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# Trade wars and asset prices

Guillermo Carlomagno *, Elías Albagli

*Central Bank of Chile, Chile*

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ABSTRACT

We study the effects of the trade war between the US and China since 2018 on international asset markets, including bonds, stocks, and exchange rates for a sample of 36 developed and emerging economies. Our main methodology relies on event studies, based on a daily indicator of trade-war news. We find that the typical negative news significantly and persistently compresses 10-y yields in the US and other developed economies, leading to cumulative effects close to −67 bp and −32 bp over the sample period, respectively. These episodes also significantly reduce stock markets in the US and other developed economies, with cumulative effects close to 23 and 14% for developed and emerging economies, respectively. For emerging countries, we also find significant currency depreciations against the USD. These results are confirmed with an alternative identification based on time-varying volatility. Given the relatively larger impacts on long-term yields, as well as the differential effects across developed and emerging markets, we conclude that the dominant channel behind the effects are increases in global risk aversion, as opposed to worsening perspectives about global growth prospects.



## 1. Introduction

The shift in international trade policies initiated by president Trump in 2018 constitute the most protectionist policy shock since the 1930 Smoot-Hawley Act, and the 1971 so called ‘Nixon shock’ (Fajgelbaum et al., 2020). Only in 2018 the U.S imposed new import tariffs on more than $250 billion of imports, with rates varying between 10% and 50%, being China the most affected country. As a consequence, China retaliated with tariff increases on almost $95 million of U.S exports, which represents more than 70% of total Chinese imports from the US (see UNCTAD, 2018). The specialized media and academic literature have denominated these circumstances as a ‘trade-war’.

In spite of the short time span since the beginning of the trade-war, there is already a body of literature analyzing its consequences for the global economy. Much of this literature analyzes the effects of tariffs on domestic prices, output, welfare, international trade, and other domestic variables for the U.S and/or China (see e.g., Amiti et al., 2019; Fajgelbaum et al., 2020 and Bolt et al., 2019, and the references therein). The consensus of this literature is that the trade-war has increased domestic prices and reduced output, trade, and aggregated welfare.

Analysis at the firm-level are consistent with the aggregate effects: tariffs have had a negative impact on market returns, sales, and employment levels on firms exposed to international trade and value chains (see Huang et al., 2019 for an analysis

* Corresponding author.
*E-mail addresses:* gcarlomagno@bcentral.cl (G. Carlomagno), ealbagli@bcentral.cl (E. Albagli).

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focused for U.S and Chinese firms). Additionally, the results of Crowley et al. (2018) show that uncertainty about future tariffs negatively affect the decision of Chinese firms to enter into new export markets.

This paper aims to quantify and interpret the impact of the trade-war on global financial markets. We believe this is an important endeavor for at least three reasons. First, being forward looking, asset markets can provide valuable information about the present value of output disruptions from the potential regime change of international trade relations brought about by the events following the election of President Trump. Second, it is not obvious even what the sign of such effects ought to be for the different parties affected, especially countries not directly involved in the trade war–i.e., some may even benefit from trade diversion. Thirdly, despite the change in US administration following the November 2020 election, many commentators fear that the trade war ignited in 2018 may prove to be the spark of a longer-lasting paradigm change in international trade relations, away from low tariffs and a multilateral rules-based system[1]. In this sense, gaining insights about the impact of the trade war on global financial markets may prove valuable beyond a mere ex-post explanation and quantification of one of the dominant forces shaping markets in between 2018 and the outset of the COVID 19 crisis.

Compared with the empirical literature on the real effects of trade disruptions, the evidence about the impacts of the trade war on financial markets is rather scarce. With a specific interest on the U.S, Burggraf et al. (2020) perform a simple event study on the effects of Trump's tweets on the S&P and the VIX, and find significant negative effects on the former, and positive on the latter. Huynh and Burggraf (2020) analyze how the trade-war affected the co-movement of stock markets in U.S, China, and G7 economies. They find that the trade-war increased comovements at the left tails of return's distributions. Amstad et al. (2021) construct a trade sentiment index based on Chinese media, and use it to estimate the impact of trade sentiment on 60 global equity markets. With some heterogeneity related to the exposure to global value chains, the authors find that a negative trade sentiment shock has a negative impact on stock markets. Interestingly, among the most affected countries are those which could potentially benefit from trade diversion (Hong Kong, Taiwan, Korea, Indonesia, Vietnam, and Singapore). Matveev et al. (2020) find that the Trump's tweets related to the imposition of tariffs on Canada and Mexico lead to an appreciation of the U.S. dollar with respect to both currencies.

There is also an older strand of literature that analyzes the impact of trade fluctuations on financial markets. Several authors have focused on the importance of trade linkages in explaining asset prices co-movements between countries (see e.g., Forbes and Chinn, 2004 and Kali and Reyes, 2010), and how those linkages boost financial contagion of crises (see e.g.,Hernández and Valdés, 2001; Forbes, 2004 and Kali and Reyes, 2010). Finally, there is also research showing that trade relationships influence patterns of foreign portfolio allocations (see Agarwal et al., 2020, and the references therein).

Our paper contributes to this emerging literature in three key respects. First, instead of focusing on particular events of the trade-war –most papers focus on the effects of specific increases in tariffs–, our interest is in quantifying the impact of the unobservable 'trade-tension' generated by this war. To do so, we identify days with relevant trade-war news, and perform empirical analysis using a standard event-study methodology as well as one based on conditional heteroskedasticity (described in detail momentarily). Second, we do not restrict the analysis to the U.S and China, but consider a wider range of developed and emerging markets. Thirdly, we differentiate from other studies analyzing the effects of the trade-war on financial markets (see e.g., Burggraf et al., 2020 and the references therein) by increasing the variety of assets considered, which allows us to go further into the interpretation about the underlying forces triggered by the trade tension.

Based on the Peterson Institute for International Economics registry (PIEE) registry, we construct an indicator of 'news related to the conflict' and adopt two alternative methodological approaches to measure the mentioned impact. Our first and main empirical methodology is based on event studies, which implicitly assume that the news affect the level of trade tension. As a robustness, we follow the heteroskedasticity-based estimator proposed by Rigobon and Sack (2004). The empirical results obtained with both strategies are similar and suggest that the trade-tension generates an increase in the preference for safer assets: it reduces long yields in developed economies, reduces the price of equity all over the world, depreciates the currencies of emerging markets, and increases the VIX. These are the typical developments after a global risk-off shock (see Caballero and Kamber, 2019). Furthermore, the interpretation of risk aversion as the dominant force shaping markets in response to trade tensions –in contrast to, say, pessimism about global growth prospects– is also favored by considering a wider set of assets. In particular, we find that the effects on short-term rates, when statistically significant, are much smaller than on long-term ones, suggesting a larger role for movements in term-premia than changes in expectations about future monetary policy rates driven by economic growth and the associated inflationary pressures. This interpretation maps well with the structural shoks interpretation defined by the recent work of Cieslak and Pang (2021).

The magnitude of the effects can be substantial. We estimate that the effect of the typical event on 10-year bond yields is close to $-3$ bp for the US, $-1.6$ bp for developed central economies, and $-1.2$ for other developed economies. For the period we consider (Jan-03–2018/ Jan-31–2020), these figures imply a cumulative effect on 10-year bond yields of $-67$ bp, $-37$ bp, and $-28$ bp, for the US, central, and other developed economies, respectively. For the case of US, these magnitudes represent more than 70% of the reduction observed in the sample. For emerging economies, the typical effect on yields is $-0.55$ bp for those with investment grade, and +1.1 bp without investment grade. However the evidence is mixed across these countries and estimates are not always significant.

Regarding stock returns, the typical (cumulative) effects are close to $-1\%$ ($-23\%$) for the US and other developed economies, and $-0.6\%$ ($-14\%$) for emerging markets, a result that is compatible both with other empirical papers and with theo-

[1] See for example"The US-China trading relationship will be fraught for years to come", *The Economist*, May 11th 2019 edition.

retical models that take into account general equilibrium macroeconomic effects, but substantially higher than the direct *treatment* effects of tariffs considered by static models (see Amiti et al., 2021). Also, we find that exchange rates in emerging markets depreciate significantly, with the typical (cumulative) effect around 0.15% (32%). Finally, we find a strong effect of the trade-tension on the VIX (typical effect of +18%), but evidence on the WTI and gold prices is not robust.

While our methodological approach is best suited for establishing the short-term effects of trade tensions on financial markets, we also make some headway into its persistence by building on the dynamic impulse-response functions of the conditional heteroskedasticity methodology. Indeed, we find that most of the effects that are significant on-impact, do not reverse after the shock, at least within the next month, which suggests that the trade war may have generated long lasting consequences on the levels of financial assets. Nonetheless, we believe that more research is necessary to make strong claims about long-run effects on financial markets.

The rest of the paper is organized as follows. Sections 2 and 3 describe the methodology and data used, respectively. Section 4 presents our main results, and Section 5 includes robustness checks. Section 6 concludes.

## 2. Estimation strategies

As mentioned in the introduction, we construct an indicator of 'news related to the conflict' and adopt two alternative methodological approaches to measure the impact of the trade-tension on financial markets. In this section we describe the two approaches.

In both approaches we assume that the news–e.g., tweets of president Trump–are exogenous events to the financial markets. This assumption is standard in the literature (all the papers quoted above assume exogeneity[2]), and derives from the fact that these news are unpredictable by market participants, even a few seconds before they occur–a *tweeter silver lining* in terms of event study identification.

This unpredictability from market participants does not imply that news are serially independent. In fact, it is actually the case that some of the news we consider are evident retaliations to threats coming from the other side of the war. Moreover, it is probably the case that all of the announcements are part of a general strategic behavior of actions and reactions. This does not invalidate our exogeneity assumption, as it only requires that the timing of the news (and sign in the case of the event study) cannot be predicted by market participants with reasonable accuracy.

The main challenge for the empirical analysis relates to the fact that the trade-tension is unobservable. In the first methodological approximation, we tackle this issue by assuming that the level of trade-tension can take only two values: one high, on days with news (and possibly some days after), and one low, the rest of the days. Under this assumption we conduct an event study (Section 2.1).

In the second methodological approach, we adopt the heteroskedasticity-based estimator proposed by Rigobon and Sack (2004), and adapted by Rigobon and Sack (2005) to identify the effects of the risk of war between the U.S and Irak on financial markets (see also Wright, 2012). The main advantage of this strategy is that it allows us to identify the effects of the trade-tension without the need of quantifying it. As in the first approach, the identifying assumption also consists of assuming two states of the nature determined by the arrival of news. However, in this case, it is the variance of the trade-tension, not the level, that distinguishes the two states. Other important advantages of this approach are that the identification assumptions are empirically testable, and that it is robust to inaccuracies in the determination of days with news. To the best of our knowledge, this is the first attempt to apply this estimator to measure the impact of the trade-war on any economic variable.

### 2.1. Event study

The event study approach is based on two assumptions: (a) news are exogenous, and (b) there are only two states of the nature for the level of trade-tension, one high on days with news and other low for the other days. In Appendix A we show that under those assumptions the difference in the expected returns in days with news *vs.* days without news represents the total (direct plus indirect) causal effect of the trade-tension on the expected returns. As Eq. (A.8) indicates, an estimator for this causal effect can be obtained as the difference between the mean returns in days with and without news.

The main advantage of this methodological approach is that event studies have an intuitive interpretation, and can easily summarize the effects of the typical event included in the sample. Moreover, they can be tested with standard econometric approaches that can be performed either at the individual country level, as well as in panel regressions as useful measures of average effects. Given that we include a large sample of countries in our study, the ability of grouping them into standard categories (such as developed vs. emerging) and providing estimates for group averages is a useful way of dealing with the high dimensionality of the results, which is why we refer to the event study as as our 'baseline' methodology.

However, the event-study methodology also has some drawbacks. First, from the analysis Appendix A, it is evident that this approach requires all of the news to have the same 'sign' (i.e., all of the news must increase the trade-tension), so it requires a subjective determination of the news' signs. Although for 'big' news this may not be an issue, the sign of other

[2] See also Bianchi et al., 2021, who assume exogeneity to identify the impact of president Trump tweets criticizing the Fed on market expectations about future monetary policy

news could be crystal-clear only ex-post.[3] We tackle this problem by making an ad hoc classification of the news from the perspective of the US stock market. In this way, we classify news that can be expected to have a negative impact on US stock returns as negative, and those with expected positive impact as positive.

Secondly, events studies cannot distinguish the direct from the indirect effects –that is, those working indirectly through the effects of the event on other variables, which in turn affect a particular variable of interest. Thirdly, by assuming only two possible levels of trade-tension, the estimations represent an average of potentially highly heterogeneous changes of the trade-tension on financial markets. To overcome these concerns, the next subsection presents the heteroskedasticity-based estimator, which largely overcomes these limitations.

### 2.2. Heteroskedasticity-based estimation

The heteroskedasticity-based estimator proposed by Rigobon and Sack (2004) and Rigobon and Sack (2005) overcomes the difficulties of the previous one. As it is based on restrictions placed on second-moments of returns, it does not require a subjective classification of the trade tension events, nor does it restrict only two-levels of trade tensions, as it only requires that the volatility driving the distribution of returns be different in days with trade tension news as compared to regular days. Also, because it is based on a structural identification of the parameter of interest, it can distinguish between direct and indirect effects. These are advantages not only in comparison to event study approaches, but also to strategies using textual analysis to quantify the level of trade tension, which crucially depend on a correct quantification of the tension, and cannot distinguish direct and indirect effects.

Moreover, this strategy is robust to inaccuracies in the determination of days with news. This is an overriding property in our framework because our choice of the days with news is somehow arbitrary, so that a slightly different selection of days may also be sensible. Rigobon and Sack (2005) show analytically that, provided that selection is not so misguided to remove the heteroskedasticity, changes to the definition of days with news and/or days without news does not affect point estimates. In Section 5 we change the set of war and non-war days and show that estimations remain close to the original ones.

In what follows we describe how we adopt this estimator to measure the impact of the trade-tension on financial markets. To the best of our knowledge this is the first attempt in the literature to apply this kind of estimator to quantify the impact of the trade-war. Other papers regarding the effects of the trade-war on financial returns have adopted more limited approaches, for a narrower set of assets and countries (see e.g., Burggraf et al., 2020 and the references therein).

Implementing this strategy requires defining a set of 'trade-war-days' and a set of 'non-trade-war-days', and two basic assumptions: (a) All parameters are stable in those sets of days, except for the variance of the unobservable trade-tension ($\omega_t$). We assume that this parameter can take only two values: one high, $\sigma^2_{\omega,H}$, which takes palace in trade-war-days, and one low, $\sigma^2_{\omega,L}$, that takes place the non-trade-war-days. (b) The trade-tension factor is exogenous (same as in Section 2.1).

We define trade-war-days as the days with news (without subjectively classifying them as positive or negative), and non-trade-war-days as the closest days previous to the days of the news. The starting point of this methodological approach is similar to that of the event study (Eq. (A.1)) but for the returns of two assets ($r_{1t}$ and $r_{2t}$):

$$\begin{bmatrix} r_{1t} \\ r_{2t} \end{bmatrix} = \begin{bmatrix} c_1 \\ c_2 \end{bmatrix} + \begin{bmatrix} \beta_1 & \lambda_1 \\ \beta_2 & \lambda_2 \end{bmatrix} \begin{bmatrix} \omega_t \\ X_t \end{bmatrix} + \begin{bmatrix} e_{1t} \\ e_{2t} \end{bmatrix}, \tag{1}$$

where $e_{it}$ is a random noise independent of $\omega_t$ and $X_t, \forall\, t$. $\beta_i$ is a coefficient that captures the effect of the trade-tension on the return $r_i$, and $\lambda_i$ is a $1 \times N$ vector of coefficients that represents the effect of the $N$ variables included in vector $X_t$ on $r_{it}$.

Let $Y_t = [\omega_t, X_t]'$, and define $\mathbb{1}_{n,t}$ as the indicator function that takes the value 1 in days with news, and $\Sigma_{Y,w}$ and $\Sigma_{Y,\bar{w}}$ as the covariance matrix of $Y_t$ in trade-war-days and non-trade-war-days, respectively. The the identifying assumption a) can be formalized in the following two conditions:

$$Var([e_{1t}, e_{2t}] | \mathbb{1}_{n,t} = 1) = Var([e_{1t}, e_{2t}] | \mathbb{1}_{n,t} = 0), \text{ and} \tag{2}$$

$$\Sigma_{Y,w} - \Sigma_{Y,\bar{w}} = \begin{bmatrix} \rho & 0 & 0 & \dots & 0 \\ 0 & 0 & 0 & \dots & 0 \\ 0 & 0 & 0 & \dots & 0 \\ \vdots & & & & \\ 0 & 0 & 0 & \dots & 0 \end{bmatrix}_{(N+1)\times(N+1)}, \tag{3}$$

where $\rho$ represents the difference in the variance of the trade-tension factor in trade-war-days with respect to non-trade-war-days.

Define now the covariance matrix of the returns $r_1$ and $r_2$ for trade-war-days and non-trade-war-days as $\Omega_w$ and $\Omega_{\bar{w}}$, respectively. Then, from expressions (1)–(3), it follows that:

[3] As an example, consider the following case: On February the $2^{nd}$ 2018 the Chinese government initiated anti-dumping investigations on U.S sorghum exports, is this a positive or negative news for Chinese financial markets?

$$\Delta\Omega \equiv \Omega_w - \Omega_{\bar{w}} = \rho \begin{pmatrix} (\beta_1)^2 & \beta_1\beta_2 \\ \beta_1\beta_2 & (\beta_2)^2 \end{pmatrix}. \tag{4}$$

Given that $\omega_t$ is an unobserved continuous variable, a normalization is required for identifying our parameters of interest ($\beta_i$). Following Rigobon and Sack (2005), we impose $\beta_1 = 1$. Thus, $r_{1t}$ is to be interpreted as the *return of reference*, and our parameter of interest ($\beta_2$), as the response of $r_2$ to the trade-tension factor *relative* to the response of $r_1$. In the empirical application of Section 4, we use the same return of reference for all of the variables we consider. Using this normalization, we get:

$$\Delta\Omega = \rho \begin{pmatrix} 1 & \beta_2 \\ \beta_2 & (\beta_2)^2 \end{pmatrix} \tag{5}$$

Hence, the change in the covariance matrix of the returns can be used to estimate $\beta_2$. Intuitively, we may compute the change in the covariance matrix in the sample ($\Delta\hat{\Omega}$), and get an estimator of $\beta_2$ ($\hat{\beta}_2$) either as: (a) $\frac{\Delta\hat{\Omega}_{22}}{\Delta\hat{\Omega}_{21}}$, or as (b) $\frac{\Delta\hat{\Omega}_{21}}{\Delta\hat{\Omega}_{11}}$, where $\Delta\hat{\Omega}_{ij}$ represents the element $(i,j)$ of $\Delta\hat{\Omega}$.

As Rigobon and Sack (2004) show, this estimator can be obtained either by instrumental variables or GMM. For the IV estimation, define the instrument $\kappa_{1t}$, as $r_{1t}$ in trade-war-days, and minus $r_{1t}$ in non-trade-war-days. As the authors show, the estimator $\hat{\beta}_2$ indicated in option b above can be obtained by means of a IV regression of $r_{2t}$ on $r_{1t}$ using $\kappa_{1t}$ as the instrument[4].

In the GMM estimation, parameters $\rho$ and $\beta_2$ are chosen to minimize:

$$[vech(\Delta\hat{\Omega} - \rho\beta\beta')]' W [vech(\Delta\hat{\Omega} - \rho\beta\beta')], \tag{6}$$

where $\beta = [1, \beta_2]'$.

For the sake of brevity we do not include more estimation details here. Full details can be found in Sections 3.1 and 3.2 of Rigobon and Sack (2004).

## 3. Data

### 3.1. Assets and countries

Our sample consist of 36 countries, four country-specific assets, and three commodities and global risk indicators. The countries are classified in four groups: Developed central, other developed, emerging with investment grade, and emerging without investment grade (see details in Table 1). In all cases we use daily data for the period January-03–2018/ January-31–2020.[5] For sovereign bonds, returns are defined as the difference of yields (the price change with inverted sign), for the other assets, returns are defined as the log difference of the levels.

### 3.2. News related to the trade-war

We use the record of trade-war-related news collected by the PIEE.[6] In the period under analysis, the record contains a total of 75 news, which the PIIE classifies in five different types of *battles* (as they call them) related to the trade-war. We focus on the battle about *Unfair Trade Practices for Technology and Intellectual Property*, which contains the most relevant news of the trade-war between the U.S and China. However, as some definitely significant news are included in other battles (e.g., the kickoff of the trade-war on March-22–2018) we also include some other news.

Table 2 summarizes the list of news we consider in the estimations. The first column indicates the date of the news, the second one specifies whether we classified it as positive or negative, and the third column includes a brief description of the event. Shadowed dates indicate news that appeared during a weekend, which we assign to the following Monday. We have a total of 27 news, out of which 23 are classified as negative, and 4 as positive.

Figs. 1 and 2 illustrate the relevance of these news for U.S financial markets. Almost all of the big falls of the 10-year bond yield since January-2018 are related to the trade-war. This observation is also true for the $S\&P500$: With the exception of the sell-off period that took place at the end of 2018, almost all of the big falls since January-2018 are related to the trade-war.

[4] Rigobon and Sack (2004) also show that the an alternative to this strategy is defining an instrument $\kappa_{2t}$, which is analogous to $\kappa_{1t}$, but for $r_{2t}$. Then the estimator $\hat{\beta}_2$ would be obtained by means of a IV regression of $r_{1t}$ on $r_{1t}$ using $\kappa_{2t}$ as the instrument. In order to avoid the heterogeneity that may arise due to variations in the instruments' quality across counties, in this paper, we do not explore this possibility.

[5] After January 2020, due to the beginning of the COVID crisis, global interest rates were heavily affected by the aggressive conventional and unconventional monetary policies deployed by central banks, specially at the US. In particular, the zero lower bound was a biding restriction for short rates. This may affect our estimations, so that we decided not to use data beyond that period.

[6] See the PIIE registry here.

**Table 1**
Assets and countries in the sample.

| | |
|---|---|
| **Developed central economies** | Germany, Japan, UK, US. |
| **Developed not central economies** | Canada, France, Hong Kong, Italy, Norway, Spain, Sweden, Switzerland. |
| **Emerging with inv. grade** | Bulgaria, Chile, China, Colombia, Croatia, Czech Rep., India, Indonesia, Israel, Malaysia, Mexico, Peru, Poland, Romania, Russia, South Korea, Thailand. |
| **Emerging without inv. grade** | Brazil, Hungary, Nigeria, South Africa, Sri Lanka, Turkey, Ukraine, Vietnam. |
| **Country specific assets** | Yield of the 10 year government bond, Yield of the 1 year government bond, stock market returns (S&P for the US), exchange rate in local currency per US dollar. |
| **Commodities and risk indicators** | WTI price, Gold price and VIX. |

**Table 2**
News related to the trade-war.

| | Sign | Event |
|---|---|---|
| 22-01-2018 | - | Trump imposes safeguard tariffs |
| 05-02-2018 | - | China investigates US sorghum exports |
| 16-02-2018 | - | Results of the national security investigation |
| 01-03-2018 | - | Tariff announcement |
| 22-03-2018 | - | More tariff exemptions<br>Research results on unfair business practices |
| 23-03-2018 | - | Rates are effective |
| 03-04-2018 | - | USA threatens tariffs |
| 04-04-2018 | - | China threatens retaliation against cars, planes and agriculture |
| 05-04-2018 | - | USA Consider additional $ 100 billion tariffs |
| 29-05-2018* | - | White house plans rates after short inertia |
| 18-06-2018 | - | Trump calls for more tariffs |
| 22-06-2018 | - | EU retaliates on iconic American products |
| 20-07-2018 | - | Trump threatens tariffs on all imports from China |
| 01-08-2018 | - | Trump wants 25% tariffs instead of 10% |
| 03-12-2018 | + | Agreement in the G20 between Trumnp and Xi |
| 25-02-2019 | + | Rate increase delayed |
| 06-05-2019 | - | Trump tweets intention to raise the tariffs on $200 bn of Chinese goods to 25% |
| 10-05-2019 | - | USA increase tariff rate in the previous list |
| 13-05-2019 | - | China plans to escalate the tariff rate |
| 01-08-2019 | - | USA Announces Tariff Increase on Almost All Remaining Imports from China |
| 05-08-2019** | - | China halts purchases of U.S. agricultural products |
| 13-08-2020 | + | Trump postpones some of the 10% tariffs on the $300 billion goods list until Dec. 15 |
| 23-08-2019 | - | China announces new tariffs for the US |
| 20-09-2019** | - | Chinese trade officials canceled farm visits and Trump called the nation a threat |
| 07-10-2019** | - | The U.S. Commerce Department puts 28 Chinese companies on its “entity list” |
| 11-10-2019 | + | Trump Cancels October Tariffs, Points to “Phase One” of Deal with China |
| 24-01-2020 | - | Trump Broadens Tariffs |

Notes:
- * In this date there is no data for the *S&P*500 so it is excluded in the analysis of this asset.
- ** Theses dates are not included in the PIIE registry, but are highlighted by Reuters as relevant events (see the Reuters’ time line).
- Shadowed dates indicate news that appeared during a weekend, which we assign to the following Monday.

## 4. Empirical results

In this section, we present the results of the event study and of the heroskedasticity-based estimator. As we show, the orders of magnitude of both estimations are similar, which we interpret as an indication that our results are robust to modifications of the identification strategy.

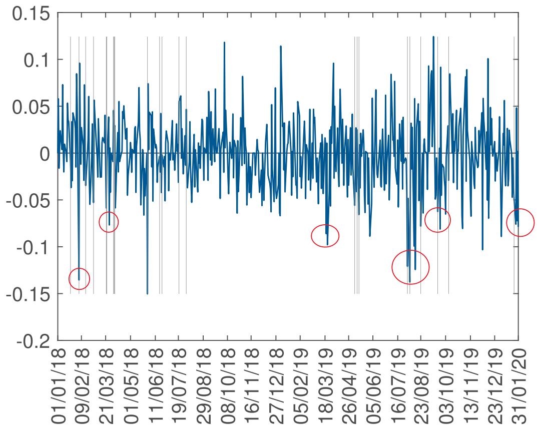


**Fig. 1.** Impact of the trade-war on U.S 10-year treasury bond yield (daily variation in basis points). Period Jan-03–2018/ January-31–2020. Notes: - Most relevant news marked in red circles - Date March-22–2019 is marked in red but is not included in PIIE's record. This date there was a tweet by president Trump related to the suspension of sanctions to North Korea which, according to press releases, generated turmoil in the financial markets due to its relation with Chinese cargo ships. However, given that this event is not included in our source, Table 2 doe not include it, and we do not consider it for the statistical analysis of next section. Adding this event to the analysis amplifies the estimated effects of the trade-war on financial markets, these results are not reported but are available upon request.

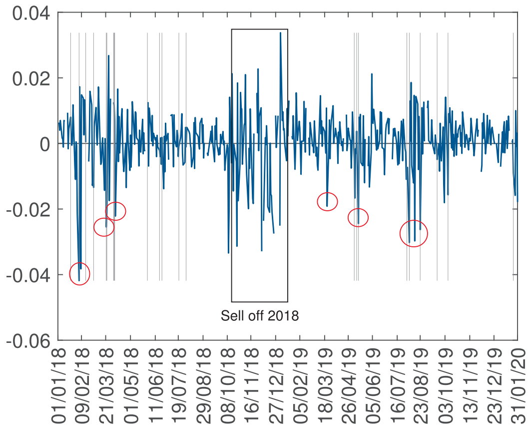


**Fig. 2.** Impact of the trade-war on the S&P (daily variation in percentage). Period Jan-03–2018/ January-31–2020. Notes: - See notes to Fig. 1.

### 4.1. Event study

We now present the outcomes of applying the strategy described in Section 2.1, as well as some extensions of empirical interest. The difference in expected returns between days with and without news can be estimated as the OLS estimator of $\gamma$ in the following equation:

$$r_{it} = c_i + \gamma \mathbb{1}_{n,t} + r^*_{it}. \tag{7}$$

One limitation of not using intra-day data is that news may occur after markets' closure, what is especially relevant for eastern economies when news are originated in the US. If all the news occurred after the markets close, the news' indicator in Eq. (7) should be lagged one period. In Appendix B we show that, in order to cover both possibilities (news coming before and after markets' closures), we can analyze the returns accumulated in two days as in the following equation:

$$R_{it} = c_i + \gamma \mathbb{1}_{n,t-1} + R^*_{it}, \tag{8}$$

where $R_{it} = r_{it} + r_{it-1}$ (see lines corresponding to DGPs iii and iv of Table B.1 in Appendix B).

In what follows we focus on Eq. (8) (results for Eq. (7) are included in Appendix C) and adopt two estimation strategies: (a) Estimate Eq. (8) as a panel regression using countries as the cross-sectional units, so that we will have four separate regressions; one for the 10-year bond, one for the 1-year bond, one for the stock market, and another one for the exchange rate. We report these results in Table 3. (b) Estimate Eq. (8) as an individual regression for each asset of each country, and for each global asset. These results are included in Table 4.

**Table 3**
Mean effects of a higher trade-tension from panel regressions (in bps for sovereign bonds and percentage points for the rest of the assets).

| | (−) | (+) | all |
|---|---|---|---|
| | *10y Bond* | | |
| Dev. Centr | −1.58*** | −0.89* | −1.49*** |
| Dev. Not Centr | −1.23*** | −2.37 | −1.42*** |
| EME Inv Gr. | −0.55 | −0.84 | −0.59 |
| EME Not Inv Gr. | 1.11 | −1.2 | 0.75 |
| All | −0.51 | −1.31** | −0.63* |
| | *1y Bond* | | |
| Dev. Centr | 0.05 | −0.52 | −0.04 |
| Dev. Not Centr | 0.12 | −0.66** | −0.01 |
| EME Inv Gr. | −0.08 | 0.23 | −0.03 |
| EME Not Inv Gr. | 1.97 | 0.37 | 1.73 |
| All | 0.37 | −0.06 | 0.3 |
| | *Stock Mkt* | | |
| Dev. Centr | −1.08*** | −0.09 | −0.95*** |
| Dev. Not Centr | −0.99*** | 0.17 | −0.82*** |
| EME Inv Gr. | −0.57*** | 0.65*** | −0.39*** |
| EME Not Inv Gr. | −0.56*** | 0.18 | −0.46*** |
| All | −0.72*** | 0.36*** | −0.56*** |
| | *Exch. Rate* | | |
| Dev. Centr | −0.09 | −0.22 | −0.11 |
| Dev. Not Centr | −0.02 | −0.14 | −0.04 |
| EME Inv Gr. | 0.14*** | −0.18* | 0.09** |
| EME Not Inv Gr. | 0.12* | 0.09 | 0.11* |
| All | 0.09** | −0.11 | 0.06** |

Notes:
Columns (−), (+), and *all* include results for negative, positive, and all news, respectively.
Asterisks (***, **, *) indicate statistical significance at the 1%, 5% and 10%, respectively.
See Table 1 for the definition of the country groups.

Focus first on Table 3. An increase in the trade-tension associated with negative news (columns '(−)' of the table) has a negative and highly statistically significant effect on 10-year bond yields in developed countries. The effect on long yields of central economies is 1.3 times as large as that for the other developed: −1.58 bp and −1.23 bp, respectively. In emerging markets, on the other hand, effects are not statistically significant. Short yields (1 year) do not seem to be significantly affected by the trade tension in any group of countries.

Regarding stock markets, the effect on all groups is negative and statistically significant. The largest effects take place in central economies (−1.08%), followed by the other developed countries (−0.99%), and emerging economies (−0.57%).

Exchange rates are defined as USD per domestic currency–so that a higher exchange rate stands for a depreciation of the domestic currency against the USD. While there are no statistically significant deviations with respect to the USD for developed countries, emerging countries exhibit currency depreciations that are significant both for countries with investment grade (IG) and without it (non-IG).

Regarding positive news, they are not statistically significant in general, except for the stock markets of emerging economies with investment grade, for which we find a significant positive effect. However, as we show below, this result is driven by a few particular cases. It is noteworthy that the general irrelevance of positive news, plus its relative small number, implies that considering only negative news or all of the news at the same time makes little difference (compare columns '(−)' and 'all' of Table 3).

Table 4 confirms that previous conclusions are also valid when considering countries individually. The higher trade tension (columns '(−)' of the table) reduces the 10-year bond yields of central and developed economies. Among emerging markets, we find a marginally significant compression in yields for some IG countries, while the impact on non-IG are generally positive though not significant. Consistently with this observation, Italy is the only developed economy in which bond yields reacts positively to trade tensions, and is also the lowest rated country in the category. Yields on 1 year bonds show non-significant reactions to the trade tension, except for a fall in US bonds, but in smaller magnitude compared to movements in long-term yields.

Stock markets react to higher trade tensions negatively in all of the countries considered, with the unique exceptions of India, Brazil and Ukraine which show positive but non-significant effects. In general, the effects are higher in developed than in emerging markets. In the case of China, and consistent with being the target of the war, its stock market reaction (−1.57%) is one of the largest in the entire sample (for the US the effect is −0.97%).

Exchange rates of developed markets show mixed, but rather small reactions to the trade tension. The Japanese yen constitutes the only exception: consistent with the safe-haven role usually attributed to the said currency, it appreciates 0.34%

**Table 4**
Mean effects of a higher trade-tension from individual regressions (in bps for sovereign bonds and percentage points for the rest of the assets).

| | 10 year Bond | | | Stock Mkt | | | Exch. Rate | | | 1 year Bond | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | (−) | (+) | all | (−) | (+) | all | (−) | (+) | all | (−) | (+) | all |
| *Developed Central* | | | | | | | | | | | | |
| Japan | 0.03 | 0.38 | 0.09 | −1.33*** | −0.61 | −1.27*** | −0.34*** | 0.05 | −0.29** | 0.06 | 0.21 | 0.08 |
| UK | −1.8 | −0.42 | −1.6 | −0.8** | −0.02 | −0.67** | 0.19 | −0.64* | 0.07 | −0.52 | 0.9 | −0.3 |
| Germany | −1.61* | −1.51 | −1.61** | −1.23*** | 0.69 | −0.94*** | −0.11 | −0.08 | −0.1 | 0.08 | 0.14 | 0.09 |
| US | −2.93** | −2 | −2.82** | −0.97*** | −0.7 | −0.93*** | | | | −2.48*** | 1.39 | −1.91** |
| *Developed* | | | | | | | | | | | | |
| Canada | −2.44** | −1.36 | −2.3** | −0.55*** | −0.6 | −0.56*** | −0.05 | 0.02 | −0.04 | −0.31 | 0.91 | −0.12 |
| Italy | 1.18 | −10.95** | −0.67 | −1.05** | 0.64 | −0.79** | −0.11 | −0.08 | −0.1 | 3.41* | −7.69* | 1.74 |
| France | −0.92 | −2.12 | −1.11 | −1.1*** | 0.27 | −0.89*** | −0.11 | −0.08 | −0.1 | −0.05 | −0.53 | −0.12 |
| Norway | −2.23** | 0.3 | −1.86** | −0.54* | −0.1 | −0.47 | 0.06 | −0.4 | −0.01 | 0.33 | 0.02 | 0.24 |
| Sweden | −1.51 | 3.63* | −0.74 | −1.23*** | −0.4 | −1.11*** | 0.05 | −0.34 | −0.01 | −0.26 | 0.7 | −0.12 |
| Switzerland | −1.59** | −2.81 | −1.79** | −1.07*** | 0.27 | −0.87*** | −0.19 | 0.08 | −0.15 | 0.04 | −1.03 | −0.13 |
| Hong Kong | −2.02 | −1.03 | −1.93 | −1.83*** | 0.98 | −1.3*** | 0.02* | −0.06* | 0.01 | – | – | – |
| Spain | −0.82 | −3.6 | −1.26 | −0.75** | 0.1 | −0.62* | −0.11 | −0.08 | −0.1 | 0.59 | −0.2 | 0.47 |
| *Emerging with investment grade* | | | | | | | | | | | | |
| China | −1.79** | 0.83 | −1.3* | −1.57*** | 2.25*** | −0.84** | 0.4*** | −0.8*** | 0.18** | – | 1.56 | −0.67 |
| South Korea | −2.04 | – | −2.04 | −1.26*** | 0.66 | −0.94*** | 0.54*** | −0.82*** | 0.34*** | – | 0.54 | −0.94 |
| Czech Rep. | −2.25** | 0.8 | −1.8* | −0.47* | 0.68 | −0.29 | −0.02 | −0.08 | −0.03 | −1.03 | −0.78 | −1 |
| Chile | −2.56 | 3.69 | −1.86 | −0.8** | 0.41 | −0.62** | 0.52** | −0.64 | 0.34* | −1.66 | 3.03 | −0.09 |
| Thailand | −1.34 | −1.37 | −1.35 | −0.78*** | 1.18 | −0.61** | −0.03 | −0.19 | −0.05 | −0.57 | −3.17* | −0.8 |
| Poland | −2.82** | 0.32 | −2.35* | −0.97** | 0.8 | −0.69* | 0.01 | −0.11 | −0.01 | −0.18 | 0 | −0.15 |
| Mexico | −4.06** | −3.07 | −3.95** | −0.13 | −0.42 | −0.18 | −0.02 | 0.03 | −0.01 | −2.7 | 4.24 | −1.66 |
| Israel | −0.27 | – | −0.27 | −1.05*** | −1.96 | −1.14*** | 0.02 | 0.52** | 0.09 | 0.06 | −2.53 | −0.08 |
| India | −2.02 | −3.41 | −2.11 | 0.15 | 0.41 | 0.18 | 0.19 | 0.43 | 0.22** | −1.33 | 1.07 | −0.73 |
| Indonesia | 3.73* | −3.21 | 2.67 | −0.48* | 1.12* | −0.23 | 0.23** | −0.19 | 0.16 | 5.58** | 2.26 | 5.12** |
| Malaysia | −0.16 | −1.22 | −0.3 | −0.56*** | 0.66 | −0.41** | 0.13* | −0.36* | 0.06 | −0.08 | 0.29 | −0.03 |
| Colombia | 1.92 | −5.67 | 1.3 | −0.38 | 1.39** | −0.16 | 0.41* | −0.7 | 0.24 | – | – | – |
| Romania | −1.07 | −7.43 | −1.57 | −0.47 | 0.03 | −0.41 | −0.1 | −0.12 | −0.1 | 0.6 | −2.89 | 0.15 |
| Peru | 0.45 | 0.44 | 0.45 | −0.57** | 0.65 | −0.39 | 0.1 | 0.07 | 0.1 | −0.21 | 0.74 | −0.06 |
| Russia | 2.75 | −0.84 | 2.16 | −0.5 | −0.18 | −0.45 | 0.29 | 0.14 | 0.27 | 1.95 | −2.52 | 1.21 |
| Croasia | −0.27 | −0.04 | −0.23 | −0.1 | 0.81 | 0.05 | −0.06 | −0.06 | −0.06 | −0.2 | −2.43 | −0.57 |
| Bulgaria | 1.06* | 2.58* | 1.31** | −0.28 | 0.92** | −0.07 | −0.11 | −0.08 | −0.1 | −0.34 | 0.69 | −0.17 |
| *Emerging without investment grade* | | | | | | | | | | | | |
| Hungary | −0.82 | −4.53 | −1.4 | −0.75** | 0.87 | −0.5 | −0.03 | −0.15 | −0.05 | 0.14 | −0.35 | 0.02 |
| South Africa | 3.74 | −1.63 | 2.41 | −1.03*** | 0.41 | −0.8** | 0.36 | −0.6 | 0.22 | – | – | – |
| Brazil | 2.23 | 2.58 | 2.31 | −0.36 | −0.12 | −0.33 | 0.45* | 0.29 | 0.43** | −3.42 | 2.22 | −2.59 |
| Vietnam | 1.89 | −1.55 | 1.34 | −0.83** | 0.81 | −0.56* | 0.06*** | −0.05 | 0.05** | 3.22 | −4.16 | 1.98 |
| Turkey | 2.51 | −4.83 | 1.37 | −0.85** | −1.58 | −0.94** | 0.09 | 1.2 | 0.26 | 3.26 | −0.32 | 2.75 |
| Nigeria | – | – | – | −0.49 | 0.21 | −0.4 | −0.04 | −0.25 | −0.08 | −7.84 | −22.49 | −11.33 |
| Sri Lanka | −2.77 | −2.58 | −2.77* | −0.22 | 0.26 | −0.16 | 0.04 | 0.13 | 0.05 | −1.92 | −0.8 | −1.8 |
| Ukraine | 3.75 | 1.77 | 3.44 | 0.06 | 0.17 | 0.08 | −0.02 | 0.14 | 0 | 19.11* | −2.8 | 13.74 |
| *Global Assets* | | | | | | | | | | | | |
| | (−) | (+) | all | | | | | | | | | |
| VIX | 10.15*** | 2.36 | 8.97*** | | | | | | | | | |
| WTI | −0.57 | 0.71 | −0.36 | | | | | | | | | |
| GOLD | 0.36 | 0.18 | 0.33 | | | | | | | | | |

Notes:
See notes to Table 3.

on average during negative news days. In emerging markets, exchange rates reactions range from almost none, to depreciation between 0.3% and 0.54% in cases the cases of China, Colombia, Brazil, Chile and South Korea.

Finally, as the last line of Table 4 shows, the increase in the trade tension significantly increases the volatility of financial markets as measured by the VIX. The effects on the prices of the WTI and Gold have the expected signs (negative and positive, respectively), but are not statistically significant.

Regarding positive news, they are typically not statistically significant, probably because of its small number. The cases of the Italian 10-year bond yield, and the Chinese stock market and exchange rate, are the most remarkable exceptions. While the Italian bond yield decreases on average 11 bp after positive news, Chinese stock market increases 2.25% and the Yuan appreciates 0.8%.

Overall, these result indicate that, by pushing down long-term yields of central economies and stock markets all over the world, depreciating currencies of emerging markets, and increasing the VIX, the higher trade-tension seems to trigger the typical developments of a global risk-off shock (see Caballero and Kamber, 2019). However, these asset prices reactions may also be compatible with a negative shock to the economic outlook caused by a deterioration of the world trade, at least for some countries.

As mentioned in the introduction, we try to disentangle these two channels by mapping our results to the structural shocks defined by Cieslak and Pang (2021). The authors define four structural shocks that drive movements in local yields and stock market returns: (i) a *growth (economic outlook) shock*, after which returns and yields move in the same direction, and the effect on short yields is larger than on long-term yields, (ii) a *monetary policy shock*, after which returns and yields move in the opposite direction, and the effect on short-term yields is also larger than on long-term yields, (iii) a *risk premium-hedging shock* (risk-off shock), after which returns and yields move in the same direction, but now the effect on long-term yields is larger than on short-term yields, and (iv) a *risk premium-common shock*, after which returns and yields move in the opposite direction, and the effect on long-term yields is again larger than on short-term yields.

Therefore, using the definitions of Cieslak and Pang (2021), we may distinguish the channel through which the trade war affects asset prices by comparing the effects of the trade tension on long and short-term rates. Specifically, the risk-aversion interpretation should be favored if the effect of the trade tension is comparably larger on long-term yields, while the economic outlook shock would be the preferred interpretation if the converse happens. Since our results consistently show largely economically insignificant reaction in short-term yields, and when they are they tend to be significantly smaller than movements in long-term yields, they favor increases in risk-aversion as the dominant interpretation behind the effects of trade tensions. This is reinforced by the fact that countries that could potentially benefit from trade diversions (Souuth Korea, Thailand, Indonesia, and Vietnam) also show a negative stock market reaction. Moreover, this interpretation is also consistent with the results of Amstad et al. (2021), who analyze stock markets reactions to a higher trade tension and show that some of the most affected countries are those which could potentially benefit from trade diversions (Hong Kong, Taiwan, South Korea, Indonesia, Vietnam, and Singapore).

Interestingly, among emerging markets, there are cases for which while stock markets fall, bond yields do not show significant movements, so that the shock cannot be directly mapped to Cieslak and Pang (2021) definitions. However, as currencies generally depreciate (exchange rate is not included in Cieslak-Pang model), we interpret the shock as a risk premium-hedging shock in which the safe assets are central economies' bonds, as opposed to local ones.

There are yet other emerging markets for which stock markets fall and exchange rates depreciate, but yields rise. For these countries, it is generally the case that long yields rise more than short rates do, which again tends to favor the risk-aversion hypothesis. Thus, while one cannot rule out some role for worsening economic prospects as relevant drivers of financial returns throughout the trade war, the available evidence consistently favors the risk-aversion interpretation as the dominant force, both for developed and emerging economies.

Finally, it is important to note that although our results are quantitatively similar to other in empirical papers, they are orders of magnitude higher than those which would be compatible with the real effects predicted by classical trade models. Amiti et al. (2021) tackle this puzzle in detail. First, they show that movements in stock prices can be linked to expected movements in productivity and aggregate welfare. Next, they use daily and three-day stock market reactions to trade-war events to estimate the effects of those events on expected productivity and welfare. The authors find an average US stock market decline close to 1% after each trade-war event, very similar to our results. Considering the set of twelve events they include in the analysis (we include 23), they find that this figure is compatible with an expected decline in total factor productivity and aggregate welfare of 9.5% and 7.8%, respectively.

Very importantly, Amiti et al. (2021) decompose these effects into a *general macro effect*, that affects all companies in the same way, and a *treatment effect*, that affects companies depending on their exposure to the affected tariffs. They find that almost all of the effect of the trade war takes place through the general macro channel, as it accounts for 92% of the effect on welfare. As the authors argue, the small treatment effect they find is compatible with results from classic trade models, which are static and do not capture general dynamic macroeconomic effects or increased uncertainty. In fact, they calibrate the dynamic model of trade by Perla et al. (2021) to map the stock prices movements caused by the trade war onto welfare effects, and find a result close to the aforementioned 7.8%.

This is consistent with our finding that the trade war generated negative stock market reactions even in countries which could be expected to benefit from trade diversions–a result also found by other authors (see e.g., Amstad et al., 2021)–, which suggests that expected dynamic macroeconomic effects through higher uncertainty, lower productivity, higher prices, lower wages, and lower welfare may dominate the direct effects of tariffs.

All in all, our estimated magnitudes are compatible both with other empirical papers and with theoretical models that take into account general equilibrium macroeconomic effects, which tend to be substantially higher than the direct *treatment* effects of tariffs considered by static models.

### *4.2. Heteroskedasticity-based estimation*

We now present the results of the estimates based on heteroskedasticity described in Section 2.2. As we mentioned there, the identification strategy requires using the return of an asset as a reference ($r_{1t}$). Following Rigobon and Sack (2005) we use the US 2-year bond yield for that purpose. Thus, for each asset in each economy, we consider the bi-variate model described in Section 2.2, including the yield of the US 2-year bond as the reference series.

**Table 5**
Estimates based on the IV implementation of heteroskedasticity based estimator (normalized to cause a drop of 3.79 bp of the 2-yr treasury yield). Average effects of increased trade tensions - in percentage points for exchanges and exchange rates and in base points for bonds.

| | 10y Bond | Stock Mkt | Exch. Rate | 1y Bond |
|---|---|---|---|---|
| | | *Developed central* | | |
| Japan | −0.03 | −0.35* | −0.33*** | 0.17 |
| UK | −1.5** | −0.44** | 0.13 | −0.79* |
| Germany | −0.89** | −0.38** | 0.03 | −0.42** |
| US | −3.82*** | −1.16*** | | −2.26*** |
| | | *Developed not central* | | |
| Canada | −2.06*** | −0.29** | 0.14* | −0.56* |
| Italy | 4.14*** | −0.45** | 0.03 | 4.74*** |
| France | −0.26 | −0.35** | 0.03 | 0.25* |
| Norway | −1.39*** | −0.19 | 0.26*** | −0.2 |
| Sweden | −0.5 | −0.36** | 0.16 | −0.03 |
| Switzerland | −0.56 | −0.52*** | −0.14** | −0.78* |
| Hong Kong | −0.5 | −0.38* | 0.02*** | – |
| Spain | 0.49 | −0.32** | 0.03 | 1.44*** |
| | | *Emerging with investment grade* | | |
| China | −0.49 | −0.38 | 0.16*** | −0.49 |
| South Korea | 0.62 | −0.27* | 0.17** | 0.95 |
| Czech Rep. | −0.32 | −0.38*** | 0.15* | −0.66 |
| Chile | −1.49 | −0.53*** | 0.31*** | −1.81* |
| Thailand | −0.4 | −0.3 | 0.06 | −0.17 |
| Poland | −0.02 | −0.54*** | 0.14 | 0.17 |
| Mexico | 0.52 | −0.25* | 0.39*** | 0.23 |
| Israel | −1.1 | 0.01 | 0.12* | 0.34 |
| India | 1.7** | −0.19 | 0.19** | – |
| Indonesia | 2.12 | −0.39** | 0.11* | 1.6 |
| Malaysia | 1.04 | −0.05 | 0.09* | 0.04 |
| Colombia | 2.35** | −0.68*** | 0.47*** | – |
| Romania | 0.28 | −0.24 | 0.06 | 1.23 |
| Peru | 0.17 | −0.37*** | 0.15*** | −0.03 |
| Russia | −0.05 | −0.12 | 0.28*** | −0.93 |
| Croasia | −0.44** | −0.02 | 0.07 | 0.54 |
| Bulgaria | −0.22 | −0.17* | 0.02 | 0.2 |
| | | *Emerging without investment grade* | | |
| Hungary | 0.51 | −0.44** | 0.12 | −0.21 |
| South Africa | 2.84** | −0.32* | 0.51*** | – |
| Brazil | 3.54** | −0.5** | 0.35** | 0.52 |
| Vietnam | −0.53 | −0.25 | 0.01 | −0.71 |
| Turkey | 1.54 | −0.38* | 0.13 | 3.11 |
| Nigeria | – | −0.05 | −0.07 | – |
| Sri Lanka | −2.01 | −0.16* | 0.03 | −1.35 |
| Ukraine | 2.71 | 0.1 | 0.07 | 9.81* |
| | | *Global Assets* | | |
| VIX | 18.63*** | | | |
| Oil | −1.61*** | | | |
| Gold | −0.09 | | | |

Notes:
- Asterisks (***, **, *) indicate statistical significance at the 1%, 5% and 10%, respectively.

Table 5 includes the results for the IV strategy. Results are normalized to the average drop in the yield of the 2-year treasury bill in days with negative news (3.8 bp). Thus, reported coefficients represent the impact of an increase in the trade-tension that generates a reduction in the yield of the 2-year treasury bill of 3.8 bp.

The first relevant observation is that signs and orders of magnitude coincide with those of Table 4. Consider, for example, the case of the US: While in Table 4 we reported that an increase in the trade-tension reduces the 10-year treasury yield by 2.93 bp, and the S&P by 0.97%, Table 5 indicates that an increase in the trade tension in a magnitude that generates an *average* drop on the 2-year treasury yield, reduces the 10-year yield in 3.8 bp, and the S&P in 1.2%.

For central economies we find that the effect of the trade tension on long rates almost doubles that on short ones. Again, Italy appears as a special case; it is the only developed economy whose long yield increases due to the higher trade tension. As in Table 4, for emerging markets, the evidence is mixed, but for those countries in which there is a significant negative effect on long rates (all of them have investment grade), this effect is larger than that on the short-term rate. We also find that the drop in long-term rates is generally larger for central economies than for the emerging ones. In reference to global assets, we also find that a higher trade tension increases the VIX, reduces the WTI price, and increases the price of gold, albeit not significantly in the latter case.

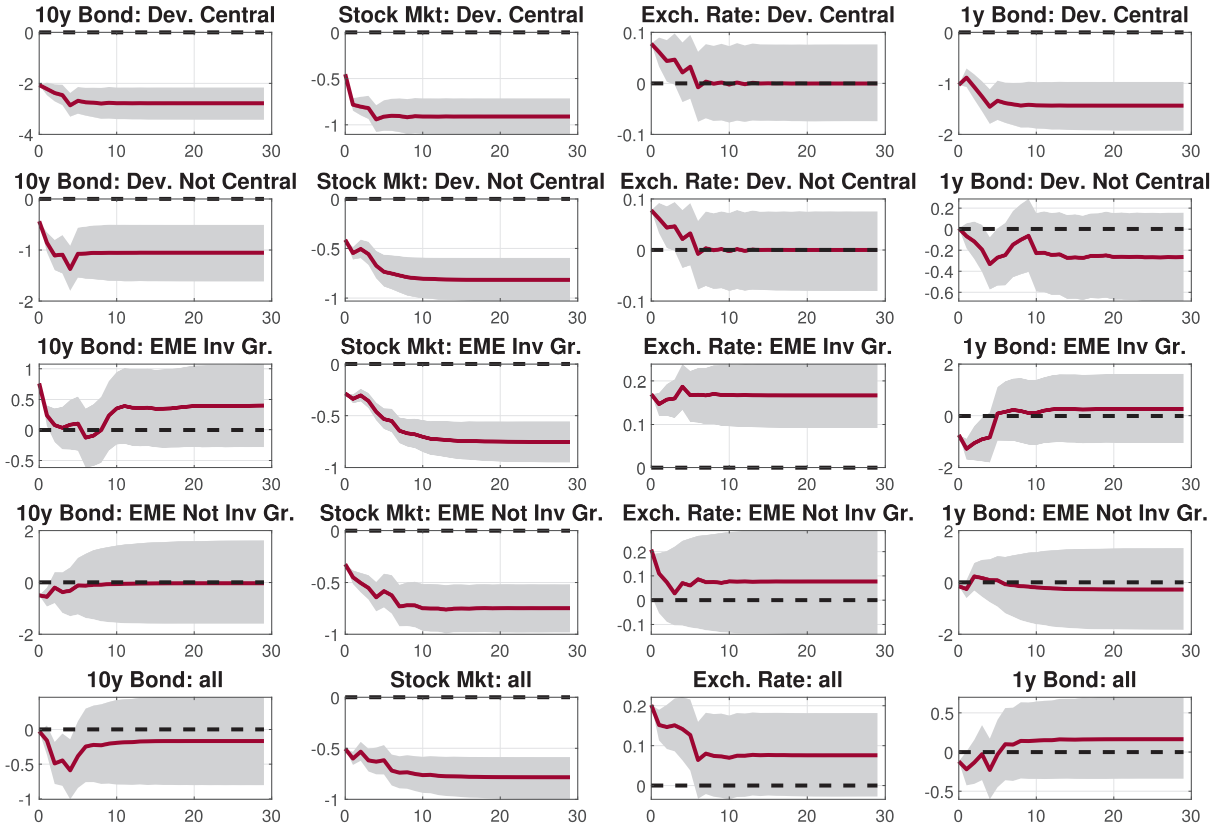


**Fig. 3.** Cumulative structural impulse-responses to trade-war shocks identified via heteroskedasticity in bi-variate VARs (median country, normalized to cause a drop of 3.79 bp of the 2-yr treasury yield). Notes: Shaded areas indicate 90% bootstrap confidence intervals.

Overall, the results point to the same direction as those reported under the event-study methodology: the trade war significantly increases global financial risk, which increases volatility, reduces long-term yields of central economies more significantly than short-term rates, reduces stock markets' returns all over the world, and depreciates the currency of emerging economies. Again, the reported effects tend to favor the risk-aversion hypothesis as the dominant force shaping markets around trade tension events.

Before concluding this section, a word about the normalization used in Table 5 is in order. As we mentioned above, figures in the table are normalized to the average drop of the 2-year U.S bond yield on days with negative news (3.8 bp). If instead of the average negative news, we consider 'big' negative news and normalize by, say, the ninth decile of the change in the 2-year U.S bond yield (12 bp), figures in Table 5 would be multiplied by a factor of 3.2.

### 4.3. Dynamic effects

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

As the event study approach is not well suited for analyzing dynamic effects (see e.g, Kothari and Warner, 2007), we focus on a slight modification of the heteroskedasticity-based approach. For every pair of returns ($r_{1t}$ and $r_{2t}$ in Eq. (1), with $r_{2t}$ being the US 2-yr treasury yield), we specify a VAR model with the number of lags selected by the BIC. Under the assumptions in Section 4.2, and defining $\Sigma_1$ and $\Sigma_0$ as the covariance matrix of the reduced form errors in trade-war and non-trade-war days, respectively, the difference $\Sigma_1 - \Sigma_0$ can be used to identify the structural shock related to the trade tension (see Wright, 2012 for the details).

We compute structural impulse-responses to a trade-war shock, and conduct inference by applying the non-parametric bootstrap procedure of Kilian (1998). For each group of countries (developed central, developed not-central, emerging with investment grade, and emerging without investment grade) Fig. 3 reports cumulative impulse-responses for the median

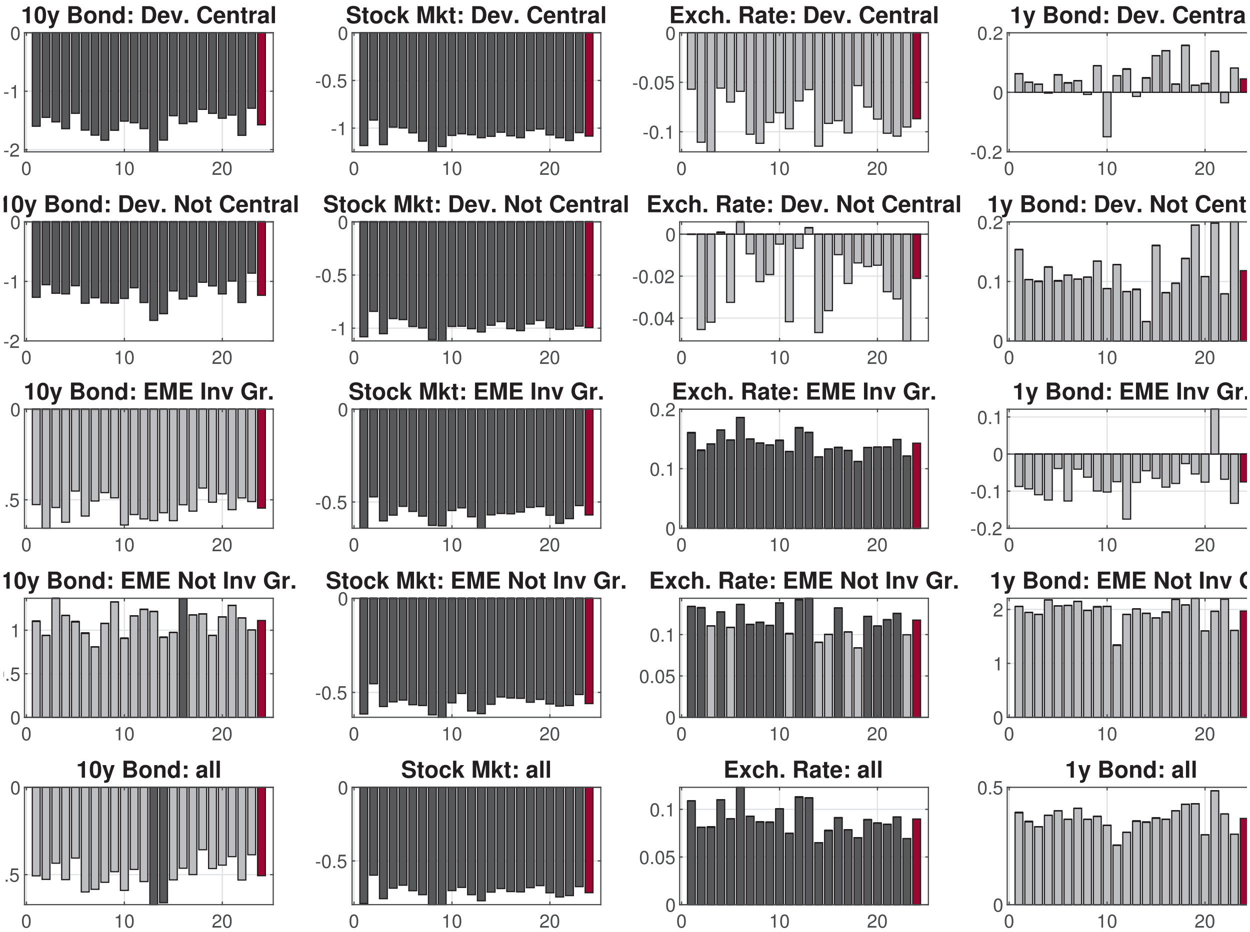


**Fig. 4.** Estimated coefficients from the panel regressions –Eq. (8)– leaving out 1 news. Notes: - Red: estimated coefficients considering all of the events (coefficients reported in Table 3). - Dark grey: significant (at the 10% of significance) estimated coefficients leaving out one news. - Light grey: not significant (at the 10% of significance) estimated coefficients leaving out one news.

country[7]. A first important result to note is that initial effects are typically close to those of Table 3, which constitutes a further signal of the robustness.

Regarding dynamic effects, for the cases in which initial effects are statistically significant, they do not tend to die out. In fact, in all of the cases in which initial effects are not zero, we find that they are still different from zero well into 30 days after the shock. As mentioned in the introduction we do not use these results to make a strong claim about long-run effects of the trade war, because our framework is not will suited for that. We just use them show that initial impacts do not die out shorty after the news.

Considering the results of Table 3 and the number of negative events recorded, we conclude that those figures imply a cumulative effect on 10-year bond yields of $-67$ bp, $-37$ bp, and $-28$ bp, for the US, developed central, and developed not-central economies, respectively. For the case of US, these magnitudes represent more than 70% of the reduction observed in the sample. Regarding stock returns, the cumulative effects are close to $-23\%$ for the US and other developed economies, and $-14\%$ for emerging markets. Finally, for the exchange rates of emerging markets, the cumulative effect is a depreciation close to 3%, both for IG and non-IG countries. The exchange rates of developed economies do not show significant effects.

## 5. Further robustness

The similarity of the results obtained from the alternative methodological approaches described above already provide important support to the robustness of the results. In this section, we present additional evidence in this regard. We first reproduce the event study leaving out some news. Then, we reproduce the heteroskedasticity-based estimations but with

[7] *Median country* is defined as the country located at the median of the cumulative IRF five days after the shock. If instead of using the same country for all horizons we consider the median country at each horizon, main conclusions do not change.

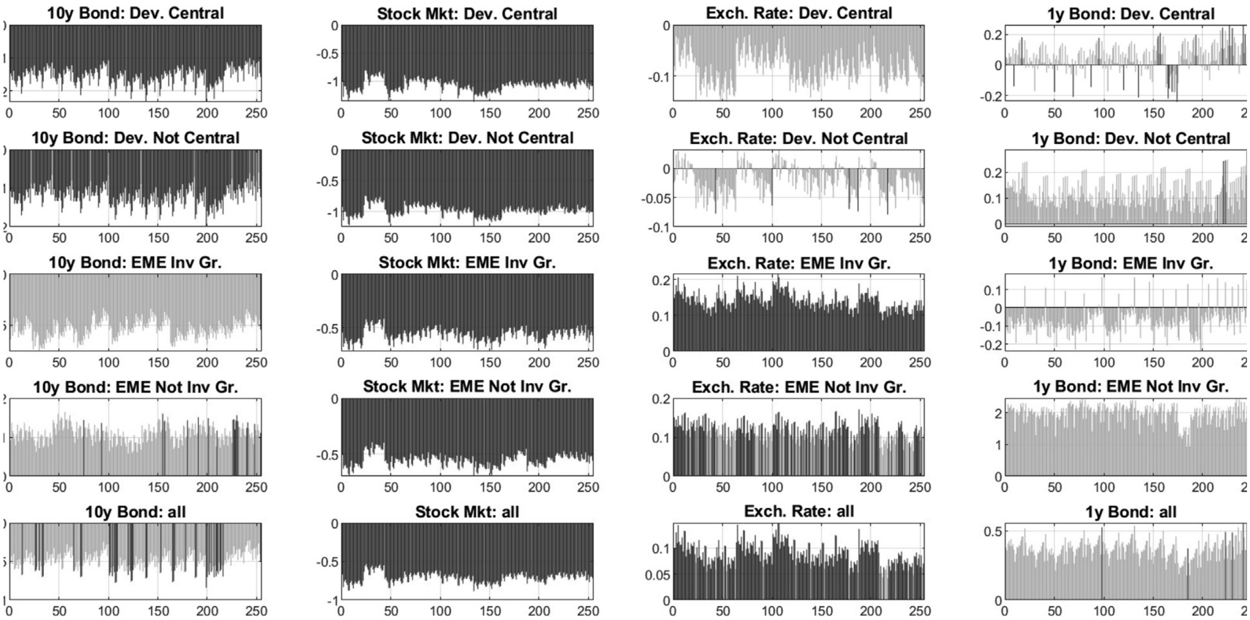


**Fig. 5.** Estimated coefficients from the panel regressions –Eq. (8)– leaving out 2 news. Notes: - Red: estimated coefficients considering all of the events (coefficients reported in Table 3). - Dark grey: significant (at the 10% of significance) estimated coefficients leaving out a pair of news. - Light grey: not significant (at the 10% of significance) estimated coefficients leaving out a pair of news.

the GMM strategy, using other reference asset as a normalization, and considering slight alterations in the definition of trade-war and non-trade war days. Finally, we provide a robustness check on the estimation of dynamic effects by applying local projection methods.

### *5.1. Further robustness on the event-study approach*

The event study includes 23 news associated to the trade-war. In order to check whether our results are driven by a single relevant event, we repeat the estimations reported in Table 3 leaving out each one of the news, one at a time. Thus, for each asset of each group of countries, we end up with 24 estimations: one that includes all of the news, and 23 other estimations, each one excluding a specific news. Results are reported in Fig. 4.

Each bar of the figure represents the coefficient estimated leaving out one news, except for the red ones, which correspond to the coefficients estimated with all of the news (those reported in Table 3). Dark (light) grey colors indicate statistical (not) significance at the 10% of significance. As it is clear from the figure, coefficients that were reported as statistically significant in Table 3, remain rather stable and statistically significant. This implies that our results are not driven by a single 'big' news.

In Fig. 5 we repeat the exercise but leaving out pairs of news. We consider every possible pair and leave them out from the estimations, one pair at a time. Thus, for each asset of each group of countries, we end up with 254 estimations: one that includes all of the news, and 253 combinations that iteratively exclude a specific pair. Inasmuch as coefficients that were reported as statistically significant in Table 3 remain rather stable and statistically significant, we conclude that the results are not driven by a pair of 'big' news.

### *5.2. Further robustness on heteroskedasticity-based strategy*

One potential shortcoming of this approach is the assumption of a two, and only two, state structure. This is an arbitrary assumption, as it could be fairly argued that different types of news may cause different types of market reactions. If this is the case, our approach would be capturing only average effects of the trade tension, leaving out potentially relevant heterogeneity. Nonetheless, the empirical validity of our identification assumptions can be formally tested because our model is overidentified (it has three moment conditions but only two unknown parameters). Indeed, it can be formally shown that if there were more than two relevant states, a standard test of overidentifying restrictions should reject the null that the restrictions are valid (see Rigobon and Sack, 2005).

Table 6 includes the results of the GMM estimation, and the p-values of the Sargan's tests for overidentifying restrictions, which are never rejected, except for Peru's stock market. This implies that our identification assumptions cannot be rejected by the data, which can be intuitively seen in the fact that point estimates in Table 6 are typically close to those of Table 5.

**Table 6**
Estimates based on the GMM implementation of heteroskedasticity based estimator (normalized to cause a drop of 3.79 bp of the 2-yr treasury yield). Average effects of increased trade tensions - in percentage points for exchanges and exchange rates and in bps for bonds.

| | 10y Bond | | Stock Mkt | | Exch. Rate | | 1y Bond | |
|---|---|---|---|---|---|---|---|---|
| | | | | *Developed central* | | | | |
| | $\beta_2$ | Sg. Pv | $\beta_2$ | Sg. Pv | $\beta_2$ | Sg. Pv | $\beta_2$ | Sg. Pv |
| Japan | −0.02 | 0.45 | −0.38** | 0.27 | −0.32*** | 0.73 | 0.16 | 0.87 |
| UK | −1.47** | 0.61 | −0.46** | 0.63 | 0.13 | 0.27 | −0.77* | 0.83 |
| Germany | −0.94** | 0.61 | −0.36* | 0.33 | 0.04 | 0.44 | −0.44** | 0.73 |
| US | −3.78*** | 0.63 | −1.05*** | 0.59 | | | −2.23*** | 0.63 |
| | | | | *Developed not central* | | | | |
| | $\beta_2$ | Sg. Pv | $\beta_2$ | Sg. Pv | $\beta_2$ | Sg. Pv | $\beta_2$ | Sg. Pv |
| Canada | −1.77*** | 0.20 | −0.29*** | 0.66 | 0.12 | 0.53 | −0.43 | 0.29 |
| Italy | 4.41*** | 0.68 | −0.43** | 0.39 | 0.04 | 0.44 | 5.47*** | 0.38 |
| France | −0.33 | 0.67 | −0.34** | 0.87 | 0.04 | 0.44 | 0.17 | 0.44 |
| Norway | −1.35*** | 0.51 | −0.11 | 0.08 | 0.24*** | 0.49 | 0.06 | 0.17 |
| Sweden | −0.61 | 0.45 | −0.36** | 0.57 | 0.15 | 0.71 | −0.11 | 0.59 |
| Switzerland | −0.67 | 0.47 | −0.55*** | 0.23 | −0.14** | 0.40 | −0.4 | 0.23 |
| Hong Kong | −0.54 | 0.54 | −0.39* | 0.56 | 0.02*** | 0.10 | – | – |
| Spain | 0.41 | 0.35 | −0.25* | 0.13 | 0.04 | 0.44 | 2*** | 0.09 |
| | | | | *Emerging with investment grade* | | | | |
| | $\beta_2$ | Sg. Pv | $\beta_2$ | Sg. Pv | $\beta_2$ | Sg. Pv | $\beta_2$ | Sg. Pv |
| China | −0.5 | 0.75 | −0.47* | 0.3 | 0.17*** | 0.26 | −0.48 | 0.92 |
| South Korea | 0.58 | 0.92 | −0.26 | 0.5 | 0.17** | 0.75 | −0.17 | 0.21 |
| Czech Rep. | −0.37 | 0.58 | −0.44*** | 0.4 | 0.16** | 0.74 | 0.12 | 0.09 |
| Chile | −1.24 | 0.46 | −0.51*** | 0.5 | 0.28*** | 0.26 | −1.69** | 0.90 |
| Thailand | −0.34 | 0.54 | −0.37** | 0.2 | 0.07 | 0.73 | −0.17 | 1.00 |
| Poland | 0.22 | 0.42 | −0.56*** | 0.8 | 0.15* | 0.77 | 0.29 | 0.24 |
| Mexico | 0.6 | 0.77 | −0.25* | 0.3 | 0.31*** | 0.17 | 0.31 | 0.24 |
| Israel | −0.94 | 0.65 | 0.09 | 0.4 | 0.06 | 0.08 | 0.13 | 0.05 |
| India | 1.76** | 0.82 | −0.25 | 0.3 | 0.22*** | 0.39 | – | – |
| Indonesia | 2.1 | 0.52 | −0.38** | 0.9 | 0.11* | 0.67 | 2 | 0.36 |
| Malaysia | −0.3 | 0.06 | −0.03 | 0.7 | 0.09* | 0.80 | 0.02 | 0.47 |
| Colombia | 3.14*** | 0.35 | −0.68*** | 1.0 | 0.42*** | 0.38 | – | – |
| Romania | −0.24 | 0.06 | −0.25 | 0.4 | 0.07 | 0.50 | 1.09 | 0.87 |
| Peru | −0.17 | 0.25 | −0.32*** | 0.0 | 0.14*** | 0.57 | −0.05 | 0.94 |
| Russia | 0.09 | 0.63 | −0.15 | 0.1 | 0.26** | 0.06 | −0.34 | 0.43 |
| Croasia | −0.46** | 0.89 | −0.01 | 0.6 | 0.07 | 0.69 | 0.45 | 0.64 |
| Bulgaria | −0.15 | 0.45 | −0.15* | 0.7 | 0.03 | 0.42 | −0.15 | 0.27 |
| | | | | *Emerging without investment grade* | | | | |
| | $\beta_2$ | Sg. Pv | $\beta_2$ | Sg. Pv | $\beta_2$ | Sg. Pv | $\beta_2$ | Sg. Pv |
| Hungary | −0.26 | 0.26 | −0.56*** | 0.14 | 0.13 | 0.51 | −0.15 | 0.08 |
| South Africa | 2.4** | 0.52 | −0.33* | 0.93 | 0.48*** | 0.56 | – | – |
| Brazil | 3.13** | 0.60 | −0.4* | 0.19 | 0.27** | 0.19 | −0.38 | 0.41 |
| Vietnam | −0.68 | 0.55 | −0.31 | 0.10 | 0.01 | 0.88 | −1.04 | 0.52 |
| Turkey | 1.51 | 0.95 | −0.27 | 0.11 | 0.15 | 0.82 | 2.93 | 0.87 |
| Nigeria | – | – | −0.06 | 0.29 | −0.03 | 0.42 | – | 0.56 |
| Sri Lanka | −1.49 | 0.17 | −0.14 | 0.51 | 0.03 | 0.18 | −2.04 | 0.16 |
| Ukraine | 2.75 | 0.74 | 0.04 | 0.29 | 0.08 | 0.46 | 8.46* | 0.73 |
| | | | | *Global Assets* | | | | |
| | $\beta_2$ | Sg. Pv | | | | | | |
| VIX | 21.16*** | 0.60 | | | | | | |
| Oil | −1.71*** | 0.76 | | | | | | |
| Gold | −0.08 | 0.51 | | | | | | |

Notes:
- Asterisks (***, **, *) indicate statistical significance at the 1%, 5% and 10%, respectively.

In Table 7, instead of following Rigobon and Sack (2005) in using the 2-year treasury yield as the reference asset, we use the 5-year yield. Signs are basically unchanged with respect to Table 5, magnitudes remain close, and values that appear as statistically significant/not-significant are essentially the same in both tables.

Finally, in Table 8 we consider small alterations in the definition of trade-war and non-trade-war days. Columns ‘orig’ contain original results (those of Table 5). Columns ‘+W’ duplicate trade-war days by also considering as trade-war days the days following the original ones. Columns ‘+NW’ duplicates non-trade-war days by also considering as non-trade-war days the days previous to the original ones. Point estimates are generally close to each other, and main conclusions do not change. Hence, as discussed in Section 2.2, we conclude that results are robust to slight alterations in the definition of war and non-war days.

**Table 7**
Estimates based on the IV implementation of heteroskedasticity based estimator (normalized to cause a drop of 3.78 bp of the 5-yr treasury yield). Average effects of increased trade tensions - in percentage points for exchanges and exchange rates and in bps for bonds.

| | 10y Bond | Stock Mkt | Exch. Rate | 1y Bond |
|---|---|---|---|---|
| | | *Developed dentral* | | |
| Japan | −0.11 | −0.27* | −0.27*** | 0.13 |
| UK | −1.74*** | −0.32** | 0.14* | −0.91** |
| Germany | −0.88** | −0.4** | 0.02 | −0.39** |
| US | −3.26*** | −0.98*** | | −2.03*** |
| | | *Developed dentral* | | |
| Canada | −2.28*** | −0.25** | 0.13* | −0.69*** |
| Italy | 3.67*** | −0.41** | 0.02 | 4.17*** |
| France | −0.36 | −0.33** | 0.02 | 0.22* |
| Norway | −1.24*** | −0.18 | 0.19** | −0.11 |
| Sweden | −0.55* | −0.34** | 0.11 | −0.08 |
| Switzerland | −0.71* | −0.53*** | −0.12** | −0.69** |
| Hong Kong | −0.52 | −0.29 | 0.01* | – |
| Spain | 0.46 | −0.33*** | 0.02 | 1.4*** |
| | | *Emerging with investment grade* | | |
| China | −0.52 | −0.28 | 0.11** | −0.34 |
| South Korea | 0.15 | −0.23 | 0.13* | 0.92 |
| Czech Rep. | −0.29 | −0.35*** | 0.13** | −0.12 |
| Chile | −1.17 | −0.48*** | 0.27*** | −1.6* |
| Thailand | −0.41 | −0.28* | 0.05 | −0.22 |
| Poland | −0.17 | −0.52*** | 0.14* | 0.25 |
| Mexico | 0.24 | −0.19* | 0.3*** | −0.07 |
| Israel | −1.03* | −0.05 | 0.11** | 0.3 |
| India | 1.22* | −0.22 | 0.16*** | – |
| Indonesia | 1.33 | −0.28* | 0.06 | 1.46 |
| Malaysia | 0.55 | 0.02 | 0.06 | 0.05 |
| Colombia | 2.07** | −0.49** | 0.4*** | – |
| Romania | 0.42 | −0.32** | 0.05 | 1.58 |
| Peru | 0.04 | −0.36*** | 0.11*** | −0.2 |
| Russia | 0.1 | −0.1 | 0.21** | −0.61 |
| Croasia | −0.28* | −0.01 | 0.04 | 0.32 |
| Bulgaria | −0.18 | −0.07 | 0.01 | 0.07 |
| | | *Emerging without investment grade* | | |
| Hungary | 0.42 | −0.45*** | 0.11 | −0.07 |
| South Africa | 1.97** | −0.26 | 0.46*** | – |
| Brazil | 3.16** | −0.39** | 0.26** | 0.22 |
| Vietnam | −0.46 | −0.05 | 0.01 | −1.78 |
| Turkey | 0.83 | −0.33** | 0.06 | 1.8 |
| Nigeria | – | −0.04 | −0.08 | – |
| Sri Lanka | −1.67* | −0.15** | 0.01 | −1.53 |
| Ukraine | 1.61 | 0.09 | 0.04 | 5.74 |
| | | *Global Assets* | | |
| VIX | 18.63*** | | | |
| Oil | −1.61*** | | | |
| Gold | −0.09 | | | |

Notes:
- Asterisks (***, **, *) indicate statistical significance at the 1%, 5% and 10%, respectively.

## *5.3. Further robustness on dynamic effects*

As mentioned in Section 4.3, the event study approach is well suited for analyzing the immediate impact, but not for addressing persistence. We now provide additional evidence about dynamic effects computed as the cumulative impulse-responses to negative news from Eq. (8) using the local projection method of Jordà (2005) (see also Plagborg-Møller and Wolf, 2021 for an analysis of efficiency and consistency of impulse-response functions constructed from local projections). These results are reported in Fig. 6. Consistent with Fig. 3, effects that appear to be statistically significant on impact, do not fade out as the days go by. The only exception are stock returns, which do revert towards zero after about a week after the event. This results reinforces the need of further exploring the dynamic effects of trade tensions using robust methodologies especially designed for the said purpose.

**Table 8**
Estimates based on the IV implementation of heteroskedasticity based estimator (normalized to cause a drop of 3.79 bp of the 2-yr treasury yield). Average effects of increased trade tensions - in percentage points for exchanges and exchange rates and in bps for bonds.

| | **10 year Bond** | | | **Stock Mkt** | | | **Exch. Rate** | | | **1 year Bond** | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| *Developed central* | | | | | | | | | | | | |
| | Orig | +W | +NW | Orig | +W | +NW | Orig | +W | +NW | Orig | +W | +NW |
| Japan | −0.03 | −0.09 | −0.02 | −0.35* | −0.05 | −0.35** | −0.33*** | −0.31*** | −0.33*** | 0.17 | 0.11 | 0.17 |
| UK | −1.5** | −1.56*** | −1.51** | −0.44** | −0.32** | −0.42** | 0.13 | 0.09 | 0.13 | −0.79* | −0.87*** | −0.8* |
| Germany | −0.89** | −1.26*** | −0.88* | −0.38** | −0.37** | −0.38** | 0.03 | 0.06 | 0.03 | −0.42** | −0.6*** | −0.43** |
| US | −3.82*** | −4.08*** | −3.83*** | −1.16*** | −1.11*** | −1.17*** | | | | −2.26*** | −2.09*** | −2.27*** |
| *Developed not central* | | | | | | | | | | | | |
| | Orig | +W | +NW | Orig | +W | +NW | Orig | +W | +NW | Orig | +W | +NW |
| Canada | −2.06*** | −2.52*** | −2.06*** | −0.29** | −0.34*** | −0.29*** | 0.14* | 0.17*** | 0.13* | −0.56* | −0.8*** | −0.53* |
| Italy | 4.14*** | 3.07*** | 4.27*** | −0.45** | −0.45*** | −0.44** | 0.03 | 0.06 | 0.03 | 4.74*** | 3.23*** | 4.93*** |
| France | −0.26 | −0.6 | −0.24 | −0.35** | −0.29** | −0.34** | 0.03 | 0.06 | 0.03 | 0.25* | 0.07 | 0.26* |
| Norway | −1.39*** | −1.2*** | −1.39*** | −0.19 | −0.19 | −0.19 | 0.26*** | 0.29*** | 0.26*** | −0.2 | −0.16 | −0.2 |
| Sweden | −0.5 | −0.71* | −0.47 | −0.36** | −0.31** | −0.36** | 0.16 | 0.21*** | 0.15 | −0.03 | −0.16 | −0.02 |
| Switzerland | −0.56 | −0.95** | −0.54 | −0.52*** | −0.3** | −0.52*** | −0.14** | −0.12*** | −0.14** | −0.78* | −0.71*** | −0.78* |
| Hong Kong | −0.5 | 0.78 | −0.35 | −0.38* | −0.06 | −0.38* | 0.02*** | 0.01** | 0.02*** | – | – | – |
| Spain | 0.49 | 0.18 | 0.54 | −0.32** | −0.28** | −0.31** | 0.03 | 0.06 | 0.03 | 1.44*** | 1.1*** | 1.46*** |
| *Emerging with investment grade* | | | | | | | | | | | | |
| | Orig | +W | +NW | Orig | +W | +NW | Orig | +W | +NW | Orig | +W | +NW |
| China | −0.49 | −0.34 | −0.48 | −0.38 | −0.13 | −0.38 | 0.16*** | 0.11** | 0.16*** | −0.49 | −0.2 | −0.55 |
| South Korea | 0.62 | 0.57 | 0.62 | −0.27* | −0.08 | −0.27* | 0.17** | 0.1 | 0.16* | 0.95 | 0.55 | 0.95 |
| Czech Rep. | −0.32 | −0.23 | −0.32 | −0.38*** | −0.26*** | −0.39*** | 0.15* | 0.18*** | 0.15* | −0.66 | −0.25 | −0.66 |
| Chile | −1.49 | −1.26* | −1.49 | −0.53*** | −0.43*** | −0.52*** | 0.31*** | 0.29*** | 0.31*** | −1.81* | −1.53** | −1.81* |
| Thailand | −0.4 | −0.44 | −0.42 | −0.3 | −0.27* | −0.31* | 0.06 | 0.08** | 0.05 | −0.17 | −0.29 | −0.22 |
| Poland | −0.02 | −0.4 | −0.04 | −0.54*** | −0.37** | −0.53*** | 0.14 | 0.2*** | 0.14* | 0.17 | 0.04 | 0.18 |
| Mexico | 0.52 | 0.13 | 0.49 | −0.25* | −0.32*** | −0.25* | 0.39*** | 0.41*** | 0.4*** | 0.23 | −0.23 | 0.15 |
| Israel | −1.1 | −0.93 | −1.1 | 0.01 | −0.08 | 0.01 | 0.12* | 0.13** | 0.11* | 0.34 | 0.23 | 0.34 |
| India | 1.7** | 1.42** | 1.75** | −0.19 | −0.11 | −0.19 | 0.19** | 0.15*** | 0.18** | – | – | – |
| Indonesia | 2.12 | 1.57 | 2.12 | −0.39** | −0.18 | −0.39** | 0.11* | 0.06 | 0.11* | 1.6 | 1.47 | 1.6 |
| Malaysia | 1.04 | 0.6 | 1.04 | −0.05 | 0.07 | −0.05 | 0.09* | 0.05 | 0.09* | 0.04 | 0.03 | 0 |
| Colombia | 2.35** | 1.49* | 2.32** | −0.68*** | −0.61*** | −0.67*** | 0.47*** | 0.43*** | 0.46*** | – | – | – |
| Romania | 0.28 | 0.1 | 0.26 | −0.24 | −0.04 | −0.26 | 0.06 | 0.08 | 0.06 | 1.23 | 0.48 | 1.26 |
| Peru | 0.17 | 0.49 | 0.17 | −0.37*** | −0.38*** | −0.36*** | 0.15*** | 0.14*** | 0.15*** | −0.03 | 0.08 | −0.04 |
| Russia | −0.05 | 0.06 | −0.08 | −0.12 | −0.14 | −0.11 | 0.28*** | 0.36*** | 0.27*** | −0.93 | −0.36 | −1.07 |
| Croasia | −0.44** | −0.25* | −0.44** | −0.02 | 0.04 | −0.02 | 0.07 | 0.08 | 0.07 | 0.54 | 0.47** | 0.52 |
| Bulgaria | −0.22 | 0.01 | −0.21 | −0.17* | −0.02 | −0.16 | 0.02 | 0.06 | 0.02 | 0.2 | 0.07 | 0.13 |
| *Emerging without investment grade* | | | | | | | | | | | | |
| | Orig | +W | +NW | Orig | +W | +NW | Orig | +W | +NW | Orig | +W | +NW |
| Hungary | 0.51 | −0.12 | 0.49 | −0.44** | −0.29** | −0.43** | 0.12 | 0.17** | 0.13 | −0.21 | −0.2 | −0.22 |
| South Africa | 2.84** | 2.73*** | 2.84** | −0.32* | −0.27* | −0.31 | 0.51*** | 0.56*** | 0.51*** | – | – | – |
| Brazil | 3.54** | 3.11** | 3.71** | −0.5** | −0.63*** | −0.5** | 0.35** | 0.35*** | 0.37** | 0.52 | 0.55 | 0.78 |
| Vietnam | −0.53 | −0.63 | −0.51 | −0.25 | −0.07 | −0.27 | 0.01 | −0.01 | 0.01 | −0.71 | −1.21 | −0.69 |

(*continued on next page*)

**Table 8** (*continued*)

| | 10 year Bond | | | Stock Mkt | | | Exch. Rate | | | 1 year Bond | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| Turkey | 1.54 | 1.5 | 1.49 | −0.38* | −0.25 | −0.38* | 0.13 | 0.22* | 0.13 | 3.11 | 3.13 | 3.09 |
| Nigeria | – | – | – | −0.05 | 0 | −0.05 | −0.07 | −0.07* | −0.07 | – | – | – |
| Sri Lanka | −2.01 | −2.01* | −2.01 | −0.16* | −0.21** | −0.16* | 0.03 | 0.01 | 0.03 | −1.35 | −0.87 | −1.35 |
| Ukraine | 2.71 | 2.57* | 2.72 | 0.1 | 0.14 | 0.06 | 0.07 | 0.09 | 0.07 | 9.81* | 9.98** | 9.67* |
| | | | | | | | *Global Assets* | | | | | |
| | Orig | +W | +NW | | | | | | | | | |
| VIX | 18.63*** | 15.6*** | 18.73*** | | | | | | | | | |
| Oil | −1.61*** | −1.19*** | −1.58*** | | | | | | | | | |
| Gold | −0.09 | −0.12 | −0.08 | | | | | | | | | |

Notes:
- Asterisks (***, **, *) indicate statistical significance at the 1%, 5% and 10%, respectively.

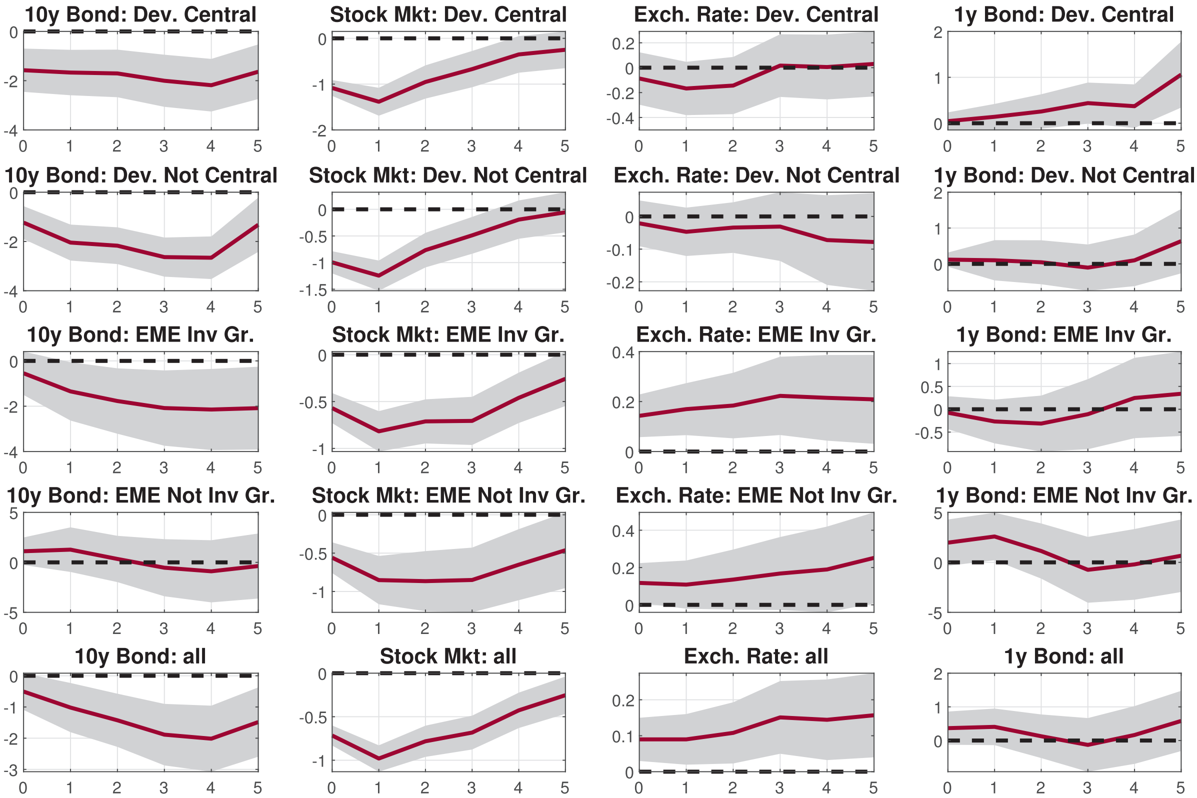


**Fig. 6.** Cumulative impulse-responses to negative news from local projections. Notes: Shaded areas indicate 90% confidence intervals.

## 6. Concluding remarks

This paper analyzed the impact of the trade-war between the United States and China on world financial markets. Our results show that trade-war related news significantly compresses bond yields of developed and some emerging economies with investment grade, lowers stock markets' returns all over the world, depreciate currencies of emerging economies and increases the VIX. Moreover, given the relatively smaller effect on short-term yields, these results favor an interpretation in which the main driving force is an increase in global risk-aversion, over the alternative mechanism working through a worsening of economic growth prospects.

Our results are robust to a number of alternative specifications, including event-studies as well as estimations based on conditional heteroskedasticity. An extension of these methodologies to dynamic effects show that most of these results persist at least throughout the next few weeks after the events. We therefore conclude that the trade-war has been a significant driver of asset market returns in 2018 and 2019. Moreover, these results may have important implications going forward, to the extent that this episode provides just the first taste of a longer-lasting paradigm change in international trade relations, away from the low-tariff, multilateral rules-based system that characterized the preceding decades. As time goes by, further research will be necessary to properly assess the long-run effects of the trade war, not only on asset prices and its effects on the real economy, but also on global financial market structures.

## Data availability

Data will be made available on request.

## Appendix A. Event Study: Methodological details

In this appendix we show that, under two basic assumptions, the difference in the expected returns in days with news *vs.* days without news represents the total (direct plus indirect) causal effect of the unobserved trade-tension on the expected returns.

The two basic assumptions are: (a) the news and the trade-tension factor are exogenous, (b) there are only two states of the nature for the level of trade-tension, one high on days with news and other low for the other days.

Assume further that the daily return of a financial asset $i$ ($r_{it}$) is generated by:

$$r_{it} = c_i + \beta_i \omega_t + \lambda_i X_{it} + e_{it}, \tag{A.1}$$

where $\omega_t$ is the unobservable level of trade-tension, $\beta_i$ is the coefficient that captures the effect of the trade-tension on the return, $X_{it}$ is an $N \times 1$ vector containing other determinants of the return (e.g., variables related to monetary policy, macroeconomic factors, etc.), $\lambda_i$ is a $1 \times N$ vector of coefficients, and $e_{it}$ is a random noise with zero mean, serially uncorrelated, but not necessarily *iid*. Using assumption b we can write:

$$\omega_t = \psi_1 \mathbb{1}_{n,t} + \psi_0 \mathbb{1}_{\bar{n}t}, \tag{A.2}$$

where $\psi_1$ is the level of trade-tension in days with news, $\mathbb{1}_{n,t}$ is the indicator variable that takes value one in days with news and zero in the rest of the days, $\psi_0$ is the level of trade-tension in days without news ($\psi_0 < \psi_1$), and $\mathbb{1}_{\bar{n}t} = 1 - \mathbb{1}_{n,t}$.

As it is clear from Eq. (A.2), this approach requires all of the news to have the same "sign" (i.e., all of the news must increase the trade-tension).

Plugging Eq. (A.2) into Eq. (A.1) we get:

$$r_{it} = c_i + \beta_i \psi_1 \mathbb{1}_{n,t} + \beta_i \psi_0 \mathbb{1}_{\bar{n}t} + \lambda_i X_{it} + e_{it}. \tag{A.3}$$

Defining $c_i^* \equiv c_i + \beta_i \psi_0$, and using $\mathbb{1}_{\bar{n}t} = 1 - \mathbb{1}_{n,t}$, we get a statistically equivalent, but estimable, version of Eq. (A.3):

$$r_{it} = c_i^* + \beta_i(\psi_1 - \psi_0)\mathbb{1}_{n,t} + \lambda_i X_{it} + e_{it}. \tag{A.4}$$

From Eq. (A.4), the expected return in days without news is:

$$E[r_{it} \mid \mathbb{1}_{n,t} = 0] = c_i^* + \lambda_i E[X_{it} \mid \mathbb{1}_{n,t} = 0]. \tag{A.5}$$

For days with news we get:

$$E[r_{it} \mid \mathbb{1}_{n,t} = 1] = c_i^* + \lambda_i E[X_{it} \mid \mathbb{1}_{n,t} = 1] + \beta_i(\psi_1 - \psi_0). \tag{A.6}$$

The exogeneity assumption implies that the order of causality between $\mathbb{1}_{n,t}$ and $X_{it}$, if it exists, goes from $\mathbb{1}_{n,t}$ to $X_{it}$ and not the other way around[8]. Thus, we may write:

$$X_{it} = \alpha \mathbb{1}_{n,t} + X_{it}^*, \tag{A.7}$$

where $X_{it}^*$ is the value of $X_{it}$ in days without news, and $E[\mathbb{1}_{n,t} \times X_{it}^*] = 0\ \forall t$. Hence, using expressions A.5, A.6 and A.7, we get:

$$E[r_{it} \mid \mathbb{1}_{n,t} = 1] - E[r_{it} \mid \mathbb{1}_{n,t} = 0] = \lambda_i \alpha + \beta_i(\psi_1 - \psi_0). \tag{A.8}$$

Eq. (A.8) implies that total impact of the increase in the trade-tension on the expected return is $\lambda_i \alpha + \beta_i(\psi_1 - \psi_0)$, where $\beta_i(\psi_1 - \psi_0)$ represents the direct impact, and $\lambda_i \alpha$ is the indirect impact that may take place through a change in the value of $X_{it}$ caused by the higher trade-tension.

## Appendix B. More general DGP for the returns

In this appendix we analyze how different estimation strategies would perform when the true DGP departs from Eqs. A.1 and A.2 (those that give rise to Eq. (7)). We are interested in two particular situations. First, news may have lagged impacts on markets, what is not considered in the DGP described in Section 2.1. Second, as noted in Section 4.1, using daily data could be problematic – especially for western economies – because news may arrive after markets closures.

To tackle these two issues we generalize Eq. (A.2) as follows:

$$\omega_t = \psi_1 \mathbb{1}_{n,t} + \psi_2 \mathbb{1}_{n,t-1} + \rho_1 \mathbb{1}_{n^*,t} + \rho_2 \mathbb{1}_{n^*,t-1} + \psi_0 \mathbb{1}_{\bar{n}t}, \tag{B.1}$$

where, now, $\mathbb{1}_{n,t}$ is the indicator function that takes value 1 the days with news that take palace before markets' closure – type-1 news. The indicator function $\mathbb{1}_{n^*,t}$ takes value one on the days following those with news that arrived after markets' closures – type-2 news –, and $\mathbb{1}_{\bar{n}t} = 1 - \mathbb{1}_{n,t} - \mathbb{1}_{n^*,t}$. We define the number of type-1 and type-2 news as $q$ and $q^*$, respectively, and assume that the trade-tension generated by both types of news is the same, what implies $\psi_1 = \rho_1$, and $\psi_2 = \rho_2$.

Using similar arguments as those in Section 2.1, it can be shown that the contemporaneous and lagged impacts of each kind of news, can be estimated with the following regression:

$$r_t = c + \gamma_0(\mathbb{1}_{n,t} + \mathbb{1}_{n,t}^*) + \gamma_1(\mathbb{1}_{n,t-1} + \mathbb{1}_{n^*,t-1}) + r_t^*, \tag{B.2}$$

where $\gamma_0$ represents the difference in the expected return in days with news (if the news is of type 1, or in the day after if it is of type 2) with respect days without news, and $\gamma_1$ represents the difference in the expected returns the day after the news (if the news is of type 1, or two days after if it is of type 2).

Note that, since we just observe the day of the news, we cannot separate the the terms of the indicator function $\mathbb{1}_{n,t} + \mathbb{1}_{n,t+1}^*$.

We analyze four alternative DGPs that arise from imposing different restrictions on Eq. (B.2):

[8] We are referring to the concept of *strong exogeneity* defined by Engle et al. (1983).

i) There are some news coming after the market closes, and there are also lagged effects, so that: $\mathbb{1}_{n^*,t} \neq 0$, for some $t$, and $\gamma_1 \neq 0$.
ii) All news take place while markets are open, but there are lagged effects, so that: $\mathbb{1}_{n^*,t} = 0$, for all $t$, and $\gamma_1 \neq 0$,
iii) There are some news coming after the market closes, but no lagged effects, so that: $\mathbb{1}_{n^*,t} \neq 0$, for some $t$, and $\gamma_1 = 0$,
iv) All news take place while markets are open, and there are no lagged effects, so that: $\mathbb{1}_{n^*,t} = 0$, for all $t$, and $\gamma_1 = 0$,

For each DGP we consider three possible estimation strategies:

a) $r_t$ with lags: $r_t = c + \lambda_{r,0}(\mathbb{1}_{n,t} + \mathbb{1}_{n^*,t+1}) + \lambda_{r,1}(\mathbb{1}_{n,t-1} + \mathbb{1}_{n^*,t}) + \epsilon_t$
b) $R_t$ with no lags: $R_t = c_* + \lambda_R(\mathbb{1}_{n,t-1} + \mathbb{1}_{n^*,t}) + \epsilon_t$
c) $r_t$ with no lags: $r_t = c_* + \lambda_r(\mathbb{1}_{n,t} + \mathbb{1}_{n^*,t+1}) + \epsilon_t$

Table B.1 includes the bias of the estimators for $\gamma_0, \gamma_1$, and $\gamma_0 + \gamma_1$ for each DGP and estimation equation.

Focusing first on cases with no lagged effects (DGPs iii and iv), strategy c is the most efficient for DGP iv, but delivers a downwards estimation bias for DGP iii. On the other hand, strategy (b) delivers un-biased estimators for both DGPs. Thus, if lagged effects could be ruled out before hand, a valid strategy would be obtaining estimators from equations b and c and keeping the maximum.

In real applications, lagged effects cannot be disregarded without testing. For DGP ii, strategy (a) is the most efficient, but both strategies (a) and (b) deliver un-biased estimators of $\gamma_0 + \gamma_1$. For DGP i biases are un-avoidable: while in strategy (b) the bias is proportional to the relative amount of type-1 news ($\frac{q}{q+q^*}$), in strategy (a) the bias is proportional to the relative number of type-2 news ($\frac{q^*}{q+q^*}$).

For Asian economies, whose time difference with Washington DC is around 12 hours, $\frac{q}{q+q^*}$ would be close to zero and strategy (b) would be the most appropriate. For economies located in the Americas, the opposite is true, so that strategy (a) would be preferred. For European economies, for which the time difference with DC is around 6 hours, we can expect $q \simeq q^*$. In the extreme case that $q = q^*$ the difference between the estimator (b) and (a) would be un-biased, though un-efficient, estimator of $\gamma_0 + \gamma_1$.

**Table B.1**
Estimators' bias for different DGPs and estimation strategies.

| | Parameter | Estimated equation<br>(a) $r_t$ with lags | (b) $R_t$ with no lags | (c) $r_t$ with no lags |
|---|---|---|---|---|
| DGP i | $\gamma_0$ | $-\frac{q^*\gamma_0}{q+q^*}$ | – | $-\frac{q^*\gamma_0}{q+q^*}$ |
| | $\gamma_1$ | $\frac{q^*(\gamma_0-\gamma_1)}{q+q^*}$ | – | – |
| | $\gamma_0+\gamma_1$ | $-\frac{q^*\gamma_1}{q+q^*}$ | $\frac{q\gamma_1}{q+q^*}$ | $-(\gamma_1 + \frac{q^*\gamma_0}{q+q^*})$ |
| DGP ii | $\gamma_0$ | 0 | – | 0 |
| | $\gamma_1$ | 0 | – | $-\gamma_1$ |
| | $\gamma_0+\gamma_1$ | 0 | 0 | $-\gamma_1$ |
| DGP iii | $\gamma_0$ | $-\frac{q^*\gamma_0}{q+q^*}$ | 0 | $-\frac{q^*\gamma_0}{q+q^*}$ |
| DGP iv | $\gamma_0$ | 0 | 0 | 0 |

Notes:
Calculations assume that there are no consecutive news.

## Appendix C. Event study with Eq. (7)

See Table C.1.

**Table C.1**
Mean effects of a higher trade-tension from panel regressions (in bps for sovereign bonds and percentage points for the rest of the assets).

| | (−) | (+) | all |
|---|---|---|---|
| | *10y Bond* | | |
| Dev. Centr | −1.46** | 1.46** | −1.02** |
| Dev. Not Centr | −0.5 | −0.04 | −0.43* |
| EME Inv Gr. | 0.08 | 0.49 | 0.14 |
| EME Not Inv Gr. | 0.36 | −2.35* | −0.09 |
| All | −0.21 | −0.15 | −0.2 |

(*continued on next page*)

**Table C.1** (*continued*)

| | (−) | (+) | all |
|---|---|---|---|
| | | *1y Bond* | |
| Dev. Centr | −1.46** | 1.46** | −1.02** |
| Dev. Not Centr | −0.5 | −0.04 | −0.43* |
| EME Inv Gr. | 0.08 | 0.49 | 0.14 |
| EME Not Inv Gr. | 0.36 | −2.35* | −0.09 |
| All | −0.21 | −0.15 | −0.2 |
| | | *Stock Mkt* | |
| Dev. Centr | −1.46** | 1.46** | −1.02** |
| Dev. Not Centr | −0.5 | −0.04 | −0.43* |
| EME Inv Gr. | 0.08 | 0.49 | 0.14 |
| EME Not Inv Gr. | 0.36 | −2.35* | −0.09 |
| All | −0.21 | −0.15 | −0.2 |
| | | *Exch. Rate* | |
| Dev. Centr | −1.46** | 1.46** | −1.02** |
| Dev. Not Centr | −0.5 | −0.04 | −0.43* |
| EME Inv Gr. | 0.08 | 0.49 | 0.14 |
| EME Not Inv Gr. | 0.36 | −2.35* | −0.09 |
| All | −0.21 | −0.15 | −0.2 |

Notes:
Columns (−), (+), and *all* include results for negative, positive, and all news, respectively.
Asterisks (***, **, *) indicate statistical significance at the 1%, 5% and 10%, respectively.
See Table 1 for the definition of the country groups.

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