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The global financial cycle and macroeconomic tail risks

Johannes Beutel
(Deutsche Bundesbank)

Lorenz Emter
(European Central Bank and Trinity College Dublin)

Norbert Metiu
(Deutsche Bundesbank)

Esteban Prieto
(Deutsche Bundesbank, CEPR, and IAW)

Yves Schüler
(Deutsche Bundesbank)

Deutsche Bundesbank, Wilhelm-Epstein-Straße 14, 60431 Frankfurt am Main,
Postfach 10 06 02, 60006 Frankfurt am Main

Tel +49 69 9566-0

Please address all orders in writing to: Deutsche Bundesbank,
Press and Public Relations Division, at the above address or via fax +49 69 9566-3077

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Non-technical summary

Research question

Since the global financial crisis, policymakers have become increasingly concerned with the risk of severe economic declines, i.e., macroeconomic tail risk. In this paper, we investigate the impact of an unexpected tightening of U.S. financial conditions and monetary policy on macroeconomic tail risks around the world. In addition, we explore the country characteristics that can explain the different responses of macroeconomic tail risks to these U.S. shocks.

Contribution

A recent literature emphasizes that domestic macroeconomic tail risk increases when domestic financial conditions tighten. However, at the same time, domestic financial conditions are to a substantial degree determined by global forces, as documented in a literature studying the "global financial cycle" and "global liquidity". Global financial conditions, in turn, are greatly influenced by U.S. financial conditions and monetary policy. Yet, the questions of how abrupt changes in U.S. financial conditions and monetary policy affect macroeconomic tail risks in other countries and which country characteristics increase the vulnerability to such changes have received little attention in the literature.

Results

We find that an unexpected tightening of U.S. financial conditions or monetary policy increases macroeconomic tail risks internationally. Furthermore, we find that macroeconomic tail risks rise significantly more strongly for countries with fixed exchange rates, higher foreign currency exposure, and a higher level of household leverage. This suggests that exchange rate and macroprudential policies can mitigate macroeconomic tail risk.

Nichttechnische Zusammenfassung

Fragestellung

Seit der globalen Finanzkrise beschäftigen sich Entscheidungsträger vermehrt mit dem Risiko eines starken Konjunktureinbruchs, auch makroökonomisches Tail Risk genannt. In der vorliegenden Studie analysieren wir, welchen Einfluss abrupte Verschärfungen der finanziellen Bedingungen und der Geldpolitik in den Vereinigten Staaten auf makroökonomische Tail Risks weltweit haben. Zudem untersuchen wir, welche Eigenschaften von Ländern die unterschiedlichen Ausprägungen der Tails Risks erklären können.

Beitrag

Neuere Studien zeigen, dass das makroökomische Tail Risk in einem Land ansteigt, wenn sich die dortigen finanziellen Bedingungen verschlechtern. Zugleich zeigt die Literatur über den globalen Finanzzyklus oder globale Liquidität, dass inländische finanzielle Bedingungen in großen Teilen durch globale Kräfte bestimmt werden. Die globalen finanziellen Bedingungen werden wiederum stark durch die finanziellen Bedingungen und die Geldpolitik in den Vereinigten Staaten bestimmt. Ungeachtet dessen fanden zwei Fragen bislang nur wenig Beachtung: 1) Beeinflussen abrupte Veränderungen der finanziellen Bedingungen und der Geldpolitik in den Vereinigten Staaten die makroönomischen Tail Risks anderer Länder? 2) Welche Eigenschaften von Ländern können sie vor solchen Ereignissen schützen?

Ergebnisse

Unsere Ergebnisse zeigen, dass eine unerwartete Verschlechterung der finanziellen Bedingungen oder eine Straffung der Geldpolitik in den Vereinigten Staaten makroökomische Tail Risks weltweit erhöhen. Zudem stellen wir fest, dass das makroökonomische Tail Risk weit mehr in Ländern ansteigt, die sich durch feste Wechselkurse, ein höheres Fremdwährungsrisiko und eine höhere Verschuldungsquote privater Haushalte auszeichnen. Dies legt nahe, dass Wechselkurspolitik und makroprudenzielle Politik das makroökonomische Tail Risk abschwächen können.

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Johannes Beutel
Deutsche Bundesbank

Lorenz Emter
European Central Bank
and Trinity College Dublin

Norbert Metiu
Deutsche Bundesbank

Esteban Prieto
Deutsche Bundesbank, CEPR, and IAW

Yves Schüler
Deutsche Bundesbank

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Abstract

We study the link between the global financial cycle and macroeconomic tail risks using quantile vector autoregressions. Contractionary shocks to financial conditions and monetary policy in the United States cause elevated downside risks to growth around the world. By tightening financial conditions globally, these shocks affect the left tail of the conditional output growth distribution more strongly than the center of the distribution. This effect is particularly pronounced for countries with less flexible exchange rate arrangements, higher foreign currency exposures, and higher levels of private sector leverage, suggesting that exchange rate policies and macroprudential policies can mitigate downside risks to growth.

Keywords: Financial shocks; Monetary policy; Global financial cycle; Growth-at-Risk; International spillovers; Quantile VAR

JEL Classification: C32, E23, E32, E44, F44

*Contact addresses: johannes.beutel@bundesbank.de; lorenz.emter@ecb.europa.eu; norbert.metiu@bundesbank.de; esteban.prieto@bundesbank.de; yves.schueler@bundesbank.de. We thank Georgios Georgiadis, Michal Franta, Alan M. Taylor, Felix Thierfelder, and seminar participants at the European Central Bank, the Central Bank of Ireland, the Deutsche Bundesbank, the Trinity College Dublin, the EEA 2020 and the 4th ESCB Research Cluster 3 Workshop for valuable comments and suggestions. The views expressed in this paper are those of the authors and do not necessarily represent the views of the Deutsche Bundesbank, the Central Bank of Ireland, the European Central Bank or the Eurosystem.

1. Introduction

Since the global financial crisis, policymakers have become increasingly concerned with the risk of large negative output growth realizations. As a result, a rapidly growing body of research on downside risks to growth has emerged (e.g., Adrian et al., 2019; Brownlees and Souza, 2021). This literature emphasizes that macroeconomic tail risks increase when domestic financial conditions tighten. At the same time, domestic financial conditions are to a substantial degree determined by global forces, as documented in a literature studying the "global financial cycle" and "global liquidity" (e.g., Eickmeier et al., 2014; Rey, 2015). Global financial conditions, in turn, are greatly influenced by financial conditions and monetary policy in the United States (e.g., Bruno and Shin, 2015; Miranda-Agrippino and Rey, 2020). Yet, the questions of how unexpected changes in U.S. financial conditions and monetary policy affect macroeconomic tail risks in other countries and which country characteristics increase the vulnerability to such changes have received little attention in the literature.

In this paper, we investigate the effects of shocks to U.S. financial conditions and monetary policy on downside risks to growth around the world. To model the international transmission of U.S. financial and monetary policy shocks to tail risks, we estimate Bayesian quantile vector autoregressions on data for 44 advanced and emerging economies. We compute quantile impulse responses for the median, the upper tail and the lower tail of each country's conditional GDP growth distribution, using the excess bond premium (EBP) proposed by Gilchrist and Zakrajsek (2012) as a measure of exogenous changes in U.S. financial conditions and a measure of U.S. monetary policy shocks derived from Miranda-Agrippino and Ricco (2020). We then relate the quantile impulse responses to various country characteristics.

Three main results emerge. First, we find that an exogenous tightening in U.S. financial conditions raises macroeconomic tail risks internationally. After four quarters, the effect of the shock on the lower tail is on average four times stronger than on the median of the conditional distribution of GDP growth, while the impact on the upper tail is positive and far less pronounced. Second, we find that an unexpected tightening in U.S. monetary policy also has stronger effects on the lower tail of the conditional GDP growth distribution than on the median and the upper tail. Hence, following U.S. financial and monetary policy shocks, the conditional distribution of GDP growth changes its shape and becomes more negatively skewed as the lower tail widens. Third, we find that certain country characteristics matter significantly for the international transmission of these shocks on the *lower tail* of the conditional GDP growth distribution. Specifically, following a shock to U.S. financial conditions, downside risks to growth increase significantly more strongly for countries with fixed exchange rates, higher foreign currency exposure, and a higher level of household leverage. Following a U.S. monetary policy shock, macroeconomic tail risks increase more in countries with a relatively less flexible exchange rate regime. These characteristics, if at all, only weakly matter for the *median* of the conditional GDP growth distribution.

Our findings provide a new angle on how shocks that originate in the U.S. affect foreign economies by documenting their impact on macroeconomic *tail* risks. This represents a novel contribution to an extensive literature on the international transmission of U.S. financial shocks and to the literature on international spillovers from U.S. monetary policy, which have focused on the effect of these shocks on the mean.[1] By taking an international perspective, we also bring a new dimension to the literature on downside risks to growth, which has thus far studied the relationship between macro tail risks and *domestic* financial conditions. For instance, Adrian et al. (2019) show that tighter U.S.

[1] On the international transmission of financial shocks, see, e.g., Helbling et al. (2011), Fink and Schüler (2015), Eickmeier and Ng (2015), Abbate et al. (2016), Metiu et al. (2016), Cesa-Bianchi et al. (2018), and Born and Enders (2019). On the international transmission of monetary policy shocks, see, e.g., Kim (2001), Mackowiak (2007), Bruno and Shin (2015), Rey (2015), Fratzscher et al. (2017), Degasperi et al. (2020), and Miranda-Agrippino and Rey (2020).

financial conditions are associated with higher downside risks to U.S. GDP growth.[2] We contribute to a recent strand of this literature that seeks to move from reduced-form models to modeling the structural shocks that drive downside risks, which has analyzed the domestic relationship between shocks and macroeconomic outcomes (e.g., Loria et al., 2019; Chavleishvili and Manganelli, 2019).

Our paper is related to a long-standing literature on international monetary policy dependence in open economies. The classical "trilemma" (or "impossible trinity") of international macroeconomics postulates that countries face a trade-off among the objectives of exchange rate stability, free capital mobility, and independent monetary policy (e.g., Mundell, 1960; Fleming, 1962). Considerable evidence suggests that floating exchange rates enable countries with an open capital account to pursue an independent monetary policy (e.g., Obstfeld et al., 2005; Klein and Shambaugh, 2015; Bekaert and Mehl, 2019). Moreover, evidence also shows that emerging market economies with flexible exchange rates are better able to use their own monetary policies to dampen the influence of foreign financial and monetary developments on domestic financial variables (see, e.g., Obstfeld et al., 2019; Obstfeld, 2021). By contrast, Rey (2015) argues that the scale of financial integration in recent decades has turned the trilemma into a dilemma in which, regardless of the exchange rate regime, national monetary policies are constrained by global co-movement in gross capital flows, banking sector leverage, credit, and asset prices – i.e., the "global financial cycle". These variables are, in turn, significantly influenced by U.S. monetary policy (see, e.g., Bruno and Shin, 2015; Miranda-Agrippino and Rey, 2020). Rey (2015) argues that, while most individual countries cannot control the global financial cycle, national macroprudential policies can be used to increase resilience of the domestic financial system.

Our results provide valuable insights for the trilemma versus dilemma debate and its policy implications. We find that a fixed exchange rate regime does significantly increase

[2]Related papers include: Cecchetti (2008); De Nicolò and Lucchetta (2017); Aikman et al. (2019); Plagborg-Møller et al. (2020); Brownlees and Souza (2021); Adrian et al. (2022).

a country's vulnerability to large downturns in GDP triggered by U.S. financial and monetary policy shocks. This suggests that the classical trilemma still exists and may help to explain episodes such as the 1997 Asian crisis, in which the impossible trinity of fixed exchange rates, unconstrained capital mobility, and an independent monetary policy contributed to large capital inflows followed by a sudden stop, a financial crisis, and a large decline in GDP growth (e.g., Gourinchas and Obstfeld, 2012). Our finding that downside risks to growth are larger for countries with fixed exchange rates and high foreign currency exposure is consistent with the Asian experience. Our result that laxer limits on household debt are associated with higher downside risks to growth mirrors the macroeconomic developments following the 2008 financial crisis, consistent with evidence on the role of household debt for GDP growth (e.g., Mian et al., 2017). These findings suggest that macroprudential policies limiting the extent of foreign currency exposure and the amount of household leverage could mitigate the risk of large economic downturns, consistent with structural models of leverage and macroprudential policies (e.g., Bianchi, 2011; Farhi and Werning, 2016; Korinek and Simsek, 2016).

Our paper proceeds as follows. Section 2 presents the econometric methodology, Section 3 describes the empirical specification used in the analysis, Section 4 presents the results, and Section 5 concludes.

2. Methodology

In this section, we first describe the general VAR framework for modeling the conditional mean of a vector of endogenous variables. We then introduce a quantile VAR that allows us to model the conditional distribution of the endogenous variables in terms of its conditional quantiles, and to derive quantile impulse response functions. Finally, the section closes with a description of the multi-country setup used to study the international transmission of U.S. shocks to macro tail risks.

2.1. The general framework

Consider the following reduced-form VAR of finite order p:

$$y_t = \mu + B_1 y_{t-1} + \ldots + B_p y_{t-p} + u_t, \tag{1}$$

where y_t is a $k \times 1$ vector of time series observed at $t = 1, \ldots, T$, μ is an $k \times 1$ vector of constant terms, B_l are $k \times k$ coefficient matrices for $l = 1, \ldots, p$, and u_t denotes a $k \times 1$ vector of forecast errors where $u_t \sim N(0, \Sigma_u)$ with covariance matrix $\Sigma_u = E(u_t u_t^{'})$. Equation (1) can be written as a first-order VAR using its companion form (see, e.g., Lütkepohl, 2005, Section 2.1.1):

$$\mathbf{Y}_t = \boldsymbol{\mu} + \mathbf{B}\mathbf{Y}_{t-1} + \mathbf{U}_t. \tag{2}$$

The impulse response function (IRF) of $\mathbf{Y}_{t+h}$ to a vector of innovations $\mathbf{U}_t^{\star}$ at horizon $h = 0, \ldots, H$ can be expressed as follows (see, e.g., Koop et al., 1996; Jorda, 2005):

$$\begin{aligned} IRF(h, \mathbf{Y}_{t-1}, \mathbf{U}_t^{\star}) &= E_{t-1}^{\star}\left[E_t\left[\ldots E_{t+h-1}\left[\mathbf{Y}_{t+h}\right]\right]\right] - E_{t-1}^{0}\left[E_t\left[\ldots E_{t+h-1}\left[\mathbf{Y}_{t+h}\right]\right]\right] \\ &= \mathbf{B}^h \mathbf{U}_t^{\star}, \end{aligned} \tag{3}$$

where $E_{t-1}^{\star}[\cdot]$ denotes expectations conditional on the innovation vector $\mathbf{U}_t^{\star}$ (and the information set up to $t-1$), that is, $E_{t-1}^{\star}[\cdot] = E[\cdot|\mathbf{Y}_{t-1}, \mathbf{U}_t = \mathbf{U}_t^{\star}]$; and $E_{t-1}^{0}[\cdot]$ denotes expectations conditional on no shock in period t, that is, $E_{t-1}^{0}[\cdot] = E[\cdot|\mathbf{Y}_{t-1}, \mathbf{U}_t = \mathbf{0}]$; while $E_t[\cdot] = E[\cdot|\mathbf{Y}_t]$. Hence, the IRF in Equation (3) can be interpreted as the difference between a path conditional on a shock in period t and a path conditional on no shock in period t. Each path is obtained by iterating on the respective conditional expectation of $\mathbf{Y}_t$. In this framework, innovations shift the entire conditional distribution of $\mathbf{Y}_t$. Since higher moments of the innovations are assumed not to vary over time, the IRF of the conditional mean and the IRF of any quantile of the conditional distribution are identical.

In such a world, there is no separate role for studying tail risks.

2.2. Quantile vector autoregression

The results by Adrian et al. (2019) and others suggest that there are dynamics beyond the conditional mean affecting the relationship between financial and macroeconomic variables which are not captured by standard VARs. Therefore, we generalize the setup in a way that allows us to study tail risks.

Specifically, we consider quantiles of the conditional distribution of $\mathbf{Y}_t$ through a quantile VAR (see, e.g., Schüler 2020). In companion form, the model can be written as

$$\mathbf{Y}_t = \boldsymbol{\mu}_\tau + \mathbf{B}_\tau \mathbf{Y}_{t-1} + \mathbf{U}_{t|\tau}. \tag{4}$$

This representation generalizes the VAR in Equation (2) by allowing the parameters $\boldsymbol{\mu}_\tau$ and $\mathbf{B}_\tau$ and the vector of innovations $\mathbf{U}_{t|\tau}$ to vary depending on a $k \times 1$ vector of quantiles τ. The quantile VAR in Equation (4) provides a framework to model the conditional distribution of $\mathbf{Y}_t$ via its conditional quantiles:

$$Q_\tau \left[\mathbf{Y}_t | \mathbf{Y}_{t-1}\right] = \boldsymbol{\mu}_\tau + \mathbf{B}_\tau \mathbf{Y}_{t-1}. \tag{5}$$

We estimate the model in Equation (4) using Bayesian methods.[3] For further details on the estimation procedure, see Appendix A.

[3] We estimate the reduced-form parameters for fixed quantile values τ by drawing from a multivariate Laplace distribution using a Metropolis-within-Gibbs sampler with uninformative priors. The priors are $\beta \sim \mathcal{N}(0, I_{k(kp+1)} \cdot 10)$ and $\Sigma \sim \mathcal{IW}(k, I_k)$, with β being the column vector $\text{vec}((\mu_\tau, B_{1|\tau}, \ldots, B_{p|\tau})')$. The Metropolis-within-Gibbs sampler takes 10,000 draws and discards the first 5,000 draws as burn-in draws. We check trace plots and the quantile conditions on the error terms to assure the convergence of the sampler.

2.2.1. Quantile impulse response function

To derive the quantile impulse response function (QIRF), consider the effects of an innovation vector $\mathbf{U}_t^\star$ on the τ'th conditional quantile of $\mathbf{Y}_t$, given parameters $\boldsymbol{\mu}_\tau$ and $\mathbf{B}_\tau$. The effects on impact are known with certainty and are given by $\mathbf{U}_t^\star$. This means that the effects on all conditional quantiles on impact coincide with the effect on the conditional mean. Thus, the QIRF at horizon $h = 0$ is given by:

$$\begin{aligned} QIRF(h=0, \mathbf{Y}_{t-1}, \mathbf{U}_t^\star, \tau) &= Q_\tau[\mathbf{Y}_t|\mathbf{Y}_{t-1}, \mathbf{U}_{t|\tau} = \mathbf{U}_t^\star] - Q_\tau[\mathbf{Y}_t|\mathbf{Y}_{t-1}, \mathbf{U}_{t|\tau} = \mathbf{0}] \\ &= \mathbf{U}_t^\star \end{aligned} \tag{6}$$

One period after the innovation, the impulse response becomes quantile-specific:

$$\begin{aligned} QIRF(h=1, \mathbf{Y}_{t-1}, \mathbf{U}_t^\star, \tau) &= Q_\tau[\mathbf{Y}_{t+1}|\mathbf{Y}_{t-1}, \mathbf{U}_{t|\tau} = \mathbf{U}_t^\star] - Q_\tau[\mathbf{Y}_{t+1}|\mathbf{Y}_{t-1}, \mathbf{U}_{t|\tau} = \mathbf{0}] \\ &= \mathbf{B}_\tau \mathbf{U}_t^\star. \end{aligned} \tag{7}$$

By further iterating on the quantile VAR, we can express the QIRF at horizon h analogously to the IRF in Equation (3) as:

$$\begin{aligned} QIRF(h, \mathbf{Y}_{t-1}, \mathbf{U}_t^\star, \tau) &= Q_{\tau,t-1}^\star \left[Q_{\tau,t}\left[\dots Q_{\tau,t+h-1}\left[\mathbf{Y}_{t+h}\right]\right]\right] \\ &\quad - Q_{\tau,t-1}^0 \left[Q_{\tau,t}\left[\dots Q_{\tau,t+h-1}\left[\mathbf{Y}_{t+h}\right]\right]\right] \\ &= \mathbf{B}_\tau^h \mathbf{U}_t^\star, \end{aligned} \tag{8}$$

where $Q_{\tau,t-1}^\star[\cdot]$ denotes the τ'th quantile conditional on the innovation vector $\mathbf{U}_t^\star$, that is, $Q_{\tau,t-1}^\star[\cdot] = Q_{\tau,}[\cdot|\mathbf{Y}_{t-1}, \mathbf{U}_{t|\tau} = \mathbf{U}_t^\star]$; and $Q_{\tau,t-1}^0[\cdot]$ denotes the τ'th quantile conditional on no shock in period t, that is, $Q_{\tau,t-1}^0[\cdot] = Q_\tau[\cdot|\mathbf{Y}_{t-1}, \mathbf{U}_{t|\tau} = \mathbf{0}]$; while $Q_{\tau,t}[\cdot] = Q_\tau[\cdot|\mathbf{Y}_t]$. Hence, in the spirit of Koop et al. (1996), the QIRF can be interpreted as the difference between a path conditional on a shock in period t and a path conditional on no shock in period t.[4] Each path is obtained by iterating on the respective conditional quantile of $\mathbf{Y}_t$. The QIRF differs from the effect on the conditional mean if and only if the relevant parameters in the matrix $\mathbf{B}_\tau$ of the QVAR in Equation (4) differ from the parameters in the matrix $\mathbf{B}$ of the VAR in Equation (2). If this is the case, there is a role for studying tail risks.

In contrast to the conventional IRF, the QIRF allows for estimating the effects on different conditional quantiles. Specifically, let $\tau_{t:t+h} = (\tau_t, \ldots, \tau_{t+h})$ $(k \times h + 1)$, where τ_t is a $k \times 1$ vector of quantiles chosen for period t. Then, $\tau_{t:t+h}$ defines a path of quantiles for each variable contained in the vector $\mathbf{Y}_t$.

2.2.2. Orthogonalization of quantile VAR innovations

In the context of conventional VARs, the IRF in Equation (3) can be expressed in terms of mutually uncorrelated structural shocks, given a mapping between u_t and structural shocks ε_t. For instance, such a mapping can be obtained by applying the Cholesky decomposition to the covariance matrix of the reduced-form residuals.

To perform structural analysis based on the quantile VAR model, we need to transform the error terms using a similar approach. Following Schüler (2020), we use a co-exceedance matrix in the spirit of Blomqvist (1950) and Koenker and Bassett (1990) that captures the co-variation of residuals from the quantile VAR around quantiles, just as Σ_u summarizes the common fluctuations of residuals around the conditional mean in conventional VARs. The co-exceedance matrix is given by:

$$\boldsymbol{\Omega}_\tau = (\omega_{ij}) \equiv \frac{\mathrm{E}[\psi_{\tau_i}(u_{it|\tau_i})\psi_{\tau_j}(u_{jt|\tau_j})]}{f_{u_{it|\tau_i}}(0) f_{u_{jt|\tau_j}}(0)}, \tag{9}$$

where $\psi_{\tau_i}(u_{it|\tau_i}) \equiv \tau_i - \mathbb{1}(u_{it|\tau_i} < 0)$, with $\mathbb{1}$ being the indicator function, $i, j \in \{1, \ldots, k\}$, and $u_{t|\tau} = (u_{1t|\tau_1}, \ldots, u_{kt|\tau_k})'$. Furthermore, $f_{u_{it|\tau_i}}(0)$ denotes the probability density func-

[4] Thus, while the level of the 10% (90%) conditional quantile of a variable of interest is always below (above) the 50% conditional quantile, this does not have to be the case for its QIRF, which measures how the quantile path changes in response to the shock.

tion of $u_{it|\tau_i}$ evaluated at 0. The product of $\psi_{\tau_i}(u_{it|\tau_i})$ and $\psi_{\tau_j}(u_{jt|\tau_j})$ measures the strength with which two error terms jointly exceed their respective quantiles contemporaneously.

By applying the Cholesky decomposition to the co-exceedance matrix, we obtain orthogonal shocks $\varepsilon_{t|\tau} = (\varepsilon_{1t|\tau}, \ldots, \varepsilon_{nt|\tau})'$, where each $\varepsilon_{it|\tau}$ depends on the full vector of quantiles τ. Specifically, the Cholesky decomposition $\boldsymbol{\Omega}_\tau = P_\tau P_\tau'$ yields

$$\varepsilon_{t|\tau} = P_\tau^{-1} \widetilde{\psi}_\tau(u_{t|\tau}), \tag{10}$$

where $\widetilde{\psi}_\tau(u_t) = (\psi_{\tau_1}(u_{1t|\tau_1})/f_{u_{1t|\tau_1}}(0), \ldots, \psi_{\tau_k}(u_{kt|\tau_k})/f_{u_{kt|\tau_k}}(0))'$.

2.2.3. Relation to other approaches

Our approach to derive quantile impulse responses is based on Schüler (2020). An alternative approach is proposed by Chavleishvili and Manganelli (2019). They develop a recursive quantile VAR model in which the contemporaneous relation between the endogenous variables is estimated explicitly. We show below that our results would essentially be identical when using the methodology of Chavleishvili and Manganelli (2019).

Another approach to compute quantile impulse responses has been proposed by Loria et al. (2019). They suggest a two-step approach which consists of (i) estimating the quantile of GDP growth conditional on financial conditions using quantile regressions, and (ii) using linear local projections to obtain impulse responses of the conditional quantile. In a multi-country setup, this two-step approach requires measuring local financial conditions for all countries in the sample. The necessary data for doing so is however only available for a limited set of countries for a sufficiently long period. For instance, Adrian et al. (2022) compute quantiles of GDP growth conditional on local financial conditions for 11 advanced economies for a sample starting in the early 1980s. The quantile VAR approach is able to accommodate more countries, in our case 44 advanced and emerging countries, in a more flexible way compared to the two-step local projection approach.

2.3. Multi-country analysis

Our objective is to study cross-country differences in the transmission of U.S. financial and monetary policy shocks to macroeconomic tail risks, and to shed light on the country characteristics that help to explain these differences. We proceed in two stages. In the first stage, we use a multi-country approach similar to, e.g., Cesa-Bianchi et al. (2018), Degasperi et al. (2020), and Metiu (2021). For a specific vector of quantiles τ, we estimate a quantile VAR for each country and compute country-specific QIRFs. To obtain a summary measure across all countries, we compute the point-for-point median across country-specific QIRFs, which can be interpreted as the median group estimator as in Degasperi et al. (2020).[5] Around the median effects, we report one-standard-deviation bands that reflect cross-country heterogeneity.

In the second stage, conditional on a specific vector of quantiles, we investigate the sources of heterogeneity in QIRFs across countries using a regression-based approach as in, e.g., Dedola and Lippi (2005) and Dedola et al. (2017). In line with previous studies, we use a 4-quarter sum of the QIRFs as a summary measure of the strength of the estimated effects for each country. We regress this measure on various country characteristics to explain cross-country variation in the QIRFs for different quantiles of interest.

[5]Following Degasperi et al. (2020), we use the median because it is more robust to the impact of potential outliers. However, we obtain similar results when using the mean to aggregate across countries as in, e.g., Pesaran and Smith (1995), Gambacorta et al. (2014), Cesa-Bianchi et al. (2018), and Metiu (2021).

3. Empirical specification

We analyze quarterly time series data for 44 advanced and emerging economies for the period between 1980:Q1 and 2018:Q4.[6] The baseline quantile VAR for country n includes four variables in the following order: a measure of U.S. financial conditions; the log-difference of real GDP of country n; the log-difference of CPI of country n; and the short-term interest rate of country n (where $n = 1, \ldots, 44$). We measure U.S. financial conditions using the Gilchrist-Zakrajsek EBP, retrieved from Favara et al. (2016) (see also Metiu et al., 2016; Born and Enders, 2019). The EBP captures "variation in the average price of bearing exposure to U.S. corporate credit risk, above and beyond the compensation for expected defaults" (Gilchrist and Zakrajsek, 2012, p. 1700). For each country, we collect data for real GDP and CPI from the OECD and the IMF.[7] Interest rate data come from Eurostat, OECD, IMF, and Emter et al. (2019).

Shock identification: We use this baseline specification to estimate the effects of U.S. financial shocks, which implies the identifying assumption that the EBP may contemporaneously affect macroeconomic outcomes in individual countries, while developments outside of the United States may affect the EBP only with a lag of one quarter (see also Rey, 2015; Cesa-Bianchi et al., 2018). For our recursive quantile VAR, this timing restriction implies that the EBP is ordered before the country-specific variables. As a

[6] The countries in our (unbalanced) sample are: Australia (AUS), Austria (AUT), Belgium (BEL), Canada (CAN), Switzerland (CHE), Chile (CHL), China (CHN), Colombia (COL), Czechia (CZE), Germany (DEU), Denmark (DNK), Spain (ESP), Estonia (EST), Finland (FIN), France (FRA), United Kingdom (GBR), Greece (GRC), Hong Kong (HKG), Hungary (HUN), Indonesia (IDN), Irland (IRL), Island (ISL), Israel (ISR), Italy (ITA), Japan (JPN), South Korea (KOR), Lithuania (LTU), Luxembourg (LUX), Latvia (LVA), Mexico (MEX), Malaysia (MYS), Netherlands (NLD), Norway (NOR), New Zealand (NZL), Philippines (PHL), Poland (POL), Portugal (PRT), Russia (RUS), Singapore (SGP), Slovakia (SVK), Slovenia (SVN), Sweden (SWE), Thailand (THA), and South Africa (ZAF).

[7] If for either of those variables no seasonally adjusted series are available, we apply the U.S. Census Bureau's X-13 seasonal adjustment procedure.

robustness check, we show that our results continue to hold when ordering the EBP last.

We estimate the effects of U.S. monetary policy shocks by adding to each country-specific quantile VAR a measure of U.S. monetary policy shocks derived from a high-frequency instrument that accounts for informational rigidities (Miranda-Agrippino and Ricco, 2020).[8] Since the shock series derived from Miranda-Agrippino and Ricco (2020) is an exogenous shock sequence, we order it before the EBP and the country-specific block, applying the same recursive identification strategy.

Parameterization of the quantile VAR: To characterize the QIRF for the lower tail of the conditional distribution of real GDP growth, say $y_{j,t}$, we consider the 10% quantile. Specifically, we set:

$$\tau_{i=j,t} = \tau_{i=j,t+1} = \ldots = \tau_{i=j,t+H} = 0.1, \tag{11}$$

while all other variables take on median paths:

$$\tau_{i\neq j,t} = \tau_{i\neq j,t+1} = \ldots = \tau_{i\neq j,t+H} = 0.5 \tag{12}$$

for $i = 1, \ldots, k$. To characterize the upper tail of $y_{j,t}$ we set $\tau_{i=j,\cdot} = 0.9$ instead of 0.1 in Equation (11). These specifications of the matrix $\tau_{t:t+h}$ allow us to study the dynamics of the system at the lower (upper) tail of GDP growth, while keeping changes – relative to our median benchmark ($\tau_{i,t+h} = 0.5$ for all i and h) – to a minimum (see,

[8] We estimate a VAR for U.S. data using the baseline model specification of Miranda-Agrippino and Ricco (2020) that includes the EBP. We extract the realized monetary policy shock sequence from the VAR using an instrumental variables approach with a proxy for U.S. monetary policy shocks proposed by Miranda-Agrippino and Ricco (2020) that we obtain from the authors' website (see: http://silviamirandaagrippino.com/).

e.g., Montes-Rojas, 2019; Schüler, 2020).[9]

Our approach can easily accommodate other scenarios. There are two dimensions along which alternative assumptions could be considered. First, one could change assumptions on the behavior of the co-variates of the system, for instance, by assuming upper (or lower) tail responses for these variables as well (i.e. different assumptions on $\tau_{i\neq j,t} = \tau_{i\neq j,t+1} = \ldots = \tau_{i\neq j,t+H}$, in the i-dimension). In contrast to such setups, our approach is conservative in the sense that we would expect even stronger tail responses if we assumed that, for instance, financial conditions were at their upper tail, i.e. especially tight, when GDP growth is at its lower tail (see, e.g. Schüler 2020). Second, one could modify the number of tail realizations of GDP growth after a shock (i.e. different assumptions for $\tau_{i=j,t+h}$, in the h-dimension). Clearly, in order to study tail responses at least some tail realizations have to be considered. We consider two quarters to be the minimum of subsequent tail realizations that still allow us to study interesting propagation dynamics. In a robustness check, we therefore study QIRFS when the conditional quantile of GDP growth is at its tail for two quarters after the shock, and switches to the median thereafter. We find that our results are robust to using this conservative approach (see Section 4).

4. Results

This section describes the results, beginning with the effects of U.S. financial and monetary policy shocks on macroeconomic tail risks in foreign economies. We then analyze whether the effects are systemically related to certain country characteristics.

[9] As shown above, the vector of quantiles also affects the identification of structural shocks (see Equation 10). Considering just the lower tail of conditional GDP growth (and not changing the conditional quantile of the other variables) minimizes the changes with respect to the median benchmark in this regard as well.

4.1. Effects of U.S. financial shocks on foreign output growth

Figure 1 displays the international response of GDP growth for the median (black solid line), the 10% quantile (red solid line), and the 90% quantile (blue solid line) of the conditional GDP growth distribution following an unexpected tightening in U.S. financial conditions, associated with an orthogonal one-standard-deviation increase in the EBP. The solid lines reflect the cross-country median of the QIRFs of GDP growth. The shaded areas represent one-standard-deviation bands of the cross-country distribution of the QIRFs of GDP growth, indicating heterogeneity in the QIRFs across countries.

A tightening in U.S. financial conditions leads to a reduction in GDP growth abroad for the median of the GDP growth distribution, for instance, considering the cross-country median (black solid line).[10] The effect of the shock on the upper tail (90% quantile) is positive and less pronounced than the effect on the median. By contrast, the effect on the lower tail (10% quantile) is substantially stronger than the effect on the median. After four quarters, the effect on the lower tail is roughly four times stronger than on the median. Thus, the conditional distribution of GDP growth changes its shape and becomes more negatively skewed as the lower tail widens in response to the shock. The U.S. financial shock accounts for more than three times as much of the variation in GDP growth for the 10% quantile than for the median or the 90% quantile at horizons associated with business cycle frequencies (see Figure B.2 in Appendix B).

The shaded areas around the cross-country median in Figure 1 convey considerable cross-country heterogeneity. The individual QIRFs for each country (median and lower tail) are shown in Figure 2, along with posterior 68% probability bands. Consistent with Figure 1, the individual QIRFs display substantial heterogeneity across countries. At the

[10] Figure B.1 in Appendix B depicts the impulse responses of the endogenous variables to an EBP shock obtained via Cholesky decomposition from a standard recursive VAR. The shock reduces the conditional means of GDP growth, the inflation rate and the short-term interest rate, in line with findings by, e.g., Rey (2015). In Figure 1, we omit the QIRFs of the remaining variables because these are in line with the effects on the conditional mean.

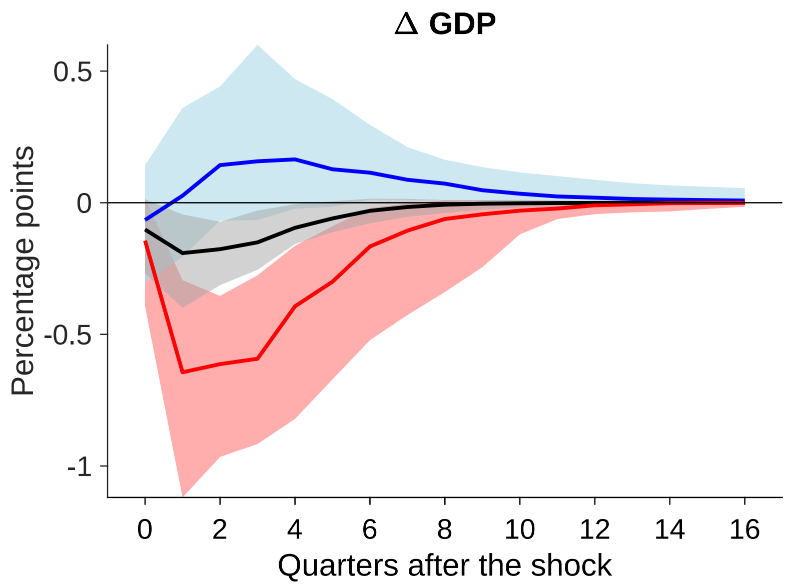

Figure 1: Effects of a U.S. financial shock: Cross-country heterogeneity

Notes: The figure depicts the cross-country median of the quantile impulse response functions for the median (black solid line), the 10% quantile (red solid line), and the 90% quantile (blue solid line) of the conditional GDP growth distribution. Shaded areas correspond to one standard deviation around the cross-country median, indicating heterogeneity in the estimated impulse responses across countries.

same time, the QIRFs for each country can be estimated fairly precisely. For nearly all countries and time horizons, the probability bands around the conditional 50% and 10% quantiles of GDP growth do not overlap and support the conclusion that the effects of a U.S. financial shock are stronger for the conditional 10% quantile.

The asymmetric effects of a shock to U.S. financial conditions on the conditional GDP growth distribution hold in a variety of robustness checks (see Figure 3). First, we obtain virtually identical results when using the quantile VAR model proposed by Chavleishvili and Manganelli (2019), in which the contemporaneous relation between the endogenous variables is estimated explicitly. Second, the results are robust to using a balanced sample that starts in 1996:Q1. Third, using the Chicago Fed's National Financial Conditions Index (NFCI) as a measure for U.S. financial conditions instead of the EBP leads to the same conclusions. Fourth, our results also hold when using PPP-GDP weighted median group estimates. Fifth, the results remain unchanged when amending the quantile VAR with a U.S. block consisting of GDP growth, CPI inflation, and the federal funds rate. The

U.S. variables are ordered above the EBP in order to remove any remaining confounding factors from the U.S. financial shock arising from other U.S. shocks. Sixth, the results are robust to ordering the EBP last. Finally, the results are robust to considering the conditional 10% (90%) quantile of GDP growth only up to two quarters after the shock and the 50% quantile thereafter. This robustness check represents a lower bound on the effects of a U.S. financial shock on macroeconomic tail risks.[11]

[11]For our sample period the average duration of a U.S. recession is around 11 months. Financial recessions typically last longer (e.g., Jorda et al., 2013), especially in emerging economies (e.g., Reinhart and Rogoff, 2014).

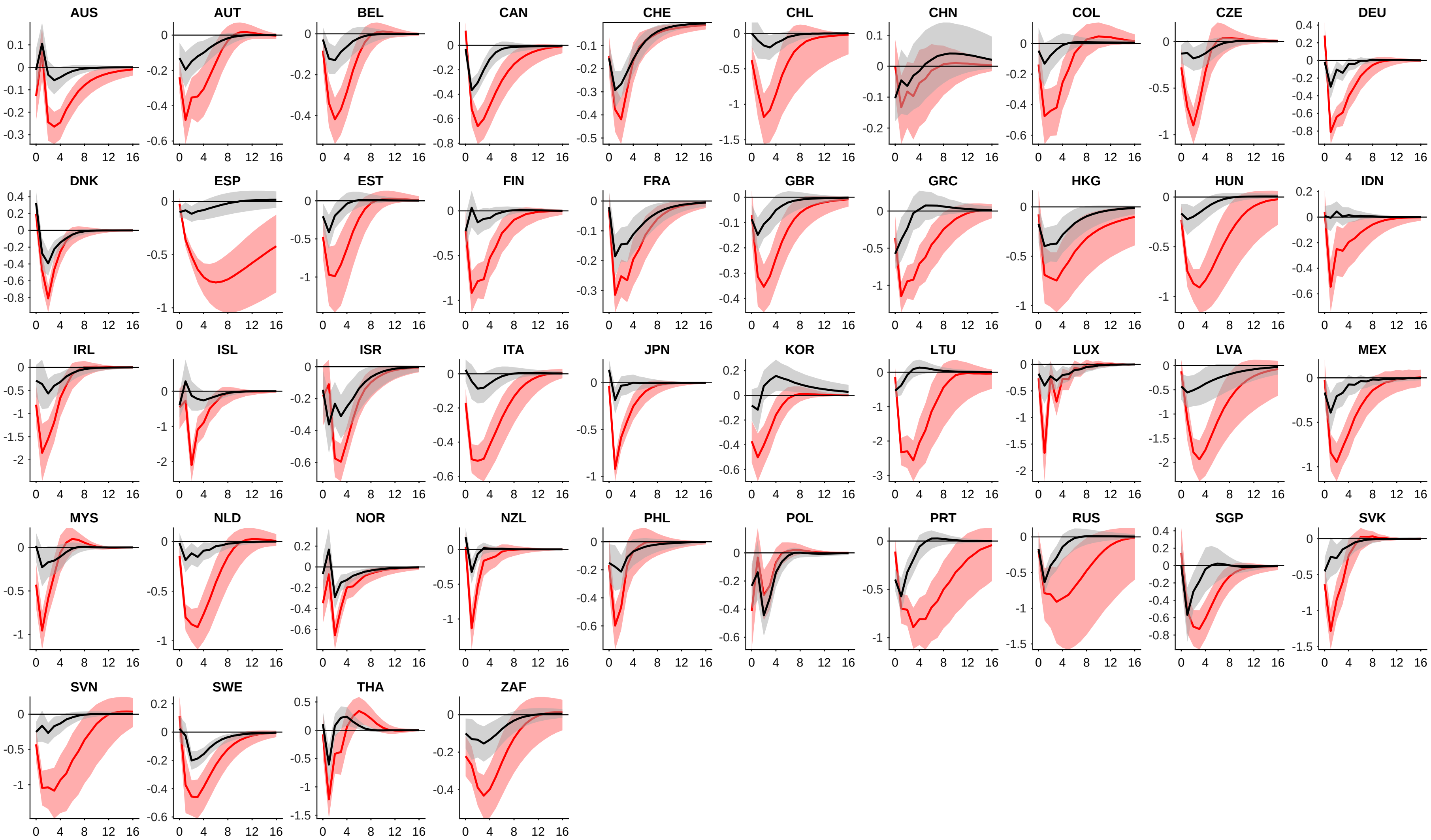

Figure 2: Effects of a U.S. financial shock: Individual country responses

Notes: Median of the posterior estimate of quantile impulse response functions (QIRFs) (solid lines) with posterior 68% probability bands (shaded areas). The black lines represent the QIRFs for the median of the conditional GDP growth distribution, and the red lines are the QIRFs for the 10% quantile of the conditional GDP growth distribution. The x-axis shows quarters after the shock. Shocks are scaled to one standard deviation of the unconditional EBP.

4.2. Effects of U.S. monetary policy shocks on foreign output growth

A number of recent studies have highlighted that U.S. monetary policy is an important driver of the global financial cycle (e.g., Rey, 2015; Bruno and Shin, 2015; Miranda-Agrippino and Rey, 2020). We contribute to this literature by examining the international transmission of U.S. monetary shocks on the tails of the conditional GDP growth distribution.

Figure 4 depicts the international response of GDP growth for the median, the 10% quantile, and the 90% quantile of the GDP growth distribution following an unanticipated tightening in U.S. monetary policy.[12] The impact of the shock on the lower tail (10% quantile) of conditional GDP growth is substantially stronger than the effect on the median or the upper tail (90% quantile). Our estimates thus indicate that an unexpected monetary policy tightening in the United States leads to an increase in downside risks to growth around the world. The size of the impulse response is smaller than for a shock to U.S. financial conditions, in line with monetary policy being only one of several factors that impact financial conditions.

Robustness checks confirm that U.S. monetary policy has stronger effects on the lower tail of the conditional GDP growth distribution (see Figure 5). In particular, estimates yield robust conclusions upon using the recursive quantile VAR model proposed by Chavleishvili and Manganelli (2019), a balanced sample starting in 1996:Q1, the PPP-GDP weighted median group estimator, and considering the conditional 10% (90%) quantile of GDP growth only up to two quarters after the shock and the 50% quantile thereafter.[13]

[12] Individual country IRFs can be found in Figure B.3 in the appendix.

[13] Varying the variable ordering or purging the U.S. monetary policy shock series from U.S. variables would be redundant because it is an exogenous shock sequence extracted from an identified VAR model.

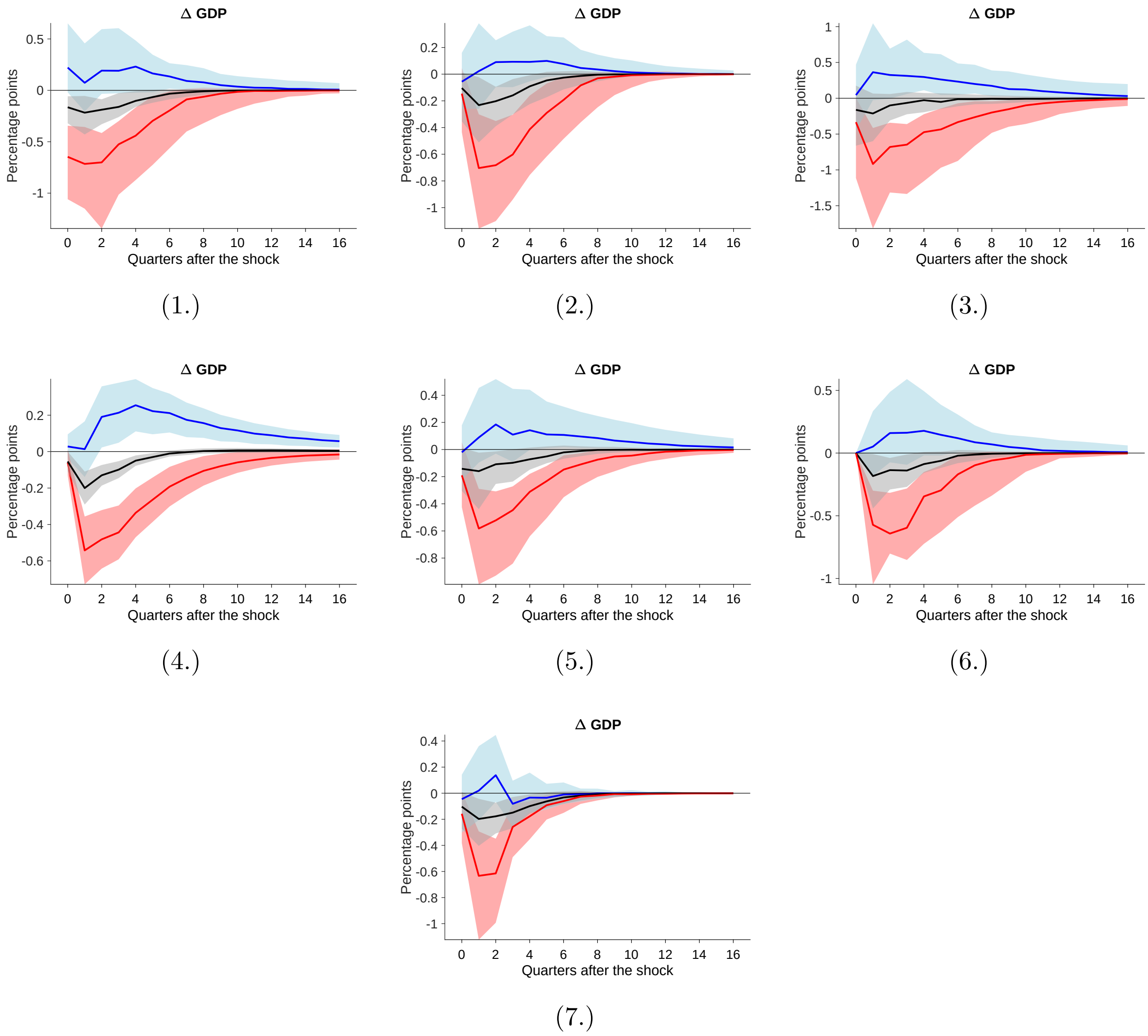

Figure 3: Effects of a U.S. financial shock: Robustness checks

Notes: The figure depicts the cross-country median of the quantile impulse response functions for the median (black solid line), the 10% quantile (red solid line), and the 90% quantile (blue solid line) of the conditional GDP growth distribution. Shaded areas correspond to one standard deviation around the cross-country median, indicating heterogeneity in the estimated QIRFs across countries. Results are shown for: a recursive quantile VAR following Chavleishvili and Manganelli (2019) (Panel 1); a balanced sample starting in 1996:Q1 (Panel 2); NFCI instead of EBP (Panel 3); PPP-GDP weighted median group estimator (Panel 4); a quantile VAR augmented with U.S. GDP growth, U.S. inflation, and the U.S. short-term interest rate (Panel 5); the EBP ordered last (Panel 6); and a robustness check in which, after two quarters following the shock, the quantile value for GDP growth switches from 10% or 90% to 50% (Panel 7).

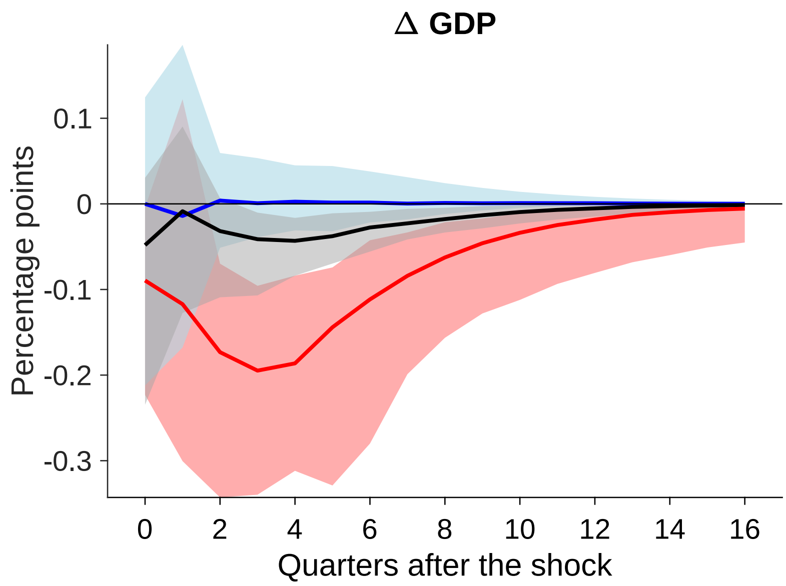

Figure 4: Effects of a U.S. monetary policy shock: Cross-country heterogeneity

Notes: The figure depicts the point-for-point median of the quantile impulse response functions for the median (black solid line), the 10% quantile (red solid line), and the 90% quantile (blue solid line) of the conditional GDP growth distribution. Shaded areas correspond to one standard deviation around the point-for-point median, indicating heterogeneity in the estimated impulse responses across countries.

4.3. Country characteristics and responses to U.S. financial and monetary policy shocks

Are there country characteristics which increase the vulnerability to downside risks from U.S. financial and monetary policy shocks? In order to address this question, we relate the size of individual countries' GDP growth responses to several country characteristics in a cross-sectional regression framework in the spirit of, e.g., Dedola and Lippi (2005) and Dedola et al. (2017).

Following Dedola et al. (2017) and Cesa-Bianchi et al. (2018), we consider the following country characteristics: the degree of *de facto* financial openness following Lane and Milesi-Ferretti (2007); the flexibility of the exchange rate regime, measured according to the classification by Ilzetzki et al. (2019), in which countries with more freely floating currencies are assigned higher values; the level of leverage in the household sector, measured by homeownership-weighted maximum loan-to-value (LTV) ratios following

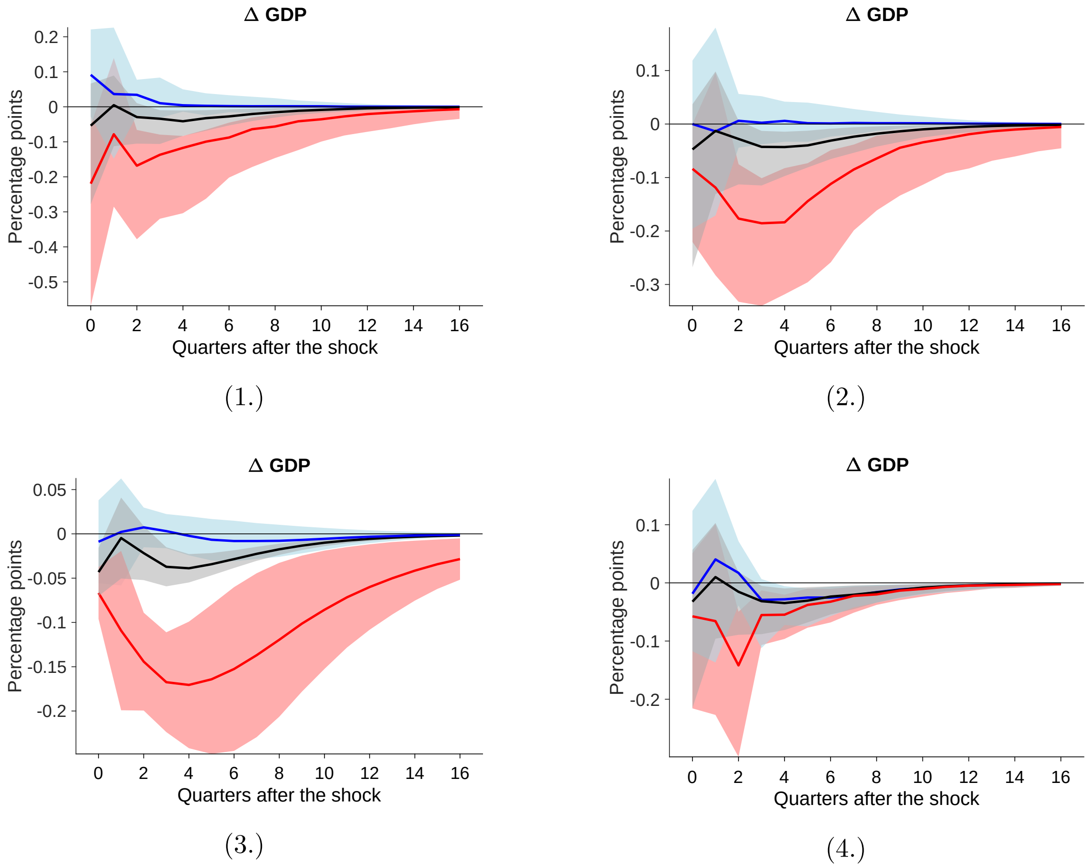

Figure 5: Effects of a U.S. monetary policy shock: Robustness checks

Notes: The figure depicts the cross-country median of the quantile impulse response functions for the median (black solid line), the 10% quantile (red solid line), and the 90% quantile (blue solid line) of the conditional GDP growth distribution. Shaded areas correspond to one standard deviation around the cross-country median, indicating heterogeneity in the estimated impulse responses across countries. Results are shown for: a recursive quantile VAR following Chavleishvili and Manganelli (2019) (Panel 1); a balanced panel starting in 1996:Q1 (Panel 2); the PPP-GDP weighted median group estimator (Panel 3); and a robustness check in which, after two quarters following the shock, the quantile value for GDP growth switches from 10% (90%) to 50% (Panel 4).

Cesa-Bianchi et al. (2018);[14] financial market development, measured by the IMF financial market development index; and the extent of foreign currency exposure, measured as cross-border bank claims in foreign currency over total cross-border bank claims following Cesa-Bianchi et al. (2018).[15] All country characteristics enter the regressions as time averages over the sample period between 1980:Q1 and 2018:Q4. To the extent that not all country characteristics are available over the full sample period, they enter the regressions as time averages over slightly shorter time periods. Our dependent variable is the impact of either the U.S. financial shock or the U.S. monetary policy shock on the 10% quantile of GDP growth cumulated over the first four quarters. For comparison, a second specification shows corresponding estimates on the 50% quantile.

Table 1 shows the OLS estimates for the U.S. financial shock. The first column contains our main specification. We do not find a statistically significant relationship between the effects of U.S. financial shocks and country characteristics at the conditional median of GDP growth (lower panel), similar to the evidence presented by Dedola et al. (2017) and Miranda-Agrippino and Rey (2020).[16] By contrast, for the 10% conditional quantile of GDP growth (upper panel), we find that countries with a relatively more flexible exchange rate regime exhibit a significantly more moderate (i.e., less negative) tail response of GDP growth to a U.S. financial shock. This is consistent with a mechanism by which countries with floating exchange rates face lower macroeconomic tail risks from U.S. financial shocks given that they can, in principle, pursue an independent monetary policy (e.g., Obstfeld et al., 2005; Klein and Shambaugh, 2015; Bekaert and Mehl, 2019)

[14]We obtain maximum LTV ratios from Alam et al. (2019) and home ownership rates from the Housing Finance Information Network.

[15]Numbers on FX exposures follow Cesa-Bianchi et al. (2018) and have been cross-checked with values reported in Bénétrix et al. (2019). Numbers in Cesa-Bianchi et al. (2018) are based on "[...] cross-border bank claims in foreign currency over total cross-border bank claims." Cross-border bank claims refer to "[...] foreign claims (all instruments, in all currencies) of all BIS reporting banks vis-à-vis all sectors." The correlation between this FX measure and the one in Bénétrix et al. (2019) is 0.9. We use the numbers in Cesa-Bianchi et al. (2018) to maximize our country sample.

[16]A similar result is obtained for the 90% quantile (which is therefore not shown).

and therefore dampen the influence of foreign financial and monetary developments on domestic financial variables (Obstfeld, 2021). From the perspective of our results, this mechanism dominates any potentially stabilising effects of a pegged regime that insulates the economy from large swings in the exchange rate.

A higher share of debt denominated in foreign currency is associated with a stronger (i.e., more negative) tail response. Countries with significant amounts of foreign currency debt may face stronger outflows of foreign capital when U.S. financial conditions tighten. In the case of a floating exchange rate, adverse balance sheet effects in response to a currency depreciation may represent an additional channel through which countries with a high share of foreign currency debt may suffer from a U.S. financial shock (e.g., Lane and Shambaugh, 2010; Cesa-Bianchi et al., 2018).

The homeownership-weighted maximum LTV ratio – a key determinant of domestic leverage (Cesa-Bianchi et al., 2018) – is associated with a stronger (i.e., more negative) GDP response on the tail, implying that economies with higher structural levels of leverage face higher downside risks from a shock to U.S. financial conditions. This finding is consistent with the debt-driven consumption channel for business cycle dynamics highlighted by Mian et al. (2017), as well as with amplification effects due to financial frictions (e.g., Guerrieri and Iacoviello, 2017).[17]

Higher financial openness is not significantly associated with increased downside risk from U.S. financial shocks. One the one hand, financially more open economies might be more prone to spillovers from external shocks given their more integrated nature with the global financial system (Eichengreen and Gupta, 2016). On the other hand, an open financial account could enhance resilience since a more diversified portfolio of creditors could insulate the economy from shocks to specific lenders (Edwards, 2004). Our results

[17] Cesa-Bianchi et al. (2018) find a positive association between the share of foreign currency debt as well as maximum LTV ratios and the extent to which consumption in different countries responds to international credit supply shocks in conventional VARs. They do not consider the effects on GDP growth or tail effects. The latter may be amplified by financial frictions that become binding at the tails.

VARIABLES	QIRF of ΔGDP ($Q_{10\%}$)				
	(1)	(2)	(3)	(4)	(5)
Financial openness (de facto)	-0.001	-0.008	-0.001	-0.001	-0.001
	(0.006)	(0.056)	(0.006)	(0.006)	(0.006)
Exchange rate regime	0.158***	0.091**	0.159***	0.166***	0.161***
	(0.048)	(0.039)	(0.052)	(0.051)	(0.053)
Maximum LTV x home ownership	-0.032**	-0.024**	-0.032**	-0.035**	-0.034**
	(0.015)	(0.010)	(0.015)	(0.015)	(0.016)
Financial market development	-0.014	-0.010	-0.014	-0.014	-0.015
	(0.017)	(0.014)	(0.017)	(0.017)	(0.017)
Foreign currency exposure	-0.054***	-0.037**	-0.054***	-0.053***	-0.054***
	(0.018)	(0.014)	(0.018)	(0.018)	(0.018)
US financial links			-0.002		0.012
			(0.020)		(0.029)
US trade links				-0.007	-0.013
				(0.013)	(0.019)
Constant	0.770	0.361	0.795	0.977	0.976
	(1.783)	(1.287)	(1.828)	(1.841)	(1.862)
Observations	44	40	44	44	44
R-squared	0.45	0.39	0.45	0.46	0.46
VARIABLES	QIRF of ΔGDP ($Q_{50\%}$)				
	(1)	(2)	(3)	(4)	(5)
Financial openness (de facto)	-0.002	-0.071**	-0.002	-0.002	-0.002
	(0.002)	(0.028)	(0.003)	(0.003)	(0.002)
Exchange rate regime	0.030	0.003	0.029	0.036*	0.030
	(0.020)	(0.019)	(0.022)	(0.021)	(0.022)
Maximum LTV x home ownership	-0.009	-0.009*	-0.009	-0.011*	-0.010
	(0.006)	(0.005)	(0.006)	(0.006)	(0.006)
Financial market development	-0.001	0.006	-0.001	-0.001	-0.001
	(0.007)	(0.007)	(0.007)	(0.007)	(0.007)
Foreign currency exposure	-0.007	-0.001	-0.007	-0.006	-0.007
	(0.007)	(0.007)	(0.007)	(0.007)	(0.007)
US financial links			0.001		0.013
			(0.008)		(0.012)
US trade links				-0.005	-0.011
				(0.005)	(0.008)
Constant	-0.077	-0.178	-0.087	0.069	0.067
	(0.746)	(0.629)	(0.765)	(0.765)	(0.764)
Observations	44	40	44	44	44
R-squared	0.19	0.29	0.19	0.21	0.23

Table 1: Country characteristics and responses to U.S. financial shock

Notes: Regressions of the 4-quarter sum of QIRFs for the 10% and 50% quantile on country characteristics. Column (1) reports our baseline estimates. Column (2) omits potential outliers by excluding the countries IRE, LTU, LUX, LVA; (3) includes U.S. financial links; (4) includes U.S. trade links; (5) both U.S. linkages. Robust standard errors in parentheses. ***, **, and * denote significance at the 1%, 5%, and 10% levels, respectively. LTV denotes loan to value.

suggest that these effects, conditional on controlling for the share of foreign currency debt, are either small or offset each other. Hence, taken together, our results suggest that it is not financial openness *per se* but rather exposure to certain debt instruments that makes an economy vulnerable to downside risk from U.S. financial shocks. Finally, while one may have conjectured that a more developed financial system should provide better domestic risk sharing opportunities and serve to reduce the vulnerability to foreign financial shocks, our evidence suggests that it may not effectively insulate an economy from the effects of U.S. financial shocks on macroeconomic tail risks.

Jointly, these country characteristics explain a sizable share of the cross-country variation in the GDP growth responses of the lower tail ($R^2 = .45$). The regression results remain intact when potential outliers are excluded, and when we control for direct links to the U.S. via trade and financial channels (columns (2) to (5)). Our results are consistent with experiences from past financial crises that were associated with an abrupt and sizable tightening in global financial conditions. For instance, Berkmen et al. (2012) show that the global financial crisis had a particularly detrimental effect on output in countries with high levels of leverage in the domestic financial system, while exchange rate flexibility helped to cushion the impact.

We now turn to the question of whether the vulnerability to downside risks from U.S. monetary policy shocks varies systematically across the same country characteristics. Estimates from a regression of the size of individual countries' cumulated GDP growth responses on country characteristics are shown in Table 2.

For the conditional median, the response of GDP growth does not display a robust relationship with *any* of the considered country characteristics, consistent with studies using standard VAR models which do not find a clear-cut relationship (e.g., Dedola et al., 2017; Miranda-Agrippino and Rey, 2020). By contrast, the coefficient on the exchange rate variable turns statistically significant when considering the lower tail of conditional GDP growth. In countries with relatively more flexible exchange rates, the lower tail of the GDP growth distribution responds less (i.e., becomes less negative) following a U.S.

monetary policy shock. This suggests that the classical trilemma does exist, meaning that economies with flexible exchange rates are better able to use their own monetary policies to shield themselves from foreign monetary policy shocks (e.g., Obstfeld et al., 2005; Klein and Shambaugh, 2015; Bekaert and Mehl, 2019; Obstfeld et al., 2019; Obstfeld, 2021).

We run additional specifications to test the robustness of the results concerning the GDP growth responses for the lower tail, which are reported in Table 3 in Appendix B. First, we test alternative measures of financial openness. Columns (1) and (2) show that the results are robust to using a *de jure* measure of financial openness (Chinn and Ito, 2006) and a measure of *de facto* financial openness based on the ratio of cross-border banking claims to GDP. Second, the results are robust to excluding financial centers as defined by Lane and Milesi-Ferretti (2018) (column 3).[18] Third, our results are not sensitive to including the share of countries' non-U.S. trade invoiced in U.S. dollars, measured according to Boz et al. (2020). Our country sample is reduced to 36 in this case, which increases the standard error on the homeownership-weighted maximum LTV ratio and renders it insignificant for the financial shock, but the coefficient estimate remains unchanged (column 4). Fourth, we obtain very similar results when using an alternative exchange rate classification for members of the euro area (column 5).[19] Finally, our results are robust to using the GDP responses obtained from an alternative path for future quantiles in which 10% quantile dynamics for GDP growth are only considered up

[18]The financial openness variable is statistically significant in this specification. This might highlight the fact that financial centers have high gross foreign asset and liability positions that are often not related to the domestic sector to a substantial degree.

[19]Ilzetzki et al. (2019) classify the euro area as a system of fixed exchange rates. Since we are considering U.S. financial shocks, exchange rate flexibility vis-à-vis the U.S. dollar might be important for the transmission of the shock. While the euro is freely floating against the U.S. dollar, the size of the exchange rate adjustment following a U.S. financial shock might be subject to an individual country's weight in the euro area aggregate. Therefore, we re-code the exchange rate regime as follows: For the country with the highest GDP weight (in PPP terms) we set it to fully flexible, and we record reduced flexibility with decreasing weight in aggregate euro area GDP. We obtain similar results when we use the correlation with euro area growth or inflation as weights instead of GDP.

VARIABLES	QIRF of ΔGDP ($Q_{10\%}$)				
	(1)	(3)	(3)	(4)	(5)
Financial openness (de facto)	0.001	0.063*	0.001	0.001	0.001
	(0.003)	(0.037)	(0.003)	(0.003)	(0.003)
Exchange rate regime	0.073***	0.072***	0.058**	0.070***	0.059**
	(0.024)	(0.026)	(0.026)	(0.026)	(0.026)
Maximum LTV x home ownership	-0.001	0.002	0.002	-0.001	0.001
	(0.007)	(0.007)	(0.008)	(0.008)	(0.008)
Financial market development	0.006	-0.001	0.006	0.006	0.006
	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)
Foreign currency exposure	-0.014	-0.012	-0.015	-0.014	-0.015*
	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)
US financial links			0.016		0.026*
			(0.010)		(0.014)
US trade links				0.002	-0.010
				(0.007)	(0.009)
Constant	-0.916	-0.942	-1.131	-0.987	-0.991
	(0.909)	(0.841)	(0.903)	(0.940)	(0.912)
Observations	44	40	44	44	44
R-squared	0.35	0.27	0.39	0.35	0.40
VARIABLES	QIRF of ΔGDP ($Q_{50\%}$)				
	(1)	(3)	(3)	(4)	(5)
Financial openness (de facto)	-0.001	-0.012	-0.001	-0.001	-0.001
	(0.001)	(0.016)	(0.001)	(0.001)	(0.001)
Exchange rate regime	0.020*	0.011	0.015	0.020*	0.016
	(0.011)	(0.011)	(0.011)	(0.011)	(0.011)
Maximum LTV x home ownership	-0.006*	-0.005	-0.005	-0.006	-0.005
	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)
Financial market development	-0.001	-0.001	-0.001	-0.001	-0.002
	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)
Foreign currency exposure	-0.007*	-0.004	-0.007*	-0.007	-0.007*
	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)
US financial links			0.005		0.010
			(0.004)		(0.006)
US trade links				0.000	-0.004
				(0.003)	(0.004)
Constant	0.344	0.312	0.274	0.333	0.332
	(0.396)	(0.375)	(0.399)	(0.410)	(0.403)
Observations	44	40	44	44	44
R-squared	0.29	0.18	0.31	0.29	0.33

Table 2: Country characteristics and responses to U.S. monetary policy shock

Notes: Regressions of the 4-quarter sum of QIRFs for the 10% and 50% quantile on country characteristics. Column (1) reports our baseline estimates. Column (2) omits potential outliers by excluding the countries IRE, LTU, LUX, LVA; (3) includes U.S. financial links; (4) includes U.S. trade links; (5) both U.S. linkages. Robust standard errors in parentheses. ***, **, and * denote significance at the 1%, 5%, and 10% levels, respectively. LTV denotes loan to value.

to horizon $h = 2$ and thereafter replaced by median dynamics (column 6).

5. Conclusion

Financial conditions and monetary policy in the United States are important determinants of macroeconomic tail risks across the globe. In this paper, we show that an unexpected tightening in U.S. financial conditions leads to a reduction of the lower tail of the conditional GDP growth distribution in individual countries that is about four times larger than the reduction of the median. We obtain similar results for U.S. monetary policy shocks. These findings connect the literature on downside risks to growth with the literature on the global financial cycle.

Our results indicate that the strength of the GDP growth response systematically varies with certain country characteristics for the lower tail of the conditional GDP growth distribution but not for the median. Policymakers concerned with the possibility of large negative output growth realizations should therefore pay particular attention to policy choices that expose their economies to elevated GDP tail risks arising from external shocks. Countries with high foreign currency liability exposures and high levels of household sector leverage are particularly vulnerable to downside risks to growth arising from U.S. financial shocks. Macroprudential policies that limit foreign currency exposure and household indebtedness might thus help to contain downside risks to growth. Countries with less flexible exchange rates are more vulnerable to GDP tail risks arising from both U.S. financial and monetary policy shocks. This suggests that open economies are better equipped to weather external shocks under flexible exchange rate regimes that allow national monetary and macroprudential policies to limit downside risks to growth.

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A. Estimation procedure

This section summarizes the estimation procedure. For further details and proofs, see Schüler (2020).

A.1. Multivariate Laplace distribution for multiple equation quantile regression

The Bayesian estimation of the quantile VAR exploits a multivariate Laplace distribution, as proposed by Schüler (2020). This approach extends earlier contributions that use the univariate Laplace distribution for single equation quantile regression (see, e.g., Koenker and Machado, 1999; Yu and Moyeed, 2001).

To estimate $(\mu_\tau, B_{1|\tau}, \ldots, B_{p|\tau})$, we assume that

$$u_{t|\tau} \sim \mathcal{L}_k(Cm_\tau, C\Sigma_\tau C'), \tag{13}$$

where $\mathcal{L}_k$ denotes the general multivariate Laplace distribution. m_τ and the diagonal elements of Σ_τ are defined as

$$m_\tau = (m_j) = \frac{1 - 2\tau_j}{\tau_j(1-\tau_j)} \quad \text{and} \quad \text{diag}(\Sigma_\tau) = (\sigma^2_{jj}) = \frac{2}{\tau_j(1-\tau_j)}. \tag{14}$$

Furthermore, C is a positive definite matrix of size $(k \times k)$ defined as $\text{diag}(c_1, \ldots, c_d)$.

Schüler (2020) proposes to use a Metropolis-within-Gibbs sampler, exploiting a mixture representation of the multivariate Laplace distribution, which is given by

$$u_{t|\tau} = Cm_\tau w_t + \sqrt{w_t} C \Sigma_\tau^{1/2} z_t. \tag{15}$$

The representation allows to use commonly known results for the estimation as

$$\mathbf{y}_t | \mu_\tau, B_{1|\tau}, \ldots, B_{p|\tau}, \Sigma_\tau, C, w_t, \mathcal{F}_{t-1} \sim \mathcal{N}_k, \tag{16}$$

where w_t denotes a standard exponential random variable ($w_t \sim \mathcal{E}(1)$) and z_t a k-dimensional standard multivariate normal random variable ($z_t \sim \mathcal{N}_k(0, I_k)$), with I_k being an identity matrix of dimension k. Additionally, let $\Sigma_\tau^{1/2}$ represent the square root matrix Σ_τ that yields $\left(\Sigma_\tau^{1/2}\right)\left(\Sigma_\tau^{1/2}\right)' = \Sigma_\tau$.

A.2. Posteriors

For introducing posteriors, we use the compact form notation as given by

$$y = (I_d \otimes X)\beta_\tau + (Cm_\tau \otimes I_T)w + \left(C\Sigma_\tau^{1/2} \otimes W^{1/2}\right) z, \tag{17}$$

where $y = \mathrm{vec}(y_1, \ldots, y_T)'$ is a $(Tk \times 1)$ vector of observations, $X = (x_1', \ldots, x_T')'$ is a $(T \times (kp+1))$ matrix, where $x_t = (1, y_{t-1}', \ldots, y_{t-p}')$ represents a $(1 \times (kp+1))$ vector. Furthermore, β_τ denotes the column vector $\mathrm{vec}(\mu_\tau, B_{1|\tau}, \ldots, B_{p|\tau})$ of size $(k(kp+1) \times 1)$, $w = (w_1, \ldots, w_T)'$ is a $(T \times 1)$ vector and $W = \mathrm{diag}(w)$ reflects a $(T \times T)$ diagonal matrix. Therefore, $W^{1/2} = \mathrm{diag}(\sqrt{w_1}, \ldots, \sqrt{w_T})$. $z = \mathrm{vec}(z_1, \ldots, z_T)$ denotes a $(Tk \times 1)$ vector of multivariate standard normal random variables.

Conditional posteriors of β_τ and Σ_τ: We assume an independent normal-inverse-Wishart ($\mathcal{IW}$) prior:

$$\beta \sim \mathcal{N}(\underline{\beta}, \underline{V}) \quad \text{and} \quad \Sigma \sim \mathcal{IW}(\underline{\Sigma}, \underline{\nu}). \tag{18}$$

Prior times likelihood yields the standard posterior probability density functions:

$$\beta_\tau | y, \Sigma_\tau, C, w \sim \mathcal{N}(\overline{\beta}_\tau, \overline{V}_\tau) \quad \text{and} \quad \Sigma_\tau | y, \beta_\tau, C, w \sim \mathcal{IW}(\overline{\Sigma}_\tau, \overline{\nu}), \tag{19}$$

where

$$\overline{V}_\tau = [\underline{V} + ((C\Sigma_\tau C')^{-1} \otimes (X'W^{-1}X))]^{-1} \tag{20}$$

$$\overline{\beta}_\tau = \overline{V}_\tau[\underline{V}^{-1}\underline{\beta} + ((C\Sigma_\tau C')^{-1} \otimes X'W^{-1})(y - (Cm_\tau \otimes I_T)w)] \tag{21}$$

and

$$\overline{\nu} = \underline{\nu} + T \tag{22}$$

$$\overline{\Sigma}_\tau = \underline{\Sigma} + (C')^{-1}(Y - XB_\tau - w(Cm_\tau)')'W^{-1}(Y - XB_\tau - w(Cm_\tau)')(C)^{-1}, \tag{23}$$

where $B_\tau = (\mu_\tau, B_{1|\tau}, \ldots, B_{p|\tau})'$.

Conditional probability density function of the latent variable w_t: The conditional probability density of w_t is proportional to

$$f(w_t|y_t, B_\tau, \Sigma_\tau, C, \mathcal{F}_{t-1}) \propto w_t^{-k/2} \exp\left(-\frac{1}{2}\left(a_{t|\tau}w_t^{-1} + b_\tau w_t\right)\right), \tag{24}$$

with $a_{t|\tau} = (y_t - \mu_\tau - \sum_{i=1}^{p} B_{i|\tau}y_{t-i})'(C\Sigma_\tau C')^{-1}(y_t - \mu_\tau - \sum_{i=1}^{p} B_{i|\tau}y_{t-i})$ and $b_\tau = 2 + m_\tau'{\Sigma_\tau}^{-1}m_\tau$. This implies that w_t, conditional on the latter parameters, is proportional to a generalized inverse Gaussian with the following parameters:

$$w_t|y_t, \Sigma_\tau, C, B_\tau, \mathcal{F}_{t-1} \sim \mathcal{GIG}\left(-k/2 + 1, a_{t|\tau}, b_\tau\right). \tag{25}$$

Conditional posterior of C: We assume a noninformative prior for C, i.e. let

$$f(C) = \text{constant}. \tag{26}$$

In this case, the conditional posterior of C follows the likelihood of a $\mathcal{L}_k(Cm_\tau, \Sigma_{\tau\star})$, where $\Sigma_{\tau\star} = C\Sigma_\tau C'$. It is given by

$$f(C|y, \beta_\tau, \Sigma_\tau) \propto \prod_{t=1}^{T} \frac{2\exp\left((y_t - B_\tau' x_t')'\Sigma_{\tau\star}^{-1} Cm_\tau\right)}{(2\pi)^{k/2}|\Sigma_{\tau\star}|^{1/2}} \left(\frac{(y_t - B_\tau' x_t')'\Sigma_{\tau\star}^{-1}(y_t - B_\tau' x_t')}{2 + m_\tau' \Sigma_\tau^{-1} m_\tau}\right)^{(-k/2+1)}$$
$$K_{(-k/2+1)}\left(\sqrt{(2 + m_\tau'\Sigma_\tau^{-1} m_\tau)((y_t - B_\tau' x_t')'\Sigma_{\tau\star}^{-1}(y_t - B_\tau' x_t'))}\right), \quad (27)$$

where $K_{(-k/2+1)}(\cdot)$ reflects the modified Bessel function of the second kind of order $-k/2+1$.

A.3. Metropolis-within-Gibbs sampler

We use a Gibbs sampler to draw β_τ and w_t. The draws of the correlations contained in Σ_τ and the scaling factors in C are as follows.

For Σ_τ, we use the conditional posterior of Σ_τ and standardize each draw, such that quantile restrictions on the Laplace distribution remain fixed. Specifically, Σ_τ may be decomposed as

$$\Sigma_\tau = S_\tau R S_\tau, \quad (28)$$

where R denotes the correlation matrix with ones on the diagonal and ρ_{lk} as off-diagonal elements and $S_\tau = \text{diag}(\sigma_{\tau_1}, \ldots, \sigma_{\tau_k})$. Therefore, each draw of Σ_τ can be rearranged using

$$R = S_\tau^{-1}\Sigma_\tau S_\tau^{-1}, \quad (29)$$

meaning that the diagonal elements of Σ_τ remain unchanged. Given a new correlation matrix R, the covariance matrix Σ_τ can be updated using Equation (28).

For the draw of C, we use a random walk Metropolis-Hastings (MH) algorithm, since C appears in the mean and the variance of the conditional distribution of y_t (e.g. Equation (16)). Given a new draw of C, called C^*, and the last draw $C^{(n-1)}$, where $n \in \{1, \ldots, N\}$, the acceptance probability is (calibrated to be between 0.2 and 0.5)

$$\alpha_{\mathrm{MH},C}(C^{(n-1)}, C^{*}) = \min\left[\frac{f\left(C^{*}|y, \beta_{\tau}^{(n)}, \Sigma_{\tau}^{(n)}, w^{(n)}\right)}{f\left(C^{(n-1)}|y, \beta_{\tau}^{(n)}, \Sigma_{\tau}^{(n)}, w^{(n)}\right)}, 1\right]. \tag{30}$$

Next, we depict the algorithm, which we specify to only accept stationary draws of β_{τ}.

A. Define prior distribution for β_{τ} and Σ_{τ} and set starting values $\beta_{\tau}^{0}, \Sigma_{\tau}^{0}$ and C^{0}. Set variance of the random walk innovation used in the MH step, c.

B. Repeat for $n = 1, 2, \ldots, N$.

1. Gibbs Step 1: For $t = 1, \ldots, T$: Draw $w_t^{(n)}|y_t, \beta_{\tau}^{(n-1)}, \Sigma_{\tau}^{(n-1)}, C^{(n-1)}$.

2. Gibbs Step 2: Draw ${\beta_{\tau}}^{(n)}|y, \Sigma_{\tau}^{(n-1)}, C^{(n-1)}, w^{(n)}$.

3. Gibbs Step 3: (i) Draw $\Sigma_{\tau}^{(n)}|y, \beta_{\tau}^{(n)}, C^{(n-1)}, w^{(n)}$; (ii) Calculate $R^{(n)} = S_{\tau}^{-1}\Sigma_{\tau}^{(n)}S_{\tau}^{-1}$; (iii) Set $\Sigma_{\tau}^{(n)} = S_{\tau}R^{(n)}S_{\tau}$.

4. MH Step 1: (i) Draw $v_{**} \sim \mathcal{N}(0, cI_k)$; (ii) Calculate $(c_1^{*}, \ldots, c_k^{*})' = \left(c_1^{(n-1)}, \ldots, c_d^{(n-1)}\right)' + v_{**}$; (iii) Evaluate $\alpha_{\mathrm{MH},\mathbf{C}}$; (iv) Draw $u_{**} \sim \mathcal{U}(0, 1)$; (v) If $u_{**} \leq \alpha_{\mathrm{MH},C}$ set $\left(c_1^{(n)}, \ldots, c_k^{(n)}\right)' = (c_1^{*}, \ldots, c_k^{*})'$; (vi) else set $\left(c_1^{(n)}, \ldots, c_k^{(n)}\right)' = \left(c_1^{(n-1)}, \ldots, c_k^{(n-1)}\right)'$.

B. Additional figures and tables

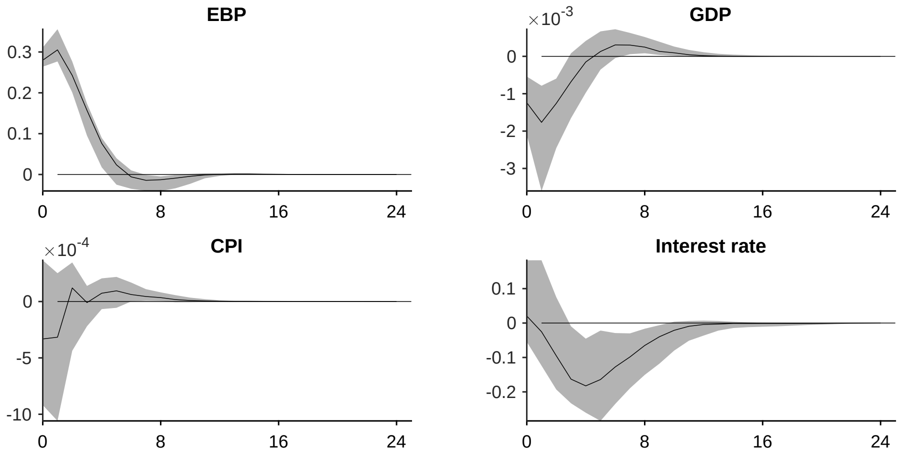

Figure B.1: Impulse responses to an EBP shock

Notes: The solid lines are the (cross-country) median group estimates of the mean IRF obtained from a standard VAR following a tightening of U.S. financial conditions scaled to one standard deviation of the unconditional EBP. Shaded areas represent one standard deviation around the median group estimator, indicating heterogeneity in responses across countries.

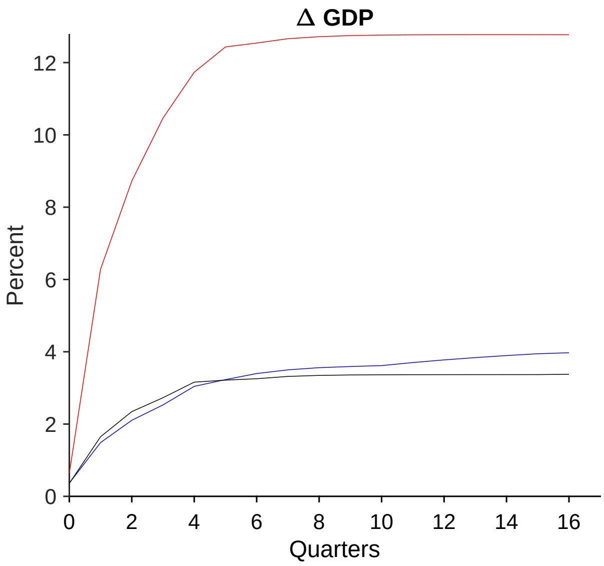

Figure B.2: Forecast error variance share of GDP growth accounted for by U.S. financial shocks.

Notes: Median group estimate of the forecast error variance share due to a one-standard-deviation increase in the EBP obtained from the baseline quantile VAR model. The black line depicts the forecast error variance share for the median of ΔGDP, the blue line shows the forecast error variance share for the 90% quantile of ΔGDP, and the red line is the forecast error variance share for the 10% quantile of ΔGDP.

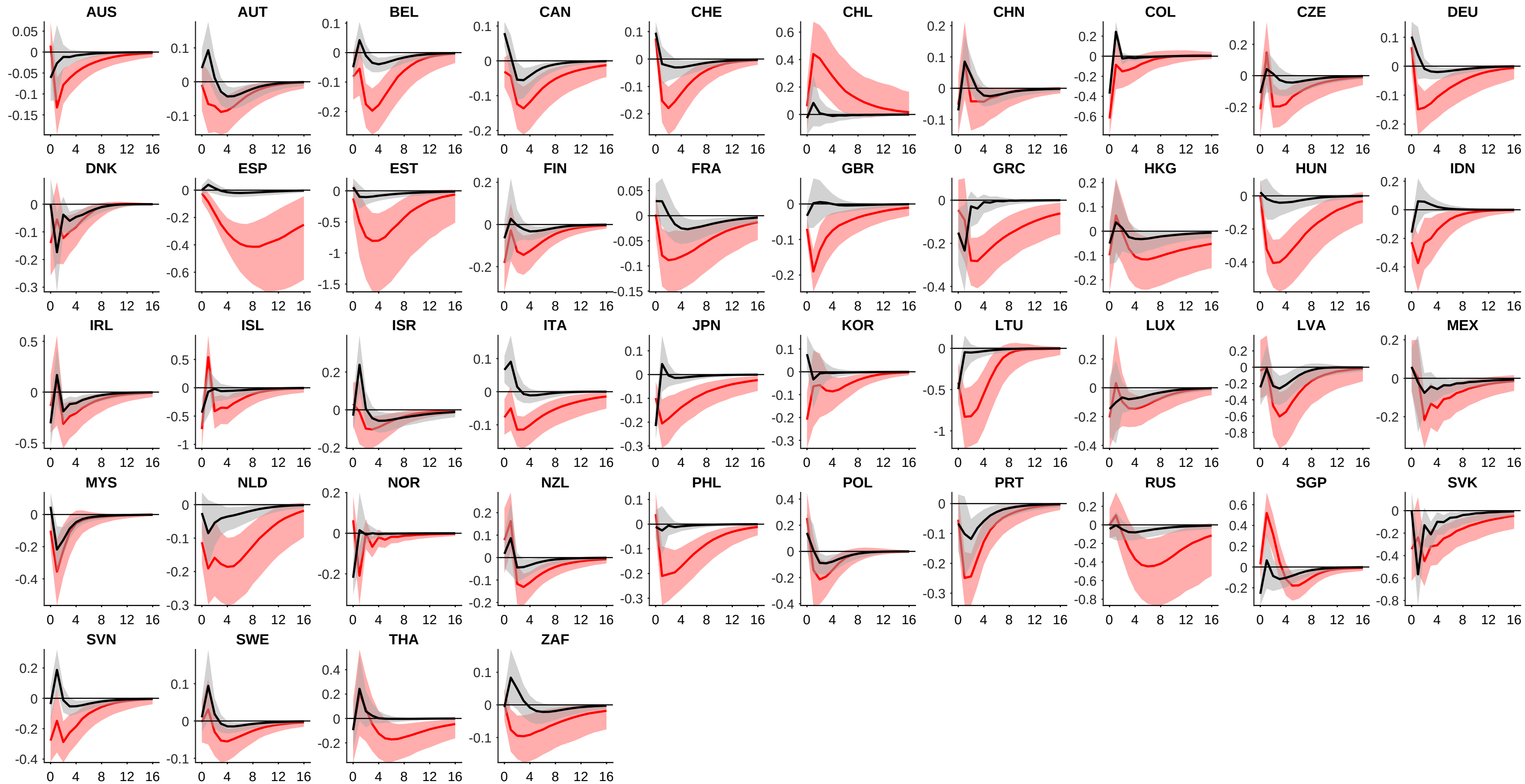

Figure B.3: Effects of a U.S. monetary policy shock: Individual country responses

Notes: Median of the posterior estimate of quantile impulse response functions (QIRFs) (solid lines) with posterior 68% probability bands (shaded areas). The black lines represent the QIRFs for the median of the conditional GDP growth distribution, and the red lines are the QIRFs for the 10% quantile of the conditional GDP growth distribution. The x-axis shows quarters after the shock.

VARIABLES	QIRF of ΔGDP to U.S. financial shock ($Q_{10\%}$)					
	(1)	(2)	(3)	(4)	(5)	(6)
Financial openness	-1.328	-0.050	-0.163**	-0.003	-0.000	-0.004
	(0.839)	(0.041)	(0.076)	(0.007)	(0.005)	(0.004)
Exchange rate regime	0.124**	0.143**	0.101*	0.184**	0.135***	0.128***
	(0.051)	(0.055)	(0.053)	(0.089)	(0.045)	(0.047)
Maximum LTV x home ownership	-0.027**	-0.029***	-0.023**	-0.024	-0.024**	-0.024***
	(0.010)	(0.010)	(0.010)	(0.022)	(0.010)	(0.008)
Financial market development	0.015	0.010	0.015	0.009	0.008	0.005
	(0.014)	(0.011)	(0.012)	(0.024)	(0.013)	(0.012)
Foreign currency exposure	-0.021***	-0.019**	-0.036**	-0.022*	-0.013*	-0.015**
	(0.008)	(0.008)	(0.016)	(0.011)	(0.007)	(0.006)
USD trade share				-0.005		
				(0.010)		
Constant	0.231	-0.588	-1.893*	-0.844	-1.170	-0.395
	(1.195)	(1.169)	(0.978)	(2.663)	(1.240)	(0.979)
Observations	44	44	37	36	44	44
R-squared	0.48	0.44	0.57	0.48	0.47	0.47
VARIABLES	QIRF of ΔGDP to U.S. monetary policy shock ($Q_{10\%}$)					
	(1)	(2)	(3)	(4)	(5)	(6)
Financial openness	-0.571	0.039	-0.027	0.005**	0.004*	0.003
	(0.365)	(0.031)	(0.042)	(0.002)	(0.002)	(0.002)
Exchange rate regime	0.052*	0.073**	0.082**	0.080**	0.056***	0.064**
	(0.027)	(0.030)	(0.031)	(0.036)	(0.020)	(0.026)
Maximum LTV x home ownership	-0.001	-0.003	0.001	0.000	-0.000	-0.002
	(0.006)	(0.006)	(0.005)	(0.007)	(0.006)	(0.006)
Financial market development	0.015**	0.009	0.011*	0.001	0.011*	0.006
	(0.006)	(0.006)	(0.006)	(0.006)	(0.006)	(0.006)
Foreign currency exposure	-0.005	-0.006	-0.007*	-0.013**	-0.002	-0.006
	(0.005)	(0.005)	(0.004)	(0.005)	(0.004)	(0.004)
USD trade share				0.007		
				(0.007)		
Constant	-0.822	-0.919	-1.176*	-0.656	-1.326*	-0.695
	(0.753)	(0.682)	(0.622)	(0.763)	(0.707)	(0.645)
Observations	44	44	37	36	44	44
R-squared	0.39	0.36	0.41	0.45	0.36	0.34

Table 3: Country characteristics and responses to U.S. financial and monetary policy shocks - Additional robustness checks

Notes: Results of regressing the four-quarter sum of different GDP growth (for the 10% quantile) responses to a one-standard-deviation U.S. financial shock (panel 1) or monetary policy shock (panel 2) on country characteristics. Columns report alternative specification with: (1) de jure measure of financial openness (Chinn and Ito, 2008); (2) financial openness measured by cross-border banking claims/GDP; (3) exclude financial centers as defined by Lane and Milesi-Ferretti (2018) (BEL, CHE, GBR, HKG, IRL, LUX, NLD, SGP); (4) including the share of countries' non-US trade invoiced in US dollar (Boz et al., 2020); (5) exchange rate regime adjusted for euro area members according to PPP-GDP weight; and (6) GDP response from an alternative path for future quantiles in which, after two quarters following the shock, the quantile value for GDP growth switches from 10% (90%) to 50%. ***, **, and * denote significance at the 1%, 5%, and 10% levels, respectively. LTV denotes loan to value.