Australian National University

Crawford School of Public Policy

CAMA

Centre for Applied Macroeconomic Analysis

The Repercussions of War Risks

CAMA Working Paper 48/2022
August 2022

Eric Tong
Reserve Bank of New Zealand, CAMA (ANU)

Abstract

I study the effects of the Russo-Ukrainian war on global financial markets of 87 developed and emerging economies. The methodology builds on Rigobon and Sack (2004) that focus on the shift in volatility on days of intense war news. I find that war risk caused considerable decline in asset prices, heightened stress in the financial system, and spike in commodity prices. However, the long-term risk-free rates remain anchored, suggesting that flight-to-safety behaviour was contained and had not morphed into an exodus of capital outflows towards the safest assets in the US. Over time, I find that the effects of a single, one-off war shock are transitory, peaking in about 15 days and dissipating within the month. This clarifies that the lasting impact of the war in the real world is attributable to a stream of war shocks rather than a persistent, singular shock. In a state-dependent model, I find that nations that share borders and have strong trade ties with the belligerents and NATO member states are more affected by war shocks. In contrast, financial markets of advanced and emerging economies do not exhibit significant differences. These results may inform strategies of nations and shape future outcomes.

Keywords

war; financial markets; conditional heteroscedasticity

JEL Classification

F30, F50, G10

Address for correspondence:

(E) cama.admin@anu.edu.au

ISSN 2206-0332

The repercussions of war risks

August 2022

Eric Tong

Reserve Bank of New Zealand, CAMA (ANU)

ABSTRACT

I study the effects of the Russo-Ukrainian war on global financial markets of 87 developed and emerging economies. The methodology builds on Rigobon and Sack (2004) that focus on the shift in volatility on days of intense war news. I find that war risk caused considerable decline in asset prices, heightened stress in the financial system, and spike in commodity prices. However, the long-term risk-free rates remain anchored, suggesting that flight-to-safety behaviour was contained and had not morphed into an exodus of capital outflows towards the safest assets in the US. Over time, I find that the effects of a single, one-off war shock are transitory, peaking in about 15 days and dissipating within the month. This clarifies that the lasting impact of the war in the real world is attributable to a stream of war shocks rather than a persistent, singular shock. In a state-dependent model, I find that nations that share borders and have strong trade ties with the belligerents and NATO member states are more affected by war shocks. In contrast, financial markets of advanced and emerging economies do not exhibit significant differences. These results may inform strategies of nations and shape future outcomes.

JEL: F30, F50, G10

Keywords: war; financial markets; conditional heteroscedasticity

I am grateful to Bob Buckle, Robert Kirkby, Matthew Ryan, an anonymous referee, and seminar participants at the University of Auckland, Auckland University of Technology, and the 4th Asia Conference on Business and Economic Studies (Vietnam 2022) for valuable feedback. The views expressed in this paper are solely the author's and do not reflect the views of the Reserve Bank of New Zealand.

Email address: eric.tong@rbnz.govt.nz

1. Introduction

Russia's invasion of Ukraine continues to be a seismic event, the full implications of which the rest of the world is struggling to grasp.[1] In light of the tragedy, this paper aims to assess the consequences of heightened war risks on international financial markets. The purposes are threefold. First, determining whether the war poses a demand or supply shock to economies is crucial, as different assessments call for different policy responses. Second, financial markets summarise all information about the potential disruptions of the war before they even take place. From this perspective, it is meaningful to see how different events shape the views of market participants about the severity and persistence of the war. Third, since the outbreak of the war, there is no shortage of analogies between the current conflict and the geopolitical tensions in the Asia-Pacific region.[2] As such, the results documented here may go beyond an ex-post explanation of an ad hoc event, and present a preface to a paradigm shift in geopolitics that matters to all.

In this paper, I apply the heteroscedasticity-based estimator proposed by Rigobon and Sack (2004) to assess the dynamic impacts of the war on global bond yields, stock prices, foreign exchange (FX) trading, commodity prices, and financial stress indicators that include implied volatility, credit default swap spreads, and the Composite Indicator of Systemic Stress (CISS) compiled by the European Central Bank (Holló *et al.* 2012). The heteroscedasticity-based estimator identifies war shocks from the shift in the second moment of financial variables on days of intense war news, which is apt given the volatile backdrop of the war. The brinkmanship exhibited prior to the outbreak of the war makes sign restriction, another common

[1] On the battlefield, where Russia may have hoped for a quick and easy victory, the war has now moved into an attritional phase. Russia cannot control Ukraine, and Ukraine cannot eject Russian forces from its territory. As of early April 2022, a supposedly unstoppable force has been met by a seemingly immovable object.

[2] Analogies were constantly made between the Russia's invasion of Ukraine and the potential China's aggression towards Taiwan. See articles from the RAND Corporation , PIIE, the Economist, and Bloomberg. At the time of writing, the presidents of the US and China exchange warnings on Taiwan, and the situation is escalated after the Chinese government threatened a military response to a planned visit to Taiwan by US House of Representative speaker Nancy Pelosi.

identification scheme, infeasible, as some war-related news can be perceived positively while others negatively, leaving an ambiguous net effect on the first moment of asset returns.[3] Identification by signs in these cases would be subjective on the part of the researcher. In contrast, by focusing on the second moment, one only needs to determine a set of days on which the variance of war-related news was elevated.

I identify war-news days by consulting the timelines compiled by Thomson Reuters, the Peterson Institute for International Economics, and the House of Commons of the UK Parliament regarding the war. These institutions focus on different aspects of the war, and combining their timelines facilitates the identification of a single war risk factor that captures the multidimensional facets of war news, such as the likelihood of war and its expected duration and costs. Depending on the variables of interest, my sample covers 10–87 countries over January 1st 2021–February 28th 2022, three days after Russia declared war on Ukraine.[4]

The results indicate that war-related news significantly affected European and global financial assets. Specifically, increases in war risk caused considerable declines in equity prices, depreciation of currencies against the dollar, heightened stress in the financial system, and a rise in commodity prices, including energy, metal and food prices. However, the long-term government bond yields are not significantly affected. This means that war risks per se had not caused the long-term costs of government borrowings to soar, potentially due to effective monetary interventions. It also means that while there was flight-to-safety behaviour, the war had not caused an exodus of capital flows from international markets towards the safest and most liquid US Treasury Notes. Governments around the world remain able to borrow at a stable, long-term risk-free rate.

[3] News can be perceived positively if they point to de-escalation of war risks, or if they imply changes in commodity prices that benefit certain commodity producers.

[4] Dependent variables include 51 countries for bond yield, 83 stock markets, 80 currencies, 9 commodity prices, 11 implied volatility indices, 7 credit default swap spreads, and 13 systemic stress indicators.

To explore the dynamic effect of war shocks, I follow Bu *et al.* (2021) to derive a series of war shocks from the heteroscedasticity-based estimator. The methodology builds on Bu *et al.* (2021), and involves running cross-sectional regressions of financial asset returns on the heteroscedasticity-based estimators on each war-news day. The idea is akin to the second-stage regressions of Fama and MacBeth (1973). Deriving a war shocks series expands the scope of analysis, enabling us to consider not only the contemporaneous effect as in the Rigobon and Sack (2004) framework, but also the dynamic effect of the war over time. This is especially beneficial when news tend to happen in Europe but assets are traded across time zones with different trading days. When I propagate the shocks and examine their trajectories, I find that the impact of war shocks usually peaks in about 10–15 days after the shocks and dissipates within the month. This result clarifies that the lasting impact of the war in the real world is attributable to a stream of war shocks rather than a persistent, singular shock. As my sample ends shortly after the outbreak of war, before actual destruction and loss of lives took place, the results reported here likely represent the lower bound impact of the crisis.

When I re-examine the results of subsamples by categorising countries along different exogenous conditions, I find that member states of the NATO alliance, and nations that share borders or have strong trade ties with the belligerents are more affected by war risks in the stock and FX markets.[5] Counterpart countries often experience little to no impact in the financial markets at all. I also find that countries with strong trade ties with the belligerents experience more distressed financial conditions, while financial markets in advanced and emerging economies do not exhibit significant differences. These results may inform countries about the strategies and coordination efforts they should take in anticipation of future geopolitical risks.

[5] NATO stands for the North Atlantic Treaty Organisation. It is a group of 30 countries from Europe and North America that exists to protect the people and territory of its members.

The rest of the paper is organised as follows. Sections 2 and 3 describe the methodology and data. Section 4 presents the main results. Section 5 discusses various robustness checks, and Section 6 concludes.

2. Empirical Approach

I use the two-step identification scheme in Bu *et al.* (2021) to identify (i) the responses of financial assets to the latent war shocks and (ii) the time series of war shocks itself. In the first step, I use the heteroscedasticity-based estimator proposed by Rigobon and Sack (2005) to measure the financial variables' responses to war shocks. Following that, I run cross-sectional regressions of daily financial returns on the estimated responses obtained in the first step on each *day of intense war news* discussed in the data section (Section 3). Doing so enables the uncovering of war shocks. This approach is akin to the two-stage regression of Fama and MacBeth (1973).

2.1. First step

Consider the returns of two assets (r_{1t} and r_{2t}) that are related to war and other shocks as follows:

$$\begin{bmatrix} r_{1t} \\ r_{2t} \end{bmatrix} = \beta \cdot \begin{bmatrix} z_{1t} \\ z_{2t} \\ z_{3t} \\ \vdots \end{bmatrix}, \tag{1}$$

where the vector $z_t = [z_{1t}\ z_{2t}\ z_{3t}\ \ldots]'$ contains all shocks that influence the financial variables, including changes in monetary and fiscal policy, technological advancement, other macroeconomic developments, and news regarding the war denoted by z_{1t}.[6]

β is a matrix that captures the overall impact of shocks on the financial variables. The elements of this matrix are:

[6] I follow Ramey (2016)'s definition of shocks, that they are (i) exogenous with respect to the other current and lagged endogenous variables in the model; (ii) uncorrelated with other exogenous shocks; and (iii) unanticipated movements in exogenous variables or news about future movements in exogenous variables.

$$\beta = \begin{bmatrix} 1 & \beta_{12} & \beta_{13} & \cdots \\ \beta_{21} & \beta_{22} & \beta_{23} & \cdots \end{bmatrix}, \tag{2}$$

where β_{ij} represents the impact of the jth factor on the ith financial variable. The first column of the matrix β captures the impact of the war risk factor on the two financial variables. Because z_{1t} is unobservable, the model is identified up to a normalisation, and the impact on the first variable is set to unity. The impact of the war-risk factor on the second variable is represented by the coefficient β_{21}, which is the parameter of interest at this stage.

To estimate β_{21}, we can separate days into two subsamples, denoted W and NW. W are the days with remarkable intensity of war-related news, and NW are the closest days before W. If all parameters are stable between W and NW except the variance of z_{1t}, the difference between the variance-covariance matrices of the shocks, $\Delta\Sigma$, is driven only by the change in the intensity of war-related news:

$$\Delta\Sigma = \Sigma_W - \Sigma_{NW} = \lambda e_1 \cdot e_1{}', \tag{3}$$

where e_1 is the first column of the identity matrix with dimension equals to the number of factors in z_t, and λ is the change in the variance of the war risk factor from non-war-news days to war-news days.[7]

From Eq. (1), we can compute the variance-covariance matrix of the two financial variables for the set of war-news days, denoted Ω_W, and likewise for a set of non-war-news days, denoted Ω_{NW}, as follows:

$$\begin{aligned} \Omega_W &= \beta \cdot \Sigma_W \cdot \beta' \\ \Omega_{NW} &= \beta \cdot \Sigma_{NW} \cdot \beta' \end{aligned} \tag{4}$$

[7] The assumption is valid in our opinion. The range of uncertainty about the war is especially large on war-news days. Take the example of the closed-door meeting between Putin and the Germany's Chancellor. They could have walked out of the room and announced major differences resolved and significantly de-escalated the risks, or they could have announced negotiations collapsed and nuclear threats would ensue. The range of potential outcomes increases on these days of intense war news. On the other hand, the range of probabilities of other shocks, such as whether Covid-19 variants would emerge or whether vaccination against them would be developed does not seem to depend on the intensity of war news on a particular day.

Under our maintained assumptions, the change in the variance-covariance matrix of the financial variables, $\Delta\Omega = \Omega_W - \Omega_{NW}$, must be driven entirely by the change in the variance of the war risk factor (λ) scaled by the responses (β_{21}):

$$\Delta\Omega = \lambda \cdot \begin{bmatrix} 1 & \beta_{21} \\ \beta_{21} & \beta_{21}^2 \end{bmatrix}. \tag{5}$$

From Eq. (5), the parameter of interest β_{21} can then be estimated by the equation:

$$\beta_{21} = \frac{\Delta\widehat{\Omega}_{21}}{\Delta\widehat{\Omega}_{11}}, \tag{6}$$

where $\Delta\widehat{\Omega}_{ij}$ denotes the (i, j) element of the matrix $\Delta\widehat{\Omega}$. As Rigobon and Sack (2004) show, this estimator can be obtained by instrumental variable, κ_{1t}, defined as r_{1t} in war-news days, and minus r_{1t} in non-war-news days.

In addition, Eq. (5) contains three restrictions on the shift in the second moments of the financial variables, which can be used to estimate the two parameters, λ and β_{21}. Following Rigobon and Sack (2004), I use a generalised method-of-moments (GMM) estimation procedure, in which the two parameters are chosen to minimise the following loss function:

$$L = \left[vech\left(\Delta\widehat{\Omega} - \lambda\beta \cdot \beta'\right)\right]' W \left[vech\left(\Delta\widehat{\Omega} - \lambda\beta \cdot \beta'\right)\right],$$

where $\beta = [1 \quad \beta_{21}]'$.[8]

2.2. Second step

The second step is relatively straightforward and borrows from the cross-sectional regression of Fama and MacBeth (1973). Suppose we have estimated a vector of $\hat{\beta}_{i1}$ for each asset i, we can uncover the unobservable war shock from cross-sectional regressions of r_{it} on the estimated impact $\hat{\beta}_{i1}$ for each war-news day t,

$$r_{it} = \alpha_i + z_{1t}\hat{\beta}_{i1} + \varepsilon_{it}, \tag{7}$$

[8] For the sake of brevity we do not include more estimation details here. Full details can be found in Sections 3 of Rigobon and Sack (2004).

where ε_{it} denotes factors unrelated to war shocks. This series of estimated coefficients is the war shock series.

2.3. Impulse responses

I make use of the Jordà (2005) local projections (LP) within a fixed-effects panel model, where inference is based on Driscoll and Kraay (1998) standard errors that allow arbitrary correlations of the error term across countries and time. In particular, I estimate the impulse responses to the war shock by projecting a variable of interest on its lags and current and lagged values of z_1. For example, the response of bond yields at horizon h is estimated from the following fixed-effects panel regression:

$$y_{i,t+h} = \alpha_i^h + \sum_{k=1}^{3} \gamma_k^h y_{i,t-k} + \sum_{k=0}^{3} \delta_k^h War\ shock_{t-k} + \sum_{k=1}^{3} \theta_k^h Control_{t-k} + \varepsilon_{i,t+h}, \quad (8)$$

where $y_{i,t+h}$ is the variable of interest, α_i^h is the country fixed effect, and $War\ shock_t$ and $Control_t$ are the war shock series and control variables, respectively. Lags of y_t and $War\ shock_t$ are included in the regression to remove any predictable movements in them. This facilitates the identification of the unanticipated news of the war. δ_k^h gives the response of the outcome variable at horizon h to a war shock at time t.

The local projection method allows for state-dependent effects in a straightforward manner.[9] Accordingly, after presenting the benchmark results estimated by Eq. (8), I estimate a version that allows the effect of war shocks to vary by (i) the geopolitical position (whether it shares borders with Russia or Ukraine, or belongs to NATO), (ii) a country's trade exposure to Russia and Ukraine, and (iii) the state of development (whether a country is labelled as advanced or emerging economy by the IMF).[10] The first state is likely exogenous to war shocks, so I estimate:

[9] Local projections are more robust to misspecification than a non-linear vector auto VAR (Zeev 2019).
[10] Classification can be found on the International Monetary Fund website.

$$y_{i,t+h} = I_i \left[\alpha_{A,i}^h + \sum_{k=1}^{3} \gamma_{A,k}^h y_{i,t-k} + \sum_{k=0}^{3} \delta_{A,k}^h War\ shock_{t-k} + \sum_{k=1}^{3} \theta_{A,k}^h Control_{t-k} \right]$$

$$+(1 - I_i) \left[\alpha_{B,i}^h + \sum_{k=1}^{3} \gamma_{B,k}^h y_{i,t-k} + \sum_{k=0}^{3} \delta_{B,k}^h War\ shock_{t-k} + \sum_{k=1}^{3} \theta_{B,k}^h Control_{t-k} \right] \quad (9)$$

$$+ \varepsilon_{i,t+h}^h,$$

where I is a dummy variable that takes the value of one if a country belongs to group (i). As states (ii) and (iii) may be interrelated, I expand Eq. (9) into:

$$y_{i,t+h} = I_i \left[\alpha_{A,i}^h + \sum_{k=1}^{3} \gamma_{A,k}^h y_{i,t-k} + \sum_{k=0}^{3} \delta_{A,k}^h War\ shock_{t-k} + \sum_{k=1}^{3} \theta_{A,k}^h Control_{t-k} \right]$$

$$+(1 - I_i) \left[\alpha_{B,i}^h + \sum_{k=1}^{3} \gamma_{B,k}^h y_{i,t-k} + \sum_{k=0}^{3} \delta_{B,k}^h War\ shock_{t-k} + \sum_{k=1}^{3} \theta_{B,k}^h Control_{t-k} \right]$$

$$+J_{i,t-1} \left[\alpha_{C,i}^h + \sum_{k=1}^{3} \gamma_{C,k}^h y_{i,t-k} + \sum_{k=0}^{3} \delta_{C,k}^h War\ shock_{t-k} + \sum_{k=1}^{3} \theta_{C,k}^h Control_{t-} \right. \quad (10)$$

$$+(1 - J_{i,t-1}) \left[\alpha_{D,i}^h + \sum_{k=1}^{3} \gamma_{D,k}^h y_{i,t-k} + \sum_{k=0}^{3} \delta_{D,k}^h War\ shock_{t-k} + \sum_{k=1}^{3} \theta_{D,k}^h Control_{t\text{-}} \right.$$

$$+ \varepsilon_{i,t+h}^h,$$

where J takes the value of one for countries that are highly exposed to trade with Russia and Ukraine (defined in the Data section). As trade intensity can be endogenous to economic development, I include as controls a dummy variable I that takes a value of 1 if it belongs to advanced economies and 0 otherwise.

Some of our variables of interest are not country-specific. In those cases, I estimate the time-series version of Eq. (8) with Newey-West standard errors instead. The number of lags included in the Newey-West correction is the horizon of the projection plus one day. This choice is based on the fact that LP residuals are autocorrelated up to the number of periods in the horizon. Appendix A outlines the autocorrelation process when the data-generating process follows an autoregressive model of order 1.

3. Data

3.1. Variables and sample

Our sample consists of six country-specific variables: 10-year government bond yields of 50 countries, stock market indices of 87 countries, exchange rates of the local currency in terms of US dollar of 83 countries, implied volatility indices of 11 countries, the 5-year credit default swap indices of seven regions, and a Composite Indicator of Systemic Stress (CISS) compiled by the European Central Bank (Holló *et al.* 2012) of 13 countries.[11] For shock identification, I use only European assets closest to the source of war news, as they respond contemporaneously to war news and are traded in the same time zone. For propagation and in assessing the dynamic impact of war shocks, I use both European and non-European assets as dependent variables in the local projections (Eqs. (8)–(10)). The list of countries and asset classes used in shocks identification is reported in Table 5, and the list of countries and asset classes used as dependent variables includes those in Table 5 and Appendix B. In all cases, I use daily data from January 1st 2021 to February 28th 2022. Data are drawn from Haver Analytics, Bloomberg and Refinitive Datastream.

In addition, I report the impact of war shocks on nine commodity prices. They include three energy prices (crude oil, coal, natural gas), three metal prices (precious metal, nickel, aluminum), and three food prices (wheat, corn, soybean oil). These data are drawn from various sources such as the Intercontinental Exchange, London Metal Exchange, Wall Street Journal, and Financial Times.

[11] My sample excludes currencies in a fixed exchange rate regime based on the classification of Ilzetzki *et al.* (2019), as in these cases, the exchange rate does not covary with news by construction.

The Composite Indicator of Systemic Stress (CISS) applies portfolio theory to the aggregation of five market-specific sub-indexes: the foreign exchange market, the equity market, the money market, the bond market, and the financial intermediaries. The aggregation takes into account time-varying cross-correlations between the five sub-indexes. As a result, the CISS puts more weight on situations in which stress prevails in several market segments simultaneously. Thus, it captures the idea that financial stress is more systemic and dangerous for an economy if financial instability spreads widely across the financial system.[12]

Four control variables are included in all regressions. They are oil price, the Citigroup Economic Surprise Index, the JP Morgan Nominal Broad Effective Exchange rate (NEER) of the US dollar, and the FTSE Global All Cap Index that comprises large, mid-sized and small company shares in developed countries and emerging markets. The latter serves as a proxy for the market portfolio, a significant explanatory factor of excess asset returns (Sharpe 1964; Lintner 1975).[13]

Apart from the bond yields, I take logs of all variables. To extract the cyclical components of the trending variables in my sample, I estimate a cubic-trend time polynomial for each variable and take the associated residuals as the corresponding variables' cyclical components as in Zeev (2019).[14]

[12] An evaluation of the CISS applied to euro area data confirms its robustness over time. The euro area CISS peaks during well-known periods of elevated financial stress, and a threshold vector autoregression shows that in a high-stress regime identified by CISS, a negative shock to the economy may trigger a downward spiral with financial and economic stress reinforcing each other over time, a finding which could be explained theoretically by financial accelerator mechanism eg. (Bernanke *et al.* 1999).
[13] The market portfolio contains all risky assets in proportion to their market value in the Capital Asset Pricing Model.
[14] Except for the economic surprise index, because surprises are cyclical in nature and hence does not make sense to further detrend. Results are essentially the same even if it is detrended.

3.2. News related to the war

I collect a list of 30 dates on which war-related events appeared to be the primary determinant of asset price movements (Table 1). These days are drawn from three sources: the House of Commons Library of the UK Parliament, the Peterson Institute for International Economics, and Reuters.[15] Combining their timelines helps cover different dimensions of the war. The original set includes 32 dates, but after removing days that overlap with the monetary policy announcements of the Federal Reserve, the European Central Bank, and other major economic announcements worldwide, we settle down at 30 events.[16]

Table 1 summarises the list of news considered in the estimations. For each piece of news, I crosscheck the timestamp to ensure that the news was the first of its kind that arrived at the traders' desks. If the news came after the European market's closure (only European assets are used in shock identification), they will be dated to the next trading day. The Excel spreadsheet (Timeline.xlsx) in the supplementary material documents the Greenwich Mean Time and the sources of the news.

[15] The hyperlinks to these sources are listed here: UK Parliament, PIIE, and Reuters.

[16] As a reference, Rigobon and Sack (2005) identify the effects of the Iraqi War in 2003 with 21 events. Carlomagno and Albagli (2022), which also use heteroscedasticity-based estimators, estimate the impulse response functions (IRF) based on 27 news. Similarly, Wright (2012) identifies monetary policy shocks by heteroscedasticity and estimate IRF in a structural vector autoregressive model (VAR) based on 28 monetary policy announcements.

Table 1 News related to the war

Date	Event
22/02/2021	President Zelenskyy imposes sanctions on Ukrainian politicians with close ties to Russian President Putin
6/04/2021	Russia announces the start of mass military drills
14/04/2021	Russia and Ukraine hold military drills. NATO criticises Russian troop build-up
22/04/2021	Russian Defence Minister announces Russia will re-deploy its forces back to their home bases by 1 May, temporarily averting the crisis
2/09/2021	President Zelenskyy presses US President Biden for a firm commitment to NATO membership
26/10/2021	Ukraine uses a Turkish Bayraktar TB2 drone in combat for the first time in eastern Ukraine, angering Russia
15/11/2021	President Zelenskyy says nearly 100,000 Russian soldiers are massed by mid-November
19/11/2021	In an unusual move, the US shares intelligence with allies about more than 150,000 Russian troops moving to Ukraine's border for a likely invasion
9/12/2021	US President Joe Biden warns Russia of sweeping Western economic sanctions if it invades Ukraine
13/12/2021	G7 Foreign Ministers and the High Representative of the EU issue a statement on Russia's military build-up and aggressive rhetoric towards Ukraine
15/12/2021	Putin discusses Ukraine tensions with France president Emmanuel Macron
17/12/2021	Russia presents security demands, including Ukraine will never gain NATO membership and NATO will give up military activity in eastern Europe
21/12/2021	Germany's Chancellor and Putin discuss Ukraine in first call
11/01/2022	US and Russian diplomats hold a day of negotiations over the fate of Ukraine. The talks are "useful" and "very professional", but no progress is made towards resolving fundamental disagreements
14/01/2022	A massive cyberattack leaves Ukrainian government websites temporarily unavailable
18/01/2022	Germany may consider halting Nord Stream 2 if Russia attacks Ukraine
25/01/2022	The US places 8,500 troops on heightened alert as NATO reinforces its eastern borders with warships and fighter jets, amid growing fears of a possible "lightning" attack by Russia to seize Kyiv
26/01/2022	The US rules out Russia's demand to halt NATO's eastward expansion, but opens to talks on arms control
28/01/2022	The Kremlin lists Russia's principal concerns as avoiding NATO expansion
2/02/2022	The US announces it will send 2,000 soldiers to Europe and reposition 1,000 from Germany to Romania to ensure the robust defence of European NATO members
4/02/2022	Presidents Xi Jinping of China and Putin sign a joint statement calling on the West to "abandon the ideologised approaches of the Cold War"
8/02/2022	Six Russian warships and a submarine pass through the Dardanelles strait, heading towards the Black Sea from the Mediterranean
11/02/2022	Russia launches its largest military exercise since the Cold War, holding joint manoeuvres with Belarus, close to the Belarus-Ukrainian border
14/02/2022	Biden tells US citizens to leave Ukraine, saying "things could go crazy quickly"
16/02/2022	President Putin confirms a partial drawdown of Russian forces near the Ukrainian border, in a step that could begin a de-escalation of tensions
17/02/2022	Russia expels the US diplomat from Moscow. The US State Department describes the move as escalatory
21/02/2022	Russia's strategic nuclear forces hold exercises overseen by Putin. The Kremlin says Russia successfully test-launched hypersonic and cruise missiles at sea and land-based targets during the exercises
22/02/2022	US and UK sanction Russian parliament members, banks and other assets in response to Putin's troop order. Germany halts the Nord Stream 2 gas pipeline project
24/02/2022	Putin authorizes "special military operations" in Ukraine. Russian forces begin missile and artillery attacks, striking major Ukrainian cities including Kyiv
25/02/2022	Ukraine's President Zelenskyy gives an early morning address and confirms multiple Russian missile strikes

To support the notion that identification by variance is more apt than identification by sign, Table 2 reports the first and second moments of financial variables on the war-news days relative to all other days over the sample period. The table shows that the average change does not differ significantly on the war-news days from the other days in the sample. One reason is that some war-news days were associated with increases in war risks while other days were associated with decreases, so the net direction of the cumulative war news is unclear. Accordingly, any attempt to identify war shocks by direction would be subjective.[17]

Table 2 Means and variances of changes in financial variables

	Mean			Standard deviation		
	War-news days (1)	Other days (2)	Diff. (1)-(2)	War-news days (1)	Other days (2)	Diff. (1)-(2)
Russia 2-year yield	0.138	0.017	0.121*	0.943	0.071	0.872***
Russia 10-year yield	0.119	0.009	0.11***	0.48	0.067	0.413***
Russia stock price	-1.445	0.05	-1.494**	8.621	1.212	7.409***
Russia stock expected volatility	3.488	0.109	3.379**	14.397	4.959	9.438***
Dollar/Ruble	-0.022	-0.033	0.011	1.776	0.642	1.133***
Gold price	0.308	0.009	0.299*	1.345	0.674	0.671***
Oil price	1.094	0.046	1.048**	2.043	2.169	-0.126
Gas futures	2.124	0.466	1.658	12.808	5.783	7.025***
Europe 10-year yield	0.001	0.002	-0.002	0.038	0.03	0.007*
Europe corporate yield spread	0.019	0.003	0.016*	0.082	0.038	0.044***
Europe stock price	-0.114	0.053	-0.167	1.234	0.867	0.366***
Dollar/Euro	-0.022	-0.033	0.011	0.534	0.341	0.193***
5-year expected inflation	0.015	0.005	0.01	0.05	0.039	0.012**

Note: bond yields, corporate yield spread and expected inflation are measured in percentage point changes; other prices are measured in percent changes. * denotes significance at 10% level, **at 1% level.

In contrast, the second moment of the variables can be informative about the effects of war risks. Indeed, what stands out from Table 2 is that the variance of financial variables increased sharply on the war-news days—by a significant amount for most of the financial variables, and by several multiples for many of them. This outcome is likely driven by the greater intensity of war-related news on those days, suggesting that those news significantly impacted these financial variables.

[17] For example, it is unclear if the discussions between President Putin and the France President on December 15th 2021 or with the Germany's Chancellor on December 21st 2021 were positive or negative news.

To formally test whether conditional variance increases on days of intense war news, I deploy a standard GARCH(1,1) model à la Bollerslev (1986):

$$
\begin{aligned}
r_t &= \mu + u_t, \\
\sigma_t^2 &= \alpha_0 + \alpha_1 u_{t-1}^2 + \phi_1 \sigma_{t-1}^2 + \gamma War_t,
\end{aligned}
\tag{11}
$$

where daily returns (r_t) earn a nonzero risk premium and the conditional variance of the residual (σ_t^2) is driven by lagged conditional variance, lagged squared residuals, and a dummy equals 1 on war-news days. If $\gamma > 0$, it reinforces the notion that variance increases on war-news days.

Fig. 1 shows that the lower bounds of most European assets used in shocks identification are above zero. To sharpen identification, I remove assets if their confidence bands are not significantly positive, although results are similar with or without the adjustment.

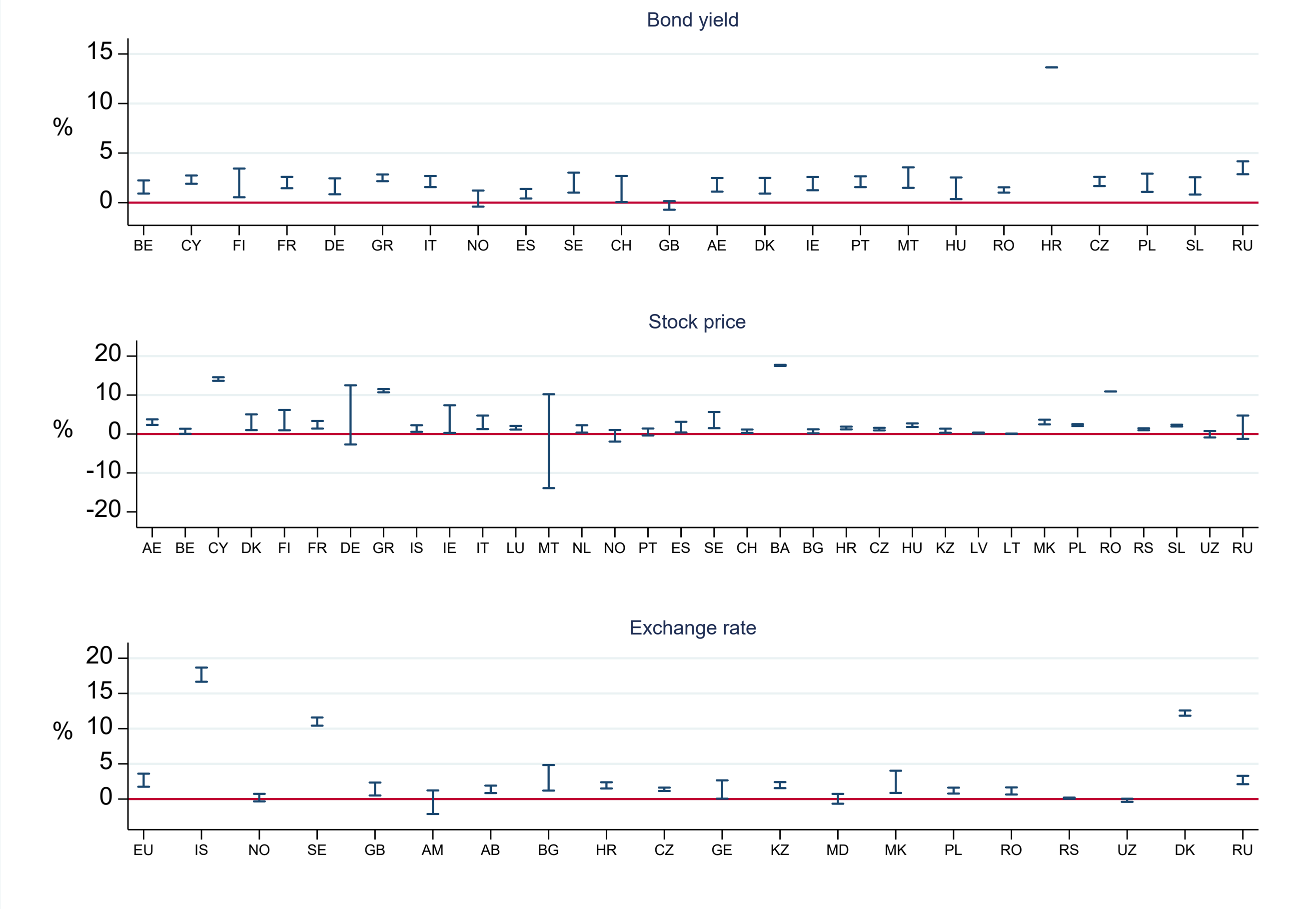

Fig. 1. 90 percent confidence bands of γ. Horizontal axis reports the Alpha-2 country code as described in the ISO 3166 international standard.

3.3. Reference asset

In section 2.1, we discussed that the identification strategy requires nominating a reference asset return (r_{1t}). Following Rigobon and Sack (2005) and Carlomagno and Albagli (2022) who use the two-year government bond yield as the reference, I use the two-year Russian local currency government bond yield as the reference return. Also known as *Obligatsyi Federal'novo Zaima* (OFZ) or Federal Loan Obligations, Russian government bonds are traded by local and foreign investors and have reflected the perspectives of domestic and global investors since 2009.[18] Foreigners' participation had increased from virtually zero in 2006 to

[18] The market was less active prior to that. From the supply side, blessed with commodity-driven fiscal surplus, the MOF had no financing pressure to issue OFZ bonds. From the demand side, high inflation led to negative real OFZ bond yields that were unattractive to institutional investors and other long-term investors. Market share of foreign investors was virtually zero in 2006.

about 30% in 2019, before reverting to about 20% as of 2021 (Lu & Yakovlev 2017) and Bank of Russia (2021). OFZ yields move in tandem with banks' cost of raising ruble through cross-currency swaps and with expected inflation (Fig. 2).[19] As such, the OFZ yields reflect market forces and allow us to gauge the views of market participants in response to the war.[20]

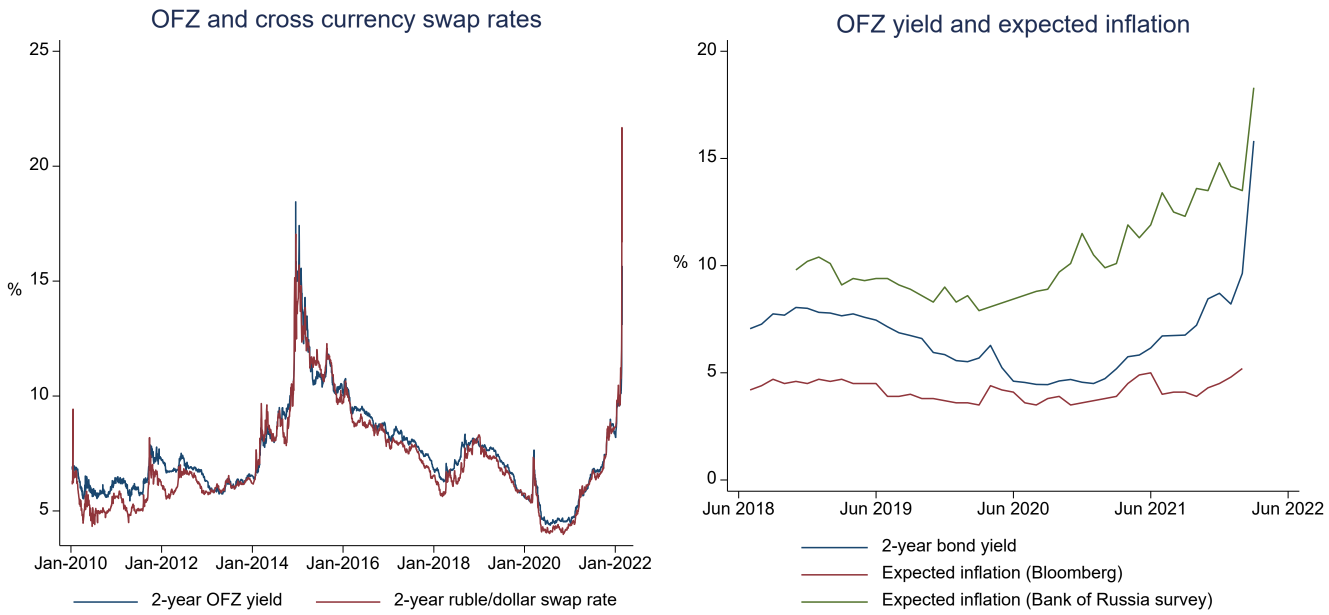

Fig. 2. Stylised facts of OFZ yields

4. Results

4.1. Contemporaneous effects

I apply the heteroscedasticity-based estimator to various Russian and European financial variables that are potentially influenced by war risk. At this stage, I confine the estimation to assets traded in similar time zones and trading days. The heteroscedasticity-based regression

[19] Russian banks raise a significant portion of their wholesale long-term funding by issuing dollar denominated bonds offshore. To hedge the currency mismatch risk where the revenue streams are denominated in ruble while the interest and principal repayments that arise from their foreign currency bond issuance are not, Russian banks enter into agreements where they initially exchange dollar for ruble, and regularly pay ruble interest rates and receive dollar interest rates linked to LIBOR. The red line on the left panel of Fig. 1 shows the equivalent 2-year funding cost by using swaps.

[20] In contrast, a less ideal reference asset would be the short-term interest rates, as they are largely pinned down by the current monetary policy setting and very near-term policy expectations. In that case, the responses of other assets relative to the three-month rate would not have been measured very well.

requires a standard set of war-news days, and including countries with different trading days and time zones would result in an inconsistent set of days. In section 4.3, I extract a war shock series that allows us to examine the dynamic effects of war risk across time zones.[21]

Table 3 shows the estimated coefficients from the IV estimator and the GMM estimator. The size of the shock is normalised to a 1 percentage point increase in the Russia 2-year government bond yield. To put this number in perspective, the increase in the Russian yield in the last three days before the de facto declaration of war on February 24th was 4.66 percentage points (pp).[22] Accordingly, the magnitude reported here is approximately the effect of a one-off war shock within a day. In the real world, a series of war shocks are constantly occurring, resulting in persistent severity as we observe and experience.

In Table 3, the coefficients obtained under both estimators are similar, indicating that the structure assumed in our framework is supported in the data. The primary finding is that the war risk factor significantly affects Russian financial variables. An increase in war risk of the magnitude considered results in a contemporary increase in the 5-year and 10-year government bond yields by 85 and 49 basis points, respectively. Equity prices fall by about 9% within a day (which matched the actual fall), and stock volatility goes up by about 10%. The ruble depreciates by 1.4% against the dollar, with the 1-year and 5-year forward rates dropping even more, suggesting that investors had anticipated a protracted war before it began.

Table 3 Effects of war risk on Russian assets

	IV	GMM
5-year government bond yield	0.85***	0.9***
10-year government bond yield	0.49***	0.6***

[21] Carlomagno and Albagli (2022) use the returns accumulated in two days to overcome the issue of time zones. I opt to not follow this method as it does not conform to the empirical framework first set up by Rigobon and Sack (2004) and it would double-count returns on consecutive war days. Other disadvantages of this method include erroneously treating non-war returns as war-day returns and introduce more noise that weakens identification.
[22] President Putin recognised the independence of the Donbas region and ordered troops into the territories on February 21st.

Stock price (MOEX Russia Index)	-8.86***	-8.88***
Stock volatility	9.92***	9.96***
Dollar/Ruble (Spot)	-1.35***	-1.78***
Dollar/Ruble (1-year forward)	-4.01***	-4.01***
Dollar/Ruble (5-year forward)	-7.6***	-8.67***

Table reports the impact of the war risk factor (normalised to cause a rise of 1pp of the 2-yr Russia bond yield) on each financial variable. Bond yields are measured in percentage point changes; other prices are measured in percent changes. Asterisks (***, **, *) indicate statistical significance at the 1%; 5% and 10%, respectively.

Table 4 shows the spillover of war risks to European assets. The average drop in stock price induced by the war tension is about 1.1% for advanced economies and 0.8% for emerging economies (within a day). The exchange rates also depreciate by about 0.4% against the dollar among advanced economies, reflecting narrowed interest rate differential relative to the US or heightened currency risk premium (through the lens of the asset-market model of exchange rates à la Engel (2014)). By contrast, the effect on 10-year government bond yields is insignificant for many countries and small for others. This result suggests that global long-term risk-free rates remain anchored by monetary authorities of individual countries. It also implies that while there was some flight-to-safety behaviour from risky to risk-free assets, it did not amount to an exodus of capital flows from other countries to the safest and most liquid US Treasury bonds.

Table 4 Effects of war risk on European assets (based on the IV implementation of heteroscedasticity-based estimator)

	10-year bond	Stock price	Exchange rate
		Advanced economies	
Austria	-0.01	-1.88***	
Belgium	-0.02**	-0.69***	
Cyprus	0	-1.24***	
Czech Republic	0.02*	-1.13***	-0.91***
Denmark	-0.01	-0.52	-0.36***
Finland	-0.01**	-1.13***	
France	-0.01*	-1.04***	

Germany	-0.01	-1.12***	**-0.35***
Greece	0.01	-1.47***	
Iceland		-1.4***	-0.54***
Ireland	-0.01	-1.28***	
Italy	-0.02*	-1.16***	
Lithuania		-1.65***	
Luxembourg		-1.56***	
Malta	0.01		
Netherlands		-0.83***	
Norway	-0.01		-0.43***
Portugal	-0.02*	-0.7***	
Slovenia	-0.02*	-2.03***	
Spain	0.01	-0.94***	
Sweden	-0.03***	-0.9***	-0.29**
Switzerland	0	-0.74***	
United Kingdom		-1.02***	-0.52***
		Developing countries	
Albania			-0.27***
Bosnia and Herzegovina		-0.01	
Bulgaria		-0.72***	-0.35***
Croatia	0.01	-1.25***	0.13
Georgia			0.17
Hungary	0.02	-2.55***	
Kazakhstan		-0.47**	0.33*
Latvia		-0.74***	
North Macedonia		-1.97***	
Poland	0.03*	-3.26***	-0.45***
Romania	0.04***	-0.66***	-0.3***
Serbia		-0.11	-0.1

Estimates based on the IV implementation of heteroscedasticity based estimator (normalized to cause a rise of 1pp of the 2-yr Russian yield). Average effects are in basis points for bond yield and in percent change for stock price and exchange rates. Asterisks (***, **, *) indicate statistical significance at the 1%; 5% and 10%, respectively.

4.2. Importance of war risk

In our setup, the greater amount of war-related news on the specified days increases the variance of r_j by $\beta_{j1}^2 \cdot \Delta Var(r_1)$. To measure $\Delta Var(r_1)$, I use the shift in the variance of the Russian two-year yield between war-news days and non-war-news days. Together with the point estimates of β_{j1}, I obtain an estimate of the shift in the variance of each financial variable attributable to the increased volatility of the war risk factor on the war-news days.

Table 5 reports the portion of the variance of the jth financial variable attributable to war-related news, $\beta_{j1}^2 \cdot \Delta Var(r_1)/Var(r_j)$. The results show that war shocks explain, on average, about 10% of bond yield changes, 52% of stock returns, and 25% of exchange rate changes on days of intense war news, although results vary by country and asset class. These results reinforce the notion that news about the war contribute significantly to asset price volatility.[23]

Table 5 Proportion of variance of European assets returns explained by war risk on days of intense war news

	10-year bond	Stock price	Exchange rate
		Advanced economies	
Austria	2.81	61.3	
Belgium	11.9	28.25	
Cyprus	0.36	71.6	
Czech Republic	11.31	59.95	76.47
Denmark	5.34	9.52	39.64
Finland	11.09	52.98	
France	12.7	49.74	
Germany	8.03	52.65	**38.89**
Greece	2.27	65.43	
Iceland		65.07	51.12
Ireland	4.57	58.24	
Italy	12.96	53.8	
Lithuania		80.83	
Luxembourg		45.08	
Malta	3.5		
Netherlands		40.61	
Norway	6.9		25.65
Portugal	15.09	30.91	
Slovenia	9.96	77.31	
Spain	9.44	42.45	
Sweden	28	33.92	17.05
Switzerland	1.09	46.01	

[23] Note that the war shocks do not need to explain 100% of the movement, as we assume the variance of other shocks is positive—only that they remain constant between war-news and non-war-news days.

United Kingdom			69.99
		Developing countries	
Albania			27.93
Bosnia and Herzegovina		0.03	
Bulgaria		59.38	38.56
Croatia	0.19	71.65	6.2
Georgia			3.1
Hungary	3.66	82.51	
Kazakhstan		19.96	5.96
Latvia		59.25	
North Macedonia		81.02	9.15
Poland	19.08	77.78	22.76
Romania	26.07	31.88	30.33
Serbia		1.04	5.02

4.3. Dynamic effects

A natural question that arises is how persistent the effects are. This is an important question because if the effects tend to revert immediately after the news, it could be argued that war news just generate an increase in returns' volatility, but asset prices (in levels) are not much affected.

As the event study approach is not well suited for analysing dynamic effects (because of time zone and other reasons), I follow Bu *et al.* (2021) to extract a series of war shocks and propagate it to other variables of interest. To check that the shocks derived are not endogenous—which could have happened if global leaders' stances towards the war depended on financial markets performance—I follow Minesso *et al.* (2022) and regress the war shock series on contemporaneous and lagged global bond yields (US 10-year yields), global stock index (FTSE Global All-Cap Stock Price Index), and the US dollar NEER. The latter two variables are log-differenced:

$$War\ shocks_t = \mu + \sum_{i=0}^{3} \rho^i \Delta US\ yield_{t-i} + \sum_{i=0}^{3} \tau^i \Delta Stock_{t-i} + \sum_{i=0}^{3} \varphi^i \Delta NEER_{t-i} + \varepsilon_t, \quad (12)$$

where the subscript t stands for days of intense war news.

If the relevant coefficients in Eq. (12) were found to be statistically significant, this would mean that financial variables can somewhat drive war shocks, hindering their validity as exogenous shocks. The set of regressors in Eq. (12) also includes the contemporary changes in financial market variables to test for the existence of common shocks that might move both financial markets and war shocks. Results reported in Table 6 show that changes in the war shocks are not systematically predicted by developments in financial markets as measured by changes in the US yields, the global stock market, and the US dollar NEER, the latter being also a measure of global risk. Most coefficients are statistically insignificant and explain a minimal share of the volatility of the war shocks. Moreover, the results of the F-test show that coefficients are not jointly significant.

Table 6 Estimates from Eq. (12) on war-news days

	Model (1)	Model (2)	Model (3)	Model (4)
ΔUS 10-year yield$_{t-1}$	-7.542			-8.461**
	(1.58)			(2.16)
ΔUS 10-year yield$_{t-2}$	-1.945			
	(0.40)			
ΔUS 10-year yield$_{t-3}$	4.132			
	(0.96)			
ΔGlobal Stock Price Index$_{t-1}$		0.206		-0.187
		(1.00)		(1.10)
ΔGlobal Stock Price Index$_{t-2}$		-0.214		
		(0.97)		
ΔGlobal Stock Price Index$_{t-3}$		0.482**		
		(2.35)		
ΔUS NEER$_{t-1}$			-0.993	-1.482**
			(1.50)	(2.41)
ΔUS NEER$_{t-2}$			-0.741	
			(1.45)	
ΔUS NEER$_{t-3}$			0.014	
			(0.04)	
Constant	0.072	0.153	0.035	0.038
	(0.36)	(1.04)	(0.22)	(0.28)
Number of observations	26	26	26	26
Prob>F	0.2	0.09	0.23	0.1
Adjusted R^2	0.068	0.153	0.065	0.27

* $p<0.1$; ** $p<0.05$; *** $p<0.01$

Notes: Stock prices and NEER are in log-differences and 10-year yields are in first difference. Global stock price is the FTSE All-cap Global Stock Price Index. NEER is the JP Morgan Nominal Broad Effective Exchange rate of the US dollar. 26 observations are included in each model. Heteroscedastic-consistent t-statistics are reported in parenthesis below coefficients.

Fig. 3 reports the cumulative impulse response functions (IRFs) of global financial assets to an increase of Russia 2-year yield by 1 pp induced by the war shock. Although the sample of countries used to derive these IRFs is broader and covers time zones across the world, the contemporaneous effects are consistent with the average results of European countries in Table 4, which constitutes a further signal of robustness. Specifically, Fig. 3 shows that the instantaneous effect on long-term government yields is largely insignificant, and that the stock price responses are the largest among the asset classes considered. The joint reading of these results suggests that although risk-free rates have remained unchanged, risk premiums increase and depress stock prices.

Regarding dynamic effects, the initial impacts of war shocks accumulate and peak in about 10–20 days after the shock before eventually dying out. Most responses revert to zero by 30 days. In all, I record a cumulative drop of 3% and 1% in global stock price and exchange rates against the dollar worldwide for a 1 pp shock in the Russian 2-year yield.

More concerning, I document a general increase in systemic financial risks caused by the war. The lower row of Fig. 3 shows that war shock causes a persistent increase in implied volatility by about 10% for over 20 days, an increase in the credit default swap spread by about 10% for a range of countries, and the level of systemic stress as measured by logged CISS index almost doubles by about 20 days after the shock.[24] These findings corroborate the concerns of the Financial Stability Board (2022), the European Systemic Risk Board, and the IMF.

[24] I have also estimated using the raw level of CISS. In this case, it peaks at about 0.05 unit, which is about 1 standard deviation change of the series. Fig. 9 reports this result.

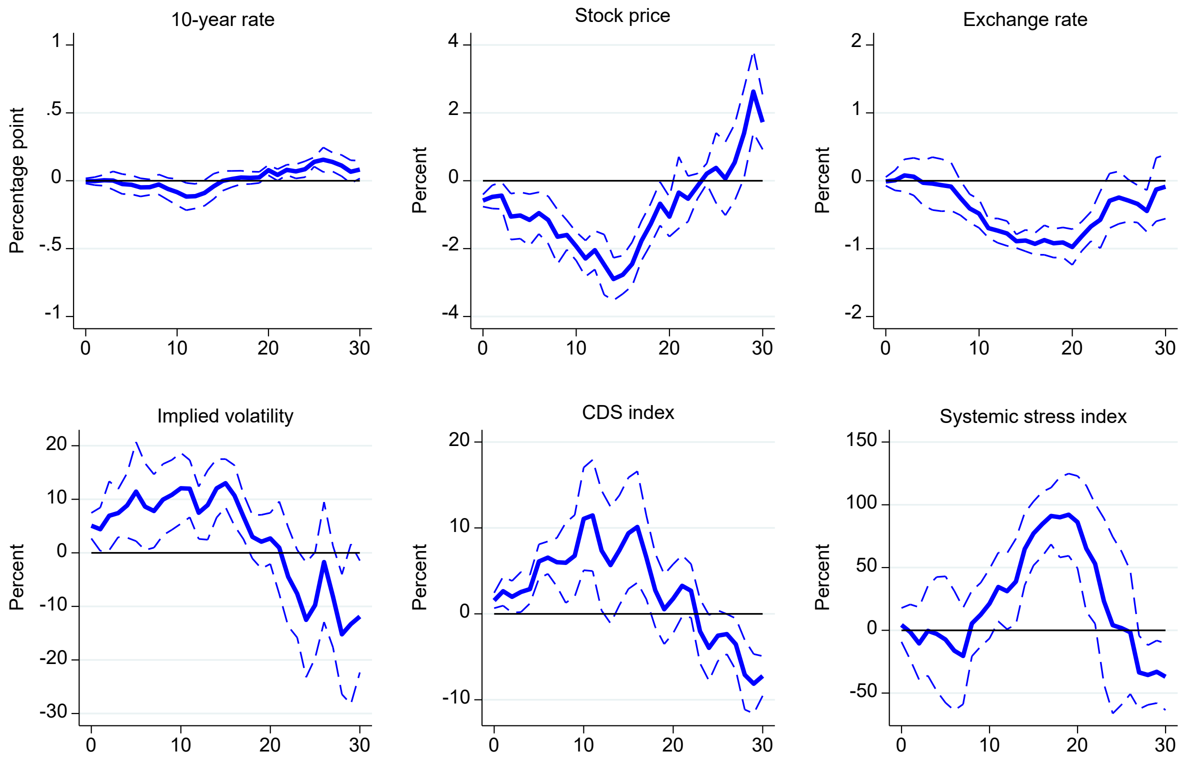

Fig. 3 Impulse response functions of global financial assets. IRFs are estimated from Eq. (8). The size of the shock is normalised to a 1 percentage point increase in the Russian 2-year yield. The 90% confidence bands are calculated from Driscoll and Kraay (1998) standard errors that allow arbitrary correlations of the error term across countries and time.

In addition, Fig. 4 shows that war shocks drive up energy prices (first row), metal prices (second row), and food prices (third row). These effects verify that investors anticipated the destruction in productive capacity and prohibition on trade brought by the war before they actually took place. These effects generally peak within 10–20 days after the shock and dissipate subsequently. Crude oil price rises by about 4% in 8 days after the shock. The magnitude is larger for coal (29% in 16 days), natural gas (20% in 7 days) and nickel (26% in 15 days).[25] By contrast, I do not find prices of corn and soybean oil increase despite the cost-of-living crisis experienced by us today.[26] As my sample stops short at the outbreak of the war and does not capture the severity of the actual destruction, the results documented in this paper

[25] For other commodity prices, precious metal and aluminium peak at 4% in 7 days, and wheat peaks at 5.7% in 20 days after the shock.
[26] In fact, prices decline initially after the shock.

likely represent a lower bound of the true damage. Overall, Fig. 4 shows that the effects on commodity prices are large but short-lived.[27] An inference is that had the conflict and destruction not taken place, normality in key markets could have been restored quickly after war risks dissipate.

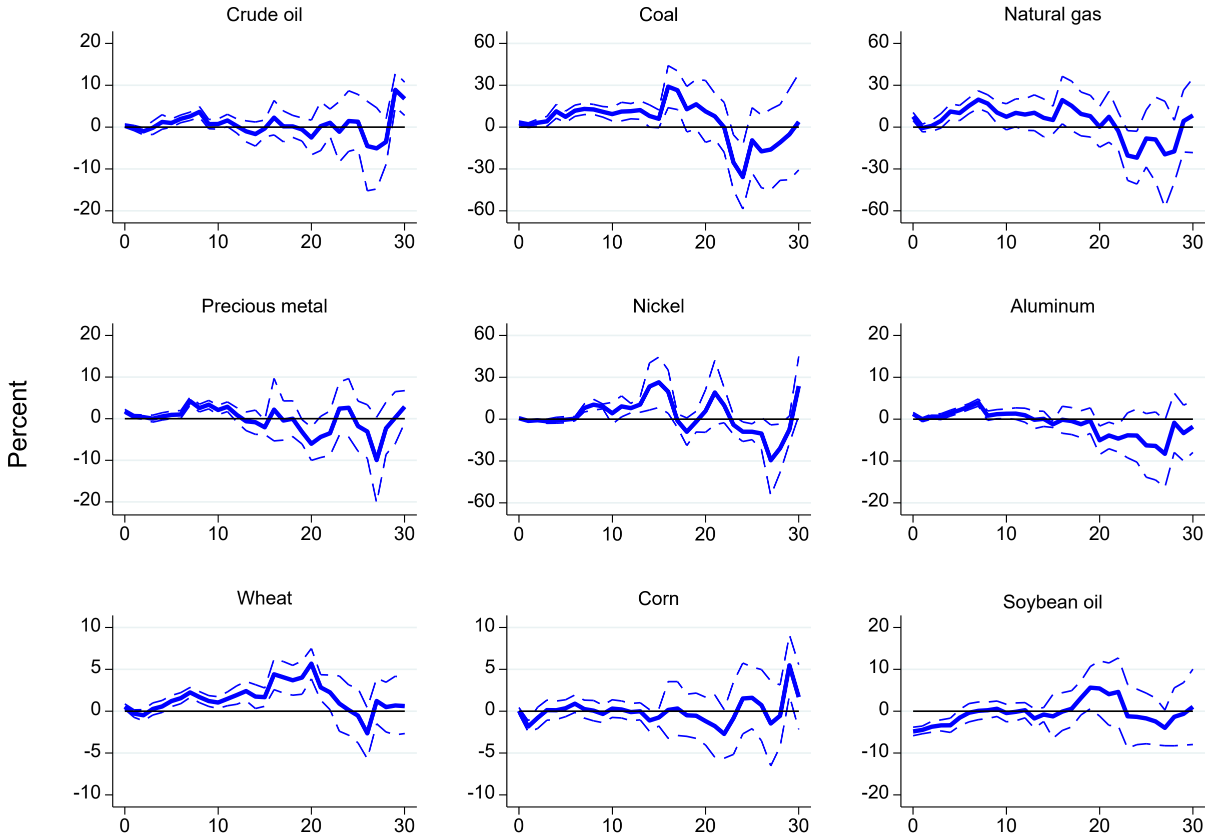

Fig. 4. Impulse response functions of global commodity prices. The size of the shock is normalised to a 1 percentage point increase in the Russia 2-year yield. The 90% confidence bands are calculated from Newey-West standard errors that allow arbitrary correlations of the error term across countries and time.

4.4. State-dependent effects

When I split the samples into (i) countries that share borders with Russia or Ukraine; or countries that belong to NATO alliance; (ii) countries with differing trade exposure to Russia and Ukraine in terms of $\frac{\text{(Exports+Imports)}}{\text{GDP}}$ obtained from the Direction of Trade Statistics, and

[27] Fig. 3 also shows that the effects on food prices are insignificant and not as large as the media depicts.

(iii) whether a country is classified as developed or emerging economy by the IMF, I find that war shocks exert nonlinear effects on countries. Fig.5 shows that countries that share borders with Russia or Ukraine—or belong to NATO—tend to experience worse performance in the stock and foreign exchange markets.[28] The trough of stock price is about -5% for countries in group (i), but only about -1% for other countries. The 90 percent confidence bands do not overlap, suggesting that the differences are significant. Similarly, neighbouring nations and NATO member states experience a currency depreciation of about 2% at the trough, while other countries experience less than 1% depreciation. These results suggest that geopolitical factors determine financial market performance during the war.

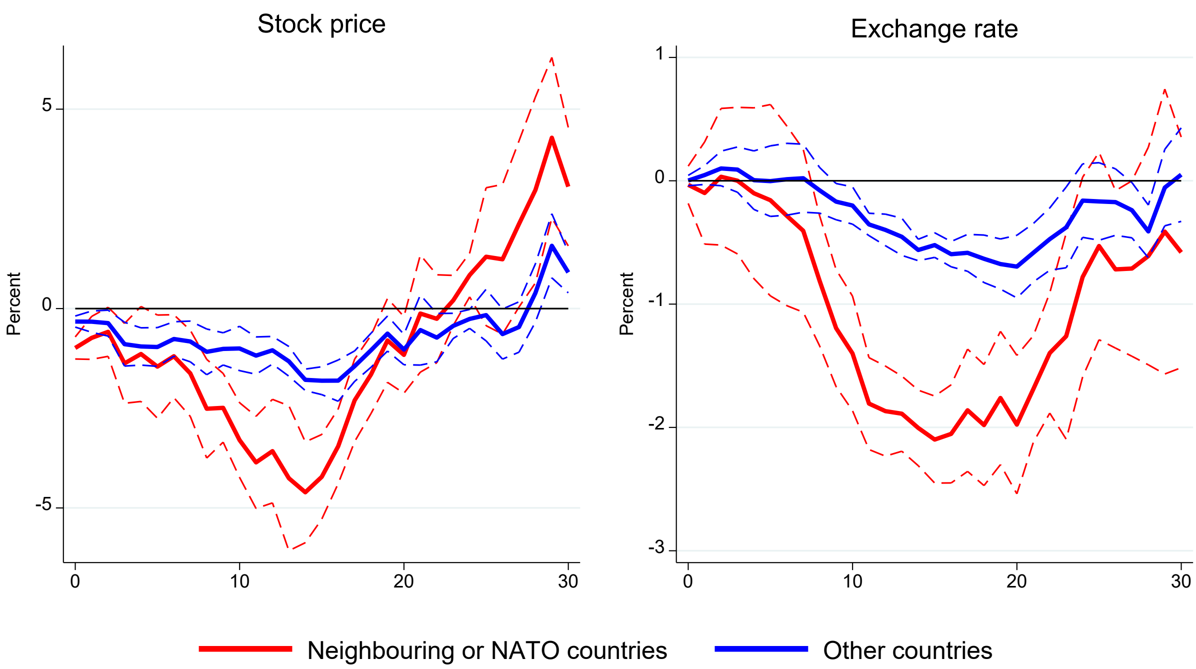

Fig.5. Effect of war on countries that share borders with Russia or Ukraine or that belong to NATO

[28] Japan, Sweden, and the US share only maritime borders with Russia. Since the outbreak of the war, the tension around these borders have intensified. The Russian Foreign Ministry has announced it will stop negotiations with Japan on a peace treaty to officially end a conflict dating back to World War II involving the disputed Kuril Islands (Source). Similarly, Sweden responded to Russia's move into Baltic Sea waters by swiftly boosting its military presence on the southeastern Swedish island of Gotland (Source).

Fig. 6 considers whether countries that trade more with the belligerents experience a larger decline in the stock and currency values. The right panel is remarkable. It shows that countries whose trade-to-GDP ratio with the belligerents are below the cross-sectional median experience nearly no depreciation in the currency.[29] The state of development is controlled for in these results. They suggest that trade exposure is a key determinant of capital movement and risk premium during a war crisis. In contrast, controlled for trade exposure, whether a country is labelled as advanced or emerging economy seems inconsequential to financial market performance, as Fig. 7 shows.

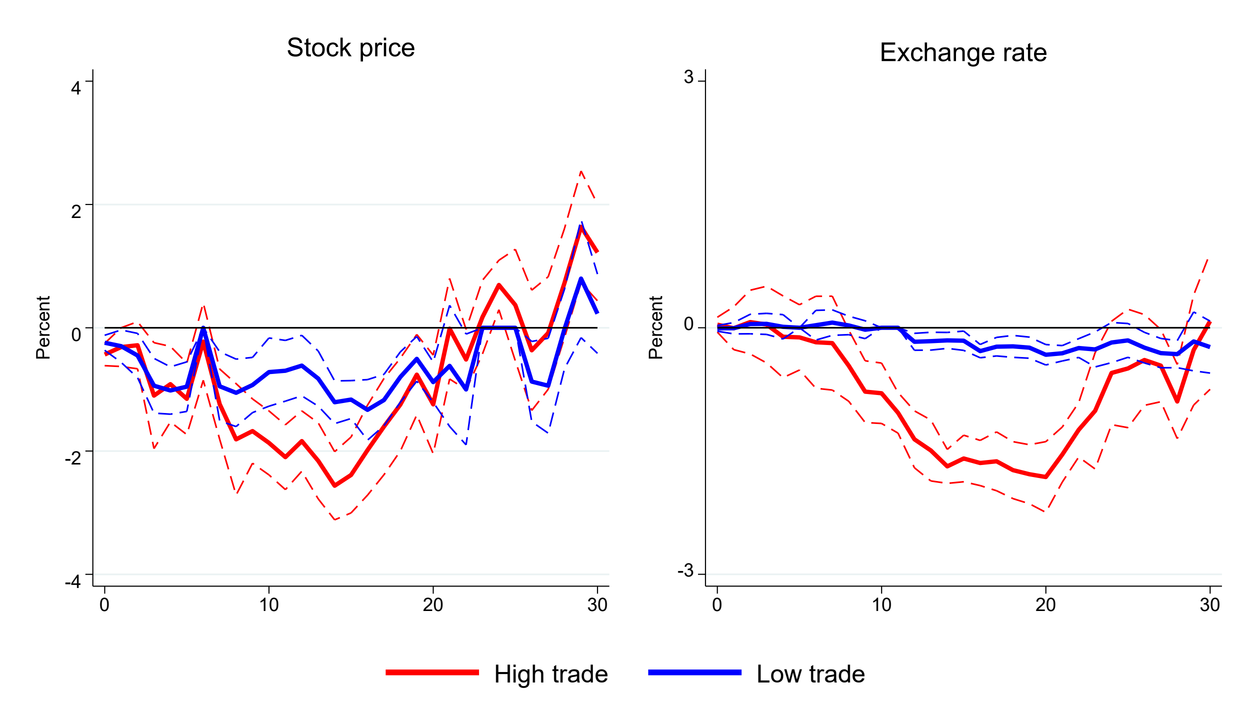

Fig. 6. Effect of war on countries with different trade exposure with Russia or Ukraine. ‘High trade’ is defined as countries whose trade-to-GDP is above the cross-sectional median of all countries, and ‘low trade’ refers to countries whose trade-to-GDP is below the cross-sectional median of all countries. These effects are controlled for the state of development of a country.

[29] Results in Fig. 6 hold when I classify high-trade and low-trade countries as the top and bottom 33rd percentile in terms of trade-to-GDP ratio.

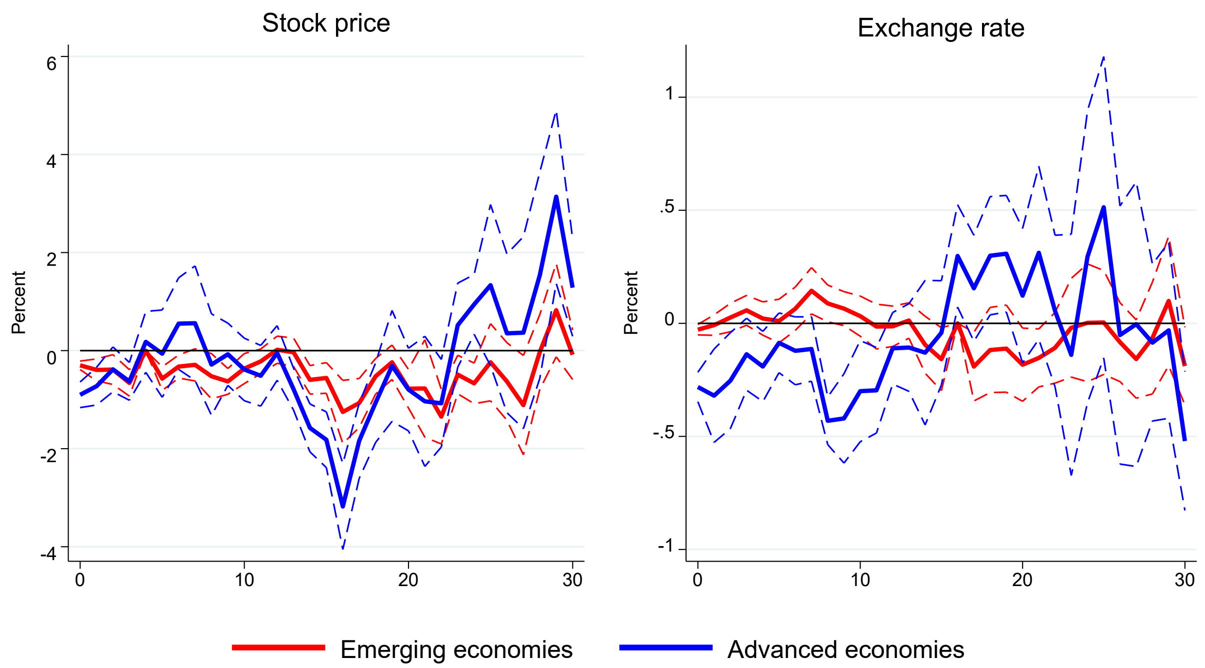

Fig. 7. Effect of war on financial markets of advanced and emerging economies

5. Robustness check

The results presented above are robust to a variety of changes to model specifications or estimation details, as reported in this section.

5.1. Multidimensional war risks

War risks may involve multiple dimensions, and compressing them into a single war shock would be invalid if the dimensions pose differentiated impacts on financial variables. In this subsection, I explore the implications of multidimensional war risks and formally test if the impact of different sources of war risks varies importantly.

To illustrate the point, assume that both z_{1t} and z_{2t} are sources of war risks in:

$$\begin{bmatrix} r_{1t} \\ r_{2t} \end{bmatrix} = \beta \cdot \begin{bmatrix} z_{1t} \\ z_{2t} \\ z_{3t} \\ \vdots \end{bmatrix}.$$

The response matrix β is given by

$$\beta = \begin{bmatrix} 1 & 1 & \beta_{13} & \dots \\ \beta_{21} & \beta_{22} & \beta_{23} & \dots \end{bmatrix},$$

where the effect of each war risks source on the reference asset is normalised to unity. Assume that the variances of both z_{1t} and z_{2t} shift on the days we identified. If both shocks are heteroscedastic, the changes in the second moments of the financial variables are:

$$\Delta var(r_1) = \Delta\sigma_{z1} + \Delta\sigma_{z2},$$
$$\Delta var(r_2) = \beta_{21}^2 \Delta\sigma_{z1} + \beta_{22}^2 \Delta\sigma_{z2},$$
$$\Delta Cov(r_1, r_2) = \beta_{21}\Delta\sigma_{z1} + \beta_{22}\Delta\sigma_{z2},$$

where $\Delta\sigma_{z1}$ and $\Delta\sigma_{z2}$ are the changes in the variance of the different dimensions of war risks. From here, two estimates can be formed:

$$\hat{d}_1 = \frac{\Delta Cov(r_1, r_2)}{\Delta var(r_1)} = \frac{\beta_{21}\Delta\sigma_{z1} + \beta_{22}\Delta\sigma_{z2}}{\Delta\sigma_{z1} + \Delta\sigma_{z2}},$$
$$\hat{d}_2 = \frac{\Delta var(r_2)}{\Delta Cov(r_1, r_2)} = \frac{\beta_{21}^2\Delta\sigma_{z1} + \beta_{22}^2\Delta\sigma_{z2}}{\beta_{21}\Delta\sigma_{z1} + \beta_{22}\Delta\sigma_{z2}}.$$

The two estimates will differ from one another unless one of the following conditions holds:

$$(1)\ \beta_{21} = \beta_{22},$$
$$(2)\ \Delta\sigma_{z1} = 0\ or\ \Delta\sigma_{z2} = 0.$$

The first case implies that the two dimensions of war risks have the same effect on the financial variables, hence they can be aggregated into a single factor. The second case explicitly shuts down one of the dimensions of war risks.

Table 7 shows that the null hypothesis of $\hat{d}_1 = \hat{d}_2$ cannot be rejected. These findings indicate that war risks are either dominated by one dimension or that different dimensions had about the same relative effects on the financial variables, regardless of the content of the news. Both cases justify our assumption that the impact of war risks can be captured by a single factor, and that it is reasonable to estimate the overall impact of all sources of war risks as we did.

Table 7 Paired t-test for null hypothesis: Mean (d1 – d2)=0

Variable	Mean	Standard errors	95% confidence bands	
d1	-0.7	0.15	-0.99	-0.41
d2	-0.56	0.22	-0.99	-0.13
d1 - d2	-0.14	0.14	-0.42	0.14
	H1: Mean(d1 - d2) < 0	H1: Mean(d1 - d2) ≠ 0	H1: Mean(d1 - d2) > 0	
p value	0.16	0.32	0.84	
Degree of freedom	70			

5.2. Event study leaving out some news

To check that our results are not driven by a single event, I repeat the estimations reported in Fig. 3 leaving out each of the 30 news in my sample, one at a time. Thus, for each asset class, we end up with 31 estimations: one that includes all of the news, and 30 other estimations, each one excluding a specific news. The results are reported in Fig. 8.

Each bar of the figure represents the estimated peaks or troughs leaving out one news, and the rightmost bar represents the full-sample estimation.[30] Blue colour indicates statistical significance at the 10% level, while red colour indicates statistical insignificance. The results show that while news near the outbreak of the war were more influential than earlier news, the coefficients remain statistically significant when we leave out any particular event.[31]

[30] First bar leaves out the first piece of news, second bar leaves out the second piece of news, and so forth.
[31] Rigobon and Sack (2005) derive that heteroscedasticity-based estimator is robust to small misspecifications in the determination of war days and misspecification on the different dimensions of war risk, such as likelihood of the war, its expected duration, and cost. The intuition of the first result is that the misspecified covariance matrices are linear combinations of the underlying true covariance matrices and hence can be estimated consistently, provided that more than half of the war news days are correctly identified. The second result comes from the fact that different dimensions of war tend to have the same relative effects on the financial variables regardless of the precise content of the news. Readers can refer to Section 5 of their paper for further details.

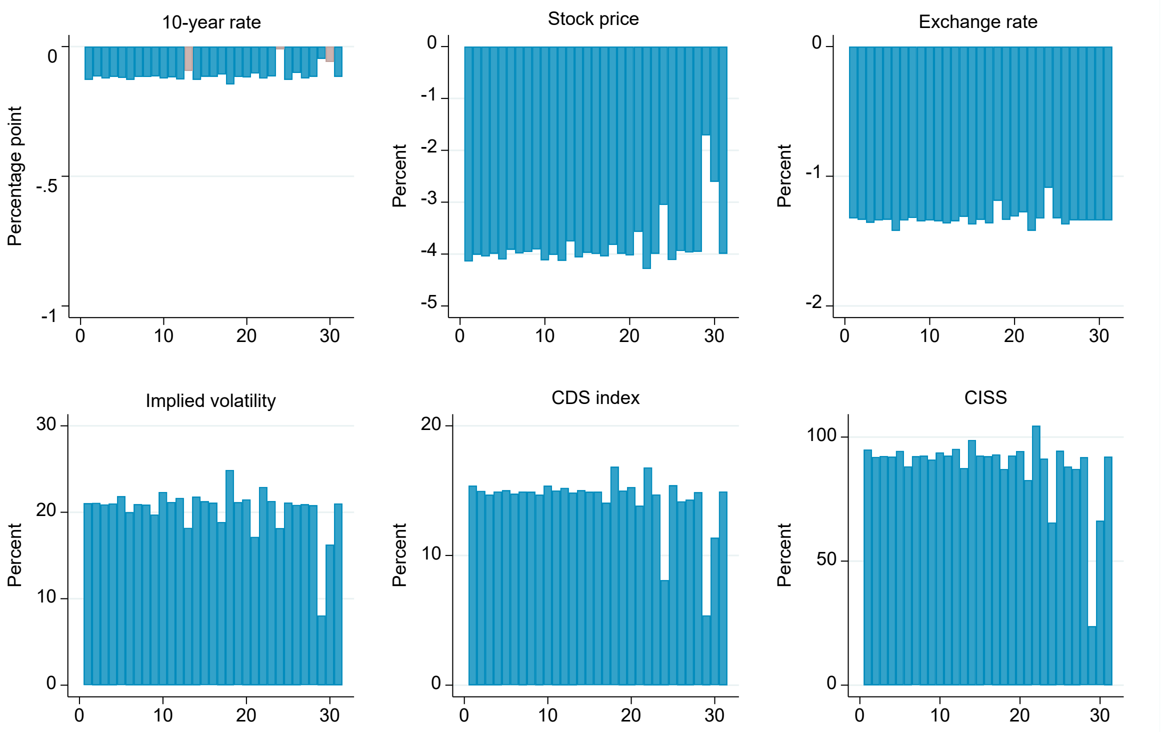

Fig. 8. Estimated peaks and troughs from the panel regression Eq. (8), leaving out 1 news. The first bar represents the estimation from omitting the first piece of news, the second bar from omitting second piece of news, and the rightmost bar represents the full-sample estimation. Blue bars denote significant at the 10% level. Red bars denotes insignificant at the 10% level.

5.3. Bootstrapped confidence bands

I re-estimate the benchmark results in Fig. 3 using nonparametric bootstrapping to compile the confidence bands. At each step of the IRF, observations are drawn with replacement for 1,000 times to compile the 90 percent confidence bands. This method is helpful when the theoretical distribution of the test statistic is unknown. Fig. 9 shows that the results are essentially the same as Fig. 3.

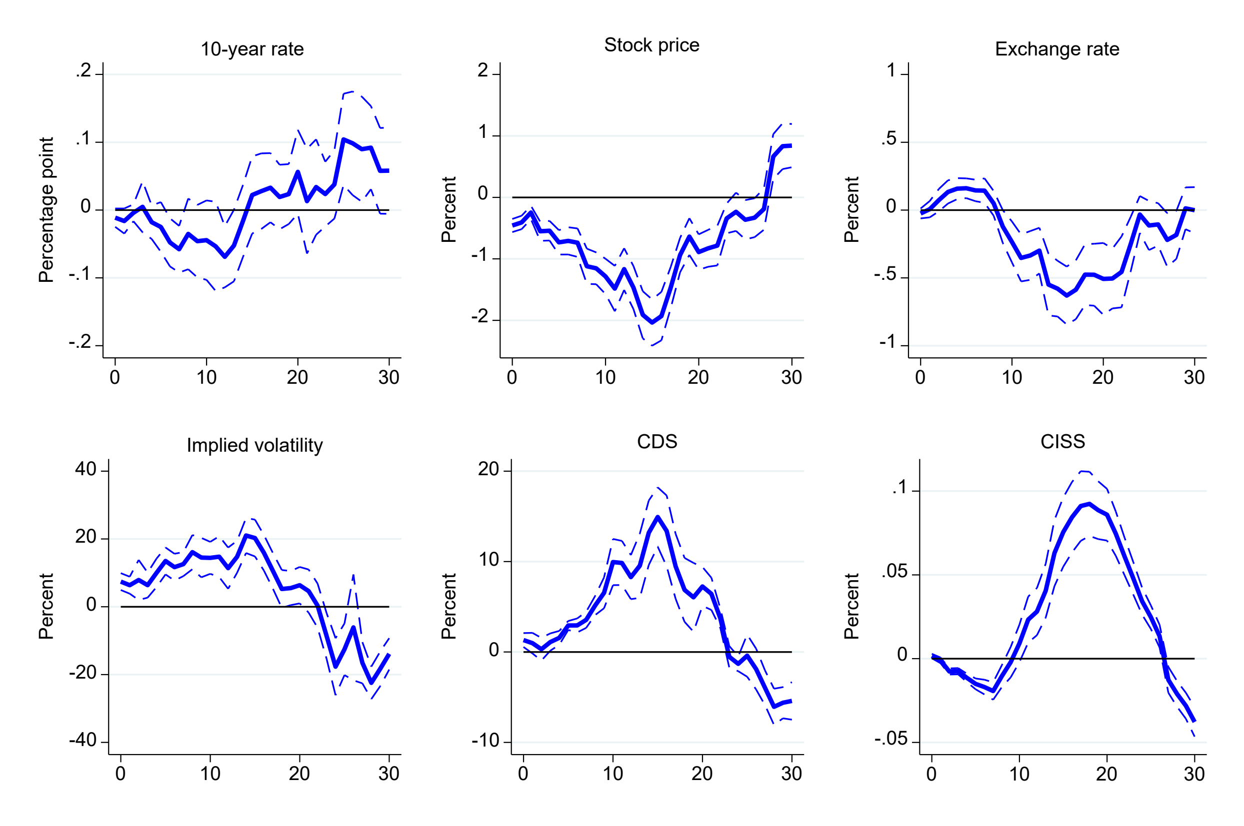

Fig. 9. Re-estimation of Fig. 3 with 90 percent confidence bands computed from nonparametric bootstrapping. Each step of the IRF is produced by bootstrapping 1000 times.

6. Conclusion

This paper analysed the impact of the Russo-Ukrainian war on world financial markets. The results show that war risks caused considerable increase in risk premium, systemic financial stress, and an increase in a range of commodity prices. However, the long-term risk-free rates remain anchored globally. The evidence suggests that greater war risks have induced investors to shift away from risky assets, but have not caused a widespread flight to the safest and most liquid assets.

I also find that war risk effects peak in about 15 days after the shock and dissipate within the month. These results should be understood as the effect of a single, one-off shock, when in reality a series of war shocks occurs on a daily basis. In state-dependent models, I find that neighbouring nations, NATO member states, and countries with strong trade ties with the belligerents are more affected by war shocks. These results are robust to a number of alternative specifications. Accordingly, they may inform the strategies of nations for managing future geopolitical risks. As time goes by, further research is needed to understand the long-run effects of the war on the real economy and other dimensions of well-being.

Appendix

A. Serial correlation in LP

Suppose the true data-generation process (DGP) follows an AR(1) model

$$y_{t+1} = Ay_t + \varepsilon_{t+1}, \tag{A1}$$

where y_t is a demeaned endogenous variables and ε_{t+1} is a white noise process with $E(\varepsilon_t) = 0$ and $Var(\varepsilon_t) = \sigma^2$. To estimate impulse responses using LP, one would estimate the impulse responses directly at each horizon with separate regressions. For instance, at horizon $h = 2$,

$$y_{t+2} = B^{(2)}y_t + e_{t+2}^{(2)}, \tag{A2}$$

where it is clear that $B^{(2)} = A^2$.

Iterate Eq. (A1) one period forward, we get

$$y_{t+2} = A^2y_t + A\varepsilon_{t+1} + \varepsilon_{t+2}. \tag{A3}$$

Substitute Eq. (A3) into (A2), we obtain

$$e_{t+2}^{(2)} = A\varepsilon_{t+1} + \varepsilon_{t+2}. \tag{A4}$$

To see that $e_t^{(2)}$ is a first-order autorrelation process, note that

$$Cov\left[e_{t+2}^{(2)}, e_{t+3}^{(2)}\right] = Cov[\varepsilon_{t+2}, A\varepsilon_{t+2}]$$

$$= A\sigma^2 \neq 0.$$

In general, $e_t^{(h)}$ follows a $(h-1)$-order autocorrelation process if the DGP is AR(1) process.

B. Countries in local projections estimations

All European countries in Table 3 are included in the estimations of these figures. In addition, countries in the following table are included for the respective variable of interest.

Variables	Countries
10-year bond yield	Australia, China, Japan, New Zealand, Singapore, South Korea, Hong Kong, India, Malaysia, Pakistan, Philippines, Sri Lanka, Taiwan, Vietnam, Egypt, Ghana, Israel, Kenya, Nigeria, South Africa, Turkey, Uganda
Stock price index	Argentina, Brazil, Canada, Chile, Colombia, Costa Rica, Jamaica, Mexico, Peru, Trinidad & Tobago, US, Venezuela, Australia, Bangladesh, Cambodia, China, Hong Kong, Indonesia, Japan, Korea, Lao PDR, Malaysia, Mongolia, Nepal, New Zealand, Pakistan, Philippines, Singapore, Sri Lanka, Taiwan, Thailand, Vietnam, Bahrain, Botswana, Egypt, Ghana, Iraq, Israel, Jordan, Kenya, Kuwait, Malawi, Mauritius, Nigeria, Oman, Qatar, Saudi Arabia, South Africa, Tunisia, Turkey, United Arab Emirates, Zambia,
Exchange rate	Argentina, Bolivia, Brazil, Canada, Chile, Colombia, Costa Rica, Dominican Republic, Guatemala, Honduras, Jamaica, Mexico, Nicaragua, Paraguay, Peru, Trinidad & Tobago, Uruguay, Venezuela, Australia, Bangladesh, Cambodia, China, India, Indonesia, Japan, Korea, Lao PDR, Macao, Malaysia, Mongolia, Myanmar, Nepal, New Zealand, Pakistan, Philippines, Singapore, Sri Lanka, Taiwan, Thailand, Vietnam, Angola, Bahrain, Botswana, Egypt, Ghana, Iran, Israel, Kenya, Kuwait, Lebanon, Mauritius, Morocco, Mozambique, Namibia, Oman, Nigeria, Qatar, Rwanda, South Africa, Tanzania, Tunisia, Turkey, Uganda, United Arab Emirates, Zambia
Implied volatility	Euro area, Germany, Spain, Switzerland, Australia, Hong Kong, India, Japan, South Africa, United States, emerging markets
CDS index	Europe, the US, emerging markets, Japan, Asia ex-Japan, Australia, and the UK
CISS	Australia, Belgium, China, Finland, France, Germany, Ireland, Italy, Netherlands, Portugal, Spain, the UK, and the US

Note: implied volatilities of the euro area, Germany and Spain are included in the same regression and while this may be construed as double-counting, I decide to include all to increase the sample size and on the basis that Germany and Spain alone may not capture volatilities of the euro area. Emerging markets implied volatilities is proxied by CBOE EM ETF Volatility Index (VXEEM) and similar reasoning is used to consider emerging markets, India and South Africa together.

References

Bernanke, B.S., Gertler, M., Gilchrist, S., 1999. The Financial Accelerator in a Quantitative Business Cycle Framework. Handbook of Macroeconomics 1, 1341-1393

Bollerslev, T., 1986. Generalized Autoregressive Conditional Heteroskedasticity. Journal of Econometrics 31, 307-327

Bu, C., Rogers, J., Wu, W., 2021. A Unified Measure of Fed Monetary Policy Shocks. Journal of Monetary Economics 118, 331-349

Carlomagno, G., Albagli, E., 2022. Trade Wars and Asset Prices. Journal of International Money and Finance 124, 102631

Driscoll, J.C., Kraay, A.C., 1998. Consistent Covariance Matrix Estimation with Spatially Dependent Panel Data. Review of Economics and Statistics 80, 549-560

Engel, C., 2014. Exchange Rates and Interest Parity. Handbook of International Economics 4, 453-522

Fama, E.F., MacBeth, J.D., 1973. Risk, Return, and Equilibrium: Empirical Tests. Journal of Political Economy 81, 607-636

Holló, D., Kremer, M., Lo Duca, M., 2012. Ciss-a Composite Indicator of Systemic Stress in the Financial System. European Central Bank Working Paper 1426

Ilzetzki, E., Reinhart, C.M., Rogoff, K.S., 2019. Exchange Arrangements Entering the Twenty-First Century: Which Anchor Will Hold? The Quarterly Journal of Economics 134, 599-646

Jordà, Ò., 2005. Estimation and Inference of Impulse Responses by Local Projections. American Economic Review 95, 161-182

Lintner, J., 1975. The Valuation of Risk Assets and the Selection of Risky Investments in Stock Portfolios and Capital Budgets. In: Stochastic Optimization Models in Finance. Elsevier, pp. 131-155.

Lu, Y., Yakovlev, D., 2017. Exploring the Role of Foreign Investors in Russia's Local Currency Government Bond (Ofz) Market. International Monetary Fund Working Paper 17/28

Minesso, M.F., Kurcz, F., Pagliari, M.S., 2022. Do Words Hurt More Than Actions? The Impact of Trade Tensions on Financial Markets. Journal of Applied Econometrics

Ramey, V.A., 2016. Macroeconomic Shocks and Their Propagation. Handbook of Macroeconomics 2, 71-162

Rigobon, R., Sack, B., 2004. The Impact of Monetary Policy on Asset Prices. Journal of Monetary Economics 51, 1553-1575

Rigobon, R., Sack, B., 2005. The Effects of War Risk on Us Financial Markets. Journal of Banking and Finance 29, 1769-1789

Sharpe, W.F., 1964. Capital Asset Prices: A Theory of Market Equilibrium under Conditions of Risk. The Journal of Finance 19, 425-442

Wright, J.H., 2012. What Does Monetary Policy Do to Long-Term Interest Rates at the Zero Lower Bound? The Economic Journal 122, F447-F466

Zeev, N.B., 2019. Global Credit Supply Shocks and Exchange Rate Regimes. Journal of International Economics 116, 1-32