place in a given country, we follow the same methodology as for global wars. The severity of the conflict at the war site is defined analogously as

$$Severity_{i,t} = \frac{Casualties_{i,t}}{Pop_{i,t-1}}, \tag{3.5}$$

where $Casualties_{i,t}$ is the total number of casualties incurred over the entire duration of all wars starting on the soil of country i in year t (0 otherwise), and $Pop_{i,t}$ is the population of country i in year t.

We trace out the dynamic effects of war by estimating a simple set of local projections of the following form:

$$x_{i,t+k} = \alpha_k + \beta_k \times Severity_{i,t} + \zeta_k' Z_{i,t} + \varepsilon_{i,t+k}, \tag{3.6}$$

where $x_{i,t+k}$ is either: the log return of country i in year $t+k$ or the long-difference in the log yield from $t-1$ to $t+k$. $Z_{i,t}$ denotes the controls, which comprise three lags of the outcome and the war-shock variable. We compute standard errors that are robust to heteroskedasticity as well as serial and cross-sectional correlation (Driscoll and Kraay, 1998).

For binary variables, such as realized defaults and crises dummies, we estimate the logit analogue of Equation (3.6) with the binary variable in $t+k$ as the dependent variable and the domestic war shock as the key regressor, controlling for three lags of both the binary outcome variable and the key regressor. We cluster standard errors at the country level. Throughout, we show the average marginal effects of a war shock on the binary outcome variables. As above, the war onset uses realized cumulative casualties. The resulting impulse responses therefore characterize the dynamics associated with wars ex post rather than predictable return patterns observable to investors at the onset of conflict.

3.3.2 Descriptive statistics and results

Figure 6 shows a few stylized facts about bond returns in the context of war onsets. The top panel shows the average year-on-year log returns of war sites around war onsets in an event window spanning 11 years. The event study shows no discernible pre-trend in the years leading up to the war onset. On impact, however, average returns drop sharply, by about 6% across all wars. For severe wars, defined as those with casualties greater than 1 percent of the local pre-war population, the average decline is almost 20%.

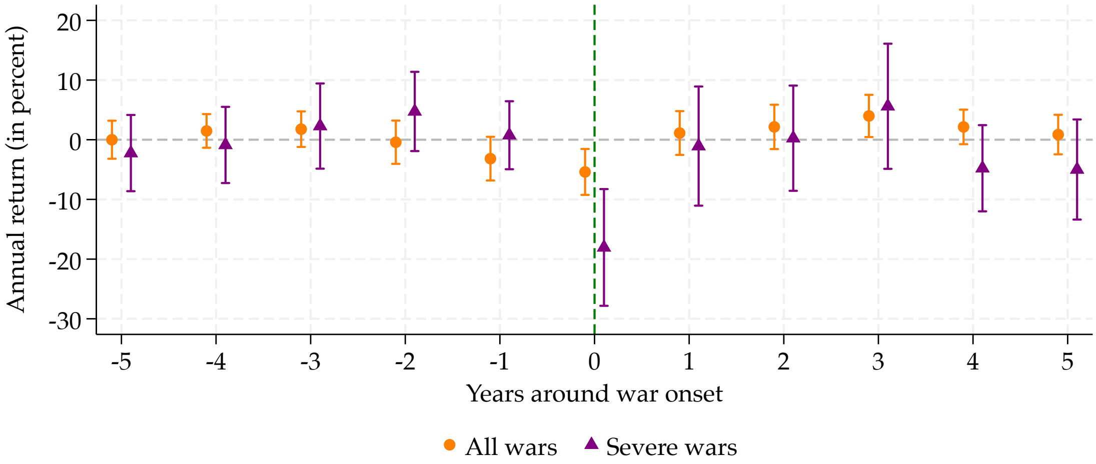
The bottom panels of Figure 6 show the probability and distribution of bond collapses around war onset. We define a collapse, following Mishkin and White (2002), as an annual return below -20%. The bottom-left panel shows a markedly higher collapse probability in onset years than in other years: in the year of onset, war-site countries face a collapse probability above 20%—nearly four times the rate observed in other years. The bottom-right panel documents that wars on a country's territory account for a disproportionate share of extreme losses: almost one quarter of all collapses in our sample occur within five years following a war onset on a country's own soil. This number is particularly high in light of the low probability of becoming a war site (3% annually).

Table 6 compares the effects of wars with those of typical domestic economic crises on country-level sovereign bond returns. In column (1), we include only a war-site dummy indicating whether a war starts on a country's own territory. The estimates imply that war onsets are significantly associated with a decline in returns of 12 percentage points. In column (2), we replace the dummy with our quantitative war-shock measure. A one-standard-deviation increase in war severity—corresponding to casualties amounting to 1.6% of the local population—leads to a drawdown in returns of about 9 percentage points.

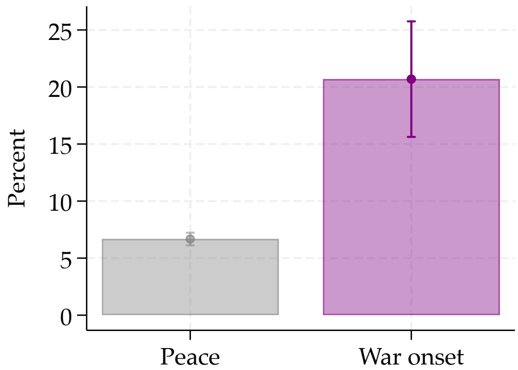
Columns (3), (4), and (5) report the response of returns to the onset of stock-market crashes, external debt crises, and banking crises, respectively. Among these, only external debt crises are associated with a statistically significant effect, with returns falling by roughly 19 percentage points. In column (6), we turn to the year-on-year change in the domestic economic crises tally—defined as the

Figure 6: Country-level bond returns and domestic wars

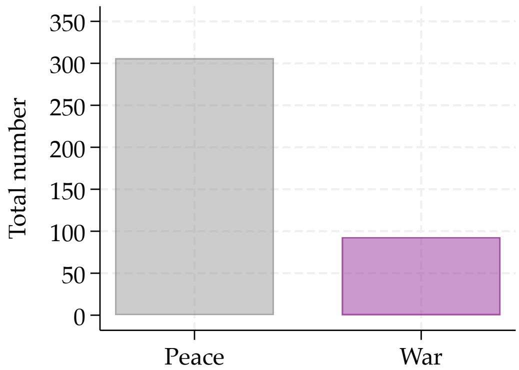
a) Average bond returns around war onset



b) Annual probability of bond collapse



c) Distribution of bond collapses



Note: Top panel a) shows average year-on-year log returns (y-axis) in years around war onset (x-axis). Orange circles denote averages around all wars and purple triangles denote averages around severe war onsets (casualties/local population $\geq$ 1ppt). Bars denote 90% (z-based) error bands. Bottom left panel b) shows annual probability of bond collapse conditional on war starting on own soil ("war onset") or not ("peace"). Bottom right panel (c) shows distribution of bond collapses associated with war (all collapses within five years after war onset) and peace. Bond collapses defined as instances with an annual return below -20% (Mishkin and White, 2002).

total number of ongoing currency, inflation, and banking crises, external and domestic sovereign debt crises, and stock-market crashes. Each additional crisis is associated with a reduction in returns of almost 5 percentage points. Lastly, column (7) includes both the domestic war shock and the crisis indicators. Both the war-shock variable and external debt crises remain statistically significant at the 5% and 1% levels, respectively.

Wars are rare, so it is not surprising that they only capture a modest amount

Table 6: Country-level bond returns, domestic wars, and economic crises

	EXBI-C log return ($r_{i,t}$)						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
War site (dummy)	-12.43***						
	(3.727)						
War shock (1 SD)		-9.012**					-12.16**
		(4.364)					(5.125)
Stock crash			1.425				0.551
			(3.696)				(3.347)
Ext. debt crisis				-19.00***			-14.96***
				(4.499)			(5.314)
Banking crisis					-5.942		-5.550
					(4.468)		(4.671)
ΔDomestic crises						-4.661***	
						(1.392)	
Constant	7.035***	6.799***	7.264***	7.641***	7.324***	7.048***	8.778***
	(2.318)	(2.284)	(2.374)	(2.367)	(2.364)	(2.373)	(2.695)
R^2	0.006	0.004	0.000	0.012	0.001	0.013	0.016
Observations	5,366	5,366	3,831	4,607	4,606	4,503	3,320

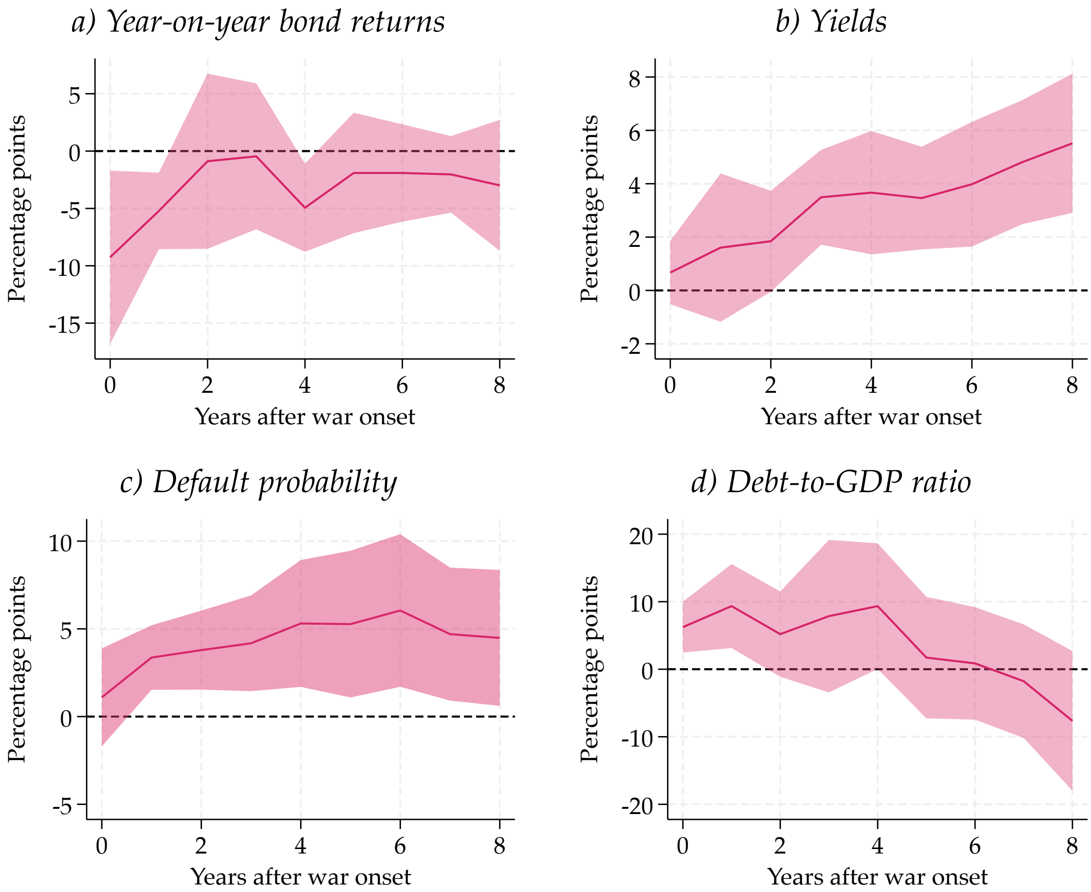
Note: Table shows results of estimating regression (3.6) for $k = 0$ (response of returns in year of war onset) and considers additional regressors (onsets of stock markets crash, external debt crises, banking crises and year-on-year difference of domestic economic crises tally; crises tally denotes total of ongoing currency, inflation, and banking crises, external and domestic sovereign-debt crises, and stock-market crashes; see Reinhart and Rogoff (2009)). No lagged controls included to facilitate interpretation of share of variation explained (R^2). Standard errors robust to heteroskedasticity as well as serial and cross-sectional correlation (Driscoll and Kraay, 1998).

of the variation in bond returns. However, across specifications, external debt crises and wars explain the largest portions of the variation, among the variables we consider.[16] This is consistent with the view that, at the local level, wars are rare and therefore only explain a tiny portion of the overall variation. However, when they occur, they are highly consequential disasters. This mirrors the bottom panels Figure 6 above, which suggests that a quarter of all bond collapses are associated with wars. In this way, even if wars explain little variation in overall returns, they explain a substantial share of bond collapses—which, by definition, happen to be rare themselves.

In Figure 7, we report the dynamic responses to domestic war onsets. The top-

[16]Note that Table 6 does not include lagged controls to facilitate the interpretation of the R^2. The local projections below outline the results of the more demanding specifications.

Figure 7: Country-level response to 1 SD domestic war shock (Casualties/Local population ≈ 1.6%)



Note: Figure depicts results of estimating Equation (3.6), showing impulse responses to a one-standard-deviation domestic war shock (casualties/local population ≈ 1.6%). Top-left panel shows response of year-on-year bond returns in percentage points; Top-right panel shows response of yields in percentage points. Bottom-left panel shows average marginal effect on probability of being in default in percentage points. Bottom-right panel shows response of debt-to-GDP ratio in percentage points. Shaded area denotes 90% confidence interval. Data on macroeconomic outcomes for up to 60 countries sourced from Federle et al. (2026). Standard errors robust to heteroskedasticity as well as serial and cross-sectional correlation (Driscoll and Kraay, 1998).

left panel shows how year-on-year bond returns react to a one-standard-deviation war shock. The profile features a sharp decline of returns in the first two years and somewhat attenuated returns thereafter.[17] In contrast, the top-right panel shows that yields rise following war onsets and remain elevated, in line with the

[17] As in the global analysis above, this persistence does not imply return predictability. Rather, it reflects the forward-looking nature of the war-onset shock, which assigns to year t the full casualties ultimately incurred over the entire course of conflicts that begin in t, scaled by the country's pre-war population, even though the full extent of the casualties are not known, and may not be anticipated, at the start of the conflict.

notion of financing becoming more costly in the wake of wars.

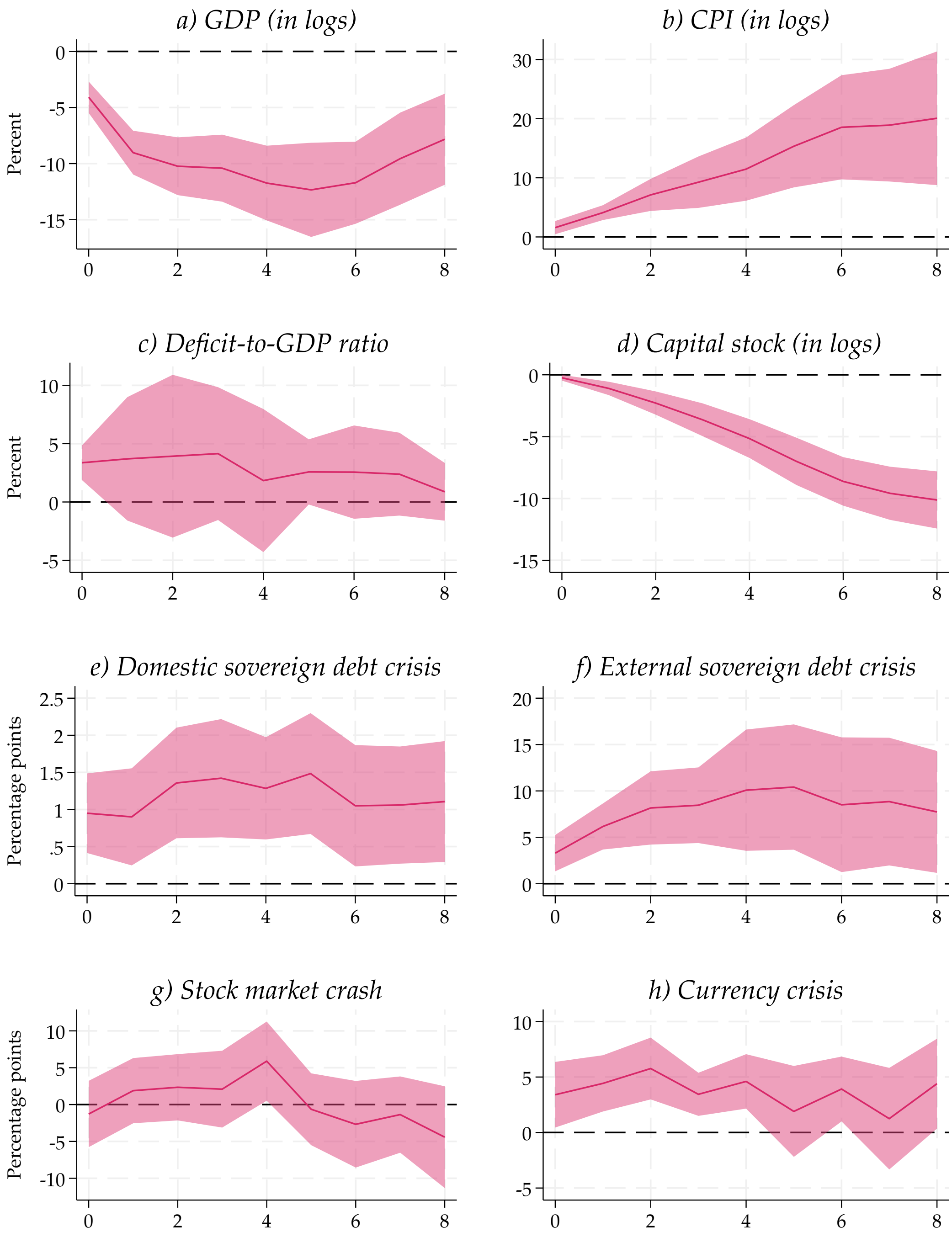
The bottom panels show how these market reactions relate to default risk. The bottom-left panel plots the average marginal effects from logit local projections in which the outcome is a forward default indicator and the regressor is the domestic war shock. A one-standard-deviation war shock raises the probability of being in default by about 7 percentage points in year 6. Even after 8 years, the default probability remains substantially elevated. The bottom-right panel shows a pronounced increase in a country's debt-to-GDP ratio following war onsets. On impact, a one-standard-deviation war shock increases the ratio by more than 5 percentage points. The effect rises gradually and peaks in the second year, by which time debt-to-GDP has risen by 10 percentage points on average. The response converges back in the fifth year following the war onset.

The increased debt-to-GDP ratio alludes to the channels driving the market response. Figure 8 provides a more systematic overview of the potential macroeconomic and financial drivers behind the sovereign risk during war. The figure highlights that wars are generalized macro-financial disasters that trigger multifaceted crisis dynamics. It again plots impulse responses to a one-standard-deviation domestic war shock. The top four panels report local-projection estimates from the linear specification in Equation (3.6), while the bottom four panels show average marginal effects from logit models for the probability of ending up in different types of crises.

The top row summarizes the real macroeconomic consequences of war. Panel (a) shows that GDP falls persistently by about 12 percent following a war shock, with no evidence of a quick rebound over the eight-year horizon. Panel (b) documents a pronounced and lasting increase in consumer prices of some 20%, consistent with war-time inflation and post-war inflationary pressures. In panel (c), the fiscal deficit widens markedly in the wake of the shock and remains elevated, indicating sustained fiscal stress as governments finance military operations and reconstruction. Panel (d) shows that the capital stock declines sharply as well and stays below its pre-war path, reflecting the destruction of physical assets and depressed investment.

The two bottom rows link these macro dislocations to episodes of open financial distress. While panel (e) shows that the probability of a domestic sovereign debt crisis rises substantially after war onset, a direct comparison with panel (f) reveals that the effect on external sovereign debt crises is almost 10 times higher. Lastly, panels (g) and (h) document elevated risks of stock-market crashes and currency crises, respectively. These responses are obtained from logit

Figure 8: Wars are macro-financial multi-crises
(Response to 1 SD domestic war shock: Casualties/Local population ≈ 1.6%)



Note: Panels a) - d) depict results of estimating Equation (3.6), showing impulse responses to a one-standard-deviation domestic war shock (casualties/local population ≈ 1.6%). Standard errors robust to heteroskedasticity as well as serial and cross-sectional correlation (Driscoll and Kraay, 1998). Panels e) - h) show logit analogues of Equation (3.6), and plotted lines represent average marginal effects of the war shock on the probability of various crises. Data on macroeconomic outcomes for up to 60 countries sourced from Federle et al. (2026). Standard errors clustered at the country level.

analogues of Equation (3.6), and the plotted lines represent the average marginal effects of the war shock on the probability of each crisis type. Taken together, the figure highlights that the bond markets' response to domestic wars is not associated with a single channel, but the result of broad-based macro-financial turmoil in which output collapses, inflation surges, fiscal positions deteriorate, and multiple crises types become more likely at the same time.

Robustness tests show that these findings hold beyond the most severe episodes. Even when we remove the 5% of most destructive events, in terms of war severity, our results remain remarkably similar, both in terms of returns and default probabilities; see Figure B.5 in the Online Appendix. We also arrive at the same patterns, irrespective of subsampling to the set of narratively identified interstate wars or other wars; see Figures B.7 and B.8 in the Online Appendix. Lastly, we show that our conclusions regarding the share of bond collapses by wars are robust to another definition requiring a negative return of 40% over two years (Greenwood, Shleifer and You, 2019); see Figure B.9 in the Online Appendix.

Just as in our global analysis above, identification relies on wars being started for reasons that are exogenous to bond returns. To this end, we refer first to the top panel of Figure 6 above, which suggests that wars are generally not to be anticipated by the capital markets. Additionally, we show in our robustness that the observed patterns persist when we restrict our analysis to a subsample of interstate wars that have been narratively identified as initiated for reasons orthogonal to short-run domestic economic conditions (Federle et al., 2026); see Figure B.7 in the Appendix. Naturally, however, confidence bands widen in these specifications, as we drop about two-thirds of the treated episodes.

It has been argued that the international bond market was less sensitive to political and military shocks in the decades before 1914, in part because of the gold standard and the increased financial integration at the time (Ferguson, 2006). We test whether the bond-market response to war differs across regimes by re-estimating Equation (3.6) on two subsamples, covering 1870–1914 and 1915–2022. Figure B.6 in the Appendix reports the results. Across both samples, wars significantly depress bond returns, albeit more strongly in the recent period. Conversely, the response of sovereign default is more pronounced in the earlier period, though the confidence bands widen.

We have also investigated whether the impact of war on bond returns is a function of pre-war fragility, as suggested in the multiple-equilibria framework of Calvo (1988). Table B.2 in the Appendix allows the response to war to vary with the pre-war level of yields in both continuous and threshold specifications. Point estimates on the interaction terms are small and statistically indistinguishable from zero. We interpret this as consistent with the multicrisis nature of domestic wars documented in Figure 8: the shock is sufficiently broad-based that its effect on returns is largely independent of pre-war conditions.

4 Military threats and bond returns

In this section, we turn our attention to the question of whether war threats are reflected in bond returns, and how quickly, and how effectively financial markets anticipate the consequences of war.

4.1 Data

To examine the effects of military threats, we draw on data on Militarized Interstate Disputes (MID) from the Correlates of War Project (Jones, Bremer and Singer, 1996; Maoz et al., 2019).[18] The dataset records 10,358 directed dyadic disputes between 1816 and 2014, where a "dyad" denotes a pair of states engaged in a militarized interaction.[19] Each dispute is coded by its highest hostility level in each year: (1) threat to use force, (2) display of force, (3) use of force—actual fighting short of 1,000 battle-related deaths—and (4) interstate war—hostilities with at least 1,000 battle-related deaths. The data provides an exact daily account for when disputes started and which country was the initiator. In our baseline classification, we group categories (1) and (2) as "threats," capturing coercive signaling absent direct combat, and treat categories (3) and (4) as realized force and war.

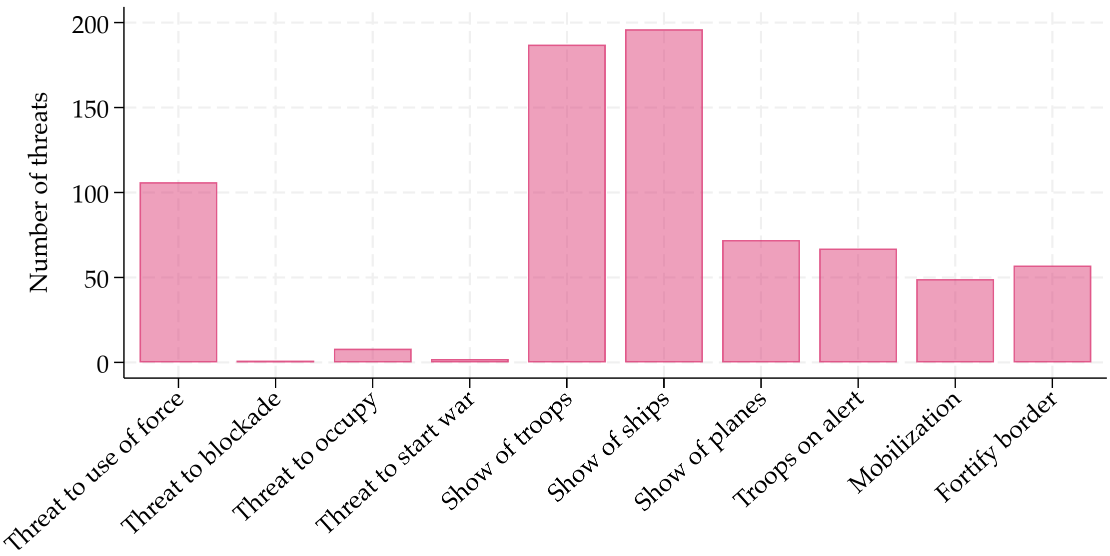
The MID framework requires threats to be explicit and to take the form of an ultimatum with a coercive component (Jones, Bremer and Singer, 1996). They are direct announcements to react with military force or war if the counterparty "fails to act, or does not refrain from acting in a specified manner." Diffuse warnings of "dire consequences" or "fire and fury" are not treated as threats (Jones, Bremer and Singer (1996) and Palmer et al. (2022)).

Figure 9 plots the distribution of militarized threats in the restricted sample

[18] The MID data has also been used in Hess and Orphanides (2001), Martin, Mayer and Thoenig (2008), Caselli, Morelli and Rohner (2015), Couttenier et al. (2019), and Federle, Rohner and Schularick (2026), among many others

[19] We drop two broader conflicts pertaining to the fall of the Milošević regime in Yugoslavia as these highly multilateralized conflicts alone exhibit more than 100 dyads.

Figure 9: Number of threats by type



Note: Figure shows number of country-level threats by threat type. Sample restricted to observations where bond return data in month of threat onset is available.

used for our bond-return analysis, i.e., disputes for which we have bond returns of the participating country in the month the MID begins. Threats are disaggregated into explicit threats and demonstrations of force. The figure indicates that the sample is weighted toward display- and mobilization-type actions, with comparatively fewer explicit threats to start war, blockade, or occupy.

Most threats are issued by Russia, the US, the UK, China, and France. The most frequently threatened countries are Russia, Turkey, the US, China, and Japan.[20] Both lists are dominated by major powers rather than small or peripheral states, reflecting these countries' central role in great-power rivalries.

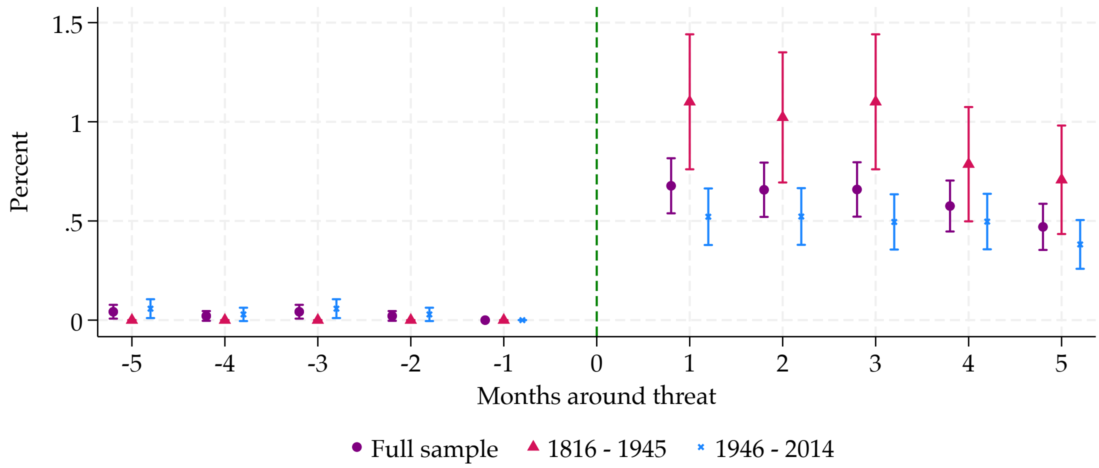
4.1.1 Threats predict conflict

Our analysis of how military threats affect sovereign bond returns assumes that threats work by altering the perceived probability of future war—especially the risk of becoming a war site—and thereby lowering expected future cash flows. If threats raise the likelihood of armed conflict, investors should price in higher default risk and lower bond values as soon as the threats are issued. We therefore begin by examining how threats map into subsequent war and escalation. This exercise is similar to Ferguson (2006), who shows that investors reacted strongly to wars and revolutions during the nineteenth century, but appeared increasingly

[20]We provide more details on the distribution of threats across countries in Figure C.1 in the Appendix.

insensitive to geopolitical crises after 1880, which he interprets as evidence that the outbreak of World War I was largely unanticipated by financial markets. Compared to Ferguson (2006) our data allow us to more systematically analyze the impact of explicit threats over a larger sample.

Figure 10: War onset probability around threats



Note: Figure shows average monthly probability of war onset (y-axis) in months around threats (x-axis) in percent for three samples: 1816–2014 (purple circles); 1816–1945 (red triangles); 1946–2014 (blue crosses). Vertical bars denote z-based 90% confidence intervals.

Figure 10 provides some initial evidence in this regard. It shows the monthly probability of any pair of countries in the world going to war around a bilateral threat for three periods: the full sample, a subsample between 1816 and 1945, and a subsample in the post-World War period. As can be inferred from the figure, the probability of going to war prior to a threat is almost equal to zero. After the threat, however, the probability rises by between 0.5 and 1 percentage points and remains substantially elevated in the months after, suggesting that a considerable number of wars are preceded by threats.

To examine the relationship between threats and wars in a more formal setting, we estimate a set of Poisson-Pseudo Maximum Likelihood regressions of the following form:

$$x_{i,j,t} = exp\left[\gamma_{i,j} + \eta_t + \delta Threat_{i,j,t-1}\right] + \varepsilon_{i,j,t} \quad (4.1)$$

where $x_{i,j,t}$ is either an undirected dummy indicating whether country i and j started going to war in month t (as defined by fighting in excess of 1,000 battle-related casualties) or started to use force against each other (as defined by fighting short of 1,000 battle-related casualties). The variables $\gamma_{i,j}$ and η_t denote

Table 7: Threats and the probability of participating in war

	Onset $hostility = 3$ (Use of force)		Onset $hostility = 4$ (War)	
$Threat_{i,j,t-1}$	1.307***	0.292**	2.540***	3.351*
	(0.323)	(0.117)	(0.892)	(1.916)
Dyad FE		✓		✓
Year FE		✓		✓
Observations	1,817,622	1,817,622	241,920	241,920

Note: Table shows average marginal effects obtained via estimating Equation (4.1). Dependent variable is either onset of force or onset of war in a given country-pair-month. Independent variable is lagged threat. ***$p < 0.01$, **$p < 0.05$, *$p < 0.1$.

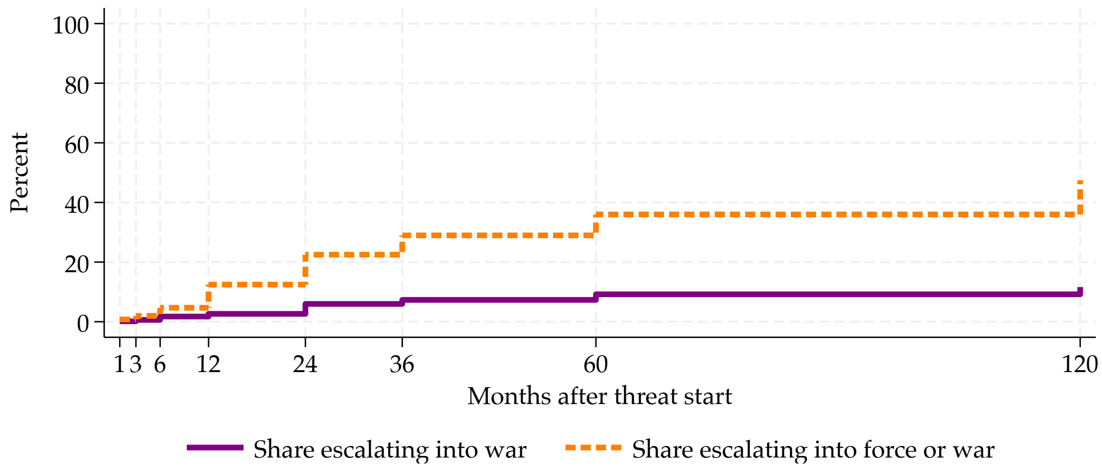
country-pair and time fixed effects, respectively. $Threat_{i,j,t}$ denotes whether there was an active threat between country i and country j in month t. Lastly, $\varepsilon_{i,j,t}$ is the error term. Standard errors are two-way clustered at country and year level.

Table 7 shows the average marginal effects resulting from estimating these regressions. Columns (1) and (2) show the effect of threat on the onset of force, and columns (3) and (4) show the effects on the onset of wars. Across all specifications, we find threats lead to a significant increase in violence. The most demanding specification in column (4) suggests that a threat increases the probability of war in the next month by 3.4 percentage points.

While threats can convert to wars, only a subset of them actually do. The share of threats converting into the use of force or war over time is shown in Figure 11. The solid purple line represents the cumulative fraction of threats that have turned into war in percent on the y-axis against the number of months that have passed after the threat was initiated on the x-axis. Even after 36 months, only about 5% of threats transform into full-blown wars. However, about 30% turn into some form of violent action, as shown by the orange dashed line.

When countries go to war, both sides count as participants, so participation is a symmetric measure of escalation. But when we focus on war sites—the places where fighting actually occurs—the pattern can be asymmetric, since military action often takes place on only one party's territory. For instance, aside from a few isolated attacks on overseas territories and Pearl Harbor, the United States has not experienced significant fighting on its own soil in the last 150 years, even though it has taken part in many wars. We now ask whether being the *initiator* or the *target* of a threat also affects the likelihood of becoming a war site. To

Figure 11: Share of threats escalating to military force or war



Note: Figure shows cumulative share of threats converting to war (solid purple line) or the use of force or war (dashed orange line) in percent (y-axis) in the months following a threat onset (x-axis).

study this, we estimate the following monadic PPML regression:

$$Site_{i,t} = exp\left[\alpha_i + \eta_t + \beta \times \mathbb{1}Threatening_{i,t-1} + \gamma \times \mathbb{1}Threatened_{i,t-1}\right] + \varepsilon_{i,t} \quad (4.2)$$

where $Site_{i,t}$ is a dummy indicating whether country i becomes a war site in year t, α_i and η_t denote country and year fixed effects, $\mathbb{1}\text{Threatening}_{i,t}$ indicates whether i is threatening at least one new country in year t, and $\mathbb{1}\text{Threatened}_{i,t}$ indicates whether i is threatened by at least one new country in year t.[21] $\varepsilon_{i,t}$ is the error term, and standard errors are two-way clustered at the country and year levels.

Table 8 reports the average marginal effects from estimating Equation (4.2). Column (1) reports the response to an unconditional threat, that is, without differentiating between the threatened and threatening country. The coefficient indicates that a new threat is associated with a 1.6 percentage point increase in the probability of becoming a war site in the following year. When we add country and year fixed effects in column (2), the implied effect rises to 2.8 percentage points. These effects are statistically significant at the 1% and 5% level, respectively. In columns (3) and (4), we allow the effect of threats to differ between the country that initiates the threat and the country that is threatened. In this specification, only the threatened country exhibits a statistically significant

[21] Data on war sites are only available at the yearly level. We run this regression on a balanced panel of all countries that became a war site at least once over the sample period. Reciprocal threats, where sender and receiver cannot be clearly separated, are excluded.

Table 8: Threats and the probability of becoming war site

	Becoming war site			
	(1)	(2)	(3)	(4)
Threat, unconditional	1.641***	2.811**		
	(0.598)	(1.244)		
Threatening			0.764	1.227
			(0.696)	(1.426)
Threatened			1.607**	2.541**
			(0.727)	(1.183)
Country FE		✓		✓
Year FE		✓		✓
Observations	13,968	7,275	13,968	7,275

Note: Table shows average marginal effects obtained via estimating Equation (4.2). Dependent variable is onset of war on own territory. Independent variable is lagged threat. $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$.

increase in the probability of becoming a war site.[22] In particular, being a victim of a threat is associated with an increased probability of becoming a war site in the following year of some 2.5 percentage points. The effect is significant at the 5% level.

To summarize, at a dyadic level, threats between two countries are significantly associated with subsequent wars between those countries. At the monadic level, we can show that there is heterogeneity across the sender and receiver of threats: threats raise the probability that the target becomes a war site, but we find no corresponding effect for the threatening country. These insights set the stage for our subsequent analysis: Because threats raise the probability of becoming a war site only for the threatened state, we expect them to lower bond returns more in threatened countries than in threatening ones.

4.1.2 Threats lead to bond price declines

In this section, we examine how threats affect bond returns and let the response differ across the threatening country that initiated the threat and the threatened country. For this purpose, we again estimate a set of local projections of the

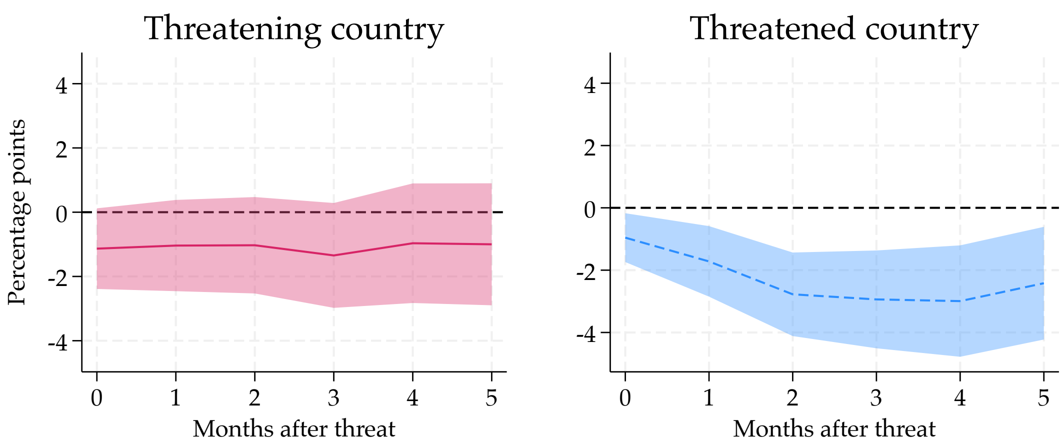
[22]This aligns broadly with studies showing that the initiators of *war* are less likely to be those who experience violence on their soil (Federle et al., 2026).

following form:

$$x_{i,t+k} - x_{i,t-1} = \alpha_k + \beta_k \times \mathbb{1}Threatening_{i,t} + \gamma_k \times \mathbb{1}Threatened_{i,t} + \delta_k' Z_{i,t} + \varepsilon_{i,t+k}, \tag{4.3}$$

where $x_{i,t}$ is the cumulative log-return series of country i in month t, $\mathbb{1}Threatening_{i,t}$ indicates whether i is threatening at least one new country in t and $\mathbb{1}Threatened_{i,t}$ indicates whether i is threatened by at least one new country in t. Variable $Z_{i,t}$ denotes the controls comprising three lags of returns and the explanatory variables. Lastly, $\varepsilon_{i,t+k}$ is the error term. Throughout, we compute standard errors that are robust to heteroskedasticity as well as serial and cross-sectional correlation (Driscoll and Kraay, 1998).

Figure 12: Response of returns in threatening and threatened countries



Note: Figure depicts results of estimating Equation (4.3), showing impulse responses of cumulative returns to threat onsets. Left panel shows response of threatening country (which initiates threat); Right panel shows response of threatened country. Shaded area denotes 90% confidence interval. Standard errors robust to heteroskedasticity as well as serial and cross-sectional correlation (Driscoll and Kraay, 1998).

Figure 12 shows the estimated response. The coefficients for the threatening country are shown on the left panel. As can be inferred from the figure, the threatening country does not exhibit any significant response in terms of returns to threats. In turn, the threatened country sees a substantial decline in returns of about 1 percentage point on impact, which culminates about 3 months after the threat with a negative cumulative return of about 3 percentage points.

Note that we already have estimates on how threats shift the war probability and on the loss given wars, which allows for a back-of-the-envelope calculation using a simple two-state risk-neutral asset pricing model: Let P be the price of a given bond, X its peacetime payoff, r the discount rate, ω the probability of war, and ζ the fractional loss given war. The price function is then given by

$$P = \frac{1}{1+r}\left[(1-\omega)X + \omega(1-\zeta)X\right] = \frac{X}{1+r}(1-\omega\zeta). \tag{4.4}$$

Differentiating with respect to the war probability ω yields

$$\frac{\partial P}{\partial \omega} = -\zeta\frac{X}{1+r}. \tag{4.5}$$

For small war probabilities[23] the proportional effect of the price change given a change in the war probability $\Delta\omega$ is

$$\frac{\Delta P}{P} = -\frac{\zeta}{1-\omega\zeta}\Delta\omega \approx -\zeta\Delta\omega. \tag{4.6}$$

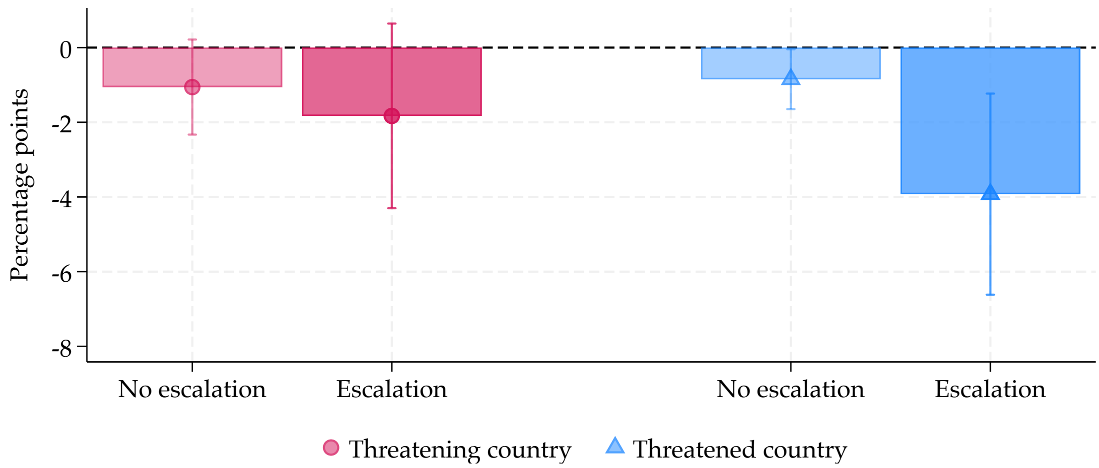
Thus, an increase in war probability lowers prices approximately in proportion to the product of the conditional loss and the probability shock. Empirically, our PPML estimates above suggest an average marginal effect of being threatened on the probability of becoming a war site of 2.5 percentage points, implying $\widehat{\Delta\omega}_{i,t} = 0.025$. Becoming a war site in an *interstate war* is on average associated with negative real returns of 28% in the first two years of the war, such that $\hat{\zeta} = 0.28$.[24] Plugging these values into (4.6) yields an expected market reaction in returns to threats of about -0.7%, which is right in the ballpark of the observed price response for threatened countries in the month that the threat is issued. In contrast, we cannot reject the null of $\widehat{\Delta\omega}_{i,t} = 0$ for threatening countries, providing a natural explanation for why we do not find significant changes to bond returns for these in response to threats.

While these calculations hold for the *average threat*, individual threats can differ along many dimensions, and some are arguably more credible or severe than others. In principle, markets should react more strongly to threats that are ex post revealed to be more serious, provided the market can assess heterogeneity in threats at the time they are issued. We test this hypothesis by re-estimating the specification above while distinguishing between new threats that escalate into military violence within three months of the initial threat and those that

[23]The unconditional probability of becoming a war site in our sample is 3% per year.

[24]The number resembles the coefficient obtained from a linear model relating the 2-year country-level bond return to a dummy indicating the onset of an interstate war on a country's own soil.

Figure 13: Market response to threats conditioned on ex post escalation



Note: Figure depicts results of estimating an extended specification of Equation (4.3), further differentiating between threats that escalated to higher forms of violence within 3 months and those that did not. It shows returns in month of threat in percentage points on the y-axis, differentiating both between threatening/threatened countries and escalation/no escalation.

do not. Figure 13 summarizes the results. It shows the market reaction in the month the threat is issued for both threatening and threatened countries, separately for threats that later escalate into the use of force or war and those that remain non-violent. In both groups of countries, returns respond more negatively to threats that are followed by higher forms of violence. However, the difference is statistically significant only for threatened countries: here, threats that escalate are associated with an average return decline of about 4 percentage points, compared with roughly 1 percentage point for threats that do not escalate. A two-sided Wald test rejects equality of the two responses at the 10 percent level (p-value = 0.071). We think of this as corroborating evidence of the threats in our sample triggering a market response via a revised assessment on the expected probability of domestic war.

A natural question is whether the asymmetry we document reflects differences in power between the two countries. The data suggest that it does. Across all threat onsets in our sample, economically larger countries are substantially more likely to initiate threats than smaller countries. In a logit specification, being the larger economy in the dyad increases the probability of issuing the threat by roughly 17.5 percentage points. The result is unchanged when we replace the size dummy with the log GDP difference between countries. Although smaller states sometimes threaten larger ones, military threats predominantly flow from stronger to weaker economies. This pattern is consistent with the view that military threats operate as an instrument of geoeconomic coercion used by powerful states, helping explain why the financial costs fall primarily on the threatened rather than the threatening country.

Robustness test show that these findings are persistent across different episodes. Similar to our analysis of domestic wars, we examine whether the gold standard and the pre-1914 era's degree of financial integration shaped the bond-market response to threats (see also Ferguson, 2006). To this end, we re-estimate Equation (4.3) on two subsamples, covering 1822–1914 and 1915–2014, and document similar patterns; see Figure C.2 in the Appendix.

5 Conclusion

Wars are a persistent feature of the global economy, and sovereign debt markets price them accordingly. Using two centuries of data, we show that wars are a major driver of sovereign borrowing costs, investor losses, and default risk, both globally and at the country level. From the perspective of financial markets, wars are not isolated geopolitical events. They are macro-financial *multicrises* that simultaneously disrupt output, public finances, inflation, and trade.

While wars are rare for individual countries, they account for a disproportionate share of major bond market collapses. In our data, roughly one quarter of all severe sovereign bond crashes occur in the aftermath of war. These episodes leave persistent scars on sovereign balance sheets and financial markets long after the fighting begins.

We also show that financial markets respond not only to realized wars, but also to military threats. Explicit threats and displays of force depress sovereign bond prices in targeted countries, while leaving threatening countries largely unaffected. Military coercion therefore operates not only through direct force, but also through financial channels. Stronger states are able to impose economic costs on weaker rivals, even before conflict occurs.

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Online Appendix to
Wars, Threats, and the Sovereign Bond Market

May 2026

A Complementary descriptive statistics

Table A.1: EXBI composition in 2022 by country

Country	GDP (current USD bn)	Market value (current USD bn)	Weight (in %)	MV/GDP (in %)
Saudi Arabia	1,108.6	80.8	7.86	7.28
Mexico	1,465.3	77.3	7.53	5.28
Indonesia	1,318.7	76.9	7.48	5.83
Turkey	903.7	68.6	6.68	7.59
United Arab Emirates	502.7	57.4	5.59	11.41
Qatar	235.7	51.4	5.00	21.80
Russia	2,258.4	48.0	4.68	2.13
Brazil	1,950.8	43.8	4.26	2.24
Philippines	403.9	36.6	3.57	9.07
Colombia	345.4	29.4	2.86	8.51
Oman	109.9	28.6	2.79	26.05
Egypt	409.3	25.8	2.51	6.30
Peru	246.0	25.2	2.45	10.23
Dominican Republic	113.5	22.4	2.18	19.71
Bahrain	44.4	22.0	2.14	49.51
Argentina	631.9	21.6	2.10	3.41
China	17,888.2	20.8	2.03	0.12
South Africa	406.2	19.3	1.88	4.75
Panama	76.5	19.3	1.88	25.18
Chile	302.4	18.4	1.79	6.09
Uruguay	70.2	17.7	1.72	25.15
Ukraine	161.6	14.8	1.44	9.13
Nigeria	475.1	13.6	1.32	2.86
Hungary	177.3	12.0	1.16	6.74
Ecuador	116.6	11.8	1.15	10.09
Romania	296.1	10.4	1.01	3.50
Ghana	74.3	9.9	0.96	13.32
Poland	693.7	8.9	0.87	1.29

Continued on next page

Table A.1 – continued

Country	GDP (current USD bn)	Market value (current USD bn)	Weight (in %)	MV/GDP (in %)
Angola	113.3	7.5	0.73	6.64
Kazakhstan	225.5	7.3	0.71	3.24
Pakistan	325.3	7.2	0.70	2.22
Kenya	114.4	7.0	0.68	6.11
Sri Lanka	74.8	6.7	0.65	8.89
Guatemala	95.6	6.1	0.59	6.34
Paraguay	42.0	5.6	0.54	13.24
Costa Rica	69.2	5.1	0.50	7.44
Jamaica	17.1	5.1	0.50	30.01
Morocco	131.0	4.9	0.48	3.75
Côte d'Ivoire	70.2	4.8	0.47	6.85
Kuwait	183.9	4.8	0.46	2.59
Jordan	48.7	4.7	0.46	9.74
Malaysia	406.9	4.5	0.43	1.10
El Salvador	32.0	4.3	0.42	13.54
Mongolia	17.1	4.0	0.39	23.43
Gabon	20.4	4.0	0.39	19.37
Iraq	264.2	3.6	0.35	1.38
Croatia	71.0	3.4	0.34	4.85
Belarus	74.0	2.7	0.27	3.68
Senegal	27.6	2.5	0.24	8.96
Azerbaijan	78.8	2.4	0.23	2.98
Zambia	29.1	2.2	0.22	7.71
Uzbekistan	81.1	2.1	0.21	2.61
Trinidad & Tobago	28.5	2.1	0.20	7.29
Bolivia	44.0	1.9	0.19	4.36
Lebanon	434.6	1.8	0.18	0.42
Honduras	31.4	1.8	0.17	5.63
Armenia	19.5	1.7	0.16	8.49
Slovakia	115.7	1.5	0.15	1.31
Lithuania	70.9	1.5	0.15	2.12

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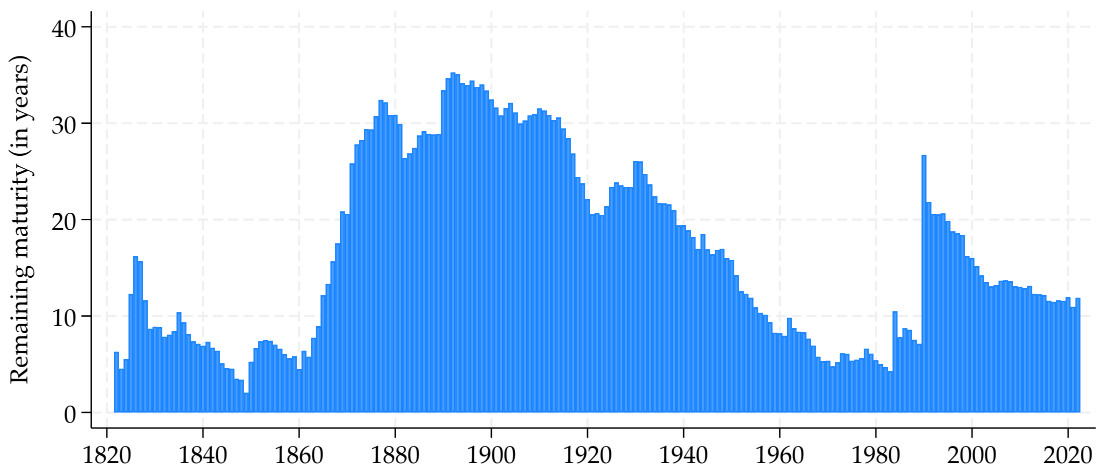
Table A.1 – continued

Country	GDP (current USD bn)	Market value (current USD bn)	Weight (in %)	MV/GDP (in %)
Venezuela	42.4	1.4	0.13	3.26
Vietnam	410.3	1.1	0.10	0.26
Cameroon	44.3	0.8	0.08	1.78
Mozambique	18.9	0.8	0.08	4.14
Namibia	12.6	0.8	0.08	6.19
Ethiopia	119.0	0.8	0.07	0.64
Rwanda	13.3	0.6	0.06	4.49
Barbados	6.3	0.5	0.05	8.39
Georgia	25.0	0.5	0.05	1.92
Papua New Guinea	31.6	0.5	0.05	1.49
Tajikistan	10.7	0.4	0.04	4.07
Suriname	3.6	0.4	0.04	11.78
Belize	2.8	0.3	0.03	9.52

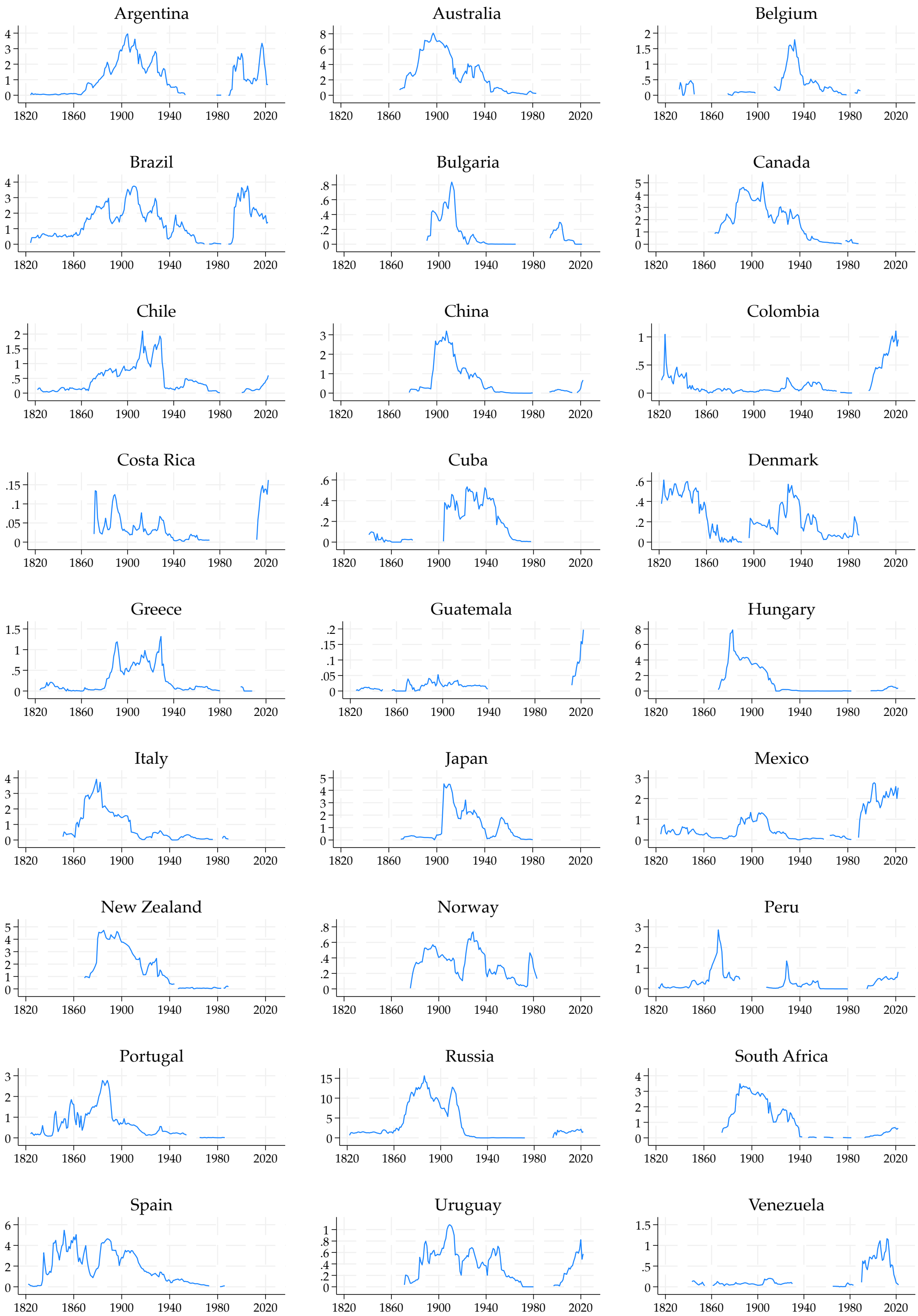
Note: Table shows market value of external bonds by country in 2022. All figures in current USD.

Figure A.1: EXBI: Remaining Maturity



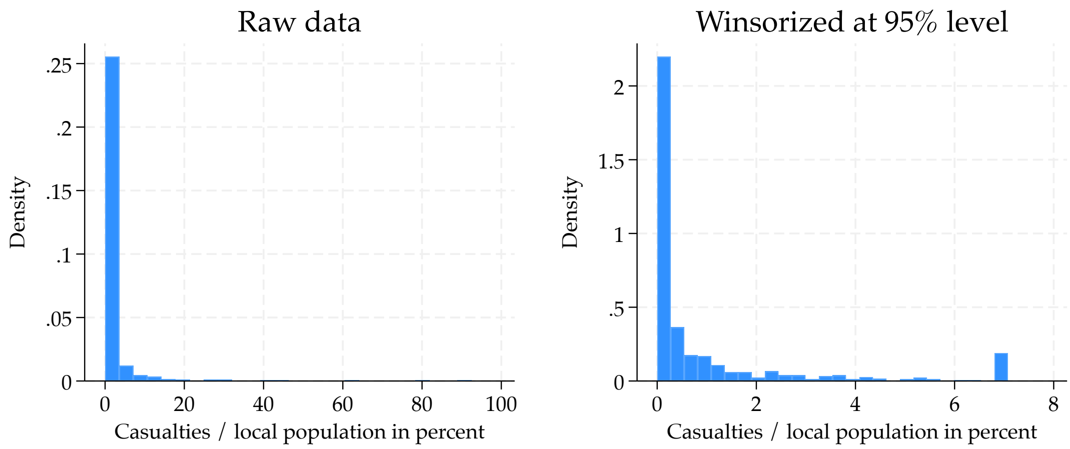
Note: Figure shows remaining maturity of bonds in EXBI over time. The gap during the 1980s is because the major lending instrument had been syndicated bank loans rather than international sovereign bonds at that time.

Figure A.2: Sovereign bond market capitalization by countries



Note: Market capitalization of sovereign bonds issued by different countries, expressed as a percentage of UK GDP.

Figure A.3: Distribution of war severity



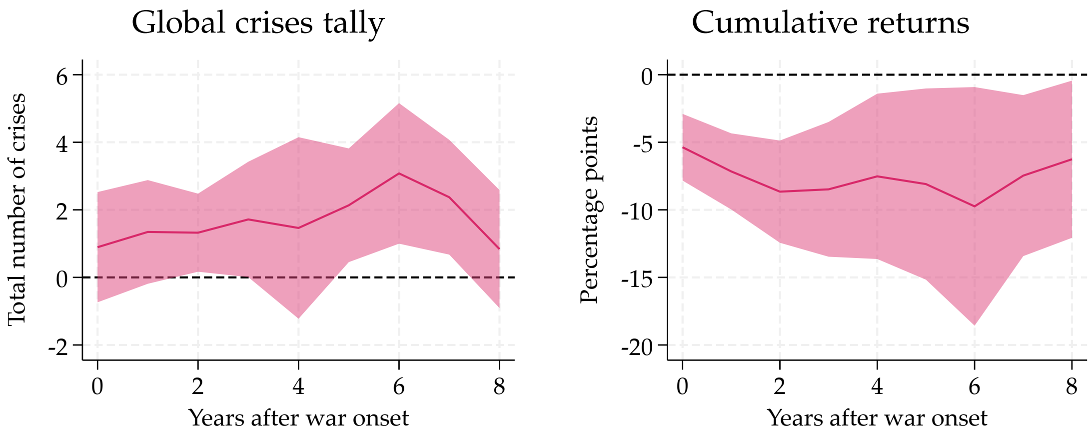
Note: Left panel shows histogram of war severity in sample, right panel shows same metric but winsorized at 95% level. Y-axis shows density, x-axis shows severity (casualties/local population) in percent.

B Wars and bond returns

B.1 Global wars and global returns

Figure B.1: Response of returns and crises to a 1 SD global war shock
Without winsorization 1 SD: Casualties/World Population ≈ 8bps

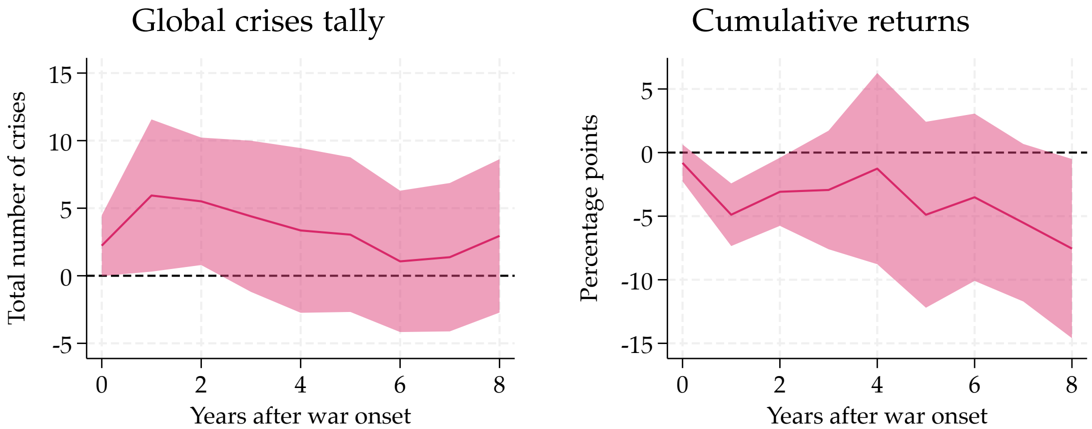
(Casualties/World population ≈ 2bps)



Note: Figure depicts results of estimating Equation (3.4), showing impulse responses to a one-standard-deviation unwinsorized global war shock (casualties/world population ≈ 8bps; allocation to weight and severity in equal parts). Left panel shows response of global crises tally (total of ongoing currency, inflation, and banking crises, external and domestic sovereign-debt crises, and stock-market crashes); Right panel shows response of cumulative log return series in percentage points. Shaded area denotes 90% confidence interval. Standard errors robust to heteroscedasticity and serial correlation with three lags (Newey and West, 1986).

Figure B.2: Response of returns and crises to a 1 SD global war shock
Restricted sample: 5% of most severe episodes excluded

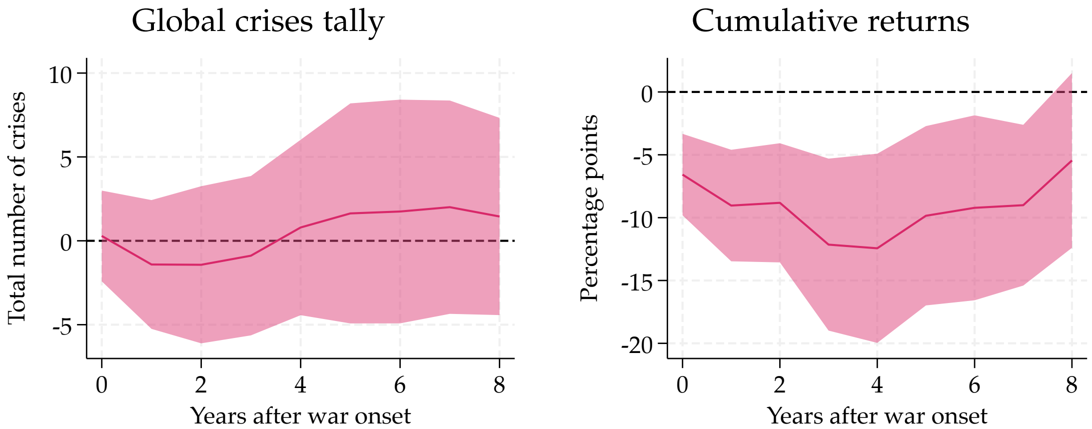
(Casualties/World population ≈ 2bps)



Note: Figure depicts results of estimating Equation (3.4), showing impulse responses to a one-standard-deviation global war shock (casualties/world population ≈ 2bps; allocation to weight and severity in equal parts). Left panel shows response of global crises tally (total of ongoing currency, inflation, and banking crises, external and domestic sovereign-debt crises, and stock-market crashes); Right panel shows response of cumulative log return series in percentage points. Shaded area denotes 90% confidence interval. Standard errors robust to heteroscedasticity and serial correlation with three lags (Newey and West, 1986).

Figure B.3: Response of returns and crises to a 1 SD global war shock
Restricted sample: only narratively identified interstate wars

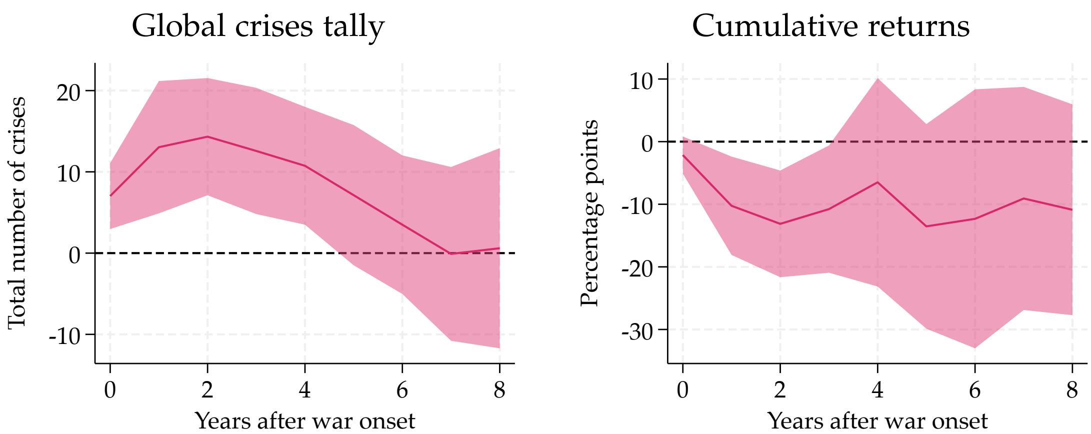
(Casualties/World population ≈ 2bps)



Note: Figure depicts results of estimating Equation (3.4), showing impulse responses to a one-standard-deviation global war shock (casualties/world population ≈ 2bps; allocation to weight and severity in equal parts). Left panel shows response of global crises tally (total of ongoing currency, inflation, and banking crises, external and domestic sovereign-debt crises, and stock-market crashes); Right panel shows response of cumulative log return series in percentage points. Shaded area denotes 90% confidence interval. Standard errors robust to heteroscedasticity and serial correlation with three lags (Newey and West, 1986).

Figure B.4: Response of returns and crises to a 1 SD global war shock
Restricted sample: only other wars

(Casualties/World population ≈ 2bps)



Note: Figure depicts results of estimating Equation (3.4), showing impulse responses to a one-standard-deviation global war shock (casualties/world population ≈ 2bps; allocation to weight and severity in equal parts). Left panel shows response of global crises tally (total of ongoing currency, inflation, and banking crises, external and domestic sovereign-debt crises, and stock-market crashes); Right panel shows response of cumulative log return series in percentage points. Shaded area denotes 90% confidence interval. Standard errors robust to heteroscedasticity and serial correlation with three lags (Newey and West, 1986).

Table B.1: Global bond returns, war, and other geopolitical risk indices ($k = 3$)

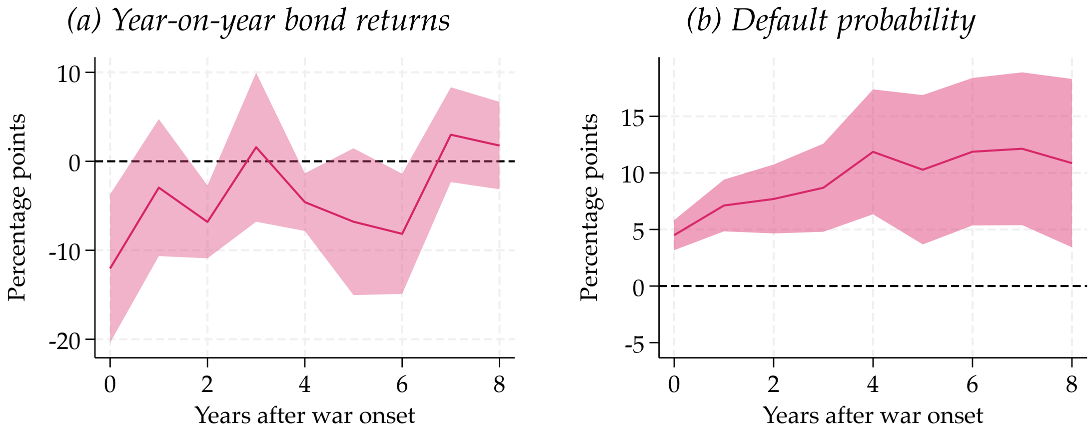
	Cumulative EXBI log return ($R_{t+k} - R_{t-1}$)							
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
GWI (1 SD)	-8.890**						-8.887**	-8.121**
	(4.175)						(4.142)	(3.844)
Acts (Caldara & Iacoviello, 2022)		-0.854						-0.209
		(4.476)						(3.782)
Threats (Caldara & Iacoviello, 2022)		-3.953*						-4.053
		(2.271)						(5.461)
War discourse (Hirshleifer et al., 2025)			-4.002					1.118
			(4.021)					(6.942)
ΔVIX proxy, USA				-0.269				
				(0.298)				
ΔVIX proxy, UK					-0.329			
					(0.456)			
ΔVIX proxy, spliced						-0.0943	-0.0970	-0.131
						(0.314)	(0.244)	(0.247)
Constant	18.49***	15.14***	16.59***	15.49***	15.48***	15.46***	18.45***	17.07***
	(2.388)	(3.038)	(3.212)	(3.338)	(3.359)	(3.352)	(2.408)	(3.742)
R^2	0.115	0.055	0.034	0.008	0.006	0.001	0.116	0.141
Observations	88	88	88	88	88	88	88	88

Note: Table shows results of estimating regression (3.4) for $k = 3$. No lagged controls included to facilitate interpretation of share of variation explained (R^2). Standard errors are reported in parentheses and robust to heteroscedasticity and serial correlation (Newey and West, 1986). ***$p < 0.01$, **$p < 0.05$, *$p < 0.1$.

B.2 Domestic wars and country-level returns

Figure B.5: Response of returns and bond defaults to a 1 SD local war shock
Restricted sample: 5% of most severe episodes excluded

(Casualties/Local population ≈ 1.6%)

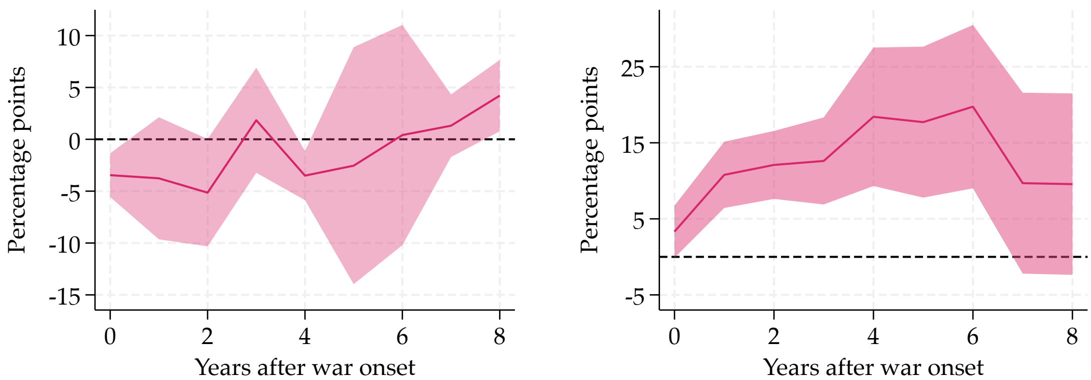


Note: Figure depicts results of estimating Equation (3.6), showing impulse responses to a one-standard-deviation domestic war shock (casualties/local population ≈ 1.6%). Left panel shows response of year-on-year bond returns in percentage points; Right panel shows average marginal effect on probability of being in default in percentage points. Shaded area denotes 90% confidence interval. Standard errors robust to heteroskedasticity as well as serial and cross-sectional correlation (Driscoll and Kraay, 1998).

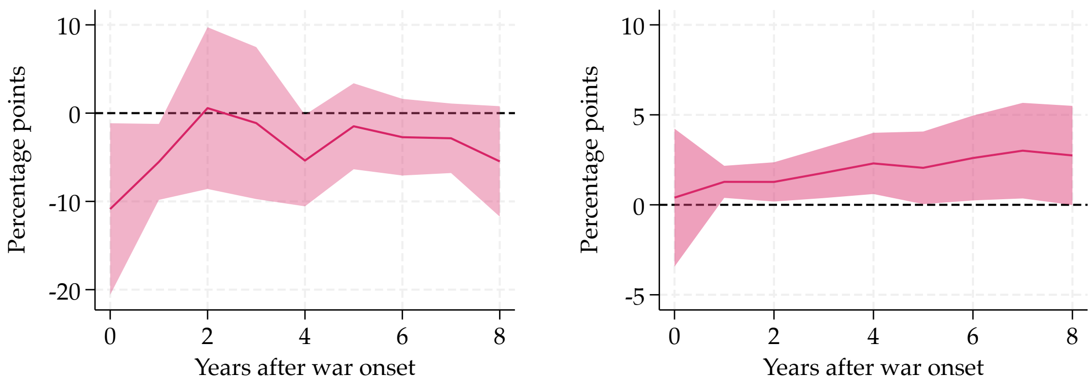
Figure B.6: Response of returns and bond defaults to a 1 SD local war shock (Casualties/Local population ≈ 1.6%)

(a) Year-on-year bond returns *(b) Default probability*

Restricted sample: 1870–1914



Restricted sample: 1915–2022

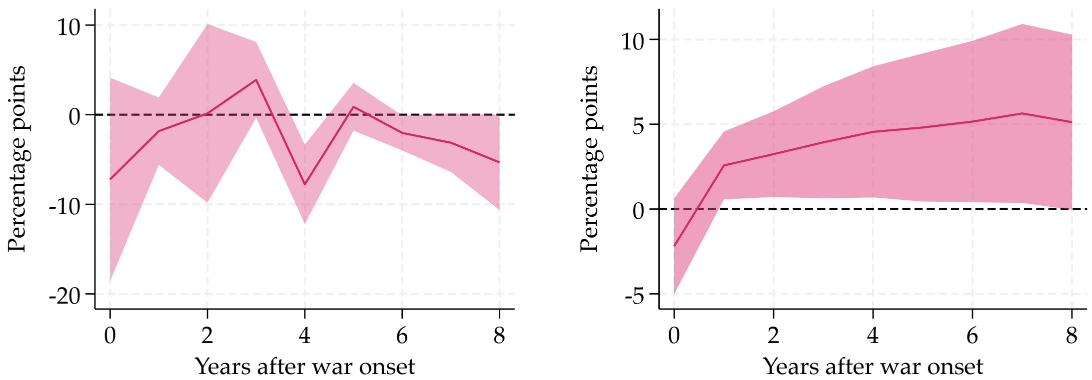


Note: Figure depicts results of estimating Equation (3.6), showing impulse responses to a one-standard-deviation domestic war shock (casualties/local population ≈ 1.6%). Left panels show response of year-on-year bond returns in percentage points; Right panels show average marginal effect on probability of being in default in percentage points. Shaded area denotes 90% confidence interval. Standard errors robust to heteroskedasticity as well as serial and cross-sectional correlation (Driscoll and Kraay, 1998).

Figure B.7: Response of returns and bond defaults to a 1 SD local war shock
Restricted sample: only narratively identified interstate wars

(Casualties/Local population ≈ 1.6%)

(a) Year-on-year bond returns *(b) Default probability*

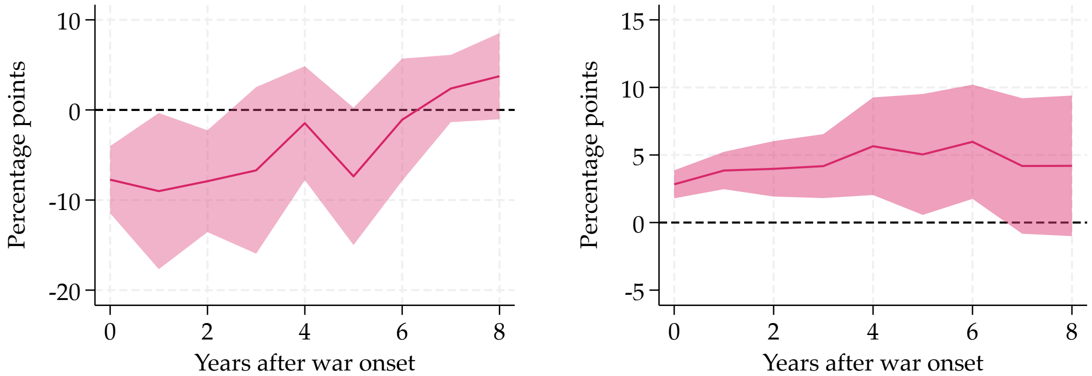


Note: Figure depicts results of estimating Equation (3.6), showing impulse responses to a one-standard-deviation domestic war shock (casualties/local population ≈ 1.6%). Left panel shows response of year-on-year bond returns in percentage points; Right panel shows average marginal effect on probability of being in default in percentage points. Shaded area denotes 90% confidence interval. Standard errors robust to heteroskedasticity as well as serial and cross-sectional correlation (Driscoll and Kraay, 1998).

Figure B.8: Response of returns and bond defaults to a 1 SD local war shock
Restricted sample: only other wars

(Casualties/Local population ≈ 1.6%)

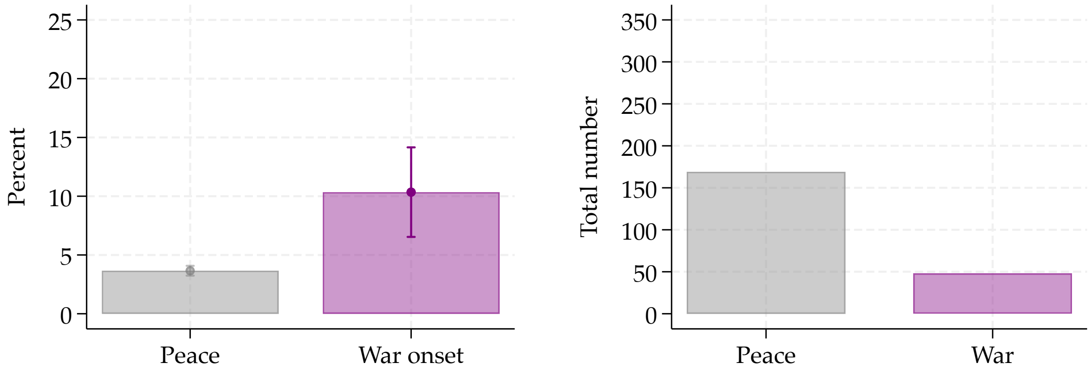
(a) Year-on-year bond returns — *(b) Default probability*



Note: Figure depicts results of estimating Equation (3.6), showing impulse responses to a one-standard-deviation domestic war shock (casualties/local population ≈ 1.6%). Left panel shows response of year-on-year bond returns in percentage points; Right panel shows average marginal effect on probability of being in default in percentage points. Shaded area denotes 90% confidence interval. Standard errors robust to heteroskedasticity as well as serial and cross-sectional correlation (Driscoll and Kraay, 1998).

Figure B.9: Alternative definition of bond collapses
Bond collapses: -40% over 2 years (Greenwood, Shleifer and You, 2019)

a) Annual probability of bond collapse — *b) Distribution of bond collapses*



Note: Left panel (a) shows annual probability of bond collapse conditional on war starting on own soil ("war onset") or not ("peace"). Right panel (b) shows distribution of bond collapses associated with war (all collapses within five years after war onset) and peace. Bond collapses defined as instances with a 2-year return below -40% (Greenwood, Shleifer and You, 2019).

Table B.2: Country-level bond returns, domestic wars, and pre-war distress

	EXBI-C log return ($r_{i,t}$)	
	(1)	(2)
War shock$_{i,t}$	-10.73**	-12.13***
	(5.051)	(3.267)
War shock$_{i,t}$ × Yield$_{i,t-1}$	2.776	
	(15.01)	
Yield$_{i,t-1}$	1.222	
	(2.339)	
War shock$_{i,t}$ × Yield$^{middle}_{i,t-1}$		2.746
		(4.725)
War shock$_{i,t}$ × Yield$^{high}_{i,t-1}$		2.186
		(8.932)
Yield$^{high}_{i,t-1}$		-0.891
		(5.492)
Yield$^{middle}_{i,t-1}$		1.773
		(4.548)
Constant	6.475***	5.965
	(2.463)	(5.225)
R^2	0.005	0.006
Observations	5,173	5,191

Note: Table shows results of estimating regression (3.6) for $k = 0$ (response of returns in year of war onset) and allows responses to differ depending on pre-war level of yields. $Yield^{high}_{i,t}$ is dummy indicating whether yields of i in t are among top 25% of yields in sample. $Yield^{middle}_{i,t}$ analogously signifies whether yields are between 25th and 75th percentile. No lagged controls apart from outlined coefficients included. Standard errors robust to heteroskedasticity as well as serial and cross-sectional correlation (Driscoll and Kraay, 1998).

C Threats and bond returns

Figure C.1: Most frequent threatening and threatened countries

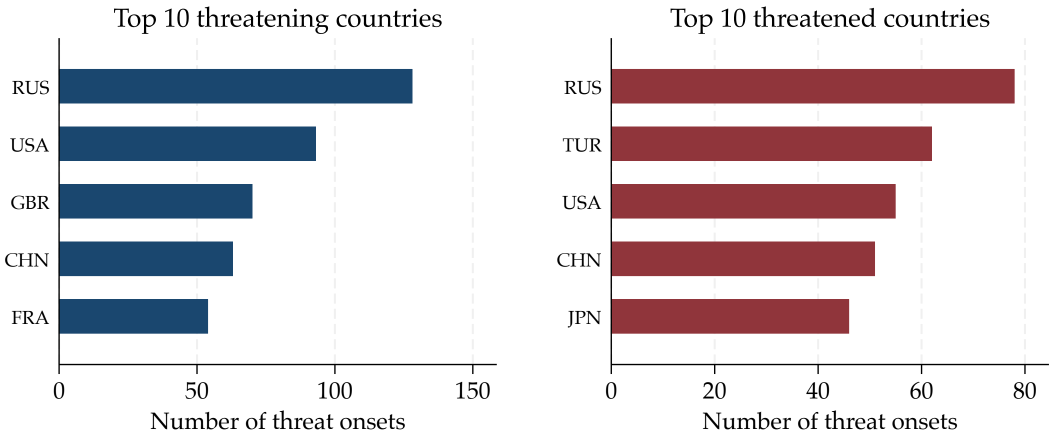
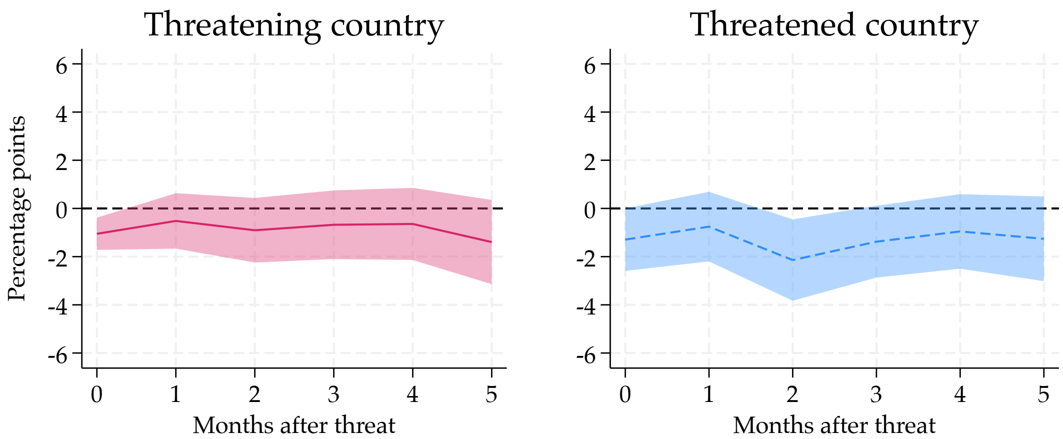
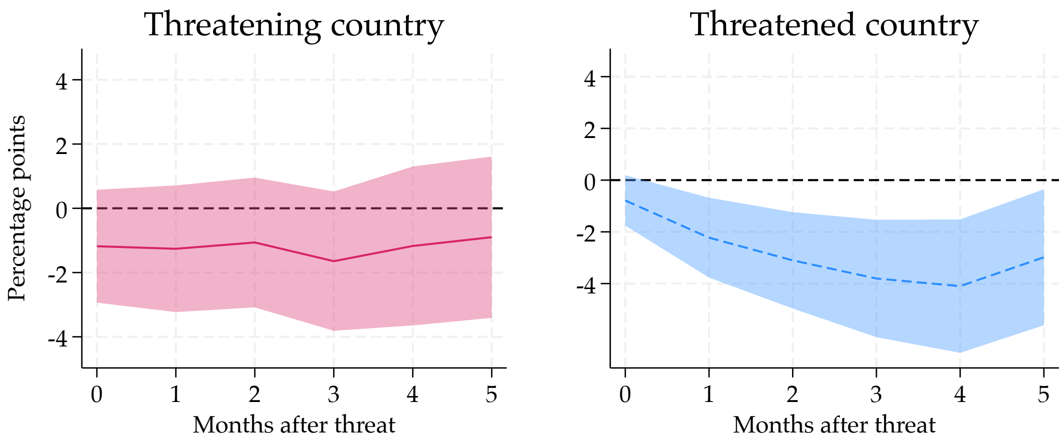


Figure C.2: Response of returns in threatening and threatened countries

a) Restricted panel: 1822–1914



b) Restricted panel: 1915–2014



Note: Figure depicts results of estimating Equation (4.3), showing impulse responses of cumulative returns to threat onsets. Left panels show response of threatening country (which initiates threat); Right panels show response of threatened country. Shaded area denotes 90% confidence interval. Top panel is restricted to period 1822–1914, bottom panel is restricted to period 1915–2014. Standard errors robust to heteroskedasticity as well as serial and cross-sectional correlation (Driscoll and Kraay, 1998).