Board of Governors of the Federal Reserve System

International Finance Discussion Papers

ISSN 1073-2500 (Print)
ISSN 2767-4509 (Online)

Number 1371

February 2023

The US, Economic News, and the Global Financial Cycle

Christoph E. Boehm and T. Niklas Kroner

Please cite this paper as:
Boehm, Christoph E., and T. Niklas Kroner (2023). "The US, Economic News, and the Global Financial Cycle," International Finance Discussion Papers 1371. Washington: Board of Governors of the Federal Reserve System, https://doi.org/10.17016/IFDP.2023.1371.

The US, Economic News, and the Global Financial Cycle*

Christoph E. Boehm
UT Austin and NBER

T. Niklas Kroner
Federal Reserve Board

This draft: February 15, 2023
First draft: January 10, 2020

Abstract

We provide evidence for a causal link between the US economy and the global financial cycle. Using intraday data, we show that US macroeconomic news releases have large and significant effects on global risky asset prices. Stock price indexes of 27 countries, the VIX, and commodity prices all jump instantaneously upon news releases. The responses of stock indexes co-move across countries and are large—often comparable in size to the response of the S&P 500. Further, US macroeconomic news explains on average 23 percent of the quarterly variation in foreign stock markets. The joint behavior of stock prices, bond yields, and risk premia suggests that systematic US monetary policy reactions to news do not drive the estimated effects. Instead, the evidence points to a direct effect on investors' risk-taking capacity. Our findings show that a byproduct of the United States' central position in the global financial system is that news about its business cycle has large effects on global financial conditions.

JEL Codes: E44, E52, F40, G12, G14, G15,

Keywords: Global Financial Cycle; Macroeconomic announcements; International spillovers; Stock returns; VIX; Monetary Policy; High-frequency event study

*We thank the editor (Kurt Mitman), three anonymous referees, and Ambrogio Cesa-Bianchi, Olivier Coibion, Charles Engel, Benjamin Hebert, Zhengyang Jiang, Luciana Juvenal, Şebnem Kalemli-Özcan, Benjamin Knox, Andrei Levchenko, Guido Lorenzoni, Matteo Maggiori, Silvia Miranda-Agrippino, Peter Morrow, Nitya Pandalai-Nayar, Marco Pinchetti, Alessandro Rebucci, Helene Rey, Jesse Schreger, Eric van Wincoop, Francesco Zanetti, Tony Zhang, as well as seminar and conference participants at UT Austin, Bocconi, Maryland, Fed Board, Carleton, KU Leuven, IWH Halle, Stanford GSB, Notre Dame, ASSA 2020, CEA 2021, EEA-ESEM 2021, EWMES 2020, NASMES 2021, RES 2021, SMYE 2021, SED 2021, GEA 2022, CFM International Macro Conference 2022, and NBER SI 2022 for helpful comments. We thank Olivier Coibion, Stefano Eusepi, Nitya Pandalai-Nayar, Aysegul Sahin, and the UT Austin Department of Economics for financial support to purchase the proprietary data used in this paper. We thank Domenico Giannone, Refet Gürkaynak, Burçin Kısacıkoğlu, Chiara Scotti, Clara Vega, and Jonathan Wright for generously sharing data and programs with us. We also thank Gregory Weitzner for helping us with accessing parts of the data. A previous version of this paper was circulated under the title "What does high frequency identification tell us about the transmission and synchronization of business cycles?" The views expressed are those of the authors and do not necessarily reflect those of the Federal Reserve Board or the Federal Reserve System.

Email: `chris.e.boehm@gmail.com` and `t.niklas.kroner@gmail.com`.

1 Introduction

> *The global financial cycle appears in co-movements of gross flows, asset prices, leverage, and credit creation, which are all closely linked to fluctuations in the VIX. But what are its drivers?*
>
> — Rey (2013)

In an influential speech at the Jackson Hole Symposium in 2013, Rey (2013) provides evidence for the global co-movement of capital flows, risky asset prices, credit growth, and leverage. According to Rey, this co-movement—which she calls the *global financial cycle*—constitutes an external source of financial and macroeconomic volatility for countries with open capital accounts. In episodes of favorable international financial conditions, these countries experience capital inflows, buildups of credit and leverage, and appreciations in risky asset prices, ultimately resulting in macroeconomic expansion. In episodes of retrenchment, however, capital flows reverse, credit and leverage contract, and risky asset prices plummet. Historically, these episodes of retrenchment are often associated with economic crises.

Some observers, however, have challenged Rey's interpretation of the global financial cycle. Since the observed co-movements of capital flows, risky asset prices, credit growth, and leverage across countries are ultimately correlations, alternative interpretations are also possible. Bernanke (2017) discusses several of these alternatives and notes, among other things, that the global financial cycle could be driven by common shocks—shocks that directly affect multiple countries simultaneously. In addition, even if the global financial cycle reflects the transmission of shocks across countries, it is generally not clear where these shocks originate and which mechanisms govern their transmission.

In this paper, we show that US business cycle shocks are important drivers of the global financial cycle. We do so by studying the effects of US macroeconomic news releases on international asset markets. These news releases have large effects on international equity prices and the VIX—a close proxy of the global financial cycle.[1] They also induce the co-movement characteristic of the global financial cycle and explain a sizable fraction of its variation. Identifying this novel driver allows us to narrow the set of possible interpretations of the global financial cycle. In particular, we provide evidence that common shocks are unlikely to play an important role in this context. Rather, the estimated effects predominantly reflect the transmission of US-specific shocks to foreign economies. Further, the systematic conduct of US monetary policy is not the main mechanism through which US news affects foreign asset prices. The evidence instead points to a direct effect on the risk-taking behavior of international investors. Our paper complements prior work by Miranda-Agrippino and Rey (2020). Whereas they emphasize the contribution of US monetary policy shocks to the global financial cycle, we document that non-monetary US news also plays a central role in driving the global financial cycle.

Establishing a causal link between any potential driving force and the global financial cycle

[1] The VIX is the 30-day option-implied volatility index of the S&P 500.

is econometrically challenging. By its very nature, the global financial cycle is characterized by fast-moving financial variables such as risky asset prices and capital flows. At this point, it is well understood that identification strategies can fail at isolating the true underlying disturbances, if they do not account for the fact that financial markets respond quickly to new information (e.g., Gertler and Karadi, 2015; Ludvigson, Ma, and Ng, 2021).[2] In this paper, we resolve this identification problem by implementing a high-frequency event study. In particular, we analyze the intraday effects of US macroeconomic news surprises such as those associated with the nonfarm payroll employment release published monthly by the Bureau of Labor Statistics. While surprises about US macroeconomic variables are not structural shocks and care must be taken when interpreting their effects, this research design allows us to causally attribute asset price movements to these surprises. Of course, this research design also limits us to study asset prices as outcomes. Since the VIX has been shown to be a close proxy of the global financial cycle and since the co-movement of risky asset prices is a defining feature of the global financial cycle (Miranda-Agrippino and Rey, 2020), we view this research design nonetheless as a natural step to better understand the global financial cycle. Prior work has established that scheduled macroeconomic announcements are a unique source of variation to study asset price movements (e.g., Faust et al., 2007).

We begin our analysis with studying the effects of US macro news on major stock indexes of 27 countries from 1996 to 2019. Within a 30-minute window, these stock indexes show a statistically significant response and strongly co-move across countries. For instance, a positive surprise about nonfarm payroll employment generates a statistically significant increase in stock prices in all but one of the countries in our sample. We also document significant effects on the VIX and other implied volatility measures as well as commodity prices, which are often interpreted as indicators of risk appetite (Etula, 2013; Miranda-Agrippino and Rey, 2020).

High-frequency analyses often face the limitation that it is difficult to assess the economic importance of the identified relationship. We address this concern and demonstrate that the effects of US macroeconomic news on risky asset prices are both large and constitute an important driving force. The effects are large in the sense that international stock prices respond by a similar magnitude as the US stock market. Using the method by Altavilla, Giannone, and Modugno (2017), we further show that US macro news explains a sizable fraction of the variation in international stock prices at lower frequencies. On average, US macro news explains 23 percent of the quarterly variation in foreign equity prices once non-headline news is taken into account (Gürkaynak, Kısacıkoğlu, and Wright, 2020). This magnitude is comparable with its explanatory power for the S&P 500. US macroeconomic news further explains around 15 and 25 percent of the quarterly variation in the VIX and commodity prices, respectively. The concern that effects identified with high-frequency methods dissipate quickly therefore does not apply in our context.

The remainder of the paper interprets these findings and sheds light on the underlying mechanisms. We start by proposing a test for the presence of global common shocks to

[2] Miranda-Agrippino and Rey (2020) resolve this simultaneity problem by identifying monetary policy shocks from high-frequency asset price responses around Federal Reserve monetary policy releases.

address Bernanke's (2017) observation discussed above. Intuitively, if global common shocks drove international business cycles and stock markets, news releases in other countries should also be informative about the global state. Consequently, market participants should observe foreign macroeconomic news releases—even in small countries—and the US stock market should respond to this news. Our analysis shows that this is not the case. The S&P 500 essentially does not respond to foreign news releases. The evidence thus suggests a limited role of global common shocks and instead points to the transmission of US-specific shocks.

The same evidence also highlights a striking asymmetry: While US news has strong effects on foreign stock markets, foreign news has essentially no effect on the US. We carefully examine the robustness of this result and show that it is neither explained by lower timeliness of foreign news nor by lower measurement quality of foreign data. As an additional check, we confirm a similar asymmetry in the effects of monetary policy shocks. Unlike macroeconomic news releases, monetary policy shocks are known to be country-specific, that is, they have no common component. Thus, their effects are ideal for corroborating our interpretation of the asymmetry results for macro news. We find that US monetary policy shocks have effects on international equity markets that are approximately three times as large as equally sized shocks of the European Central Bank and the Bank of England.[3] These findings underscore the US' central position in the global monetary and financial system.

Lastly, we investigate the transmission channels through which US macro news affects foreign stock prices. To guide our analysis, we make use of the decompositions of (i) stock prices into a risk-free rate, an equity (risk) premium, and a growth expectations component (Boyd, Hu, and Jagannathan, 2005), and (ii) of bond yields into a risk-free rate and a term (risk) premium component. We then study the joint responses of bond yields and stock markets, both for the US and foreign markets. While bond yields do respond to US macroeconomic news, these responses can generally not explain the observed changes in stock prices. Instead, the evidence suggests that US news affects stock prices predominantly through growth expectations or risk premia. In the second step, we use direct measures for the equity premium (Martin, 2017), the term premium (Adrian, Crump, and Moench, 2013) as well as growth expectations (Gormsen and Koijen, 2020). Their responses to US macro news show that the behavior of risk premia is critical. Specifically, the estimates indicate that the equity premium is quantitatively more important for stocks than expected future dividends. Further, after news revealing greater-than-expected economic activity—such as positive surprises in nonfarm payroll employment—the equity premium falls while the term premium rises. Such news thus appears to induce a shift from less risky securities such as bonds to more risky securities such as stocks. Systematic US monetary policy reactions to news can for the most part *not* explain our findings.

Related literature Our paper relates to various topics in international finance and macroeconomics. First, our paper relates to work studying the global financial cycle. Important antecedents of Rey's (2013) seminal work include Diaz-Alejandro (1983, 1984), Calvo, Lei-

[3]For related findings, see Brusa, Savor, and Wilson (2020), Ca'Zorzi et al. (2020), and Miranda-Agrippino and Nenova (2022), among others.

derman, and Reinhart (1993, 1996), Reinhart and Reinhart (2008) and many others. These papers suggest a role for external and/or common drivers of countries' financial conditions. Following Rey (2013), several papers emphasize increased financial synchronization over recent decades, and discuss their implications (e.g., Bruno and Shin, 2015b; Obstfeld, 2015; Jordà et al., 2019).[4] Prior work has also shown that US monetary policy shocks affect global financial conditions. Bruno and Shin (2015a) provide evidence that US monetary policy affects the risk-taking behavior of international banks, Jordà et al. (2019) argue that US monetary policy drives global risk appetite and equity prices, and Miranda-Agrippino and Rey (2020) demonstrate that contractionary US monetary policy shocks worsen global financial conditions by affecting risky asset prices, leverage of global financial intermediaries, and international credit flows. We show that US macroeconomic news is a second causal driver of the global financial cycle, and that the outsized role of US-specific shocks is a broader phenomenon, not limited to monetary policy.[5]

More broadly, our paper relates to work studying the central role of the US in the international monetary and financial system—as reviewed in Gourinchas, Rey, and Sauzet (2019). Gourinchas and Rey (2007) emphasize the role of the US as world banker (or venture capitalist), Maggiori, Neiman, and Schreger (2020) document a dollar bias of international investors, and Goldberg and Tille (2008), Gopinath (2015), and Gopinath et al. (2020) document and study the importance of the US dollar as the dominant currency of trade invoicing. Our results show that an additional byproduct of the US' central position in the global financial system is that US macroeconomic news has large and persistent effects on global financial conditions while other countries' macro news has, if any, much smaller effects.

Lastly, our paper relates to prior work studying the high-frequency effects of US macroeconomic news releases on international financial markets.[6] Andersen et al. (2007) and Faust et al. (2007) analyze the effects of US news on financial markets in Germany and the United Kingdom. Ehrmann, Fratzscher, and Rigobon (2011) identify shocks through heteroscedasticity and study the interdependence of asset markets between the US and the Euro Area for multiple assets. We contribute to this literature in multiple ways. First, our sample contains a broader set of countries, including developing ones, while using intraday variation in asset prices. Second, we document the synchronized nature of foreign stock price responses in this large sample of countries and thereby establish a link between the US economy and the global financial cycle. Third, building on Altavilla, Giannone, and Modugno (2017) and Gürkaynak,

[4] Cerutti, Claessens, and Rose (2019) argue that common factors explain a relatively small fraction of the variation in international capital flows. Monnet and Puy (2019) study a broad sample of countries since 1950 and find that co-movement has increased for asset prices, but not for credit. They also study the effects of US monetary, fiscal, uncertainty, productivity shocks on the global financial cycle—with mixed results.

[5] Additional recent papers on the global financial cycle include Kalemli-Özcan (2019); Acalin and Rebucci (2020); Bekaert, Hoerova, and Xu (2020); Jiang, Krishnamurthy, and Lustig (2020); Miranda-Agrippino and Rey (2021); Davis and Van Wincoop (2021); Chari, Stedman, and Forbes (2022); Di Giovanni et al. (2022).

[6] A large set of papers study the effect of US macroeconomic releases on domestic financial markets (McQueen and Roley, 1993; Balduzzi, Elton, and Green, 2001; Gürkaynak, Sack, and Swanson, 2005b; Boyd, Hu, and Jagannathan, 2005; Rigobon and Sack, 2008; Beechey and Wright, 2009; Swanson and Williams, 2014; Gilbert et al., 2017; Law, Song, and Yaron, 2018; Gürkaynak, Kısacıkoğlu, and Wright, 2020). See Gürkaynak and Wright (2013) for a survey on high-frequency event studies in macroeconomics.

Kısacıkoğlu, and Wright (2020), we show that US macroeconomic news has persistent effects on international stock markets and explains a sizable fraction of their quarterly variation. Fourth, we document new properties of the transmission mechanism of US news to foreign markets.

Roadmap The remainder of the paper is structured as follows. The next section introduces our research design and discusses how to interpret the relationship between the measured surprises, the observed asset price responses, and the unobserved structural shocks. Section 3 introduces the data. We analyze the high-frequency effects of US news on international asset markets in Section 4. In Section 5, we demonstrate that the effects of US news on international asset prices are persistent and explain a sizable fraction of their quarterly variation. In Section 6, we document the asymmetric effects of US and foreign macro news, and discuss the role of global common shocks. Section 7 investigates the underlying channels of US macro news and Section 8 summarizes additional robustness checks and extensions. Section 9 concludes.

2 Research design

We are interested in assessing the effects of shocks, which drive the US business cycle, on global financial conditions. Since identifying structural disturbances is difficult and often requires strong assumptions, we instead study the effects of surprises about US macroeconomic news releases. This section discusses how to interpret these surprises and their effects on international asset prices.

Surprises Consider the release of US macroeconomic variable y at time t. For instance, the Bureau of Labor Statistics publishes nonfarm payroll employment typically at 8:30 am on the first Friday of each month. In this example, nonfarm payroll employment is the macroeconomic series of interest (y), and the announcement time t is 8:30 am on a given day. We construct surprises by subtracting from the US macroeconomic series y its forecast, that is,

$$s^{y}_{US,t} = \frac{y_{US,t} - E\left[y_{US,t}|\mathcal{I}_{t-\Delta^-}\right]}{\hat{\sigma}^{y}_{US}}, \tag{1}$$

where $y_{US,t}$ is the released value and $E\left[\cdot|\mathcal{I}_{t-\Delta^-}\right]$ is the expectation conditional on information available just prior to the release. To make the magnitudes of surprises comparable across macroeconomic series y, we also divide by the sample standard deviation of $y_{US,t} - E\left[y_{US,t}|\mathcal{I}_{t-\Delta^-}\right]$, denoted by $\hat{\sigma}^{y}_{US}$.

As equation (1) makes clear, macroeconomic surprises are by construction forecast errors and thus—up to a first order—linear combinations of structural shocks. Our research design therefore differs from common macroeconometric approaches, which attempt to directly identify structural disturbances: It is silent on the precise nature of structural shocks that generate the surprise.

Estimating equation Let i index countries and let $q_{i,t}$ denote the log of country i's asset price of interest. We study the effects of US macroeconomic surprises on a variety of international asset prices by estimating equations of the form

$$\Delta q_{i,t} = \gamma_i^y s_{US,t}^y + \varepsilon_{i,t}, \tag{2}$$

where we omit the constant and controls for simplicity. In this specification, Δ denotes a 30-minute change around the announcement time t. The error term $\varepsilon_{i,t}$ includes the effects of unmeasured news and/or noise on the asset price of interest.

The coefficient γ_i^y captures the effect of surprise $s_{US,t}^y$ on asset price $q_{i,t}$. It can be consistently estimated by Ordinary Least Squares (OLS) if the error term $\varepsilon_{i,t}$ is uncorrelated with the surprise. A large literature in macroeconomics and finance has argued that for sufficiently narrow windows around the release, this is likely the case. In this section, we proceed under the assumption that this condition holds. We will return to this question in Sections 3 and 4.

Interpretation of γ_i^y Under the identification assumption, the estimate of γ_i^y measures a causal effect. It is causal in the sense that we can unambiguously attribute systematic asset price responses to the surprises. Since surprises are not structural shocks, but linear combinations of structural shocks, the question arises how to interpret estimates of γ_i^y.

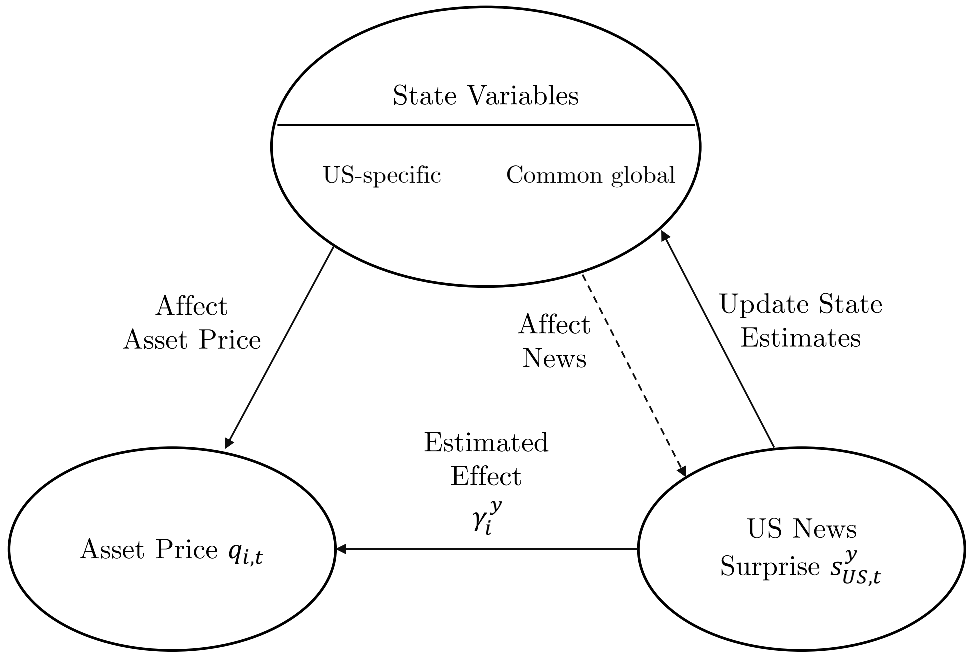
Building on Faust et al. (2007), we present a simple conceptual framework in Online Appendix A, which delivers estimating equation (2). In this framework, the coefficient γ_i^y captures the following intuition, which is illustrated in Figure 1. First, upon observing the surprise, market participants update their estimates of all state variables that generate economic fluctuations in the model. The solid arrow from the surprise to the state variables depicts this updating in the figure. Second, asset prices then respond to surprises because they depend on market participants' state estimates. This dependence is indicated by the solid arrow from state variables to the asset price $q_{i,t}$. The coefficient γ_i^y thus reflects both the updating of the state estimates and the dependence of the asset price on the state variables.

To build intuition, consider the following example. Suppose that shocks to total factor productivity (TFP), among other shocks, drive the US business cycle. Suppose further that market participants observe a positive surprise about US nonfarm payroll employment. Since this surprise may reflect a positive innovation to TFP, market participants may revise their estimate of TFP upwards upon observing the surprise. Higher expected productivity, in turn, may indicate greater expected future cash flows and thus lead to an increase in stock prices. Thus, the stock price responds to the news release because market participants update their TFP estimate and the stock price depends on TFP. We emphasize that the framework in Online Appendix A is agnostic on the set of structural disturbances that drive business cycle fluctuations and requires minimal assumptions on economic behavior.

If all underlying structural disturbances that drive the surprise $s_{US,t}^y$ originated in—and were specific to—the US, estimates of γ_i^y would reflect the *transmission of US-specific shocks* to country i's asset price $q_{i,t}$. However, the framework also makes clear that this need not be the case. It is also possible that the US and other countries are subject to *global common shocks*.[7] By directly affecting all countries' macroeconomic outcomes, including the US',

[7]Shocks are defined as global common if they are *exogenous structural disturbances directly affecting all countries.*

Figure 1: Interpretation of Country's i Asset Price Response to US News



Notes: This Figure illustrates the discussion in the text. Solid arrows display relevant relationships at the time of the news release, as captured by equation (2). The dashed arrow indicates that the relationship is predetermined at the time of the release.

such shocks could be reflected in the measured US surprises (see Figure 1). Foreign stock markets may respond to these surprises, because they reveal information about global common fundamentals (the common state vector). Prior work has acknowledged that global common shocks could drive business cycle co-movement (e.g., Canova and Marrinan, 1998; Canova, 2005). Further, Bernanke (2017, p. 23) notes that common shocks could drive the global financial cycle.[8]

The role of monetary policy For the interpretation of our results below, we briefly discuss the role of monetary policy *shocks* and monetary policy *reactions* to observed surprises.

Even though we can generally not infer structural shocks from observed surprises, we can rule out that monetary policy *shocks* are reflected in macroeconomic surprises. Any US monetary policy news is usually assumed to be fully revealed by Federal Open Market Committee (FOMC) announcements (Kuttner, 2001; Gürkaynak, Sack, and Swanson, 2005a), or

This definition is equivalent to modeling countries' shocks as being contemporaneously correlated (this is the definition adopted in Canova and Marrinan, 1998). In contrast, country-specific shocks are uncorrelated across countries. As an example, suppose that all countries in a model produce with production functions, which have a common productivity component. Exogenous fluctuations in this common productivity component would constitute an instance of such a global common shock (e.g., the baseline calibration in Backus, Kehoe, and Kydland (1992) has contemporaneously correlated productivity disturbances). Note that global common shocks generally differ from common pricing factors as frequently studied in the empirical asset pricing literature. In general equilibrium models, such pricing factors need not be exogenous.

[8]Online Appendix A discusses the possibility that shocks are specific to countries other than the US. To the extent that other countries are *small* relative to the US, such shocks are unlikely to play an important role.

other communication channels such as speeches by Fed officials (Cieslak, Morse, and Vissing-Jorgensen, 2019). Our macroeconomic surprises should therefore not reveal any new information about monetary policy. Since macroeconomic announcement times generally differ from Fed release times, our narrow 30-minute window also rules out that monetary policy news and macroeconomic news are conflated in our analysis. Hence, macroeconomic surprises should not reflect monetary policy *shocks*.

However, expected systematic monetary policy *reactions* as implied by a Taylor-type rule will affect how asset prices respond to surprises. For instance, upon observing a positive surprise about CPI inflation, the stock price response will depend on how aggressively market participants expect the Federal Reserve to respond to higher inflation. All else equal, the greater the expected increase in the policy rate, the more US stock prices should fall. We provide a more detailed discussion of this and other channels in Section 7.

Summary In summary, surprises are forecast errors and hence linear combinations of structural shocks. While our research design allows us to causally attribute asset price movements to these surprises, we can generally not identify the underlying structural shocks. Further, US macroeconomic surprises need not reflect US-specific structural shocks. It is also possible that foreign asset prices respond to US news releases because they reveal information about the global common state.

Relative to previous work on the global financial cycle, the key advantage of our research design is that it isolates conditional variation—from US macroeconomic surprises. We will use this variation (i) to show that shocks which drive the US business cycle also drive global financial conditions, and (ii) to study the mechanisms through which these shocks affect international asset prices. We will also propose a test for the presence of common shocks, which is specific to this research design. This test suggests that global common shocks are unlikely to be important in our context, and that the estimated effects predominantly capture the transmission of shocks from the US.

3 Data

In this section, we provide a brief overview of the data used for our main analysis.

3.1 US Macroeconomic News

The data on macroeconomic news releases comes from Bloomberg's US Economic Calendar. For each macroeconomic release, Bloomberg reports, among other things, release date and time, released value, and the median market expectation prior to the release. Table 1 provides an overview of the 12 major macroeconomic news series we focus on in Sections 4 and 7. This selection is inspired by previous studies in the literature (e.g., Faust et al., 2007; Rigobon and Sack, 2008; Gürkaynak, Kısacıkoğlu, and Wright, 2020). We treat different releases for the same macroeconomic variable—for instance, the advanced, second, and third release of GDP—as separate news series. For the interpretation of our results, it is often instructive to group the 12 major series into those providing information on US real economic activity and

Table 1: Overview of Major US Macroeconomic News

Announcement	Release Time	Frequency	Category	Observations
Capacity Utilization	9:15 am	Monthly	Real Activity	274
CB Consumer Confidence	10:00 am	Monthly	Real Activity	273
Core CPI	8:30 am	Monthly	Price	275
Core PPI	8:30 am	Monthly	Price	275
Durable Goods Orders	8:30 am	Monthly	Real Activity	266
GDP A	8:30 am	Quarterly	Real Activity	91
Initial Jobless Claims	8:30 am	Weekly	Real Activity	1166
ISM Mfg Index	10:00 am	Monthly	Real Activity	277
New Home Sales	10:00 am	Monthly	Real Activity	267
Nonfarm Payrolls	8:30 am	Monthly	Real Activity	274
Retail Sales	8:30 am	Monthly	Real Activity	275
UM Consumer Sentiment P	10:00 am	Monthly	Real Activity	247

Notes: This table displays the 12 major macroeconomic series we focus on in most of the paper. Online Appendix Table B1 shows the full set of series considered in the paper. The sample ranges from October 1996 to December 2019. *Frequency* refers to the frequency of the data releases and *Observations* to the number of observations (surprises) of a macroeconomic series in our sample. *Category* specifies if the news release is predominantly informative about real activity or prices. Abbreviations: A—advanced; P—preliminary; Mfg—Manufacturing; CB—Chicago Board; UM—University of Michigan; ISM—Institute for Supply Management.

those providing information on prices (Beechey and Wright, 2009).[9]

When studying the explanatory power of US macroeconomic news in Section 5 we use *all* available US macroeconomic news series. These are listed in Online Appendix Table B1. As discussed below, we will also use this broader set of announcements as controls. For more details on the macro news data, see Online Appendix B.1.

We use the median market expectation of the release as our measure of $E\left[y_{US,t}|\mathcal{I}_{t-\Delta^-}\right]$ when constructing surprises based on equation (1). Since Bloomberg allows forecasters to update their prediction up until the release time, these forecasts should reflect all publicly available information at the time. As noted above, surprises are standardized so that the coefficient γ_i^y measures the effect of a one standard deviation surprise. For ease of interpretation, we flip the sign of Initial Jobless Claims surprises. A positive sign thus corresponds to positive news about real economic activity—consistent with the other releases.

Online Appendix Figure C1 shows the resulting time series of standardized surprises for each macroeconomic variable. Reassuringly, all series of surprises are centered at zero. Further, there is no discernible pattern of autocorrelation, and there is no systematic trend in the standard deviation of surprises. Some series such as Initial Jobless Claims and Retail Sales display somewhat higher volatility during recessions. In contrast, other series such as Core PPI and New Home Sales, have lower volatility during downturns. Overall, there is no indication that using these surprises as our identifying variation is econometrically problematic.

[9]As discussed in Section 2, it is possible that both categories provide information about the same underlying macroeconomic shocks. The classification into price and real activity news should therefore be regarded as pragmatic rather than conceptual. It turns out that this grouping is useful for summarizing and interpreting our findings.

Table 2: Intraday Data on International Stock Markets

Name	Ticker	Sample	Country	ISO
MERVAL	.MERV	1996–2019	Argentina	ARG
ATX	.ATX	1996–2019	Austria	AUT
BEL 20	.BFX	1996–2019	Belgium	BEL
Bovespa	.BVSP	1996–2019	Brazil	BRA
S&P/TSX	.GSPTSE*	2000–2019	Canada	CAN
SMI	.SSMI	1996–2019	Switzerland	CHE
IPSA	.SPIPSA*	1996–2019	Chile	CHL
PX	.PX*	1999–2019	Czech Republic	CZE
DAX	.GDAXI	1996–2019	Germany	DEU
OMX Copenhagen 20	.OMXCXC20PI*	2000–2019	Denmark	DNK
IBEX 35	.IBEX	1996–2019	Spain	ESP
OMX Helsinki 25	.HEX25	2001–2019	Finland	FIN
CAC 40	.FCHI	1996–2019	France	FRA
FTSE 100	.FTSE	1996–2019	United Kingdom	GBR
FTSE/Athex Large Cap	.ATF	1997–2019	Greece	GRC
BUX	.BUX	1997–2019	Hungary	HUN
ISEQ	.ISEQ	1996–2019	Ireland	IRL
FTSE MIB	.FTMIB*	1996–2019	Italy	ITA
S&P/BMV IPC	.MXX	1996–2019	Mexico	MEX
AEX	.AEX	1996–2019	Netherlands	NLD
OBX	.OBX	1996–2019	Norway	NOR
WIG20	.WIG20	1997–2019	Poland	POL
PSI-20	.PSI20	1996–2019	Portugal	PRT
MOEX Russia	.IMOEX*	2001–2019	Russia	RUS
OMX Stockholm 30	.OMX	1996–2019	Sweden	SWE
BIST 30	.XU030	1997–2019	Turkey	TUR
FTSE/JSE Top 40	.JTOPI	2002–2019	South Africa	ZAF

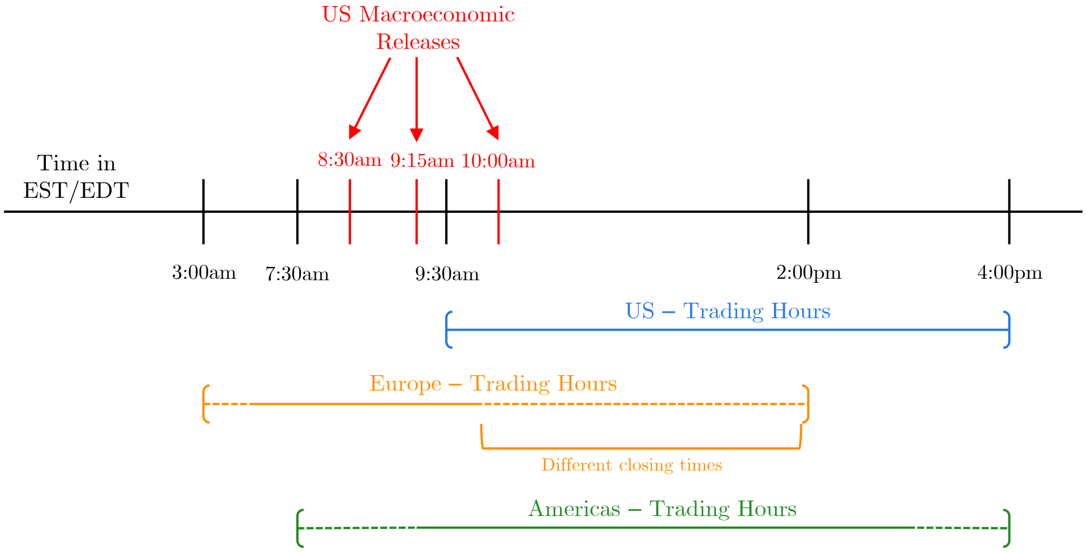
Notes: The table shows the stock market indexes used in our analysis. The data is from *Thomson Reuters Tick History*. For all series, the sample period ends in December 2019. *For Canada, Chile, the Czech Republic, Denmark, Italy and Russia, the ticker of the stock index changes over our sample period. Hence, we also use the previous tickers, which are .TSE300 for Canada, .IPSA and .SPCLXIPSA for Chile, .PX50 for the Czech Republic, .KFMX for Denmark, .MIB30 and .SPMIB for Italy, and .MCX for Russia. *Ticker* refers to the Reuters Instrument Code (RIC), and *ISO* denotes the 3-digit ISO country code.

3.2 Financial Data

The data on asset prices comes from the *Thomson Reuters Tick History* dataset and is obtained via *Refinitiv*. We use intraday data for most analyses. As shown by prior work—mostly in a domestic context—moving from daily to intraday data leads to lower risk of confounding by other news releases, and to increased precision by mitigating noise. Using intraday data is likely even more important when studying the effects on international markets since most countries are more open than the US. A country's stock market is driven by domestic *and* foreign news, making US news releases just one among many sources of information throughout the trading day.

Our primary outcomes of interest are minute-by-minute series of 27 countries' major stock indexes. Table 2 provides an overview of these. The table also shows the sample periods over

Figure 2: US Macroeconomic Releases and International Stock Market Trading Hours



Notes: This figure provides an overview of release times and trading hours of stock markets in our sample. Note that the trading hours of South Africa and Turkey are represented by the European trading hours.

which these indexes are available to us. For Canada, Chile, the Czech Republic, Denmark, Italy, and Russia, the stock indexes change their ticker symbols during the sample period. In these cases, we merge the series with their predecessors in a consistent fashion. We inspect each data series for potential misquotes, and remove them if necessary. Throughout the paper, we use a country's 3-digit ISO code to refer to its stock index (e.g., DEU instead of DAX). Besides the data on international stock markets, we use intraday data on various other asset prices. We defer a more detailed discussion to the relevant sections below. Online Appendix B.2 provides an overview of all financial instruments employed throughout the paper.

Our intraday analysis of international equity markets requires that the time window around a particular news release lies within the trading hours of the respective foreign stock market. The country composition of our sample reflects this constraint. For instance, Asian and Australian equity markets are closed during almost all release times and are thus not included in our sample. When comparing US and foreign stock price responses, we rely on data on E-mini S&P 500 futures, which are traded outside of regular trading hours. Hence, we do not need to limit our analysis to announcements for which US markets are open. Figure 2 visualizes the timing of news releases and trading hours for the stock markets in our sample. Further, Online Appendix Table B4 summarizes which countries' equity markets are open for each of the 12 main announcements.

4 High-Frequency Effects of US Macro News

In this section, we implement a high-frequency event study and estimate the effect of US macroeconomic releases on risky asset prices. Due to their importance for the global financial cycle, we are interested in the effects on international stock indexes, the VIX and other implied volatility measures, as well as commodity prices. We show that all of these asset prices strongly respond to US news. Importantly, we document that US news releases induce co-movement of international equity markets.

4.1 International Stock Markets

4.1.1 Pooled Effects

We begin our empirical analysis with demonstrating that international stock indexes respond to the release of news about the US economy. As discussed in Section 2, we estimate pooled regressions of the form

$$\Delta q_{i,t} = \alpha_i + \gamma^y s^y_{US,t} + \sum_{k \neq y} \gamma^k s^k_{US,t} + \varepsilon_{i,t}, \tag{3}$$

where $\Delta q_{i,t} = q_{i,t+20} - q_{i,t-10}$ is the 30-minute log-change of country i's stock market index.[10] Further, $s^y_{US,t}$ is the surprise of interest and $\varepsilon_{i,t}$ captures the effects of unmeasured news and/or noise. Note that the pooled effect γ^y is informative about the *average* effect on international stock markets. It masks, however, potential heterogeneity in the responses of the 27 stock indexes in our sample. Since such heterogeneity (or the lack thereof) is of interest for our research question, we study the country-specific effects below.

We include other surprises about US macroeconomic variables, $s^k_{US,t}$, which are published within the time window we study, as controls. For instance, the Bureau of Labor Statistics publishes Nonfarm Payrolls together with the Unemployment Rate (and other macroeconomic variables) as part of the US employment report. Attributing asset price changes solely to the surprise about Nonfarm Payrolls could therefore be misleading. Note that we consider all 66 announcements as listed in Online Appendix Table B1 as controls, except for those, which by construction convey the same information as the release of interest.[11]

The identification assumption for the consistent estimation of γ^y holds that, conditional on controls, error $\varepsilon_{i,t}$ is uncorrelated with the surprise $s^y_{US,t}$. To account for the fact that surprises on the right-hand side are US-specific and thus perfectly correlated across foreign countries, we two-way cluster standard errors by announcement and by country.

Table 3 shows the estimates of γ^y for the 12 major macroeconomic releases. Two results emerge from the table. First, all announcements have a significant effect at the one percent level with the exception of the Capacity Utilization announcement, which is significant at the

[10] More precisely, $\Delta q_{i,t} = \log((Q_{i,t+15} + ... + Q_{i,t+25})/11) - \log((Q_{i,t-15} + ... + Q_{i,t-5})/11)$, where $Q_{i,t}$ is country i's stock market index, and then express this change in basis points.

[11] For instance, Capacity Utilization is constructed by dividing Industrial production by a slow-moving estimate of capacity. When studying the effect of Capacity Utilization on international equity markets, we therefore exclude Industrial Production from the set of controls. Including Industrial Production as a control would make the coefficient on Capacity Utilization difficult to interpret—due to collinearity problems. To avoid such collinearity problems, we choose the set of controls as follows: For Core CPI and Core PPI, we exclude CPI and PPI, respectively. For Durable Goods Orders, we exclude Durable Goods Orders Excluding Transportation (Durable Ex Transportation). For Nonfarm Payrolls, we exclude Private and Manufacturing Nonfarm Payrolls (Private and Mfg Payrolls). For Retail Sales, we exclude Retail Sales Excluding Autos (Retail Sales Ex Auto).

Table 3: Effects of US News on International Stock Markets

	Capacity Utilization	CB Consumer Confidence	Core CPI	Core PPI	Durable Goods Orders	GDP A
Stock Index (bp)						
News	5.36**	12.35***	-8.84***	-4.87***	5.63***	17.60***
	(2.28)	(2.02)	(1.89)	(1.29)	(1.60)	(3.36)
R^2	0.04	0.13	0.10	0.15	0.10	0.26
Observations	6054	6041	5717	5828	5610	1911
	Initial Jobless Claims $\cdot(-1)$	ISM Mfg Index	New Home Sales	Nonfarm Payrolls	Retail Sales	UM Consumer Sentiment P
Stock Index (bp)						
News	4.89***	11.71***	4.23***	17.06***	10.52***	5.61***
	(0.73)	(2.24)	(1.40)	(2.99)	(1.68)	(1.54)
R^2	0.09	0.12	0.03	0.13	0.15	0.04
Observations	24334	5548	5908	5688	5786	5726

Notes: This table presents estimates of γ^y of equation (3) for each of the 12 macroeconomic announcements. The stock index changes are expressed in basis points. Standard errors are two-way clustered by announcement and by country, and reported in parentheses. ***, **, and * indicate significance at the 1, 5, and 10 percent level.

five percent level. Second, positive news about US real activity leads to an increase in stock prices. As we will discuss in Section 7 below, this effect is consistent with increased risk-taking of international investors and/or higher expected future dividends after such surprises. In contrast, inflation surprises—as captured by positive surprises in the Core CPI and Core PPI—lead to a decrease in stock prices. We show in Section 7 that this result is at least in part driven by higher interest rates.

Kurov et al. (2019) have documented that some asset prices drift prior to certain US macroeconomic news releases. Such drifts may reflect information leakage or superior forecasting ability relative to the median forecast and cast doubt on market efficiency—which our analysis relies on. As Online Appendix Figure C2 shows, international equity prices do not drift prior to the news releases we study (at least not during the time window relevant for our analysis). This is in line with Lucca and Moench (2015) who also do not find evidence for pre-announcement drifts around US macro releases.

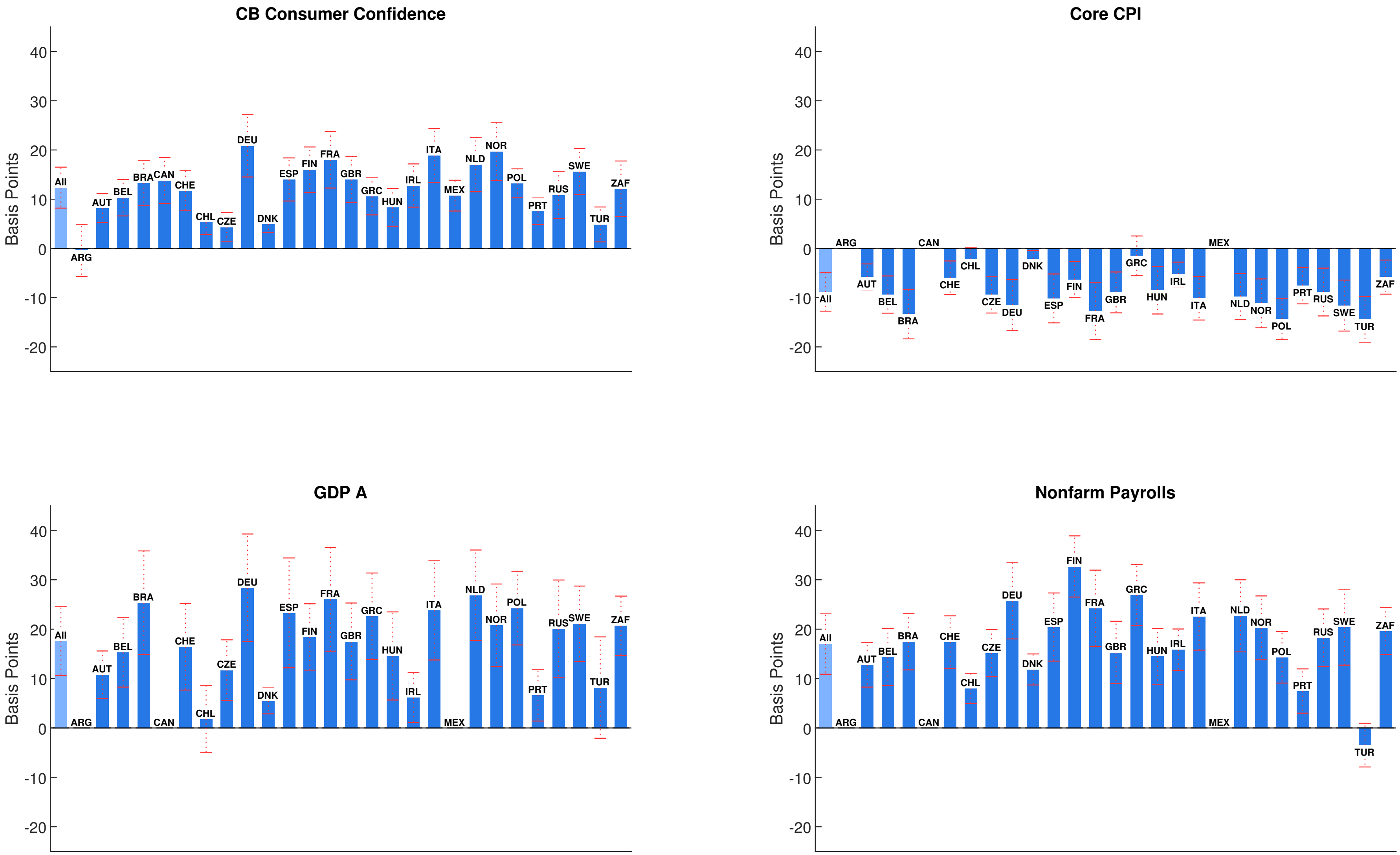
4.1.2 Cross-country Heterogeneity

We next study country-specific effects and show that US macroeconomic news induces co-movement across markets. In particular, we estimate

$$\Delta q_{i,t} = \alpha_i + \gamma_i^y s_{US,t}^y + \sum_{k \neq y} \gamma_i^k s_{US,t}^k + \varepsilon_{i,t}, \tag{4}$$

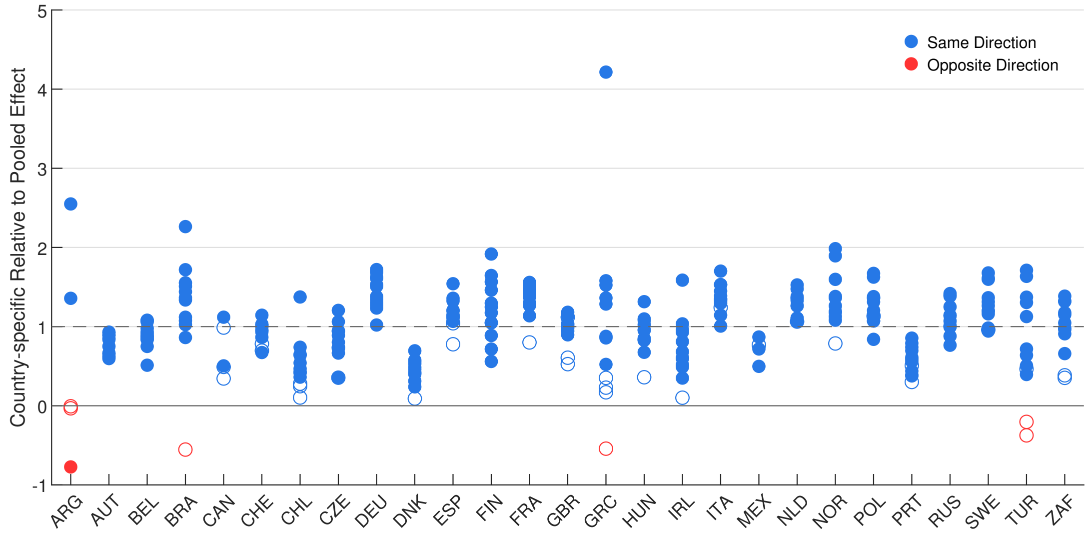
where $\Delta q_{i,t} = q_{i,t+20} - q_{i,t-10}$. Different from equation (3), the coefficients γ_i^y and γ_i^k are now specific to each country.

Figure 3: Effects of US News on International Stock Markets by Country



Notes: This figure shows the stock index responses for four selected announcements. The stock index changes are expressed in basis points. The light blue bar shows the pooled effect, i.e., the estimate of common coefficient γ^y of equation (3), while the dark blue bars show the country-specific effect, i.e., the estimate of γ_i^y of equation (4). Missing country bars depict cases in which the country is dropped because it had less than 24 observations for a given announcement. The red error bands depict 95 percent confidence intervals, where standard errors are two-way clustered by announcement and by country. Analogous bar charts for all news releases are shown in Online Appendix Figure C3.

Figure 4: Countries' Stock Market Responses Relative to Pooled Response



Notes: The figure plots the country-specific stock index responses relative to the pooled response for all 12 announcements, or formally, $\hat{\gamma}_i^y/\hat{\gamma}^y$, where the estimates are obtained from estimating equations (3) and (4). Blue (red) circles indicate that the country's response has the same (opposite) sign as the pooled effect. Filled circles indicate significance at the 5 percent level while an empty circle indicates an insignificant effect. For a given announcement, country-specific estimates obtained from fewer than 24 observations are dropped.

Figure 3 illustrates countries' stock index responses for four of the 12 announcements. Strikingly, for a given announcement the sign of the response is identical for all countries whenever statistically significant. That is, US macroeconomic news not only affects international stock markets but they also lead to *correlated* asset price responses. This co-movement of risky asset prices is a defining feature of the global financial cycle (Miranda-Agrippino and Rey, 2020).

Figure 4 summarizes this finding for all 12 announcements by plotting the country-specific effect $\hat{\gamma}_i^y$ relative to the pooled effect $\hat{\gamma}^y$ (estimated from equation (3)). Circles above zero indicate cases in which the country-specific effect has the same sign as the pooled effect. The fact that almost all circles are positive confirms the results of Figure 3. Figure 4 also illustrates systematic heterogeneity in responsiveness across countries. While the Netherlands, for example, responds more strongly than the average for all 12 announcements, countries such as Austria, Denmark, and Portugal always respond less than the average. We explore this point in Supplementary Appendix S6 where we examine whether this responsiveness correlates with observables such as financial openness.

4.1.3 Assessing the Magnitude

While our high-frequency event study above allows us to establish a causal relationship between US news and foreign stock markets, it comes at the cost that the economic significance of this finding is not immediately obvious. To shed light on this question, we next assess the effect size by comparing it to a benchmark. In particular, we compare the foreign stock price response to the response of the S&P 500.

To do so, we estimate equation (3) after replacing the left hand side with $\Delta q_{US,t} - \Delta q_{i,t}$,

Table 4: Effects on US Stock Market Relative to International Markets

	Capacity Utilization	CB Consumer Confidence	Core CPI	Core PPI	Durable Goods Orders	GDP A
Stock Index Diff. (bp)						
News	-0.44	3.45**	-4.67***	-0.73	-1.01	-0.95
	(1.10)	(1.34)	(1.18)	(0.81)	(0.87)	(2.00)
R^2	0.00	0.04	0.05	0.02	0.04	0.05
Observations	5535	5953	5575	5668	5610	1871
	Initial Jobless Claims $\cdot(-1)$	ISM Mfg Index	New Home Sales	Nonfarm Payrolls	Retail Sales	UM Consumer Sentiment P
Stock Index Diff. (bp)						
News	0.59	4.13**	-0.58	2.83	-1.13	-1.68
	(0.44)	(1.88)	(0.90)	(2.28)	(1.16)	(1.15)
R^2	0.01	0.06	0.01	0.03	0.03	0.01
Observations	24122	5432	5893	5578	5593	5087

Notes: This table presents estimates of γ^y as defined in equation (3) after replacing the left hand side with $\Delta q_{US,t} - \Delta q_{i,t}$ for each of the 12 macroeconomic announcements. The stock index changes are expressed in basis points. Standard errors are two-way clustered by announcement and by country, and reported in parentheses. ***, **, and * indicate significance at the 1, 5, and 10 percent level.

where $\Delta q_{US,t}$ is the 30-minute log-change in the front-month E-mini S&P 500 futures contract, and $\Delta q_{i,t}$ is the 30-minute log-change of country i's stock market index as above. A positive coefficient γ^y now indicates that the response of the S&P 500 is greater than the response of the foreign stock price index. We follow earlier studies and use E-mini S&P 500 futures contracts for this analysis (e.g., Hasbrouck, 2003). These are highly liquid, traded outside of regular hours, and thus available for all announcements.

Table 4 shows the estimates. Strikingly, we find evidence that the US stock market responds differently from foreign stock markets for only 3 out of 12 announcements. In absolute terms, the US response is greater for the CB Consumer Confidence, the Core CPI, and the ISM Manufacturing Index. (Recall that stock markets respond negatively to Core CPI announcements.) In the remaining cases, we can neither reject the null hypothesis of equally-sized responses, nor do the insignificant point estimates suggest a greater response of the S&P 500. For news about real activity, the insignificant point estimates are often negative, if at all hinting at greater responses of foreign equity markets. In sum, foreign stock price responses to US news are often comparable in magnitude to the response of US stock prices.

4.2 The VIX and Other Risky Asset Prices

In this section, we estimate the effects of US macro news on the VIX, a measure of risk aversion and uncertainty, as well as other risky asset prices. Declines in the VIX are typically interpreted as signalling increasing willingness of investors to take risk. Various papers highlight the important role of the VIX for international financial markets. Rey (2013) shows that the VIX is a close proxy of the global financial cycle, Forbes and Warnock (2012) emphasize the correlation of the VIX with international capital flows, and Bruno and Shin (2015a) link it to global banks' leverage.

Analogous to specification (3), we estimate the effect of US news on the 30-minute log-change in the VIX:

$$\Delta q_t = \alpha + \gamma^y s^y_{US,t} + \sum_{k \neq y} \gamma^k s^k_{US,t} + \varepsilon_t, \tag{5}$$

where $s^y_{US,t}$ is the announcement surprise of interest, $s^k_{US,t}$ are other surprises released in the same time window, and $\Delta q_t = q_{t+20} - q_{t-10}$ is the 30-minute log-change in the VIX. If the stock market is not open at the announcement time, we instead use changes in the front-month VIX futures contract.[12] Since VIX futures are available for the relevant trading hours only since 2011, the sample sizes are often smaller than before (see Online Appendix Table B3). Due in part to the small samples sizes for the VIX, we also study the VSTOXX, which is the implied volatility index for the Euro Area stock index STOXX 50. As shown in Miranda-Agrippino and Rey (2021), this index is also highly correlated with the global financial cycle and high-frequency data is available for all announcements from 2005 onwards.

Table 5 reports the estimates of these regressions. 9 out of 12 announcements show a strong and significant effect on the VIX. Positive news about real economic activity leads to a reduction in the VIX, confirming that US macroeconomic news drives the global financial cycle. A comparison to the estimates in Table 3 makes clear that after most announcements stock prices co-move negatively with the VIX. The estimates for the VSTOXX confirm this co-movement (and are significant throughout). To the extent that the implied volatility indexes serve as a rough proxy for the equity premium (Martin, 2017), this negative co-movement suggests that changes in the equity risk premium drive part of the stock price response. We provide more evidence on this in Section 7. In Online Appendix Table C1, we also report results for the implied volatility indexes of Germany (VDAX), the United Kingdom (VFTSE), and France (VCAC). The effects of US macro news are robust across these measures.[13]

Lastly, in Supplementary Appendix S1, we study the effects on commodity prices as additional measures of risky asset prices. For the majority of news releases, we find a significant effect on a common factor extracted from several commodity prices. The signs are as expected. Positive (negative) news about real activity leads to an increase (decrease) in commodity prices. Thus, our findings for other risky asset prices confirm that US macro news drives the global financial cycle.

5 Explanatory Power of US Macro News at Lower Frequencies

In this section, we demonstrate that the effects of US news on international stock markets are persistent and explain a sizable share of their variation.

Headline news We apply Altavilla, Giannone, and Modugno's (2017) method to assess the explanatory power of US macro news and thus switch from our earlier intraday event study approach in the previous section to a daily time series analysis. In a first step, we estimate

[12] In our sample, the correlation of the daily returns of the VIX and front-month VIX futures contract is 78 percent.

[13] In unreported robustness checks, we have confirmed that the results in Table 5 do not change fundamentally when we drop the zero lower bound episode from the sample.

Table 5: Effects of US News on VIX and VSTOXX

	Capacity Utilization	CB Consumer Confidence	Core CPI	Core PPI	Durable Goods Orders	GDP A
VIX (bp)						
News	-15.66	-65.29***	37.14***	-5.21	-5.42	-45.65***
	(11.59)	(12.55)	(13.24)	(8.50)	(5.74)	(16.20)
R^2	0.03	0.13	0.21	0.40	0.26	0.35
Observations	108	270	105	108	108	36
VSTOXX (bp)						
News	-25.61**	-50.99***	46.23***	24.82**	-23.13**	-94.80***
	(12.24)	(12.17)	(11.80)	(10.47)	(11.06)	(20.19)
R^2	0.07	0.07	0.15	0.30	0.12	0.32
Observations	175	175	175	175	174	59
	Initial Jobless Claims $\cdot(-1)$	ISM Mfg Index	New Home Sales	Nonfarm Payrolls	Retail Sales	UM Consumer Sentiment P
VIX (bp)						
News	-15.09**	-66.21***	-25.38*	-118.04***	-75.13***	-40.81***
	(6.38)	(18.08)	(13.35)	(27.15)	(18.79)	(14.95)
R^2	0.13	0.13	0.06	0.27	0.32	0.05
Observations	464	270	264	107	106	230
VSTOXX (bp)						
News	-26.51***	-101.65***	-36.83**	-158.09***	-61.44***	-41.84***
	(4.89)	(19.46)	(16.65)	(19.80)	(10.27)	(12.85)
R^2	0.14	0.27	0.12	0.32	0.30	0.07
Observations	754	163	174	171	175	176

Notes: For all 12 announcements, this table shows estimates of γ^y obtained from equation (5), where the left-hand side is the 30-minute log-change in the front-month VIX futures contract or the VSTOXX, expressed in basis points. For CB Consumer Confidence, UM Consumer Sentiment P, ISM Mfg Index, and New Home Sales, we are able to use the VIX instead of the VIX futures due to the late announcement time. Heteroskedasticity-robust standard errors are reported in parentheses. ***, **, and * indicate significance at the 1, 5, and 10 percent level.

the specification

$$\Delta q_{i,d} = \alpha_i + \sum_k \beta_i^k s_{US,d}^k + \varepsilon_{i,d}. \tag{6}$$

Here, d indexes time in days and $\Delta q_{i,d}$ is the daily return of country i's stock price index as measured by the log-difference from market closing to market closing. The sum on the right-hand side now includes *all* available announcements as listed in Online Appendix Table B1. By focusing on daily log-returns, we circumvent the problem that some foreign markets are closed for some announcements. Hence, the set of US news releases that drive foreign asset prices in specification (6) is identical for all countries.[14] Note that all coefficients are country-specific. A surprise $s_{US,d}^k$ takes the value 0 if no news is released on a given day. Since the coverage of news releases is incomplete in the late 1990s, the sample period now ranges

[14]Relative to Altavilla, Giannone, and Modugno (2017), our set of announcements includes more macroeconomic news releases. However, we exclude news about monetary policy.

from January 1, 2000 to December 31, 2019.

Next, we define the daily headline news index $hni_{i,d}$ as the fitted value from equation (6), and aggregate this predicted value to the desired time horizon h (in days), $hni_{i,d}^{(h)} = \sum_{j=0}^{h-1} hni_{i,d-j}$. Letting $\Delta q_{i,d}^{(h)} = q_{i,d} - q_{i,d-h} = \sum_{j=0}^{h-1} \Delta q_{i,d-j}$ denote the h-day log-return of stock index q_i, we estimate in a second step the specification

$$\Delta q_{i,d}^{(h)} = \alpha_i^{(h)} + \beta_i^{(h)} hni_{i,d}^{(h)} + \varepsilon_{i,d}^{(h)}. \tag{7}$$

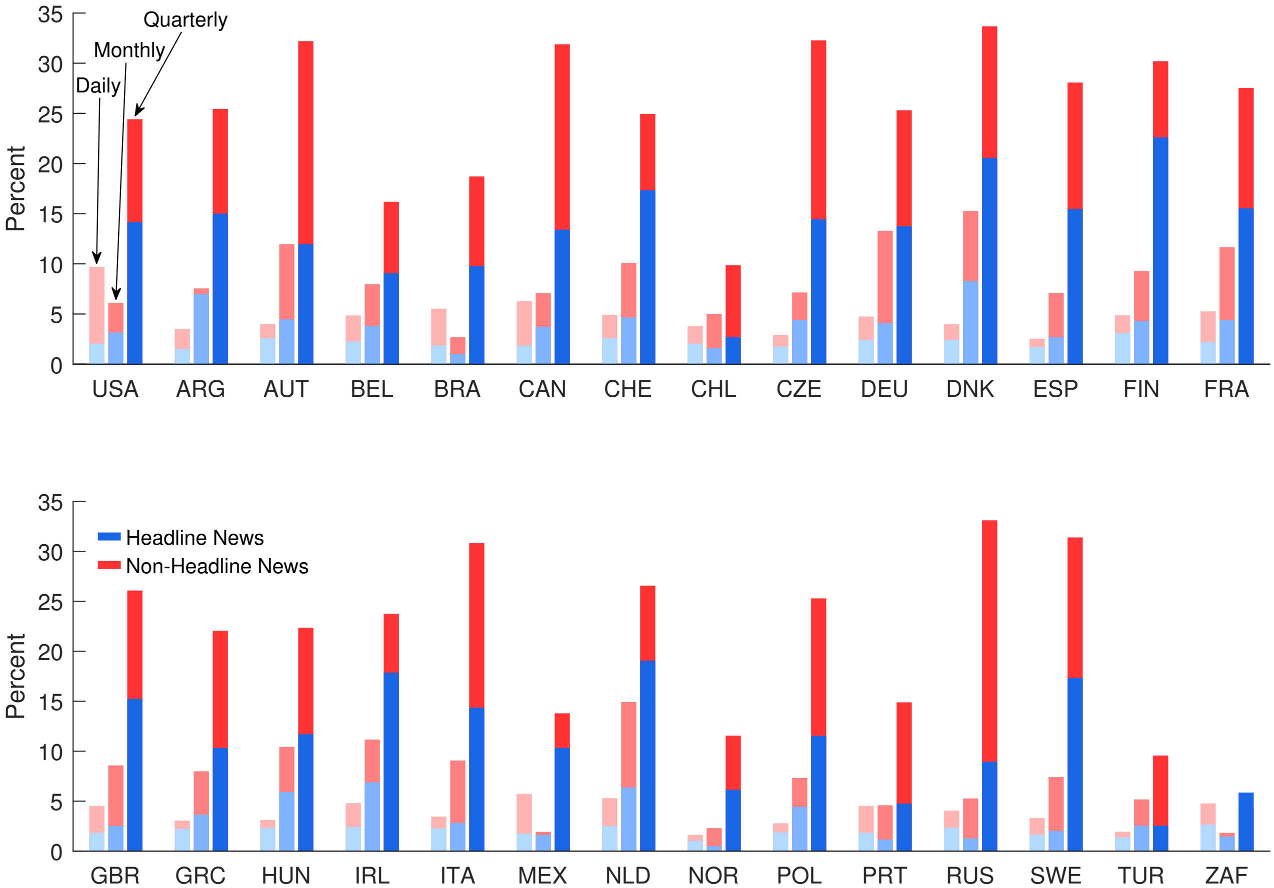
The statistic of primary interest is the R-squared of regression (7). It measures the explanatory power of the headline US macroeconomic news releases at aggregation horizon h and is therefore informative about how persistent the effects of macroeconomic news are relative to residual driving forces. Additionally, if the coefficient $\beta_i^{q,h}$ is greater (smaller) than one, macroeconomic news exerts a delayed (mean-reverting) effect. As in Altavilla, Giannone, and Modugno (2017), we consider aggregation to the monthly and quarterly frequency.

Non-headline news Following Gürkaynak, Kısacıkoğlu, and Wright (2020), we also incorporate the effects of "non-headline news" into our measurement of explanatory power. This news describes a part of macro releases, which is not captured by the surprises we have studied so far. Non-headline news is therefore latent, that is, it is not observed by the econometrician. However, as market participants observe such news, it can affect asset prices. For example, the Bureau of Labor Statistics publishes the nonfarm payroll employment number as part of the US employment report, which varies in length between 20 and 40 pages over our sample period. These pages contain additional macroeconomic data, for which no survey expectations exists, as well as text to provide context and details. All of this information potentially qualifies as non-headline news.

Gürkaynak, Kısacıkoğlu, and Wright (2020) propose an estimation procedure to recover non-headline news factors using the Kalman filter and demonstrate that they are important for explaining the observed asset price reactions around macroeconomic announcements. We closely follow their procedure. With the estimated non-headline news factors in hand, we can add them as additional regressors into equation (6). The fitted value is then a daily *broad news index*, and we can obtain the combined explanatory power of headline and non-headline news from a modified version of equation (7). Details on the estimation as well as robustness checks are available in Supplementary Appendix S2.

Results Figure 5 shows the daily, monthly, and quarterly R-squared for the foreign stock indexes by country. The blue bars display the contributions of headline news while the red bars display the contributions of non-headline news. The figure shows that the explanatory power of US news for foreign stock indexes increases at lower frequencies for both headline and non-headline news. In an overwhelming number of cases, the R-squared at the quarterly frequency exceeds the R-squared at the monthly frequency, which in turn, exceeds the R-squared at the daily frequency. The explanatory power of US news is sizable at the quarterly frequency, often explaining between 15 and 35 percent of the variation. On average, US news explains 23 percent of the quarterly variation. For comparison, we repeat the analysis for the

Figure 5: Daily, Monthly, and Quarterly R-Squared for Stock Indexes



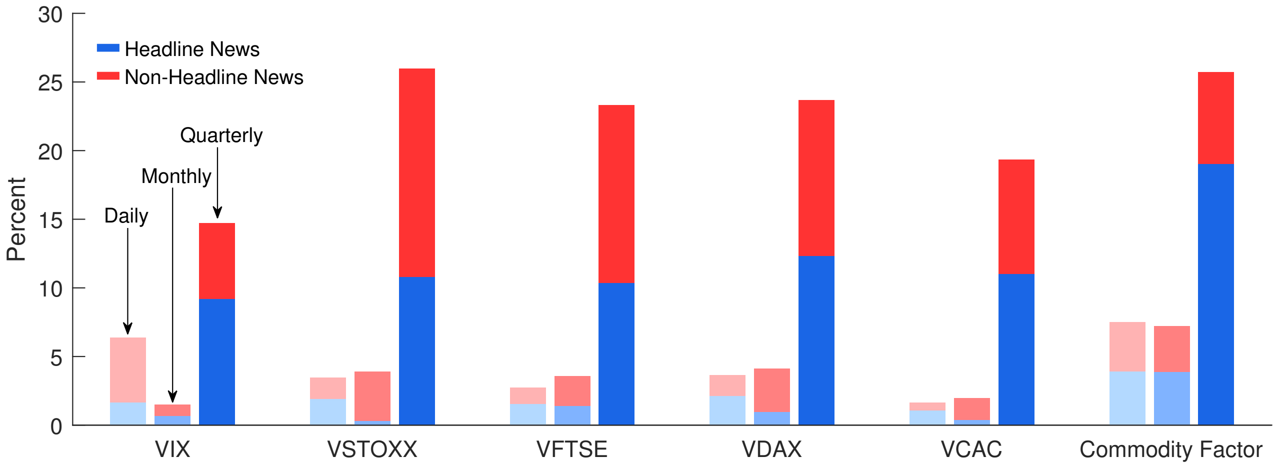
Notes: For each country's stock index, this figure plots the R-squared of equations (6) for the daily frequency, and the R-squared of equations (7) for the monthly and quarterly frequency. The left, middle, and right bar for each country indicate, respectively, the R-squared of the daily, monthly, and quarterly regression. For a given country and frequency, the blue bar represents the R-squared of the headline surprises of US macroeconomic news, whereas the red bar displays the increment in R-squared once non-headline news is included. The sample runs from January 1, 2000 to December 31, 2019.

S&P 500, and report the R-squared first in Figure 5. US macroeconomic news explains an even greater share of stock price movements in several foreign countries than it does in the US.

The increased R-squared at lower frequencies imply that the effects of US macroeconomic news are more persistent than residual driving forces of international stock prices. Online Appendix Table C2 reports the monthly and quarterly estimates of $\beta_i^{(h)}$ from equation (7), and shows that at least part of this persistence is due to delayed effects of the macroeconomic news. For several countries, we can reject the null hypothesis that $\beta_i^{(h)} = 1$.

Overall, the explanatory power of US macro news for international stock markets at lower frequencies is striking. Reassuringly, our estimates for headline news and the US market are similar to those by Altavilla, Giannone, and Modugno (2017). We also repeat this exercise for US dollar-denominated foreign exchange rates. The results, shown in Online Appendix

Figure 6: Daily, Monthly, and Quarterly R-Squared for Volatility and Commodity Indexes



Notes: This figure plots the R-squared of equations (6) for the daily frequency, and the R-squared of equations (7) for the monthly and quarterly frequency, where we now use log-returns of the volatility indexes or the commodity factor instead of country's i stock index. The left, middle, and right bar indicate the R-squared of the daily, monthly, and quarterly regression, respectively. For a given country and frequency, the blue bar represents the R-squared of the headline surprises of US macroeconomic news, whereas the red bar displays the increment in R-squared once non-headline news is included. The sample runs from January 1, 2000 to December 31, 2019 for the volatility indexes, and from May 7, 2007 to December 31, 2019 for the commodity factor.

Figure C4, make clear that the methodology does not mechanically lead to an increase in the R-squared at lower frequencies. The explanatory power for exchange rates is typically very small.[15]

We further repeat the analysis for the VIX, the international VIX analogues (VSTOXX, VDAX, VFTSE, VCAC), and the commodity price factor (constructed as described in Supplementary Appendix S1).[16] To do so, we simply replace $q_{i,d}$ in equations (6) and (7) with the respective index or commodity price factor. Figure 6 shows the resulting daily, monthly, and quarterly R-squared. Similar to the estimates for stock indexes, the explanatory power increases at lower frequencies. At the quarterly frequency, US macroeconomic news explains typically between 15 and 25 percent of the variation in the implied volatility measures, as well as 25 percent in the commodity factor.

Lastly, we note that while incorporating non-headline news leads to a sizable increase in explanatory power, our estimates should be interpreted as conservative. The reason is that as in Gürkaynak, Kısacıkoğlu, and Wright (2020), we extract our non-headline news factor exclusively from the US yield curve.[17] International stock or bond market data likely captures additional information that could raise the explanatory power of non-headline news, but we

[15] Also note that we have sufficiently many observations for all news releases that overfitting concerns should not apply when estimating equation (6). Observation counts for all announcements are shown in Online Appendix Table B1. See also the out-of-sample check in Altavilla, Giannone, and Modugno (2017, pp. 40-41).

[16] To improve the sample coverage, we obtain daily data from Bloomberg for the VDAX, VFTSE, and VCAC.

[17] We only use US data in our estimation to keep our factors close to those extracted by Gürkaynak, Kısacıkoğlu, and Wright (2020) who provide extensive evidence that they are well identified. Also note that yields are preferred for the factor estimation (in comparison to stock returns), since the assumption of a time-invariant announcement effect, which is key for the identification of the factor, is more likely to hold for yields.

do not use this information here.

6 The Asymmetric Effects of US and Foreign Macro News

As we discussed in Section 2, shocks to global common state variables could principally drive the observed responses of foreign equity markets. If this was the case, US macro news releases would not impact foreign markets by transmitting US-specific shocks, but by revealing information about the global common state. In this section, we document that the effects of US and foreign macro news are highly asymmetric and provide an interpretation suggesting a limited role for common shocks.

6.1 Effects of Foreign Macro News on the US

We begin with analyzing the effects of foreign countries' macroeconomic news releases. The intuition of this exercise is as follows: Suppose that US macroeconomic surprises were driven by common global shocks. Then, our results in Section 4 and Section 5 would imply that a sizable fraction of the variation in global equity markets was driven by variation in common global state variables. Market participants should then seek other sources of information about the common global state, including foreign macroeconomic news releases, and hence global equity markets should also respond to other countries' macroeconomic releases.

A test for the presence of common shocks To test for the presence of common shocks, we study the effect of foreign news releases on the US stock market. In particular, we regress the log-change in the S&P 500 on foreign macroeconomic surprises. We show formally in Online Appendix A.3 that the estimated coefficient reflects the presence of common shocks: If countries' macroeconomic and financial variables were driven by common global state variables, other countries' macroeconomic releases should generally be informative about the common state vector. Further, under this premise, the S&P 500 and other international asset prices should respond to foreign macroeconomic surprises.

We add two remarks on the interpretation of the estimated coefficient. First, we study the effect of foreign news on the US rather than a third country, because the US is large. If the US was not large relative to the foreign country, the estimated effect could also reflect the transmission of foreign shocks to the US. For the interpretation of our estimates below, we must therefore keep in mind that, all else equal, smaller foreign countries offer a sharper test for the presence of common shocks.[18]

Second, our test for the presence of common shocks requires that macroeconomic series and their releases in foreign countries are similar to those in the US. Specifically, they (i) should

[18] As we discuss in Online Appendix A.3, the estimated coefficient could also reflect that market participants learn about the US state vector by observing macroeconomic news in country i. Since the US is large relative to country i, shocks in the US are likely to have an effect on country i's macroeconomic outcomes. As a result, country i's surprises could be informative about US-specific shocks. While this possibility cannot be ruled out *a priori*, we don't view it as particularly plausible either. Since US shocks presumably affect foreign macroeconomic outcomes with a lag and many indicators of US macroeconomic performance become available in a timely fashion, it is rather unlikely that this indirect channel of learning about the US state is active in practice. Further, if it was active, we would expect to find an effect of foreign news on US stock prices whereas our results below show that this is not the case.

be released in a similarly timely fashion, they (ii) need to be of comparably high measurement quality, and (iii) information leakages prior to the official release should be limited. If either of these criteria were violated, news about foreign macroeconomic aggregates would be of questionable use to learn about *any* state variable and asset prices should respond less strongly or not at all.

We therefore consider major macroeconomic news releases in the non-US G7 countries (i.e., Canada, France, Germany, Italy, Japan, and the United Kingdom). While differences to US macroeconomic news releases likely exist, these countries' news releases are *a priori* most likely to satisfy the above criteria (i) to (iii). Further, we perform several checks below, which suggest that neither timeliness, nor measurement quality, nor information leakages are major concerns for our analysis. Online Appendix Table B2 provides information on the foreign macroeconomic announcements. We consider 10 major releases per country.

We next estimate specifications analogous to equation (3), now with the 30-minute log-change in the S&P 500 on the left-hand side (as measured by the front-month E-mini S&P 500 futures contract) and the foreign macroeconomic surprise on the right-hand side. We control for other surprises released within the same time window, including releases of US news. As before, the surprises are standardized, so that the coefficients measure the effect size of a one standard deviation surprise.

Results The results in Table 6 reveal a striking asymmetry. Foreign news releases have essentially no effect on the US stock market. Out of 60 news releases, 8 have statistically significant effects on the S&P 500 at the 10 percent level—just 2 more than predicted by chance. Of these 8 significant effects 3 are for German macroeconomic news releases. Since Germany is closely integrated with the US and may not be small in comparison, it is also possible that these effects reflect the transmission of shocks—rather than the presence of common shocks.

In addition, the effect sizes are approximately an order of magnitude smaller than those of US news on foreign markets. The largest estimated effect in Table 6 suggests that a one standard deviation surprise in the advance release of GDP in the UK moves the S&P 500 by 4.42 basis points. This contrasts with a pooled effect of 17.60 basis points of the US advance GDP release on foreign countries (see Table 3) and almost 30 basis points for some countries as shown in Figure 3. Since the UK is a major financial center, it is again possible that this significant effect reflects the transmission of shocks—similar to the effects for Germany. Taken together, these results suggest a very limited role of global common shocks. They further highlight the unique position of the US economy in the global financial system: The effects of macroeconomic news are highly asymmetric.

Lastly, note that our findings above do not generally rule out the presence of common shocks as drivers of international financial and/or macroeconomic variables. Our findings only suggest that the effects of US news on foreign markets predominantly reflect US-specific shocks, rather than shocks common to all countries.

Table 6: Effects of Foreign News on US Stock Market

Canada	Capacity Utilization	Core CPI	GDP	Housing Starts	Intl. Trade	IPPI	Mfg Sales	PMI	Retail Sales	Unemployment Rate
S&P 500 (bp)										
News	0.82	1.69*	1.31	-1.96	0.64	1.19	-1.16	2.23	0.64	0.09
	(2.12)	(0.87)	(1.36)	(1.22)	(1.52)	(1.13)	(1.95)	(2.50)	(1.04)	(1.18)
Observations	78	220	81	230	259	253	264	193	263	264
Effect on Exchange Rate	No	Yes	Yes	Yes	Yes	No	Yes	Yes	Yes	Yes

France	BoF Industry Sentiment	Consumer Confidence	CPI P	GDP P	Industrial Production	Mfg Confidence	PPI	Production Outlook	Trade Balance	Unemployment Rate
S&P 500 (bp)										
News	2.75**	-0.06	0.57	-0.74	-0.93	-0.63	1.67	-0.12	0.18	0.28
	(1.20)	(0.70)	(0.58)	(1.57)	(1.12)	(0.87)	(1.34)	(0.97)	(0.80)	(0.81)
Observations	135	229	231	84	246	214	153	179	243	150
Effect on Exchange Rate	Yes	Yes	No	No	No	No	No	No	No	No

Germany	CPI P	GDP	GfK Consumer Confidence	IFO Business Climate	Industrial Production	PPI	Retail Sales	Trade Balance	Unemployment Change	ZEW Survey Expectations
S&P 500 (bp)										
News	-0.69	3.49**	0.69	0.98	2.38*	1.29	0.53	0.46	1.22	2.42***
	(1.69)	(1.49)	(0.90)	(1.44)	(1.29)	(0.88)	(0.75)	(0.85)	(1.11)	(0.87)
Observations	240	78	159	253	255	236	229	238	261	211
Effect on Exchange Rate	No	Yes	No	Yes	Yes	No	Yes	Yes	No	Yes

Continued on next page.

Italy	Consumer Confidence	CPI P	GDP F	Industrial Production	Industrial Sales	Mfg Confidence	PPI	Trade Balance	Retail Sales	Unemployment Rate
S&P 500 (bp)										
News	-0.38	-0.44	-0.91	0.78	4.24*	-0.70	-0.10	0.68	0.92	-0.47
	(1.02)	(0.70)	(1.55)	(0.89)	(2.37)	(1.22)	(1.52)	(1.51)	(0.82)	(0.99)
Observations	218	243	77	236	62	231	175	75	171	141
Effect on Exchange Rate	No	No	Yes	No	Yes	No	No	No	No	No

Japan	BoJ Mfg Index	BoJ Mfg Outlook	Consumer Confidence	CPI	Exports	GDP P	Industrial Production P	PPI	Retail Sales	Unemployment Rate
S&P 500 (bp)										
News	1.01	-3.51	-0.31	-0.22	-0.94	1.03	0.23	-0.90	0.34	0.20
	(1.12)	(3.06)	(0.49)	(0.36)	(1.11)	(1.54)	(0.44)	(0.76)	(0.65)	(0.42)
Observations	80	59	150	204	129	79	230	226	195	224
Effect on Exchange Rate	Yes	Yes	No	No	No	No	Yes	No	No	Yes

United Kingdom	Core CPI	Core PPI	Exports	GDP A	GfK Consumer Confidence	House Price Index	Industrial Production	Jobless Claims	Retail Sales	Unemployment Rate
S&P 500 (bp)										
News	0.94	-0.15	-0.18	4.42**	0.03	0.39	-0.27	0.48	1.78**	-1.18
	(0.99)	(1.00)	(1.34)	(1.75)	(0.54)	(0.67)	(0.91)	(0.70)	(0.74)	(0.90)
Observations	172	168	59	85	205	187	256	217	118	211
Effect on Exchange Rate	Yes	No	No	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Notes: The table presents the response of the S&P 500 to foreign macroeconomic news releases. For each non-US G7 country, this table shows estimates of ζ^y obtained from specification

$$\Delta q_{US,t} = \alpha_i + \zeta_i^y s_{i,t}^y + \sum_{k \neq y} \zeta_i^k s_{i,t}^k + \sum_{w} \zeta_{US}^w s_{US,t}^w + \varepsilon_{i,t},$$

where $s_{i,t}^y$ is the surprise of interest, $s_{i,t}^k$ and $s_{US,t}^w$ are other surprises of country i and the US released in the same time window, and $\Delta q_{US,t}$ is the 30-minute log-change of the front-month E-mini S&P 500 futures contract, expressed in basis points. *Effect on Exchange Rate* indicates whether the news release has a significant effect on the US dollar exchange rate at the 10 percent level. Online Appendix Table C3 shows the associated estimates. Online Appendix Table B2 provides details on the foreign news releases. Note that the observations reported in Online Appendix Table B2 can differ from those reported here, because the E-mini S&P 500 futures data is not always available. Heteroskedasticity-robust standard errors reported in parentheses. ***, **, and * indicate significance at the 1, 5, and 10 percent level.

6.2 Alternative Explanations for Asymmetry

We next check alternative explanations for the asymmetric effects of US and foreign macroeconomic news releases. As noted above, one may expect less timely news releases and releases with lower measurement quality to lead to smaller effect sizes (Gilbert et al., 2017). Further, information leakages could imply that measured surprises only contain outdated or irrelevant information, which should not move financial markets. Hence, if foreign surprises were less timely, of low measurement quality, or subject to leakages, their effects on the S&P 500 could be small despite the presence of common shocks.

Timeliness We first ask whether a lack of timeliness can explain the small effects of foreign news releases. To do so, we use the reporting lag of a macroeconomic series, a widely used proxy of timeliness (e.g., Fleming and Remolona, 1997). The smaller the reporting lag, the more timely is the release. More specifically, and following Gilbert et al. (2017), we define the reporting lag of a series as the difference between the announcement day and the last day of the reference period averaged over the sample.[19] Negative reporting lags exist for releases for which the reference period is in the future.[20] The data for this measure comes from Bloomberg, see Online Appendix B for details.

The left panel of Figure 7 plots estimated effect sizes (i.e., in absolute value) for the twelve US and 60 foreign releases against the measure of timeliness. The figure shows that most foreign news releases are approximately as timely as US releases and hence US releases are not special in terms of their timeliness. Further, while greater timeliness correlates positively with the price impact (or effect size) of news releases in the US as shown by other papers (e.g., Fleming and Remolona, 1997), timeliness cannot explain much of the differences in effect sizes between US and foreign news releases. The magnitudes of US releases are clear outliers. The fact that many foreign news releases in our sample are relatively timely is in line with Cascaldi-Garcia et al. (2021) who also show this for Germany, France, and Italy.

Measurement quality A second potential explanation of our findings is that US statistical authorities measure macroeconomic outcomes with greater precision than their foreign counterparts. To check this concern, we follow Gilbert (2011) and construct a proxy of measurement quality as the difference between the initial released value (used to construct the surprises) and its final revised value (a proxy for the true value of the macroeconomic series). A greater average revision magnitude suggests lower measurement quality of the initial re-

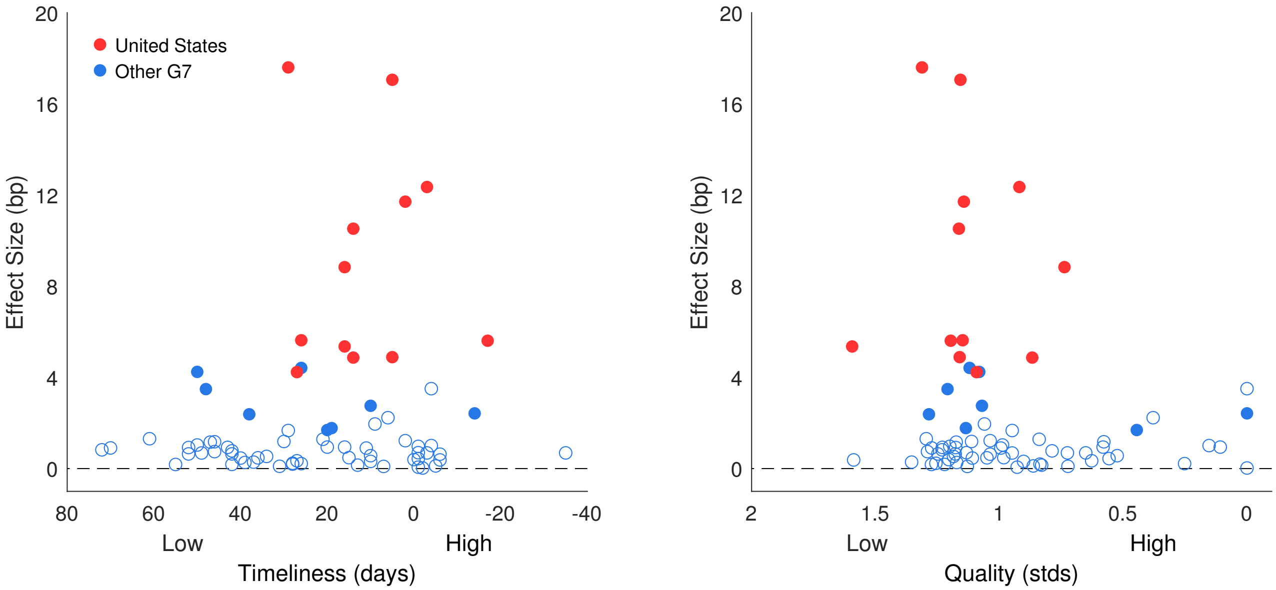
[19]Specifically, the reporting lag of series y in country i is

$$rl_i^y = \frac{1}{N_i^y} \sum_{n=1}^{N_i^y} \left(ann_{i,n}^y - ref_{i,n}^y \right), \tag{8}$$

where $ann_{i,n}^y$ and $ref_{i,n}^y$ refer to the announcement day and the last day of the reference period of the n^{th} release in our sample, and N_i^y denotes the total number of announcements for series y.

[20]Such negative reporting lags arise for several surveys. For instance, the preliminary release of the University of Michigan Consumer Sentiment Index has a negative average reporting lag (of 17 days), because the results are published before the end of the reference period.

Figure 7: Relation of Effect Size to Timeliness and Quality of Releases



Notes: This Figure shows how the effect size of a release relates to its timeliness (left panel), as well as its quality (right panel). Timeliness is measured by the reporting lag as defined in equation (8) and is in units of days. Quality is proxied by the revision magnitude as defined in equation (9) and is in units of standard deviations. For the US releases (red), the effect size corresponds to the absolute value of the coefficients shown in Table 3. For the foreign releases (blue), the coefficients in Table 6 are used. Filled circles correspond to effects which are significant at the 10 percent level.

lease.[21] As the right panel of Figure 7 shows, US news releases do not have a higher average quality than foreign news releases. Further, this measure of quality cannot explain much of the differences in effect sizes between US and foreign news releases. These findings are in line with Gilbert et al. (2017) who come to a similar conclusion for US releases.

Effects on domestic markets As a third check, we estimate the effects of foreign macroeconomic surprises on their respective domestic financial markets. Specifically, we study the effects on the local currencies' US-dollar denominated exchange rate.[22] A significant effect of a foreign macroeconomic news release on the local exchange rate implies that the news release in question contains market-relevant information and suggest that information leakages are not a major concern. Table 6 shows that out of the 60 foreign macroeconomic surprises under consideration 30 have a significant effect on the exchange rate at the 10 percent level. We report details on these estimates in Online Appendix Table C3.

[21] Following Gilbert (2011), we define the revision magnitude as the average absolute value of the difference between final revised and initial released number in the sample. To be precise, the revision magnitude of series y in country i is

$$rm_i^y = \frac{1}{N_i^y} \sum_{n=1}^{N_i^y} \frac{|y_{i,n}^F - y_{i,n}|}{\sigma_{|y_{i,n}^F - y_{i,n}|}}, \tag{9}$$

where $y_{i,n}$ and $y_{i,n}^F$ refer to the initial and final revised number of release n, $\sigma_{|y_{i,n}^F - y_{i,n}|}$ refers to the standard deviations of the absolute value of the difference. In Online Appendix Figure C6, we show that our results are robust to an alternative measure of revision magnitude.

[22] We perform this check on exchange rates rather than alternative financial instruments due to their extended trading hours, liquidity, and data availability.

Taken together, these checks suggest that concerns about measurement quality, timeliness, and information leakages do not explain the differences in the estimated effects documented above.

6.3 Transmission of US versus Foreign Monetary Policy Shocks

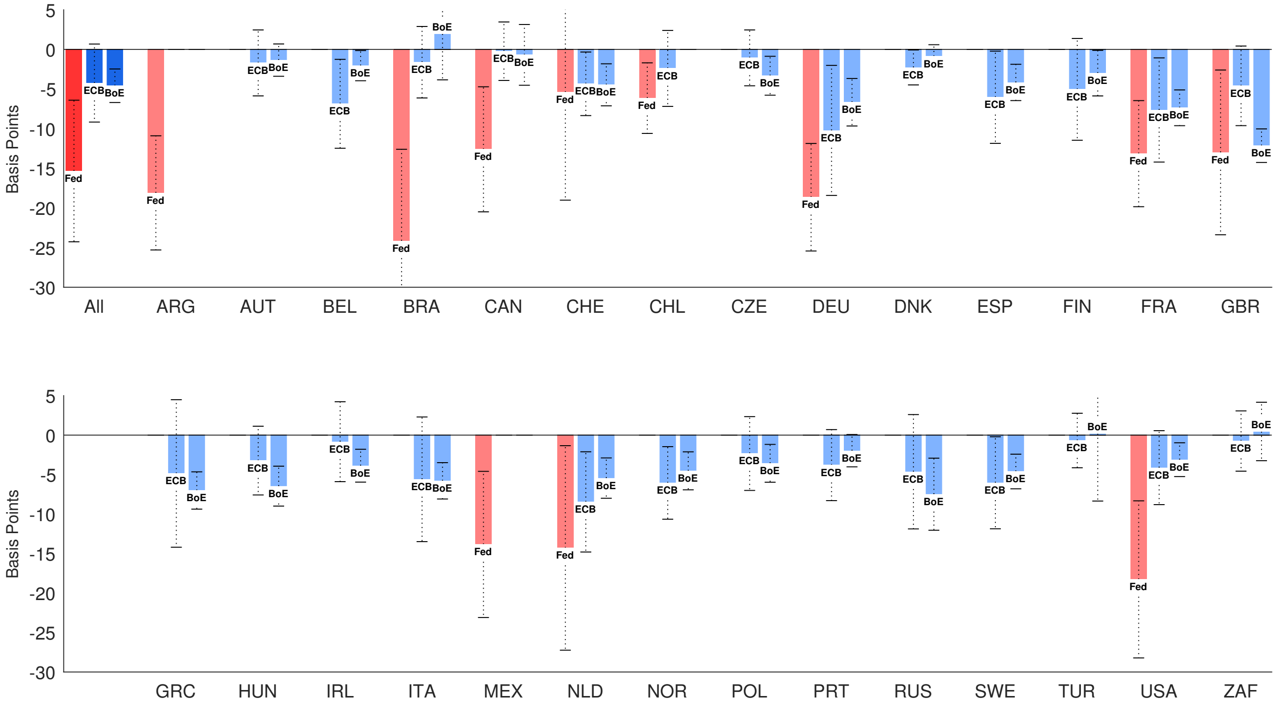
We next contrast the international transmission of monetary policy shocks of the Federal Reserve (Fed) with that of other central banks—where we focus on the European Central Bank (ECB) and the Bank of England (BoE). The rationale behind this exercise is that these shocks are well-identified, they are country-specific and therefore not contaminated by a global common component, and they contribute to business cycle fluctuations similar to other macroeconomic disturbances. There are also no concerns about differences in timeliness or measurement quality relative to the US. Hence, this exercise allows us to provide further evidence on the comparatively strong transmission of US-specific shocks. The evidence we present is based on an analogous set of event study regressions where we now use measures of central bank's policy surprises instead of the macroeconomic surprises as the right-hand-side variable of interest.

Critical for this exercise is the construction of comparable monetary policy shocks for the Fed, the ECB, and the BoE. To capture the different dimensions of monetary policy, we focus on three types of shocks: shocks to the *target rate*, *forward guidance* shocks, and *quantitative easing* shocks. The construction of the three shocks closely follows Swanson (2021) and is based on 30-minute changes in the yield curve around central bank announcements. Supplementary Appendix S3.1 provides details on the construction of each series. Since Fed announcements occur outside of the trading hours for many countries in our dataset, we switch out the stock indexes with the corresponding front-month futures contracts, which are traded outside of regular trading hours, where possible. This is the case for Brazil, Canada, Switzerland, Germany, France, United Kingdom, and the Netherlands.[23] We also include the S&P 500 so that we have domestic responses for all three central banks in the sample.

Figure 8 shows the pooled and country-specific effects of an increase in the target rate for each central bank. All shocks are measured in standard deviations of the respective series to ensure comparability of the magnitudes across central banks. Consistent with standard theory, the pooled effects are negative. An unexpected increase in the target rate leads to a decrease in stock markets for all three central banks. Quantitatively, however, the Fed's target rate shocks have an effect that is more than three times as large as the corresponding effects of the ECB and the BoE. Further, the pooled effects of the Fed's and the BoE's target rate shocks are significant at the 5 percent level, while the ECB's effect is estimated with less precision. The country-specific effects reported in the figure reveal that there is no instance in which for a given country the effect size of the ECB or the BoE exceeds that of the Fed. This point implies that the pooled effects shown first in the figure are not driven by the composition of countries. Overall, the results in Figure 8 show that US monetary policy shocks have a

[23] Note that stock index futures are available for more countries. However, only those we switched out are traded at Fed announcement times over a sufficiently long period of our sample.

Figure 8: Effects of Monetary Policy Shocks on International Stock Markets



Notes: The figure shows the effects of contractionary target rate shocks by the Federal Reserve (Fed), the European Central Bank (ECB), and the Bank of England (BoE) on international stock markets. The leftmost bars in the first row show the pooled effects for each central bank. The remaining bars represent the effects of a given central bank's shock on a given country's stock market. Missing country bars depict cases in which the country is dropped because it had less than 24 observations for a given shock. The coefficients are estimated analogously to equations (3) and (4). Stock index changes are expressed in basis points. The shocks are in standard deviations. The black error bands depict 95 percent confidence intervals, where standard errors are two-way clustered by announcement and by country. Analogous bar charts for forward guidance and quantitative easing shocks are shown in Supplementary Appendix Figure S3.4.

substantially larger impact on international stock markets, and hence are consistent with our previous interpretation that the outsized effect of US macro news is driven by the transmission of US-specific shocks as opposed to the presence of common shocks.

Supplementary Appendix Figure S3.4 presents the results for the forward guidance and quantitative easing shocks. The figure demonstrates that the transmission of unconventional monetary policy shocks is also substantially greater for the Fed than for the ECB and the BoE. The effects, however, are less precisely estimated. Further, we present several robustness checks in Supplementary Appendix S3.3. First, we show that the asymmetry documented above is robust to normalizing the shocks by their effects on the domestic yield curve as opposed to the standard deviation of surprises. Second, the results hold when purifying the shocks of information effects, i.e., the idea that a tightening might signal good news about the economy if the central bank in question has superior information. The results also indicate that information effects are potentially responsible for the noisy estimates in the case of unconventional monetary policy shocks. Lastly, our main findings are robust to using alternative shock series from the literature.

These results are broadly consistent with those of prior research. To our knowledge, the most closely related papers are Brusa, Savor, and Wilson (2020), Ca'Zorzi et al. (2020), and Miranda-Agrippino and Nenova (2022). Brusa, Savor, and Wilson (2020) find that the Fed has a uniquely strong impact on global equities compared to the BoE, the ECB, and the Bank of Japan. Ca'Zorzi et al. (2020) show that conventional policy shocks by the Fed have a greater impact on the Euro Area and the rest of the world than do shocks of the ECB. Lastly, Miranda-Agrippino and Nenova (2022) compare international spillovers of unconventional policy shocks by the Fed and ECB. While the transmission is qualitatively similar, it is substantially stronger for the Fed.

7 Transmission Channels of US Macro News

This section studies the channels underlying the foreign stock price reactions to US macro news in greater detail. To do so, we analyze the effects of news releases on a broader set of asset prices and draw on theory to interpret these findings.

7.1 Framework

To learn about the dominant transmission channels at work, we introduce a conceptual framework which allows us to summarize and distinguish theoretical channels in a concise fashion. This framework consists of two classic decompositions. The first one is the decomposition of the stock price into its three fundamental *components*: a risk-free interest rate, a risk premium, and a growth expectations component. Following Boyd, Hu, and Jagannathan (2005), we write

$$\Delta q_{i,t} \approx c_i \Big(\underbrace{\Delta g_{i,t}}_{\text{growth expectations}} - \underbrace{\Delta ep_{i,t}}_{\text{equity premium}} - \underbrace{\Delta r^f_{i,t}}_{\text{risk-free rate}} \Big), \tag{10}$$

where $\Delta q_{i,t}$ is the observed change in the stock price index, $\Delta g_{i,t}$ is the change in the weighted average of expected future growth rates of cash flows, $\Delta ep_{i,t}$ is the change of the equity (risk) premium, $\Delta r^f_{i,t}$ is the change in the interest rate on long-term risk-free claims, and c_i is a positive constant (the price-dividend ratio).

The second part of our framework is the decomposition of the long-term bond yield, $\Delta r_{i,t}$, into a risk-free rate component, $\Delta r^f_{i,t}$, and a term (risk) premium component, $\Delta tp_{i,t}$ (e.g., Singleton, 2006, p. 359):

$$\Delta r_{i,t} = \underbrace{\Delta r^f_{i,t}}_{\text{risk-free rate}} + \underbrace{\Delta tp_{i,t}}_{\text{term premium}} . \tag{11}$$

Note that $r^f_{i,t}$ captures the average of expected future short-term risk-free rates over the maturity of the claim, which relates $r^f_{i,t}$ directly to the expected future conduct of monetary policy.

Equipped with equations (10) and (11), we are now in a position to interpret the effects of US macro news releases as the result of one or multiple *channels* at work. To do so, we introduce the following four theoretical channels, which are inspired by Cieslak and Pang (2021): a

Table 7: Asset Price Predictions of Different Channels

Channel	Stock Return	Yield Change	Stock-Yield Co-movement
Growth Expectations	+ (Growth Exp. ↑)	+ (Risk-free Rate ↑)	+
Monetary Policy	+ (Risk-free Rate ↓)	− (Risk-free Rate ↓)	−
Common Premium	+ (Equity Premium ↓)	− (Term Premium ↓)	−
Hedging Premium	+ (Equity Premium ↓)	+ (Term Premium ↑)	+

Notes: This table summarizes the predictions of different channels following the categorization by Cieslak and Pang (2021). The terms in brackets describe the dominant force behind each prediction. All channels are signed to generate a positive stock return. See text for more details.

growth expectations channel, a monetary policy channel, a common premium channel, and a hedging premium channel.[24] These channels drive the joint behavior of stock prices and government bond yields. Note that they are not necessarily mutually exclusive, that is, multiple channels can be active simultaneously.[25] Importantly, the channels allow us to describe interpretable economic phenomena ("narratives") that are well understood from prior work and can easily be compared across studies. We next describe the key properties of each channel where we continue to closely follow Cieslak and Pang (2021). Table 7 provides an overview. Each channel is signed to generate a positive stock return.

First, the *growth expectations channel* increases stock prices ($q_{i,t} \uparrow$) through a rise in growth expectations ($g_{i,t} \uparrow$). Bond yields ($r_{i,t} \uparrow$) increase as well due to an expected increase in future policy rates ($r^f_{i,t} \uparrow$). For stock prices, however, the risk-free rate component is not dominant as it is assumed to respond less than one-for-one with growth expectations (see equation (10)). This is consistent with existing empirical evidence (e.g., Coibion and Gorodnichenko, 2012).

Second, the *monetary policy channel* leads to an increase in stock prices ($q_{i,t} \uparrow$) through a decrease in the risk-free rate ($r^f_{i,t} \downarrow$), which also leads to a decrease in bond yields ($r_{i,t} \downarrow$). These predictions for stock prices and bond yields are documented in a wide range of empirical papers (e.g., Rigobon and Sack, 2004). We defer a more detailed discussion of what this channel may and may not capture to Section 7.2 below.

Lastly, we consider two types of risk premium channels. These are motivated by the idea that changes in risk premia can lead to both positive or negative co-movement in bond and stock markets (e.g., Bansal and Shaliastovich, 2013; Campbell, Pflueger, and Viceira, 2020). The *common premium channel* pushes the equity and term premium in the same direction. It increases stock prices ($q_{i,t} \uparrow$) through a decrease in the equity premium ($ep_{i,t} \downarrow$). In this case, it would also decrease bond yields ($r_{i,t} \downarrow$) through its effect on the term premium ($tp_{i,t} \downarrow$).

[24] A slight difference relative to Cieslak and Pang (2021) is that they refer to these four categories as "shocks" in a VAR, while we refer to them as "channels" that are potentially active after the release of US macro news.

[25] It is straightforward to verify that in a vector space with the dimensions (i) growth expectations, (ii) risk-free rate, (iii) equity premium, and (iv) term premium, these four channels constitute a basis. This implies that any change in stock returns and bond yields can be expressed as a linear combination of these four channels (or basis vectors). In this sense the four channels are collectively exhaustive.

In contrast, the *hedging premium channel* generates a movement of the equity and term premium in opposite directions. This channel increases stock prices ($q_{i,t} \uparrow$) through a decrease in the equity premium ($ep_{i,t} \downarrow$). At the same time, bond yields would increase ($r_{i,t} \uparrow$) due to the rise in the term premium ($tp_{i,t} \uparrow$). In this example, investors hedge less and take on more risk.[26] The opposite case of more hedging and less risk-taking can be thought of as a "flight-to-safety" effect (see, e.g., Baele et al., 2020).

The objective for the remainder of this section is to identify the dominant channel(s) from the co-movements of asset prices around US macro news releases. To do so, we start in Section 7.2 by studying the stock-bond co-movement. As shown in the rightmost column of Table 7, the co-movement of stock returns and bond yields allows us to rule out that two channels are dominant. Subsequently, we study in Section 7.3 the effect of US macro news on the individual components in equations (10) and (11). This analysis allows us to further tighten the interpretation. As some of these components are not perfectly observable at high frequencies, we proxy for them with several daily measures proposed in the literature.

7.2 Stock-Bond Co-Movement

7.2.1 International Markets

We now estimate the effects of the US macro releases on government bond yields, starting with international markets. To do so, we re-estimate equation (3) after replacing the dependent variable with the 30-minute change of country i's 10-year government bond yield in local currency. We focus on 10-year government bonds compared to bonds of other maturities because they provide a standard measure of long-term interest rates and are available for all countries in our sample.[27] We exclude bond market data during sovereign debt crises in Argentina and Greece. Table 8 reports the results. For convenience, the table also reports the previously obtained estimates for stock indexes from Table 3.

As Table 8 shows, foreign bond yields increase significantly after all 12 releases. For instance, a positive one standard deviation surprise in Nonfarm Payrolls raises foreign long-term interest rates, on average, by 1.69 basis points. Importantly, for all 10 releases about US real activity, a positive surprise raises international stock prices *despite* the increase in international long-term bond yields. The positive co-movement of stock returns and bond yields for real activity news implies that neither the monetary policy nor the common premium channel is dominant (see Table 7). In contrast, positive inflation surprises (Core CPI and Core PPI) raise long-term bond yields while lowering international stock prices. This suggests that the monetary policy or the common premium channel is dominant for inflation surprises. In Online Appendix C4, we show analogous results for the 1-year bond yield. For these bonds a

[26] Note that we avoid the term "risk-taking channel" throughout. Although we believe that this term accurately describes the phenomenon at work, it is often used to describe increased risk-taking following expansionary monetary policy shocks. In our context, this channel can be active independently of monetary policy.

[27] We rely on yields calculated by *Thomson Reuters*, which are based on bond prices from "external" sources. This ensures consistency in the yield calculations across countries. The corresponding identifiers are ending with $=RR$, e.g., $AR10YT = RR$ for the Argentinian 10-year government bond yield. Online Appendix Table B3 provides an overview of the employed instruments.

Table 8: Effects on International Stock and Bond Markets

	Capacity Utilization	CB Consumer Confidence	Core CPI	Core PPI	Durable Goods Orders	GDP A
Stock Index (bp)						
News	5.36**	12.35***	-8.84***	-4.87***	5.63***	17.60***
	(2.28)	(2.02)	(1.89)	(1.29)	(1.60)	(3.36)
R^2	0.04	0.13	0.10	0.15	0.10	0.26
Observations	6054	6041	5717	5828	5610	1911
10-Year Bond Yield (bp)						
News	0.21***	0.53***	0.61***	0.43***	0.29***	0.87***
	(0.06)	(0.08)	(0.11)	(0.07)	(0.10)	(0.16)
R^2	0.02	0.10	0.05	0.10	0.04	0.19
Observations	4540	4308	4456	4561	4373	1420
	Initial Jobless Claims $\cdot(-1)$	ISM Mfg Index	New Home Sales	Nonfarm Payrolls	Retail Sales	UM Consumer Sentiment P
Stock Index (bp)						
News	4.89***	11.71***	4.23***	17.06***	10.52***	5.61***
	(0.73)	(2.24)	(1.40)	(2.99)	(1.68)	(1.54)
R^2	0.09	0.12	0.03	0.13	0.15	0.04
Observations	24334	5548	5908	5688	5786	5726
10-Year Bond Yield (bp)						
News	0.28***	0.89***	0.28***	1.69***	0.47***	0.28***
	(0.04)	(0.09)	(0.06)	(0.20)	(0.08)	(0.06)
R^2	0.02	0.17	0.03	0.23	0.15	0.03
Observations	19228	4069	4232	4491	4525	4101

Notes: The table presents results from the pooled regression for stock indexes, as shown in Table 3, as well as the effects on 10-year government bond yields, obtained from estimating γ^y of equation (3) after replacing the left-hand variable with the 30-minute change of country i's 10-year government bond yield in local currency. The units are in basis points. Standard errors are two-way clustered by announcement and by country, and reported in parentheses. ***, **, and * indicate significance at the 1, 5, and 10 percent level.

qualitatively similar picture emerges, although the estimates are generally less precise because the 1-yield is not available for all countries in our sample.

With these results in hand, it is important to understand whether real activity or price news drives our findings in the previous sections. As we show in Online Appendix Figure C5, price news explains only a relatively small fraction of the quarterly variation in foreign stock prices. To obtain these results, we re-run the explanatory power exercise from Section 5 separately for price and real activity news.[28] It turns out that more than 80 percent of the quarterly explanatory power of foreign stock prices comes from news about US real activity. Taken together, the evidence from international bond markets thus indicates that the growth expectations or the hedging premium channel are dominant for our findings.

[28] For a classification of all news releases into the real activity and price category, see Online Appendix Table B1.

7.2.2 US Markets

We next repeat this exercise for US markets. Specifically, we estimate specification (5), where we now use the S&P 500 and the 10-year Treasury yield on the left-hand-side. Table 9 shows the results. The estimates have the same signs as those for international markets—although the effects on Treasury yields are quantitatively larger. As for international markets, stock prices and bond yields positively co-move following news about US real activity, implying a dominant growth expectations or hedging premium channel. For price news, this co-movement is negative, implying a dominant monetary policy or common premium channel. As real activity news captures the large majority of all variation, the evidence here again suggests a dominant growth expectations or hedging premium channel. As an extension, we present results for yields of different maturities in Online Appendix Table C5. Consistent with the evidence from prior work (e.g., Gürkaynak, Kısacıkoğlu, and Wright, 2020), the entire US yield curve shifts in the same direction as the 10-year Treasury yield.

7.2.3 Discussion: The role of US monetary policy

We next discuss the role of US monetary policy reactions for driving our results. First, it is important to emphasize that ruling out a dominant monetary policy channel does *not* imply that monetary policy is passive or unimportant. Our results show that markets indeed perceive a strong endogenous monetary policy response (see Table 9 and Online Appendix Table C5). Below we further verify that the yield responses are predominantly driven by the risk-free rate component (as opposed to the term premium, see Section 7.3). However, this risk-free rate component is not the dominant force for stock prices and hence we can rule out that the monetary policy channel, as defined above, is dominant.

The second remark concerns the definition of the monetary policy channel itself. While we assumed in Section 7.1 that monetary policy affects stock prices mostly through its effect on the risk-free rate, this assumption is not critical for ruling out a dominant monetary policy channel based on the stock-bond co-movement. Since a monetary-policy-induced increase in the risk-free rate leads to an increase in the equity premium and diminished growth expectations (Bernanke and Kuttner, 2005), stock prices will fall while bond yields rise. Following US real activity news, however, this co-movement is positive implying that an expected monetary policy reaction cannot explain the effects on stock markets.

One way to rationalize the results with a dominant monetary policy channel is that, for example, the expected endogenous interest rate increases following US macro news are associated with increases in stock prices. Such a mechanism is also referred to as the signaling channel of monetary policy or as interest rate movements triggering information effects. While the importance of this signalling/information effects channel is still debated, several papers find that it is, *on average*, not dominating the effect on stock markets.[29] Hence, in our assessment, it does not appear that the monetary policy channel is the dominant driving force of

[29] Specifically, without separating "pure" monetary policy shocks and information shocks, interest rates and stock prices move, on average, in opposite directions (see, e.g., Figure 2 Panel B in Jarociński and Karadi (2020) and Table 5 in Nakamura and Steinsson (2018)). See also Cieslak and Schrimpf (2019) Table 8 for a variance decomposition of yield changes and stock returns around Fed communication events.

Table 9: Effects of US News on US Stock and Bond Markets

	Capacity Utilization	CB Consumer Confidence	Core CPI	Core PPI	Durable Goods Orders	GDP A
S&P 500 (bp)						
News	4.98*	15.93***	-14.07***	-5.83***	4.57**	17.69***
	(2.92)	(2.64)	(2.64)	(1.69)	(1.90)	(3.89)
R^2	0.05	0.20	0.23	0.21	0.24	0.43
Observations	244	265	258	264	265	87
10-Year Treasury Yield (bp)						
News	0.45***	1.15***	1.31***	0.98***	0.44*	1.56***
	(0.10)	(0.17)	(0.23)	(0.15)	(0.26)	(0.34)
R^2	0.09	0.37	0.22	0.36	0.25	0.30
Observations	270	195	264	274	187	90
	Initial Jobless Claims $\cdot(-1)$	ISM Mfg Index	New Home Sales	Nonfarm Payrolls	Retail Sales	UM Consumer Sentiment P
S&P 500 (bp)						
News	5.33***	14.84***	3.31	18.62***	9.09***	3.85*
	(0.76)	(3.39)	(2.05)	(3.76)	(2.05)	(2.28)
R^2	0.22	0.18	0.06	0.15	0.21	0.02
Observations	1143	265	264	266	259	216
10-Year Treasury Yield (bp)						
News	0.59***	2.14***	0.73***	4.18***	1.46***	0.60***
	(0.07)	(0.18)	(0.13)	(0.42)	(0.21)	(0.12)
R^2	0.22	0.47	0.27	0.46	0.37	0.13
Observations	1025	273	190	274	271	243

Notes: The table presents regression results for the S&P 500 and 10-year Treasury yields, obtained from estimating γ^y of equation (5) after replacing the left-hand-side variable with the 30-minute log-change in the front-month E-mini S&P 500 futures contract or the change in the yield based on 10-year Treasury futures contracts. The units are in basis points. Heteroskedasticity-robust standard errors are reported in parentheses. ***, **, and * indicate significance at the 1, 5, and 10 percent level.

stock prices following US real activity news.

7.3 Effects on the Components of Stock Prices and Bond Yields

So far, we have not been able to estimate the effects on equity and term premia, even though they are crucial for obtaining a better understanding of the underlying mechanisms. The reason for this is that, to our knowledge, measures of these premia only exist at the daily frequency and not intraday. Further, it is well understood that estimating the effects on daily—as opposed to intraday—data adds a substantial amount of noise so that surprises of a single announcement series typically lack the statistical power necessary to obtain informative estimates.

In this section, we overcome this issue by creating a combined real activity news and a combined price news series using the aggregation method by McCoy et al. (2020). As we have shown, positive news about US real activity tends to raise stock prices while lowering bond yields, and higher-than-expected inflation tends to lower stock prices while raising bond yields. This homogeneity within each release category allows us to retain key information on the dominant channels when bundling individual releases into combined daily series.

Specifically, the daily series of real activity news is constructed as the weighted average of real activity surprises:

$$s_d^{RA} = \sum_{k=1}^{K_d^{RA}} w_d^k \, s_{US,d}^k, \tag{12}$$

where K_d^{RA} is the number of all real activity releases on day d (as listed in Online Appendix Table B1), and w_d^k denotes the relative weight of series k. This weight is obtained by dividing the Bloomberg relevance value W_d^k of series k by the sum of all relevance values, where the sum is taken over all series within the real activity category on day d, i.e., $w_d^k = W_d^k / \sum_{k=1}^{K_d^{RA}} W_d^k$.[30,31] The daily series of price news s_d^P is constructed analogously.

With both daily series in hand, we estimate the following regression for various asset prices:

$$\Delta q_{US,d} = \alpha + \beta s_d^{RA} + \gamma s_d^P + \delta \Delta q_{US,d-1} + \varepsilon_{US,d}, \tag{13}$$

where $\Delta q_{US,d}$ is the daily return of asset price q. As in Section 5, the sample runs from January 1, 2000 to December 31, 2019. We begin by confirming that this daily regression produces estimates consistent with those of our earlier intraday analysis. To do so, we study the effects on the S&P 500, the VIX, and the 10-year Treasury yield. The first three columns of Panel A of Table 10 show that this is indeed the case. After positive real activity news, the S&P 500 rises, the VIX falls, and the 10-year Treasury yield rises as well. The effects of price news are also consistent with our earlier estimates although they are estimated with less precision. This difference in statistical power is expected as the index of price news is constructed from only 15 series while the index on real activity news is constructed from 51 series.

As noted above, an advantage of moving to the daily frequency is that better measures of the four components determining stock prices and bond yields (see equations (10) and (11)) are then available. In the fourth and fifth column of Panel A of Table 10, we decompose the 10-year Treasury yield into an average of expected future short rates and the term premium using the series from Adrian, Crump, and Moench (2013).[32] The estimates make clear that the 10-year Treasury yield rises after positive real activity and price news, in part because of an expected increase in future short-term rates and in part because of an increase in the term premium.

We next turn to the effects on the equity premium and expected future dividends. To do so, we use Martin's (2017) measure of the 1-year equity premium, as well as the 1-year expected dividend, which we construct from S&P 500 dividend futures contracts (following

[30] The Bloomberg relevance index measures the relative popularity of a given announcement series within Bloomberg. More specifically, it measures how many Bloomberg users set an alert for a given announcement series relative to all alerts set for a given country.

[31] Note that we flip the sign for the Unemployment Rate, Initial Jobless Claims, Continuing Claims, Government Budget Balance, and Current Account Balance, such that a positive surprise corresponds to greater-than-expected real activity for all announcements.

[32] Details on all daily series we use in this section are provided in Online Appendix B.3.

Table 10: Effects of US News on Daily Returns

	S&P 500	VIX	10-Year Treasury Yield	10-Year Risk-Free Rate	10-Year Term Premium
Panel A: Daily Return (bp)					
Real Activity News	7.80***	-29.37***	1.07***	0.61***	0.47***
	(2.23)	(10.84)	(0.10)	(0.06)	(0.08)
Price News	-0.63	5.36	0.53***	0.24***	0.28**
	(3.14)	(15.70)	(0.15)	(0.09)	(0.14)
R^2	0.01	0.01	0.03	0.03	0.01
Observations	4976	4976	4918	4918	4918
	1-Year Equity Premium	**1-Year Growth Exp.**	**1-Year Treasury Yield**	**1-Year Risk-Free Rate**	**1-Year Term Premium**
Panel B: Daily Return (bp)					
Real Activity News	-1.09**	2.12***	0.78***	0.54***	0.24***
	(0.48)	(0.70)	(0.07)	(0.07)	(0.03)
Price News	-0.53	0.37	0.27***	0.19**	0.08
	(0.71)	(1.11)	(0.09)	(0.09)	(0.05)
Panel C: Stock Price Elasticity					
	-0.95	0.02		-0.95	
Panel D: Daily Return $\times$ Stock Price Elasticity (bp)					
Real Activity News	1.08**	0.04***		-0.52***	
	(0.44)	(0.01)		(0.07)	
Price News	0.50	0.01		-0.18**	
	(0.64)	(0.02)		(0.08)	
R^2	0.02	0.02	0.03	0.02	0.02
Observations	3635	1076	4918	4918	4918

Notes: The table shows the effects of the real activity news index as defined in equation (12) and the price news index, defined analogously, on various asset prices. Panel A and B show estimates of β and γ of equation (13) for different dependent variables. Log-changes are used for the S&P 500 and the VIX while changes in levels are used for the other asset prices. All units are in basis points. See text and Online Appendix B.3 for details. Panel C reports the average stock price elasticity (or semi-elasticity for a change in levels). These are constructed as in Knox and Vissing-Jorgensen (2022). See also Online Appendix B.3 for details. Panel D shows the contributions of the 1-year equity premium, the 1-year growth expectations, and the 1-year risk-free rate to the effect on the S&P 500 (as shown in Panel A). Heteroskedasticity-robust standard errors are reported in parentheses. ***, **, and * indicate significance at the 1, 5, and 10 percent level.

Gormsen and Koijen, 2020).[33] To obtain a complete picture at the 1-year horizon, we further study the 1-year Treasury yield and its decomposition into risk-free rate and term premium. Panel B of Table 10 shows the results for all measures at the 1-year horizon.

Positive real activity news decreases the 1-year equity premium while raising the term premium. Hence, better-than-expected real activity news appears to lower preferences to hedge risk and to increase preferences to take risk. This result implies that the hedging premium channel is potentially the key mechanism behind our results. At the same time, a

[33] More precisely, the measure of the equity premium is Martin's (2017) lower bound. However, Martin (2017) shows that this bound is quite tight.

positive surprise about US real activity also increases the 1-year expected dividend, which is consistent with a dominant growth expectations channel. Note that price news has no significant effects on the equity premium. Since the decomposition in Panel B of Table 10 is qualitative as opposed to quantitative, it is not possible to infer without additional information whether the hedging premium channel or the growth expectations channel is ultimately the dominant driving force of our observed stock price responses.

To make some progress, we calculate elasticities of the stock price with respect to expected dividends, the risk-free rate, and the equity premium.[34] To do so, we use the method of Knox and Vissing-Jorgensen (2022). Compared to earlier approaches, their method focuses on observables and requires less structure. While their decomposition method still requires some assumptions and is incomplete in our application in the sense that there is a large unexplained residual, it allows us to compare the magnitudes of the partial effects reported in Panel B of Table 10 to one another. The elasticities are reported in Panel C of Table 10.[35] Multiplying these elasticities with the estimated partial effects from Panel B delivers the partial contributions of the 1-year equity premium, the 1-year dividend, and the 1-year risk-free rate to the overall S&P 500 response of 7.80 basis points reported in Panel A. These products are reported in Panel D.

According to this "1-year decomposition", the equity premium has the largest effect on the S&P 500, accounting for approximately one basis point. The 1-year expected dividend contributes positively, but the effect is quantitatively small. Further, the 1-year risk-free rate contributes negatively. The intuition for why the effect of the equity premium is greater than that for dividends is the following: Changes in the 1-year equity premium affect the discounting at all maturities above one year as future dividends are discounted by the cumulative discount rate. On the other hand, the change in the expected dividend for a given year does not mechanically affect expected dividends of other years.

In conclusion, while data limitations prevent us from carrying out a complete decomposition, the available evidence suggests that the hedging premium channel is most important in our context and potentially the dominant driving force of the observed effects. We cannot fully rule out, however, that the growth expectations channel is important as well.

8 Extensions and Robustness

In this section, we briefly discuss several extensions and robustness checks.

US dollar exchange rates The US dollar exchange rate is a key variable in international finance and a potential amplification mechanism of cross-border financial spillovers as shown by Bruno and Shin (2015b). We therefore investigate in Supplementary Appendix S5 the role of the US dollar exchange rate in the transmission of US macro news. To do so, we re-estimate our pooled specification (3) after replacing the dependent variable with the US dollar denominated local exchange rate. For inflation news, the dollar appreciates while for

[34] To be precise, semi-elasticities are constructed for variables which are not in log-changes.

[35] For details on the construction of elasticities, see Online Appendix B.3.

real activity news the dollar either appreciates or is not affected significantly. For real activity news, which captures the large majority of variation, the co-movement with the stock response is inconsistent with a dollar-based mechanism proposed by Bruno and Shin (2015b). These findings therefore suggest that the exchange rate response is not central for understanding the direct effect of US macro news on stock prices. Of course, this does not conflict with the view that the US dollar is central for understanding the global financial system and that the dollar may also be important for understanding the asymmetry documented above.

State-dependent effects Prior work has shown that the effects of news on equity prices vary over the business cycle (e.g., McQueen and Roley, 1993; Boyd, Hu, and Jagannathan, 2005; Andersen et al., 2007; Goldberg and Grisse, 2013; Gürkaynak, Kısacıkoğlu, and Wright, 2020; Gardner, Scotti, and Vega, 2022; Elenev et al., 2022, among others). This raises the question whether our estimates are driven by very large effects in extreme phases of the business cycle and are otherwise absent. In Supplementary Appendix S4, we extend our analysis and allow for time-varying effects. We estimate a specification in which the effect of US news on foreign equity prices can vary with the level of several US and foreign variables. Consistent with prior work, we find that the effect size increases during bad times. In particular, in episodes of high US unemployment and in periods in which the US economy is perceived as doing poorly, as measured by Gardner, Scotti, and Vega's (2022) FOMC Sentiment Index, the effects tend to be largest. Our results also show that the effect size varies more with the state of the US economy than with the state of the foreign economy. However, and most importantly in the context of our analysis, the appendix shows that the effects reported in Table 3 are present in normal times and not driven by large effects in the extreme episodes of our sample period.

Cross-sectional heterogeneity Recall that Figure 4 showed that some countries' stock markets, including Germany's, France's, Italy's, and the Netherlands', respond systematically more strongly to US macroeconomic news than stock markets in Austria, Denmark or Portugal. In Supplementary Appendix S6, we return to this heterogeneity and ask whether it correlates with observables. Perhaps surprisingly, we find no robust correlation of the effect size with (i) a measure of financial integration, (ii) a measure of trade integration, (iii) a measure of industry dissimilarity, or (iv) an exposure measure to dollar valuation effects—once we control for other determinants of the effect size such as the state-dependent effects discussed above. While this evidence does not rule out the existence of any of the mechanisms (i)-(iv), it does suggest that they are not sufficiently salient to be statistically detectable in our sample. In our view, understanding the heterogeneity in effect size across countries is an interesting topic for future research.

9 Conclusion

Prior work has convincingly established that capital flows, risky asset prices, credit growth, and leverage co-move globally. Since much of the evidence in the literature is based on correlations, however, the interpretation of this co-movement is often not clear. Bernanke (2017), for instance, questions that the US economy is an important source of the disturbances driving the global financial cycle.

In this paper, we contribute to our understanding of the global financial cycle by establishing a causal link between the US economy and a large set of global risky asset prices. US macroeconomic news has strong and synchronous effects on foreign stock markets, the VIX and other implied volatility measures, as well as commodity prices. It also explains a sizable fraction of their variation. Since the co-movement of these risky asset prices is a defining feature of the global financial cycle, we interpret our findings as evidence that shocks driving the US business cycle also drive the global financial cycle.

We also document a striking asymmetry between the effects of US macro news and foreign macro news. While US macro news has large effects on foreign stock markets, foreign macro news has essentially no effect on the US stock market. This finding highlights the US' central position in the global financial system, and suggests a limited role for global common shocks. Consequently, and providing a partial answer to Bernanke's (2017) conjecture mentioned above, our evidence *does* indicate that US-specific shocks drive international financial conditions.

Our results are consistent with and complementary to those in Miranda-Agrippino and Rey (2020). This suggests that the common elements across findings may help guide future modeling efforts. In our assessment, the most salient of these are the following. First, both papers identify drivers of the global financial cycle and the origin of the shock is the US. Hence, features of the US economy—whether size or other—are likely central to understanding the driving forces of the global financial cycle. Second, in both cases the effects of the respective shocks on risk-taking is the key driving force of international risky asset prices. The evidence therefore points to a class of models that can generate time variation in measured global risk-premia.

Lastly, a central question arises from our and prior work on the global financial cycle: Is the size of the US sufficient or are other features necessary to explain the US' role for the global financial cycle? Since economic size and, for example, the special role of the US dollar are likely interdependent and not easily separable from other characteristics, this question is empirically difficult to answer. Our evidence only provides a loose indication: ECB policy shocks tend to have smaller effects on international equity prices than monetary policy shocks of the Federal Reserve, even though the size of the Euro Area is comparable to the US according to some measures. This may suggest that other features specific to the US determine its importance for the global financial cycle. It is clear, however, that more research is needed to answer this question satisfactorily.

References

Acalin, Julien and Alessandro Rebucci. 2020. "Global Business and Financial Cycles: A Tale of Two Capital Account Regimes." Working Paper 27739, National Bureau of Economic Research.

Adrian, Tobias, Richard K Crump, and Emanuel Moench. 2013. "Pricing the term structure with linear regressions." *Journal of Financial Economics* 110 (1):110–138.

Altavilla, Carlo, Domenico Giannone, and Michele Modugno. 2017. "Low frequency effects of macroeconomic news on government bond yields." *Journal of Monetary Economics* 92:31 – 46.

Andersen, Torben G., Tim Bollerslev, Francis X. Diebold, and Clara Vega. 2007. "Real-time price discovery in global stock, bond and foreign exchange markets." *Journal of International Economics* 73 (2):251 – 277.

Backus, David K, Patrick J Kehoe, and Finn E Kydland. 1992. "International real business cycles." *Journal of political Economy* 100 (4):745–775.

Baele, Lieven, Geert Bekaert, Koen Inghelbrecht, and Min Wei. 2020. "Flights to safety." *The Review of Financial Studies* 33 (2):689–746.

Balduzzi, Pierluigi, Edwin J Elton, and T Clifton Green. 2001. "Economic news and bond prices: Evidence from the US Treasury market." *Journal of financial and Quantitative analysis* 36 (4):523–543.

Bansal, Ravi and Ivan Shaliastovich. 2013. "A long-run risks explanation of predictability puzzles in bond and currency markets." *The Review of Financial Studies* 26 (1):1–33.

Beechey, Meredith J and Jonathan H Wright. 2009. "The high-frequency impact of news on long-term yields and forward rates: Is it real?" *Journal of Monetary Economics* 56 (4):535–544.

Bekaert, Geert, Marie Hoerova, and Nancy R Xu. 2020. "Risk, Uncertainty and Monetary Policy in a Global World." *Available at SSRN 3599583* .

Bernanke, Ben S. 2017. "Federal reserve policy in an international context." *IMF Economic Review* 65 (1):1–32.

Bernanke, Ben S and Kenneth N Kuttner. 2005. "What explains the stock market's reaction to Federal Reserve policy?" *The Journal of finance* 60 (3):1221–1257.

Boyd, John H, Jian Hu, and Ravi Jagannathan. 2005. "The stock market's reaction to unemployment news: Why bad news is usually good for stocks." *The Journal of Finance* 60 (2):649–672.

Bruno, Valentina and Hyun Song Shin. 2015a. "Capital flows and the risk-taking channel of monetary policy." *Journal of Monetary Economics* 71:119–132.

———. 2015b. "Cross-border banking and global liquidity." *The Review of Economic Studies* 82 (2):535–564.

Brusa, Francesca, Pavel Savor, and Mungo Wilson. 2020. "One central bank to rule them all." *Review of Finance* 24 (2):263–304.

Calvo, Guillermo A, Leonardo Leiderman, and Carmen M Reinhart. 1993. "Capital inflows and real exchange rate appreciation in Latin America: the role of external factors." *Staff Papers* 40 (1):108–151.

———. 1996. "Inflows of Capital to Developing Countries in the 1990s." *Journal of economic perspectives* 10 (2):123–139.

Campbell, John Y, Carolin Pflueger, and Luis M Viceira. 2020. "Macroeconomic drivers of bond and equity risks." *Journal of Political Economy* 128 (8):3148–3185.

Canova, Fabio. 2005. "The transmission of US shocks to Latin America." *Journal of Applied Econometrics* 20 (2):229–251.

Canova, Fabio and Jane Marrinan. 1998. "Sources and propagation of international output cycles: common shocks or transmission?" *Journal of International Economics* 46 (1):133–166.

Cascaldi-Garcia, Danilo, Thiago RT Ferreira, Domenico Giannone, and Michele Modugno. 2021. "Back to the Present: Learning about the Euro Area through a Now-casting Model." *International Finance Discussion Paper* (1313).

Ca'Zorzi, Michele, Luca Dedola, Georgios Georgiadis, Marek Jarocinski, Livio Stracca, and Georg Strasser. 2020. "Monetary policy and its transmission in a globalised world." .

Cerutti, Eugenio, Stijn Claessens, and Andrew K Rose. 2019. "How important is the global financial cycle? Evidence from capital flows." *IMF Economic Review* 67 (1):24–60.

Chari, Anusha, Karlye Dilts Stedman, and Kristin Forbes. 2022. "Spillovers at the extremes: The macroprudential stance and vulnerability to the global financial cycle." Tech. rep., National Bureau of Economic Research.

Cieslak, Anna, Adair Morse, and Annette Vissing-Jorgensen. 2019. "Stock returns over the FOMC cycle." *The Journal of Finance* 74 (5):2201–2248.

Cieslak, Anna and Hao Pang. 2021. "Common shocks in stocks and bonds." *Journal of Financial Economics* 142 (2):880–904.

Cieslak, Anna and Andreas Schrimpf. 2019. "Non-monetary news in central bank communication." *Journal of International Economics* 118:293–315.

Coibion, Olivier and Yuriy Gorodnichenko. 2012. "Why are target interest rate changes so persistent?" *American Economic Journal: Macroeconomics* 4 (4):126–62.

Davis, J Scott and Eric Van Wincoop. 2021. "A Theory of the Global Financial Cycle." Tech. rep., National Bureau of Economic Research.

Di Giovanni, Julian, Şebnem Kalemli-Özcan, Mehmet Fatih Ulu, and Yusuf Soner Baskaya. 2022. "International spillovers and local credit cycles." *The Review of Economic Studies* 89 (2):733–773.

Diaz-Alejandro, Carlos. 1983. "Stories of the 1930s for the 1980s." In *Financial policies and the world capital market: The problem of Latin American countries*. University of Chicago Press, 5–40.

Diaz-Alejandro, Carlos F. 1984. "Latin American debt: I don't think we are in Kansas anymore." *Brookings papers on economic activity* 1984 (2):335–403.

Ehrmann, Michael, Marcel Fratzscher, and Roberto Rigobon. 2011. "Stocks, bonds, money markets and exchange rates: measuring international financial transmission." *Journal of Applied Econometrics* 26 (6):948–974.

Elenev, Vadim, Tzuo Hann Law, Dongho Song, and Amir Yaron. 2022. "Fearing the fed: How wall street reads main street." *Available at SSRN 3092629* .

Etula, Erkko. 2013. "Broker-dealer risk appetite and commodity returns." *Journal of Financial Econometrics* 11 (3):486–521.

Faust, Jon, John H. Rogers, Shing-Yi B. Wang, and Jonathan H. Wright. 2007. "The high-frequency response of exchange rates and interest rates to macroeconomic announcements." *Journal of Monetary Economics* 54 (4):1051 – 1068.

Fleming, Michael J and Eli M Remolona. 1997. "What moves the bond market?" *Economic policy review* 3 (4).

Forbes, Kristin J and Francis E Warnock. 2012. "Capital flow waves: Surges, stops, flight, and retrenchment." *Journal of international economics* 88 (2):235–251.

Gardner, Ben, Chiara Scotti, and Clara Vega. 2022. "Words speak as loudly as actions: Central bank communication and the response of equity prices to macroeconomic announcements." *Journal of Econometrics* 231 (2):387–409.

Gertler, Mark and Peter Karadi. 2015. "Monetary policy surprises, credit costs, and economic activity." *American Economic Journal: Macroeconomics* 7 (1):44–76.

Gilbert, Thomas. 2011. "Information aggregation around macroeconomic announcements: Revisions matter." *Journal of Financial Economics* 101 (1):114–131.

Gilbert, Thomas, Chiara Scotti, Georg Strasser, and Clara Vega. 2017. "Is the intrinsic value of a macroeconomic news announcement related to its asset price impact?" *Journal of Monetary Economics* 92:78 – 95.

Goldberg, Linda S and Christian Grisse. 2013. "Time variation in asset price responses to macro announcements." Tech. rep., National Bureau of Economic Research.

Goldberg, Linda S and Cédric Tille. 2008. "Vehicle currency use in international trade." *Journal of international Economics* 76 (2):177–192.

Gopinath, Gita. 2015. "The international price system." Tech. rep., National Bureau of Economic Research.

Gopinath, Gita, Emine Boz, Camila Casas, Federico J. Díez, Pierre-Olivier Gourinchas, and Mikkel Plagborg-Møller. 2020. "Dominant Currency Paradigm." *American Economic Review* 110 (3):677–719.

Gormsen, Niels Joachim and Ralph SJ Koijen. 2020. "Coronavirus: Impact on stock prices and growth expectations." *The Review of Asset Pricing Studies* 10 (4):574–597.

Gourinchas, Pierre-Olivier and Helene Rey. 2007. "From world banker to world venture capitalist: US external adjustment and the exorbitant privilege." In *G7 current account imbalances: sustainability and adjustment*. University of Chicago Press, 11–66.

Gourinchas, Pierre-Olivier, Hélène Rey, and Maxime Sauzet. 2019. "The international monetary and financial system." *Annual Review of Economics* 11:859–893.

Gürkaynak, Refet, Brian Sack, and Eric Swanson. 2005a. "Do Actions Speak Louder Than Words? The Response of Asset Prices to Monetary Policy Actions and Statements." *International Journal of Central Banking* 1 (1).

Gürkaynak, Refet S, Burçin Kısacıkoğlu, and Jonathan H Wright. 2020. "Missing Events in Event Studies: Identifying the Effects of Partially Measured News Surprises." *American Economic Review* 110 (12):3871–3912.

Gürkaynak, Refet S., Brian Sack, and Eric Swanson. 2005b. "The Sensitivity of Long-Term Interest Rates to Economic News: Evidence and Implications for Macroeconomic Models." *American Economic Review* 95 (1):425–436.

Gürkaynak, Refet S. and Jonathan H. Wright. 2013. "Identification and Inference Using Event Studies." *The Manchester School* 81 (S1):48–65.

Hasbrouck, Joel. 2003. "Intraday price formation in US equity index markets." *The Journal of Finance* 58 (6):2375–2400.

Jarociński, Marek and Peter Karadi. 2020. "Deconstructing Monetary Policy Surprises—The Role of Information Shocks." *American Economic Journal: Macroeconomics* 12 (2):1–43.

Jiang, Zhengyang, Arvind Krishnamurthy, and Hanno Lustig. 2020. "Dollar safety and the global financial cycle." Tech. rep., National Bureau of Economic Research.

Jordà, Òscar, Moritz Schularick, Alan M Taylor, and Felix Ward. 2019. "Global financial cycles and risk premiums." *IMF Economic Review* 67 (1):109–150.

Kalemli-Özcan, Şebnem. 2019. "US Monetary Policy and International Risk Spillovers." *Federal Reserve Bank of Kansas City* Proceedings - Economic Policy Symposium - Jackson Hole.

Knox, Benjamin and Annette Vissing-Jorgensen. 2022. "A stock return decomposition using observables." .

Kurov, Alexander, Alessio Sancetta, Georg Strasser, and Marketa Halova Wolfe. 2019. "Price Drift Before U.S. Macroeconomic News: Private Information about Public Announcements?" *Journal of Financial and Quantitative Analysis* 54 (1):449–479.

Kuttner, Kenneth N. 2001. "Monetary policy surprises and interest rates: Evidence from the Fed funds futures market." *Journal of monetary economics* 47 (3):523–544.

Law, Tzuo Hann, Dongho Song, and Amir Yaron. 2018. "Fearing the fed: How wall street reads main street." *Available at SSRN 3092629* .

Lucca, David O and Emanuel Moench. 2015. "The pre-FOMC announcement drift." *The Journal of Finance* 70 (1):329–371.

Ludvigson, Sydney C., Sai Ma, and Serena Ng. 2021. "Uncertainty and Business Cycles: Exogenous Impulse or Endogenous Response?" *American Economic Journal: Macroeconomics* 13 (4):369–410.

Maggiori, Matteo, Brent Neiman, and Jesse Schreger. 2020. "International Currencies and Capital Allocation." *Journal of Political Economy* 128 (6):2019–2066.

Martin, Ian. 2017. "What is the Expected Return on the Market?" *The Quarterly Journal of Economics* 132 (1):367–433.

McCoy, Jack, Michele Modugno, Berardino Palazzo, and Steven A Sharpe. 2020. "Macroeconomic News and Stock Prices Over the FOMC Cycle." *FEDS Notes* (2020-10):14–1.

McQueen, Grant and V Vance Roley. 1993. "Stock prices, news, and business conditions." *The Review of Financial Studies* 6 (3):683–707.

Miranda-Agrippino, Silvia and Tsvetelina Nenova. 2022. "A tale of two global monetary policies." *Journal of International Economics* 136:103606.

Miranda-Agrippino, Silvia and Hélene Rey. 2020. "US monetary policy and the global financial cycle." *The Review of Economic Studies* 87 (6):2754–2776.

Miranda-Agrippino, Silvia and Hélène Rey. 2021. "The global financial cycle." Tech. rep., National Bureau of Economic Research.

Monnet, Eric and Mr Damien Puy. 2019. *One Ring to Rule Them All? New Evidence on World Cycles*. International Monetary Fund.

Nakamura, Emi and Jón Steinsson. 2018. "High-frequency identification of monetary non-neutrality: the information effect." *The Quarterly Journal of Economics* 133 (3):1283–1330.

Obstfeld, Maurice. 2015. "Trilemmas and trade-offs: living with financial globalisation." .

Reinhart, Carmen M. and Vincent R. Reinhart. 2008. "Capital Flow Bonanzas: An Encompassing View of the Past and Present." *NBER International Seminar on Macroeconomics* 5 (1):9–62.

Rey, Helene. 2013. "Dilemma not trilemma: the global cycle and monetary policy independence." Proceedings - Economic Policy Symposium - Jackson Hole, Federal Reserve Bank of Kansas City.

Rigobon, Roberto and Brian Sack. 2004. "The impact of monetary policy on asset prices." *Journal of monetary economics* 51 (8):1553–1575.

———. 2008. "Noisy macroeconomic announcements, monetary policy, and asset prices." In *Asset prices and monetary policy*. University of Chicago Press, 335–370.

Singleton, Kenneth J. 2006. "Empirical Dynamic Asset Pricing: Model Specification and Econometric Assessment."

Swanson, Eric T. 2021. "Measuring the effects of federal reserve forward guidance and asset purchases on financial markets." *Journal of Monetary Economics* .

Swanson, Eric T. and John C. Williams. 2014. "Measuring the Effect of the Zero Lower Bound on Medium- and Longer-Term Interest Rates." *American Economic Review* 104 (10):3154–85.

Online Appendix
for
The US, Economic News, and the Global Financial Cycle*

February 15, 2023

Christoph E. Boehm
UT Austin and NBER

T. Niklas Kroner
Federal Reserve Board

Table of Contents

A A Structural Framework to Interpret the Results 2
A.1 Setup 2
A.2 Discussion 3
A.3 Foreign Macroeconomic News 5
B Data Appendix 7
B.1 Macroeconomic News Releases 7
B.2 Intraday Financial Markets Data 9
B.3 Daily Financial Markets Data 12
B.4 Overview of Data Usage 14
C Additional Results 15
References 27

*The views expressed are those of the authors and do not necessarily reflect those of the Federal Reserve Board or the Federal Reserve System.
Email: chris.e.boehm@gmail.com and t.niklas.kroner@gmail.com.

A A Structural Framework to Interpret the Results

The following exposition extends the framework in Faust et al. (2007) to the international setting.

A.1 Setup

We adopt the high-frequency setup from Section 4, and denote by t the release time. The time window around the release is $[t-\Delta^-, t+\Delta^+]$, where Δ^- and Δ^+ are short time periods. We are interested in the effect of news about a US macroeconomic variable $y_{US,\tau}$ on an asset price q_i in country i. τ is a generic time index.

Letting $\mathcal{I}_{t-\Delta^-}$ denote agents' (common) information set prior to the news release, the *surprise* about the US macroeconomic variable is $s^y_{US,t} = y_{US,t} - E\left[y_{US,t}|\mathcal{I}_{t-\Delta^-}\right]$, where $E\left[\cdot|\mathcal{I}_{t-\Delta^-}\right]$ denotes the expectation conditional on information set $\mathcal{I}_{t-\Delta^-}$. Consistent with recent evidence (Gürkaynak, Kısacıkoğlu, and Wright, 2020), we assume that $s^y_{US,t}$ is measured without error. We denote the set of new information that becomes available in the time window we study by $\mathcal{N}_{[t-\Delta^-,t+\Delta^+]}$. It includes, in particular, news on the macroeconomic variable $y_{US,t}$, but also other news. Asset prices at time $t+\Delta^+$ are then based on the information set $\mathcal{I}_{t+\Delta^+} = \mathcal{I}_{t-\Delta^-} \cup \mathcal{N}_{[t-\Delta^-,t+\Delta^+]}$.

We assume a log-linear multi-country world with a unique equilibrium. Countries are indexed by i, j, and k, and $\mathcal{C}$ denotes the set of countries. The state variables of the economy are elements of the *vectors* $x_{j,\tau}$ and $x_{glob,\tau}$. State variables specific to country $j \in \mathcal{C}$ are included in the vector $x_{j,\tau}$ and global state variables are included in the vector $x_{glob,\tau}$. For instance, a component of total factor productivity (TFP) specific to the US is an element in vector $x_{US,\tau}$, while the global TFP component is included in $x_{glob,\tau}$. We are agnostic as to which state variables drive the business cycle and explicitly allow for news shocks in the spirit of Beaudry and Portier (2006). All structural shocks are uncorrelated.

The price of an asset of interest in country i can then be written as

$$q_{i,\tau} = E\left[\sum_{k\in\mathcal{C}} a^q_{i,k} x_{k,\tau} + a^q_{i,glob}\, x_{glob,\tau}|\mathcal{I}_\tau\right], \tag{A1}$$

where $a^q_{i,k}$, $k \in \mathcal{C}$, and $a_{glob,i}$ are coefficient vectors that depend on the specification of the model. They capture, respectively, how the asset price $q_{i,\tau}$ is affected by the country-specific state variables in $x_{k,\tau}$ and the global state variables in $x_{glob,\tau}$. Similarly, we can express country j's macroeconomic variable y of interest as

$$y_{j,\tau} = \sum_{k\in\mathcal{C}} a^y_{j,k} x_{k,\tau} + a^y_{j,glob}\, x_{glob,\tau}. \tag{A2}$$

For most of the paper, we are interested in US macroeconomic variables so that $j = US$.

Under the assumption that $x_{k,t+\Delta^+} = x_{k,t-\Delta^-}$ for all k and $x_{glob,t+\Delta^+} = x_{glob,t-\Delta^-}$ for small Δ^-, Δ^+, we can write the change in asset price $q_{i,\tau}$ over the window we study as

$$\begin{aligned}\Delta q_{i,t} &= q_{i,t+\Delta^+} - q_{i,t-\Delta^-} \\ &= \sum_{k\in\mathcal{C}} a^q_{i,k}\left(E\left[x_{k,t+\Delta^+}|\mathcal{I}_{t+\Delta^+}\right] - E\left[x_{k,t+\Delta^+}|\mathcal{I}_{t-\Delta^-}\right]\right) \\ &\qquad + a^q_{i,glob}\left(E\left[x_{glob,t+\Delta^+}|\mathcal{I}_{t+\Delta^+}\right] - E\left[x_{glob,t+\Delta^+}|\mathcal{I}_{t-\Delta^-}\right]\right).\end{aligned} \tag{A3}$$

In words, when new information becomes available, market participants change their expectations about the state of the economy, which in turn, changes asset price $q_{i,t}$.

We next use the fact that $\mathcal{I}_{t+\Delta^+} = \mathcal{I}_{t-\Delta^-} \cup \mathcal{N}_{[t-\Delta^-,t+\Delta^+]}$, and parameterize the conditional expectations in equation (A3),

$$E\left[x_{k,t+\Delta^+}|\mathcal{I}_{t+\Delta^+}\right] - E\left[x_{k,t+\Delta^+}|\mathcal{I}_{t-\Delta^-}\right] = b_k^y s_{US,t}^y + u_{k,t}, \qquad \text{for } k \in \mathcal{C}, \tag{A4}$$

$$E\left[x_{glob,t+\Delta^+}|\mathcal{I}_{t+\Delta^+}\right] - E\left[x_{glob,t+\Delta^+}|\mathcal{I}_{t-\Delta^-}\right] = b_{glob}^y s_{US,t}^y + u_{glob,t}. \tag{A5}$$

These expressions make explicit that market participants use the surprise about US macroeconomic news, as well as other information that becomes available within the time window (as captured by $u_{k,t}$ and $u_{glob,t}$), to update their expectations about the state of the world economy. To the extent that the US macroeconomic news release is informative about the state, the *vectors* b_k^y and b_{glob}^y contain nonzero elements. For instance, higher-than-expected US Nonfarm Payrolls may lead market participants to update their expectation of the US-specific component of TFP. In this case, the relevant element in b_{US}^y is nonzero. If the surprise is not useful for estimating particular state variables, then the relevant entries in b_k^y and b_{glob}^y are zero.

We make no specific assumptions on how agents update their estimate of the state. They could, for instance, use the Kalman filter, but we do not impose this assumption. We only require that the estimation of the unobserved state requires a nonzero correlation between the observed macroeconomic variable and the state of interest. Formally, we require

Assumption 1. *For all $k \in \mathcal{C} \cup \{glob\}$: $b_k^y \neq 0 \Rightarrow a_{US,k}^y \neq 0$.*

Plugging equations (A4) and (A5) into equation (A3) gives

$$\Delta q_{i,t} = \left(\sum_{k\in\mathcal{C}} a_{i,k}^q b_k^y + a_{i,glob}^q b_{glob}^y\right) s_{US,t}^y + \varepsilon_{i,t}, \tag{A6}$$

where $\varepsilon_{i,t} = \sum_{k\in\mathcal{C}} a_{i,k}^q u_{k,t}^y + a_{i,glob}^q u_{glob,t}$. Letting $\gamma_i := \sum_{k\in\mathcal{C}} a_{i,k}^q b_k^y + a_{i,glob}^q b_{glob}^y$, delivers our estimating equation (4).

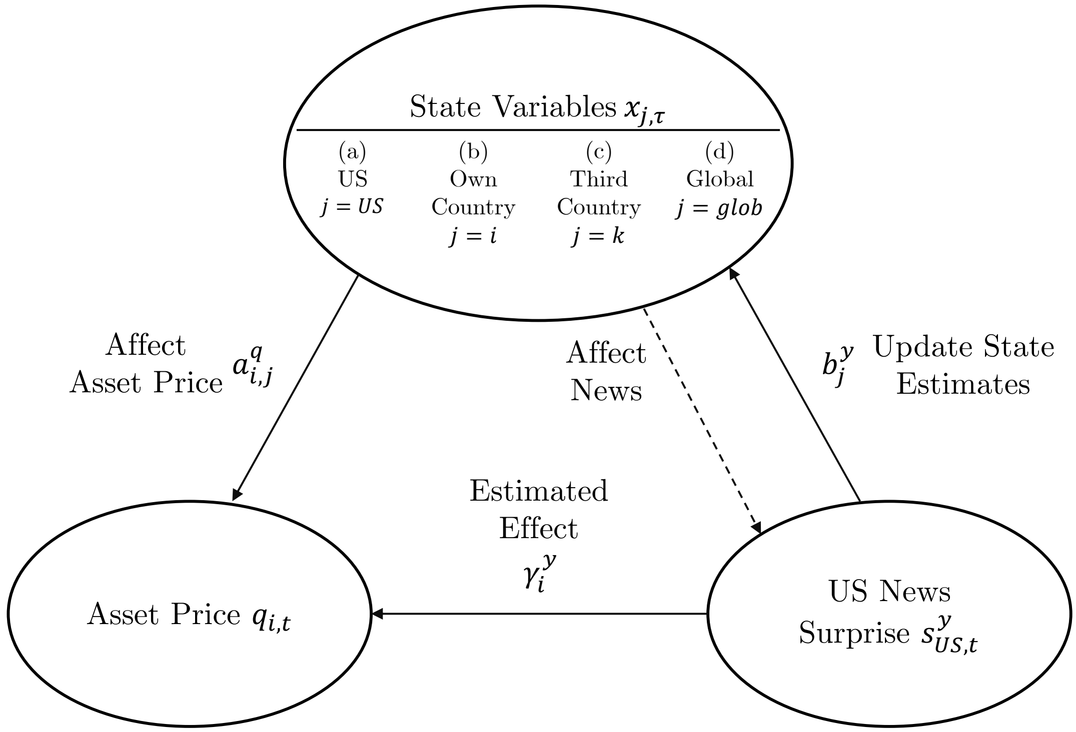
A.2 Discussion

For a given asset price $q_{i,t}$ and surprise $s_{US,t}^y$, equation (A6) highlights that a country's response reflects two components. First, the response reflects the asset price's dependence on the true unobserved state, as captured by $a_{i,k}^q$ and $a_{i,glob}^q$. Second, the response reflects market participants' updates about the state of the world, as measured by vectors b_k^y and b_{glob}^y. If market participants use the newly available information to update only some state variables, and country i's asset price does not depend on the state variables being updated, then the asset price should not systematically respond to the surprise. The nonzero responses that we identified in Section 4 thus imply that market participants update their belief about states, which country i's asset price depends on.

We next split the asset price response in equation (A6) by country into four different components,

$$\Delta q_{i,t} = \left(\underbrace{a_{i,US}^q b_{US}^y}_{\text{(a)}} + \underbrace{a_{i,i}^q b_i^y}_{\text{(b)}} + \underbrace{\sum_{j\neq US,i} a_{i,j}^q b_j^y}_{\text{(c)}} + \underbrace{a_{i,glob}^q b_{glob}^y}_{\text{(d)}}\right) s_{US,t}^y + \varepsilon_{i,t}. \tag{A7}$$

Figure A1: Interpretation of Country's i Asset Price Response to US News (with details)



Notes: The figure illustrates the discussion in the text. Solid arrows display relevant relationships at the time of the news release, as captured by equation (A7). The dashed arrow indicates that the relationship is predetermined at the time of the release.

This breakdown reflects the origins of disturbances. Term (a) captures economic disturbances originating in the US. If, for instance, the change in US TFP affects US macroeconomic variable $y_{US,\tau}$, market participants who observe the surprise $s^y_{US,t}$ may update their estimate of US TFP. This would be captured by a nonzero element in vector b^y_{US}. At the same time the change in US TFP may affect foreign asset price $q_{i,t}$—as captured by a nonzero entry in vector $a^q_{i,US}$. The asset price in country i only responds to a change in US TFP if both market participants update their expectation of US TFP *and* US TFP indeed affects the asset price in country i. More generally, term (a) captures this logic for all US state variables and thus reflects country i's asset price responses to disturbances originating in the US.

Term (b) in the above expression reflects changes in state variables, which originate in country i. In order for an innovation to the state in country i to affect i's own asset price *through the US macroeconomic surprise*, it would have to the case that market participants learn about i's state by studying US macroeconomic news. Similarly, term (c) captures disturbances, which originate in a third country j, and affect both US macro news as well as the asset price in country i. Lastly, term (d) reflects changes in the global state vector. Such disturbances may affect US macroeconomic surprises, and as a result market participants may use these surprises to estimate these global state variables. Figure A1 illustrates this intuition.

A reasonable assumption in the context of our analysis is that surprises in US macroeconomic variables are not used to update state variables that are specific to countries other than the US.

That is, $b^y_k = 0$ for $k \notin \{US, glob\}$. This assumption implies that it is not the case that market participants use US payroll employment to forecast the country-specific component of Belgian TFP. Under Assumption 1, a sufficient condition for this to hold is that countries other than the US are *small* relative to the US. Continuing with the earlier example, a change in Belgian TFP has no impact on US macroeconomic variables, and hence, the forecaster would find no useful correlation to predict Belgian TFP when new information about the US macroeconomy becomes available. Formally, Assumption 1 immediately implies that $a^y_{US,BEL} = 0 \Rightarrow b^y_{BEL} = 0$. The premise is satisfied because Belgium is small relative to the US.

Under this assumption, equation (A6) becomes

$$\Delta q_{i,t} = \left(\underbrace{a^q_{i,US} b^y_{US}}_{\text{transmission from US}} + \underbrace{a^q_{i,glob} b^y_{glob}}_{\text{common shock}} \right) s^y_{US,t} + \varepsilon_{i,t}. \tag{A8}$$

This estimating equation makes clear that a significant coefficient on the US macroeconomic surprise reflects two different components. First, if the surprise leads to an update of market participants' expectations on US state variables (as captured by nonzero elements in the vector b^y_{US}), and if changes in US state variables impact the foreign asset price (the vector $a^q_{i,US}$ contains nonzero elements), then the inner product $a^q_{i,US} b^y_{US}$ can be different from zero. This component thus reflects *transmission* of macroeconomic shocks from the US to country i. Second, the surprise $s^y_{US,t}$ may be useful to forecast global state variables (b^y_{glob} contains nonzero elements). In this case, a significant coefficient on the surprise reflects that country i is impacted by a *common shock*.

This discussion helps interpret our estimates in Section 4. While foreign stock prices strongly respond to the release of US macroeconomic news, this does not necessarily imply the transmission of US shocks to foreign countries. It is also possible that the US and other countries are subject to common shocks. These common shocks affect US macroeconomic outcomes and are therefore reflected in the measured surprises. Foreign stock markets respond to these surprises, because they reveal information about the common state vector.

A.3 Foreign Macroeconomic News

To test for the presence of common shocks, we study the effect of foreign news releases on the US stock market. In particular, we regress the log-change in the S&P 500 on foreign macroeconomic surprises,

$$\Delta q_{US,t} = \zeta^y_i s^y_{i,t} + \varepsilon_{i,t}, \tag{A9}$$

where we omit the constant and controls for clarity. Analogous to Section A.1, it is possible to obtain a structural interpretation of the estimated coefficient ζ^y_i. In particular, we can write

$$\zeta^y_i = a^q_{US,US} b^y_{US,i} + a^q_{US,i} b^y_{i,i} + \sum_{k \neq US,i} a^q_{US,k} b^y_{k,i} + a^q_{US,glob} b^y_{glob,i}, \tag{A10}$$

where the vectors $b^y_{j,i}$ ($b^y_{glob,i}$) are now specific to country i, and capture how market participants update their estimate of country j's state $x_{j,t}$ (the global state $x_{glob,t}$) upon observing news in country i. Further, vectors $a^q_{US,k}$ ($a^q_{US,glob}$) capture how country k's (the global) state affects the US stock market.

Studying the effects of foreign news on the US stock market—rather than on a third country—has a key advantage. Since most countries are small relative to the US, the interpretation of coefficient (A10) simplifies considerably. In particular, under the assumptions that (i) country i is small relative to the US so that $a^q_{US,i} = 0$, and (ii) country i's news does not affect the US stock market through third countries ($a^q_{US,k} b^y_{k,i} = 0$ for all k),[1] the estimated coefficient simplifies to

$$\zeta^y_i = \underbrace{a^q_{US,US} b^y_{US,i}}_{\text{(a)}} + \underbrace{a^q_{US,glob} b^y_{glob,i}}_{\text{(b)}}. \tag{A11}$$

These remaining two terms reflect the following intuition. First, term (a) reflects the possibility that market participants learn about the US state vector by observing foreign macroeconomic news. Since the US is large relative to country i, shocks in the US are likely to have an effect on country i's macroeconomic outcomes. As a result, country i's surprises could be informative about US-specific shocks. While this possibility cannot be ruled out *a priori*, we don't view it as particularly plausible either. Since US shocks presumably affect foreign macroeconomic outcomes with a lag and many indicators of US macroeconomic performance become available in a timely fashion, it is rather unlikely that this indirect channel of learning about the US state is active in practice.

Second, term (b) reflects the presence of common shocks. As noted earlier, if countries' macroeconomic and financial variables were driven by common global state variables, other countries' macroeconomic releases should generally be informative about it. Further, this state should drive international asset prices, including the S&P 500.

[1] The second assumption is satisfied if third countries are small relative to the US so that $a^q_{US,k} = 0$ or if market participants do not update their estimate of country k's state vector upon observing country i's macroeconomic news ($b^y_{k,i} = 0$).

B Data Appendix

In this Appendix, we provide an overview of the main datasets used in the paper. In Section B.1, we describe the data on macro news releases. In Section B.2 and Section B.3, we provide details on the intraday and daily financial markets data, respectively. In Section B.4, we discuss which data is used in which part of the paper.

B.1 Macroeconomic News Releases

Data Series For a given release, we use the following series in the paper, which if not otherwise noted are taken directly from Bloomberg:

- *Announcement time*
- *Forecast*: median survey estimate of professional forecasters in Bloomberg
- *Initial released number*: released number at time of announcement
- *Final revised number*: final revised number as of 2022
- *Reference period*: period which released number is referencing to (e.g., month X for a monthly release)
- *Surprise*: constructed from *forecast* and *initial released number* as shown in equation (1)
- *Category*: manual selection into real activity or price news based on Beechey and Wright (2009)
- *Reporting lag*: measure of inverse timeliness constructed from announcement time and reference period (see equation (8))
- *Revision magnitude*: measure of inverse quality constructed from *initial released number* and *final revised number* (see equation (9))
- *Relevance*: measure of relative popularity reflecting how many people within Bloomberg set an alert for a certain release relative to all alerts set for a given country. The measure is between 0 and 100 as it is measured in percent.

Sample Construction For both US and foreign countries, we obtain the final set of news releases based on the following two criteria: First, at least 50 observations with both initial released number and forecast are available in order to construct a surprise. Second, relevance of the series is greater than or equal to 30. We end up in total with 66 announcement series for the US, 23 for Canada, 16 for France, 23 for Germany, 16 for Italy, 50 for Japan, and 43 for the United Kingdom. Table B1 lists all 66 releases for the US. Table B2 provides an overview of the 60 major releases of the other G7 countries. Note that for each announcement, we remove surprises which are more than 6 standard deviations in absolute value.

Table B1: Overview of All US Macroeconomic News

Announcement	Frequency	Category	Observations	Announcement	Frequency	Category	Observations
ADP Employment	Monthly	Real Activity	160	ISM Chicago Index	Monthly	Real Activity	275
Average Hourly Earnings	Monthly	Price	258	ISM Mfg Index	Monthly	Real Activity	277
Building Permits	Monthly	Real Activity	208	ISM Non-Mfg Index	Monthly	Real Activity	251
Business Inventories	Monthly	Real Activity	269	ISM Prices Paid	Monthly	Price	234
CB Consumer Confidence	Monthly	Real Activity	273	Import Price Index	Monthly	Price	253
CB Leading Economic Index	Monthly	Real Activity	272	Industrial Production	Monthly	Real Activity	277
CPI	Monthly	Price	277	Initial Jobless Claims	Weekly	Real Activity	1166
Capacity Utilization	Monthly	Real Activity	274	Mfg Payrolls	Monthly	Real Activity	252
Capital Goods Orders	Monthly	Real Activity	112	NAHB Housing Market Index	Monthly	Real Activity	201
Capital Goods Shipments	Monthly	Real Activity	95	NFIB Small Business Optimism	Monthly	Real Activity	118
Chicago Fed Nat Activity Index	Monthly	Real Activity	107	NY Fed Mfg Index	Monthly	Real Activity	206
Construction Spending	Monthly	Real Activity	252	Net Long-term TIC Flows	Monthly	Real Activity	117
Consumer Credit	Monthly	Real Activity	277	New Home Sales	Monthly	Real Activity	267
Continuing Claims	Weekly	Real Activity	863	Nonfarm Payrolls	Monthly	Real Activity	274
Core CPI	Monthly	Price	275	Nonfarm Productivity F	Quarterly	Real Activity	86
Core PCE Price Index	Monthly	Price	174	Nonfarm Productivity P	Quarterly	Real Activity	87
Core PPI	Monthly	Price	275	PPI	Monthly	Price	263
Current Account Balance	Quarterly	Real Activity	87	Pending Home Sales	Monthly	Real Activity	176
Dallas Fed Mfg Index	Monthly	Real Activity	131	Personal Consumption Expenditure	Monthly	Real Activity	273
Durable Goods Orders	Monthly	Real Activity	266	Personal Income	Monthly	Real Activity	274
Durables Ex Transportation	Monthly	Real Activity	217	Philly Fed Business Outlook	Monthly	Real Activity	273
Employment Cost Index	Quarterly	Price	91	Private Payrolls	Monthly	Real Activity	116
Existing Home Sales	Monthly	Real Activity	178	Retail Sales	Monthly	Real Activity	275
FHFA House Price Index	Monthly	Price	138	Retail Sales Ex Auto	Monthly	Real Activity	270
Factory Orders	Monthly	Real Activity	277	Richmond Fed Mfg Index	Monthly	Real Activity	170
GDP A	Quarterly	Real Activity	91	Total Vehicle Sales	Monthly	Real Activity	82
GDP Price Index A	Quarterly	Price	87	Trade Balance	Monthly	Real Activity	277
GDP Price Index S	Quarterly	Price	87	UM Consumer Sentiment F	Monthly	Real Activity	248
GDP Price Index T	Quarterly	Price	85	UM Consumer Sentiment P	Monthly	Real Activity	247
GDP S	Quarterly	Real Activity	90	Unemployment Rate	Monthly	Real Activity	273
GDP T	Quarterly	Real Activity	91	Unit Labor Costs F	Quarterly	Price	81
Government Budget Balance	Monthly	Real Activity	273	Unit Labor Costs P	Quarterly	Price	81
Housing Starts	Monthly	Real Activity	260	Wholesale Inventories	Monthly	Real Activity	270

Notes: This table provides information on all US macroeconomic series utilized in the paper. The sample ranges from October 1996 to December 2019. *Observations* refers to number of observations (surprises) of a macroeconomic series in the sample and *Frequency* to the frequency of the data releases. Abbreviations: A—advanced; S—second; T—third; P—preliminary; F—final; Mfg—Manufacturing; ADP—Automatic Data Processing Inc; CB—Chicago Board; ISM—Institute for Supply Management; UM—University of Michigan; NFIB—National Federation of Independent Business; NAHB—National Association of Home Builders.

Table B2: Overview of Major Foreign Macroeconomic News

Announcement	Frequency	Observations	Announcement	Frequency	Observations
Canada			*Italy*		
Capacity Utilization	Quarterly	79	Consumer Confidence	Monthly	221
Core CPI	Monthly	226	CPI P	Monthly	259
GDP	Quarterly	81	GDP F	Quarterly	80
Housing Starts	Monthly	233	Industrial Production	Monthly	248
Intl. (Merchandise) Trade	Monthly	273	Industrial Sales	Monthly	63
IPPI (Industrial Product Price Index)	Monthly	255	Mfg Confidence	Monthly	233
Mfg Sales	Monthly	273	PPI	Monthly	190
PMI (Purchasing Managers Index)	Monthly	195	Retail Sales	Monthly	173
Retail Sales	Monthly	266	Trade Balance	Monthly	76
Unemployment Rate	Monthly	274	Unemployment Rate	Monthly	146
France			*Japan*		
BoF Industry Sentiment	Monthly	135	BoJ (Tankan) Mfg Index	Quarterly	86
Consumer Confidence	Monthly	237	BoJ (Tankan) Mfg Outlook	Quarterly	60
CPI P	Monthly	259	Consumer Confidence	Monthly	153
GDP P	Quarterly	89	CPI	Monthly	219
Industrial Production	Monthly	271	Exports	Monthly	130
Mfg Confidence	Monthly	218	GDP P	Quarterly	89
PPI	Monthly	159	Industrial Production P	Monthly	239
Production Outlook	Monthly	187	PPI	Monthly	237
Trade Balance	Monthly	270	Retail Sales	Monthly	199
Unemployment Rate	Monthly/Quarterly	174	Unemployment (Jobless) Rate	Monthly	239
Germany			*United Kingdom*		
CPI P	Monthly	242	Core CPI	Monthly	172
GDP	Quarterly	90	Core PPI (Output)	Monthly	168
GfK Consumer Confidence	Monthly	159	Exports	Quarterly	59
IFO Business Climate	Monthly	271	GDP A	Quarterly	86
Industrial Production	Monthly	270	GfK Consumer Confidence	Monthly	205
PPI	Monthly	275	House Price Index	Monthly	187
Retail Sales	Monthly	255	Industrial Production	Monthly	275
Trade Balance	Monthly	273	Jobless Claims	Monthly	240
Unemployment Change	Monthly	274	Retail Sales	Monthly	118
ZEW Survey Expectations	Monthly	214	Unemployment Rate	Monthly	211

Notes: This table provides information on the macroeconomic series of non-US G7 countries utilized in Section 7. The data is obtained from Bloomberg's Economic Calendar and the sample ranges from October 1996 to December 2019. *Observations* refers to number of observations (surprises) of a macroeconomic series in the sample and *Frequency* to the frequency of the data releases. Note that the reported number of observations in Table 6 is smaller than the one reported here due to the unavailability of the E-mini S&P 500 futures on certain dates. Abbreviations: A—advanced; BoF—Bank of France; BoJ—Bank of Japan; F—final; GFK—Society for Consumer Research; IFO—Institute for Economic Research; ILO—International Labor Organization; Mfg—Manufacturing; P—preliminary.

B.2 Intraday Financial Markets Data

All intraday data on asset prices comes from *Thomson Reuters Tick History* dataset and is obtained via *Refinitiv*. We inspect each data series for potential misquotes, and remove them if necessary. As discussed in Section 3, our sample of countries is based on the trading hours, market liquidity, and availability of historical data. Table B3 provides an overview of the full dataset. Table B4 provides an overview of which stock markets are open for each of the twelve major US macro releases. Table B5 displays an overview of the other intraday data series used throughout the paper. Note that the intraday data which is used in the context of the monetary policy shocks is detailed in Table S3.1.

Table B3: Overview of Intraday Data on International Financial Markets

Country	ISO	Stock Index		Dollar Exchange Rate		1-Y Govt. Bond Yield		10-Y Govt. Bond Yield	
		Ticker	Sample	Ticker	Sample	Ticker	Sample	Ticker	Sample
Argentina	ARG	.MERV	1996–2019	ARS=	1996–2019			AR10YT=RR	1999–2017
Brazil	BRA	.BVSP	1996–2019	BRL=	1996–2019	BR1YT=RR	2007–2019	BR10YT=RR	1998–2019
Canada	CAN	.TSE300/.GSPTSE	2000–2019	CAD=	1996–2019			CA10YT=RR	1996–2019
Switzerland	CHE	.SSMI	1996–2019	CHF=	1996–2019	CH1YT=RR	2002–2019	CH10YT=RR	1996–2019
Chile	CHL	.IPSA/.SPCLXIPSA/.SPIPSA	1996–2019	CLP=	1996–2019			CL10YT=RR	2007–2019
Czech Republic	CZE	.PX50/.PX	1999–2019	CZE=	1996–2019	CZ1YT=RR	1998–2019	CZ10YT=RR	2000–2019
Denmark	DNK	.KFMX/.OMXCXC20PI	2000–2019			DK1YT=RR	1996-2017	DK10YT=RR	1996–2019
United Kingdom	GBR	.FTSE	1996–2019	GBP=	1996–2019	GB1YT=RR	1996-2019	GB10YT=RR	1996–2019
Hungary	HUN	.BUX	1997–2019	HUF=	1996–2019			HU10YT=RR	1999–2019
Mexico	MEX	.MXX	1996–2019	MXN=	1996–2019			MX10YT=RR	2002–2019
Norway	NOR	.OBX	1996–2019	NOK=	1996–2019			NO10YT=RR	1996–2019
Poland	POL	.WIG20	1997–2019	PLN=	1996–2019			PL10YT=RR	1999–2019
Russia	RUS	.MCX/.IMOEX	2001–2019	RUB=	1998–2019	RU1YT=RR	2001–2019	RU10YT=RR	2003–2019
Sweden	SWE	.OMX	1996–2019	SEK=	1996–2019			SE10YT=RR	1996–2019
Turkey	TUR	.XU030	1997–2019	TRY=	2004–2019			TR10YT=RR	2010–2019
South Africa	ZAF	.JTOPI	2002–2019	ZAR=	1996–2019			ZA10YT=RR	1997–2019
Euro Area	EUR			EUR=	1999–2019				
Austria	AUT	.ATX	1996–2019			AT1YT=RR	2002–2019	AT10YT=RR	1996–2019
Belgium	BEL	.BFX	1996–2019			BE1YT=RR	2004–2019	BE10YT=RR	1996–2019
Germany	DEU	.GDAXI	1996–2019			DE1YT=RR	2004–2019	DE10YT=RR	1996–2019
Spain	ESP	.IBEX	1996–2019			ES1YT=RR	2010–2019	ES10YT=RR	1996–2019
Finland	FIN	.HEX25	2001–2019					FI10YT=RR	1996–2019
France	FRA	.FCHI	1996–2019			FR1YT=RR	1996–2019	FR10YT=RR	1996–2019
Greece	GRC	.ATF	1997–2019					GR10YT=RR	1998–2019
Ireland	IRL	.ISEQ	1996–2019			IE1YT=RR	1998–2019	IE10YT=RR	1998–2019
Italy	ITA	.MIB30/.SPMIB/.FTMIB	1996–2019			IT1YT=RR	1996–98,09–2019	IT10YT=RR	1996–2019
Netherlands	NLD	.AEX	1996–2019			NL1YT=RR	1996–2019	NL10YT=RR	1996–2019
Portugal	PRT	.PSI20	1996–2019			PT1YT=RR	2004–2019	PT10YT=RR	1996–2019

Notes: This table gives an overview of part of the cross-country intraday data from *Thomson Reuters Tick History* utilized in the paper. For all series the sample period ends in December 2019. *Ticker* refers to the Reuters Instrument Code (RIC). For a given country, the table provides details of the major stock index, US exchange rate, and 10-year government bond yield with the respective samples periods. For members of the Euro Area, we do not use country-specific exchange rates prior to the inception of the currency union due to the short sample length. Further, we drop Denmark from the sample since the Danish Krone is tightly and credibly pegged to the Euro. Abbreviations: ISO—3 digit ISO country code.

Table B4: Overview of Open/Closed Equity Markets during US Macroeconomic News Announcements

Event	ARG	AUT	BEL	BRA	CAN	CHE	CHL	CZE	DEU	DNK	ESP	FIN	FRA	GBR
Capacity Utilization	Open	Open	Open	Open	Closed	Open	Open	Open	Open	Open	Open	Open	Open	Open
CB Consumer Confidence	Open	Open	Open	Open	Open	Open	Open	Open	Open	Open	Open	Open	Open	Open
Core CPI	Closed	Open	Open	Open	Closed	Open	Open	Open	Open	Open	Open	Open	Open	Open
Core PPI	Closed	Open	Open	Open	Closed	Open	Open	Open	Open	Open	Open	Open	Open	Open
Durable Goods Orders	Closed	Open	Open	Open	Closed	Open	Open	Open	Open	Open	Open	Open	Open	Open
GDP A	Closed	Open	Open	Open	Closed	Open	Open	Open	Open	Open	Open	Open	Open	Open
Initial Jobless Claims	Closed	Open	Open	Open	Closed	Open	Open	Open	Open	Open	Open	Open	Open	Open
ISM Mfg Index	Open	Open	Open	Open	Open	Open	Open	Open	Open	Open	Open	Open	Open	Open
New Home Sales	Open	Open	Open	Open	Open	Open	Open	Open	Open	Open	Open	Open	Open	Open
Nonfarm Payrolls	Closed	Open	Open	Open	Closed	Open	Open	Open	Open	Open	Open	Open	Open	Open
Retail Sales	Closed	Open	Open	Open	Closed	Open	Open	Open	Open	Open	Open	Open	Open	Open
UM Consumer Sentiment P	Open	Open	Open	Open	Open	Open	Open	Open	Open	Open	Open	Open	Open	Open

	GRC	HUN	IRL	ITA	MEX	NLD	NOR	POL	PRT	RUS	SWE	TUR	ZAF
Capacity Utilization	Open	Open	Open	Open	Open	Open	Open	Open	Open	Open	Open	Open	Open
CB Consumer Confidence	Open	Open	Open	Open	Open	Open	Open	Open	Open	Open	Open	Open	Open
Core CPI	Open	Open	Open	Open	Closed	Open	Open	Open	Open	Open	Open	Open	Open
Core PPI	Open	Open	Open	Open	Closed	Open	Open	Open	Open	Open	Open	Open	Open
Durable Goods Orders	Open	Open	Open	Open	Closed	Open	Open	Open	Open	Open	Open	Open	Open
GDP A	Open	Open	Open	Open	Closed	Open	Open	Open	Open	Open	Open	Open	Open
Initial Jobless Claims	Open	Open	Open	Open	Closed	Open	Open	Open	Open	Open	Open	Open	Open
ISM Mfg Index	Open	Open	Open	Open	Open	Open	Open	Open	Open	Open	Open	Open	Open
New Home Sales	Open	Open	Open	Open	Open	Open	Open	Open	Open	Open	Open	Open	Open
Nonfarm Payrolls	Open	Open	Open	Open	Closed	Open	Open	Open	Open	Open	Open	Open	Open
Retail Sales	Open	Open	Open	Open	Closed	Open	Open	Open	Open	Open	Open	Open	Open
UM Consumer Sentiment P	Open	Open	Open	Open	Open	Open	Open	Open	Open	Open	Open	Open	Open

Notes: *Green* indicates that the corresponding equity market is usually open at the time of the news release. *Orange* indicates that the equity market is usually open but that the news release is around market opening or closing. In the case of Brazil, it indicates that the news release moves outside the trading hours during the US daylight saving time since Sao Paulo, the location of the Brazilian stock market, does not observe daylight saving time. *Red* indicates that the equity market is usually closed at the release time.

Table B5: Overview of Other Intraday Financial Data

Name	Ticker	Sample
Stock Index Futures		
E-mini S&P 500 Futures	ESc1	1997–2019
AEX Futures (NLD)	AEXc1	1997–2019
CAC 40 Futures (FRA)	FCEc1	1999–2019
DAX Futures (DEU)	FDXc1	1996–2019
FTSE 100 Futures (GBR)	FFIc1	1998–2019
SMI Futures (CHE)	FSMIc1	1998–2019
Bovespa Futures (BRA)	INDc1	1996–2019
S&P/TSX 60 Futures (CAN)	SXFc1	1999–2019
Volatility Indexes		
VIX	.VIX	1996–2019
VIX Futures	VXc1:VE/VXc1	2011–2019
VSTOXX	.V2TX	2005–2019
VDAX	.V1XI	2005–2019
VFTSE	.VFTSE	2006–2019
VCAC	.VCAC	2007–2019
Interest Rates		
1- & 4-Quarter Eurodollar Futures	EDcm1/EDcm4	1996–2019
2-Year Treasury Futures	TUc1/TUc2	1996–2019
10-Year Treasury Futures	TYc1/TYc2	1996–2019
Commodity Indexes		
S&P GSCI Agriculture	.SPGSAG	2007–2019
S&P GSCI Energy	.SPGSEN	2007–2019
S&P GSCI Industrial Metals	.SPGSINTR	2007–2019

Notes: This table gives an overview of additional intraday data series utilized in the paper, complementing Table B3. The data comes from *Thomson Reuters Tick History*. For all series, the sample period ends in December 2019. *Ticker* refers to the Reuters Instrument Code (RIC). Abbreviations: ISO—3 digit ISO country code.

B.3 Daily Financial Markets Data

This section provides details on the daily data employed in the paper. Table B6 documents for each series its source, sample period, and reference paper if applicable. Based on these series, we construct a proxy for the equity premium and for growth expectations, as well as stock price (semi-)elasticities. All of these are used in our analysis in Section 7.3. We next discuss the construction of the variables of interest. Note that since the associated analysis exclusively focuses on the US, we omit the country subscript for brevity.

We start with the equity premium. Under the assumption that Martin's (2017) lower bound on the equity premium binds (as argued by Martin, 2017), the *1-year equity premium* on day d, i.e., the expected excess return over the next year, can be calculated as

$$ep_{d,1} = (1 + r^f_{d,1}) svix^2_{d,1}, \tag{B1}$$

where $svix_{d,1}$ is the 1-year SVIX on day d and $r^f_{d,1}$ is the expected risk-free rate over the next 1-year on day d. As shown in Table B6, we take the former series directly from Martin (2017) and the

Table B6: Overview of Daily Data

Series	Reference	Source	Sample
S&P 500		CRSP via WRDS	1996–2019
VIX		CBOE via WRDS	1996–2019
VSTOXX		Bloomberg (Ticker: V2X)	1999–2019
VDAX		Bloomberg (Ticker: V1X)	1992–2019
VFTSE		Bloomberg (Ticker: VFTSE)	2000–2019
VCAC		Bloomberg (Ticker: VCAC)	2000–2019
S&P 500 Dividend Futures		Bloomberg (Tickers: ASD1–ASD10)	2015/2017–2019
SVIX	Martin (2017)	Martin and Wagner (2019)	1996–2014
Treasury Yields	Gürkaynak, Sack, and Wright (2007)	Federal Reserve Bank	1996–2019
Expected Short Rates & Term Premia	Adrian, Crump, and Moench (2013)	Federal Reserve Bank of New York	1996–2019

Notes: This table provides an overview of the daily data series employed in the paper including literature reference where applicable, source, and available sample period.

latter from Adrian, Crump, and Moench (2013). With the equity premium in hand, it is convenient to define the 1-year gross discount rate

$$\theta_{d,1} = 1 + r^f_{d,1} + ep_{d,1}. \tag{B2}$$

$\theta_{d,1}$ discounts the next year's expected dividend of the market which we define next.

To proxy for *1-year growth expectations* on day d, we employ the next year's expected dividend, which can be expressed as

$$\begin{aligned} div_{d,1} &= \frac{(1+\theta_{d,1})f_{d,1}}{1+r^f_{d,1}}, \\ &= \frac{((1+r^f_{d,1}) + (1+r^f_{d,1})svix^2_{d,1})f_{d,1}}{1+r^f_{d,1}}, \\ &= \left(1+svix^2_{d,1}\right) f_{d,1}, \end{aligned} \tag{B3}$$

where $div_{d,1} \equiv E_d[div_{d+365}]$, and $f_{d,1}$ is the price of 1-year dividend futures contract at date d.[2] The first equality shows the relationship between expected dividend and dividend futures contract as shown in Gormsen and Koijen (2020). Plugging in equations (B1) and (B2) yields the last term.

We next turn to the construction of *stock price (semi-)elasticities*. To do so, we define the returns that we use in Section 7.3. These are $\Delta q_{d,1} \equiv q_{d,1} - q_{d-1,1}$, $\Delta ep_{d,1} \equiv ep_{d,1} - ep_{d-1,1}$, $\Delta r^f_{d,1} \equiv r^f_{d,1} - r^f_{d-1,1}$, and $\Delta div_{d,1} \equiv \frac{div_{d,1} - div_{d-1,1}}{div_{d-1,1}}$. Under the assumptions discussed in Knox and Vissing-Jorgensen (2022), we can construct *stock price (semi-)elasticities* as follows. The elasticity of the S&P 500 with respect to the next year's expected dividend $div_{d,1}$ is given by

$$\frac{\Delta q_d}{\Delta div_{d,1}} = \frac{f_{d-1,1}}{(1+r^f_{d-1,1})q_{d-1}},$$

[2]Note that the price of 1-year fixed-horizon dividend futures contract $f_{d,1}$ is interpolated based on the prices of current and the next year S&P 500 Annual Dividend Index futures contracts which have an annual fixed expiration. The underlying security is the S&P 500 Annual Dividend Points Index which tracks the total dividends from S&P 500 constituents over a year before resetting to zero at the end of each year.

where the right-hand side is the weight on the 1-year dividend strip on day $d-1$.

The semi-elasticity of the stock price with respect to 1-year equity premium and risk-free rate is given by

$$\frac{\Delta q_d}{\Delta r^f_{d,1}} = \frac{\Delta q_d}{\Delta ep_{d,1}} = -\frac{1}{\theta_{d-1,1}}.$$

With the (semi-)elasticities in hand, we next turn to several practical issues we face. Since the SVIX is not available to us over the entire sample, we set the SVIX to its sample average on missing days when we construct the elasticity (changes in the equity premium are only based on days for which the SVIX is available). This allows us to construct the discount rate $\theta_{d,1}$ for our entire period. Further, as we do not have data on the SVIX and the dividend futures contract for an overlapping sample period, we assume that daily changes in the SVIX are roughly zero. Based on equation (B3), this allows us to construct changes in the expected dividends directly from price changes in the dividend futures contract, i.e., $\Delta div_{d,1} = \Delta f_{d,1}$. While this induces a small bias, we know in case of our analysis in Section 7 in which direction this bias goes. As Panel B of Table 10 shows that positive real activity news decreases the 1-year equity premium and increases the risk-free rate, we can infer from equation (B1) that it decreases the SVIX. Hence, the response of the price of the futures contract will slightly overstate the effect of real activity news on expected dividends.

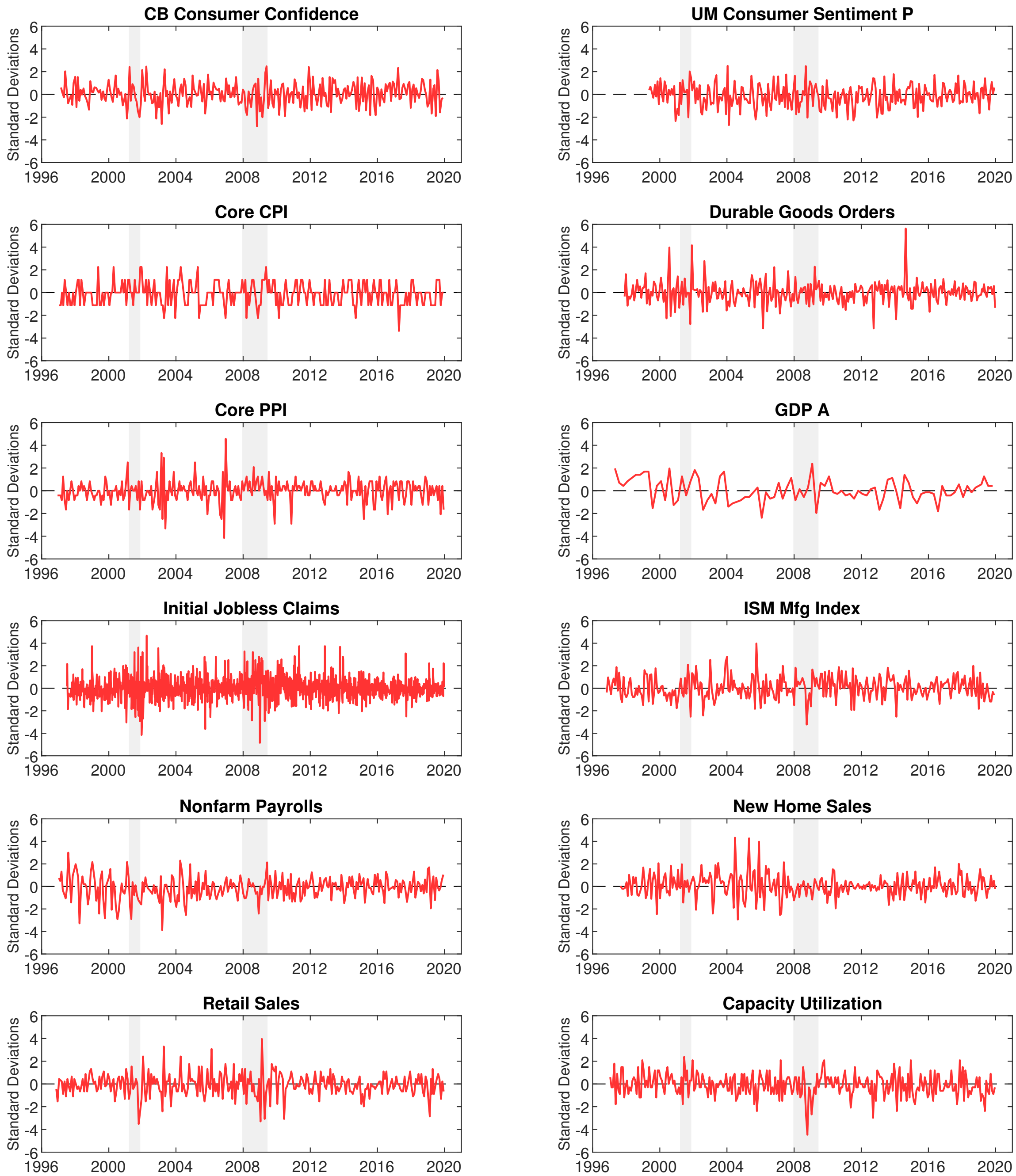
B.4 Overview of Data Usage

Main Text In Section 4, we employ news *surprises* for US releases (see Table B1), as well as the intraday data on international stock indexes (see Table 2). We also use S&P 500 futures and volatility indexes (see Table B5). In Section 5, we use the same financial data at lower frequencies except for the US, where we directly use the daily S&P 500 (see Table B6). We also substitute the volatility indexes with the daily versions from Bloomberg to extend the sample (see Table B6). In Section 6, we use *surprises* for foreign countries (see Table B2). We also employ the *reporting lag* and the *revision magnitude* for both US and foreign news releases. In Section 7, we use the US *surprises* again, as well as the *relevance* indexes to construct the daily series. Further, we use the daily financial market data discussed in Section B.3.

Appendices In Appendix C, we present multiple results which use various different series. If the data reference is not clear from the main text, the notes below the figure or table provide the data source. In Supplementary Appendix S1, we employ the commodity indexes (see Table B5), as well as the news *surprises* for US releases. In Supplementary Appendix S2, we use the news *surprises* for US releases and the daily data on Treasury yields (see Table B6). In Supplementary Appendix S3, we use additional data from *Thomson Reuters Tick History* which is detailed in the appendix. In Supplementary Appendix S4, we employ data to gauge the state of the US and foreign business cycles in addition to the US news *surprises*. Details on the data are provided there. In Supplementary Appendix S5, we employ news *surprises* for US releases, as well as the intraday data on international stock indexes and US dollar exchange rates (see Table B3). In Supplementary Appendix S6, we use, besides US news *surprises*, external sources for measures of cross-country linkages. Details on the data are provided in that appendix.

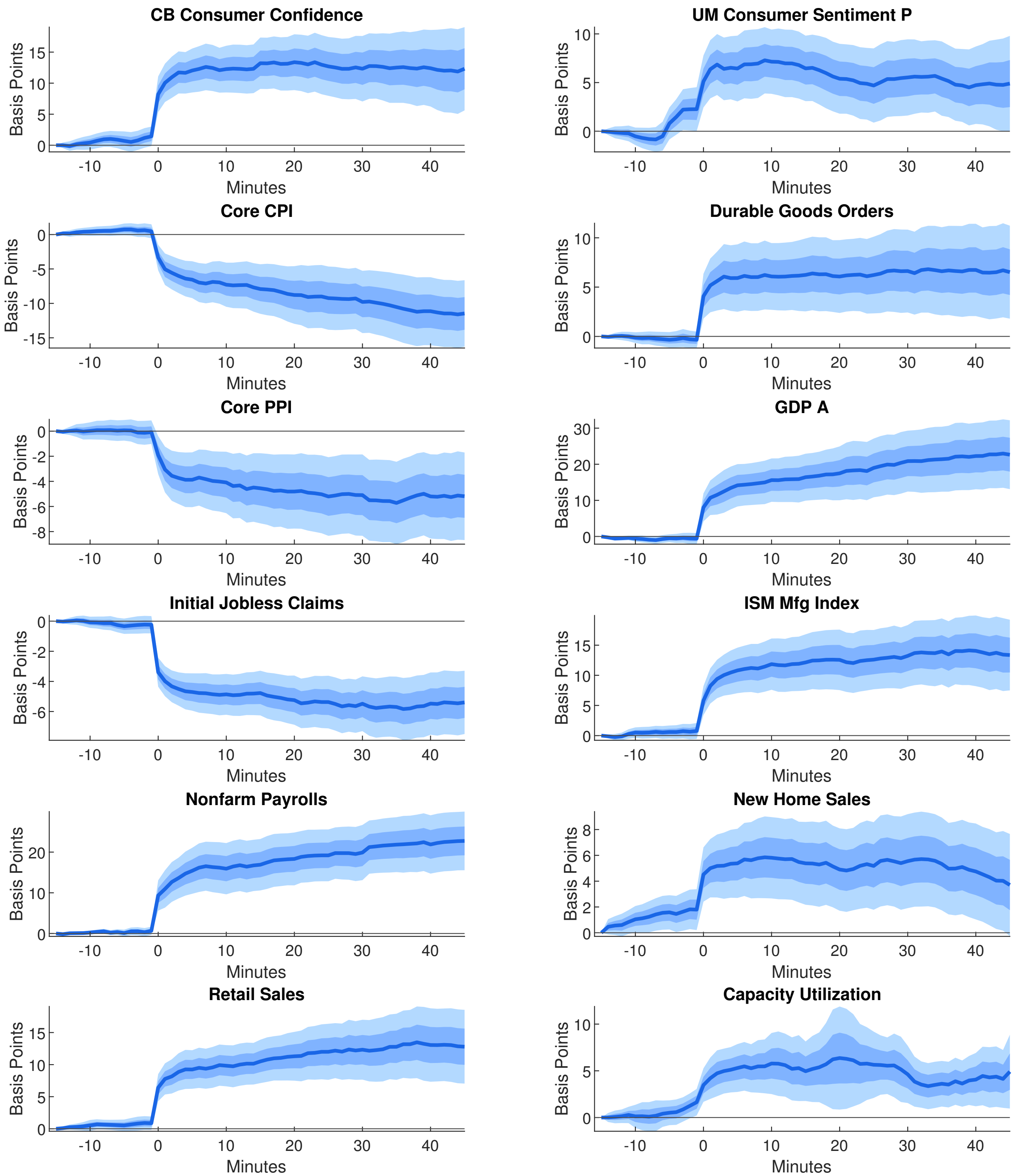
C Additional Results

Figure C1: Time Series of Standardized Surprises



Notes: This figure shows the standardized surprises for the 12 major macroeconomic series over the sample period. The construction follows equation (1) in the text. Shaded areas indicate NBER recession periods.

Figure C2: Impulse Response Functions for Major Announcements

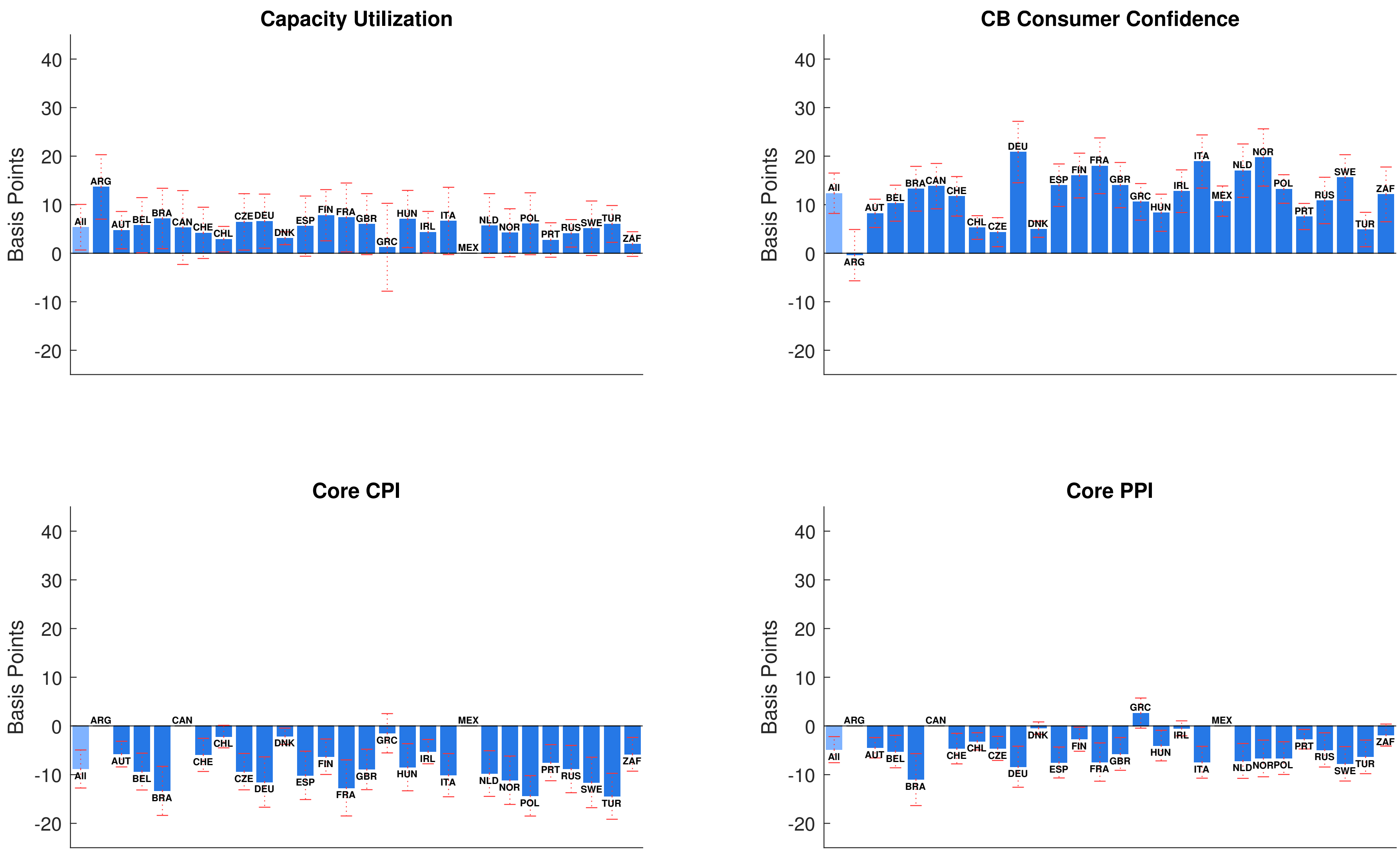


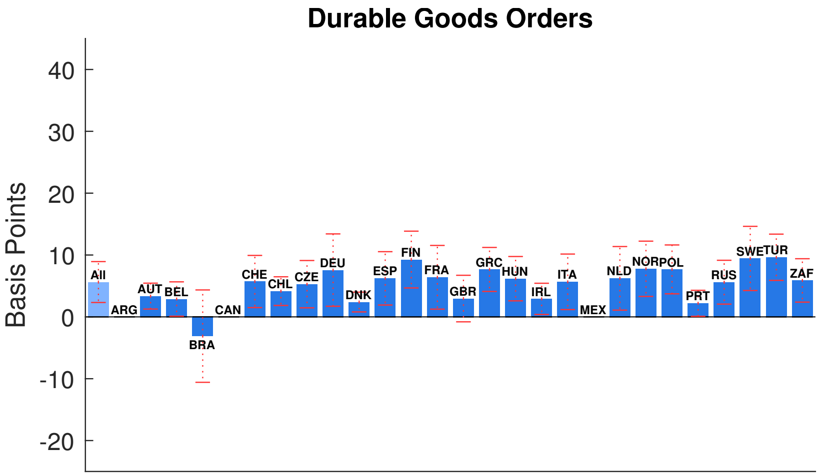
Notes: This figure displays impulse response functions for stock indexes over a 60-minute window for a given news release, estimated from specification

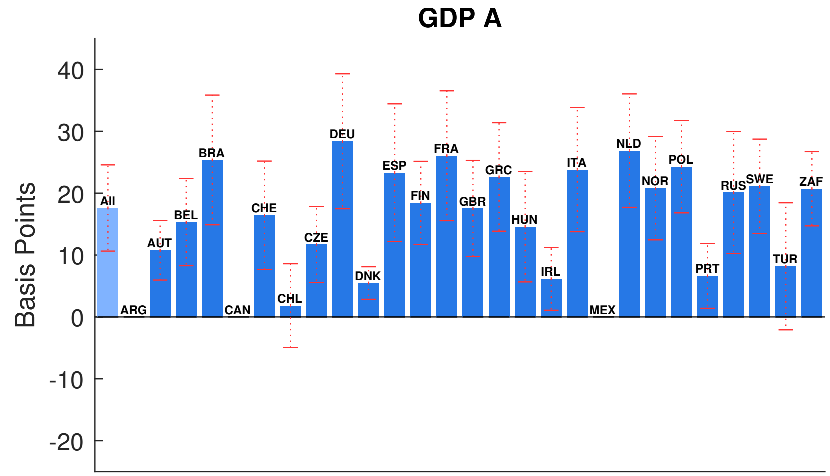
$$q_{i,t+h} - q_{i,t-15} = \alpha_i + \gamma^y s^y_{US,t} + \sum_{k \neq y} \gamma^k s^k_{US,t} + \varepsilon_{i,t},$$

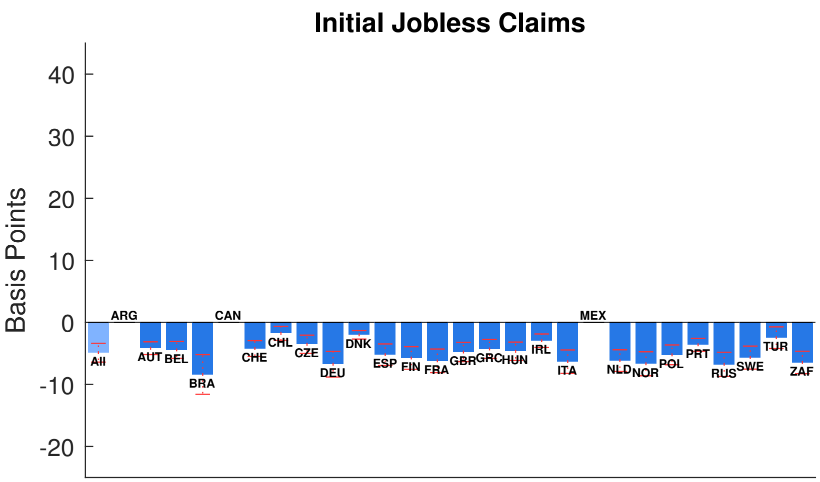
where $q_{i,t}$ is the log price index and $h = -14, ..., 45$. The stock index changes are expressed in basis points. The dark and light blue bands display the 68 percent and 95 percent confidence bands, respectively. Standard errors are two-way clustered by announcement and by country.

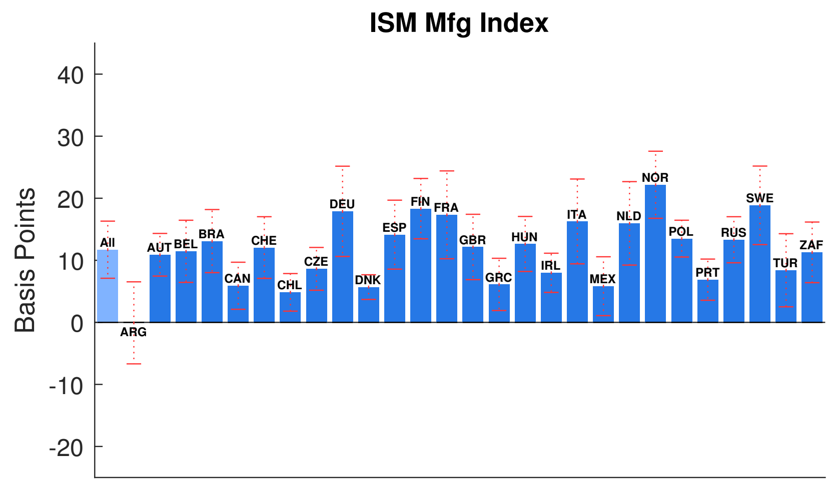
Figure C3: Effects of US News on International Stock Markets by Country

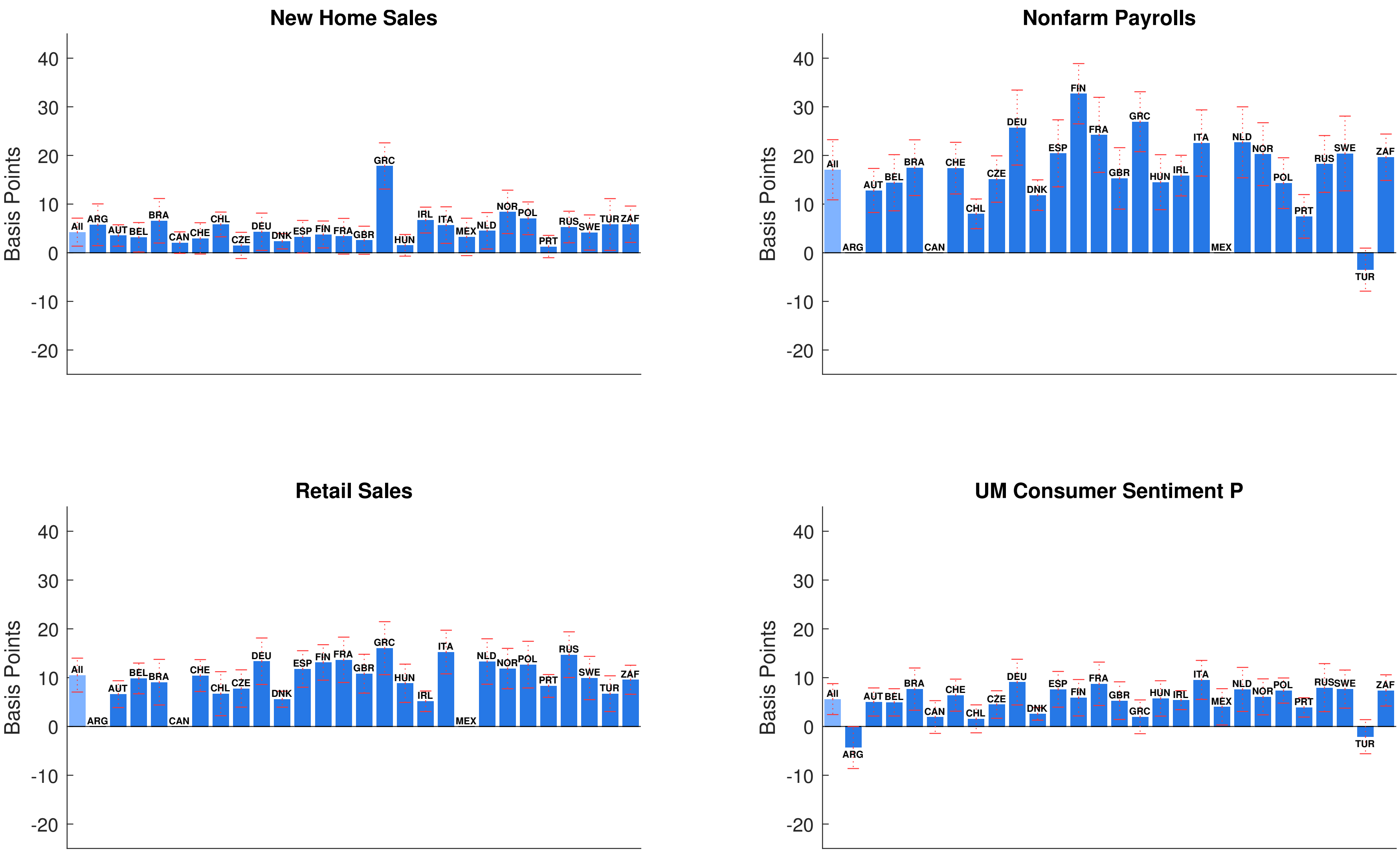



Durable Goods Orders
Basis Points
40
30
20
10
0
-10
-20
All ARG AUT BEL BRA CAN CHE CHL CZE DEU DNK ESP FIN FRA GBR GRC HUN IRL ITA MEX NLD NOR POL PRT RUS SWE TUR ZAF


GDP A
Basis Points
40
30
20
10
0
-10
-20
All ARG AUT BEL BRA CAN CHE CHL CZE DEU DNK ESP FIN FRA GBR GRC HUN IRL ITA MEX NLD NOR POL PRT RUS SWE TUR ZAF


Initial Jobless Claims
Basis Points
40
30
20
10
0
-10
-20
All ARG AUT BEL BRA CAN CHE CHL CZE DEU DNK ESP FIN FRA GBR GRC HUN IRL ITA MEX NLD NOR POL PRT RUS SWE TUR ZAF


ISM Mfg Index
Basis Points
40
30
20
10
0
-10
-20
All ARG AUT BEL BRA CAN CHE CHL CZE DEU DNK ESP FIN FRA GBR GRC HUN IRL ITA MEX NLD NOR POL PRT RUS SWE TUR ZAF



This figure shows the equity market responses for all releases. For a given announcement, the light blue bar represents the pooled effect, i.e., the estimate of common coefficient γ^y of equation (3), while the dark blue bars represent the country-specific effects, i.e., the estimates of γ_i^y obtained from estimating equation (4). Missing country bars indicate cases in which the country is dropped because it had fewer than 24 observations for a given announcement. The red error bands depict 95 percent confidence intervals, where standard errors are two-way clustered by announcement and by country.

Table C1: Effects of US News on Other Implied Volatility Indexes

	Capacity Utilization	CB Consumer Confidence	Core CPI	Core PPI	Durable Goods Orders	GDP A
VDAX (bp)						
News	-20.10***	-40.05***	35.66***	24.91**	-27.08***	-89.30***
	(6.68)	(8.84)	(11.56)	(10.71)	(9.78)	(14.40)
R^2	0.06	0.14	0.12	0.27	0.16	0.35
Observations	175	171	175	175	173	59
VCAC (bp)						
News	-33.28*	-33.38**	43.42*	7.56	-15.79	-54.12*
	(16.89)	(16.33)	(25.96)	(18.92)	(11.33)	(28.85)
R^2	0.06	0.08	0.08	0.20	0.13	0.15
Observations	146	145	146	146	145	49
VFTSE (bp)						
News	-22.74	-46.16***	3.02	-31.82	4.11	-106.77***
	(18.39)	(17.33)	(15.83)	(28.55)	(13.24)	(24.89)
R^2	0.02	0.15	0.03	0.07	0.17	0.47
Observations	128	121	124	124	126	41
	Initial Jobless Claims $\cdot(-1)$	ISM Mfg Index	New Home Sales	Nonfarm Payrolls	Retail Sales	UM Consumer Sentiment P
VDAX (bp)						
News	-24.19***	-85.38***	-33.69**	-137.03***	-49.86***	-45.76***
	(4.46)	(18.00)	(14.98)	(18.37)	(7.96)	(12.29)
R^2	0.13	0.23	0.11	0.28	0.26	0.10
Observations	751	162	173	171	175	176
VCAC (bp)						
News	-43.40***	-94.30***	-34.65	-149.67***	-59.67***	-21.42
	(11.74)	(21.66)	(24.28)	(26.55)	(19.14)	(26.10)
R^2	0.08	0.18	0.09	0.30	0.16	0.02
Observations	629	136	143	143	146	147
VFTSE (bp)						
News	-30.91***	-79.87***	-31.58	-59.98	-35.56	-71.54***
	(8.54)	(23.78)	(20.62)	(54.67)	(32.59)	(17.48)
R^2	0.12	0.18	0.06	0.09	0.09	0.11
Observations	541	112	122	121	122	124

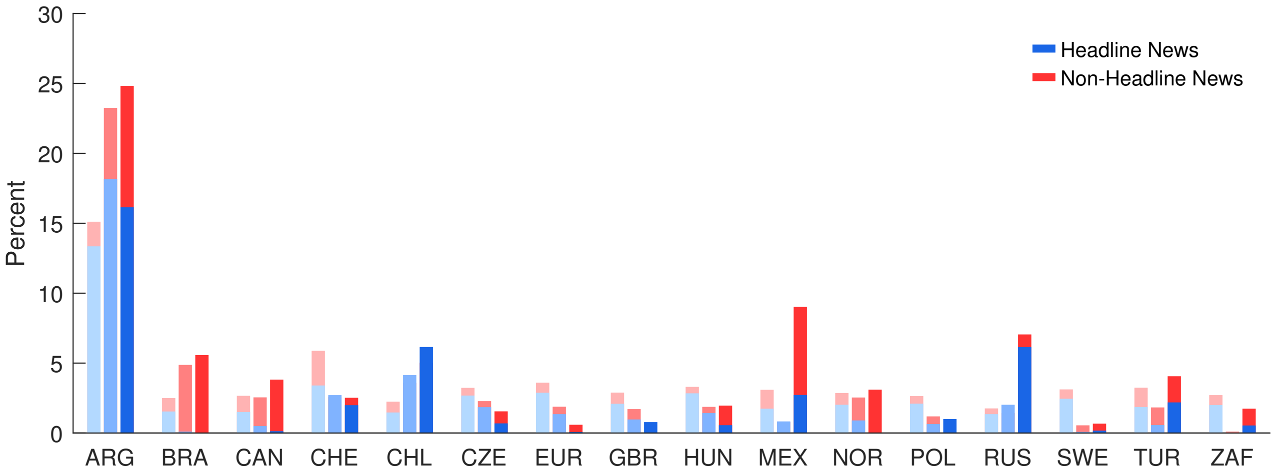
Notes: For all 12 announcements, this table shows estimates of γ^y obtained from equation (5), where the left-hand side is the 30-minute log-change in the VFTSE, the VDAX, or the VCAC. Heteroskedasticity-robust standard errors are reported in parentheses. ***, **, and * indicate significance at the 1, 5, and 10 percent level.

Table C2: Low Frequency Analysis—Stock Indexes

Coefficient $\beta_i^{(h)}$	USA	ARG	AUT	BEL	BRA	CAN	CHE	CHL	CZE	DEU	DNK	ESP	FIN	FRA
Headline News Index														
1-month	1.05	2.37	1.28	1.13	0.75	1.18	0.89	0.69	1.70	1.11	2.28	1.10	0.93	1.06
	(0.41)	(0.44)	(0.71)	(0.81)	(0.37)	(0.45)	(0.35)	(0.45)	(0.64)	(0.38)	(0.99)	(0.56)	(0.35)	(0.43)
1-quarter	2.10	3.06	2.37	1.78	2.35	2.15	1.52	0.81	2.95	2.02	4.04	2.36	2.09	1.96
	(0.56)	(0.53)	(0.95)	(1.03)	(0.63)	(0.76)	(0.46)	(0.44)	(0.92)	(0.73)	(1.71)	(0.82)	(0.52)	(0.67)
Broad News Index														
1-month	0.61	1.59	1.55	1.09	0.63	0.85	0.94	0.97	1.56	1.40	2.26	1.41	1.08	1.05
	(0.17)	(0.32)	(0.55)	(0.34)	(0.22)	(0.18)	(0.16)	(0.32)	(0.27)	(0.23)	(0.63)	(0.39)	(0.26)	(0.17)
1-quarter	1.27	2.75	2.88	1.60	1.75	1.89	1.38	1.33	3.16	1.94	3.66	2.47	2.00	1.65
	(0.25)	(0.28)	(0.72)	(0.37)	(0.27)	(0.35)	(0.20)	(0.50)	(0.38)	(0.32)	(0.81)	(0.49)	(0.41)	(0.21)
	GBR	GRC	HUN	IRL	ITA	MEX	NLD	NOR	POL	PRT	RUS	SWE	TUR	ZAF
Headline News Index														
1-month	0.82	1.40	1.53	1.46	0.94	0.92	1.29	0.62	1.61	0.84	0.58	0.86	1.38	0.58
	(0.44)	(0.54)	(0.66)	(0.76)	(0.55)	(0.46)	(0.51)	(0.57)	(0.47)	(0.69)	(0.35)	(0.42)	(0.61)	(0.35)
1-quarter	1.88	2.42	2.10	2.66	1.94	2.29	2.10	1.91	2.40	2.08	1.52	2.39	1.17	1.03
	(0.53)	(0.60)	(1.09)	(1.02)	(0.73)	(0.78)	(0.62)	(0.57)	(0.66)	(1.26)	(0.46)	(0.75)	(0.79)	(0.49)
Broad News Index														
1-month	0.95	1.65	1.71	1.28	1.28	0.52	1.31	1.02	1.65	1.00	0.90	1.13	1.60	0.48
	(0.20)	(0.36)	(0.46)	(0.39)	(0.34)	(0.31)	(0.25)	(0.50)	(0.32)	(0.33)	(0.40)	(0.28)	(0.46)	(0.28)
1-quarter	1.54	2.81	2.42	2.17	2.13	1.37	1.68	2.07	2.74	2.01	2.46	2.26	1.93	0.68
	(0.21)	(0.60)	(0.88)	(0.52)	(0.27)	(0.45)	(0.24)	(0.46)	(0.50)	(0.33)	(0.61)	(0.53)	(0.70)	(0.32)

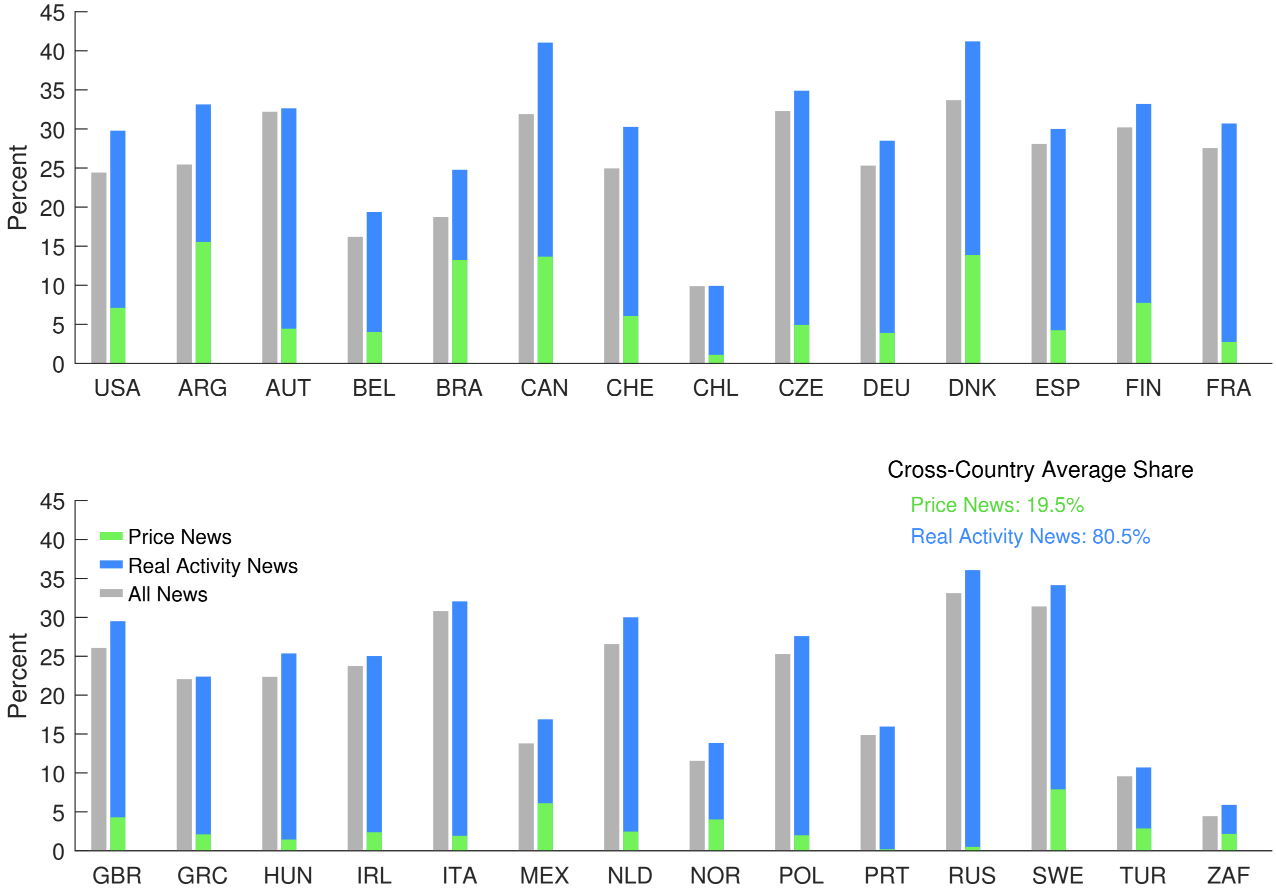
Notes: This table reports for each country the coefficients $\beta_i^{(h)}$ of equations (7) and (S2.3) for stock indexes at the monthly and quarterly frequency. The estimates of equation (7) are displayed under "Headline News Index" whereas results of equation (S2.3) are shown under "Broad News Index". The corresponding R-squared are illustrated in Figure 5. The sample ranges from January 1, 2000 to December 31, 2019. Newey-West standard errors are reported in parentheses. For the US, we use the S&P 500. Daily data on the S&P 500 is obtained from the Center of Research in Security Prices (CRSP).

Figure C4: Daily, Monthly, and Quarterly R-Squared for US Dollar Exchange Rates



Notes: For each US dollar-denominated exchange rate, this figure plots the R-squared of equations (6) and (S2.2) for the daily frequency, and the R-squared of equations (7) and (S2.3) for the monthly and quarterly frequency. The left, middle, and right bars indicate the R-squared of the daily, monthly, and quarterly regression, respectively. For a given country and frequency, the blue bar represents the R-squared of the headline surprises of US macroeconomic news, whereas the red bar displays the increment in R-squared once non-headline news is included. The sample runs from January 1, 2000 to December 31, 2019.

Figure C5: Quarterly R-Squared for Stock Indexes—Price vs. Real Activity



Notes: For each country's stock index, this figure plots the quarterly R-squared as shown in Figure 5 in grey, as well as a decomposition into the relative contributions of price (green) and real activity news (blue). These are constructed by calculating the fitted values of the daily regression separately for price and real activity news using the estimates from the baseline analysis. While the daily fitted values are orthogonal to one another, those at the monthly and quarterly frequency need not be. Indeed, the combined explanatory power of price and real activity news is larger then the total, indicating that there is overlapping information in the two categories. The sample runs from January 1, 2000 to December 31, 2019. Appendix Table B1 provides an overview of the news releases and their classification into the two groups.

Table C3: Effects of Foreign News on US Dollar Exchange Rates

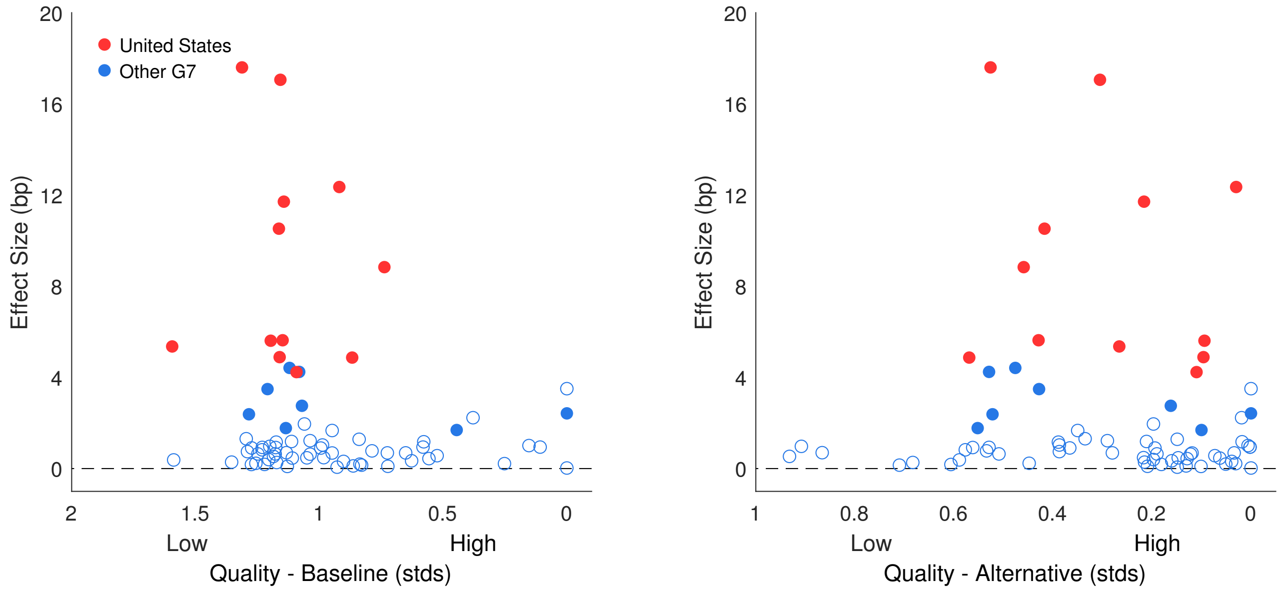
Canada	Capacity Utilization	Core CPI	GDP	Housing Starts	Intl. Trade	IPPI	Mfg Sales	PMI	Retail Sales	Unemployment Rate
Exchange Rate (bp)										
News	1.02	9.06***	10.43***	2.09**	9.70***	1.42	3.77***	8.23***	6.10***	-7.21***
	(1.82)	(1.70)	(2.27)	(0.89)	(1.69)	(1.08)	(0.86)	(1.30)	(1.73)	(1.65)
Observations	79	225	81	231	270	253	272	193	265	274
France	BoF Industry Sentiment	Consumer Confidence	CPI P	GDP P	Industrial Production	Mfg Confidence	PPI	Production Outlook	Trade Balance	Unemployment Rate
Exchange Rate (bp)										
News	1.35*	2.39*	0.51	2.11	-0.33	-0.54	0.55	0.41	0.66	-0.78
	(0.77)	(1.30)	(0.73)	(1.37)	(0.61)	(0.77)	(0.98)	(1.08)	(0.63)	(0.77)
Observations	135	237	258	89	268	217	158	185	268	173
Germany	CPI P	GDP	GfK Consumer Confidence	IFO Business Climate	Industrial Production	PPI	Retail Sales	Trade Balance	Unemployment Change	ZEW Survey Expectations
Exchange Rate (bp)										
News	1.90	6.10***	-0.29	8.65***	1.70***	-0.03	1.56***	1.57**	0.70	3.31***
	(1.28)	(1.04)	(0.83)	(1.17)	(0.59)	(0.58)	(0.58)	(0.72)	(0.90)	(0.75)
Observations	242	89	159	269	267	274	255	273	274	213
Italy	Consumer Confidence	CPI P	GDP F	Industrial Production	Industrial Sales	Mfg Confidence	PPI	Trade Balance	Retail Sales	Unemployment Rate
Exchange Rate (bp)										
News	0.11	0.52	1.34*	0.14	3.70*	0.08	0.71	-0.40	0.26	-0.22
	(0.63)	(0.72)	(0.71)	(0.83)	(2.18)	(0.96)	(1.03)	(1.53)	(0.71)	(1.02)
Observations	221	256	78	246	63	233	189	75	173	145
Japan	BoJ Mfg Index	BoJ Mfg Outlook	Consumer Confidence	CPI	Exports	GDP P	Industrial Production P	PPI	Retail Sales	Unemployment Rate
Exchange Rate (bp)										
News	3.17**	5.52***	-0.34	0.34	0.53	3.40	1.56**	0.11	0.34	-1.81**
	(1.52)	(1.96)	(0.55)	(0.92)	(0.78)	(2.33)	(0.78)	(0.63)	(0.61)	(0.80)
Observations	84	60	153	215	130	89	237	230	199	234
United Kingdom	Core CPI	Core PPI	Exports	GDP A	GfK Consumer Confidence	House Price Index	Industrial Production	Jobless Claims	Retail Sales	Unemployment Rate
Exchange Rate (bp)										
News	10.91***	0.60	-0.10	20.19***	0.78*	3.47***	2.70**	-3.15*	12.94***	-6.09***
	(1.63)	(1.68)	(1.98)	(3.40)	(0.41)	(1.22)	(1.09)	(1.66)	(1.65)	(1.22)
Observations	172	168	59	86	205	186	273	239	118	211

Notes: The table presents the response of the US dollar exchange to foreign macroeconomic news releases. For each non-US G7 country, this table shows estimates of ζ^y obtained from specification

$$\Delta q_{US,t} = \alpha_i + \zeta_i^y s_{i,t}^y + \sum_{k \neq y} \zeta_i^k s_{i,t}^k + \sum_{w} \zeta_{US}^w s_{US,t}^w + \varepsilon_{i,t},$$

where $s_{i,t}^y$ is the surprise of interest, $s_{i,t}^k$ and $s_{US,t}^w$ are other surprises of country i and the US released within the same time window, and $\Delta q_{US,t}$ is the 30-minute change of country i's US dollar denominated exchange rate. Exchange rates are expressed in US dollars so that an increase reflects a depreciation of the US dollar relative to the foreign currency. The units are in basis points. Heteroskedasticity-robust standard errors reported in parentheses. ***, **, and * indicate significance at the 1, 5, and 10 percent level.

Figure C6: Relation of Effect Size to Quality of Releases—Robustness



Notes: This Figure shows how the effect size of a release relates to its quality. The left panel shows the relationship when quality is proxied by the revision magnitude as defined in equation (9). The right panel shows the relationship with an alternative measure of quality, which is defined as $\frac{1}{N_i^y}\sum_{n=1}^{N_i^y}\frac{|y_{i,n}^F - y_{i,n}|}{\sigma_{y_{i,n}^F}}$, where $y_{i,n}$ and $y_{i,n}^F$ denote the initial and final revised value of release n, $\sigma_{y_{i,n}^F}$ denotes the standard deviation of the final revised time series, and N^y denotes the total number of announcements for series y in our sample. For US releases (red) the effect size corresponds to the absolute value of the coefficients shown in Table 3. For the foreign releases (blue), the coefficients in Table 6 are used. Filled circles indicate significance at the 10 percent level.

Table C4: Effects of US News on 1-Year Bond Yield

	Capacity Utilization	CB Consumer Confidence	Core CPI	Core PPI	Durable Goods Orders	GDP A
1-Year Bond Yield (bp)						
News	0.05	0.10	-0.05	0.20**	0.07	0.31**
	(0.07)	(0.20)	(0.12)	(0.08)	(0.10)	(0.13)
R^2	0.02	0.02	0.01	0.03	0.01	0.04
Observations	1894	1916	1916	1935	1884	584
	Initial Jobless Claims $\cdot(-1)$	ISM Mfg Index	New Home Sales	Nonfarm Payrolls	Retail Sales	UM Consumer Sentiment P
1-Year Bond Yield (bp)						
News	0.27***	0.37***	-0.40	1.13***	-0.03	0.13
	(0.07)	(0.09)	(0.37)	(0.20)	(0.10)	(0.13)
R^2	0.01	0.05	0.10	0.05	0.02	0.04
Observations	8468	1844	1888	2005	1951	1899

Notes: This table presents estimates of γ^y of equation (3) for each of the 12 macroeconomic announcements. The units are in basis points. Standard errors are clustered by announcement and reported in parentheses. ***, **, and * indicate significance at the 1, 5, and 10 percent level.

Table C5: Effects of US News on US Yield Curve

	Capacity Utilization	CB Consumer Confidence	Core CPI	Core PPI	Durable Goods Orders	GDP A
1-Q Eurodollar Rate (bp)						
News	0.23***	0.50***	0.50***	0.41***	0.21***	0.61***
	(0.05)	(0.17)	(0.09)	(0.07)	(0.08)	(0.15)
R^2	0.08	0.15	0.18	0.21	0.18	0.23
Observations	231	239	258	261	256	89
4-Q Eurodollar Rate (bp)						
News	0.52***	1.18***	1.48***	1.02***	0.68***	1.65***
	(0.12)	(0.22)	(0.24)	(0.18)	(0.24)	(0.37)
R^2	0.10	0.27	0.22	0.33	0.22	0.32
Observations	263	259	267	274	260	88
2-Y Treasury Yield (bp)						
News	0.46***	0.96***	1.19***	0.80***	0.57***	1.42***
	(0.10)	(0.21)	(0.22)	(0.14)	(0.20)	(0.32)
R^2	0.13	0.24	0.20	0.36	0.21	0.33
Observations	244	240	265	270	253	89
10-Y Treasury Yield (bp)						
News	0.45***	1.15***	1.31***	0.98***	0.44*	1.56***
	(0.10)	(0.17)	(0.23)	(0.15)	(0.26)	(0.34)
R^2	0.09	0.37	0.22	0.36	0.25	0.30
Observations	270	195	264	274	187	90
	Initial Jobless Claims $\cdot(-1)$	ISM Mfg Index	New Home Sales	Nonfarm Payrolls	Retail Sales	UM Consumer Sentiment P
1-Q Eurodollar Rate (bp)						
News	0.27***	0.69***	0.21***	1.54***	0.46***	0.23***
	(0.04)	(0.08)	(0.06)	(0.17)	(0.10)	(0.07)
R^2	0.12	0.32	0.14	0.37	0.23	0.07
Observations	1108	259	243	273	263	227
4-Q Eurodollar Rate (bp)						
News	0.66***	2.01***	0.77***	4.71***	1.37***	0.63***
	(0.07)	(0.23)	(0.15)	(0.51)	(0.24)	(0.12)
R^2	0.22	0.36	0.25	0.45	0.29	0.12
Observations	1146	268	259	274	271	242
2-Y Treasury Yield (bp)						
News	0.58***	1.79***	0.64***	4.15***	1.23***	0.50***
	(0.07)	(0.21)	(0.12)	(0.44)	(0.18)	(0.11)
R^2	0.23	0.40	0.25	0.47	0.33	0.10
Observations	1111	249	239	270	268	233
10-Y Treasury Yield (bp)						
News	0.59***	2.14***	0.73***	4.18***	1.46***	0.60***
	(0.07)	(0.18)	(0.13)	(0.42)	(0.21)	(0.12)
R^2	0.22	0.47	0.27	0.46	0.37	0.13
Observations	1025	273	190	274	271	243

Notes: For all 12 announcements, this table shows estimates of γ^y obtained from the following specification:

$$\Delta q_t = \alpha + \gamma^y s^y_{US,t} + \sum_{k \neq y} \gamma^k s^k_{US,t} + \varepsilon_t,$$

where $s^y_{US,t}$ is the surprise of interest, $s^k_{US,t}$ are other surprises released in the same time window, and Δq_t is the 30-minute change in the yield of interest. The dependent variables are constructed as in Gürkaynak, Kısacıkoğlu, and Wright (2020). See Boehm and Kroner (2021) for more details on this. The units of the dependent variables are in basis points. Heteroskedasticity-robust standard errors are reported in parentheses. ***, **, and * indicate significance at the 1, 5, and 10 percent level.

References

Adrian, Tobias, Richard K Crump, and Emanuel Moench. 2013. "Pricing the term structure with linear regressions." *Journal of Financial Economics* 110 (1):110–138.

Beaudry, Paul and Franck Portier. 2006. "Stock prices, news, and economic fluctuations." *American Economic Review* 96 (4):1293–1307.

Beechey, Meredith J and Jonathan H Wright. 2009. "The high-frequency impact of news on long-term yields and forward rates: Is it real?" *Journal of Monetary Economics* 56 (4):535–544.

Boehm, Christoph and Niklas Kroner. 2021. "Beyond the Yield Curve: Understanding the Effect of FOMC Announcements on the Stock Market." *Available at SSRN 3812524* .

Faust, Jon, John H. Rogers, Shing-Yi B. Wang, and Jonathan H. Wright. 2007. "The high-frequency response of exchange rates and interest rates to macroeconomic announcements." *Journal of Monetary Economics* 54 (4):1051 – 1068.

Gormsen, Niels Joachim and Ralph SJ Koijen. 2020. "Coronavirus: Impact on stock prices and growth expectations." *The Review of Asset Pricing Studies* 10 (4):574–597.

Gürkaynak, Refet S, Burçin Kısacıkoğlu, and Jonathan H Wright. 2020. "Missing Events in Event Studies: Identifying the Effects of Partially Measured News Surprises." *American Economic Review* 110 (12):3871–3912.

Gürkaynak, Refet S, Brian Sack, and Jonathan H Wright. 2007. "The US Treasury yield curve: 1961 to the present." *Journal of monetary Economics* 54 (8):2291–2304.

Knox, Benjamin and Annette Vissing-Jorgensen. 2022. "A stock return decomposition using observables." .

Martin, Ian. 2017. "What is the Expected Return on the Market?" *The Quarterly Journal of Economics* 132 (1):367–433.

Martin, Ian WR and Christian Wagner. 2019. "What is the Expected Return on a Stock?" *The Journal of Finance* 74 (4):1887–1929.

Supplementary Appendix
for
The US, Economic News, and the Global Financial Cycle*

February 15, 2023

Christoph E. Boehm
UT Austin and NBER

T. Niklas Kroner
Federal Reserve Board

Table of Contents

S1 Commodity Prices 2

S2 Non-Headline News 4

S2.1 Factor Estimation 4

S2.2 Explanatory Power of Headline and Non-Headline News 5

S2.3 Alternative Specifications 6

S3 Monetary Policy Analysis 9

S3.1 Construction of Shocks 9

S3.2 Additional Results 13

S3.3 Robustness 13

S4 State-Dependent Effects of US Macro News 17

S5 The Role of the US Dollar Exchange Rate 21

S6 Inspecting the Cross-Sectional Heterogeneity 23

References 28

*The views expressed are those of the authors and do not necessarily reflect those of the Federal Reserve Board or the Federal Reserve System.
Email: chris.e.boehm@gmail.com and t.niklas.kroner@gmail.com.

S1 Commodity Prices

To ensure that our results hold for a large set of risky asset prices, we study the effect of US macro news on commodity prices in this appendix. Gorton and Rouwenhorst (2006) show that commodities and equities have similar return profiles. Bastourre et al. (2012) and Etula (2013) emphasize the relationship of commodity prices and risk appetite. In our analysis, we focus on three commodity classes: energy, agriculture, and industrial metals and measure them using the corresponding S&P GS commodity sector indexes.[1] Table S1.1 provides additional information on the three indexes.

As documented by prior research, commodity prices co-move over time, and can be summarized by common factors (Pindyck and Rotemberg, 1990; Byrne, Fazio, and Fiess, 2013; Alquist, Bhattarai, and Coibion, 2019). Bastourre et al. (2012) find that such a commodity factor is also informative about global risk-taking capacity. We follow this literature and use principal component analysis on the 30-minute log-changes in the commodity indexes around the 12 macroeconomic announcements of interest. Table S1.2 summarizes the results. The first common factor explains around 55 percent of the variation, and loads with the same sign on all three commodity indexes. Hence, this factor captures the co-movement of commodity prices. The second factor, which explains 30 percent of the variation, loads positively on agricultural commodities, and negatively on energy commodities and industrial metals. This factor primarily explains variation of the agricultural index and is relatively unimportant for energy and industrial metals.

We proceed with studying the effects of US news on the first common factor within a 30-minute window of the release. Table S1.3 shows the results. For the majority of news releases, we find a significant effect on the factor. Further, the signs are as expected. Positive (negative) news about real activity leads to an increase (decrease) in commodity prices. Our results are in line with Kurov and Stan (2018), but differ somewhat from Kilian and Vega (2011). The former paper finds significant effects of macroeconomic news on energy prices using intraday data similar to us, whereas the latter, employing daily data, does not find significant effects.

Table S1.1: Compositions of Commodity Indexes

Energy		Industrial Metals		Agriculture	
WTI Crude Oil	0.41	LME Aluminium	0.35	Chicago Wheat	0.18
Brent Crude Oil	0.30	LME Copper	0.41	Kansas Wheat	0.08
RBOB Gasoline	0.07	LME Lead	0.06	Corn	0.31
Heating Oil	0.07	LME Nickel	0.08	Soybeans	0.20
Gasoil	0.10	LME Zinc	0.11	Cotton	0.08
Natural Gas	0.05			Sugar	0.10
				Coffee	0.04
				Cocoa	0.02

Notes: This table shows the underlying commodity prices and corresponding weights for each of the three S&P GS commodity indexes.

[1] Following the previous literature, we exclude precious metals as they behave differently compared to other commodities (Chinn and Coibion, 2014). We also exclude livestock commodities since intraday data is not available to us for early-morning (8:30 ET) announcements from 2014 onwards.

Table S1.2: Results of Principal Component Analysis

	Loadings		Explained Variance		
	Factor 1	Factor 2	Factor 1	Factor 2	Total
Energy	0.65	-0.27	0.70	0.07	0.76
Industrial Metals	0.65	-0.27	0.70	0.07	0.76
Agriculture	0.39	0.92	0.25	0.75	1.00
Total			0.55	0.30	0.84

Notes: This table shows the loadings and explained variance of the first two factors of the commodity data. They are estimated using principal components on 30-minute changes of the S&P GS energy, industrial metals, and agriculture commodity index around the 12 macroeconomic announcements.

Table S1.3: Effects of US News on Commodity Prices

	Capacity Utilization	CB Consumer Confidence	Core CPI	Core PPI	Durable Goods Orders	GDP A
Commodity Factor (bp)						
News	0.72	18.12***	-3.75	-1.58	6.90*	24.34**
	(3.87)	(4.80)	(3.70)	(2.99)	(3.57)	(11.01)
R^2	0.00	0.15	0.12	0.11	0.17	0.31
Observations	152	151	151	152	151	50
	Initial Jobless Claims $\cdot(-1)$	ISM Mfg Index	New Home Sales	Nonfarm Payrolls	Retail Sales	UM Consumer Sentiment P
Commodity Factor (bp)						
News	7.15***	15.63***	11.66**	38.42***	15.15***	0.37
	(1.74)	(4.29)	(4.64)	(8.68)	(3.20)	(4.11)
R^2	0.12	0.23	0.11	0.24	0.24	0.01
Observations	658	151	151	148	151	152

Notes: For all 12 announcements, this table shows estimates of γ^y obtained from the following specificaton:

$$\Delta q_t = \alpha + \gamma^y s^y_{US,t} + \sum_{k \neq y} \gamma^k s^k_{US,t} + \varepsilon_t,$$

where $s^y_{US,t}$ is the surprise of interest, $s^k_{US,t}$ are other surprises released within the same time window, and $\Delta q_t = q_{t+20} - q_{t-10}$ is the 30-minute log-change in the commodity factor estimated from 30-minute changes in the energy, industrial metals, and agriculture commodities. See text and Supplementary Appendix Table S1.2 for details on the construction of the factor. Heteroskedasticity-robust standard errors are reported in parentheses. ***, **, and * indicate significance at the 1, 5, and 10 percent level.

S2 Non-Headline News

In this section, we provide details on the estimation of the non-headline factors used in Section 5. We also show that the key finding, which is that these non-headline factors increase the explanatory power for international stock markets, is robust to different specification choices.

S2.1 Factor Estimation

Gürkaynak, Kısacıkoğlu, and Wright (2020) argue that macro announcements elicit effects on asset prices beyond the headline variables, which are measured through surveys. Following their approach, we estimate these effects for the twelve major announcements $l \in L$, which we focus on in our paper. Specifically, we estimate the following specification

$$\Delta i_{US,d} = \alpha + \sum_k \beta^k s^k_{US,d} + \sum_l d^l_d \gamma^l f^l_{US,d} + \varepsilon_d, \tag{S2.1}$$

where ε_d is i.i.d. normal with mean zero and diagonal variance-covariance matrix. The factors $\{f^l_{US,d}\}^L_{l=1}$ are all standard normal and independent over time and of one another, d^l_d is a dummy that is 1 if announcement l occurs on day d. On the left-hand side, we use a vector of daily changes in two-, five-, and ten-year US yields, i.e., $\Delta i_{US,d} \equiv \{\Delta i^2_{US,d}, \Delta i^5_{US,d}, \Delta i^{10}_{US,d}\}$, as used by Gürkaynak, Kısacıkoğlu, and Wright (2020) in their lower frequency analysis.[2]

The latent factor $f^l_{US,d}$ captures the effects of announcement l beyond the surprise $s^l_{US,d}$ in the headline variable. Note that as some macroeconomic series come out simultaneously—for example, nonfarm payrolls is released jointly with other numbers such as the unemployment rate—an estimated latent factor can complement more than one headline surprise. While one could in principle incorporate non-headline news for all announcements, we restrict ourselves to these twelve as Gürkaynak, Kısacıkoğlu, and Wright (2020) show that the latent factors are well identified for major macro announcements as we consider them here. That being said, we consider alternative specifications in the next section and show that our baseline results are generally robust.

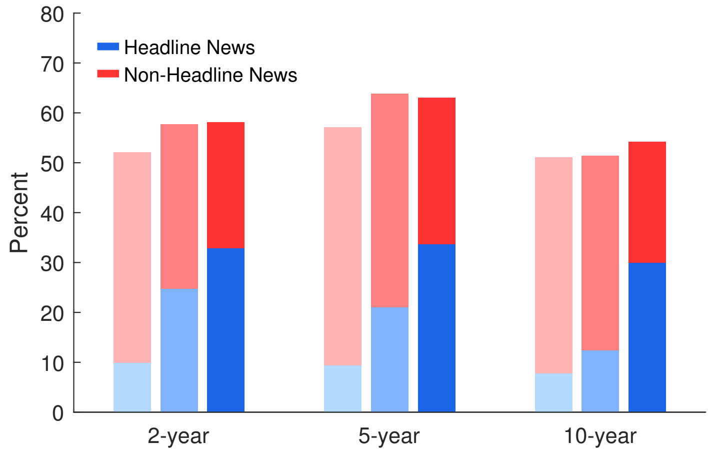
We estimate equation (S2.1) via the Kalman filter approach of Gürkaynak, Kısacıkoğlu, and Wright (2020). As a in "traditional" heteroskedasiticty identification (e.g., Rigobon, 2003), the latent factors are estimated by exploiting the difference in the variances on announcement and non-announcement days after taking out the variation attributable to the headline surprises.[3] While the sample period and the set of announcements differ, our results are similar to Gürkaynak, Kısacıkoğlu, and Wright (2020) in that the latent factors explain almost all of the remaining variation in yields on announcement days (not reported). In the following, we report the overall explanatory power (that is, for announcement and non-announcement days) for the US yield curve. To do so, we estimate versions of equations (S2.2) and (S2.3) below with the daily yield changes $\Delta i_{US,d}$ on the left hand side. While we estimate our factors, i.e., equation (S2.1), for a extended sample starting in 1997, we implement this exercise for the same sample as in Section 5, i.e., starting in 2000.

Figure S2.1 shows the results of this analysis. US headline macro news has increasing explanatory power for the US yield curve at lower frequencies consistent with the findings by Altavilla, Giannone, and Modugno (2017). Comparing our results to Gürkaynak, Kısacıkoğlu, and Wright (2020, Table

[2]Following Gürkaynak, Kısacıkoğlu, and Wright (2020), we use the daily zero coupon yields from Gürkaynak, Sack, and Wright (2007) for this exercise.

[3]To mitigate complications arising from monetary policy, we exclude days of FOMC releases in our set of announcement and non-announcement days.

Figure S2.1: Daily, Monthly, and Quarterly R-Squared for US Treasury Yields



Notes: This figure plots the R-squared of equations (6) for the daily frequency, and the R-squared of equations (7) and (S2.3) for the monthly and quarterly frequency, where we now use two-, five-, and ten-year US Treasury yields instead of country i's stock index. The left, middle, and right bar indicates the R-squared of the daily, monthly, and quarterly regression, respectively. For a given country and frequency, the blue bar represents the R-squared of the headline surprises of US macroeconomic news, whereas the red bar displays the increment in R-squared once non-headline news is included. The sample runs from January 1, 2000 to December 31, 2019.

15), we see that the results are very similar. They also find that while non-headline news increases the explanatory power substantially at lower frequencies, the relative contribution decreases. The total explanatory power is somewhat higher in our case. This mostly comes the fact that we consider a broader set of headline announcements, resulting in higher explanatory power of headline news. Overall, our findings are consistent with previous results in the literature.

S2.2 Explanatory Power of Headline and Non-Headline News

We use the latent factors for our explanatory power estimates in Section 5. To do so, we estimate the following specification:

$$\Delta q_{i,d} = \alpha_i + \sum_k \beta_i^k s_{US,d}^k + \sum_l \gamma_i^l f_{US,d}^l + \varepsilon_{i,d}, \tag{S2.2}$$

where $f_{US,d}^l$ is the latent non-headline news factor of major announcement l, estimated from equation (S2.1) above. Based on equation (S2.2), we define the daily broad news index $bni_{i,d}$ as the fitted value, and aggregate it to the desired time horizon h (in days), $bni_{i,d}^{(h)} = \sum_{j=0}^{h-1} bni_{i,d-j}$. Analogous to the procedure for headline news, we then calculate the R-squared of specification

$$\Delta q_{i,d}^{(h)} = \alpha_i^{(h)} + \beta_i^{(h)} bni_{i,d}^{(h)} + \varepsilon_{i,d}^{(h)} \tag{S2.3}$$

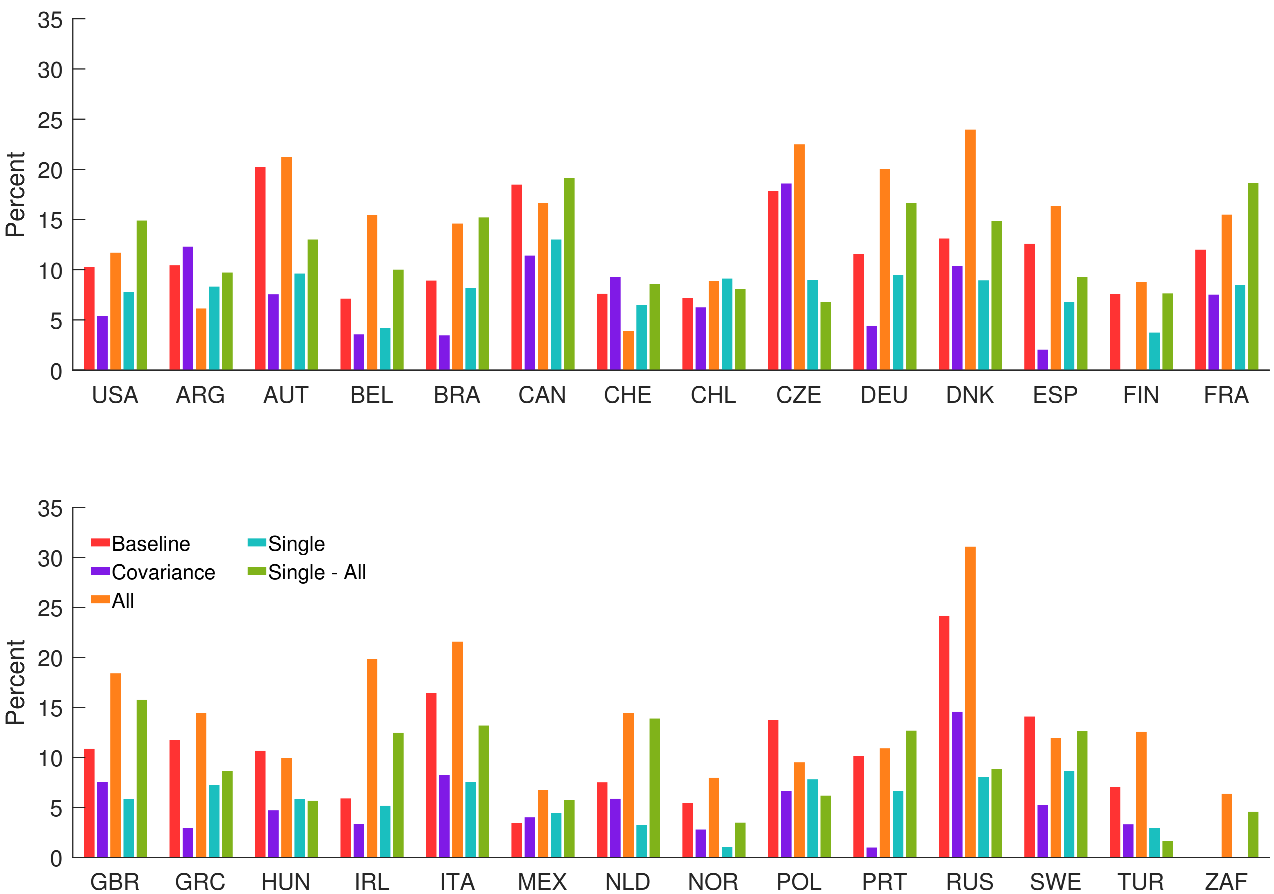
to measure the joint explanatory power of headline and non-headline news at the monthly and quarterly frequency. Note the red bars in Figure 5, which display the R-squared of non-headline news, are estimated as the difference in R-squared values of equations (S2.2) and (6) for the daily frequency, and the difference in R-squared values of equations (S2.3) and (7) for the monthly and quarterly frequency.

S2.3 Alternative Specifications

In this section, we look at alternative ways of estimating non-headline news and compare the results with the baseline specification. In particular, we do this by repeating the explanatory exercise in Section 5 for each specification. Figure S2.2 shows the comparison for the stock indexes, and Figure S2.3 for the volatility and commodity indexes. In what follows, we go over each alternative specification as well as the corresponding results, and discuss how they compare to the baseline.

In the first one, labeled as *covariance* in Figure S2.2 and S2.3, we follow the robustness check of Gürkaynak, Kısacıkoğlu, and Wright (2020) and allow for an unrestricted variance-covariance matrix of ε_d in equation (S2.1). This specification allows for the possibility of ever-present factors, i.e., drivers which lead to systematic movements on announcement and non-announcement days. Looking at Figure S2.2, the explanatory power falls for some countries compared to the baseline, while it increases for others. On average, the specification finds a smaller role for non-headline news which is broadly consistent with the findings by Gürkaynak, Kısacıkoğlu, and Wright (2020). Similar conclusions can be drawn from Figure S2.3.

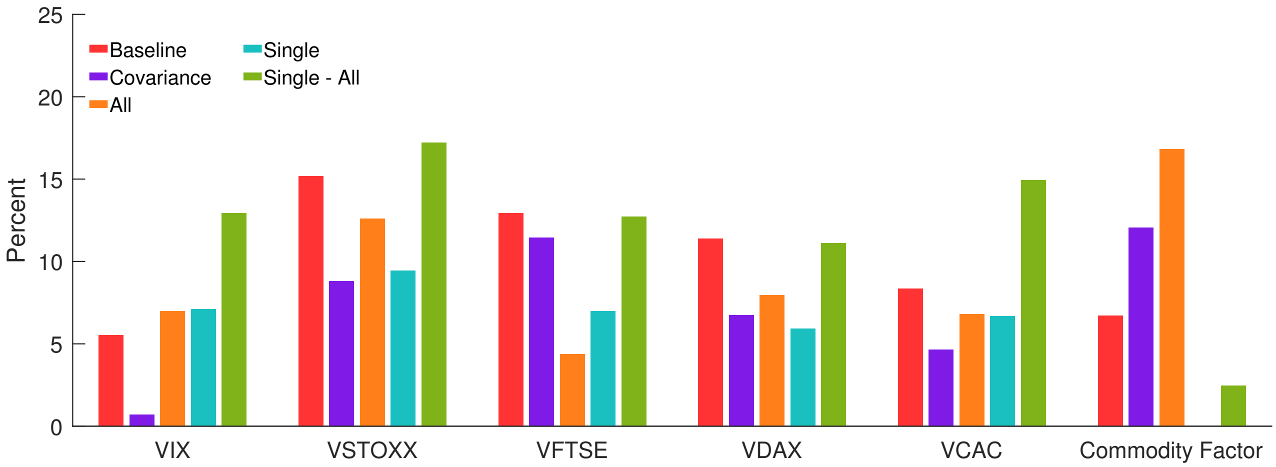
Figure S2.2: Quarterly R-Squared of Non-Headline News for Stock Indexes



Notes: For each country's stock index, this figure plots the increment in R-squared of non-headline news for the quarterly frequency. The red bars (most left) correspond to the *Baseline* specification and are the same as the red bars in Figure 5. The purple, orange, turquoise, and green bars depict alternative specifications *Covariance*, *All*, *Single*, and *Single—All*, respectively. These specifications are explained in Supplementary Appendix S2.3. The sample runs from January 1, 2000 to December 31, 2019.

In the second specification, labeled as *all* in Figure S2.2 and S2.3, we estimate equation (S2.1) with non-headline factors for all 66 announcement series. As some series are released jointly, we end up with 45 factors. As expected, this leads to an increase in the explanatory power in the vast majority of outcomes we consider. Note that an reduction in the R-squared is possible as the non-headline factors are likely not as precisely estimated for minor announcements. If they pick up noise, this can lead to a reduction in the explanatory power at the quarterly frequency.

Figure S2.3: Quarterly R-Squared of Non-Headline News for Volatility and Commodity Indexes



Notes: For each country's asset price, this figure plots the increment in R-squared of non-headline news for the quarterly frequency. The red (leftmost) bars correspond to the *Baseline* specification and are the same as the red bars in Figure 5. The purple, orange, turquoise, and green bars depict alternative specifications *Covariance*, *All*, *Single*, and *Single—All*, respectively. These specifications are explained in Supplementary Appendix S2.3. The sample runs from January 1, 2000 to December 31, 2019 for the volatility indexes, and from May 7, 2007 to December 31, 2019 for the commodity factor.

In the third specification, labeled as *single* in Figure S2.2 and S2.3, we estimate a single non-headline factor for all twelve major announcements. Hence, this restricts the effect on the US yield curve to be the same across announcements. Note that this is the specification for which Gürkaynak, Kısacıkoğlu, and Wright (2020) run their lower frequency analysis. Despite being estimated over a different sample and using a different set of announcement series (e.g., Gürkaynak, Kısacıkoğlu, and Wright (2020) include FOMC announcements in their estimation) our factor has a correlation of 0.84 with their factor for the overlapping announcements. As illustrated in Figures S2.2 and S2.3, this specification leads to reduced explanatory power compared to the baseline in the large majority of cases. This implies that the common factor assumption is likely too restrictive to understand the international effects.

For completeness, we lastly estimate a single common factor for all announcements. The results are labeled as *single—all* in Figures S2.2 and S2.3. While this specification leads to an increase in explanatory power compared to the *single* specification, it is generally smaller than the *all* specification—again indicating that the common factor assumption is too restrictive in our context.

With these results in hand, we briefly discuss why we chose the current *baseline* specification as it is. While specifications *all* and *single—all* lead to greater R-squared values, additional unreported checks indicate that these values are not very robust. This likely comes from the fact that in the former case many of the factors are not well identified, and that in the latter case the common factor is identified from a relatively small set of non-announcement days. Further, the *single* specification seems to restrictive—as discussed earlier. Lastly, while we view the *covariance* specification as similarly justifiable, we already consider our entire estimation as conservative since it is only based on US yields. In light of that, we decided to go with our current baseline, which leads to slightly greater R-squared values.

S3 Monetary Policy Analysis

S3.1 Construction of Shocks

For each central bank, we use high-frequency surprises in interest rates around monetary policy announcements to construct monetary policy shocks. Following Gürkaynak, Sack, and Swanson (2005) and Swanson (2021), we construct three shocks: a *target rate shock*, a *forward guidance shock*, and a *quantitative easing shock*. We next describe the shock construction for each central bank.

S3.1.1 Fed Dataset

For the Federal Reserve, we use scheduled FOMC announcements from January 1996 till December 2019. We focus on scheduled releases because unscheduled meetings are potentially accompanied with exceptional financial market responses. Our sample covers 190 announcements. Following Swanson (2021), our shocks are based on eight variables ($MP1$, $MP2$, $ED2$, $ED3$, $ED4$, $T2$, $T5$, and $T10$), which capture interest rates for maturities of up to 10 years. The shocks are constructed from 30-minute changes in interest rate futures contracts and are standard in the literature. All data comes from *Thomson Reuters Tick History*. The dataset is also used in Boehm and Kroner (2021). In that paper, we provide details on the shock construction and show that the 30-minute changes align well with those of prior work. See Table S3.1 for more details.

Following Swanson (2021), we construct three monetary policy shocks. To do so, we first extract three factors via principal components from the eight variables. Using the Cragg and Donald (1997) test, we confirm that the data is best explained by three factors. We rotate these factors such that only one factor loads on changes in the current federal funds rate, which we refer to as the target rate shock. The other two factors have no effect on the federal funds rate. To disentangle them, we impose that one factor minimizes the variation in the data prior to the zero lower bound period starting on December 16, 2008. We call this factor the quantitative easing shock. We refer to the last factor as the forward guidance shock. For details on how to impose these restrictions, see Swanson (2021).

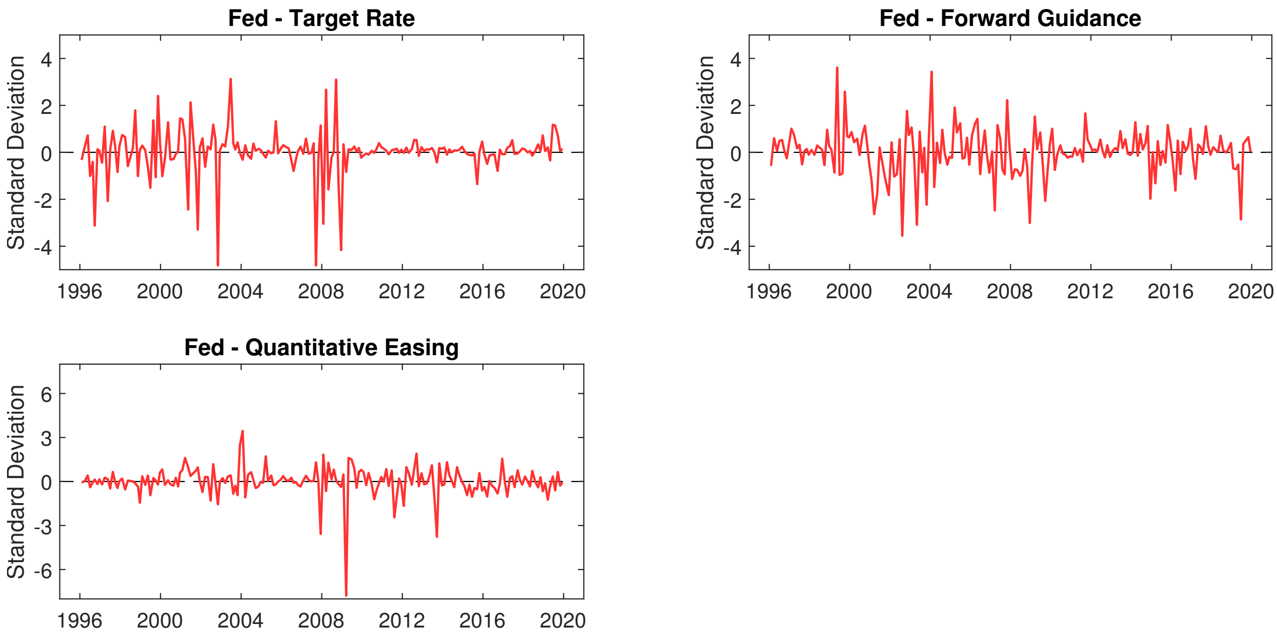
The resulting time series of each shock are shown in Figure S3.1. We also compare our shocks to those from Swanson (2021). For the overlapping sample, the correlations are 97 percent for the target rate shock, 87 percent for the forward guidance shock, and 78 percent for the quantitative easing shock. Further, we show below in Supplementary Appendix S3.3 that our main findings are robust to directly using the shocks by Swanson (2021).

S3.1.2 ECB Dataset

To construct the shocks for the Euro Area, we use an updated version of the high-frequency event study dataset by Altavilla et al. (2019). Due to the announcement structure of the ECB, we have a press release, as well as a press conference window. We have 195 press releases and 190 press conferences between January 2002 and December 2019. For each of the two releases, we construct 30-minute changes in asset prices following Altavilla et al. (2019). We use the seven variables (OIS_{1M}, OIS_{3M}, OIS_{6M}, OIS_{1Y}, OIS_{2Y}, OIS_{5Y}, OIS_{10Y}). Note that the maturities of these contracts match those in the other datasets relatively well. See Table S3.1 for more details.

Following Altavilla et al. (2019), we extract one factor for the press release window, which we refer to as the target shock. For the press conference window, we extract three factors, apply the restrictions as in Swanson (2021), and use the two factors that have no effect on the short rate. We refer to these as the forward guidance and quantitative easing shocks.

Figure S3.1: Times Series of US Monetary Policy Shocks



Notes: This figure shows the time series of the three monetary policy shocks of the Federal Reserve. The units are in standard deviations.

Figure S3.2 shows the time series of each shock. We also compare our shocks to those constructed by Altavilla et al. (2019). For the overlapping sample, the correlations are 99 percent for the target rate shock, 79 percent for the forward guidance shock, and 84 percent for the quantitative easing shock. Further, we show below in Supplementary Appendix S3.3 that our main findings are robust to directly using the shocks from Altavilla et al. (2019).

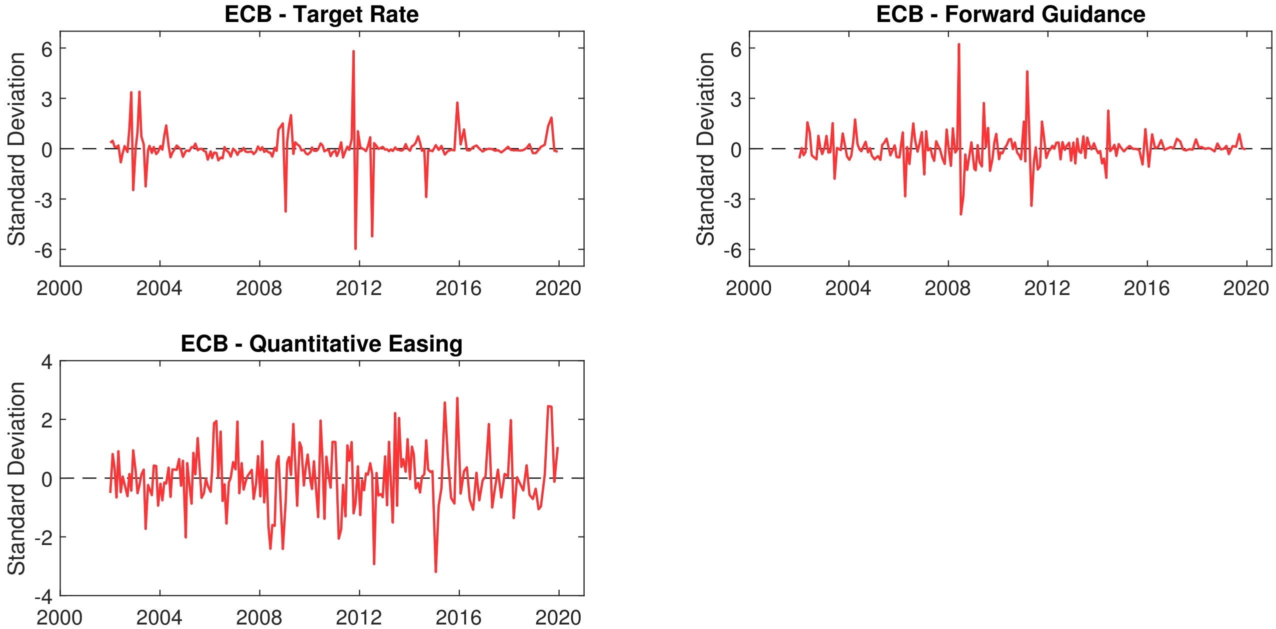
S3.1.3 BoE Dataset

For the Bank of England, we focus on scheduled Monetary Policy Committee (MPC) announcements. The sample ranges from June 1997, when the Bank of England became independent, to December 2019. The dates and times are from Bloomberg, as well as the Bank of England online archive on news, publications and events (`www.bankofengland.co.uk/news`). We drop the exceptional 150 basis points rate cut on November 6, 2008, leaving us with 256 announcements.

The construction of the shocks is based on seven variables, the first four short Sterling futures contracts ($FSS1$–$FSS4$), as well as the 2-year, 5-year, and 10-year Gilt yields ($G2$, $G5$, and $G10$). All data comes from *Thomson Reuters Tick History* and each variable is constructed as a 30-minute change around announcements. See Table S3.1 for more details. We then again construct three monetary policy shocks following the procedure of Swanson (2021). We start by showing that the data is best explained by three dimensions using the Cragg and Donald (1997) test, and subsequently extract three principal components. The restrictions to obtain the target rate, forward guidance, and quantitative easing shocks are similar to those described above for the US. For the BoE shocks, the sample for which the explained variation by the quantitative easing shock is minimized ends in February 5, 2009, the last MPC meeting before the asset purchasing program started.

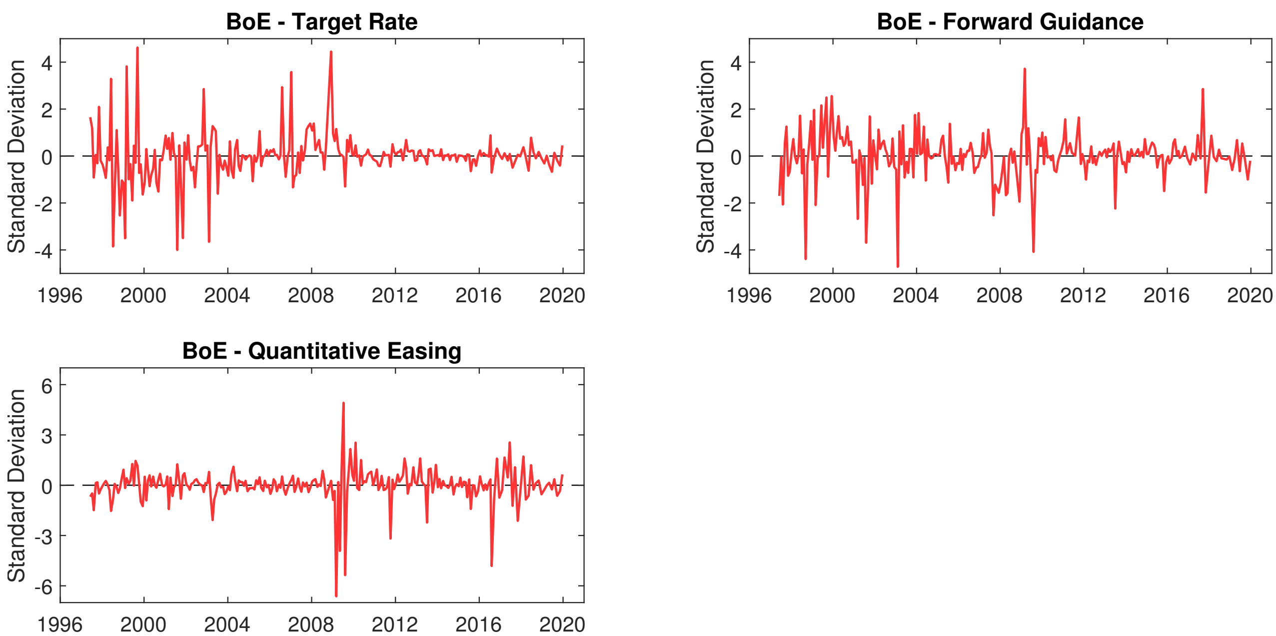
Figure S3.3 shows the time series for each shock. Broadly, the shocks are consistent with the idea that forward guidance and quantitative easing played a more dominant role since the Great

Figure S3.2: Times Series of ECB Monetary Policy Shocks



Notes: This figure shows the time series of the three monetary policy shocks of the European Central Bank. The units are in standard deviations.

Figure S3.3: Times Series of BoE Monetary Policy Shocks



Notes: This figure shows the time series of the three monetary policy shocks of the Bank of England. The units are in standard deviations.

Recession. While we do not have access to comparable shock series from a previous paper, other papers have used some of the underlying 30-minute changes as shocks. Our series of changes in the nearest Sterling futures contract ($FSS1$) has a 89 percent correlation with the series by Miranda-

Agrippino (2016), and a 91 percent correlation with the series by Gerko and Rey (2017), where for the latter comparison our series is aggregated to the monthly level. Our series of changes in the second nearest Sterling futures contract ($FSS2$), aggregated to the monthly level, has a 91 percent correlation with the series by Cesa-Bianchi, Thwaites, and Vicondoa (2020). All of these shock series from the previous literature correspond most closely to our target rate shock.

Table S3.1: Intraday Data for Monetary Policy Shocks

Variable in Text	Underlying Instruments	Ticker	Sample
Fed Shocks			
$MP1$	Federal Funds Rate Futures	FFc1–FFc2	1996–2019
$MP2$	Federal Funds Rate Futures	FFc3–FFc4	1996–2019
$ED2$	Eurodollar Futures	EDcm2	1996–2019
$ED3$	Eurodollar Futures	EDcm3	1996–2019
$ED4$	Eurodollar Futures	EDcm4	1996–2019
$T2$	2-Year Treasury Futures	TUc1/TUc2	1996–2019
$T5$	5-Year Treasury Futures	FVc1/FVc2	1996–2019
$T10$	10-Year Treasury Futures	TYc1/TYc2	1996–2019
ECB Shocks			
OIS_{1M}	1-Month Overnight Index Swap Rate	EUREON1M=	2002–2019
OIS_{3M}	3-Month Overnight Index Swap Rate	EUREON3M=	2002–2019
OIS_{6M}	6-Month Overnight Index Swap Rate	EUREON6M=	2002–2019
OIS_{1Y}	1-Year Overnight Index Swap Rate	EUREON1Y=	2002–2019
OIS_{2Y}	2-Year Overnight Index Swap Rate	EUREON2Y=	2002–2019
OIS_{5Y}	5-Year Overnight Index Swap Rate	EUREON5Y=*	2002–2019
OIS_{10Y}	10-Year Overnight Index Swap Rate	EUREON10Y=*	2002–2019
BoE Shocks			
$FSS1$	1-Quarter Short Sterling Futures	FSScm1/FSSc1–FFc3	1997–2019
$FSS2$	2-Quarter Short Sterling Futures	FSScm2/FSSc4	1997–2019
$FSS3$	3-Quarter Short Sterling Futures	FSScm3/FSSc5	1997–2019
$FSS4$	4-Quarter Short Sterling Futures	FSScm4/FSSc6	1997–2019
$G2$	2-Year Gilt Yield	GB2YT=RR	1997–2019
$G5$	5-Year Gilt Yield	GB5YT=RR	1997–2019
$G10$	10-Year Gilt Yield	GB10YT=RR	1997–2019
Stock Indexes (Figure S3.6)			
S&P 500		.SPX	1996–2019
STOXX 50 Index		.STOXX50E	2002–2019
FTSE 100		.FTSE	1997–2019

Yield Curve (Figure S3.5)	Fed	ECB	BoE
3-Month Yield	US3MT=X	EUREON3M=	GB3MT=RR
1-Year Yield	US1YT=X	EUREON1Y=	GB1YT=RR
2-Year Yield	US2YT=X	EUREON2Y=	GB2YT=RR
5-Year Yield	US5YT=X	EUREON5Y=*	GB5YT=RR
10-Year Yield	US10YT=X	EUREON10Y=*	GB10YT=RR

Notes: This table provides an overview of the intraday data from *Thomson Reuters Tick History* used to construct the monetary policy shocks. *Ticker* refers to the Reuters Instrument Code (RIC). For Fed shocks, details are provided in Boehm and Kroner (2021). For ECB shocks, the data comes from Altavilla et al. (2019) where we are providing the underlying data as shown in their Appendix Table B.1. The *Stock Indexes* and the *Yield Curve* panel refer to the additional data used for Figure S3.6 and Figure S3.5, respectively. *Following Altavilla et al. (2019), we use German bond yields of the corresponding maturity before 2011 as the 2-year and 5-year OIS rates are not available.

S3.2 Additional Results

The first two rows of Figure S3.4 show the effects of forward guidance shocks on international stock markets. As the pooled effects show (labelled "All"), a one standard deviation contractionary forward guidance shock of the Fed reduces international stock prices by approximately 10 basis points. This effect is statistically significant at the 5 percent level. This contrasts to the forward guidance shocks of the ECB and the BoE, which have substantially smaller effects that are not significant at conventional levels. The country-specific effects shown in the figure are of varying sizes and significance. An important feature of these estimates, however, is that whenever we can estimate the effects of multiple central banks on a given countries' stock market, the point estimates for the Fed are greater (in absolute value) than those of the ECB and the BoE. Similar to the conclusions from the target rate shock, the results in Figure S3.4 are consistent with our previous interpretation that the outsized effect of US macro news is driven by the transmission of US-specific shocks as opposed to the presence of common shocks.

Rows three and four of Figure S3.4 show analogous effects of quantitative easing shocks. While the relative magnitudes of the effects display a pattern across central banks that is qualitatively similar to that of target rate and forward guidance shocks, almost all effects are imprecisely estimated. The usefulness of these shock series for comparing effect sizes across central banks is therefore limited.

S3.3 Robustness

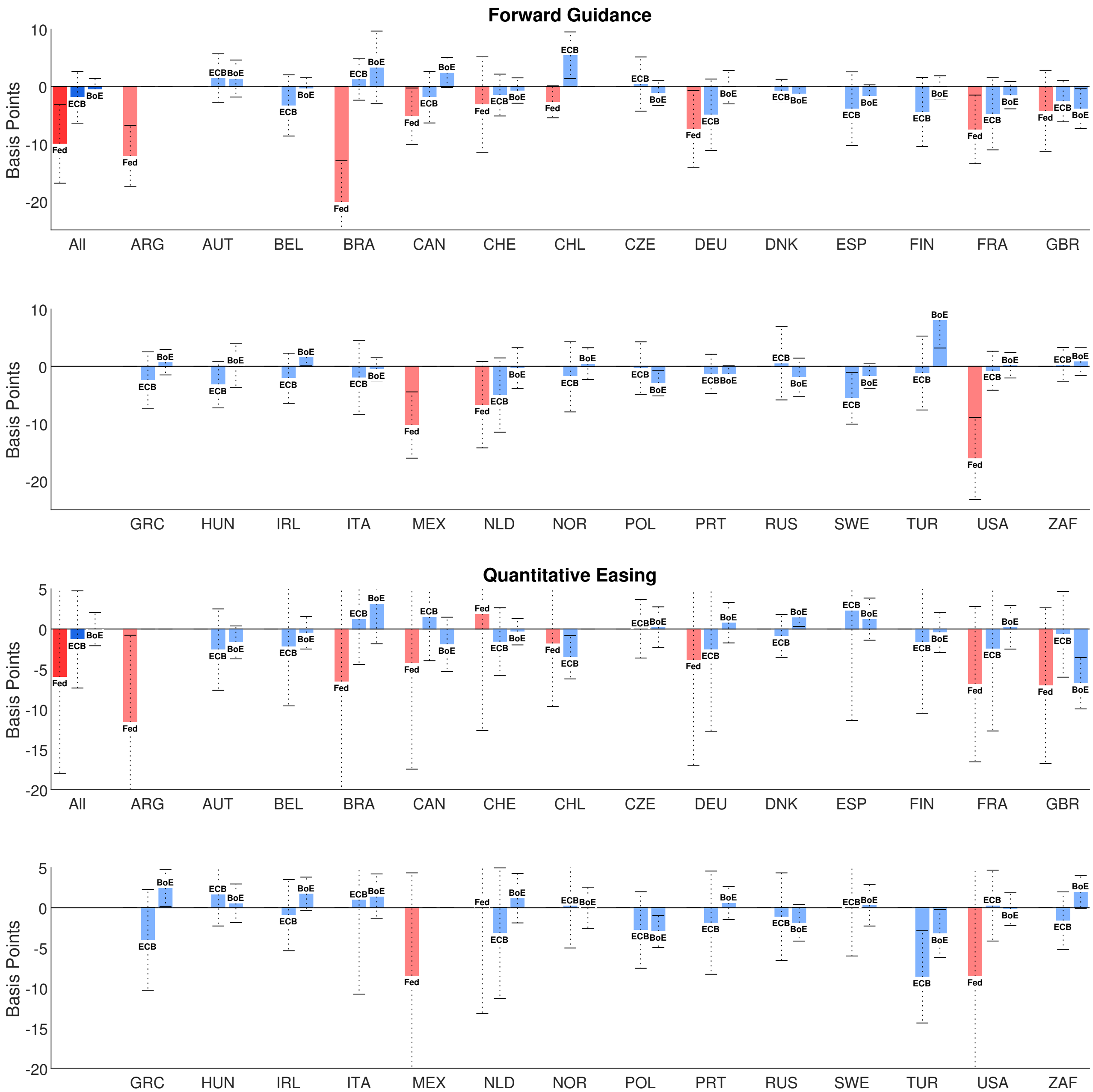
S3.3.1 Unit of Comparison

To ensure the comparability of shock magnitudes across central banks in the baseline analysis, we expressed each shock in units of standard deviations—like the macroeconomic news surprises. The idea here is to compare the average effects of a typical one standard deviation monetary policy surprise on financial markets. However, since central bank policies are generally difficult to compare, an alternative is to normalize the shocks in terms of their effects on the domestic yield curve. We next present results of this alternative strategy as a robustness check.

The top row of Figure S3.5 shows the loadings of each shock on the domestic yield curve in a 30-minute window around announcements. These loadings are constructed from government bond yields; specifically, we regress the respective shock (in standard deviations) on the 30-minute changes of various domestic government bond yields—in separate regressions with one regressor at a time. Note that these government bond yields are not necessarily the same as the yields from which the shocks are constructed. The advantage of using government bonds in this exercise is that they allow for a direct comparison of magnitudes across central banks at the exact same maturity. (For the ECB shocks, we use OIS rates of the relevant maturity instead of government bond yields.) The conclusion from the top row of Figure S3.5 is that while the shapes are similar across central banks, the magnitudes generally differ, in particular for the BoE.

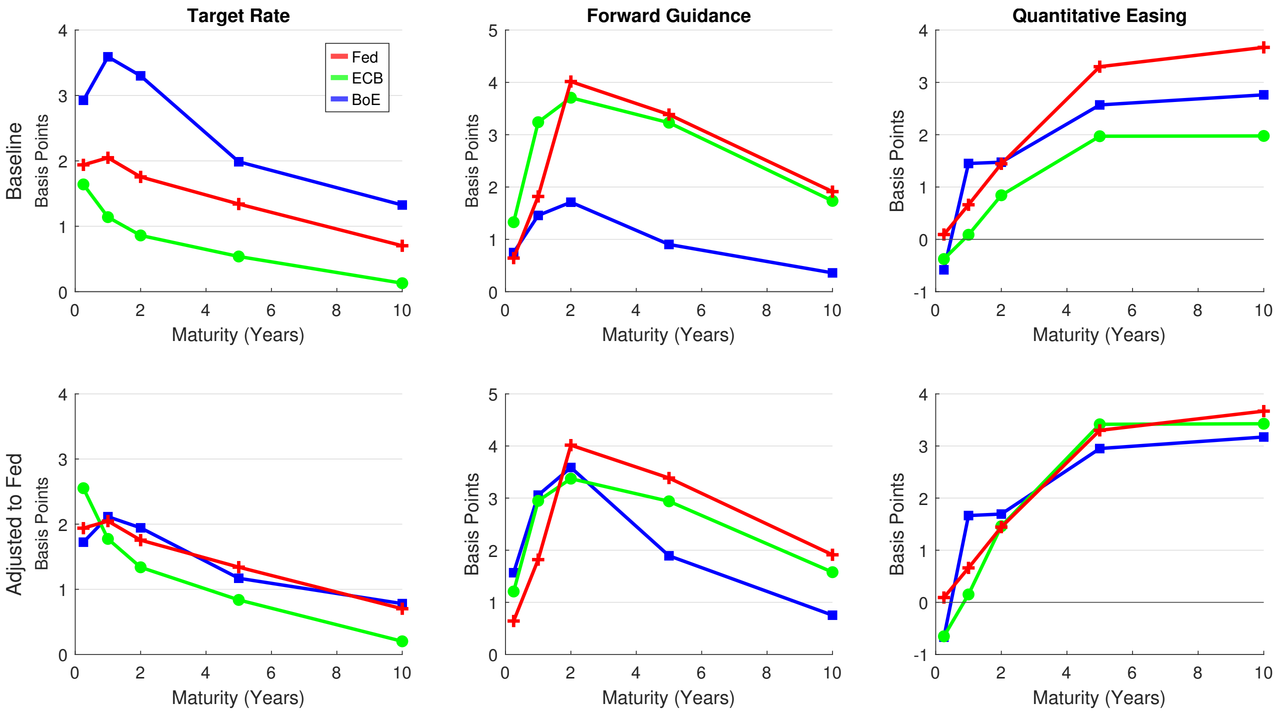
To mitigate concerns that these differences drive our results, we re-scale the ECB and BoE shocks such that the new loadings minimize the (Euclidean) distance to the respective Fed loadings, which are left unchanged. The new loadings are shown in the bottom row of Figure S3.5. The pooled effects of these shock after re-scaling are shown in the top-left panel of Figure S3.6. The results show that the asymmetry documented in Section 6.3 and Supplementary Appendix S3.2 is robust to this alternative normalization of shocks.

Figure S3.4: Effects of Unconventional Monetary Policy Shocks on International Stock Markets



Notes: This figure shows the effects of forward guidance and quantitative easing shocks of the Federal Reserve (Fed), the European Central Bank (ECB), and the Bank of England (BoE) on international stock markets. The leftmost bars in the first and third row (labelled "All") show the pooled effects across countries for each central bank. Each of the other bars represent the effect of a given central bank's shock on a country's stock market. Missing bars indicate instances in which the country is dropped because it had less than 24 observations for a given monetary policy shock. The coefficients are estimated analogously to equations (3) and (4). The units of the stock index changes are in basis points. Each shock corresponds to an increases in interest rates and is of one standard deviation in magnitude. The black error bands depict 95 percent confidence intervals, where standard errors are two-way clustered by announcement and by country.

Figure S3.5: Effects of Monetary Policy Shocks on Domestic Yield Curve



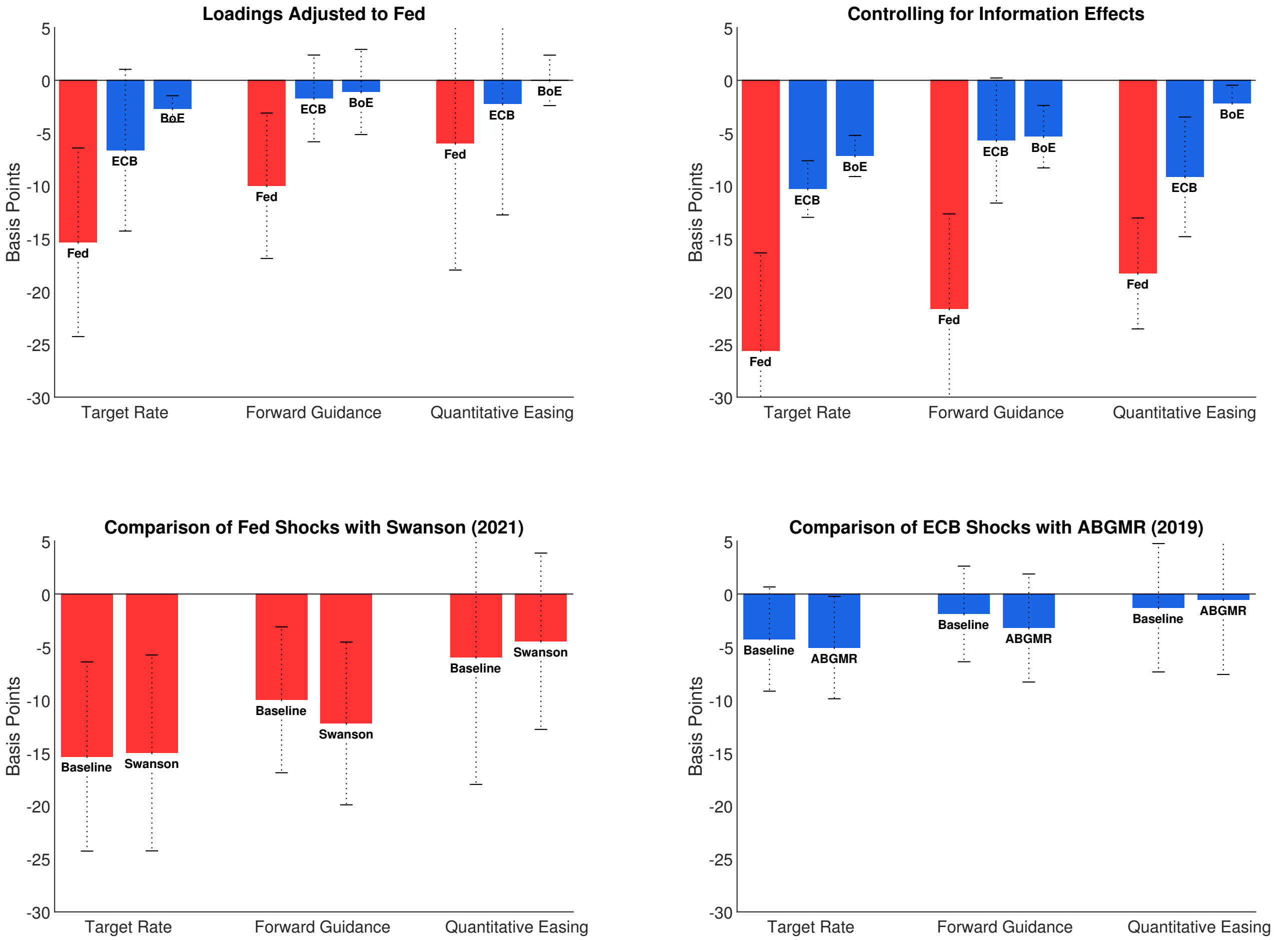
Notes: This figure illustrates the estimated effects of each shock on the domestic yield curve. The shown maturities are 3 months, 1 year, 2 years, 5 years, and 10 years. See the bottom panel of Table S3.1 for details on the data. The red, green, and blue lines correspond to the estimates for the Fed, the ECB, and the BoE, respectively. The top row shows the estimates for a one standard deviation shock as used in the main text. The bottom row displays the estimates for the ECB and BOE shocks after re-scaling as discussed in the text.

S3.3.2 Controlling for Information Effects

We next turn to the issue of information or signaling effects of monetary policy. The idea here is that a central bank could signal information about the state of the economy to the public through its policy. These effects would push stock markets in the opposite direction of traditional monetary policy shocks. If the strength of the information effects differ across central banks, this could potentially explain the asymmetry documented above. To check this, we follow the approach by Miranda-Agrippino and Nenova (2022), which is based on the "poor man's" identification of Jarociński and Karadi (2020), and only considers announcements for which the domestic stock market index responds negatively to contractionary shocks and positively to expansionary shocks. Here, we use the STOXX 50 index as the domestic stock market index for the ECB.

The top-right panel of Figure S3.6 shows the results of this exercise. First, and most importantly the asymmetry documented earlier is robust to controlling for information effects. Second, consistent with previous papers, the effect sizes are substantially greater and so is the precision of the estimates—in particular for the forward guidance and quantitative easing shocks. Hence, the results indicate that information effects are potentially responsible for the noisy estimates in Figure S3.2. They cannot, however, explain the asymmetry.

Figure S3.6: Effects of Monetary Policy Shocks on International Stock Markets—Robustness



Notes: This figure illustrates the results for four different robustness checks, showing the pooled effects across central banks and type of policy shocks. The top-left panel shows estimates when ECB and BoE shocks are re-scaled as described in Supplementary Appendix S3.3.1. The top-right panel shows the results when information effects are removed as described in Supplementary Appendix S3.3.2. The bottom-left panel shows the comparison of our baseline estimates with those obtained when directly using the shocks by Swanson (2021). The bottom-right panel does the analogous exercise for the ECB shocks where we now use the shocks by Altavilla et al. (2019).

S3.3.3 Comparison with Shocks of Previous Literature

As mentioned above, we also contrast our estimates with those obtained from shocks by previous papers. For the Fed, we employ the shocks by Swanson (2021). The results are shown in the bottom-left panel of Figure S3.6. For the ECB, we use the shocks by Altavilla et al. (2019) and the bottom-left panel of Figure S3.6 shows the results for that comparison. In both cases, the estimates are very similar to our baseline case.

S4 State-Dependent Effects of US Macro News

Prior work has established that the effects of news on equity prices are not stable over time (e.g., McQueen and Roley, 1993; Boyd, Hu, and Jagannathan, 2005; Andersen et al., 2007; Goldberg and Grisse, 2013; Gürkaynak, Kısacıkoğlu, and Wright, 2020; Gardner, Scotti, and Vega, 2022; Elenev et al., 2022, among others). In this appendix, we extend our analysis to allow for such time-varying effects. We confirm prior findings that the effects of news on stock prices vary along several dimensions. However, we also show that the average effects we report in the main text are not driven by large effects in extreme episodes such as deep recessions or slumps, or episodes at the zero lower bound (ZLB), but are present in normal times.

Our setup in this paper differs from most prior applications since we study how news in the US affects *foreign* asset prices. For a given economic indicator of interest (e.g., recession vs. expansion), this international setting leads to the possibility that the effect size depends on the indicator's value in the US (where the news originates), its value in the foreign country (whose stock price response we study), or both. Hence, for given measure, our regression specification will allow for the effect size to vary with the value of the measure in the US and in the foreign country.

Based on the prior literature, we consider the following four measures. First, we contrast recessions and expansions using simple recession indicators (e.g., Boyd, Hu, and Jagannathan, 2005; Andersen et al., 2007). More specifically, we consider an indicator function, $\mathbf{1}_{i,t}^{rec}$, which equals one if and only if country i's economy is in recession. To measure US recession periods, we use the business cycle dates from the National Bureau of Economic Research (NBER). For the other countries, we use the dates provided by the Organisation for Economic Co-operation and Development (OECD).

Second, we will allow the effect size to depend on a measure of business cycle slack constructed from the unemployment rate (similar to, e.g., McQueen and Roley, 1993; Elenev et al., 2022; Gardner, Scotti, and Vega, 2022). Our preferred measure for whether an economy experiences slack is the empirical cumulative distribution function (cdf) of a country's unemployment rate. This function is defined as

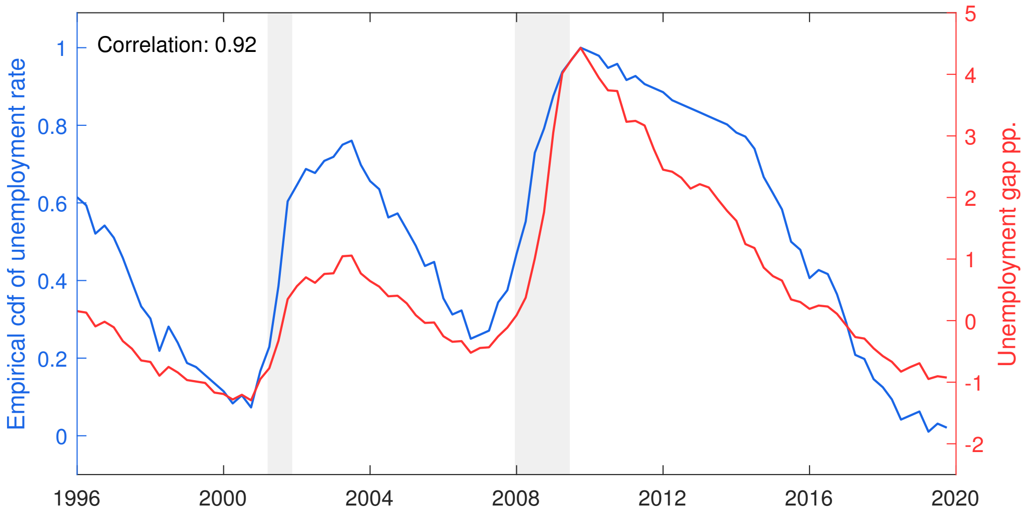
$$F_i^u(u) = \frac{1}{N_i}\sum_{\tau=1}^{N_i} \mathbf{1}\left(u_{i,\tau} \leq u\right),$$

where $u_{i,\tau}$ is the unemployment rate in country i at time τ, and N_i is the number of unemployment rate observations of country i from which the cdf is estimated. The empirical cdf maps a generic value of the unemployment rate u into the unit interval $[0, 1]$. This measure captures in a non-parametric way whether the countries' unemployment rate is high in comparison to its own history and future. Relative to other measures, the empirical cdf has a number of advantages.[4] Most importantly, it can be constructed from data on the unemployment rate alone. Relative to alternative measures such as the unemployment gap, it does not require data on the natural rate of unemployment, which is difficult to estimate and to our knowledge not available for most foreign countries in our sample.

Figure S4.1 compares the empirical cdf of the unemployment rate, evaluated at the unemployment rate at time t, $F_i^u(u_{i,t})$, to the unemployment gap for the case of the US. As in Gardner, Scotti, and Vega (2022), the unemployment gap is constructed as the difference between the unemployment

[4]For example, unlike the measure proposed by Auerbach and Gorodnichenko (2012), it does not require calibration of any parameters. Further, in contrast to the approach by Ramey and Zubairy (2018), it does not require taking a stance on a threshold value.

Figure S4.1: Comparison of Empirical cdf of Unemployment Rate and Unemployment Gap



Notes: This figure compares the empirical cdf of the unemployment rate in the US with the unemployment gap in the US. Shaded areas indicate NBER recession periods.

rate and the natural rate of unemployment.[5] The figure shows that our measure of the empirical cdf correlates very highly with the unemployment gap (the correlation is 0.92). In order to preserve the interpretation of the main effect in the regressions below, we subtract 0.5 from our measure of the empirical cdf and include the interaction term of $F^u_{i,t} := F^u_i(u_{i,t}) - 0.5$ with the surprise of interest in the regression. The data on unemployment rates is quarterly and come from the OECD.

Third, we study whether the effect size depends on whether the economy is at the ZLB (or effective lower bound). The ZLB introduces a non-linearity in the monetary reaction function and prior work has argued that time-varying responsiveness of monetary policy could drive the time-varying effects of news on equity prices (e.g., Goldberg and Grisse, 2013). Following Boehm (2020), the indicator function, $\mathbf{1}^{ZLB}_{i,t}$, equals one if and only if the countries' short-term interest rate is below 75 basis points. The data on short-term interest rates is monthly and comes from the OECD. This dataset defines the short-term rate as a three-month money market rate.[6]

Lastly, we use the FOMC Sentiment Index as constructed by Gardner, Scotti, and Vega (2022). This index is based on textual analysis of FOMC statements and captures an assessment of current and future economic conditions as perceived by the Fed. High values of the index typically occur at times when the US economy is doing well. Gardner, Scotti, and Vega (2022) show that the sensitivity of equity prices to US macro news varies strongly with this index. We de-mean this measure in order to obtain our preferred interpretation of the main effect in the regression. For ease of interpretation of the interaction effect, we also divide the de-meaned index by its standard deviation. In the regression below, this measure is denoted by $SI_{US,t}$.

[5] For the US an estimate of the natural rate of unemployment is available from the Congressional Budget Office.

[6] While data is missing for Turkey and Brazil, we confirm through other sources that neither country had a policy rate close or below 75 basis points over our sample period. Hence, we set the indicator for both countries to zero throughout.

With these measures at hand, we then estimate the following joint specification:

$$\begin{aligned}
\Delta q_{i,t} = \alpha_i + \gamma^y s^y_{US,t} + \sum_{k\neq y} \gamma^k s^k_{US,t} \\
&+ \chi^y_r s^y_{US,t} \mathbf{1}^{rec}_{US,t} + \psi^y_r s^y_{US,t} \mathbf{1}^{rec}_{i,t} + \sum_{k\neq y} \left(\chi^k_r s^k_{US,t} \mathbf{1}^{rec}_{US,t} + \psi^k_r s^k_{US,t} \mathbf{1}^{rec}_{i,t} \right) + \theta^y_r \mathbf{1}^{rec}_{US,t} + \phi^y_r \mathbf{1}^{rec}_{i,t} \\
&+ \chi^y_u s^y_{US,t} F^u_{US,t} + \psi^y_u s^y_{US,t} F^u_{i,t} + \sum_{k\neq y} \left(\chi^k_u s^k_{US,t} F^u_{US,t} + \psi^k_u s^k_{US,t} F^u_{i,t} \right) + \theta^y_u F^u_{US,t} + \phi^y_u F^u_{i,t} \qquad \text{(S4.1)} \\
&+ \chi^y_Z s^y_{US,t} \mathbf{1}^{ZLB}_{US,t} + \psi^y_Z s^y_{US,t} \mathbf{1}^{ZLB}_{i,t} + \sum_{k\neq y} \left(\chi^k_Z s^k_{US,t} \mathbf{1}^{ZLB}_{US,t} + \psi^k_Z s^k_{US,t} \mathbf{1}^{ZLB}_{i,t} \right) + \theta^y_Z \mathbf{1}^{ZLB}_{US,t} + \phi^y_Z \mathbf{1}^{ZLB}_{i,t} \\
&+ \chi^y_S s^y_{US,t} SI_{US,t} + \sum_{k\neq y} \chi^k_S s^k_{US,t} SI_{US,t} + \theta^y_S SI_{US,t} + \varepsilon_{i,t}.
\end{aligned}$$

Note that in this specification, the measures $F^u_{US,t}$, $F^u_{i,t}$, and $SI_{US,t}$ have (approximately) mean zero, and hence the main effect γ^y captures the effect of US macroeconomic surprise $s^y_{US,t}$ on the foreign asset price $q_{i,t}$ when (i) both the US and the foreign country's economy are expanding, (ii) when the two countries' unemployment rates are at their mean, (iii) when the two countries' monetary authorities are not constrained by the ZLB, and (iv) when the Fed Sentiment Index is at its mean. Note that the period over which we estimate specification (S4.1) begins in 2000 as the FOMC Sentiment Index is not available before.

Table S4.1 shows the estimates. Several interaction coefficients are statistically significant and in some cases economically large. The two most important of these are the interactions of the surprise with the empirical cdf of the US unemployment rate as well as with the FOMC Sentiment Index. The effects are larger if the US unemployment rate is high and if the FOMC Sentiment Index is low—in line with prior findings that the effects are larger during bad times. The estimates also suggest that the effect size varies more with the state of the US economy than with the state of the foreign economy.

For our analysis, the most important result in Table S4.1 is that the main effects remain similar to the estimates reported in Table 3. They also remain statistically and economically significant. Recall that given the construction of the interaction terms, these main effects capture the average effects of US news on foreign stock markets when (i) both the US and the foreign economy are in an expansion, (ii) the US and the foreign unemployment rate are at their median value, (iii) neither economy is at the ZLB, and (iv) when the FOMC Sentiment Index is at its mean. The similarity to our baseline results implies that the estimates reported in the text are not driven by very large effects in extreme business cycle states such as deep recessions, episodes of extreme slack, or times at the ZLB. They are also present in normal times.

To provide one concrete example, we discuss the case of nonfarm payrolls. As Table S4.1 shows, the main effect is 15.60 basis points per standard deviation surprise. Besides this main effect, the interaction terms of the surprise with the empirical cdf of the US unemployment rate, with the foreign ZLB indicator, and with the FOMC Sentiment Index are statistically significant. To understand the economic significance of these effects, note that, all else equal, the main effect of 15.69 basis points increases by 7.62 (= 30.47 × 0.25) basis points if the US unemployment rate is changed from its median to the 75th percentile. Further, and again holding all else equal, the effect size increases by 13.78 basis points if the foreign country's monetary authority is constrained by the ZLB. Lastly,

Table S4.1: Time-Varying Effects of US News

	Capacity Utilization	CB Consumer Confidence	Core CPI	Core PPI	Durable Goods Orders	GDP A
Stock Index (bp)						
News	6.17***	9.85***	-10.59***	-7.32***	11.80***	16.40***
	(1.99)	(2.49)	(2.20)	(1.73)	(2.58)	(3.79)
News × Recession USA	8.53	10.60	-7.44	11.83*	-14.40**	5.65
	(5.40)	(6.21)	(6.19)	(5.78)	(6.05)	(6.66)
News × Recession Foreign	-0.63	1.59	4.01**	0.47	3.20	2.37
	(1.09)	(1.73)	(1.45)	(1.36)	(2.00)	(4.34)
News × Unemployment USA	13.24*	20.55***	18.30**	0.75	25.78***	27.89**
	(7.33)	(7.36)	(6.81)	(4.68)	(6.20)	(11.45)
News × Unemployment Foreign	-2.76	-3.33*	0.36	-1.25	-5.87**	-2.10
	(1.93)	(1.94)	(2.08)	(2.10)	(2.18)	(5.22)
News × ZLB USA	-4.91	-4.98	-2.53	3.97	-9.17*	-16.95*
	(3.82)	(4.20)	(3.18)	(2.43)	(4.61)	(8.47)
News × ZLB Foreign	-0.61	2.94	6.94***	3.92**	-1.41	5.51
	(1.80)	(1.95)	(2.31)	(1.48)	(1.82)	(4.27)
News × FOMC Sentiment	-0.82	-3.12*	0.72	-0.05	-1.96	-1.29
	(1.36)	(1.81)	(1.40)	(1.34)	(1.64)	(3.17)
R^2	0.13	0.28	0.31	0.34	0.25	0.57
Observations	5215	5281	4963	5055	4957	1658

	Initial Jobless Claims $\cdot(-1)$	ISM Mfg Index	New Home Sales	Nonfarm Payrolls	Retail Sales	UM Consumer Sentiment P
Stock Index (bp)						
News	4.92***	6.95**	6.87***	15.69***	6.11**	9.11***
	(0.89)	(3.13)	(1.83)	(3.56)	(2.31)	(2.17)
News × Recession USA	-2.84	1.33	2.90	-5.51	8.02	-12.56*
	(2.13)	(9.11)	(8.73)	(8.63)	(6.39)	(7.16)
News × Recession Foreign	-1.10	3.23	-5.60***	-2.54	-2.16	-2.41
	(0.80)	(2.81)	(1.69)	(2.76)	(2.32)	(1.88)
News × Unemployment USA	5.78*	-8.58	-3.94	30.47***	4.69	13.49*
	(2.94)	(9.90)	(6.48)	(10.85)	(8.47)	(6.83)
News × Unemployment Foreign	-1.20	-3.42	-0.64	-2.13	-4.93**	-0.59
	(1.29)	(2.86)	(1.22)	(3.61)	(1.94)	(1.94)
News × ZLB USA	0.98	12.21*	10.52***	8.93	8.86*	-2.66
	(2.03)	(6.33)	(3.27)	(7.88)	(5.00)	(4.09)
News × ZLB Foreign	-1.10	3.40	-0.23	13.78***	1.93	-0.69
	(0.93)	(3.13)	(2.26)	(4.12)	(1.65)	(1.97)
News × FOMC Sentiment	-3.06***	-4.17*	-4.96***	-10.90***	-1.22	-0.86
	(0.77)	(2.24)	(1.56)	(3.53)	(2.09)	(2.08)
R^2	0.25	0.30	0.17	0.39	0.34	0.10
Observations	21470	4822	5220	4945	5036	5350

Notes: This table presents estimates of γ^y, χ^y_r, χ^y_u, χ^y_Z, χ^y_S, ψ^y_r, ψ^y_u, and ψ^y_Z obtained using specification (S4.1) with the change in stock indexes as the dependent variable. The interaction terms are constructed as discussed in the main text. Units are in basis points. Standard errors are two-way clustered by announcement and by country, and reported in parentheses. ***, **, and * indicate significance at the 1, 5, and 10 percent level.

ceteris paribus, a one standard deviation decrease in the FOMC Sentiment Index raises the effect by 10.90 basis points. Hence, these findings confirm prior work documenting that there is sizable state-dependence in the effects of US macro news on equity prices.

S5 The Role of the US Dollar Exchange Rate

In this appendix, we investigate the effect of US macro news on exchange rates, i.e., the US dollar vis-a-vis the other countries' currencies in our sample.[7] The US dollar exchange rate is a key variable in international finance (Gourinchas, Rey, and Sauzet, 2019), and a potential amplification mechanism of cross-border financial spillovers as shown by Bruno and Shin (2015). They lay out a model in which foreign firms borrow funds in US dollar but finance assets in local currency and therefore have currency mismatch. A dollar depreciation improves their balance sheets and reduces credit risk for their lenders (local banks). This reduction in credit risk, in turn, raises banks' lending capacity and therefore improves global liquidity. If the Bruno and Shin (2015) mechanism is dominant, we expect to observe a US dollar appreciation (depreciation) simultaneously with a decrease (increase) in international stock markets.

To see whether this prediction is consistent with our findings, we re-estimate the pooled regression (3), where $\Delta q_{i,t} = q_{i,t+20} - q_{i,t-10}$ is now the 30-minute change of country i's US dollar exchange rate.[8] Exchange rates are measured in US dollars per one unit of foreign currency so that a positive coefficient indicates a depreciation of the US dollar. Table S5.1 reports the results of this exercise, jointly with the previously obtained estimates for stock indexes from Table 3.

As the table demonstrates, the US dollar typically appreciates after positive surprises about both US real activity, which is in line with Andersen et al. (2007). Further, stock prices increase while the dollar appreciates. This relationship suggests that the mechanism by Bruno and Shin (2015) is not dominant. Overall, the exchange rate effect is comparatively weaker as only four out of ten announcements lead to a significant effect.

After positive news about inflation, international stock markets decrease while the dollar appreciates. These responses echo earlier findings on the effects of contractionary monetary policy shocks in the literature (Eichenbaum and Evans, 1995; Miranda-Agrippino and Rey, 2020). They are also in line with our earlier evidence of a potentially dominant interest rate channel for price news. In this case, the joint response of exchange rates and stock prices is consistent with the mechanism by Bruno and Shin (2015).

As price news only plays a minor role in our results (see Section 7), most of our evidence suggests that the exchange rate is not central to the transmission of US macro news. That being said, our results do not rule out that the international dominance of the US dollar is the source of the asymmetric effects documented in Section 6. For example, Jiang, Krishnamurthy, and Lustig (2020) build a model of the global financial cycle based on the safety of the US dollar. In their model, the response of the exchange rate to productivity shocks depends on the endogenous response of monetary policy.

[7]See Andersen et al. (2003) for prior work on the effects of macroeconomic news on US dollar exchange rates.

[8]For members of the Euro Area, we do not use country-specific exchange rates prior to the inception of the currency union due to the short samples. We further drop Denmark from the sample because the Danish Krone is tightly and credibly pegged to the Euro. See Online Appendix Table B3 for details.

Table S5.1: Effects on International Stock Markets and US Dollar Exchange Rates

	Capacity Utilization	CB Consumer Confidence	Core CPI	Core PPI	Durable Goods Orders	GDP A
Stock Index (bp)						
News	5.36**	12.35***	-8.84***	-4.87***	5.63***	17.60***
	(2.28)	(2.02)	(1.89)	(1.29)	(1.60)	(3.36)
R^2	0.04	0.13	0.10	0.15	0.10	0.26
Observations	6054	6041	5717	5828	5610	1911
Exchange Rate (bp)						
News	-0.01	-0.40	-5.77***	-3.32***	-1.40	-7.85***
	(1.09)	(1.21)	(1.33)	(0.82)	(0.81)	(2.54)
R^2	0.00	0.02	0.08	0.07	0.06	0.11
Observations	3943	3974	3812	3896	3787	1286

	Initial Jobless Claims $\cdot(-1)$	ISM Mfg Index	New Home Sales	Nonfarm Payrolls	Retail Sales	UM Consumer Sentiment P
Stock Index (bp)						
News	4.89***	11.71***	4.23***	17.06***	10.52***	5.61***
	(0.73)	(2.24)	(1.40)	(2.99)	(1.68)	(1.54)
R^2	0.09	0.12	0.03	0.13	0.15	0.04
Observations	24334	5548	5908	5688	5786	5726
Exchange Rate (bp)						
News	-0.58	-4.03**	-1.38*	-12.16***	-2.12	-0.96
	(0.50)	(1.40)	(0.73)	(2.75)	(1.47)	(0.82)
R^2	0.03	0.07	0.04	0.16	0.10	0.01
Observations	16497	3971	3915	3868	3862	3682

Notes: The table presents results of the pooled regression for stock indexes, as shown in Table 3, and US dollar denominated local exchange rates, i.e., estimates of γ^y of equation (3), where the left-hand variable is now the 30-minute change of country i's exchange rate. Exchange rates are expressed in US dollars so that an increase reflects a depreciation of the US dollar relative to the local currency. The units are in basis points. Standard errors are two-way clustered by announcement and by country, and reported in parentheses. ***, **, and * indicate significance at the 1, 5, and 10 percent level. See Online Appendix Table B3 for details on the sample.

S6 Inspecting the Cross-Sectional Heterogeneity

In this appendix, we explore the heterogeneity of responses documented in Section 4.1. As Figure 4 illustrated, some countries' stock markets, including Germany's, France's, Italy's, and the Netherlands' respond systematically more strongly to US macroeconomic news than stock markets in Austria, Denmark or Portugal. It is therefore natural to ask whether countries' responsiveness to news is correlated with observables. Perhaps surprisingly, we find no robust correlation of the effect size with financial integration, trade integration, a measure of industry dissimilarity, or exposure to dollar valuation effects after appropriately controlling for other determinants of the effect size.

We consider four different exposure measures that could plausibly impact how strongly a countries' stock market responds to US news. First, we are interested in a measure of global financial integration, an intuitive exposure measure to international financial conditions. One may expect that countries with greater financial openness respond more strongly—consistent with theoretical explanations of the global financial cycle as discussed in Rey (2016). As is common in the literature, we measure financial integration of country i in year t as

$$\text{finInt}_{i,t} = \frac{\text{FA}_{i,t} + \text{FL}_{i,t}}{\text{GDP}_{i,t}}, \tag{S6.1}$$

where $\text{FA}_{i,t}$ and $\text{FL}_{i,t}$ denote the stocks of foreign assets and liabilities, respectively. Note that $\text{FA}_{i,t}$ and $\text{FL}_{i,t}$ include asset holdings and liabilities vis-à-vis *all* countries and not only vis-à-vis the US, in line with recent work emphasizing the importance of multilateral effects (Huo, Levchenko, and Pandalai-Nayar, 2020). All series are measured in current US dollars. The data is annual and taken from Lane and Milesi-Ferretti (2007, 2017).[9]

As Figure S6.1 shows, a handful of countries experiences an enormous growth in financial integration, most notably Ireland (IRL). The main concerns with these countries are (i) that the financial integration measure (S6.1) could reflect their tax haven character rather than exposure to the global financial cycle and (ii) that extreme values of these countries' financial integration measures could unduly drive the estimates. While we have checked that the results are broadly similar in a sample including all countries (estimates not reported), we prefer a set of baseline results, which excludes the most extreme outliers (Ireland (IRL), Switzerland (CHE), the Netherlands (NLD), the United Kingdom (GBR), and Belgium (BEL)).

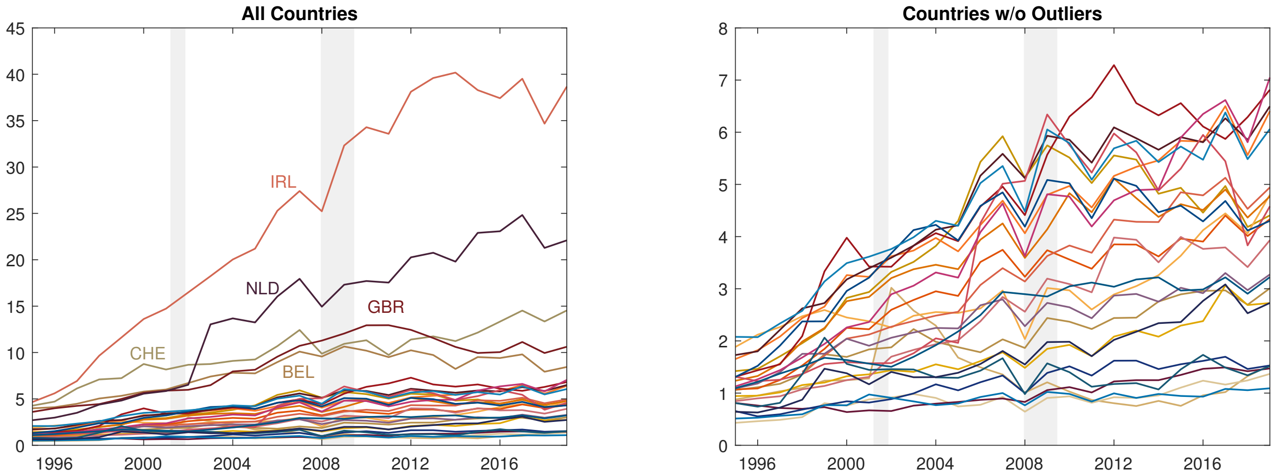
Second, we study the role of trade integration. It is known since Frankel and Rose's (1998) seminal work that countries that trade more have more correlated business cycles. This correlation suggests that trade transmits shocks across countries. Indeed, a large literature provides direct evidence for the transmission of shocks through trade linkages (see, e.g., Di Giovanni and Levchenko, 2010; Boehm, Flaaen, and Pandalai-Nayar, 2019, among many others). Again taking into account the role of multilateral effects, we calculate trade integration (or openness) for country i and year t as

$$\text{tradeInt}_{i,t} = \frac{\text{Imports}_{i,t} + \text{Exports}_{i,t}}{\text{GDP}_{i,t}}. \tag{S6.2}$$

Data on nominal imports, exports, and GDP is annual and obtained from the United Nations Statis-

[9] The asset and liability measures include portfolio equity and debt, foreign direct investment, other investment (including loans, deposits, and trade credit), financial derivatives, and reserve assets. Excluding foreign direct investment does not substantially affect our results.

Figure S6.1: Time Series of Financial Integration Measure by Country



Notes: This figure shows the time series of financial integration from 1995 to 2019. The construction of the measure follows equation (S6.1). The left hand side panel shows the time series for all countries in the sample. The right-hand side excludes the time series for the five outliers, i.e., Belgium, Ireland, Netherlands, Switzerland, and the United Kingdom. Note that the Euro Area is a separate line in both panels.

tics Division.

Third, we consider a measure of sectoral dissimilarity relative to the US. To the extent that business cycle shocks are sector-specific or have differential effects across sectors, one would expect countries with greater sectoral similarity to experience greater business cycle synchronization (Imbs, 2004). In the context of our empirical setup, US news may disproportionately capture the effects of shocks reflective of the US sectoral structure. This composition of shocks could result in greater effects on countries with an industrial structure similar to the US. We calculate country i's sectoral dissimilarity relative to the US as

$$\text{dissim}_{i,t} = \sum_k |s_{i,k,t} - s_{US,k,t}|\,,$$

where $s_{i,k,t}$ is country i's share of gross output in sector k and in year t. The data is annual and obtained from EU KLEMS and the World Input-Output Database (Timmer et al., 2015).

Fourth, we consider exposure to dollar valuation effects (e.g., Lane and Shambaugh, 2010). As Supplementary Appendix S5 shows, the US dollar tends to appreciate after positive news about US real activity or higher-than-expected prices. Such dollar appreciations raise the value of dollar assets when measured in local currency. Similarly, they raise the cost of repaying dollar liabilities when measured in local currency. The net effect on a countries' balance sheet depends on the net exposure to dollar fluctuations, which is simply the difference between dollar assets $A^{\$}_{i,t}$ and dollar liabilities $L^{\$}_{i,t}$. After scaling this difference by nominal GDP, we have

$$\text{USDnetExp}_{i,t} = \frac{A^{\$}_{i,t} - L^{\$}_{i,t}}{\text{GDP}_{i,t}}.$$

The data to construct this measure is annual and comes from Bénétrix et al. (2019).[10,11]

With these measures in hand, we then estimate the specification

$$
\begin{aligned}
\Delta q_{i,t} = \alpha_i + \gamma^y s^y_{US,t} + \sum_{k \neq y} \gamma^k s^k_{US,t} \\
+ \chi^y_F s^y_{US,t} \text{finInt}_{i,t} + \sum_{k \neq y} \chi^k_F s^k_{US,t} \times \text{finInt}_{i,t} + \theta^y_F \text{finInt}_{i,t} \\
+ \chi^y_T s^y_{US,t} \text{tradeInt}_{i,t} + \sum_{k \neq y} \chi^k_T s^k_{US,t} \times \text{tradeInt}_{i,t} + \theta^y_T \text{tradeInt}_{i,t} \\
+ \chi^y_D s^y_{US,t} \text{dissim}_{i,t} + \sum_{k \neq y} \chi^k_D s^k_{US,t} \times \text{dissim}_{i,t} + \theta^y_D \text{dissim}_{i,t} \\
+ \chi^y_\$ s^y_{US,t} \text{USDnetExp}_{i,t} + \sum_{k \neq y} \chi^k_\$ s^k_{US,t} \times \text{USDnetExp}_{i,t} + \theta^y_\$ \text{USDnetExp}_{i,t} \\
+ \text{controls} + \varepsilon_{i,t}.
\end{aligned}
\tag{S6.3}
$$

For ease of interpretation, we standardize the measures $\text{finInt}_{i,t}$, $\text{USDnetExp}_{i,t}$, $\text{tradeInt}_{i,t}$, and $\text{dissim}_{i,t}$ by first subtracting the sample mean and then by dividing by the sample standard deviation. Hence, the main effect γ^y in equation (S6.3) captures the average response and, for example, the coefficient χ^y_F captures the differential response of a country with a one standard deviation greater-than-average degree of financial integration.

Recall that we documented in Supplementary Appendix S4 that US real activity news often has greater effects on foreign stock markets when the US experiences high unemployment (as measured by the empirical cdf of the US unemployment rate) and when the US economy is doing poorly as measured by the FOMC sentiment index of Gardner, Scotti, and Vega (2022). When estimating specification (S6.3) we include both of these measures and their interaction terms with all surprises as controls. This ensures that the estimates of interest are not driven by potential correlations with these two measures.[12]

Table S6.1 shows the estimates of equation (S6.3). The only interaction coefficient that is consistently significant across announcements is the coefficient on the interaction term of the surprise of interest with trade integration. However, the effect has the unanticipated sign, suggesting that trade integration *reduces* the effect size. The interaction effects of financial integration, industry dissimilarity, and dollar exposure with the surprises of interest do not robustly differ from zero for most announcements.

[10] Assets include portfolio equity, foreign direct investment (equity and debt), portfolio debt, other investment, and reserves. Liabilities include portfolio equity, foreign direct investment (equity and debt), portfolio debt and other investment (see Bénétrix et al., 2019, for details).

[11] Similar to the financial integration measure (S6.1), several countries experience an enormous growth of the exposure measure to dollar fluctuations. Ireland, for instance, reaches a value of over 400 percent in 2017—relative to a mean value of around 19 percent. We have confirmed that the sample restriction to drop Belgium, Ireland, Netherlands, Switzerland, and the United Kingdom, motivated by Figure S6.1, also ensures that the measure $\text{USDnetExp}_{i,t}$ is not extremely right-skewed.

[12] These controls are particularly important for the coefficient on the interaction term of the surprise with trade integration. The trade integration measure (S6.2) is procyclical since both exports and imports are procyclical and more volatile than GDP. It is therefore correlated with both the empirical cdf of the US unemployment rate and the FOMC Sentiment index. When neither of these confounders is controlled for, the coefficient on the interaction term of trade integration is biased downward.

Table S6.1: Heterogeneity in Effect Size (Outliers removed)

	Capacity Utilization	CB Consumer Confidence	Core CPI	Core PPI	Durable Goods Orders	GDP A
Stock Index (bp)						
News	6.76***	11.67***	-11.45***	-5.21***	8.01***	20.03***
	(1.61)	(1.94)	(2.10)	(1.37)	(1.74)	(3.49)
Fin. Integration × News	-1.86*	0.19	1.85	2.26*	-0.74	-5.38
	(1.05)	(1.51)	(1.08)	(1.08)	(1.78)	(3.17)
Trade Integration × News	0.08	-4.20***	0.80	1.52**	-0.97*	-3.69**
	(0.98)	(1.30)	(0.65)	(0.55)	(0.49)	(1.66)
Industry Dissimilarity × News	-1.48	-1.87	-0.60	1.03	-0.17	-4.87
	(1.27)	(1.57)	(0.92)	(0.87)	(1.77)	(2.90)
Dollar Exposure × News	1.02	2.13*	-1.42*	-1.52**	1.10	2.52
	(1.35)	(1.20)	(0.72)	(0.71)	(1.01)	(1.60)
R^2	0.09	0.27	0.24	0.27	0.22	0.51
Observations	3380	3273	3190	3254	3179	1062
	Initial Jobless Claims $\cdot(-1)$	ISM Mfg Index	New Home Sales	Nonfarm Payrolls	Retail Sales	UM Consumer Sentiment P
Stock Index (bp)						
News	3.48***	13.12***	6.44***	18.70***	7.95***	6.39***
	(0.72)	(2.92)	(1.86)	(3.23)	(1.73)	(1.66)
Fin. Integration × News	0.68	2.15	2.58**	7.12**	0.26	-0.63
	(0.64)	(2.16)	(1.15)	(3.17)	(1.38)	(1.41)
Trade Integration × News	-1.00**	-2.26	-2.12**	-3.38*	-1.32	-1.40
	(0.47)	(1.35)	(0.96)	(1.84)	(0.76)	(0.84)
Industry Dissimilarity × News	0.14	-0.19	2.46***	-0.18	-0.04	-0.99
	(0.62)	(1.92)	(0.77)	(2.18)	(1.37)	(1.39)
Dollar Exposure × News	1.00*	0.81	0.91	2.08	0.54	0.46
	(0.54)	(2.01)	(0.78)	(2.26)	(0.89)	(1.13)
R^2	0.19	0.25	0.13	0.36	0.31	0.08
Observations	13767	2996	3210	3163	3241	3308

Notes: This table presents estimates of γ^y, χ^y_F, $\chi^y_\$$, χ^y_T, and χ^y_D from equation (S6.3). The sample excludes Belgium, Ireland, Netherlands, Switzerland, and the United Kingdom. The various exposure measures are defined in the text. Standard errors are two-way clustered by announcement and by country, and reported in parentheses. ***, **, and * indicate significance at the 1, 5, and 10 percent level.

It turns out that the negative correlation of the effect size with trade integration is not robust across alternative specifications. Table S6.2 shows the estimates of specification (S6.3) after replacing the average main effect γ^y with a country-specific effect γ^y_i (and similarly for the controls where we replace γ^k with γ^k_i for all k). This modification addresses endogeneity concerns arising from the possibility that the confounding variation is country-specific and time-invariant. As the table shows,

Table S6.2: Heterogeneity in Effect Size (Outliers removed & Country-specific Main Effects)

	Capacity Utilization	CB Consumer Confidence	Core CPI	Core PPI	Durable Goods Orders	GDP A
Stock Index (bp)						
Fin. Integration						
× News	-9.12**	0.35	3.83	1.46	-1.32	-6.91
	(3.83)	(4.39)	(3.63)	(2.47)	(3.35)	(6.49)
Trade Integration						
× News	7.57	-10.24	-4.80	3.82	-4.86	3.17
	(5.37)	(6.19)	(4.56)	(3.06)	(2.87)	(2.98)
Industry Dissimilarity						
× News	-2.82	5.65	0.50	0.48	3.19	-3.45
	(3.63)	(5.83)	(5.16)	(2.06)	(4.40)	(7.86)
Dollar Exposure						
× News	4.91***	1.33	-1.81	-0.76	0.86	-0.46
	(1.69)	(1.92)	(1.37)	(1.22)	(1.42)	(2.29)
R^2	0.11	0.29	0.25	0.28	0.23	0.53
Observations	3380	3273	3190	3254	3179	1062
	Initial Jobless Claims $\cdot(-1)$	ISM Mfg Index	New Home Sales	Nonfarm Payrolls	Retail Sales	UM Consumer Sentiment P
Stock Index (bp)						
Fin. Integration						
× News	2.06	1.02	8.51**	21.09***	1.15	0.19
	(1.31)	(5.11)	(3.54)	(5.30)	(3.07)	(4.37)
Trade Integration						
× News	0.96	5.80	-1.76	0.07	10.97**	-11.50*
	(2.05)	(6.74)	(4.94)	(6.07)	(4.55)	(5.95)
Industry Dissimilarity						
× News	1.35	-0.57	3.01	2.35	-4.34	3.71
	(1.36)	(3.90)	(2.91)	(6.31)	(3.37)	(3.76)
Dollar Exposure						
× News	-0.24	-0.66	-1.41	-6.01**	-2.05	1.22
	(0.80)	(1.70)	(0.99)	(2.35)	(1.56)	(2.05)
R^2	0.19	0.27	0.14	0.38	0.32	0.09
Observations	13767	2996	3210	3163	3241	3308

Notes: This table presents estimates of χ_F^y, $\chi_\$^y$, χ_T^y, and χ_D^y obtained from equation (S6.3) after replacing the main effects on the surprises γ^y and γ^k with country-specific main effects γ_i^y and γ_i^k. The sample excludes Belgium, Ireland, Netherlands, Switzerland, and the United Kingdom. The various exposure measures are defined in the text. Standard errors are two-way clustered by announcement and by country, and reported in parentheses. ***, **, and * indicate significance at the 1, 5, and 10 percent level.

the interaction term with trade integration looses its significant coefficient in all but two instances. More generally, no interaction effect in Table S6.2 differs significantly from zero systematically across announcements.

The conclusion from this appendix is that it is difficult to systematically account for the variation in effect size as documented in Figure 4 with observables. While some interaction effects are highly statistically significant and economically large for individual announcements (see, e.g., the interaction effect on the product of financial integration and the nonfarm payrolls surprise in Table S6.2), no consistent patterns emerge that are robust across announcements. Of course, this lack of a consistent pattern does not rule out the existence of any of the four channels studied here. However, it does imply that they are not sufficiently salient to be statistically detectable in our sample. In our view, understanding the heterogeneity in effect size across countries is an interesting topic for future research. Specifically, studying a broader set of channels and alternative measures for a given channel may lead to useful insights.

References

Alquist, Ron, Saroj Bhattarai, and Olivier Coibion. 2019. "Commodity-price comovement and global economic activity." *Journal of Monetary Economics* .

Altavilla, Carlo, Luca Brugnolini, Refet S Gürkaynak, Roberto Motto, and Giuseppe Ragusa. 2019. "Measuring euro area monetary policy." *Journal of Monetary Economics* 108:162–179.

Altavilla, Carlo, Domenico Giannone, and Michele Modugno. 2017. "Low frequency effects of macroeconomic news on government bond yields." *Journal of Monetary Economics* 92:31 – 46.

Andersen, Torben G., Tim Bollerslev, Francis X. Diebold, and Clara Vega. 2003. "Micro Effects of Macro Announcements: Real-Time Price Discovery in Foreign Exchange." *American Economic Review* 93 (1):38–62.

———. 2007. "Real-time price discovery in global stock, bond and foreign exchange markets." *Journal of International Economics* 73 (2):251 – 277.

Auerbach, Alan J and Yuriy Gorodnichenko. 2012. "Measuring the output responses to fiscal policy." *American Economic Journal: Economic Policy* 4 (2):1–27.

Bastourre, Diego, Jorge Carrera, Javier Ibarlucia, and Mariano Sardi. 2012. "Common drivers in emerging market spreads and commodity prices." Tech. rep., Working Paper.

Bénétrix, Agustín, Deepali Gautam, Luciana Juvenal, and Martin Schmitz. 2019. *Cross-border currency exposures*. International Monetary Fund.

Boehm, Christoph and Niklas Kroner. 2021. "Beyond the Yield Curve: Understanding the Effect of FOMC Announcements on the Stock Market." *Available at SSRN 3812524* .

Boehm, Christoph E. 2020. "Government consumption and investment: Does the composition of purchases affect the multiplier?" *Journal of Monetary Economics* 115:80–93.

Boehm, Christoph E., Aaron Flaaen, and Nitya Pandalai-Nayar. 2019. "Input Linkages and the Transmission of Shocks: Firm-Level Evidence from the 2011 Tōhoku Earthquake." *The Review of Economics and Statistics* 101 (1):60–75.

Boyd, John H, Jian Hu, and Ravi Jagannathan. 2005. "The stock market's reaction to unemployment news: Why bad news is usually good for stocks." *The Journal of Finance* 60 (2):649–672.

Bruno, Valentina and Hyun Song Shin. 2015. "Cross-border banking and global liquidity." *The Review of Economic Studies* 82 (2):535–564.

Byrne, Joseph P., Giorgio Fazio, and Norbert Fiess. 2013. "Primary commodity prices: Co-movements, common factors and fundamentals." *Journal of Development Economics* 101:16 – 26.

Cesa-Bianchi, Ambrogio, Gregory Thwaites, and Alejandro Vicondoa. 2020. "Monetary policy transmission in the United Kingdom: A high frequency identification approach." *European Economic Review* 123:103375.

Chinn, Menzie D. and Olivier Coibion. 2014. "The Predictive Content of Commodity Futures." *Journal of Futures Markets* 34 (7):607–636.

Cragg, John G and Stephen G Donald. 1997. "Inferring the rank of a matrix." *Journal of econometrics* 76 (1-2):223–250.

Di Giovanni, Julian and Andrei A Levchenko. 2010. "Putting the parts together: trade, vertical linkages, and business cycle comovement." *American Economic Journal: Macroeconomics* 2 (2):95–124.

Eichenbaum, Martin and Charles L Evans. 1995. "Some empirical evidence on the effects of shocks to monetary policy on exchange rates." *The Quarterly Journal of Economics* 110 (4):975–1009.

Elenev, Vadim, Tzuo Hann Law, Dongho Song, and Amir Yaron. 2022. "Fearing the fed: How wall street reads main street." *Available at SSRN 3092629* .

Etula, Erkko. 2013. "Broker-dealer risk appetite and commodity returns." *Journal of Financial Econometrics* 11 (3):486–521.

Frankel, Jeffrey A and Andrew K Rose. 1998. "The endogenity of the optimum currency area criteria." *The Economic Journal* 108 (449):1009–1025.

Gardner, Ben, Chiara Scotti, and Clara Vega. 2022. "Words speak as loudly as actions: Central bank communication and the response of equity prices to macroeconomic announcements." *Journal of Econometrics* 231 (2):387–409.

Gerko, Elena and Hélene Rey. 2017. "Monetary policy in the capitals of capital." *Journal of the European Economic Association* 15 (4):721–745.

Goldberg, Linda S and Christian Grisse. 2013. "Time variation in asset price responses to macro announcements." Tech. rep., National Bureau of Economic Research.

Gorton, Gary and K Geert Rouwenhorst. 2006. "Facts and fantasies about commodity futures." *Financial Analysts Journal* 62 (2):47–68.

Gourinchas, Pierre-Olivier, Hélène Rey, and Maxime Sauzet. 2019. "The international monetary and financial system." *Annual Review of Economics* 11:859–893.

Gürkaynak, Refet, Brian Sack, and Eric Swanson. 2005. "Do Actions Speak Louder Than Words? The Response of Asset Prices to Monetary Policy Actions and Statements." *International Journal of Central Banking* 1 (1).

Gürkaynak, Refet S, Burçin Kısacıkoğlu, and Jonathan H Wright. 2020. "Missing Events in Event Studies: Identifying the Effects of Partially Measured News Surprises." *American Economic Review* 110 (12):3871–3912.

Gürkaynak, Refet S, Brian Sack, and Jonathan H Wright. 2007. "The US Treasury yield curve: 1961 to the present." *Journal of monetary Economics* 54 (8):2291–2304.

Huo, Zhen, Andrei A Levchenko, and Nitya Pandalai-Nayar. 2020. "International comovement in the global production network." .

Imbs, Jean. 2004. "Trade, Finance, Specialization, and Synchronization." *The Review of Economics and Statistics* 86 (3):723–734.

Jarociński, Marek and Peter Karadi. 2020. "Deconstructing Monetary Policy Surprises—The Role of Information Shocks." *American Economic Journal: Macroeconomics* 12 (2):1–43.

Jiang, Zhengyang, Arvind Krishnamurthy, and Hanno Lustig. 2020. "Dollar safety and the global financial cycle." Tech. rep., National Bureau of Economic Research.

Kilian, Lutz and Clara Vega. 2011. "Do energy prices respond to US macroeconomic news? A test of the hypothesis of predetermined energy prices." *Review of Economics and Statistics* 93 (2):660–671.

Kurov, Alexander and Raluca Stan. 2018. "Monetary policy uncertainty and the market reaction to macroeconomic news." *Journal of Banking & Finance* 86:127–142.

Lane, Philip R and Gian Maria Milesi-Ferretti. 2007. "The external wealth of nations mark II: Revised and extended estimates of foreign assets and liabilities, 1970–2004." *Journal of international Economics* 73 (2):223–250.

———. 2017. "International financial integration in the aftermath of the global financial crisis." IMF Working Paper 17/115, International Monetary Fund.

Lane, Philip R and Jay C Shambaugh. 2010. "Financial exchange rates and international currency exposures." *American Economic Review* 100 (1):518–40.

McQueen, Grant and V Vance Roley. 1993. "Stock prices, news, and business conditions." *The Review of Financial Studies* 6 (3):683–707.

Miranda-Agrippino, Silvia. 2016. "Unsurprising shocks: information, premia, and the monetary transmission." .

Miranda-Agrippino, Silvia and Tsvetelina Nenova. 2022. "A tale of two global monetary policies." *Journal of International Economics* 136:103606.

Miranda-Agrippino, Silvia and Hélene Rey. 2020. "US monetary policy and the global financial cycle." *The Review of Economic Studies* 87 (6):2754–2776.

Pindyck, Robert S and Julio J Rotemberg. 1990. "The Excess Co-Movement of Commodity Prices." *The Economic Journal* 100 (403):1173–1189.

Ramey, Valerie A and Sarah Zubairy. 2018. "Government spending multipliers in good times and in bad: evidence from US historical data." *Journal of political economy* 126 (2):850–901.

Rey, Hélène. 2016. "International channels of transmission of monetary policy and the Mundellian trilemma." *IMF Economic Review* 64 (1):6–35.

Rigobon, Roberto. 2003. "Identification through heteroskedasticity." *Review of Economics and Statistics* 85 (4):777–792.

Swanson, Eric T. 2021. "Measuring the effects of federal reserve forward guidance and asset purchases on financial markets." *Journal of Monetary Economics* .

Timmer, Marcel P, Erik Dietzenbacher, Bart Los, Robert Stehrer, and Gaaitzen J De Vries. 2015. "An illustrated user guide to the world input–output database: the case of global automotive production." *Review of International Economics* 23 (3):575–605.