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Unconventional monetary policy and disaster risk: Evidence from the subprime and COVID–19 crises

Gustavo S. Cortes a,⁎, George P. Gao b, Felipe B. G. Silva c, Zhaogang Song d

a Warrington College of Business, University of Florida, 306 Stuzin Hall, PO Box 117168, Gainesville 32611-7168, FL, USA
b Quantitative Equities, T. Rowe Price, 100 E. Pratt St., Baltimore 21202, MD, USA
c Trulaske College of Business, University of Missouri, 425 Cornell Hall, Columbia 65211, MO, USA
d Johns Hopkins University, Carey Business School, 100 International Drive, Baltimore 21202, MD, USA

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Abstract

We compare the interventions conducted by the Federal Reserve in response to the subprime and COVID–19 crises with respect to their effectiveness in reducing disaster risk. Using model-free measures of disaster risk derived from daily options data, we document that interventions in response to both crises reduced tail risks in domestic equity markets. The spillover effects of the two crises have been markedly dissimilar. While subprime interventions are generally characterized by negative spillovers to international equity markets, policy responses to the COVID–19 crisis are generally associated with positive spillovers. We interpret these results as consistent with the different degrees of protagonism by central banks in the two episodes, emphasizing the importance of a broader participation of monetary authorities in expanding their balance sheets to counteract the effects of major crises.

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Corresponding author at: Warrington College of Business, University of Florida, 306 Stuzin Hall, PO Box 117168, Gainesville, FL 32611-7168, USA.
E-mail address: gustavo.cortes@warrington.ufl.edu (G.S. Cortes).

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“International monetary cooperation has broken down (…) The United States should worry about the effects of its policies on the rest of the world.”
— RAGHRAM G. RAJAN (2014), GOVERNOR OF THE RESERVE BANK OF INDIA

“The problem is that the previously weak speculative forces have gathered astonishingly large amount of liquidity after several rounds of QE by the major economies.”
— ZHOU XIAOCUHAN (2016), GOVERNOR OF THE PEOPLE’S BANK OF CHINA

“[The crisis] will not be overcome simply through austerity measures (…) let alone through QE policies that have triggered a ‘monetary tsunami,’ have led to a currency war, and introduced new and perverse forms of protectionism.”
— DILMA ROUSSEFF (2012), PRESIDENT OF BRAZIL
1. Introduction

Financial crises have required central banks throughout history and around the world to break the cycle of heightened risk of distress in financial markets, liquidity freezes, and declines in the real economy (Reinhart and Rogoff, 2009; Eichengreen, 2013; Aizenman, 2016). In addition to conventional monetary policies, the US and other advanced economies have also adopted new measures (e.g., large-scale asset purchases, forward-guidance) to overcome the limits on traditional policies imposed by the zero lower bound on short-term interest rates. Despite being historically rare vis-à-vis systematic interest rate policy, quantitative easing (QE) interventions were used by the Federal Reserve as early as in 1932 (e.g., Jaremski and Mathy, 2018) and between 2008 and 2014, during the Great Recession. As Bernanke (2020) advocates, unconventional monetary policy (UMP) tools should become part of the standard central bank toolkit and are expected to be even more popular in a world of chronically low interest rates.

Notwithstanding favorable arguments, the Fed’s response between 2008 and 2014 was also severely criticized by central bank governors and policy makers from major emerging market economies (EMEs). In particular, the unilateral expansion of the balance sheet of a major central bank (i.e., the Fed) relative to its peers could distort exchange rates depreciating the US dollar (Dedola et al., 2020), thereby affecting capital and trade flows. Equilibrium distortions artificially driven by rounds of QE interventions could subsequently cause a gamut of undesirable effects that are particularly harmful to EMEs, such as protectionist policies, asset market bubbles, and excessive leverage of households and firms (e.g., Aizenman et al. (2016, 2017)). Hence, a natural question is whether a multilateral expansion of the central banks’ balance sheets can mitigate these welfare losses.

In this paper, we shed light on potential welfare gains attained from a multilateral implementation of unconventional monetary policies by contrasting the subprime crisis and the ongoing recession caused by the COVID–19 pandemic. As discussed by Aizenman (2016) and Frankel (2016), despite being rare throughout history, a coordination equilibrium may occur when bad tail events induce imminent threats of financial collapse and first-order perceptions of a global depression. For such coordination to mitigate disaster risk, investors should simply take the collective actions (or expectations thereof) as credible signals that exchange rates and flows will not be substantially distorted when a broader variety of central banks actively intervene. We take these insights into consideration and formally test how tail risk perceptions react to monetary interventions in both crises.

To study the effectiveness of monetary interventions for financial stability, we use options-based measures of extreme-tail risk. Specifically, we gauge extreme–left tail expectations (disaster risk, hereafter) and extreme-right tail expectations (swift recovery, hereafter) using options written on ETFs, the underlying assets of which fall into several classes (Gao et al., 2018; Gao et al., 2019). These asset classes include domestic equity markets, international equity markets, treasuries, and domestic equity sectors.

The advantages of using option-based proxies vis-à-vis identifying instances of extreme price oscillations for underlying assets are manifold. The common approach of measuring disaster risk based on the actual frequency of extreme price movements of a given asset over a period of time bears important limitations. First, it typically yields a small number of observations given the rareness of such market crashes. Moreover, it requires an ad-hoc definition of a threshold for price movements to be deemed "extreme" (e.g., classify stock returns worse than −5% as "market crashes"), yielding an ex-post measure of tail risk. Because risk is conceptually an ex-ante phenomenon, the translation from ex-post realizations of market crashes to ex-ante expectations of risk relies on assumptions of a rational expectations equilibrium, which may be overly stringent when applied to assessments of rare disasters. Last, because this method requires the specification of a time period for which the frequency of realized returns would be evaluated, the time granularity of ex-post measures is typically coarse (e.g., a quarter or a year).

Conversely, using option prices allows us to construct a daily time series of ex-ante disaster risk that does not require subjective threshold definitions of what constitutes a market crash. Also, as in Carr and Wu (2009), the variance swap replicas are essentially model-free. Instead of imposing strong functional (parametric) assumptions, it simply relies on non-parametric estimates using option market prices. Finally, the daily frequency of our measures allows us to estimate the abnormal changes in investors’ perceptions of disaster risk surrounding key UMP announcements.

While the 2008 and the 2020 crises were inherently different in their causes, they share a combination of supply and demand effects detrimental to global economic activity. Our emphasis in this comparison, however, lies on the corresponding policy responses. Notwithstanding the use of unconventional monetary policies in both crises, the rounds of QE during the subprime crisis can be better characterized as an almost unilateral response by a protagonistic central bank (the Fed). Conversely, the response to the COVID–19 crisis was diversified. Starting in March 2020, the Fed and 20 other central banks put forth unconventional monetary policy measures to respond to the severe recession caused by the COVID–19 pandemic (see, e.g., Daehler et al., 2020; Jinjarak et al., 2021; Hartley et al., 2021). Not only did several central banks proactively...
intervene in their own domestic markets but some even engaged in greater global coordination.\(^3\) Fig. 1 depicts the evolution of major central banks’ balance sheets (normalized to 100 in January 2008), illustrating the use of UMP in response to the two crises in question. While in 2008 the response was almost exclusively driven by the Fed and the Bank of England, the 2020 crisis saw an expansion to include all major central banks (with special attention to the Bank of Canada).

Comparing the Fed’s response to the subprime crisis relative to COVID–19, we draw important lessons for the US and global economies. First, focusing on the domestic front, we find that the announcement of unconventional monetary policies by the Fed in both episodes effectively reduced perceptions of disaster risk for US (domestic) equities. We also find that the Fed’s announcements in the subprime crisis reduced the disaster risk of mid-maturity US treasuries while only marginally affecting the disaster risk for longer maturity bonds.\(^4\) Interestingly, we find that the policy response to the COVID–19 crisis had mixed effects on disaster risk for mid-maturity bonds and meaningful increases in disaster risk for longer maturity bonds. Delving into different sectors of the US domestic equity market, we find that the reduction in disaster risk following the Fed’s policy announcements in response to the COVID–19 crisis reveals a significant effect on non-financial sectors like retail and technology, accompanied by financial intermediaries such as banks.

We then analyze the international spillovers of the Fed’s interventions to global equity markets. As widely documented, there is a strong correlation between the financial conditions in the United States and in international markets (Aizenman et al. (2016, 2017); Iacoviello and Navarro (2019)). With this in mind, we estimate the spillover effects stemming from the policy interventions as the abnormal change in the co-movement that is typically observed across markets. Since the intervention dates are associated with reductions in disaster risk in the US, a positive abnormal co-movement with another country implies that the latter also experienced the desirable effect of the intervention—an effect directionally aligned with the disaster risk reduction in the US market. Conversely, a negative abnormal co-movement between the disaster risk of the US and another country indicates that the country’s disaster risk is moving abnormally in a direction contrary to the desirable US domestic effect—characterizing an undesirable spillover.

A perusal of cross-border effects of advanced economies (AEs) and emerging market economies (EMEs) in both episodes underscores key differences. Comparing disaster risk from baskets of AE equities and EME equities following the Fed’s response to the subprime crisis, we document an undesirable spillover effect for both country groups. Broadly speaking, the subprime crisis interventions are associated with abnormal co-movements between the disaster risk of US equities and the equities of AEs and EMEs in the opposite direction of the desirable domestic effects—the only exception being the longer, 6-month maturity of EME equities. To some extent, these results corroborate the concerns of undesired spillovers voiced by policy-makers outside the US. Conversely, the effects for the same AE and EME baskets during the COVID–19 interventions point to a desirable spillover of Fed policy for both country groups. The Fed’s policy announcements during COVID–19 led to abnormal tail risk co-movements for both AEs and EMEs directionally aligned with the intended domestic effects of the intervention, with pronounced effects on short-term maturities of EMEs. Narrowing down to equity markets of individual countries, we observe significantly positive disaster-risk spillovers, mostly concentrated in Asia (e.g., South Korea, Taiwan, and Malaysia).

In short, unconventional monetary policies successfully shifted downward the term structure of disaster risk for domestic equities in both crises. The spillover effects were more pronounced and aligned with a decline in disaster risk for the COVID–19 episode.

Finally, we characterize our international spillover results exploiting cross-country heterogeneity. We focus on the dimensions of the “international finance trilemma” established by an extensive literature (see Aizenman et al., 2010; Iacoviello and Navarro, 2019).
Aizenman et al., 2013). This literature documents how countries’ monetary policy independence, exchange rate stability, and financial openness modulate their exposure to international shocks. Taking these considerations into our international analysis, we find that cross-border spillovers were more pronounced for countries with greater monetary policy independence, exchange rate stability, and financial openness. Probing further, we also show that the magnitudes of the spillover effects are contingent on different dimensions of the countries’ fiscal space (Kose et al., 2017).

Our paper contributes to several strands of literature. First, we add to the literature on the international effects of US monetary policy. Aizenman et al. (2016, 2017) document how monetary and financial shocks in center economies are transmitted to the global economy, with significant macroeconomic repercussions on EMEs. Relatedly, Bauer and Neely (2014) and Banerjee et al. (2016) investigate the channels through which the Fed’s “self-centered” unconventional monetary policy affects other countries. Our paper examines a relatively unexplored dimension of international monetary policy spillovers, i.e., the perceptions of financial market investors regarding the occurrence of tail risk events in response to Fed interventions.

We also add to the literature on the effects of quantitative easing on disaster risks implied by option prices. Hattori et al. (2016) investigate the effect of the Fed’s 2008–14 QE interventions on investors’ tail risk perceptions in the S&P 500 option market and 10-year rates implied by the prices of swaptions with two years to maturity. Relative to their work, we further the understanding of tail risk and monetary policy along four dimensions. First, in addition to domestic effects of QE policies, our work investigates spillover effects of unconventional monetary policy tools to the risk of disaster in a rich set of advanced and emerging market economies. Second, regarding domestic effects, we analyze market expectations of both rare disasters and swift recoveries. Given the inherent goal of such policies of countering the negative effects of crises, our investigation sheds light not only on the effectiveness of such tools in curtailting the risk of a disaster (proxied by our rare-disaster indices), but also the expectations of an extremely fast recovery (represented by our swift-recovery indices) in asset markets. Third, we analyze a richer set of asset classes including sector equity index ETFs, allowing us to delve into the heterogeneous effects of the interventions across sectors in each crisis. Fourth, our longer time span allows us to compare the impact of the interventions put in place after 2008 with the Fed’s recent response to the COVID–19 pandemic starting in March 2020.

Lastly, our paper extends the burgeoning literature on the economic effects of epidemics following the ongoing COVID–19 global pandemic. A non-exhaustive list includes Correia et al. (2020), Coibion et al. (2020a), Coibion et al. (2020b), Schrimpf et al. (2020), Sheridan et al. (2020), Eichenbaum et al. (2021), and Spiegel (2021).

The closest paper to ours is Hartley et al. (2021). The authors study the COVID–19 QE announcements made by 21 global central banks and their effect on each country’s 10-year government bond yields, documenting that the average advanced economy QE announcement had a slightly smaller effect than past interventions during the Great Recession era. In contrast, the impact of QE announcements on the average emerging market was significantly larger. Our paper differs from their work in three dimensions. First, while they focus on the domestic effects of the intervention by the central bank in each country, we center our analysis on the Federal Reserve’s interventions and their international consequences. Second, we focus on tail risks of many different asset markets, while their analysis is centered on government bond yields. Despite these distinctions, Hartley et al.’s (2021) results are complementary to ours, as both indicate a greater sensitivity of EMEs in response to large-scale interventions—either domestically or internationally driven. Third, in addition to their insights across countries, we break down the US economy into sector ETFs. This allows us to paint a more complete picture of the domestic effects of the Fed’s interventions.

The remainder of the paper is organized as follows. Section 2 discusses the data, the methodology used to measure tail risk, and the main specifications of our empirical analysis. Section 3 presents the results of the domestic effects of the Fed’s interventions during the subprime and the COVID–19 crises. It also presents the spillover effects of each intervention. Section 4 concludes.

2. Empirical strategy

2.1. Measuring rare-disaster risks

A voluminous literature relies on the information summarized in option prices to construct measures of expected volatility of underlying assets (e.g., Hattori et al., 2016; Gao et al., 2018). To construct our proxies for expectations of abnormally large price movements for various asset classes, we follow the methodology employed by Gao et al. (2018).

The theoretical justification for the construction of ex-ante measures of high-order price movements is in the model-free implied volatility measures proposed by Carr and Madan (2001), Britten-Jones and Neuberger (2000), and Carr and Wu (2009). To isolate high-order (≥ 3) distributional moments, Gao et al. (2018) consider the difference between two variance-swap replicating portfolios: the first one accounting for mild oscillation in underlying asset prices and the second capturing large price jumps.

Consider an underlying asset whose present time t price is given by $S_t$. A variance-swap contract written for the period of $[t, T]$ is a derivative instrument whose net payoff at expiration $T$ is the difference between the realized variance (i.e., squared volatility) of $S_t$ and the expected volatility (paid by the holder upon entering the contract). The economic rationale for entering a long position in a variance swap is to hedge against price fluctuations over a time interval.
For a given time-to-maturity $\tau \equiv T - t$, such an instrument can be replicated using a continuum of out-of-the-money call options $P(S_t; K, T)$ (i.e., a continuum of strike values $K$), yielding two possible replicating portfolios:

$$\mathbb{V}^S(t; \tau) = \frac{2e^{r\tau}}{\tau} \int_{K<S_t} \frac{1}{K^2} C(S_t; K, T) dK + \frac{2e^{r\tau}}{\tau} \int_{K>S_t} \frac{1}{K} P(S_t; K, T) dK$$  \hspace{1cm} (1)$$

$$\mathbb{V}^S(t; \tau) = \frac{2e^{r\tau}}{\tau} \int_{K<S_t} \frac{1}{K^2} \log(K/S_t) C(S_t; K, T) dK + \frac{2e^{r\tau}}{\tau} \int_{K>S_t} \frac{1}{K^2} \log(K/S_t) P(S_t; K, T) dK$$  \hspace{1cm} (2)$$

The key difference between the aforementioned replication portfolios is that $\mathbb{V}^S(t; \tau)$ assigns larger (smaller) weights to more deeply out-of-the-money put (call) options than $\mathbb{V}^S(t; \tau)$. Consequently, the spread term $\mathbb{V}^S(t; \tau) - \mathbb{V}^S(t; \tau)$ should significantly isolate large deviations of $S_t$ from mild price movements, collectively accounting for the likelihood of a rare disaster (abnormally large downward movements) and swift recoveries (abnormally large upward movements, or price jumps). We consider the downward-only ($\mathbb{V}^{(S^-)}(t; \tau)$ and $\mathbb{V}^{(S^-)}(t; \tau)$) and upward-only variants ($\mathbb{V}^{(S+)}(t; \tau)$ and $\mathbb{V}^{(S+)}(t; \tau)$) of the replicating portfolios to construct our "swift-disaster index" for different asset classes:

$$\mathbb{V}^{(S^-)}(t; \tau) = \frac{2e^{r\tau}}{\tau} \int_{K<S_t} \frac{1}{K^2} P(S_t; K, T) dK$$

$$\mathbb{V}^{(S^-)}(t; \tau) = \frac{2e^{r\tau}}{\tau} \int_{K<S_t} \frac{1}{K^2} \log(K/S_t) P(S_t; K, T) dK$$

$$\mathbb{V}^{(S+)}(t; \tau) = \frac{2e^{r\tau}}{\tau} \int_{K<S_t} \frac{1}{K^2} C(S_t; K, T) dK$$

$$\mathbb{V}^{(S+)}(t; \tau) = \frac{2e^{r\tau}}{\tau} \int_{K<S_t} \frac{1}{K^2} \log(K/S_t) C(S_t; K, T) dK$$

As in Gao et al. (2018), the aforementioned rare-disaster proxy constitutes our main outcome variable of interest, as expectations of market crashes represent important state variables during major crises. Also, curtailing the probabilities of such undesirable events is arguably an important objective of monetary authorities. However, expectations of swift recoveries can be also informative, particularly in a global pandemic which can be effectively controlled by means of biomedical developments (e.g., vaccines). Thus, we employ the same logic to define the following "swift-recovery index" (i.e., right-tail proxies):

$$\mathbb{V}^{(S+)}(t; \tau) = \frac{2e^{r\tau}}{\tau} \int_{K<S_t} \frac{1}{K^2} C(S_t; K, T) dK$$

$$\mathbb{V}^{(S+)}(t; \tau) = \frac{2e^{r\tau}}{\tau} \int_{K<S_t} \frac{1}{K^2} \log(K/S_t) C(S_t; K, T) dK$$

$$\mathbb{V}^{(S+)}(t; \tau) = \frac{2e^{r\tau}}{\tau} \int_{K<S_t} \frac{1}{K^2} C(S_t; K, T) dK$$

$$\mathbb{V}^{(S+)}(t; \tau) = \frac{2e^{r\tau}}{\tau} \int_{K<S_t} \frac{1}{K^2} \log(K/S_t) C(S_t; K, T) dK$$

The advantages of relying on option prices to construct rare-disaster and swift-recovery indices are evident from the expressions above: $\mathbb{R}X^{(S^+)}(t; \tau)$ and $\mathbb{R}X^{(S^-)}(t; \tau)$ can be computed for every date $t$, as long as option prices $C(S_t; K, T)$ and $P(S_t; K, T)$ embed expectations regarding the underlying asset $S$. From a practical perspective, we numerically estimate the tail risk measure for options written on a variety of liquid ETFs, which include international equity indexes, and sector ETFs. These measures can be construed as tail risk proxies for a given market, represented by different underlying ETFs. Following Gao et al. (2018), our rare-disaster and swift-recovery measures are normalized respectively by $\mathbb{V}^{(S^-)}(t; \tau)$ and $\mathbb{V}^{(S+)}(t; \tau)$, yielding $\mathbb{R}X^{(S^+)}_t = \mathbb{R}X^{(S^+)}(t; \tau) / \mathbb{V}^{(S+)}(t; \tau)$ and $\mathbb{R}X^{(S^-)}_t = \mathbb{R}X^{(S^-)}(t; \tau) / \mathbb{V}^{(S^-)}(t; \tau)$ for different underlying ETFs $S \in \{1, 2, \ldots, N^S\}$. The algorithm performed to approximate the continuum of strikes with a finite grid and synthetically construct the integrand functions for a given time-to-maturity parameter is outlined in Appendix A.

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5 Despite the theoretical advantages of using option-implied disaster risk measures, their usefulness and informativeness depend on the existence of informationally efficient option markets written on the underlying ETFs representing a given asset class. As such, our proxies are subject to estimation errors resulting from a combination of (i) the typical tracking errors associated with ETF prices (cf. Bodurtha et al. [1995, 2006, 2020]); (ii) informational inefficiencies associated with ETF-option prices; and (iii) the numerical approximation of the integrals over a continuum of strike prices needed to estimate the $\mathbb{R}X$ proxies. Thus, for the purpose of our empirical analyses, we stress that neither the markets for the underlying ETFs nor the markets for the corresponding ETF options are equally efficient. This implies that inferences drawn from measures constructed for less liquid instruments (e.g., measures proxying for the disaster risk of some US sectors and some individual countries) should be interpreted with greater caution. Importantly, this caveat is not particularly concerning for the main markets that we focus on in this study: the US equity market, and equity indices constructed from groups of advanced economies and emerging-market economies.
2.2. Data

2.2.1. Option prices and characteristics
Daily market prices and contract characteristics of a variety of ETF options with different strikes and expiration dates are the fundamental inputs to construct our daily series of $RIX_t^{(S_t, -)}$ and $RIX_t^{(S_t, +)}$. We obtain ETF options data from OptionMetrics on US equity indices and treasuries, US industry sector indices, and international equity indices.\(^6\)

The dynamics of heightened credit risk at the onset of COVID–19 pandemic is evidenced by spikes in the rare-disaster index for the banking sector and high-yield corporate bonds in Fig. 2. Anecdotally, time series plots in Fig. 3 corroborate general perceptions on how countries were differentially affected by the COVID–19 pandemic. For example, hard-hit countries such as Italy and Spain show swift increases in disaster risk between February and March, coinciding with the global spread of the epidemic. Notably, levels of rare-disaster and swift-recovery expectations, albeit serially correlated, are asymmetric in magnitude. Such patterns contrast with the rare-disaster-risk dynamics for countries/jurisdictions which, at least during the first waves, successfully curbed the epidemic (e.g., Singapore, Hong Kong, Taiwan).

Figs. 2 and 3 depict the time series dynamics of the rare-disaster and swift-recovery measures constructed from daily option prices and spanning the period from January 2020 until June 2021, respectively for US domestic sectors and equities from different countries.\(^7\) Series are presented in an “iceberg” format, with the inverted (normal) y-axis orientation for the rare-disaster (swift-recovery) proxies represented in red (blue). At this juncture, it is important to note that the two indices are strongly correlated. This is consistent with the notion that whereas rare disasters and fast recoveries capture different underlying concepts, the likelihood of the two tail events is interdependent. Differently put, a greater likelihood of rare disasters makes a V-shaped recovery more plausible, since a swift recovery typically follows the actual realization of a market crash. Anecdotally, market rallies and asset price bubbles affect the likelihood of a market crash (Schularick and Taylor, 2012; Baron and Xiong, 2017).

2.2.2. Macroeconomic surprises
We use daily composite series of macroeconomic surprises for five important markets: the US, the Eurozone, China, Japan, and the UK. The surprise series are compiled by Citigroup, taking into consideration the actual disclosure of a macroeconomic aggregate vis-à-vis forecast consensus by professional macroeconomic analysts. More specifically, the Citi Economic Surprise Index (ESI) measures whether data releases from an economy or group of economies have beaten or missed expectations in the past 90 calendar days. A positive index reading means that releases have been better than expected and a negative reading means that releases have been worse than expected. The index is measured in basis points of aggregated and decay-adjusted standard deviations of surprises and has no natural bounds. Importantly, the components associated with different economies and regions follow the same calculation methodology, rendering the ESI useful for comparison purposes.

2.2.3. COVID–19 controls
To account for the effects of the COVID–19 pandemic on risk perceptions of economic agents, we collect data on daily infections and fatalities compiled by Johns Hopkins University’s Coronavirus Resource Center. Individual records are aggregated at the country level.

2.3. Time-series properties of disaster-risk and swift-recovery indices

To illustrate some properties of our measures, we provide more details on the time series of our tail-risk indices. We first show the volatility of changes to our indices as measured by the monthly standard deviation aggregated from daily $\Delta RIX_t$ data. Fig. 4 presents the evolution of the volatility of changes in rare-disaster ($\Delta RIX_t^{(S_t, -)}$) and swift-recovery ($\Delta RIX_t^{(S_t, +)}$) proxies with $\tau = 1$ month considering the underlying ETFs representing four major geographic equity markets: the US, Japan, the Eurozone, and the UK.

One important takeaway from Fig. 4 is that changes to the right-tail measure display higher levels of volatility than changes to the left-tail measure. This reflects the widely documented fact in the options literature that right-tail measures are naturally noisier than left-tail proxies. This happens because ETF calls are typically less liquid than ETF puts. Also, as expected, the monthly series of $\Delta RIX_t^{(S_t, -)}$ and $\Delta RIX_t^{(S_t, +)}$ indicate that the standard deviation of tail risk increases during periods of economic and financial turmoil.

\(^6\) The advanced economies’ ETF (MSCI EAFE Index, ticker EFA) includes equities in Australia, Austria, Belgium, Denmark, Finland, France, Germany, Hong Kong, Ireland, Israel, Italy, Japan, the Netherlands, New Zealand, Norway, Portugal, Singapore, Spain, Sweden, Switzerland, and the UK. The emerging market economies’ ETF (MSCI Emerging Markets, ticker EEM) is composed of equities from Argentina, Brazil, Chile, China, Colombia, the Czech Republic, Egypt, Greece, Hungary, India, Indonesia, Korea, Malaysia, Mexico, Pakistan, Peru, Philippines, Poland, Qatar, Russia, Saudi Arabia, South Africa, Taiwan, Thailand, Turkey, and United Arab Emirates. Because the calculation of the rare-disaster and swift-recovery indices require the existence of underlying ETFs with holdings defined based on a geography (or sector), not every country for which we have indices computed belongs to the AE and EME country groups (e.g., the Advanced Economies ETF does not hold Canadian assets) and vice versa (e.g., the EMEs ETF holds assets in Pakistan but we do not have the Pakistan-alone version of the indices due to the lack of an ETF with liquid options written on).

\(^7\) In Appendix B, we present the analogue “iceberg” Figures for the period of the subprime crisis.
We then show correlations of changes in the disaster-risk indices both across time and across countries. In the main-diagonal panels of Fig. 5, we present the auto-correlation functions of changes in rare-disaster indices ($D_{\text{RIX}}(S,t;s \equiv 1)$ up to 20 lags (i.e., a typical month with 20 trading days). Similarly, in the off-diagonal panels of Fig. 5, we present cross-correlation functions for series across countries.

The auto-correlation estimates in Fig. 5 suggest that series of changes in rare-disaster indices are not persistent for the sample period in question (between April 14, 2008 and June 30, 2021, the period with valid observations for all four series). Interestingly, the cross-correlation plots suggest that while some market pairs display positive and significant contemporaneous correlations (e.g. US–Eurozone, Japan–Eurozone, Japan–UK, UK–Eurozone), other pairs are surprisingly weakly correlated (e.g., US–Japan) despite their economic ties.

2.4. Baseline specifications

2.4.1. Domestic effects

Our empirical strategy follows Hattori et al.’s (2016) event study regressions. Using our $\text{RIX}$ time series for different underlying ETFs ($S \in \{1, 2, \ldots, N_s\}$) and for different time-to-maturity parameters ($\tau \in \{1, 2, 3, 6\text{ months}\}$), we evaluate the domestic effect of the policy interventions by estimating abnormal changes in these measures with the following specification:

$$
\Delta D_{\text{RIX}}(S,t;\tau) = \beta_1^{(d)} \cdot I_t + \sum_{\gamma \in \Gamma} \beta_\gamma^{(d)} \cdot \text{surprise}_t + \sum_{\gamma < \Gamma} \alpha_{\gamma,m,\tau} \cdot I_t^{(\text{month}_t = m)} \cdot I_t^{(\text{year}_t = y)} + \epsilon_t.
$$

(5)

Our dependent variable is the daily log-return of RIX ($\Delta D_{\text{RIX}}(S,t;\tau) = \log(D_{\text{RIX}}(S,t;\tau)/D_{\text{RIX}}(S,t-1,\tau))$) and our main coefficients of interest are the estimates of the indicator variable $I_t^{(d, Z)}$ which is equal to one for a set of relevant Fed announcement dates $d \in D$ pertaining to the crisis event. As in Hattori et al. (2016), we include $\text{surprise}_t$, a set of controls to account for a wide set of macroeconomic surprises that can plausibly affect disaster risk perceptions via other mechanisms unrelated to...
unconventional monetary policy announcements. Specifically, we use composite series of macroeconomic surprises compiled for five important markets: $\gamma \in \Gamma = \{\text{US, Eurozone, China, Japan, UK}\}$. To keep consistency across our estimations for multiple assets and to account for interconnections across international markets in our subsequent spillover analyses, we include the four series of macroeconomic surprises in all specifications. We also include month-year fixed effects in all specifications to account for lower-frequency omitted factors.

Finally, for the COVID–19 period, we include the log-levels of cumulative deaths and infection cases associated with the corresponding geography of the underlying ETF to account for the impact of the epidemic on investors’ expectations. 8

Throughout our analyses of domestic effects, we interpret the series of coefficients $b^{(s)}$ estimated for the four different values of $s$ as the average abnormal change in the term structure of rare-disaster risk surrounding the dates deemed relevant for unconventional monetary policy responses in both episodes. This is accomplished estimating the time-series model of Eq. (5) for different underlying ETFs (different $S$).

**2.4.2. Spillover Effects**

The empirical strategy to estimate the spillover effects of the Fed’s policy interventions to international equity markets also follows Hattori et al.’s (2016) event study approach. The key difference is the outcome variable for which we estimate the abnormal changes associated with specific relevant dates. Simply put, while in the domestic analysis we are primarily interested in the direct effect on the rare-disaster and swift-recovery indices of US markets and the different sectors of the US economy (i.e., abnormal changes in $\Delta RIX^{(c)}_{t-s}$), our focus is distinct for the international spillover analysis. In the cross-border analysis, our focus is to assess the effect of such interventions on the co-movement that is naturally observed between the disaster risk of the US equity market and the equivalent disaster risk of the equities of a given country $c$. More formally, we estimate how policy interventions cause abnormal changes in the correlation between $\Delta RIX^{(c,-)}_{t-s}$ and $\Delta RIX^{(US,-)}_{t-s}$, where $c \in C$ indexes different countries whose equity ETFs are represented in our sample.

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8 We include log-levels of both infection cases and fatalities because both metrics can be subject to underreporting issues for a variety of reasons, including availability of testing, autopsy procedures, and cross-country differences in diagnosis and procedures (see, e.g., Krantz and Rao, 2020).
Because the overarching purpose of our study is to understand the effect of unconventional monetary policies on disaster-risk expectations during major crises, in this analysis, we focus on spillovers to the rare-disaster indices of other countries alone. To do so, we estimate the following model:

$$D_{RIX}(c,\theta/C0)_{t,s} = \beta_0(s) \frac{1}{C1} I_d D_g + \beta_1(s) \frac{1}{C1} D_{RIX}(US,\theta/C0)_{t,s} \quad \text{if } d = 2D_g$$

The intuition is that the $\beta_0(s)$ coefficients capture the unconditional co-movement between the disaster risks of the US and country $c$, whereas $\beta_3(s)$ estimates of the interaction $(ARIX_{US,\theta}^{(US,\theta)} \cdot I_{[t,\theta]}^{(US,\theta)})$ represent abnormal changes to this correlation due to the Fed’s policy announcements (i.e., abnormal spillovers estimated for the dates of key policy announcements, herein represented by $I_{[t,\theta]}^{(US,\theta)} = 1$). Thus, our main estimates of interest are the $\beta_3(s)$ coefficients, which represent the term-structure of abnormal spillovers to a given country.

Positive (negative) estimates represent abnormal changes in the disaster risk of country $c$ directionally aligned (contrarian) to the domestic US effect. To understand the intuition, recall that the Fed’s intervention dates are associated with reductions in disaster risk in the US, so a positive abnormal co-movement with another country implies that it also experienced the desirable effect of the intervention. Conversely, a negative abnormal co-movement between the disaster risk of the US and that of another country indicates that the country’s disaster risk is moving abnormally in the opposite direction of the desirable US domestic effect—characterizing an undesirable spillover.

![Fig. 4. Volatility of Rare-Disaster and Swift-Recovery Indices (DRIX). This figure reports the serial evolution of the volatility of changes in rare-disaster (red) and swift-recovery (blue) expectations (standard deviation of daily $ARIX_{t}^{(US,\theta)}$ and $ARIX_{t}^{(US,\theta)}$ with $\tau = 1$ month aggregated at the month level). The plots depicted correspond to the index volatility for (i) US equities (ticker SPX, upper-left plot), (ii) Japanese equities (ticker EWJ, upper-right plot), (iii) Eurozone equities (ticker FEZ, lower-left plot), and (iv) UK equities (ticker EWU, lower-right plot).]
As in the analyses of domestic effects, we also consider the series of macroeconomic surprises $C_\gamma$ as controls for both episodes and include the log-levels of cumulative deaths and infection cases corresponding to the country $c$ in question as controls for the COVID–19 crisis.

Critically, when estimating spillover effects of the Fed's announcements to other jurisdictions during the COVID–19 crisis, it is important to consider that different central banks announced their intentions to conduct unconventional monetary policies around the time of the Fed's intervention. As such, changes in rare-disaster risks of country $c$ may also reflect policy announcements by the local monetary authority and not only the potential spillovers from the Fed's actions $C_\gamma$. We denote by an indicator variable $I_{t}^{l(c)}$ the occurrence of interventions of the local monetary authority, where $l(c)$ is the set of announcement dates by the local central bank of country $c$. The dates associated with local central bank interventions for different geographies are shown in Table 1.

Finally, we also perform a panel estimation of Eq. (6) by pooling the time series for different countries, augmenting the model to include country fixed effects. The panel estimation benefits from cross-sectional variation and helps to mitigate concerns that some of our rare-disaster indices for individual countries are not as informative as the main group ETFs (i.e., ETFs written on a basket of AEs and EMEs securities) due to liquidity differences.

3. Results

3.1. Unconventional monetary policy responses to the subprime crisis

We start our analysis of disaster risk effects around the announcements of the Fed’s response to the subprime crisis by identifying the set of relevant dates ($d \in \Delta$) for which our dummy variable in Eq. (5) should take value one. Following the literature, our starting point is the set of 17 dates of monetary easing described in Hattori et al.’s (2016) Table 1. Importantly, our main focus is not the assessment of the market-wide reduction in disaster risk for all the 17 dates associated with the
Unconventional Monetary Policy Announcements by Local Central Banks. This table presents the dates of unconventional monetary policy announcements \( l \in \mathbb{R}(c) \), by the central banks associated with jurisdiction(s) \( c \) compiled by Hartley et al. (2021). These dates are used to define the indicator variables \( l_{c,t}^{(i)} \) in the analysis of disaster-risk spillovers during the COVID–19 crisis, intended to control for the domestic (local) effects of policy announcements by these central banks. Pooled (panel) estimations are performed with \( l_{c,t}^{(i)} \) as a variable that depends on \( c \) and \( t \). Dates reported following the DD.MM.YYYY format.

| Market (ETF ticker) | Intervention Dates and Central Banks (Hartley et al., 2021) |
|---------------------|-------------------------------------------------------------|
| Advanced Economies (EFA) | 16.03.2020 (BoJ, Sweden, and Israel); 19.03.2020 (BoE, ECB, and Australia); 23.03.2020 (New Zealand and Israel); and 27.04.2020 (BoJ). |
| Emerging Market Economies (EEM) | 13.03.2020 (Croatia); 17.03.2020 (Poland); 18.03.2020, 20.03.2020, and 23.04.2020 (India); 23.03.2020 (Philippines); 24.03.2020 (Colombia); 25.03.2020 (South Africa); 26.03.2020 (South Korea); 31.03.2020 and 17.04.2020 (Turkey); 01.04.2020 (Indonesia); 07.04.2020 and 28.04.2020 (Hungary); and 21.04.2020 (Mexico). |
| Eurozone (FEZ) | 19.03.2020 (ECB). |
| Australia (EWA) | 19.03.2020 (Reserve Bank of Australia). |
| Brazil (EWZ) | None. |
| Canada (EWC) | 27.03.2020 (Bank of Canada). |
| Chile (ECH) | None. |
| China (FEZ) | None. |
| France (EWF) | 19.03.2020 (ECB). |
| Germany (EYG) | 19.03.2020 (ECB). |
| Hong Kong (EWH) | None. |
| Italy (EWI) | 19.03.2020 (ECB). |
| India (INR) | 18.03.2020, 20.03.2020, and 23.04.2020 (Reserve Bank of India). |
| Japan (EWJ) | 16.03.2020 and 27.04.2020 (BoJ). |
| Malaysia (EWM) | None. |
| Mexico (EWX) | 21.04.2020 (Banco de Mexico). |
| Russia (RSE) | None. |
| Spain (EYP) | 19.03.2020 (ECB). |
| Singapore (EWS) | None. |
| South Africa (EZA) | 25.03.2020 (South Africa Reserve Bank). |
| South Korea (EYW) | 26.03.2020 (Bank of Korea). |
| Taiwan (ETW) | None. |
| Turkey (TUR) | 31.03.2020 and 17.04.2020 (Central Bank of the Republic of Turkey). |
| United Kingdom (EWU) | 19.03.2020 (BoE). |

Table 1

subprime crisis, as this has been previously documented in Hattori et al. (2016). Instead, our main purpose is to contrast the interventions of the two episodes with respect to their domestic and international spillover effects taken together.

To do so, we first consider a variant of Eq. (5) with distinct dummies for each of the 17 dates. This procedure allows us to disaggregate the average treatment effect estimated by Hattori et al. (2016) (who use a single dummy for all dates), thereby agnostically identifying which dates are unambiguously meaningful in minimizing domestic disaster risk. The rationale is that the channel of cross-border transmission of disaster risk, if present, should occur through the domestic effects of unconventional monetary policies.

Table 2 reports coefficient estimates for the 17 dates, with Newey–West heteroskedasticity and autocorrelation (HAC) robust standard errors. As in Hattori et al. (2016), the time series regressions are estimated over the period from January 3, 2008 to November 6, 2012. A perusal of the abnormal changes for individual dates implies that, with respect to domestic equities, three dates are unambiguously associated with negative shifts in the whole term structure of disaster risk. These dates, labeled 1, 11, and 13, respectively represent: (i) the announcement by the Fed of purchases of $100 billion of agency debt and $500 billion of agency mortgage-backed securities (MBS) on November 25, 2008; (ii) the Fed’s forward guidance announcement on January 25, 2012, committing to keep a low fed funds rate at least until 2014; and (iii) the subsequent forward guidance announcement on January 25, 2012, committing to keep a low fed funds rate at least until 2014.

3.1.1. US domestic effects

We proceed with the estimation of Eq. (5) considering a single indicator variable for the three aforementioned relevant dates. We report in Fig. 6 coefficient estimates representing abnormal changes to the term structure of RIX for the US equity market (ticker SPX), US mid maturity treasuries (ticker IEF), long maturity treasuries (ticker TLT), and a variety of US sectors. Overall, results for the US equity market suggest that the policy announcement is associated with a statistically significant and economically meaningful reduction in the disaster risk for the aggregate market of US equities. Interestingly, despite the positive correlation between the two extreme-tail risk measures, we show that the policy interventions generally do not affect market perceptions of a swift recovery.

Disaster risk for mid-maturity treasuries follow a term structure reduction that closely mimics the effect for equities. Overall, results for the US equity market suggest that the policy announcement is associated with a statistically significant and economically meaningful reduction in the disaster risk for the aggregate market of US equities. Interestingly, despite the positive correlation between the two extreme-tail risk measures, we show that the policy interventions generally do not affect market perceptions of a swift recovery.

Disaster risk for mid-maturity treasuries follow a term structure reduction that closely mimics the effect for equities. Such variations partly reflect the effectiveness of the monetary policy per se in increasing (decreasing) treasury prices (yields). Nevertheless, the swift-recovery likelihood is not significantly affected for mid-maturity treasuries, possibly because bond prices are upper bounded. With respect to long maturity treasury rates, we document insignificant effects
on both extreme tail measures. The smaller disaster-risk sensitivity of US long maturity treasuries can be plausibly attributed to the “absolute debt size effect” that makes US debt securities deemed a safe store of value during periods of economic turmoil (He et al., 2016).

Finally, one can also note in Fig. 6 that the estimates for the effects on the swift-recovery indices (in blue) have much larger standard errors than the estimates for the rare-disaster indices (in red). This is consistent with the evidence previously presented in Fig. 4: for ETF options, the market for puts (used to compute the rare-disaster measures) is much more informationally efficient than the market for calls (used to compute the swift-recovery measures). Since market-wide ETFs (i.e., ETFs on the US equity market and treasuries) are more informationally efficient than sectoral ETFs, we focus our analysis on the aggregate versions of rare-disaster and swift-recovery indices.
Fig. 6. Extreme Tail Risk Term Structure and Unconventional Monetary Policy Surrounding the Subprime Crisis: US Domestic Effects. Each plot represents our estimates of abnormal changes in the term structure of extreme-left (red) and extreme-right (blue) tail risk (coefficients $\beta_{s,t}^{(l)}$) obtained from Eq. (5) for different underlying ETFs and different horizons ($\tau$). 95% confidence intervals are estimated with Newey–West HAC robust standard errors.
An analysis of sector ETFs highlights the channels through which aggregate equity disaster risk is reduced following the announcement by the Fed. Surprisingly, no sectoral ETFs display statistically significant reductions or increases in either tail risk measures. Nonetheless, we observe some non-significant, mild reductions in rare-disaster and swift-recovery expectations for sectors largely dependent on household income (consumer discretionary and, to a lesser extent, retail). Similarly, technology and biotech are associated with mild, non-significant increases in disaster risk for some maturities of their term structures, possibly due to the fact that the sectors are comprised of a large number of growth stocks. Whereas the analysis of sectoral components of rare-disaster and swift-recovery indices is worth documenting, it is important to note that options on sector ETFs are less informationally efficient than options on aggregate underlying assets. Consequently, the rare-disaster and swift-recovery results observed for individual US sectors should be interpreted with caution.

3.1.2. Spillover effects on international equities

We now turn our attention to our main inquiry of evaluating potential spillover effects following the Fed’s policy announcements. We investigate the dynamics of disaster risk co-movements between the US equity market and a variety of ETFs with underlying assets representing different markets and geographies. We proceed with time-series estimations of Eq. (6) for different countries following a top-down approach. That is, we first contrast the spillovers to disaster risk of groups of advanced and emerging market economies (our main focus), and then narrow down to ETFs whose assets are from a specific country. The ETF used to construct our main proxy for the disaster risk of advanced economies is comprised by a basket of equities from 21 countries and the emerging market economies ETF includes equities from 26 countries, as mentioned above. For liquidity and diversification reasons, such instruments are arguably the best available proxies to contrast spillover effects on the two groups of countries.

As done in the domestic analysis, to circumvent the fact that some of the individual country ETFs yield noisier proxies of rare-disaster expectations, we pool the ETFs of individual countries and estimate a panel version of Eq. (6) augmented with country fixed effects. The pooled estimation allows us to increase the power of our test of international spillovers, netting out the cross-country heterogeneity in the precision of the rare-disaster series.

Fig. 7 shows the results of the term structure of spillover effects with Newey–West HAC 95% confidence intervals and double-clustered standard errors at the country and day levels for the panel estimates. Because the policy announcements in question are associated with reductions in disaster risk of the US domestic equity market, a positive (negative) coefficient should be construed as a desirable (undesirable) spillover effect.

Overall, we show that the spillovers of a unilateral announcement of unconventional monetary policy by the Fed are undesirable even for advanced economies—an effect particularly strong for the shortest (τ = 1 month) and longest (τ = 6 months) horizons of the term structure. Emerging market economies also experience some negative spillover effects, particularly pronounced in the shorter horizon of their term structure, while benefiting from some desirable spillover effects in the longer horizon. A possible explanation for the heterogeneous effects on different duration values observed for EMEs (but not for AEs) is that the dynamics of the term structure for EMEs is partly driven by investors re-balancing their portfolios across different durations, but within the same group of EMEs. Specifically, announcements of a major injection of liquidity conducted unilaterally by the Fed may induce negative spillovers to short-term expectations as economic agents internalize destabilizing effects on exchange rates and other asset pricing distortions. However, as some institutional investors have specific mandates to hold assets from EMEs, a “reaching-for-yield” effect may arise, with capital flowing from long- to short-term maturity instruments. The immediate consequence would be concomitant negative spillovers for the short horizon and positive spillovers for the long horizon.

Delving into specific countries, we emphasize significant undesirable spillover effects for the majority of markets analyzed. Notably, negative spillovers are observed for the shorter and longer horizons of the term structure—a pattern clearly present for Germany, Hong Kong, Mexico, South Korea, and Taiwan. Finally, the pooled panel estimate (lower panel of Fig. 7) provides supporting evidence of such undesirable effects.

Collectively, these results are congruent with the notion that a unilateral QE by the central bank of a large developed economy may lead to destabilizing effects, corroborating concerns voiced by central bank governors of emerging economies that unilateral expansions of money supply may lead to a range of unintended consequences for international trade, capital flows, and financial stability. Moreover, our results suggest that even advanced economies are not immune to such effects.

3.2. Unconventional monetary policy responses to the COVID–19 crisis

In response to the unparalleled disruption caused by the COVID–19 pandemic on public health and the real economy, central banks around the world announced a range of unconventional policy measures injecting liquidity into households and firms. Along these lines, starting in mid-March 2020 the Fed initiated a sequence of policy interventions, including swap lines with core and peripheral central banks, lending facilities, and others.

9 Similarly to US sectoral ETF measures of rare disasters and swift recoveries for individual countries, we emphasize that some individual country ETFs were neither introduced nor actively traded until after the Global Financial Crisis. Hence, due to the lack of liquid ETFs for many markets, the spillover analysis of the subprime crisis is narrower than the COVID–19 analysis.
While the response to the subprime crisis was characterized by multiple Fed announcements related to three QE rounds following the collapse of Lehman Brothers, the instances of QE conducted in response to the ongoing pandemic took place within a much shorter time span. Hence, in our analysis of the COVID–19 crisis, we follow the recent literature (e.g., Haddad et al., 2021) and consider the two relevant dates \( (d \in \mathbb{D}) \) to be March 23, 2020 and April 9, 2020.  

With the establishment of the primary and secondary market corporate credit facilities (PMCCF and SMCCF) as the two clearest instances of UMPs implemented to counteract the effects of the COVID–19 pandemic, little discretion remains regarding the key dates considered to estimate domestic and spillover effects. The results in Haddad et al. (2021) support the notion that the first announcement of the PMCCF and SMCCF (on March 23, 2020) and their subsequent expansion (on April 9, 2020) should be our key dates of interest for the COVID–19 interventions.
On March 23, the Fed established the Primary Market Corporate Credit Facility (PMCCF) and the Secondary Market Corporate Credit Facility (SMCCF). Such facilities represented a significant leap beyond the policies adopted during the subprime crisis, as they effectively allowed the Fed to take on credit risk through long positions on a variety of asset classes (e.g., corporate debt). These facilities were subsequently expanded on April 9. To consider the tail risk dynamics over both normal and pandemic times, we estimate Eq. (5) considering the sample period from January 3, 2017 until June 30, 2021.11

3.2.1. US domestic effects

As in the subprime analysis, we start by investigating the effects of the Fed’s policy announcement on the aggregate and sectoral disaster risks for the US market. The time-series estimates of Eq. (5) represent the abnormal changes to the term structures of domestic rare-disaster and swift-recovery indices. Fig. 8 reports the results.

Generally speaking, the aggregate equity disaster risk in the top left panel of Fig. 8 experiences a significant reduction in response to the unveiling of the PMCCF and SMCCF—two unconventional monetary policy instruments directly designed to inject liquidity into the US corporate sector, particularly in shorter maturities. The swift-recovery index response mimics the rare-disaster index response, in line with the previously mentioned correlation of the two proxies. In other words, abnormal changes in disaster risk for announcements in response to the COVID–19 crisis are directionally consistent with the responses observed following the subprime announcements (albeit larger in magnitude).

The contrast between the two crises becomes more evident when we turn to disaster risk of treasuries in the second and third panels of Fig. 8. Specifically, we observe mixed results (short-horizon decreases coupled with long-horizon increases) for the disaster risk of mid-maturity treasuries and a significant increase in disaster risk expectation for long-maturity treasuries. Albeit far from conclusive, a possible explanation for the contrasting results with the disaster risk reduction following the subprime policies is the larger scale of the monetary expansion for the ongoing crisis. The dynamics of “treasury inconvenience yields” following the large accumulation of treasuries by market dealers during the COVID–19 crisis documented by He et al. (2021) may also explain the different patterns of treasury disaster risk. In particular during the COVID–19 crisis, the “flight-to-safety” effect is strikingly different from typical episodes of financial turmoil (e.g., how investors rebalance their portfolios between short-term and long-term bonds).

The sectoral analysis of US equities shown in Fig. 8 suggests important similarities with the disaster risk responses following the previous crisis. For instance, we document decreases in disaster risk for banks following the announcement of the corporate credit facilities. This is consistent with the inherent nature of the policies as direct bailouts to the US corporate sector that spill over to lenders. Cash infusions to non-financial firms help potentially delinquent borrowers meet their financial obligations, curtailing the credit risk of the financial sector. We also document important reductions in disaster risk for some mid-term maturities of high-yield corporate bonds, consistent with the wider scope of assets targeted by the Fed with the 2020 interventions. Non-financial sectors, in particular transportation, home construction, real estate, and technology, also experience noticeable reductions in disaster risk.12

3.2.2. Spillover effects on international equities

Turning to potential spillover effects still arising from the first announcement of the PMCCF and SMCCF, we investigate the abnormal co-movements of the disaster risk in different equity markets and the disaster risk of US equities, with Eq. (6) as our baseline model. As in our subprime analysis, we first perform time-series estimates with underlying assets representing different geographies and then pool the individual country ETFs in a panel estimation with country fixed effects. The results are presented in Fig. 9, with the lower panel representing the panel estimates.

Differently from the spillover effects arising from policy announcements related to the Fed’s unilateral response to the subprime crisis, we now report desirable (positive) spillovers for the vast majority of ETFs considered in our analysis. This is particularly true for the time-series estimates with ETFs based on baskets of countries, including advanced economies (AEs), emerging market economies (EMEs), and the Eurozone, which have better signal-to-noise ratios (i.e., are more informative) than do the proxies for individual countries.

An examination of individual countries also suggests generally positive spillover effects, including instances where the whole yield curve is unambiguously co-moving aligned with the US domestic effect. This phenomenon is observed for many important markets, advanced economies (Canada, Italy, Japan, Spain, and the UK) and emerging markets alike (Mexico, South Africa, and South Korea). However, spillover estimates obtained for other important countries are negative (e.g., Chile, Germany, and France). Pooling countries in a panel estimation to alleviate measurement error concerns also yields abnormal co-movements that are unambiguously positive along the term structure of disaster risk (lower panel), in tune with the effects obtained for the country groups.

11 For dates prior to the COVID–19 outbreak, pandemic controls (i.e., log(1 + Deaths) and log(1 + Infections)) are coded as 1 (i.e., representing the absence of COVID–19 outbreaks).

12 We reiterate that the COVID–19 analysis covers a larger set of US sectors because sector ETFs became more popular after the Global Financial Crisis. The greater availability and popularity of sector ETFs also makes the corresponding sector proxies of rare-disaster and swift-recovery expectations less noisy (more informative) vis-à-vis the subprime period.
3.2.3. Spillover effects and the role of monetary independence, exchange rate stability, and financial openness

The credibility of a shared protagonism in policy interventions is a crucial channel underlying the positive spillovers observed for the COVID–19 episode. Therefore, a natural question is whether differences in the conditions faced by monetary

![Diagrams showing extreme tail risk term structure and unconventional monetary policy surrounding the COVID–19 Crisis: US Domestic Effects.](Fig. 8)

Each plot represents our estimates of abnormal changes in the term structure of extreme-left (red) and extreme-right (blue) tail risk (coefficients $\beta^{(s)}$) obtained from Eq. (5) for different underlying ETFs and different horizons ($\tau$). 95% confidence intervals are estimated with Newey–West HAC robust standard errors.

Fig. 8. Extreme Tail Risk Term Structure and Unconventional Monetary Policy Surrounding the COVID–19 Crisis: US Domestic Effects. Each plot represents our estimates of abnormal changes in the term structure of extreme-left (red) and extreme-right (blue) tail risk (coefficients $\beta^{(s)}$) obtained from Eq. (5) for different underlying ETFs and different horizons ($\tau$). 95% confidence intervals are estimated with Newey–West HAC robust standard errors.
and fiscal authorities across countries modulated these positive spillovers (e.g., Mishkin (2008), Cortes and Paiva (2017)). To investigate this, we augment the panel estimation of Eq. (6) to include interactions with the three dimensions of Aizenman et al.’s (2010, 2013) “impossible trinity” (also known as “trilemma”). This leads to the following specification:

Fig. 9. Disaster Risk Term Structure and Unconventional Monetary Policy Surrounding the COVID–19 Crisis: Spillover Effects. This Figure shows the abnormal co-movements (coefficients $b_{ij}$ of Eq. (6)) estimated for the relevant dates of policy announcements in response to the COVID–19 crisis. 95% confidence intervals are computed with HAC robust Newey–West standard errors for the time-series estimates and with standard errors double-clustered at the geography and day levels for panel estimates (pooled sample of countries in the lower panel).
**Fig. 10.** Disaster Risk Term Structure and Unconventional Monetary Policy Surrounding the COVID–19 Crisis: Interaction of Spillover Effects with the “Trilemma” Components. This figure shows the triple interaction coefficients of abnormal co-movements estimated for the Fed’s announcement dates to the COVID–19 crisis with (i) monetary independence (Panel A); (ii) exchange rate stability (Panel B); and (iii) financial openness (Panel C). Panel D represents the estimates of the interaction effects with monetary independence excluding Eurozone countries. 95% confidence intervals with standard errors double-clustered at the geography and day levels.

\[
\begin{align*}
\Delta RIX_{c,t} & = \beta_1^{(T)} \cdot I_{t}^{(d;T)} + \beta_2^{(T)} \cdot \Delta RIX_{c,t}^{(US;T)} + \beta_3^{(T)} \cdot \Delta RIX_{c,t}^{(US;R)} + \beta_4^{(T)} \cdot I_{t}^{(d;C)} + \beta_5^{(T)} \cdot I_{t}^{(FOP)} + \beta_6^{(T)} \cdot \Delta RIX_{c,t}^{(US;P)} + \beta_7^{(T)} \cdot \Delta RIX_{c,t}^{(US;F)} + \beta_8^{(T)} \cdot \Delta RIX_{c,t}^{(US;C)} + \beta_9^{(T)} \cdot I_{t}^{(d;C)} + \beta_10^{(T)} \cdot I_{t}^{(FOP)} + \beta_11^{(T)} \cdot \log(\text{Infections}_{c,t}) + \beta_12^{(T)} \cdot \log(\text{Deaths}_{c,t}) + \beta_13^{(T)} \cdot \text{surprise}_{t} + \sum_{y \in \Omega} \sum_{m \in \mathbb{R}} \alpha(y,m,c,t) \cdot I_{t}^{(\text{monetary}-m)} + \epsilon_{t}^{(T)}
\end{align*}
\]

where $IT_c$ represents each dimension of the “impossible trinity”: monetary independence ($MI_c$), exchange rate stability ($ERS_c$), and financial openness ($FOP_c$). $IT_c$ is static (cross-country) variables for the purpose of our study. For each country represented in our sample of ETFs (cf. Fig. 9), we define $IT_c$ as the the “trilemma” component in question computed for the most recent year considering the period 2017–19.

Our coefficient of interest is $\beta_1^{(T)}$, a triple interaction representing the extent to which the spillovers documented in Fig. 9 are amplified by monetary independence, exchange rate stability, and financial openness. The panel regression results for each of the components are displayed in Fig. 10.

Overall, estimates of $\beta_1^{(T)}$ suggest that the magnitude of the (generally positive) spillovers from the Fed’s responses are more pronounced in countries with greater exchange rate stability (Panel B) and higher levels of financial openness (Panel C), but largely unrelated to monetary independence (Panel A) for the countries in our sample. It is important to note, however, that by design Eurozone countries have no monetary independence, as they belong to a monetary union. For the purposes of our study, the lack of sovereignty of Eurozone countries to define their own monetary policies does not mean that central banks in their jurisdiction (the ECB) cannot credibly conduct rounds of QE, as deemed necessary. Thus, in Panel D, we also present coefficient estimates of the pooled regression with $MI_c$ as our interaction variable, but excluding Eurozone countries from the sample. The results are in tune with greater monetary independence being associated with stronger (positive) spillover effects.

### 3.2.4. Spillover effects and the role of fiscal space

In response to the challenging nature of the COVID–19 crisis, governments put forth not only a variety of monetary policies but also fiscal policies. Rather than focusing on the expansionary fiscal policies employed by different governments (see, e.g., Benmelech and Tzur-Ilan, 2020 and Jinjarak et al., 2021), we investigate how the aforementioned spillovers of Fig. 9...

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13 We refer the reader to Aizenman et al. (2010) for more details on the justification for each trilemma component. For convenience, we briefly define here each variable. First, monetary independence ($MI_c$) is the reciprocal of the annual correlation of monthly money market interest rate in home country $j$ and base country $i$ (given by $I_i$ and $I_j$, respectively). The index is normalized so it lies between 1 (most) and 0 (least independent). Specifically: $MI_c = 1 - \frac{\text{corr}(I_j, I_i) \cdot (-1)}{1 - (-1)}$. Second, exchange rate stability ($ERS_c$) is defined as the annual standard deviations of the monthly log-change in the exchange rate between the home country and the base country. The authors then normalize the index between zero and one in the following calculation: $ERS = 0.01/0.01 + \text{stddev}(\Delta \log(\text{Exchange Rate}))$. To avoid a downward bias, they also apply a threshold commonly adopted in the literature: if the rate of monthly change in the exchange rate stayed within ±0.33% bands, they define the exchange rate as “fixed” and assign a value of one for the ERS index. Higher values of the index indicate greater exchange rate stability against the currency of the base country. Finally, financial openness ($FOP_c$) is an index of capital account openness based on information regarding restrictions in the IMF’s Annual Report on Exchange Arrangements and Exchange Restrictions. It is constructed as the first standardized principal component of the variables indicating the presence of multiple exchange rates, restrictions on current account transactions, restrictions on capital account transactions, and the requirement of the surrender of export proceeds.

14 At the time of writing, Professor Hiro Ito’s website provides data for the following periods: $MI_c$ (1960–2019); $ERS_c$ (1961–2019); and $FOP_c$ (1970–2018). Whenever available, we consider the value of 2019 for each country. If data for 2019 is missing for country $c$, we consider the data point with the most recent year (i.e., 2018 or 2017).
depend on the fiscal space of the corresponding countries, as fiscal conditions affect government multipliers, credit conditions, and the effectiveness of monetary policy (e.g., Sargent, 1999 and Silva, 2021).

To do so, we use Kose et al.’s (2017) fiscal space indicators. We follow an approach similar to the trilemma analysis, replacing $\Pi_T$ in Eq. (10) with the six indicators of government debt sustainability: (i) general government gross debt, % of GDP ($ggdyc$); (ii) primary balance, % of GDP ($pbyc$); (iii) cyclically-adjusted balance, % of potential GDP ($cbyc$); (iv) fiscal balance, % of GDP ($fbyc$); (v) general government gross debt, % of average tax revenues ($dfggdc$); and (vi) fiscal balance, % of average tax revenues ($dffbc$). Fig. 11 depicts the results of the triple interaction coefficient estimates.

Overall, the interaction coefficient estimates imply that spillovers observed at dates of Fed policy announcements are stronger when each of the six indicators is increased. These results may seem counter-intuitive at first, as higher values of these fiscal variables imply a less sustainable government debt ($ceteris paribus$). However, it is important to note the general equilibrium aspects of sovereign debt. Specifically, higher levels of government indebtedness and fiscal deficits can be construed as the outcome of an ex-ante greater debt capacity, in particular for advanced (and highly credible) economies. As such, the results of Figs. 10 and 11 are in line with positive spillovers being more prominent when the credibility of fiscal and monetary policies is high.

4. Concluding remarks

We investigate the domestic and spillover effects of unconventional monetary policies put forth by the Fed on options-based measures of disaster risk, contrasting interventions during the subprime and COVID–19 crises. The contrasting aspect we focus on is whether the monetary expansion resulted from a single central bank or multiple monetary authorities.

Broadly speaking, the Fed responded in a self-centered fashion to the subprime crisis with its rounds of QEs. In contrast, the COVID–19 crisis elicited an unprecedented multilateral response by different central banks around a short period. Differently put, the prior experience of successfully conducted rounds of QE in recent history lent credence to the possibility of a shared protagonism of central banks in response to the pandemic.

We emphasize some findings on the efficiency of such policies to minimize disaster risk. First, the domestic effects of the policy responses of both episodes are considerably similar for US equities and, to a lesser extent, for mid-maturity government bonds. Increases in disaster risk for long-term US treasuries are particularly robust following the announcements of the COVID–19 interventions. Second, we document that the main differences between the two events lie in the direction of the spillovers to international equity markets. While the announcements of subprime QEs are associated with negative spillovers, the COVID–19 crisis interventions are (mainly) characterized by positive spillovers. Finally, we show that positive spillovers are more sizeable for countries with higher levels of monetary independence, exchange rate stability and financial openness.

Fig. 11. Disaster Risk Term Structure and Unconventional Monetary Policy Surrounding the COVID–19 Crisis: Interaction of Spillover Effects with Fiscal Space Indicators. This Figure shows the interaction effects of the abnormal co-movements estimated for the relevant dates of policy announcements in response to the COVID–19 crisis with different indicators of government debt sustainability, including (i) general government gross debt, % of GDP ($ggdyc$); (ii) primary balance, % of GDP ($pbyc$); (iii) cyclically-adjusted balance, % of potential GDP ($cbyc$); (iv) fiscal balance, % of GDP ($fbyc$); (v) general government gross debt, % of average tax revenues ($dfggdc$); and (vi) fiscal balance, % of average tax revenues ($dffbc$). 95% confidence intervals are computed with standard errors double-clustered at the geography and day levels.
Our results are subject to important caveats. Naturally, the comparison of two major global crises posits critical challenges to the out-of-sample validity. The crises in question are rare events that differ not only in the number of central banks actively engaging in expansionary policy responses but also with respect to many other relevant aspects. For instance, 2008 was caused by imbalances in the financial sector, whereas 2020 resulted from deteriorating public health conditions. Notwithstanding these challenges, our results support concerns raised by central bankers and policymakers that unilateral action by large advanced economies may have unintended consequences for disaster risk in emerging markets. Along these lines, our evidence suggests that expectations of a shared protagonism in unconventional monetary policy responses to major crises can mitigate undesirable spillover effects.

**Disclaimers**

All opinions and inferences are attributable to the authors and do not represent the views of the T. Rowe Price Associates, Inc.

**Declaration of Competing Interest**

The authors declare that they have no known competing financial interests or personal relationships that could have appeared to influence the work reported in this paper.

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**Appendix A. Numerical estimation of rare disaster and swift recovery series**

As in Gao et al. (2019), we start cleaning the options data by removing options with missing Black and Scholes (1973) implied volatility, missing prices and price values less than 0.05 USD. We also remove options whose prices violate non-arbitrage restrictions and, for liquidity reasons, restrict the set of options maturing between seven and 180 days. This set of option prices and contract information are the inputs to construct the series of disaster-risk and swift-recovery measures.

For a given underlying ETF \( (S_t) \) and a given trading day \( t \) with option prices available, we estimate the tail risk proxies for a fixed “time-to-maturity” parameter \( \tau \) as follows. We identity the two option expiration dates \( (T_1, T_2) \) nearest to the synthetic maturity date of the \( \text{RIX} \) measure \( (T = t + \tau) \). Each of the adjacent option expiration dates yield a finite set of option contracts \( \{O(S_t, K^{(i)}, T_1), \text{for } i \in 1, 2, \ldots, I^{(T_1)}\} \) and \( \{O(S_t, K^{(i)}, T_2), \text{for } i \in 1, 2, \ldots, I^{(T_2)}\} \) with different strike prices (i.e., different moneyness). We disregard trading days with fewer than two option prices for a given maturity (i.e., fewer than two strike prices). In possession of the finite sets \( \{\sigma^{\text{BSM}}(S_t, K^{(i)}, T_1), \text{for } i \in 1, 2, \ldots, I^{(T_1)}\} \) and \( \{\sigma^{\text{BSM}}(S_t, K^{(i)}, T_2), \text{for } i \in 1, 2, \ldots, I^{(T_2)}\} \) of observed Black and Scholes (1973) implied volatilities for different strike/moneyness levels, for each \( T_1 \) and \( T_2 \) we construct a finite grid of 2,000 synthetic option strikes, equally spaced over a range from zero to three times the spot value—specifically, \( \{K^{(j)}\}_{j=1}^{2000} \) over \( [0, 3S_t]\). The observed Black and Scholes (1973) implied volatilities \( \{\sigma^{\text{BSM}}(S_t, K^{(j)}, T_1) \text{ and } \sigma^{\text{BSM}}(S_t, K^{(j)}, T_2)\} \) are then interpolated, generating a set of 2,000 implied volatilities \( \{\sigma^{\text{BSM}}(S_t, K^{(j)}, T_1)\}_{j=1}^{2000} \) for \( T_1 \) and 2,000 implied volatilities \( \{\sigma^{\text{BSM}}(S_t, K^{(j)}, T_2)\}_{j=1}^{2000} \) for \( T_2 \).

The two sets of 2,000 implied volatilities are then interpolated over the “maturity” dimension, i.e., linearly interpolated between \( T_1 \) and \( T_2 \) for each grid point \( j \) to obtain the implied volatility curve \( \{\sigma^{\text{BSM}}(S_t, K^{(j)}, T)\}_{j=1}^{2000} \) for the synthetic expiration \( T \) corresponding to the \( \text{RIX} \) parameter \( \tau \). We then use the BSM pricing equation to obtain the range of prices (calls and/or puts) necessary to numerically estimate the integrals of \( \text{RIX} \).

**Appendix B. Data Appendix**

Figs. B.1 and B.2

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\(^{15}\) Importantly, the selection of adjacent dates \( T_1 \) and \( T_2 \) is needed for the most general case where there are no option contracts for the maturity relative to \( \tau \). The special case where there are liquid option contracts expiring at \( T \) is equivalent to the degenerate case where \( T_1 = T_2 \), in which case the linear interpolation over expiration dates is not necessary.
Fig. B.1. Rare-Disaster and Swift-Recovery Measures around the Subprime Crisis for US Sector ETFs. This graph shows the time series of disaster risk (red areas below zero) and fast recovery expectations (blue areas above zero) for global equities from January 1, 2020 to December 31, 2010. The maturity for which tail risk is calculated in all panels is $\tau = 1$ month. (For interpretation of the references to color in this figure legend, the reader is referred to the web version of this article.)
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Fig. B.2. Rare-Disaster and Swift-Recovery Measures around the Subprime Crisis for International Equity ETFs. This graph shows the time series of disaster risk (red areas below zero) and fast recovery expectations (blue areas above zero) for global equities from January 1, 2008 to December 31, 2010. The maturity for which tail risk is calculated in all panels is $\tau = 1$ month. (For interpretation of the references to color in this figure legend, the reader is referred to the web version of this article.)
