US Monetary Policy and the Global Financial Cycle

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Abstract

US monetary policy shocks induce comovements in the international financial variables that characterize the 'Global Financial Cycle.' A single global factor that explains an important share of the variation of risky asset prices around the world decreases significantly after a US monetary tightening. Monetary contractions in the US lead to significant deleveraging of global financial intermediaries, a decline in the provision of domestic credit globally, strong retrenchments of international credit flows, and tightening of foreign financial conditions. Countries with floating exchange rate regimes are subject to similar financial spillovers.

Keywords: Monetary Policy; Global Financial Cycle; International spillovers; Identification with External Instruments

JEL Classification: E44, E52, F33, F42

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1 Introduction

Observers of balance of payment statistics and international investment positions all agree: the international financial landscape has undergone massive transformations since the 1990s. Financial globalization is upon us in a historically unprecedented way, and we have probably surpassed the pre-WWI era of financial integration celebrated by Keynes in "The Economic Consequences of the Peace." At the same time, the role of the United States as the hegemon of the international monetary system has largely remained unchanged, and has long outlived the end of Bretton Woods, as emphasized in e.g. Farhi and Maggiori (2018) and Gourinchas and Rey (2017). The rising importance of cross-border financial flows and holdings has been documented in the literature.[1] What has not been explored as much, however, are the consequences of financial globalization for the workings of national financial markets, and for the transmission of US monetary policy beyond the domestic border. How do international capital flows affect the international transmission of monetary policy? What are the effects of global banking on fluctuations in risky asset prices, and on credit growth and leverage in different economies? Using monthly data since 1980, we study how the existence of a 'Global Financial Cycle' (Rey, 2013) shapes the global financial spillovers of US monetary policy.

Monetary policy operates though multiple, complementary channels. In a standard Keynesian or neo-Keynesian world, output is demand determined in the short-run, and monetary policy stimulates aggregate consumption and investment (see Woodford, 2003 and Gali, 2008 for classic discussions). In models with frictions in capital markets, expansionary monetary policy also leads to an increase in the net worth of borrowers, either financial intermediaries or firms, which in turn boosts lending. This is the *credit channel* of monetary policy (Bernanke and Gertler, 1995). Other papers have instead analyzed the *risk-taking channel* of monetary policy in which it is the risk profile of financial intermediaries that plays a key role, and loose monetary policy relaxes leverage constraints (Borio and Zhu, 2012; Bruno and Shin, 2015a; Coimbra and Rey, 2017). In this paper, we explore empirically the *international* transmission of monetary policy that occurs through financial intermediation and global asset prices, an area that has been largely neglected

[1] See e.g. Lane and Milesi-Ferretti, 2007 and, for a recent survey, Gourinchas and Rey (2014).

by the literature.[2]

Using a dynamic factor model, we first document the existence of a *unique* global factor in international risky asset prices that explains over 20% of the variance in the data. With a global Bayesian VAR, we then study the international transmission of US monetary policy that is mediated through the reaction of asset prices, of global credit and capital inflows, and of the leverage of financial intermediaries; these are the variables that characterize the Global Financial Cycle. Our analysis is motivated by the US dollar being an important funding currency for intermediaries, and by the fact that a large portion of portfolios worldwide are denominated in dollars.[3] We identify US monetary policy shocks using an external instrument constructed from high-frequency price adjustments in the federal funds futures market around FOMC announcements, following the lead of Gürkaynak, Sack and Swanson (2005) and Gertler and Karadi (2015). At the same time, the use of a rich-information VAR ensures that we control for a wealth of other shocks, both domestic and international, to which the Fed endogenously reacts, above and beyond what is anticipated by market participants.[4]

We find evidence of powerful financial spillovers of US monetary policy to the rest of the world. When the Federal Reserve tightens, domestic demand contracts, as do prices. The domestic financial transmission is visible through the rise of corporate spreads, the contraction of lending, and the sharp fall in the prices of assets, such as housing and the stock market. But, importantly, we also document significant variations in the Global Financial Cycle, that is, the shock induces significant fluctuations in financial activity on a global scale. Risky asset prices, summarized by the single global factor, contract very significantly. This is accompanied by a deleveraging of global banks both in the US and Europe, and a surge in aggregate risk aversion in global asset markets. The supply of global credit contracts, and there is an important retrenchment of international credit flows that is particularly pronounced for the banking sector. International corporate bond spreads also rise on impact, and significantly so. These results are consistent

[2]See Rey (2016), Bernanke (2017) and Jorda, Schularick, Taylor and Ward (2018) for longer discussions.

[3]For a recent study of the international reserve currency role of the dollar see Farhi and Maggiori (2018). Gopinath (2016) analyzes the disproportionate role of the dollar in trade invoicing, and Gopinath and Stein (2017) the synergies between some of those roles.

[4]For more detailed discussions see Miranda-Agrippino (2016) and Miranda-Agrippino and Ricco (2018).

with a powerful transmission channel of US monetary policy across borders, via financial conditions. The contraction of domestic credit and international liquidity that follows the US monetary policy tightening is confirmed also for the subset of countries that have a floating exchange rate regime.

The importance of international monetary spillovers and of the world interest rate in driving capital flows has been pointed out in the classic work of Calvo et al. (1996).[5] Some recent papers have fleshed out the role of intermediaries in channeling those spillovers.[6] Our empirical results on the transmission mechanism of monetary policy via its impact on risk premia, spreads, and volatility, are related to those of Gertler and Karadi (2015) and Bekaert, Hoerova and Duca (2013) obtained in the domestic US context.[7] A small number of papers have analyzed the effect of US monetary policy on leverage and on the VIX (see e.g. Passari and Rey, 2015; Bruno and Shin, 2015b).[8] Using a rich-information Bayesian VAR permits, we believe for the first time, to jointly evaluate the response of financial, monetary and real variables, in the US and abroad. Moreover, by relying on an instrumental variable for the identification of US monetary policy shocks, we can dispense from making implausible timing restrictions on the response of our variables of interest.

The paper is organized as follows. In Section 2, we estimate a dynamic factor model on world asset prices and show that one global factor explains a large part of the common variation of the data. In Section 3, we estimate a Bayesian VAR identified using external instruments to analyze the interaction between US monetary policy and the Global Financial Cycle. Section 4 presents a simple theoretical framework featuring heterogeneous investors to interpret some of our results (Section 4.1), and microeconomic data on global

[5] Fratzscher (2012) and Forbes and Warnock (2012) have extended these findings significantly.

[6] Cetorelli and Goldberg (2012) use balance sheet data to study the role of global banks in transmitting liquidity conditions across borders. Using firm-bank loan data, Morais, Peydro and Ruiz (2015) find that a softening of foreign monetary policy increases the supply of credit of foreign banks to Mexican firms. Using credit registry data combining firm-bank level loans and interest rates data for Turkey, Baskaya, di Giovanni, Kalemli-Ozcan and Ulu (2017) show that increased capital inflows, instrumented by movements in the VIX, lead to a large decline in real borrowing rates, and to a sizeable expansion in credit supply. They find that the increase in credit creation goes mainly through a subset of the biggest banks.

[7] For a discussion on the transmission of unconventional US monetary policy on global risk premia see Rogers, Scotti and Wright (2018).

[8] These studies all rely on limited-information VARs (four to seven variables) and on recursive identification schemes to study the transmission of monetary policy shocks, it is therefore unclear whether their results survive a more robust identification of monetary policy shocks. The problem of omitted variables is also an important issue in small scale VARs (see Caldara and Herbst, 2019).

banks to give evidence of their risk-taking behavior (Section 4.2). Section 5 concludes. Details on data and procedures, and additional results are in Appendices at the end of the paper.

2 One Global Factor in World Risky Asset Prices

In order to summarize fluctuations in global financial markets we specify a Dynamic Factor Model for a large and heterogeneous panel of risky asset prices traded around the globe. The econometric specification, fully laid out in Appendix B, is very general, and allows for different global, regional and, in some specifications, sector specific factors.[9] The panel includes asset prices traded on all the major global markets, a collection of corporate bond indices, and commodities price series (excluding precious metals). The geographical areas covered are North America, Latin America, Europe, Asia Pacific, and Australia, and we use monthly data from 1990 to 2012, yielding a total of 858 different prices series.[10] Despite the heterogeneity of the asset markets considered, we find that the data support the existence of a single common global factor; moreover, this factor alone accounts for over 20% of the common variation in the price of risky assets from all continents.[11] The factor is plotted in Figure 1, solid line.

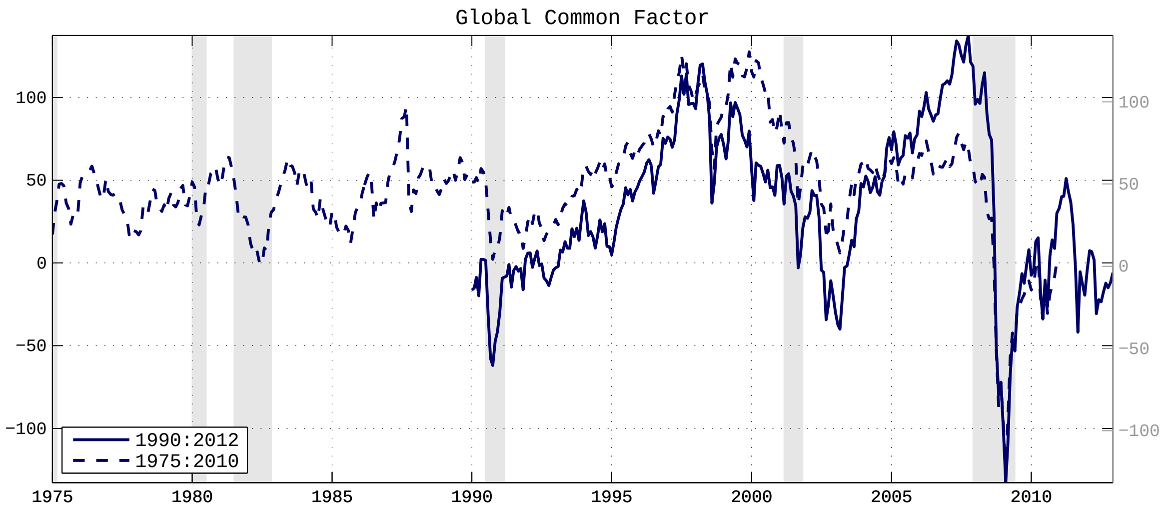
While in this instance we prefer cross-sectional heterogeneity over time length, we are conscious of the limitations that a short time span may introduce in the VAR analysis we perform in the next section. To allow more flexibility in that respect, we repeat the factor extraction on a smaller set, where only the US, Europe, Japan and commodity prices are included, but the time series go back to 1975. In this case the sample counts 303 series. The estimated global factor for the longer sample is the dashed line in Figure 1. Similar to the benchmark case, for this narrower panel too we find evidence of one global

[9] A similar specification has been adopted by Kose et al. (2003) and Kose et al. (2012) for real variables; they test the hypothesis of the existence of a world business cycle and discuss the relative importance of world, region and country specific factors in determining domestic business cycle fluctuations.

[10] All the details on the construction and composition of the panels, shares of explained variance, and test and criteria used to inform the parametrization of the model (Table B.2) are reported in Appendix B. We fit to the data a Dynamic Factor Model (Stock and Watson, 2002a,b; Bai and Ng, 2002; Forni et al., 2000, among others) where each price series is modelled as the sum of a global, a regional, and an asset-specific component. All price series are taken at monthly frequency using end of month figures.

[11] We formally test for the numbers of factors in our large panel of asset prices and find that the data support one common global factor. Results are reported in Table B.2 in the Appendix.

Figure 1: global factor in risky asset prices



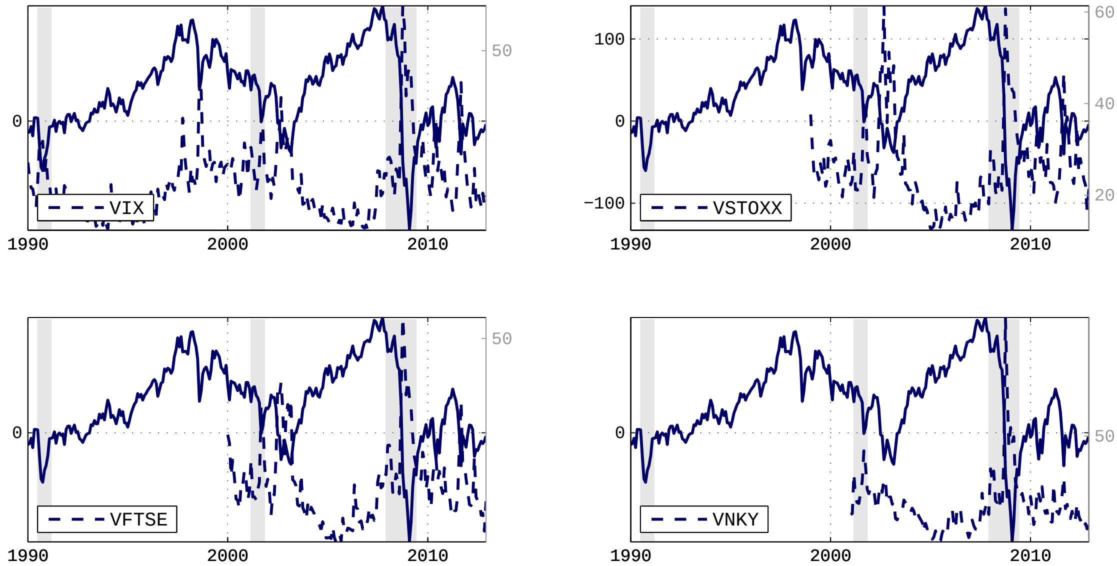
Note: The Figure plots the estimates of the global factor for the 1975:2010 sample (dotted line) together with the estimates on the wider, shorter sample 1990:2012 (solid line). Shaded areas denote NBER recession dates.

factor. In this case, however, the factor accounts for about 60% of the common variation in the data (see Table B.2). For both samples, factors are obtained via cumulation of those estimated on the stationary, first-differenced (log) price series, and are therefore consistently estimated only up to a scale and an initial value (see Bai and Ng, 2004, and Appendix B).[12] As a way of normalization, we rotate the factor such that it correlates positively with the major stock market indices in our sample, i.e. an increase in the index is interpreted as an increase in global asset prices.

Figure 1 shows that movements in the factors are consistent with both the US recession periods as identified by the NBER (shaded areas), and with major worldwide events. The index declines with all the recession episodes but remains relatively stable until the beginning of the nineties, when a sharp and sustained increase is recorded. The increase lasts until 1997-1998 when major global events like the Russian default, the LTCM bailout, the East Asian Crisis and finally the burst of the *dot-com* bubble reverse the increasing path. Starting from the beginning of 2003 the index increases again until the beginning of the third quarter of 2007. At that point, with the collapse of the

[12]This implies that positive and negative values displayed in the chart do not convey any specific information per se. Rather, it is the overall shape and the turning points that are of interest.

FIGURE 2: GLOBAL FACTOR AND VOLATILITY INDICES

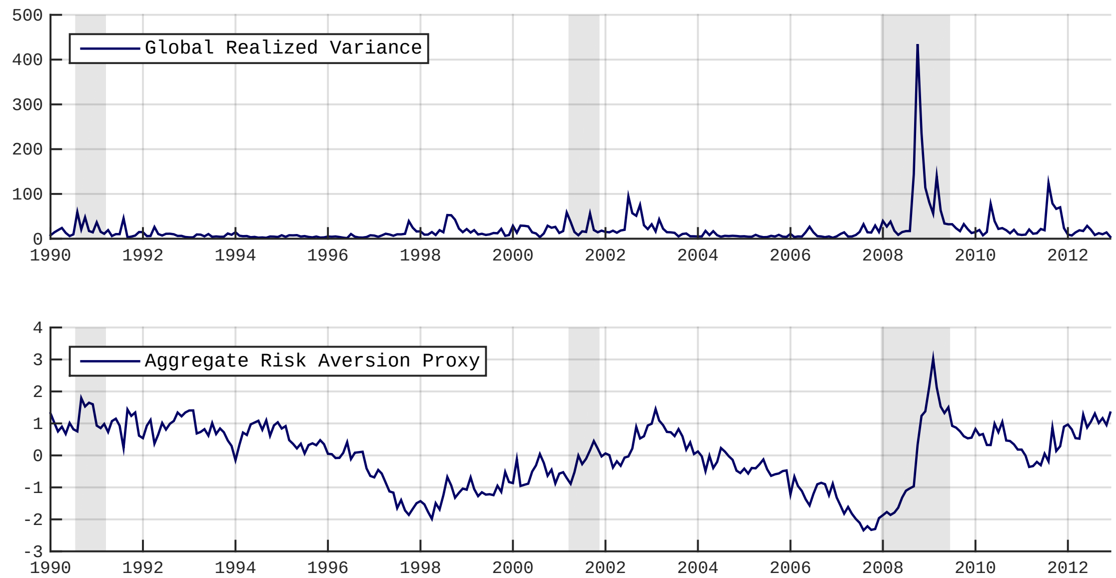


Note: Clockwise from top-left panel, the global factor (solid line) together with major volatility indices (dotted lines): VIX (US), VSTOXX (EU), VNKY (JP) and VFTSE (UK). Shaded grey areas highlight NBER recession times.

subprime market, the first signals of increased vulnerability in financial markets become visible. This led to an unprecedented plunge.

In order to provide some interpretation for our estimated global factor, we note that in a large class of asset pricing models, including in the stylized framework that we present in Section 4, the common component of risky asset prices is a function of aggregate volatility, and of the degree of aggregate risk aversion in the market. In particular, in the simple model of Section 4 with heterogeneous financial intermediaries that differ in their propensity to take on risk, the evolving distribution of wealth between different types of intermediaries gives rise to a time-varying degree of aggregate risk aversion. This interpretation of the factor, as reflecting volatility and aggregate risk aversion, is closely related to that of indices of implied volatility. In Figure 2 we highlight the comovement of our factor with the VIX, the VSTOXX, the VFTSE and the VNKY, which represent the markets included in our sample. These indices capture both the price and quantity of risk, and hence reflect both expectations about future volatility, and risk aversion. Because of our chosen normalization, we expect our factor to correlate negatively with

Figure 3: global factor decomposition



Note: [top panel] Monthly global realized variance measured using daily returns of the MSCI Index. [bottom panel] Index of aggregate risk aversion calculated as (the inverse of) the residual of the projection of the global factor onto the realized variance. Shaded grey areas highlight NBER recession times. *Source:* Global Financial Data and authors calculations.

the implied volatility indices. Indeed, this is clearly visible in the charts of Figure 2; the factor and the implied volatility indices display a remarkable common behaviour and peaks consistently coincide within the overlapping samples.[13]

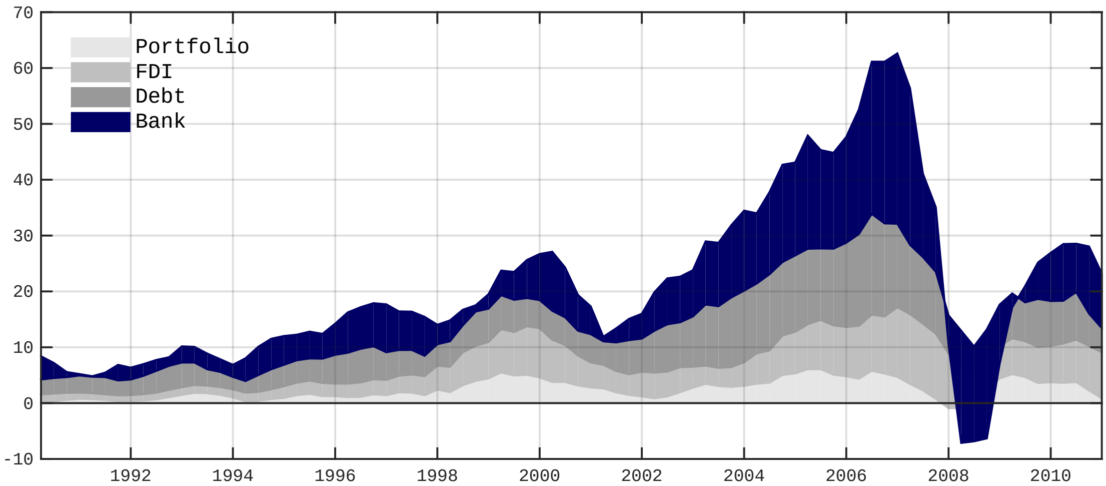
Based on the intuition offered by our simple model, we separate the aggregate risk aversion and volatility components in our global factor. We first estimate a monthly series of realized global volatility using daily returns of the global MSCI Index.[14] Second, we calculate a proxy for aggregate risk aversion as the inverse of the centred residuals of the projection of the global factor on the realized variance.[15] The results of this exercise

[13]While the comparison with the VIX is somehow facilitated by the length of the CBOE index, the same considerations extend to all other indices analyzed. Comparison with other 'risk indices' such as the GZ-spread of Gilchrist and Zakrajšek (2012) and the Baa-Aaa corporate bond spread (not reported) show that these indices also display some commonalities, even if the synchronicity is slightly less obvious.

[14]We work under the assumption that monthly realized variances calculated summing over daily returns provide a sufficiently accurate proxy of realized variance at monthly frequency (see Andersen et al. (2003)).

[15]Specifically, the proxy for aggregate risk aversion is recovered from the following regression: $GFAC_t = \alpha + \beta \ln(GRVAR_t) + \varepsilon_t$, where $GFAC_t$ is the global factor expressed in log units, and $GRVAR_t$ is the realized variance of the global MSCI Index. The construction of our proxy for aggregate risk aversion is modelled along the lines of e.g. Bekaert et al. (2013), that estimate variance risk premia

Figure 4: aggregate capital flows



Note: Global flows as a percentage of world GDP. Annual moving averages. *Source:* IFS Statistics.

are summarized in Figure 3. Our monthly measure of global realized variance is in the top panel, while our index of aggregate risk aversion is in the bottom panel. Interestingly, the degree of market risk aversion that we recover from this simple decomposition is in continuous decline between 2003 and 2007. It decreases to very low levels at a time when volatility was low, global banks were prevalent and may have been the 'marginal buyers' in international financial markets. Indeed, Shin (2012) documents the large and increasing share of banks in international financial markets over that period and until 2007; subsequently, both as a consequence of the crisis and of the changes in regulation, their relative importance has declined. For illustrative purposes, we report data relative to different types of capital flows as a percentage of world GDP in Figure 4, and we further explore the connection between the sharp increase in banking flows and the decline in global risk aversion in Section 4. After end-2007, aggregate risk aversion starts increasing to jump sharply during the financial crisis and the bankruptcy of Lehman Brothers, and remains persistently at high levels.

In more complex models than the one reported in Section 4, the common component of asset prices is not only a function of realized variances and of risk aversion but also

as the difference between a measure of the implied variance (the squared VIX) and an estimated physical expected variance, which is primarily a function of realized volatilities.

of discount rates and of expected cash-flow growth. We explore this more general case in detail in Appendix E and show that this alternative specification delivers an index of aggregate risk aversion very similar to our baseline index in Figure 3.[16]

3 US Monetary Policy and Global Financial Cycle

With the US dollar being the currency of global banking, monetary actions in the US may directly influence the Global Financial Cycle (GFC) by altering the cost of funding for major global banks, and hence their leverage decisions. US monetary policy also affects the pricing of dollar assets, both in the US and abroad, through a direct discount rate channel and/or by changing the type of marginal investors in international asset markets.[17] Furthermore, monetary conditions of the centre country can also be transmitted through cross-border capital flows, or through the internal pricing of liquidity by global banks, and influence the provision of credit outside US borders (see the corroborative evidence in Morais et al., 2015 for Mexico, and in Baskaya et al., 2017 for Turkey).

To study the effects of US monetary policy on the GFC, we use rich-information VARs that provide us with a unique framework to analyze the transmission of monetary policy beyond national borders.[18] There are a number of advantages that come with this choice. Most obviously, relying on a unique specification permits addressing the effects of US monetary policy on the GFC against the background of the response of the domestic business cycle. This acts both as a complement to the analysis, and as a disciplining device to ensure that the identified shock is in fact inducing responses that do not deviate from the standard channels of domestic monetary transmission. Moreover, the dimensionality and composition of the set of variables included in the VAR greatly reduce the problem of omitted variables that generally plagues smaller systems and is

[16]Precisely, we extract risk aversion by projecting the factor on realized variances, on discount rates in the US, Germany, the UK and Japan, and on survey forecasts for output growth 12-month-ahead in the same four countries in order to proxy for expected cash flow growth. Results are very similar. For thorough discussions on estimation of price of risk versus quantity of risk, see Bekaert et al. (2019) and Zhou (2018). We also consider in Appendix E other indices of risk aversion that have been independently developed in the literature, and show our results carry through regardless of the particular proxy used.

[17]Security-level evidence provided by Schreger et al. (2017) shows that firms who finance themselves in dollars are by and large the only ones able to attract a worldwide investors base. For a model where low funding costs lower aggregate effective risk aversion and increase leverage see Coimbra and Rey (2017).

[18]Technical details on priors and estimation of the Bayesian VAR are reported in Appendix C.

likely to invalidate the identification of the structural shocks.[19]

We start by looking at how US monetary policy affects domestic real and financial conditions in a 'closed economy VAR'. Then, we augment a small set of core domestic variables with those that characterize the GFC; namely, global credit and capital inflows, the global factor in asset prices and risk aversion, and the leverage of US and European global banks. In this first version of our empirical framework, global variables are world aggregates, and bundle together countries with different exchange rate regimes. To evaluate to what extent a floating exchange rate can provide some insulation against foreign shocks, we then repeat the analysis by specifically focusing only on the subset of 'floaters', following the IMF's de-facto classification.

3.1 Identification of US Monetary Policy Shocks

We identify US monetary policy shocks using an external instrument (Stock and Watson, 2012, 2018; Mertens and Ravn, 2013). The intuition behind this approach to identification is that the mapping between the VAR innovations and the structural shock of interest can be estimated using only moments of observables, provided that a valid instrument for such shock exists. The contemporaneous transmission coefficients are a function of the regression coefficients of the VAR residuals onto the instrument, up to a normalization. Hence, given the instrument, this method ensures that we can isolate the causal effects of a US monetary policy shock on the dynamics of our large set of variables without imposing any timing restrictions on the responses. Intuitively, if the instrument correlates with the VAR innovations only via the contemporaneous monetary policy shocks, a projection of the VAR innovations on the instrument isolates variations in the variables which are solely due to this shock (Miranda-Agrippino and Ricco, 2018).

The crucial step of this identification strategy is, naturally, the choice of the instrument. We rely on high-frequency movements in federal funds futures markets around FOMC announcements to identify the monetary policy shocks, following the lead of

[19] Bańbura et al. (2010) show that a medium-scale VAR of comparable size and composition to the one used in this paper is able to correctly recover the shocks and reproduce responses that match theoretical ones. Intuitively, the large degree of comovement among macroeconomic variables makes it possible for VARs of such size to effectively summarize the information contained in large VARs typically counting over hundred variables.

Gürkaynak et al. (2005) and Gertler and Karadi (2015). Specifically, we use 30-minutes price revisions (or surprises) around FOMC announcements in the fourth federal funds futures contracts (FF4), and we construct a monthly instrument by summing up the high-frequency surprises within each month. Because these futures have an average maturity of three months, the price revision that surrounds the FOMC monetary policy announcements captures revisions in market participants expectations about the future monetary policy stance up to a quarter ahead. As observed in Miranda-Agrippino (2016) and Miranda-Agrippino and Ricco (2017), market-based monetary surprises such as the ones we use map into the shocks only under the assumption that market participants can correctly and immediately disentangle the systematic component of policy from any observable policy action. In the presence of information asymmetries, the high-frequency surprises are also a function of the information about economic fundamentals that the central bank implicitly discloses at the time of the policy announcements.[20] Failure to account for this effect may hinder the correct identification of the shocks, resulting in severe price and real activity puzzles, particularly in small VARs. Here we address this issue by relying on the rich information in our VARs. The information set in our VARs controls for a wealth of other shocks, both domestic and international, to which the Fed endogenously reacts, and allows identification of monetary policy shocks above and beyond what is expected by market participants.

In Table 1, we report first stage IV statistics of the projection of the VAR innovation for the policy interest rate (1 year rate in our case) on our instrument (FF4). For comparison, we also include first-stage statistics obtained with the narrative instrument of Romer and Romer (2004), that we have extended up to the end of 2007 (MPN). A fist-stage F statistic below 10 is an indication of potentially weak instruments (Stock et al., 2002). The three VARs in the table are (1) a closed economy 13-variable VAR that includes only US variables; (2) a global 15-variable VAR that includes GFC variables as world aggregates; (3) and a global 15-variable VAR that focuses on the subset of countries with floating exchange rates.[21]

[20] This implicit disclosure of information is referred to as the Fed information effect in Nakamura and Steinsson (2018), and the signalling channel of monetary policy in Melosi (2017). The concept is similar to the Delphic component of forward guidance announcements in Campbell, Evans, Fisher and Justiniano (2012).

[21] All VARs are monthly and estimated with 12 lags over the sample 1980-201. Details on the compo-

Table 1: Tests for Instruments Relevance

Domestic VAR (1)		F-stat	90% posterior ci		reliability	90% posterior ci	
	FF4	17.930	[6.675	22.673]	0.496	[0.434	0.540]
	MPN	10.947	[4.264	16.246]	0.187	[0.132	0.251]
Global VAR (2)							
	FF4	14.788	[3.239	18.010]	0.530	[0.470	0.573]
	MPN	2.278	[0.106	5.698]	0.258	[0.171	0.317]
Global VAR (3)							
	FF4	14.901	[3.116	18.631]	0.529	[0.476	0.577]
	MPN	2.756	[0.139	6.216]	0.255	[0.170	0.312]

Note: First-stage F statistics, statistical reliability and 90% posterior coverage intervals. Candidate instruments are surprises in the three-months-ahead (FF4) federal fund futures and an extension to the narrative instrument of Romer and Romer (2004) up to 2007. VAR innovations are from monthly BVAR(12) estimated from 1980 to 2010. First-stage regressions are run on the overlapping sample between the VAR innovations and each instrument.

Results in Table 1 show that in a domestic context either instrument attains satisfactory levels of relevance. As we discuss in the next subsection, the two instruments also retrieve relatively similar dynamic responses to a monetary policy shock in the domestic VAR. The relevance of the narrative series deteriorate dramatically in both open economy global VARs with F statistics dropping well below 10. In contrast, the first stage IV statistics associated to the high-frequency based identification are only marginally altered in the three cases. This confirms the strong informative content of our preferred instrument.[22]

3.2 The International Transmission of US Monetary Policy through the Global Financial Cycle

We present our results in the form of dynamic responses to a US monetary policy shock that is normalized to increase the policy rate by 1% on impact. We use the 1-year rate as monetary policy variable; this, coupled with the 3-month horizon embedded in the external instrument implies that we capture standard monetary policy shocks that affect the

sition of each VAR are reported in Table 2 in the next subsection.

[22] Another paper using high frequency external instruments for the identification of US monetary policy shocks and their effects on financial markets is Ha (2016).

Table 2: Variables in VARs

Variable Name	Source	Model (1)	(2)	(3)	(4)	(5)	(6)
Industrial Production	FRED-MD	•	•	•	•	•	•
Capacity Utilization	FRED-MD	•					
Unemployment Rate	FRED-MD	•					
Housing Starts	FRED-MD	•					
CPI All	FRED-MD	•					
PCE Deflator	FRED-MD	•	•	•	•	•	•
1Y Treasury Rate	FRED-MD	•	•	•	•	•	•
Term Spread (10Y-1Y)	FRED-MD	•					
BIS Real EER	BIS	•	•	•	•	•	•
GZ Excess Bond Premium	Gilchrist and Zakrajšek (2012)	•					
Mortgage Spread	Gertler and Karadi (2015)	•					
House Price Index	Shiller (2015)	•					
S&P 500	FRED-MD	•					
Global Factor	Datastream & OC		•	•	•	•	•
Global Risk Aversion	OC		•	•	•	•	
Global Real Economic Activity Ex US	Baumeister and Hamilton (2019) & OC		•	•	•	•	
Global Domestic Credit	IMF-IFS*		•				•
Global Domestic Credit Ex US	IMF-IFS*			•			
US Total Nonrevolving Credit	FRED-MD			•		•	
Global Inflows All Sectors	BIS*		•				•
Global Inflows to Banks	BIS*			•			
Global Inflows to Non-Banks	BIS*			•			
Floaters Domestic Credit	BIS*				•	•	
Floaters Inflows All Sectors	BIS*				•		
Floaters Inflows to Banks	BIS*					•	
Floaters Inflows to Non-Banks	BIS*					•	
GZ Credit Spread	Gilchrist and Zakrajšek (2012)		•		•		•
Leverage US Brokers & Dealers	FRB Flow of Funds*		•	•	•	•	
Leverage EU Global Banks	Bankscope*		•	•	•	•	
Leverage US Banks	Bankscope*		•	•	•	•	
Leverage EU Banks	Bankscope*		•	•	•	•	
FTSE All Shares	Global Financial Data						•
GBP to 1 USD	Global Financial Data						•
UK Corporate Spread	Global Financial Data & OC						•
UK Policy Rate	Bank of England						•
DAX Index	Global Financial Data						•
EUR to 1 USD	Global Financial Data						•
GER Corporate Spread	Global Financial Data & OC						•
ECB Policy Rate	Global Financial Data & OC						•
Figures		5	6,7,8 D.1 D.2	7 D.3	9 D.4	D.5	10 D.7

Note: The table lists the variables included in the baseline domestic and global BVARs. Models correspond to (1) domestic VAR; (2) & (3) global VARs with world aggregates for GFC; (4) & (5) global VAR on subset of countries with a floating exchange rate; (6) global VAR with focus on UK and EA monetary policy and financial conditions. Variables enter the VARs in (log) levels with the exception of interest rates and spreads. OC denotes own calculations, * denotes monthly interpolation of the quarterly original variables.

fed funds rate, but also implicit and explicit Fed communication and actions that affect interest rates at longer maturities. All VARs are estimated using standard macroeconomic priors, with 12 lags at monthly frequency over the sample 1980:1 - 2010:12. Following Mertens and Ravn (2013) and Gertler and Karadi (2015), the identification step (i.e. the projection of the VAR innovations on the instrument) is run over the common sample (1990:01-2010:12).[23] We report and discuss only the IRFs for the variables of interest; full sets of IRFs are reported in Appendix D. The variables that we include in our baseline VARs are listed in Table 2, together with the composition of all the VARs we estimate for the results collected in the reminder of the section. Details on the construction of the

[23] In the appendix we report IRFs from a VAR also estimated from 1990:01 for comparison.

data are reported in the Appendix that also collects robustness tests. We report median IRFs together with 68% and 90% posterior coverage bands.

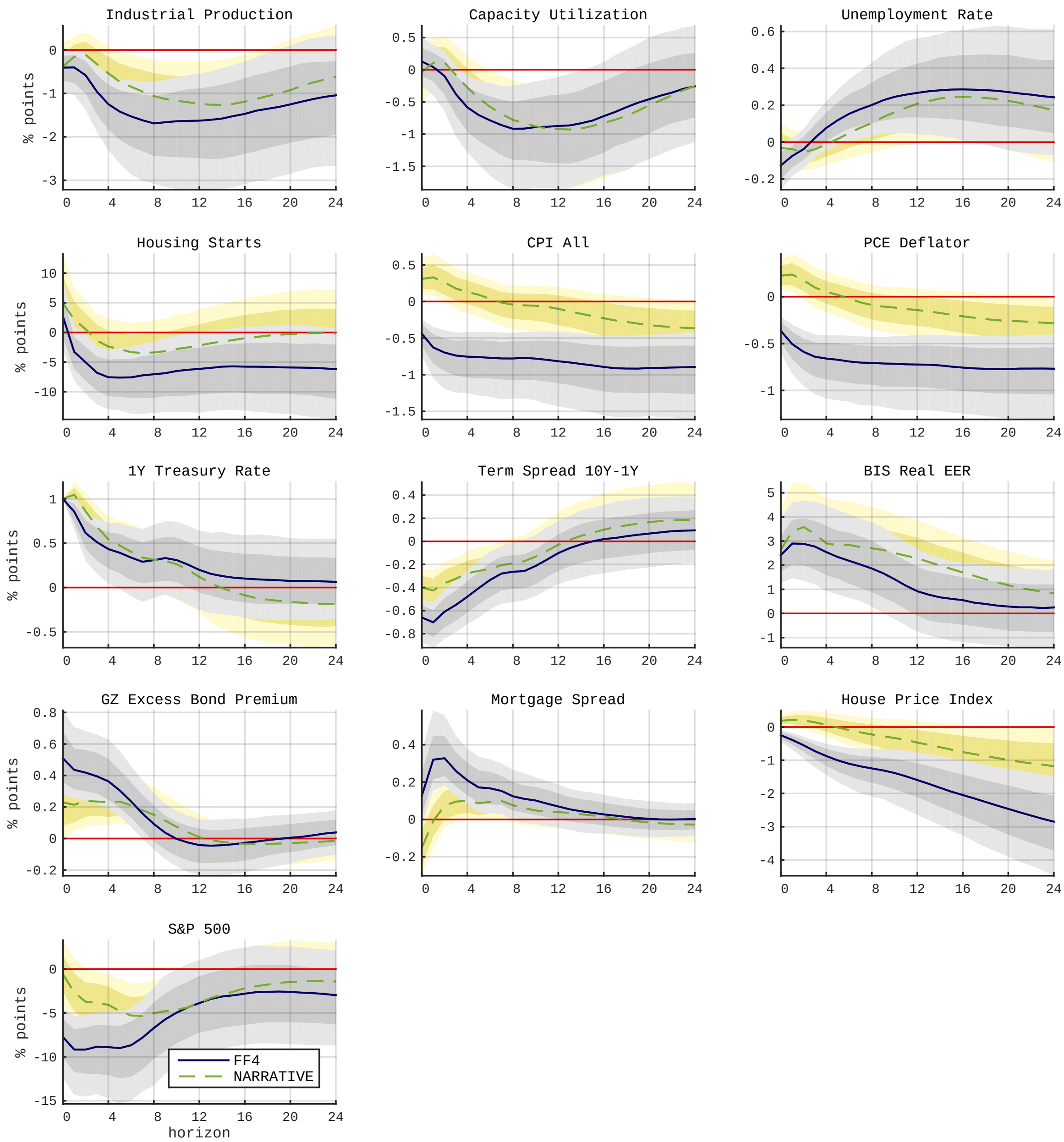
Domestic Responses We start our empirical exploration by looking at the response of the domestic financial markets and macroeconomic aggregates. To give further motivation for the choice of our instrument, Figure 5 compares the IRFs obtained with the high-frequency IV (FF4, solid lines) and the narrative IV (MPN, dashed lines). The VAR is the same in the two cases.

A contractionary monetary policy shock depresses prices and economic activity in line with the standard transmission channels. Production and capacity utilization contract, as do housing investments, while the unemployment rate rises significantly; these effects are not sudden, but build up over the horizons. Similarly, prices adjust downward. We note here that the MPN IV recovers responses that display a pronounced price puzzle. This is in contrast to our preferred identification: following an initial downward revision, prices continue to slide into negative territory, consistent with the presence of price rigidities. The shock also has important consequences for domestic financial markets. The monetary tightening at the short end decreases the term spread and induces a sudden rise in the excess bond premium variable of Gilchrist and Zakrajšek (2012) that measures corporate bond spreads net of default considerations. The response also implies increased costs of funding in the corporate market, and provides evidence of a powerful financial amplification mechanism of monetary policy shocks that operates at the domestic level. Expectations of lower economic activity and changes in discount rate are immediately priced-in in the stock market that registers a strong and sudden drop. Household finance also deteriorates substantially, with house prices falling and mortgage spreads increasing significantly.[24] Finally, the monetary contraction results in a significant appreciation of the dollar against a basket of foreign currencies.

The system of domestic dynamic responses highlights a powerful transmission of monetary policy shocks through the domestic financial markets. In the reminder of this section we will explore how monetary policy shocks spill over across borders through their effect on global financial conditions.

[24] We use the 30-year conventional mortgage spread calculated in excess of the 10-year government bond rate. We take this variable from Gertler and Karadi (2015).

Figure 5: responses of domestic business & financial cycle



Note: Closed economy responses to a contractionary US monetary policy shock that induces a 1% increase in the policy rate. [blue solid lines and grey areas] IV is the surprise in FF4 contracts, 68% & 90% posterior coverage bands. [green dashed lines and yellow areas] IV is an extension of the narrative series of Romer and Romer (2004), 68% & 90% posterior coverage bands.

Global Financial Cycle: World Aggregates We start by analyzing the responses of global asset markets, as summarized by the global factor in risky asset prices, and the implied degree of aggregate risk aversion estimated in Section 2.[25] Second, we move on to study the responses of global domestic credit and international capital flows. Our global credit variables are world aggregates that encompass countries with different exchange rate regimes.[26] Global inflows are defined as direct cross-border credit flows provided by foreign banks to both banks and non-banks in the recipient country (see Avdjiev et al., 2012). Finally, we look at banks' leverage. Here we separate US brokers/dealers and European global banks from the aggregate banking sector, due to their different risk taking behavior. Data for credit, international inflows, and leverage are originally available at quarterly frequency (see data Appendix). We convert them to monthly frequency by interpolation.[27] Full sets of responses are collected in Figures D.1 to D.3 in the Appendix. Results are robust to starting the estimation sample in January 1990.

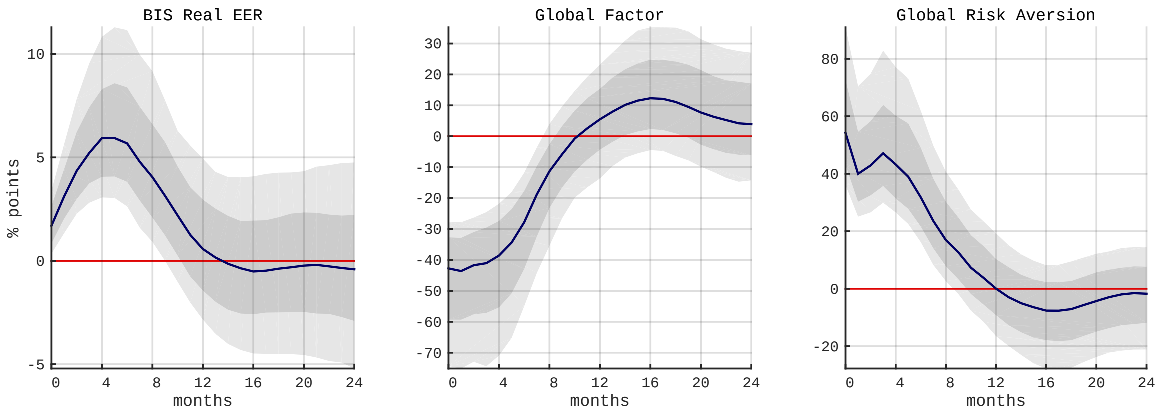
A contractionary US monetary policy shock impacts global asset markets (Figure 6). Upon realization of the monetary contraction, global risky asset prices, as summarized by the global factor, contract abruptly. While the factor has no meaningful measurement unit, we can quantify the effects on global stock markets by looking at its contribution to the overall fluctuations in the major indices. The factor explains about 20% of the common variation in our panel of international asset prices. If we assume that all asset prices loaded equally on the factor, the 40% impact fall would roughly translate into a 8% impact decrease in the local stock market. This number is consistent with both the response of the local US stock market (Figure 5), and European markets discussed at the end of the section (Figure 10). Aggregate risk aversion – i.e. the component of our factor

[25]The responses of alternative measures of risk aversion, including one that controls for discount rates and expected output (cash-flows) growth are collected in Figure E.3 in the Appendix.

[26]The countries included in our study are Argentina, Australia, Austria, Belarus, Belgium, Bolivia, Brazil, Bulgaria, Canada, Chile, Colombia, Costa Rica, Croatia, Cyprus, Czech Republic, Denmark, Ecuador, Finland, France, Germany, Greece, Hong Kong, Hungary, Iceland, Indonesia, Ireland, Italy, Japan, Latvia, Lithuania, Luxembourg, Malaysia, Malta, Mexico, Netherlands, New Zealand, Norway, Poland, Portugal, Romania, Russia, Serbia, Singapore, Slovakia, Slovenia, South Africa, South Korea, Spain, Sweden, Switzerland, Thailand, Turkey, United Kingdom and the United States.

[27]The quarterly level data are interpolated using a shape-preserving piecewise cubic interpolation; MatLab command: `y1 = interp1(t0,y0,t1,'pchip');`. Original quarterly and interpolated monthly data used in the paper are made available in the Supporting Material accompanying the paper. Results computed using alternative monthly variables (private sector liquidity instead of IMF-IFS domestic credit, and cross-border flows instead of BIS inflows, both distributed by CrossBorder Capital Ltd.) are equivalent to those discussed below and available upon request.

Figure 6: Responses of Global Asset Prices & Risk Aversion



Note: Responses to a US contractionary monetary policy shock that induces a 1% increase in the policy interest rate. Median IRFs with posterior coverage bands at 68% and 90% levels. The shock is identified using a high-frequency external instrument.

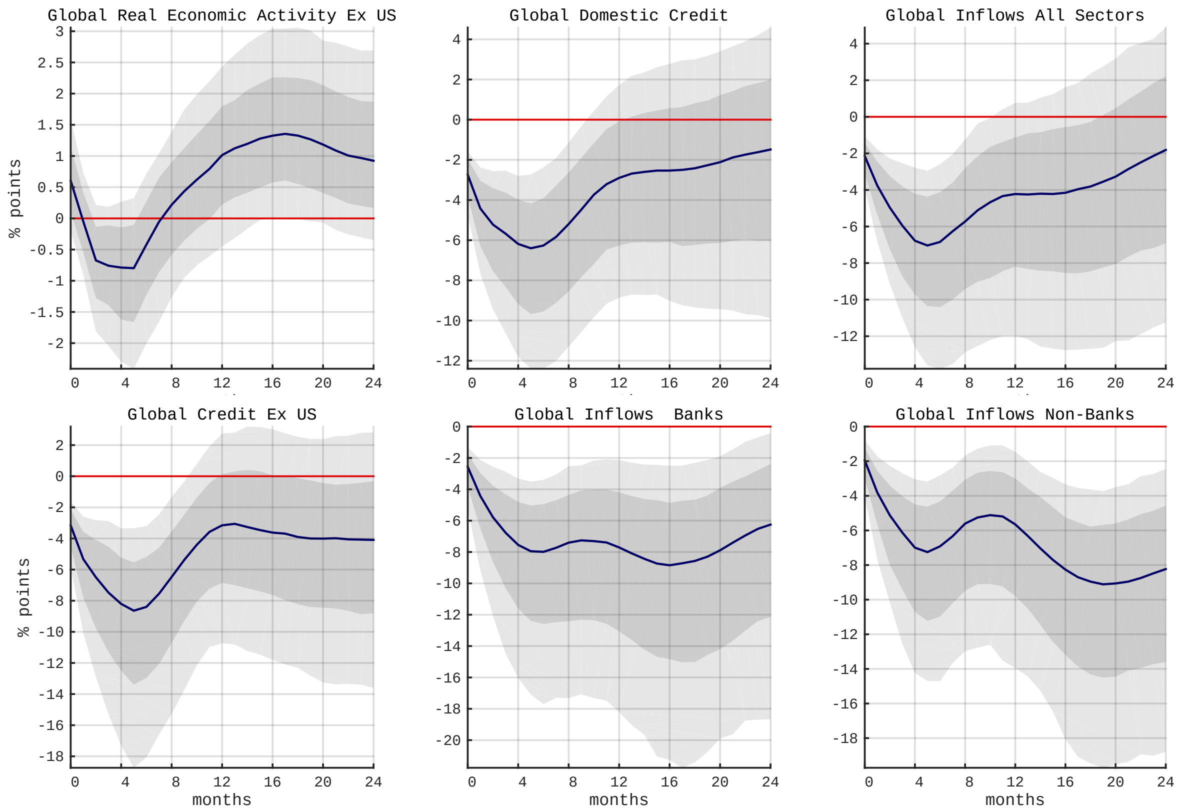
that is orthogonal to global realized variance – rises sharply.[28] The rise is consistent with the heightened levels of domestic measures of risk premia. Importantly, altering the degree of risk aversion of international investors constitutes a powerful channel for the global transmission of US monetary policy. We explore this point further when we discuss the response of global banks' leverage below. Quantifying the rise in risk aversion is less straightforward; but the shock substantially raises it by over 50% above its average trend.

Figure 7 collects the responses of global economic activity, global domestic credit, and global credit inflows. The Figure combines together responses extracted from the VARs (2) and (3) in Table 2. The US monetary policy contraction leaves global growth unchanged on impact. The inclusion of global growth here serves two purposes. First, it allows us to consider changes in global financial conditions once we have controlled for economic activity on a global scale. Second, it helps ensure that we are not confounding the effects of a US monetary policy shock with other global shocks that affect credit through their effects on growth.[29] Following a US monetary policy contraction we

[28]This result is robust to using alternative measures of risk aversion, see Appendix E.

[29]We compute global growth excluding US (Global Real Economic Activity Ex US) by using the component of the world production index of Baumeister and Hamilton (2019) that is orthogonal to the US cycle, calculated as the cyclical component of US IP. Baumeister and Hamilton (2019) use a measure of global real activity that is constructed as a weighted average of the IP indices of the OECD countries

Figure 7: Responses of Global Credit & Capital Flows

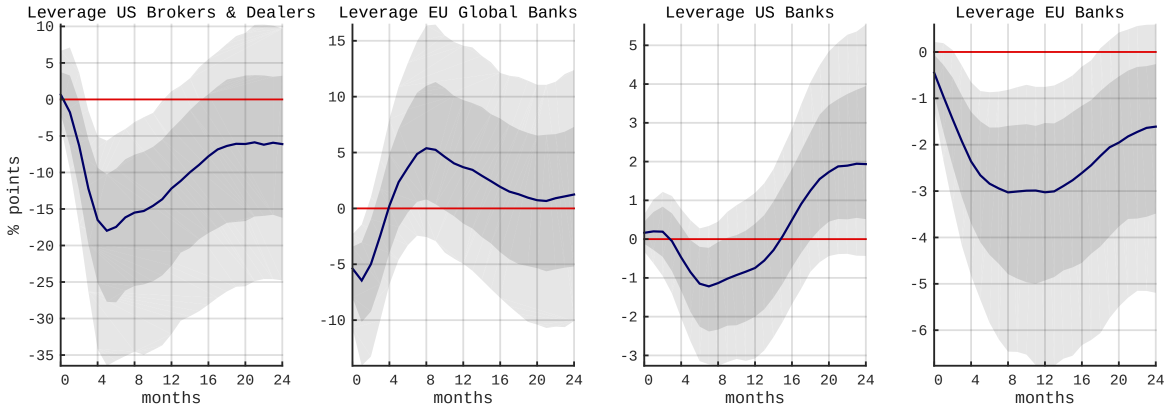


Note: Responses to a US contractionary monetary policy shock that induces a 1% increase in the policy interest rate. Median IRFs with posterior coverage bands at 68% and 90% levels. The shock is identified using a high-frequency external instrument.

register a sharp decrease in credit provision and a strong retrenchment of global capital inflows. The contraction in global domestic credit is not driven by US domestic credit, as shown in the lower left panel of the figure. Global capital inflows respond in a similar fashion: following an initial contraction, international funding flows continue to decrease to rebound at larger horizons. In the lower section of the figure we report the responses of capital inflows split by recipient type. The overall picture is consistent with a reduction of flows directed to both banking and private sectors. The decline in credit, both domestic and cross-border, whether we look at flows to banks or to non-banks, is in the order of several percentage points and thus economically significant.

+ 6 Emerging Market Economies, as an extension of a series originally maintained and distributed by the OECD. Additional details and responses of alternative measures of global growth/global real activity are reported in Figure E.1 in the Appendix.

Figure 8: Responses of Leverage of Global Banks



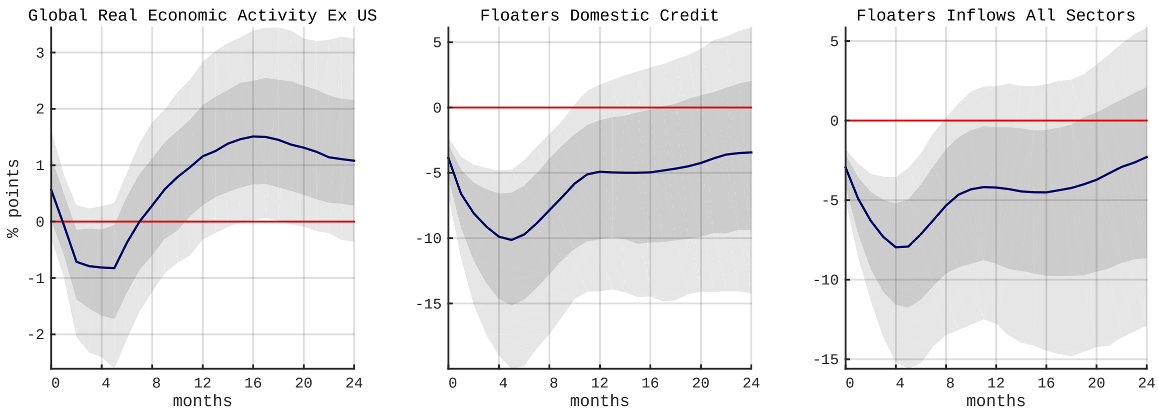
Note: Responses to a US contractionary monetary policy shock that induces a 1% increase in the policy interest rate. Median IRFs with posterior coverage bands at 68% and 90% levels. The shock is identified using a high-frequency external instrument.

Lastly, we collect the responses of banks' leverage in Figure 8. We use data on the leverage of US Security Brokers and Dealers (USBD) and Globally Systemically Important Banks (GSIBs) operating in the Euro Area and the UK. Data on total financial assets and liabilities for USBD are from the Flow of Funds of the Federal Reserve Board, while the aggregate leverage ratios for global banks in the EA and the UK are constructed using bank-level balance sheet data (details are reported in Appendix A).[30]

Consistent with declining asset prices that alter the value of banks' balance sheets, the financial leverage of global investors contracts, both among US Brokers & Dealers, and European global banks. Again, the responses are in the order of several percentage points, and hence economically relevant. The responses appear to be more delayed and more muted for the total balance sheet of the banking sector. Domestically oriented retail banks take longer to adjust, so that broader banking aggregates only react with a delay to monetary policy shocks, which instead affects more immediately the large banks with important capital market operations. The effect of US monetary policy on the whole banking sector is also less precisely estimated, and there is some variation across specifications.

[30] Adrian and Shin (2010) present evidence on the procyclicality of leverage in the domestic US context. In Section 4.2 we extend these results to an international sample of banks.

Figure 9: Responses of Global Credit & Capital Flows: Floaters



Note: Responses to a US contractionary monetary policy shock that induces a 1% increase in the policy interest rate. Median IRFs with posterior coverage bands at 68% and 90% levels. The shock is identified using a high-frequency external instrument.

Taken together, the responses collected in Figures 6 to 8 provide evidence of a powerful channel of international transmission of US monetary policy that operates mainly through global financial actors, and besides the more standard channels related to international trade. By being able to generate comovements in asset prices, credit creation and credit flows, risk appetite and financial leverage of global investors, US monetary policy can influence fluctuations in the Global Financial Cycle. This is likely the joint outcome of the dollar being the dominant currency in international financial transactions, and of the interconnectedness of global financial intermediaries.

GFC: Floaters An important question regarding Figure 7 is whether the global contraction in credit is in fact driven by countries that have a fixed or pegged exchange rate regime vis-à-vis the US dollar. In order to address this concern we restrict our sample to include only 'independently floating' countries, which we identify using use the IMF's de-facto classification.[31] We construct aggregates as the cross-sectional sum of the levels of domestic credit and capital inflows, using the same definitions as before. Full sets of

[31]Independently floating countries in our sample are Australia, Austria, Belgium, Brazil, Canada, Chile, Cyprus, Czech Republic, Finland, France, Germany, Greece, Hungary, Iceland, Ireland, Italy, Japan, Luxembourg, Malta, Mexico, Netherlands, New Zealand, Norway, Poland, Portugal, Slovenia, South Africa, Spain, Sweden, Turkey, and the United Kingdom. Source `https://www.imf.org/external/np/mfd/er/2008/eng/0408.htm`.

responses are collected in Figures D.4 and D.5 in the Appendix.

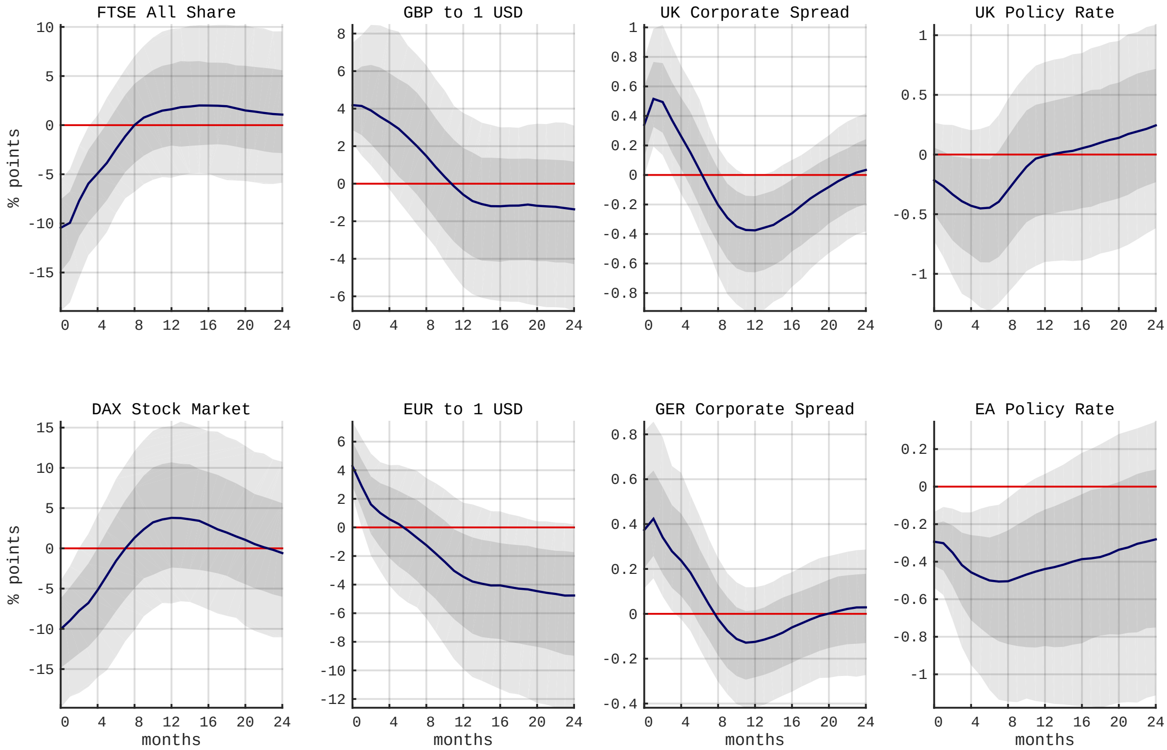
Figure 9 shows the responses of credit and capital inflows for the subset of floaters. The IRFs are obtained by replacing world aggregates with these newly constructed series in the same VAR as before (see Table 2). Again in this case we control for global economic activity. The IRFs in Figure 9 show that countries that adopt a floating exchange rate regime seem to be equally exposed to US monetary policy shocks. In fact, the magnitude of the contraction in the credit variables is very similar to that obtained over the full sample. It should be clear that these results do not imply that exchange rate regimes are equivalent. However, they do indicate that a floating exchange rate regime is not successful in providing a protective shield against US monetary policy shocks, and that fluctuations in the Global Financial Cycle can affect in a significant way all countries. We explore this point further in the next paragraph.

GFC: Currencies, Credit, and Monetary Independence We finally turn to evaluating more in detail how financial conditions transmit across borders by restricting our attention to the case of the UK and Euro Area, two important currency areas with flexible exchange rates. Full set of responses are in Figures D.6 and D.7 in the Appendix.

Figure 10 collects the responses of the local stock market indices, bilateral exchange rates vis-à-vis the dollar, corporate bond spreads, and policy interest rates for the UK (top row of the figure) and the Euro Area (bottom row of the figure).[32] We note that for all these variables the responses across the two countries are remarkably similar. Consistent with the fall in the global factor in risky asset prices, the local stock market indices plummet on impact to a very similar degree. The dollar appreciates significantly against both currencies. The exchange rate is in both cases measured as units of the foreign currency per one US dollar, such that a positive reading corresponds to an appreciation of the dollar. The appreciation is relatively short-lived in both cases, and reverts in the span of one to three quarters after the shock hits. The US monetary policy shock alters funding costs in both the UK and Euro Area, with corporate bond spreads rising very significantly and on impact in both cases. Finally, the responses of the policy rates

[32] For periods preceding the introduction of the Euro, we use the German Mark as the relevant European benchmark currency and convert it using the fixed exchange rate with the Euro chosen at the time of introduction of the common currency.

FIGURE 10: CROSS-BORDER FINANCIAL CONDITIONS & MONETARY POLICY



Note: Responses to a US contractionary monetary policy shock that induces a 1% increase in the policy interest rate. Median IRFs with posterior coverage bands at 68% and 90% levels. The shock is identified using a high-frequency external instrument.

suggest that a US contractionary monetary policy shock is likely to be followed by an endogenous easing in both the UK and the Euro Area, potentially as a response to the deterioration of the local financial conditions. While estimated with a higher degree of uncertainty in the case of the Euro area, the magnitude of the responses is very similar in the two cases, and implies an endogenous monetary easing of about 30bps. This also implies that the tightening of financial conditions in the UK and the Euro Area cannot be ascribed to a domestic monetary policy tightening, and is instead a consequence of the US monetary policy spillover.

4 Interpretation of the Results

4.1 A Simple Model with Heterogeneous Investors

The empirical results show that US monetary policy affects global banks' leverage, risky asset prices and global risk aversion. In this section, we present a stylized framework to help with the interpretation of our empirical findings; the model builds directly on the work of Zigrand et al. (2010).[33] Our illustrative model of international asset pricing features investors with heterogeneous propensities to take risk, in order to make sense of a time-varying degree of aggregate effective risk aversion.[34] The risk premium depends on the wealth distribution between leveraged global banks on the one hand, and asset managers, such as insurance companies or sovereign wealth funds, on the other hand. As the relative wealth of the two types of investors fluctuates, asset pricing will be determined mostly by one type of investors or the other.

We consider a world with two types of investors: global banks and asset managers. Global banks and asset managers account for a large part of cross-border flows, as shown in Figure 4. Global banks are leveraged entities that fund themselves in dollars for their operations in capital markets. They can borrow at the US risk-free rate and lever to buy a portfolio of world risky securities, whose returns are in dollars. They are risk-neutral investors and subject to a Value-at-Risk (VaR) constraint, which is imposed by regulation.[35] We present microeconomic evidence pertaining to the leverage and risk taking behaviour of banks in Section 4.2. The second type of investors are asset managers who, like global banks, acquire risky securities in world markets and can borrow at the US risk-free rate. Asset managers also hold a portfolio of regional assets (for example regional real estate) which is not traded in financial markets, perhaps because of information asymmetries. Asset managers are standard mean-variance investors and exhibit

[33] See also Etula (2013) and Adrian and Shin (2014).

[34] For a more realistic dynamic stochastic general equilibrium model of asset pricing with heterogeneous investors and monetary policy see Coimbra and Rey (2017). Other types of models which generate time-varying risk aversion are, for example, models with habits in consumption (see Campbell and Cochrane (1999)).

[35] Their risk neutrality is an assumption which may be justified by the fact that they benefit from an implicit bailout guarantee, either because they are universal banks, and are therefore part of a deposit guarantee scheme, or because they are too systemic to fail. Whatever the microfoundations, the crisis has provided ample evidence that global banks have taken on large amounts of risk and that this risk was not priced by creditors.

a positive degree of risk aversion that limits their desire to leverage.[36]

Global Banks

Global banks maximize the expected return of their portfolio of world risky assets subject to a Value-at-Risk (VaR) constraint.[37] The VaR imposes an upper limit on the amount a bank is predicted to lose on a portfolio with a certain probability. We denote by $\mathbf{R}_t$ the vector of excess returns of all traded risky assets in the world (in dollars). We denote by $\mathbf{x}_t^B$ the portfolio shares of a global bank, and by w_t^B the equity of the bank. The maximization problem of a global bank is

$$\max_{\mathbf{x}_t^B} \mathbb{E}_t \left(\mathbf{x}_t^{B\prime} \mathbf{R}_{t+1} \right)$$
$$\text{subject to } \mathrm{VaR}_t \leq w_t^B,$$

where VaR_t is defined as a multiple α of the standard deviation of the bank portfolio $\mathrm{VaR}_t = \alpha w_t^B \left[\mathbb{V}ar_t \left(\mathbf{x}_t^{B\prime} \mathbf{R}_{t+1} \right) \right]^{\frac{1}{2}}$.

Taking the first order condition, and using the fact that the constraint is binding (since banks are risk neutral) gives the following solution for the vector of asset demands:

$$\mathbf{x}_t^B = \frac{1}{\alpha \lambda_t} \left[\mathbb{V}ar_t(\mathbf{R}_{t+1}) \right]^{-1} \mathbb{E}_t(\mathbf{R}_{t+1}), \tag{1}$$

where we use $\mathbb{V}ar$ to denote the variance. This is formally similar to the portfolio allocation of a mean-variance investor. In Eq. (1), λ_t is the Lagrange multiplier: the VaR constraint plays the same role as risk aversion.[38]

Asset Managers

Asset managers are standard mean-variance investors with a constant degree of risk aver-

[36] The fact that only asset managers, and not the global banks, have a regional portfolio is non essential; global banks could be allowed to hold a portfolio of regional loans or assets as well. The asymmetry in risk aversion (risk neutral banks with VaR constraint and risk averse asset managers), however, is important for the results.

[37] VaR constraints have been used internally for the risk management of large banks for a long time and have entered the regulatory sphere with Basel II and III. For a microfoundation of VaR constraints, see Adrian and Shin (2014).

[38] It is possible to solve out for the Lagrange multiplier using the binding VaR constraint (see Zigrand et al., 2010). We find $\lambda_t = \left[\mathbb{E}_t(\mathbf{R}_{t+1})' \left[\mathbb{V}ar_t(\mathbf{R}_{t+1}) \right]^{-1} \mathbb{E}_t(\mathbf{R}_{t+1}) \right]^{-1/2}$.

sion equal to σ. They have access to the same set of traded assets as global banks. We call $\mathbf{x}_t^I$ the vector of portfolio weights of the asset managers in tradable risky assets. Asset managers also invest in local (regional) non-traded assets. We denote by $\mathbf{y}_t^I$ the fraction of their wealth invested in those regional assets (their net supply is $\mathbf{y}_t$). The vector of excess returns on these non tradable investments is $\mathbf{R}_t^N$. Finally, we call w_t^I the equity of asset managers. An asset manager chooses his portfolio of risky assets by maximizing

$$\max_{x_t^I} \mathbb{E}_t \left(\mathbf{x}_t^{I\prime} \mathbf{R}_{t+1} + \mathbf{y}_t^{I\prime} \mathbf{R}_{t+1}^N \right) - \frac{\sigma}{2} \mathbb{V}ar_t \left(\mathbf{x}_t^{I\prime} \mathbf{R}_{t+1} + \mathbf{y}_t^{I\prime} \mathbf{R}_{t+1}^N \right).$$

The optimal portfolio choice in risky tradable securities for an asset manager will be

$$\mathbf{x}_t^I = \frac{1}{\sigma} \left[\mathbb{V}ar_t(\mathbf{R}_{t+1}) \right]^{-1} \left[\mathbb{E}_t(\mathbf{R}_{t+1}) - \sigma \mathbb{C}ov_t(\mathbf{R}_{t+1}, \mathbf{R}_{t+1}^N) \mathbf{y}_t^I \right]. \tag{2}$$

Market clearing conditions

The market clearing condition for risky traded securities is $\mathbf{x}_t^B \frac{w_t^B}{w_t^B + w_t^I} + \mathbf{x}_t^I \frac{w_t^I}{w_t^B + w_t^I} = \mathbf{s}_t$ where $\mathbf{s}_t$ is a world vector of net asset supplies for traded assets.

Proposition 1 (Risky Asset Returns) *Using Eq. (1) and (2) and the market clearing conditions, the expected excess returns on tradable risky assets can be rewritten as the sum of a global component and a regional component:*

$$\mathbb{E}_t \left(\mathbf{R}_{t+1} \right) = \Gamma_t \mathbb{V}ar_t(\mathbf{R}_{t+1}) \; \mathbf{s}_t + \Gamma_t \mathbb{C}ov_t(\mathbf{R}_{t+1}, \mathbf{R}_{t+1}^N) \mathbf{y}_t, \tag{3}$$

where $\Gamma_t \equiv \left[\frac{w_t^B}{\alpha \lambda_t} + \frac{w_t^I}{\sigma} \right]^{-1} \left(w_t^B + w_t^I \right)$. *The global component of risky asset prices is equal to the aggregate variance scaled by the aggregate degree of effective risk aversion* Γ_t.

Γ_t is the wealth-weighted average of the 'risk aversions' of the asset managers and the global banks. It can be interpreted as the aggregate degree of effective risk aversion of the market. If all the wealth were in the hands of asset managers, for example, aggregate risk aversion would be equal to σ. Using Eq. (3) as a guiding framework, in Section 2 we extracted the global factor in world risky asset prices by writing each price series as the sum of a global, a regional and an asset specific component. We then used Eq. (3) to extract our empirical proxy for aggregate risk aversion Γ_t.[39] One possible interpretation of the decline in the measure of aggregate risk aversion observed between 2003 and 2007 in Figure 3 is therefore that it was driven by risk-neutral global banks becoming large and important for the pricing of risky assets, sustaining an increase in risky asset prices on a global scale. This trend reversed after the crisis, when instead more risk-averse asset managers became relatively bigger (see Figure 4).

Proposition 2 (Global Banks Returns) *The expected excess return of a global bank portfolio in our economy is given by*

$$\begin{aligned}\mathbb{E}_t(\mathbf{x}_t^{B\prime}\mathbf{R}_{t+1}) &= \Gamma_t \mathbb{C}ov_t(\mathbf{x}_t^{B\prime}\mathbf{R}_{t+1}, \mathbf{s}_t^{\prime}\mathbf{R}_{t+1}) + \Gamma_t \mathbb{C}ov_t(\mathbf{x}_t^{B\prime}\mathbf{R}_{t+1}, \mathbf{y}_t^{\prime}\mathbf{R}_{t+1}^N) \\ &= \beta_t^{BW}\Gamma_t \mathbb{V}ar_t(\mathbf{s}_t^{\prime}\mathbf{R}_{t+1}) + \Gamma_t \mathbb{C}ov_t(\mathbf{x}_t^{B\prime}\mathbf{R}_{t+1}, \mathbf{y}_t^{\prime}\mathbf{R}_{t+1}^N),\end{aligned} \tag{4}$$

where β_t^{BW} is the beta of a global bank with the world market portfolio.

The higher the correlation of a global bank portfolio with the world portfolio (i.e. high-β_t^{BW}), the more the bank loads on world risk, the higher the expected asset return, *ceteris paribus*.

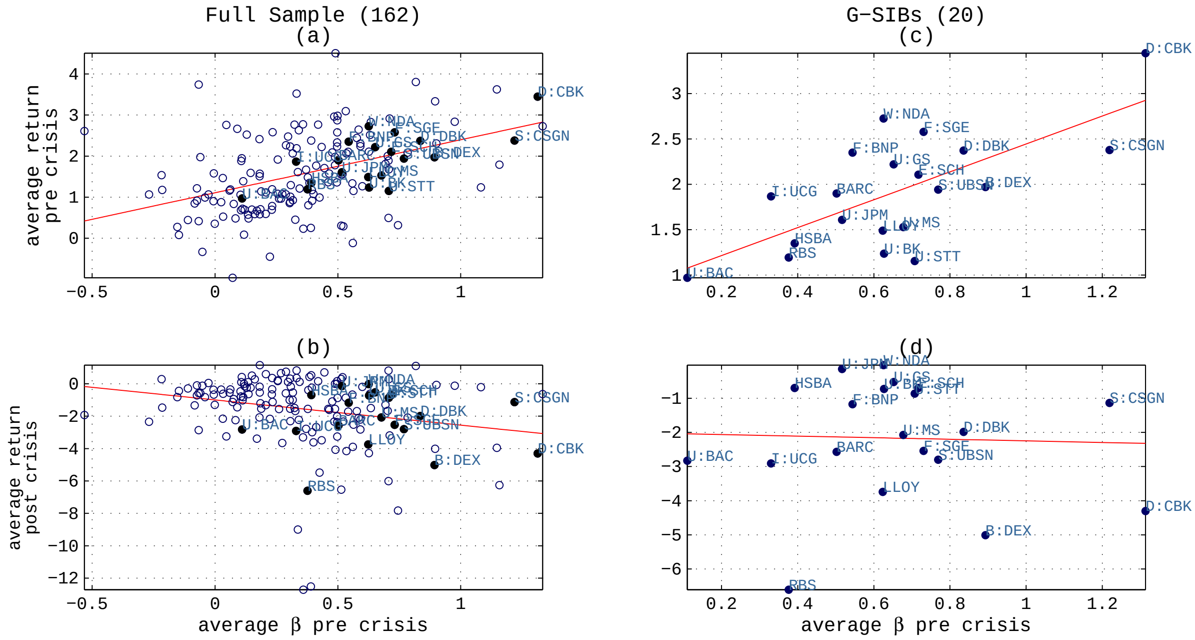
4.2 Evidence on Global Banks

In this section we use balance sheet data to provide some evidence on the risk taking behavior of banks, in line with our simple model. Adrian and Shin (2010) show that the leverage of US brokers-dealers is procyclical. Using balance sheet data for a large sample of international financial institutions (see Table A.4), we find that the positive association between leverage growth and balance sheet growth goes well beyond US borders. We report these results in Figure A.3 in the Appendix.[40] The procyclicality of leverage tends to be a stronger feature of the behavior of financial institutions that engage in global

[39] As mentioned earlier, in general our empirical proxy for aggregate risk aversion could also reflect expected dividend growth and discount rates. We control for these additional factors, estimate a more general aggregate degree of risk aversion and check that all our results go through in Appendix E.

[40] We calculate leverage along the lines of Kalemli-Ozcan et al. (2012). We use a panel 166 financial institutions in 20 countries from 2000 to 2010. We identify a subset of 21 large banks who have been classified as Globally Systemically Important Banks (GSIBs). A complete list of institutions included in our set is in Table A.4.

Figure 11: Correlation between banks' returns and loading on the global factor



Note: In each subplot, the x axis reports the average β^{BW} in the three years preceding the onset of the financial crisis (August 2007), while the y axis records average returns in percentage points. Filled blue circles highlight GSIBs within the broader population of banks (hollow circles); the sign of the correlation is visualized by a red regression line in each plot. Panels (a) and (b): banks average returns pre (2003-2007) and post (2007-2010) crisis as a function of their pre-crisis betas. Panels (c) and (d) GSIBs subsample. *Source:* Datastream, authors calculations.

capital markets operations, a subset which included in particular the former stand-alone investment banks. The same holds true for the large European (UK, Euro Area and Switzerland) universal banks, whose investment departments played a central role in channelling US dollar liquidity worldwide in the years immediately preceding the financial crisis (see Shin, 2012). Many of those large European Banks are GSIBs.

Figure 11 is the empirical counterpart of Eq. (4), and reports the correlation between the returns of each bank and their loading (β_t^{BW}) on our global factor of Section 2. Results in panels (a) and (b) are calculated over the entire population of banks, while panels (c) and (d) refer to the GSIBs subsample, and we use August 2007 to distinguish between pre and post crisis periods. Results confirm a positive association between high β_t^{BW} and high returns in the pre crisis sample. Panels (a) and (c) show that, relative to the larger population, GSIBs tend to have both higher average betas, and larger returns. This suggests that global banks were systematically loading more on world risk in the run-up to the financial crisis, and that their behaviour was delivering larger average returns, compared to the average bank in our sample. The higher loadings on risk are consistent with the build-up of leverage in the years prior to the crisis documented in Figure A.2. Panels (b) and (d) sort the banks on the x-axis according to their pre-crisis betas, but report their post crisis returns on the y axis: institutions that were loading more on global risk pre crisis suffered the largest losses after the systemic meltdown began.

5 Conclusions

This paper establishes the importance of US monetary policy as one of the drivers of the Global Financial Cycle. First, we show that a single global factor explains an important share of the common variation of a large cross section of risky asset prices around the world. Using a simple model of international asset pricing with heterogeneous intermediaries, we interpret this global factor as reflecting market volatility and aggregate risk aversion in global markets. Second, we show that US monetary policy shocks induce strong comovements in the international financial variables that characterize the Global Financial Cycle. Monetary contractions are followed by a significant deleveraging of global financial intermediaries, a rise in aggregate risk aversion, a contraction in the global factor in asset prices and a decline in global credit, a widening of corporate bond spreads and retrenchments of gross capital flows. These results also hold for the countries of our sample with floating exchange rates. This is an important result, as it challenges the degree of monetary policy sovereignty of open economies, and echoes the claim of Rey (2013) that the Mundellian trilemma may have morphed into a dilemma: as long as capital flows across borders are free, and macroprudential tools are not used, monetary conditions in any country, even one with a flexible exchange rate, are partly dictated by the monetary policy of the hegemon (the US). This of course does not mean that exchange rate regimes do not matter, as Klein and Shambaugh (2013) and Obstfeld (2015) rightly point out.[41] This international transmission mechanism of monetary policy is a priori

[41] For interesting models of the challenges of the trilemma in standard neo-Keynesian models, see Farhi and Werning, 2012, 2013.

consistent with models where financial market imperfections play an important role, e.g. via Value-at-Risk constraints, and where heterogeneous financial intermediaries price assets. It still remains to be seen whether open economy extensions of these models would be able to generate a Global Financial Cycle whose features would match the empirical regularities uncovered in this paper.[42]

[42] For a more detailed discussion of the theoretical challenges when modelling international monetary policy transmission channels, see Bernanke (2017) Rey (2016) and Coimbra and Rey (2017) .

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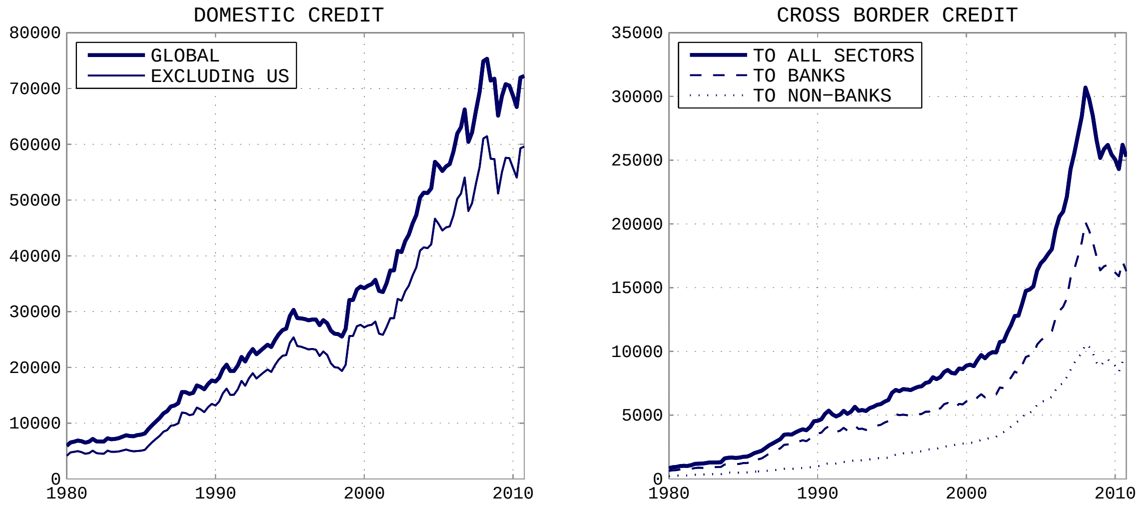
A Credit and Banking Data – For Online Publication

A.1 Domestic and Cross-Border Credit

Credit data, both domestic and cross-border, are constructed using data collected and distributed by the IMF's International Financial Statistics (IFS) and the Bank for International Settlements (BIS) databases respectively, for the countries listed in table A.1.

Figure A.1: Global Credit



Note: Global Domestic Credit and Global Cross-Border Inflows constructed as the cross sectional sum of country-specific credit variables. The unit in both plots is Billion USD.

Following Gourinchas and Obstfeld (2012) we construct National Domestic Credit for each country as the difference between Domestic Claims to All Sectors and Net Claims to Central Government reported by each country's financial institutions; however, we only consider claims of depository corporations excluding central banks. Specifically, we refer to the Other Depository Corporation Survey available within the IFS database and construct Claims to All Sectors as the sum of Claims On Private Sector, Claims on Public Non Financial Corporations, Claims on Other Financial Corporations and Claims on State And Local Government; while Net Claims to Central Government are calculated as the difference between Claims on and Liabilities to Central Government. This classification was adopted starting from 2001, prior to that date we refer to the Deposit Money Banks Survey. Raw data are quarterly and expressed in national currency, we convert them in Billion USD equivalents using end of period exchange rates again available within the IFS. Whenever there exists a discontinuity between data available under the old and new classifications we interpolate the missing observations. Global

Table A.1: List of Countries Included

North America	Latin America	Central and Eastern Europe	Western Europe	Emerging Asia	Asia Pacific	Africa and Middle East
Canada	Argentina	Belarus	Austria	China	Australia	Israel
US	Bolivia	Bulgaria	Belgium	Indonesia	Japan	South Africa
	Brazil	Croatia	Cyprus	Malaysia	Korea	
	Chile	Czech Republic	Denmark	Singapore	New Zealand	
	Colombia	Hungary	Finland	Thailand		
	Costa Rica	Latvia	France			
	Ecuador	Lithuania	Germany			
	Mexico	Poland	Greece*			
		Romania	Iceland			
		Russian Federation	Ireland			
		Slovak Republic	Italy			
		Slovenia	Luxembourg			
		Turkey	Malta			
			Netherlands			
			Norway			
			Portugal			
			Spain			
			Sweden			
			Switzerland			
			UK			

Note: Countries included in the construction of the Domestic Credit and Cross-Border Credit variables used throughout the paper. Greece is not included in the computation of Global Domestic Credit due to poor quality of original national data.

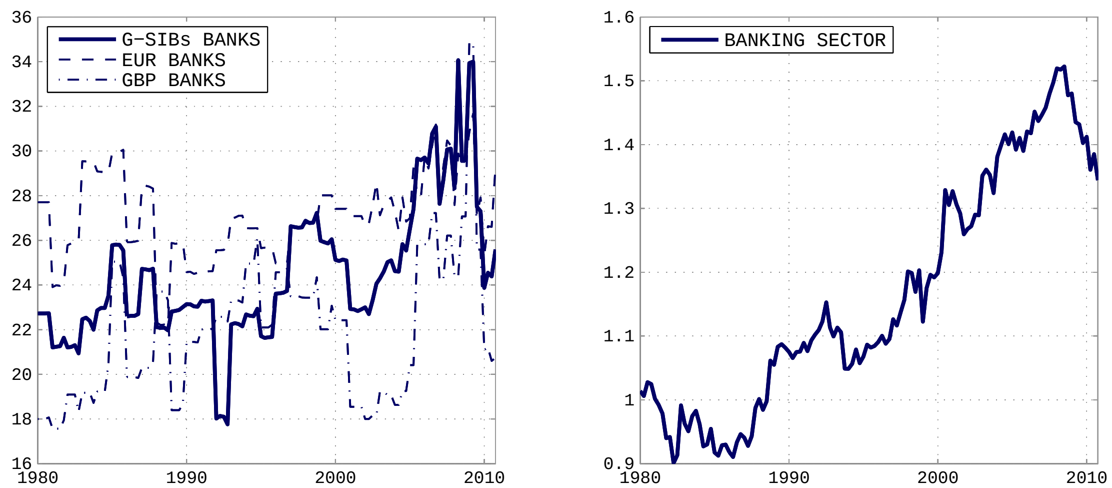
Domestic Credit is finally constructed as the cross-sectional sum of the National Domestic Credit variables.

To construct the Cross-Border Capital Inflows measures used within the paper we adopt the definition of Direct Cross-Border Credit in Avdjiev et al. (2012). We use original data available at the BIS Locational Banking Statistics Database and collected under External Positions of Reporting Banks vis-à-vis Individual Countries (Table 6). Data refer to the outstanding amount of Claims to All Sectors and Claims to Non-Bank Sector in all currencies, all instruments, declared by all BIS reporting countries with counterparty location being the individual countries in Table A.1. We then construct Claims to the Banking Sector as the difference between the two categories available. Original data are available at quarterly frequency in Million USD. Global Inflows are finally calculated as the cross-sectional sum of the national variables. Global domestic credit and global cross-border capital inflows are plotted in Figure A.1.

A.2 Banking Sector and Individual Banks Leverage data

To construct an aggregate country-level measure of banking sector leverage we follow Forbes (2012) and build it as the ratio between Claims on Private Sector and Transferable plus Other Deposits included in Broad Money of depository corporations excluding central banks. Original data are in national currencies and are taken from the Other Depository Corporations Survey; Monetary Statistics, International Financial Statistics database. The classification of deposits within the former Deposit Money Banks Survey corresponds to Demand, Time, Savings and Foreign Currency Deposits. Using these national data as a reference, we construct the European Banking Sector Leverage variable as the median leverage ratio among Austria, Belgium, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Luxembourg, Netherlands, Portugal, Spain and United Kingdom.

FIGURE A.2: EUROPEAN BANKS LEVERAGE



Note: [LEFT PANEL] Leverage ratio calculated for the European GSIBs with a detail on EUR and GBP banks using the institutions and classification in Table A.2. [RIGHT PANEL] Aggregated European banking sector leverage ratio measured as the median of European countries banking sector leverage variables following Forbes (2012).

The aggregate Leverage Ratios (defined as Total Assets over Equity) for the Global Systemic Important Banks in the Euro-Area and United-Kingdom used in the BVAR are constructed as weighted averages of individual banks data. Balance sheet Total Assets (DWTA) and Shareholders' Equity (DWSE) are from the Thomson Reuter Worldscope

Table A.2: European G-SIBs.

NAME	ISIN	GICS INDUSTRY	COUNTRY	EA LEV	UK LEV
BNP Paribas	FR0000131104	Commercial Banks	France	•	
Credit Agricole	FR0000045072	Commercial Banks	France	•	
Societe Generale	FR0000130809	Commercial Banks	France	•	
Commerzbank	DE0008032004	Commercial Banks	Germany	•	
Deutsche Bank	DE0005140008	Capital Markets	Germany	•	
Unicredit	IT0004781412	Commercial Banks	Italy	•	
ING Bank	NL0000113892	Commercial Banks	Netherlands	•	
BBVA	ES0113211835	Commercial Banks	Spain	•	
Banco Santander	ES0113900J37	Commercial Banks	Spain	•	
Nordea Group	SE0000427361	Commercial Banks	Sweden		
Credit Suisse Group	CH0012138530	Capital Markets	Switzerland		
UBS	CH0024899483	Capital Markets	Switzerland		
Royal Bank of Scotland	GB00B7T77214	Commercial Banks	UK		•
Barclays	GB0031348658	Commercial Banks	UK		•
HSBC Holdings	GB0005405286	Commercial Banks	UK		•
Lloyds Banking Group	GB0008706128	Commercial Banks	UK		•
Standard Chartered	GB0004082847	Diversified Fin'l	UK		•

Note: European Global Systemically Important Banks included in the construction of GSIBs Leverage Ratios; the last two columns highlight the components of EUR and GDP Leverage respectively.

Datastream database and available at quarterly frequency. Weights are proportional to Market Capitalisation (WC08001) downloaded from the same source. Details on the banks included and their characteristics are summarised in Table A.2 below. The aggregated banking sector leverage and the leverage ratio of the European GSIBs are plotted in Figure A.2.

Figures 11 and A.3 are built using data on individual banks total return indices excluding dividends taken from Thomson Reuters Worldscope database at quarterly frequency. Data are collected directly from banks balance sheets and Leverage Ratios are computed as the ratio between Total Assets (DWTA) and Common/Shareholders' Equity (DWSE). Total Assets include cash and due from banks, total investments, net loans, customer liability on acceptances (if included in total assets), investment in unconsolidated subsidiaries, real estate assets, net property, plant and equipment, and other assets. Descriptive statistics for bank level data and a complete list of the institutions included in the sample are provided in Tables A.3 and A.4 respectively. Although the data source is different, the calculation follows Kalemli-Ozcan et al. (2012).

Table A.3: Bank Data Summary Statistics.

(a)

	All (155)			GSIBs (25)			CommB (123)		
	A	E	L	A	E	L	A	E	L
min	0.3	0.0	1.113	60.9	2.7	6.353	0.4	0.0	4.887
max	3880.6	219.8	327.2	3880.6	219.8	163.5	3880.6	219.8	327.2
mean	251.7	12.9	18.73	1121.2	53.4	24.59	258.4	13.5	19.86
median	54.8	3.9	15.92	1108.3	39.1	22.76	55.0	3.6	17

(b)

	CapM (18)			T&MF (5)			Other Fin'l(9)		
	A	E	L	A	E	L	A	E	L
min	0.3	0.2	1.113	1.9	0.1	2.989	5.5	0.6	2.242
max	3595.1	76.9	136.2	61.2	5.7	19.5	310.0	42.8	65.13
mean	364.5	15.4	16.06	21.7	2.5	9.933	63.1	6.7	13.65
median	90.2	7.3	12.98	21.7	1.3	7.978	26.9	3.3	7.259

Note: Summary statistics for bank-level data used in the analysis. (A) Total Assets, (E) Shareholders' Equity, (L) Leverage Ratio. [panel (a)] full sample (All), Global Systemically Important Banks (GSIBs), Commercial Banks (CommB). [panel (b)] Capital Markets (CapM), Thrifts & Mortgage Finance (T&MF), Other Financial (Other Fin'l, includes Diversified Financial Services and Consumer Finance). Total assets and common equity are in Billion USD. Numbers in parentheses denote the number of banks in each category.

Table A.4: List of Financial Institutions included

ISIN Code	Bank Name	Geo Code	Country	GICS Industry	G-SIB
AT0000606306	RAIFFEISEN BANK INTL.	EU	Austria	Commercial Banks	
AT0000625108	OBERBANK	EU	Austria	Commercial Banks	
AT0000652011	ERSTE GROUP BANK	EU	Austria	Commercial Banks	
BE0003565737	KBC GROUP	EU	Belgium	Commercial Banks	
GB0005405286	HSBC HOLDING	EU	Great Britain	Commercial Banks	•
GB0008706128	LLOYDS BANKING GROUP	EU	Great Britain	Commercial Banks	•
GB0031348658	BARCLAYS	EU	Great Britain	Commercial Banks	•
GB00B7T77214	ROYAL BANK OF SCTL.GP.	EU	Great Britain	Commercial Banks	•
DK0010274414	DANSKE BANK	EU	Denmark	Commercial Banks	
DK0010307958	JYSKE BANK	EU	Denmark	Commercial Banks	
FR0000045072	CREDIT AGRICOLE	EU	France	Commercial Banks	•
FR0000031684	PARIS ORLEANS	EU	France	Capital Markets	
FR0000120685	NATIXIS	EU	France	Commercial Banks	
FR0000130809	SOCIETE GENERALE	EU	France	Commercial Banks	•
FR0000131104	BNP PARIBAS	EU	France	Commercial Banks	•
DE0008001009	DEUTSCHE POSTBANK	EU	Germany	Commercial Banks	
DE0005140008	DEUTSCHE BANK	EU	Germany	Capital Markets	•

continues on next page –

Table A.4 – continued from previous page

ISIN Code	Bank Name	Geo Code	Country	GICS Industry	G-SIB
DE000CBK1001	COMMERZBANK	EU	Germany	Commercial Banks	•
IE0000197834	ALLIED IRISH BANKS	EU	Ireland	Commercial Banks	
IE0030606259	BANK OF IRELAND	EU	Ireland	Commercial Banks	
IE00B59NXW72	PERMANENT TSB GHG.	EU	Ireland	Commercial Banks	
IT0005002883	BANCO POPOLARE	EU	Italy	Commercial Banks	
IT0003487029	UNIONE DI BANCHE ITALIAN	EU	Italy	Commercial Banks	
IT0000062957	MEDIOBANCA BC.FIN	EU	Italy	Capital Markets	
IT0000064482	BANCA POPOLARE DI MILANO	EU	Italy	Commercial Banks	
IT0000072618	INTESA SANPAOLO	EU	Italy	Commercial Banks	
IT0001005070	BANCO DI SARDEGNA RSP	EU	Italy	Commercial Banks	
IT0004984842	BANCA MONTE DEI PASCHI	EU	Italy	Commercial Banks	
IT0004781412	UNICREDIT	EU	Italy	Commercial Banks	•
NO0006000801	SPAREBANK 1 NORD-NORGE	EU	Norway	Commercial Banks	
NO0006000900	SPAREBANKEN VEST	EU	Norway	Commercial Banks	
PTBCP0AM0007	BANCO COMR.PORTUGUES R	EU	Portugal	Commercial Banks	
PTBES0AM0007	BANCO ESPIRITO SANTO	EU	Portugal	Commercial Banks	
PTBPI0AM0004	BANCO BPI	EU	Portugal	Commercial Banks	
ES0113860A34	BANCO DE SABADELL	EU	Spain	Commercial Banks	
ES0113211835	BBV.ARGENTARIA	EU	Spain	Commercial Banks	•
ES0113679I37	BANKINTER R	EU	Spain	Commercial Banks	
ES0113790226	BANCO POPULAR ESPANOL	EU	Spain	Commercial Banks	
ES0113900J37	BANCO SANTANDER	EU	Spain	Commercial Banks	•
SE0000148884	SEB A	EU	Sweden	Commercial Banks	
SE0000193120	SVENSKA HANDBKN.A	EU	Sweden	Commercial Banks	
SE0000242455	SWEDBANK A	EU	Sweden	Commercial Banks	
SE0000427361	NORDEA BANK	EU	Sweden	Commercial Banks	•
CH0012138530	CREDIT SUISSE GROUP N	EU	Switzerland	Capital Markets	•
CH0012335540	VONTOBEL HOLDING	EU	Switzerland	Capital Markets	
CH0018116472	BANK COOP	EU	Switzerland	Commercial Banks	
CH0024899483	UBS R	EU	Switzerland	Capital Markets	•
CA0636711016	BANK OF MONTREAL	AM	Canada	Commercial Banks	
CA0641491075	BK.OF NOVA SCOTIA	AM	Canada	Commercial Banks	
CA1360691010	CANADIAN IMP.BK.COM.	AM	Canada	Commercial Banks	
CA13677F1018	CANADIAN WESTERN BANK	AM	Canada	Commercial Banks	
CA51925D1069	LAURENTIAN BK.OF CANADA	AM	Canada	Commercial Banks	
CA6330671034	NAT.BK.OF CANADA	AM	Canada	Commercial Banks	
CA7800871021	ROYAL BANK OF CANADA	AM	Canada	Commercial Banks	
CA8911605092	TORONTO-DOMINION BANK	AM	Canada	Commercial Banks	
US0258161092	AMERICAN EXPRESS	AM	United States	Diversified Fin'l	
US0454871056	ASSOCIATED BANC-CORP	AM	United States	Commercial Banks	
US0462651045	ASTORIA FINL.	AM	United States	Thrifts & Mortgage	
US0549371070	BB&T	AM	United States	Commercial Banks	
US05561Q2012	BOK FINL.	AM	United States	Commercial Banks	
US0596921033	BANCORPSOUTH	AM	United States	Commercial Banks	
US0605051046	BANK OF AMERICA	AM	United States	Commercial Banks	•
US0625401098	BANK OF HAWAII	AM	United States	Commercial Banks	
US0640581007	BANK OF NEW YORK MELLON	AM	United States	Capital Markets	•
US14040H1059	CAPITAL ONE FINL.	AM	United States	Diversified Fin'l	
US1491501045	CATHAY GEN.BANCORP	AM	United States	Commercial Banks	
US1729674242	CITIGROUP	AM	United States	Commercial Banks	•
US1785661059	CITY NATIONAL	AM	United States	Commercial Banks	
US2003401070	COMERICA	AM	United States	Commercial Banks	
US2005251036	COMMERCE BCSH.	AM	United States	Commercial Banks	
US2298991090	CULLEN FO.BANKERS	AM	United States	Commercial Banks	

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Table A.4 – continued from previous page

ISIN Code	Bank Name	Geo Code	Country	GICS Industry	G-SIB
US2692464017	E*TRADE FINANCIAL	AM	United States	Capital Markets	
US27579R1041	EAST WEST BANCORP	AM	United States	Commercial Banks	
US3167731005	FIFTH THIRD BANCORP	AM	United States	Commercial Banks	
US31946M1036	FIRST CTZN.BCSH.A	AM	United States	Commercial Banks	
US3205171057	FIRST HORIZON NATIONAL	AM	United States	Commercial Banks	
US33582V1089	FIRST NIAGARA FINL.GP.	AM	United States	Commercial Banks	
US3379151026	FIRSTMERIT	AM	United States	Commercial Banks	
US3546131018	FRANKLIN RESOURCES	AM	United States	Capital Markets	
US3602711000	FULTON FINANCIAL	AM	United States	Commercial Banks	
US38141G1040	GOLDMAN SACHS GP.	AM	United States	Capital Markets	•
US4436831071	HUDSON CITY BANC.	AM	United States	Thrifts & Mortgage	
US4461501045	HUNTINGTON BCSH.	AM	United States	Commercial Banks	
US4508281080	IBERIABANK	AM	United States	Commercial Banks	
US4590441030	INTERNATIONAL BCSH.	AM	United States	Commercial Banks	
US46625H1005	JP MORGAN CHASE & CO.	AM	United States	Commercial Banks	•
US4932671088	KEYCORP	AM	United States	Commercial Banks	
US55261F1049	M&T BANK	AM	United States	Commercial Banks	
US55264U1088	MB FINANCIAL	AM	United States	Commercial Banks	
US6174464486	MORGAN STANLEY	AM	United States	Capital Markets	•
US6494451031	NEW YORK COMMUNITY BANC.	AM	United States	Thrifts & Mortgage	
US6658591044	NORTHERN TRUST	AM	United States	Capital Markets	
US6934751057	PNC FINL.SVS.GP.	AM	United States	Commercial Banks	
US7127041058	PEOPLES UNITED FINANCIAL	AM	United States	Thrifts & Mortgage	
US7429621037	PRIVATEBANCORP	AM	United States	Commercial Banks	
US7547301090	RAYMOND JAMES FINL.	AM	United States	Capital Markets	
US7591EP1005	REGIONS FINL.NEW	AM	United States	Commercial Banks	
US78442P1066	SLM	AM	United States	Diversified Fin'l	
US78486Q1013	SVB FINANCIAL GROUP	AM	United States	Commercial Banks	
US8085131055	CHARLES SCHWAB	AM	United States	Capital Markets	
US8574771031	STATE STREET	AM	United States	Capital Markets	•
US8679141031	SUNTRUST BANKS	AM	United States	Commercial Banks	
US8690991018	SUSQUEHANNA BCSH.	AM	United States	Commercial Banks	
US87161C5013	SYNOVUS FINANCIAL	AM	United States	Commercial Banks	
US8722751026	TCF FINANCIAL	AM	United States	Commercial Banks	
US87236Y1082	TD AMERITRADE HOLDING	AM	United States	Capital Markets	
US9027881088	UMB FINANCIAL	AM	United States	Commercial Banks	
US9029733048	US BANCORP	AM	United States	Commercial Banks	
US9042141039	UMPQUA HOLDINGS	AM	United States	Commercial Banks	
US9197941076	VALLEY NATIONAL BANCORP	AM	United States	Commercial Banks	
US9388241096	WASHINGTON FEDERAL	AM	United States	Thrifts & Mortgage	
US9478901096	WEBSTER FINANCIAL	AM	United States	Commercial Banks	
US9497461015	WELLS FARGO & CO	AM	United States	Commercial Banks	•
US97650W1080	WINTRUST FINANCIAL	AM	United States	Commercial Banks	
US9897011071	ZIONS BANCORP.	AM	United States	Commercial Banks	
JP3902900004	MITSUBISHI UFJ FINL.GP.	AS	Japan	Commercial Banks	•
JP3890350006	SUMITOMO MITSUI FINL.GP.	AS	Japan	Commercial Banks	•
JP3429200003	SHINKIN CENTRAL BANK PF.	AS	Japan	Commercial Banks	
JP3805010000	FUKUOKA FINANCIAL GP.	AS	Japan	Commercial Banks	
JP3842400008	HOKUHOKU FINL. GP.	AS	Japan	Commercial Banks	
JP3105040004	AIFUL	AS	Japan	Diversified Fin'l	
JP3107600003	AKITA BANK	AS	Japan	Commercial Banks	
JP3108600002	ACOM	AS	Japan	Diversified Fin'l	
JP3152400002	BANK OF IWATE	AS	Japan	Commercial Banks	
JP3175200009	OITA BANK	AS	Japan	Commercial Banks	

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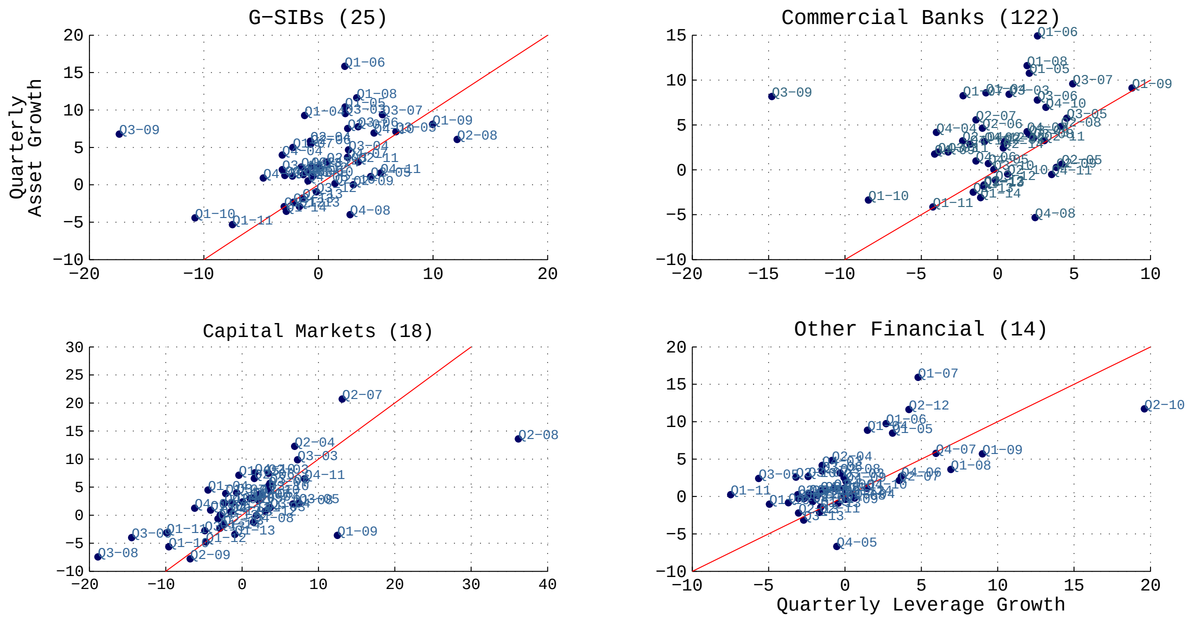
Table A.4 – continued from previous page

ISIN Code	Bank Name	Geo Code	Country	GICS Industry	G-SIB
JP3194600007	BANK OF OKINAWA	AS	Japan	Commercial Banks	
JP3200450009	ORIX	AS	Japan	Diversified Fin'l	
JP3207800008	KAGOSHIMA BANK	AS	Japan	Commercial Banks	
JP3271400008	CREDIT SAISON	AS	Japan	Diversified Fin'l	
JP3276400003	GUNMA BANK	AS	Japan	Commercial Banks	
JP3351200005	SHIZUOKA BANK	AS	Japan	Commercial Banks	
JP3352000008	77 BANK	AS	Japan	Commercial Banks	
JP3388600003	JACCS	AS	Japan	Diversified Fin'l	
JP3392200006	EIGHTEENTH BANK	AS	Japan	Commercial Banks	
JP3392600007	JUROKU BANK	AS	Japan	Commercial Banks	
JP3394200004	JOYO BANK	AS	Japan	Commercial Banks	
JP3441600008	TAIKO BANK	AS	Japan	Commercial Banks	
JP3502200003	DAIWA SECURITIES GROUP	AS	Japan	Capital Markets	
JP3511800009	CHIBA BANK	AS	Japan	Commercial Banks	
JP3520000005	CHUKYO BANK	AS	Japan	Commercial Banks	
JP3521000004	CHUGOKU BANK	AS	Japan	Commercial Banks	
JP3587000005	TOKYO TOMIN BANK	AS	Japan	Commercial Banks	
JP3601000007	TOHO BANK	AS	Japan	Commercial Banks	
JP3630500001	TOMATO BANK	AS	Japan	Commercial Banks	
JP3653400006	NANTO BANK	AS	Japan	Commercial Banks	
JP3762600009	NOMURA HDG.	AS	Japan	Capital Markets	
JP3769000005	HACHIJUNI BANK	AS	Japan	Commercial Banks	
JP3783800000	HIGO BANK	AS	Japan	Commercial Banks	
JP3786600001	HITACHI CAPITAL	AS	Japan	Diversified Fin'l	
JP3841000007	HOKUETSU BANK	AS	Japan	Commercial Banks	
JP3881200004	MIE BANK	AS	Japan	Commercial Banks	
JP3888000001	MICHINOKU BANK	AS	Japan	Commercial Banks	
JP3905850008	MINATO BANK	AS	Japan	Commercial Banks	
JP3942000005	YAMANASHI CHUO BK.	AS	Japan	Commercial Banks	
JP3955400001	BANK OF YOKOHAMA	AS	Japan	Commercial Banks	

Notes: In the first column are the ISIN identification codes followed by the institution's name, geographical location and country of reference. The last column highlights the subset of institutions which have been classified as Global Systemically Important Banks (G-SIBs) previously known as G-SIFIs (Systemically Important Financial Institutions); the classification has been adopted by the Financial Stability Board starting from November 2011 and lastly updated in November 2013.

Figure A.3: Quarterly asset growth over quarterly leverage growth across different global financial institutions



Note: he red line in each subplot is the 45 degree line. Clockwise, from top left panel, the relationship between balance sheet size and leverage for GISBs, commercial banks, institutions operating in capital markets and other financial institutions. The classification matches GICS industry codes for each entry in the sample. Source: Datastream, authors calculations.

B Dynamic Factor Model for World Risky Asset Prices – For Online Publication

Let p_t be an n-dimensional vector collecting monthly (log) asset price series $p_{i,t}$, where $p_{i,t}$ denotes the price for asset i at date t. We assume

$$p_t = \Lambda F_t + \xi_t \ . \tag{B.1}$$

F_t is an $(r \times 1)$ vector of common factors ($F_t = [f_{1,t}, \ldots, f_{r,t}]'$) that capture systematic sources of variation among prices and are loaded via the coefficients in Λ that determine how each price series reacts to the common shocks. ξ_t is a $(n \times 1)$ vector of idiosyncratic shocks $\xi_{i,t}$ that capture series-specific variability or measurement errors. We allow elements in ξ_t to display some degree of autocorrelation while we rule out pairwise correlation between assets assuming that all the co-variation is accounted for by the common component. Both the common factors and the idiosyncratic terms are assumed to be zero mean processes.

The factors are assumed to follow a VAR process of order p

$$F_t = \Phi_1 F_{t-1} + \ldots + \Phi_p F_{t-p} + \varepsilon_t, \tag{B.2}$$

where the autoregressive coefficients are collected in the p matrices $\Phi_1, \ldots, \Phi_p$, each of which is $(r \times r)$; the error term ε_t is a normally distributed zero mean process with covariance matrix Q. Any residual autocorrelation is captured by the idiosyncratic component which we assume being a collection of independent univariate autoregressive processes.

In order to distinguish between comovements at different levels of aggregation we model asset prices such that each series is a function of a global factor, a regional factor and an idiosyncratic term. We do so by allowing the vector of common shocks to include both aggregate shocks that affect all series in y_t, and shocks that affect many but not all of them:

$$p_{i,t} = \lambda_{i,g} f_t^g + \lambda_{i,m} f_t^m + \xi_{i,t} \ . \tag{B.3}$$

In Eq. (B.3) the common component ΛF_t is separated into a global factor (f_t^g) and a regional or market-specific factor (f_t^m) which is meant to capture commonalities among many but not all price series. Each $p_{i,t}$ is thus a function of a global factor loaded by all the variables in p_t, a regional or market-specific factor only loaded by those series in p_t that belong to the (geographical or sector-specific) market m, and of a series-specific factor.

Such hierarchical structure is imposed via zero restrictions on some of the elements in

Λ. In particular, we assume the common component to be partitioned into a global and several regional factors. To this aim, let the variables in y_t be such that it is possible to univocally allocate them in B different blocks or regions and, without loss of generality, assume that they are ordered according to the specific block they refer to such that $y_t = [y_t^1, y_t^2, \ldots, y_t^B]'$. Eq. (B.1) can be rewritten as

$$p_t = \begin{pmatrix} \Lambda_{1,g} & \Lambda_{1,1} & 0 & \cdots & 0 \\ \Lambda_{2,g} & 0 & \Lambda_{2,2} & & \vdots \\ \vdots & \vdots & & \ddots & 0 \\ \Lambda_{B,g} & 0 & \cdots & 0 & \Lambda_{B,B} \end{pmatrix} \begin{pmatrix} f_t^g \\ f_t^1 \\ f_t^2 \\ \vdots \\ f_t^B \end{pmatrix} + \xi_t. \tag{B.4}$$

Moreover, further restrictions are imposed on the coefficient matrices in Eq. (B.2) such that Φ_i $(i, \ldots, p)$ and Q are block diagonal.

The model in Eq. (B.1-B.2) can be cast in state space form and the unknowns consistently estimated via Maximum Likelihood (Doz et al., 2011; Engle and Watson, 1981; Reis and Watson, 2010; Bańbura et al., 2011). The algorithm is initialized using principal component estimates of the factors that are proven to provide a good approximation of the common factors when the cross sectional dimension is large.[43] We estimate the model on the price series in (log) difference and obtain the factors via cumulation.[44] We set the number of lags in the factors VAR (p) to be equal to 1. We fit to the data a model with one global and one factor per block/market; the parametrization is motivated by the results in Table B.2.

We fit the model to a vast collection of prices of different risky assets. The geographical areas covered are North America (US and Canada), Latin America (Brazil, Chile, Colombia, and Mexico), Europe (Euro Area, UK, Switzerland and the Scandinavian Countries), Asia Pacific (Japan, Hong Kong, Singapore, Korea, Taiwan), and Australia. The set of commodities considered does not include precious metals. The time span covered is from January 1990 to December 2012. In order to select the series that are included in the global set we proceed as follows: first, for each market, we pick a representative market index (i.e. S&P) and all of its components as of the end of 2012, we then select those that allow us to cover at least 80% of the cross sectional observations by the beginning of 1990, and such that by 1995 we reach a 95% coverage.[45] The procedure allows us to build

[43] Forni et al. (2000); Bai and Ng (2002); Stock and Watson (2002b,a) among others.

[44] Let $\tilde{x}_t \equiv \Delta x_t$ denote the first difference for any variable x_t, then consistent estimates of the common factors in F_t can be obtained by cumulating the factors estimated from the stationary, first-differenced model: $\tilde{p}_t = \Lambda \tilde{F}_t + \tilde{\xi}_t$. In particular, $\hat{F}_t = \sum_{s=2}^{t} \hat{\tilde{F}}_s$ and $\hat{\xi}_t = \sum_{s=2}^{t} \hat{\tilde{\xi}}_s$. Bai and Ng (2004) show that $\hat{F}_t$ is a consistent estimate of F_t up to a scale and an initial condition F_0.

[45] While estimating the Dynamic Factor Model using Maximum Likelihood does not constrain us to

Table B.1: Composition of asset price panels

	North America	Latin America	Europe	Asia Pacific	Australia	Cmdy	Corporate	Total
1975:2010	114	–	82	68	–	39	–	303
1990:2012	364	16	200	143	21	57	57	858

Note: Composition of the panels of asset prices used for the estimation of the global factor. Columns denote blocks/markets in each set, while the number in each cell corresponds to the number of elements in each block.

a final dataset with an overall cross-sectional dimension of $n = 858$. The composition is reported in Table B.1, where each category (in columns) corresponds to one of the blocks (m) within the structure imposed.

Although all series included in the set are priced in US dollars, we verify that the shape of the global factor is not influenced by this choice by estimating the same model on price series in their local currencies (i.e. the currency in which the assets are originally traded). The resulting global factor (not shown) is very similar to the one constructed from the dollar-denominated set both in terms of overall shape and of peaks and troughs that perfectly coincide throughout the time span considered. Intuitively, the robustness of the estimate of the global factor with respect to currency transformations comes directly from the structure imposed in Eq. (B.3). The blocks/markets structure imposed roughly coincides with currency areas, therefore this aspect is likely to be largely captured by the regional factors (see Table B.1).

B.1 The Number of Factors

To choose the number of global factors we use a number of criteria and tests, collected in Table B.2. The table reports the percentage of variance that is explained by the i-th eigenvalue (in decreasing order) of both the covariance matrix and the spectral density matrix, the information criteria in Bai and Ng (2002), where the residual variance of the idiosyncratic component is minimized subject to a penalty function increasing in r, and the test developed in Onatski (2009), where the null of $r-1$ factors is tested against the alternative of r common factors. The largest eigenvalue alone, in both the time and frequency domain, accounts for about 60% of the variability in the data in the longer set and about a fourth of the variation in the shorter, but more heterogeneous set; similarly, the IC criteria reach their minimum when one factor is used, and the overall picture is confirmed by the the p-values for the Onatski test.

work with a fully balanced panel, we want to ensure that none of the categories included in the set is overrepresented at any point in time.

Table B.2: Number of Global Factors.

r	% Covariance Matrix	% Spectral Density	IC_p1	IC_p2	IC_p3	Onatski (2009) Test
			(a) 1975:2010			
1	0.662	0.579	-0.207	-0.204	-0.217	0.015
2	0.117	0.112	-0.179	-0.173	-0.198	0.349
3	0.085	0.075	-0.150	-0.142	-0.179	0.360
4	0.028	0.033	-0.121	-0.110	-0.160	0.658
5	0.020	0.024	-0.093	-0.079	-0.142	0.195
			(b) 1990:2012			
1	0.215	0.241	-0.184	-0.183	-0.189	0.049
2	0.044	0.084	-0.158	-0.156	-0.169	0.064
3	0.036	0.071	-0.133	-0.129	-0.148	0.790
4	0.033	0.056	-0.107	-0.102	-0.128	0.394
5	0.025	0.049	-0.082	-0.075	-0.108	0.531

Note: For both sets and each value of r the table shows the % of variance explained by the r-th eigenvalue (in decreasing order) of the covariance matrix of the data, the % of variance explained by the r-th eigenvalue (in decreasing order) of the spectral density matrix of the data, the value of the IC_p criteria in Bai and Ng (2002) and the p-value for the Onatski (2009) test where the null of $r-1$ common factors is tested against the alternative of r common factors.

C Bayesian VAR – For Online Publication

Let Y_t denote a set of n endogenous variables, $Y_t = [y_{1t}, \ldots, y_{Nt}]'$, with n potentially large, and consider for it the following VAR(p):

$$Y_t = \mathrm{c} + A_1 Y_{t-1} + \ldots + A_p Y_{t-p} + u_t. \tag{C.1}$$

In Eq. (C.1) c is an $(n \times 1)$ vector of intercepts, the n-dimensional A_i $(i = 1, \ldots, p)$ matrices collect the autoregressive coefficients, and u_t is a normally distributed error term with zero mean and variance $\mathbb{E}(u_t u_t') = \Sigma$. We estimate the VAR using standard macroeconomic priors (Litterman, 1986; Kadiyala and Karlsson, 1997; Sims and Zha, 1998; Doan et al., 1983; Sims, 1993); in particular, we use a Normal-Inverse Wishart prior for the VAR coefficients. The Normal-Inverse Wishart prior takes the following form:

$$\Sigma \sim \mathcal{W}^{-1}(\Psi, \nu) \tag{C.2}$$

$$\beta|\Sigma \sim \mathcal{N}(b, \Sigma \otimes \Omega) \tag{C.3}$$

where β is a vector collecting all the VAR parameters, i.e. $\beta \equiv vec([\mathrm{c}, A_1, \ldots, A_p]')$. The degrees of freedom of the Inverse-Wishart are set such that the mean of the distribution exists and are equal to $\nu = n + 2$, Ψ is diagonal with elements ψ_i which are chosen to be a function of the residual variance of the regression of each variable onto its own first p lags. More specifically, the parameters in Eq. (C.2) and Eq. (C.3) are chosen to match the moments for the distribution of the coefficients in Eq. (C.1) defined by the Minnesota priors:

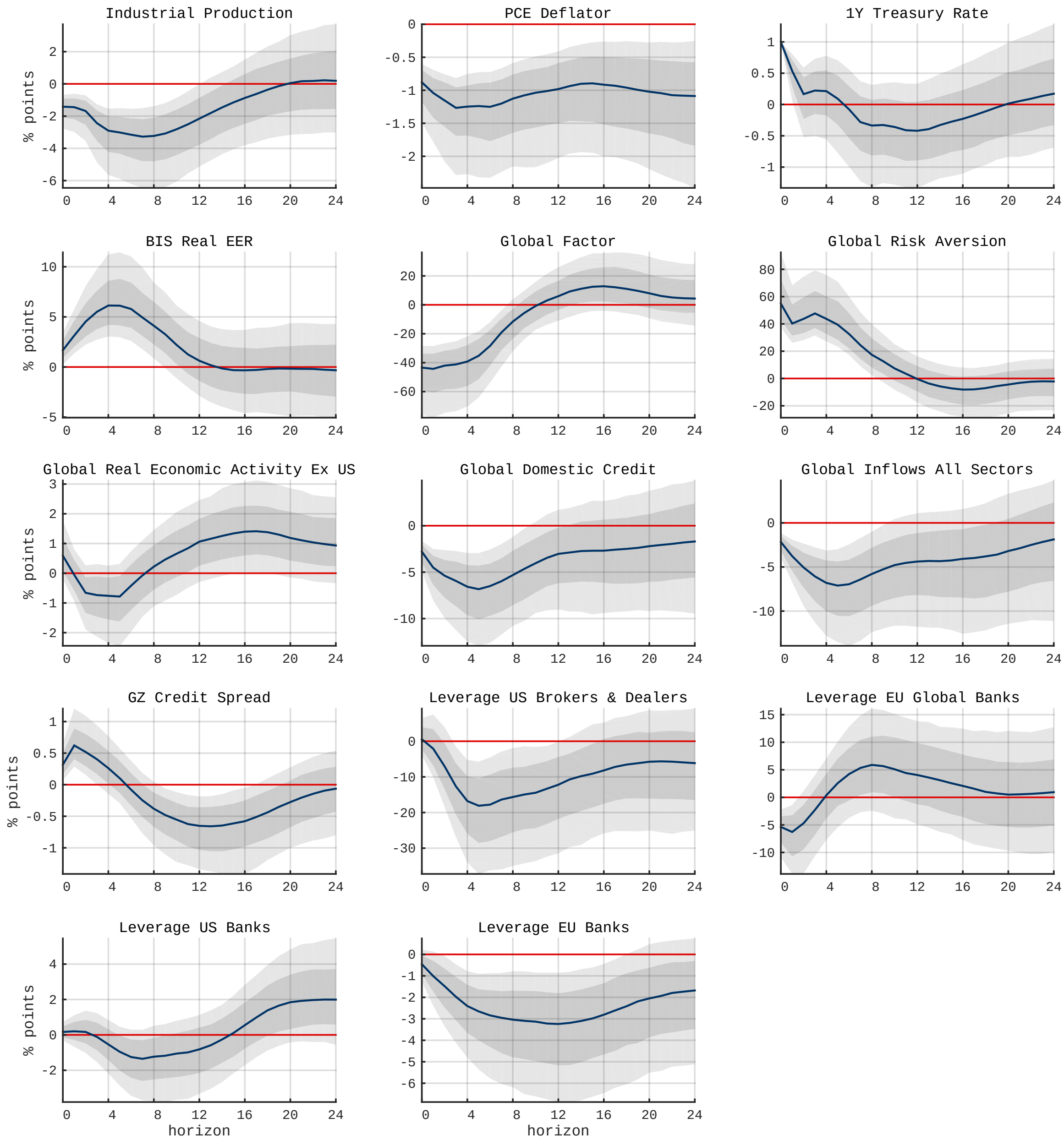
$$\mathbb{E}[(A_i)_{jk}] = \begin{cases} \delta_j & i = 1, j = k \\ 0 & otherwise \end{cases} \qquad \mathbb{V}ar[(A_i)_{jk}] = \begin{cases} \frac{\lambda^2}{i^2} & j = k \\ \frac{\lambda^2}{i^2} \frac{\sigma_k^2}{\sigma_j^2} & otherwise, \end{cases} \tag{C.4}$$

where $(A_i)_{jk}$ denotes the element in row (equation) j and column (variable) k of the coefficients matrix A at lag i ($i = 1, \ldots, p$). When $\delta_j = 1$ the random walk prior is strictly imposed on all variables; however, for those variables for which this prior is not suitable we set $\delta_j = 0$ as in Bańbura et al. (2010). On the right hand side of Eq. (C.4), the variance of the elements in A_i is assumed to be proportional to the (inverse of the) square of the lag (i^2), and to the relative variance of the variables.

λ is the hyperparameter that governs the overall tightness of the priors. We follow Giannone et al. (2015) and treat the λ as an additional model parameter which we estimate in the spirit of hierarchical modelling.

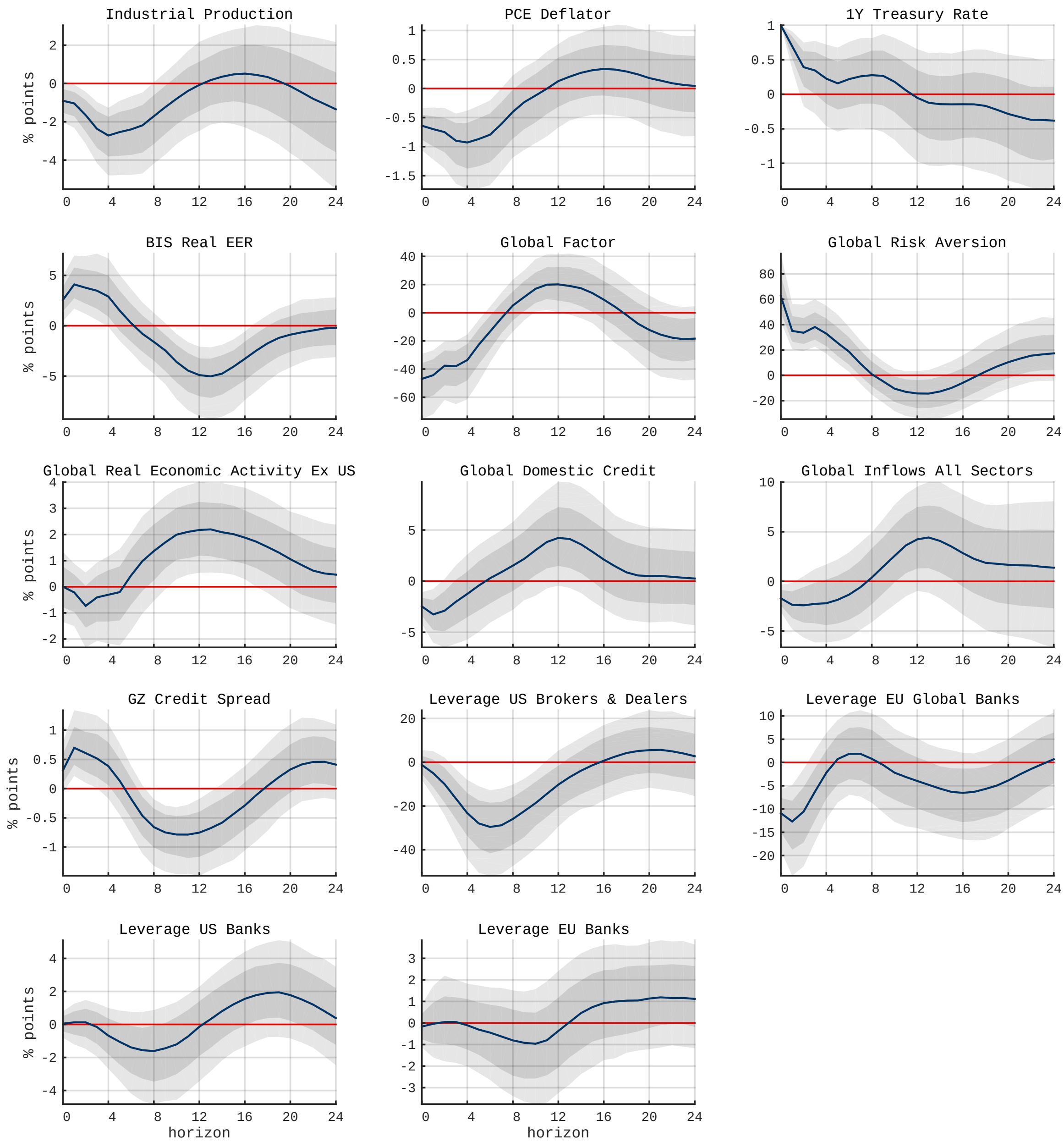
D Other Charts – For Online Publication

FIGURE D.1: GLOBAL VAR (1) - WORLD AGGREGATES



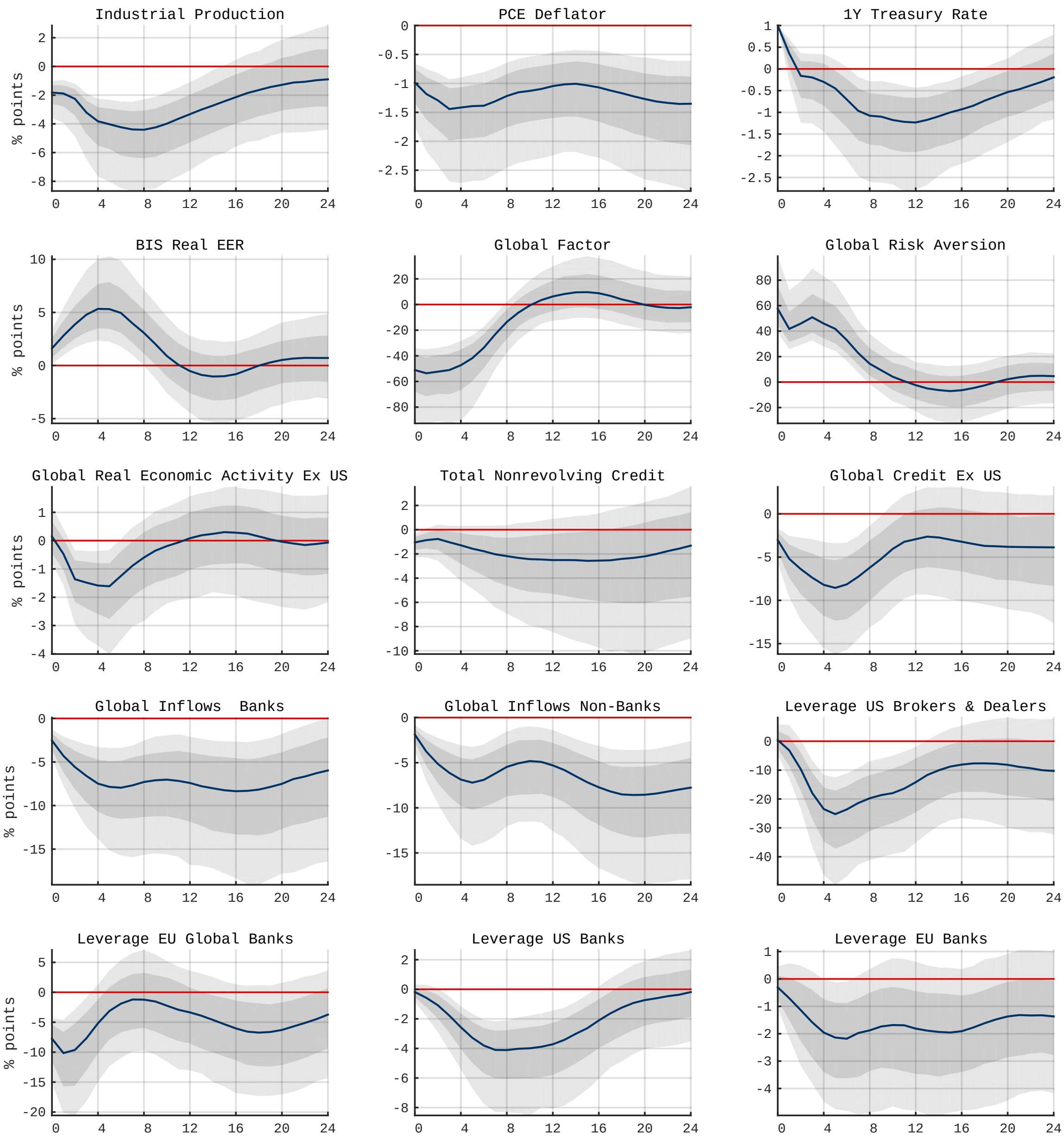
Note: Responses to a US contractionary monetary policy shock that induces a 1% increase in the policy interest rate. Median IRFs with posterior coverage bands at 68% and 90% levels. The shock is identified using a high-frequency external instrument.

Figure D.2: Global VAR (1) - World Aggregates, 1990-2010



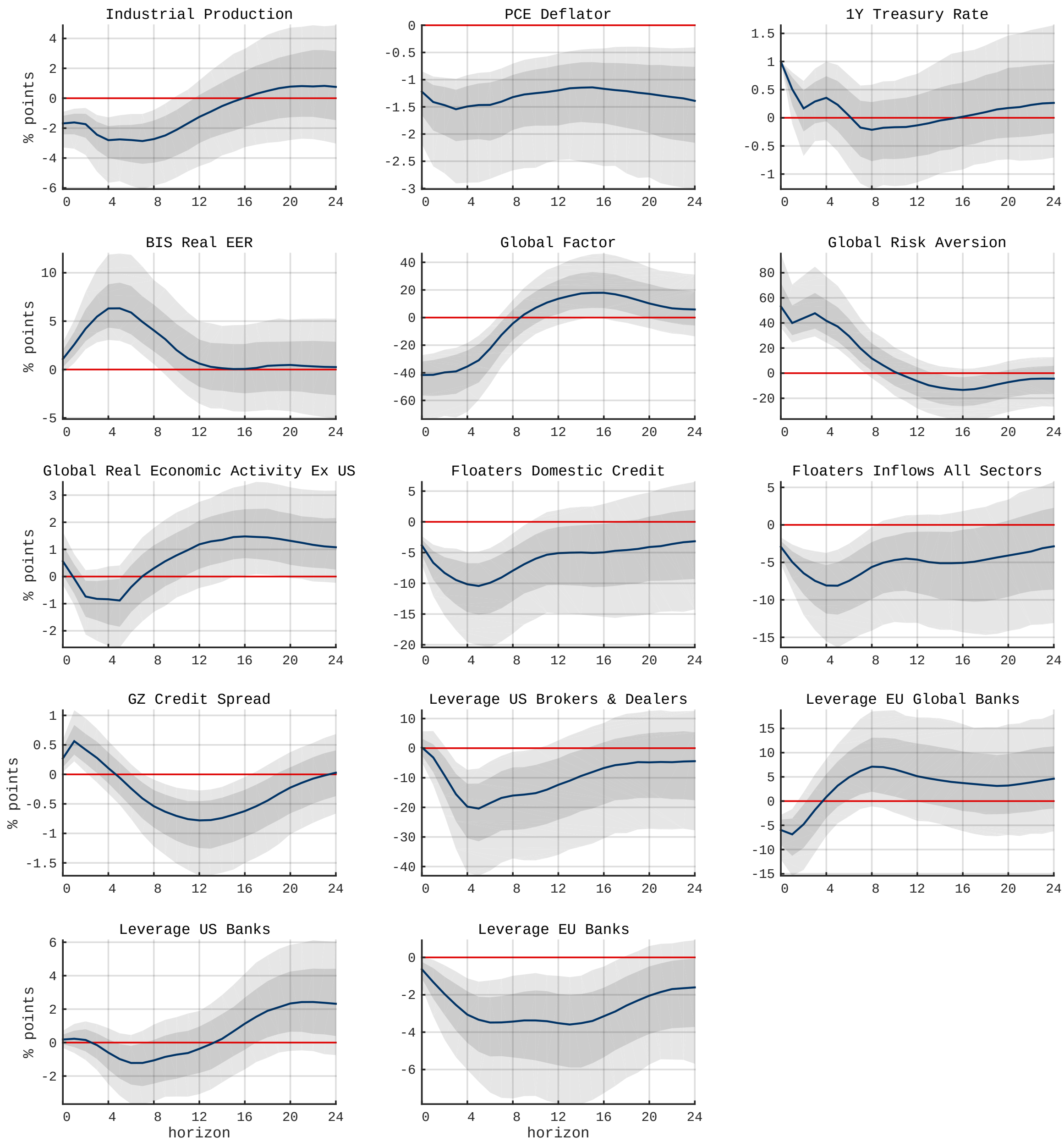
Note: Responses to a US contractionary monetary policy shock that induces a 1% increase in the policy interest rate. Median IRFs with posterior coverage bands at 68% and 90% levels. The shock is identified using a high-frequency external instrument.

Figure D.3: Global VAR (1b) - World Aggregates



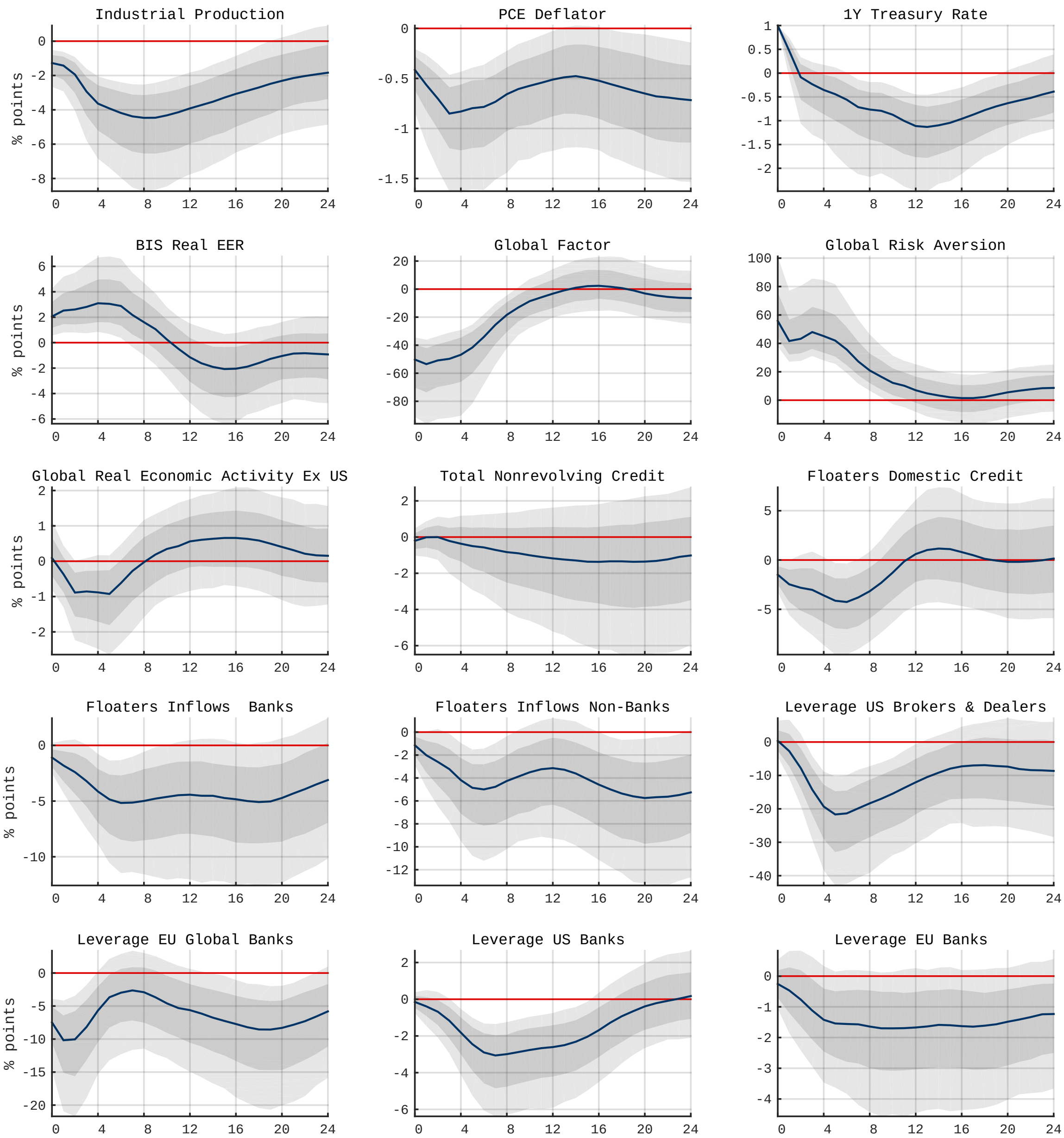
Note: Responses to a US contractionary monetary policy shock that induces a 1% increase in the policy interest rate. Median IRFs with posterior coverage bands at 68% and 90% levels. The shock is identified using a high-frequency external instrument.

Figure D.4: Global VAR (2) - FX Floaters Aggregates



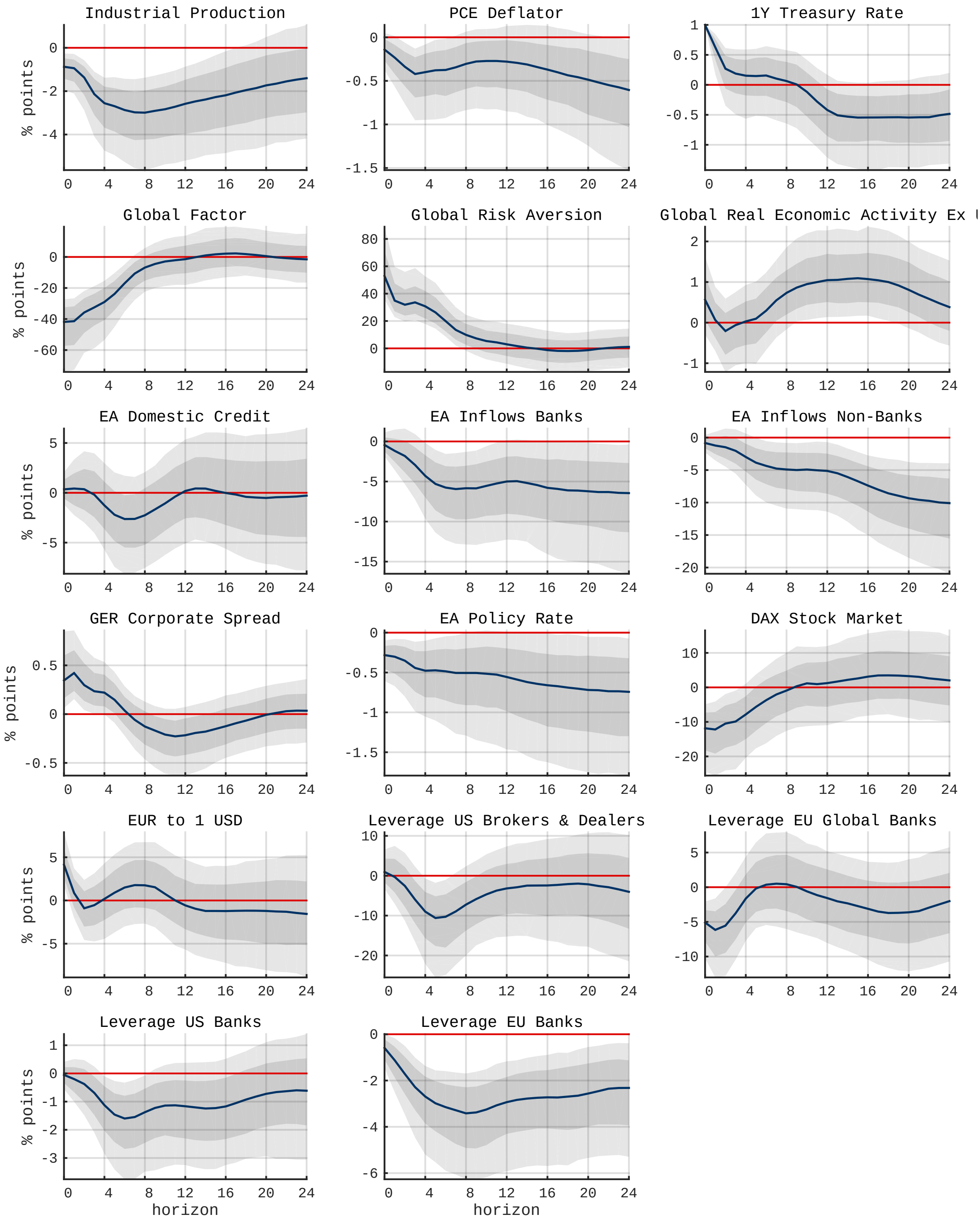
Note: Responses to a US contractionary monetary policy shock that induces a 1% increase in the policy interest rate. Median IRFs with posterior coverage bands at 68% and 90% levels. The shock is identified using a high-frequency external instrument.

FIGURE D.5: GLOBAL VAR (2B) - FX FLOATERS AGGREGATES



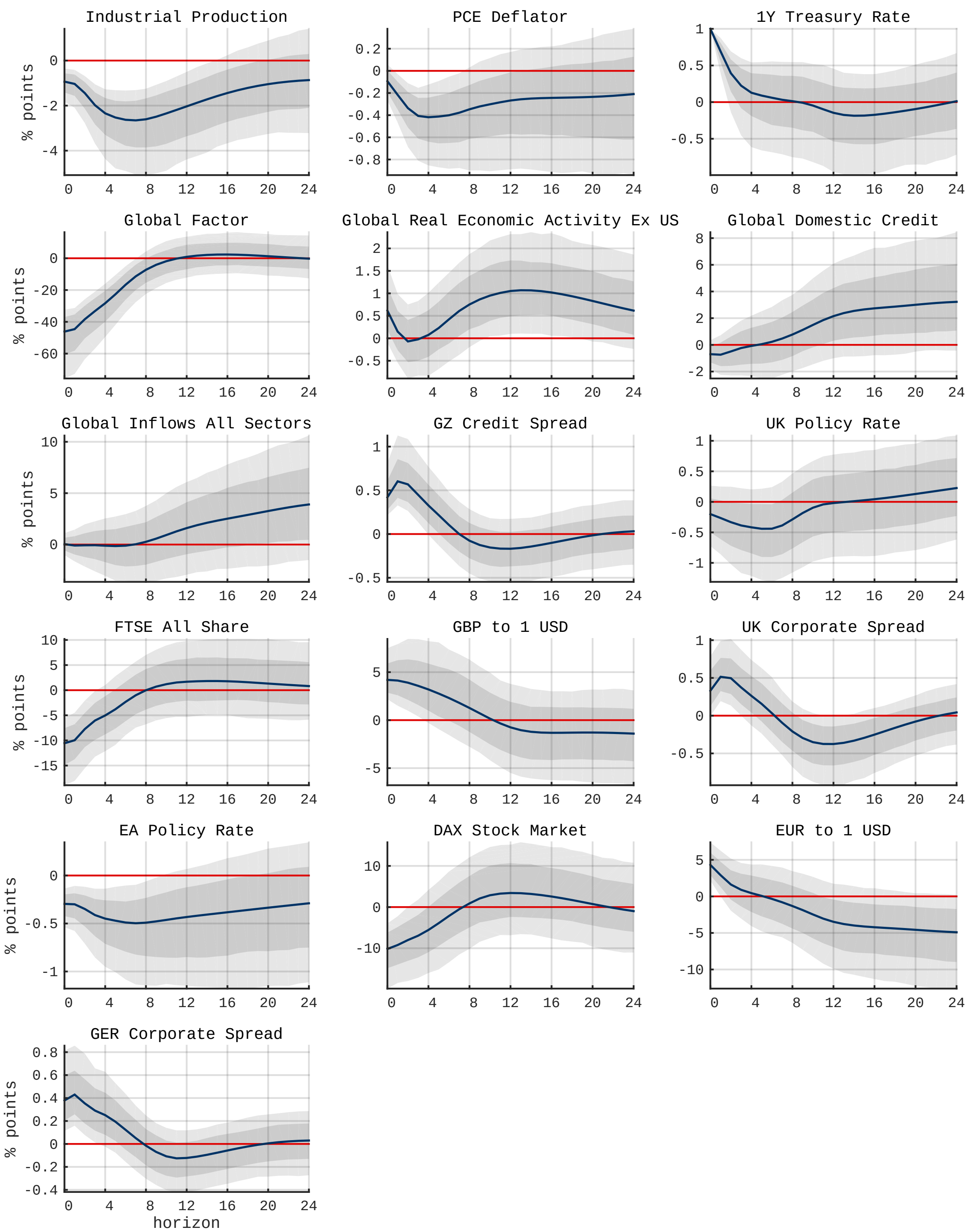
Note: Responses to a US contractionary monetary policy shock that induces a 1% increase in the policy interest rate. Median IRFs with posterior coverage bands at 68% and 90% levels. The shock is identified using a high-frequency external instrument.

Figure D.6: EA-Focus VAR - Euro Area Aggregates



Note: Responses to a US contractionary monetary policy shock that induces a 1% increase in the policy interest rate. Median IRFs with posterior coverage bands at 68% and 90% levels. The shock is identified using a high-frequency external instrument.

Figure D.7: Global VAR (1c) - World Aggregates, Focus on Responses of EA and UK Financial Conditions and Monetary Policy



Note: Responses to a US contractionary monetary policy shock that induces a 1% increase in the policy interest rate. Median IRFs with posterior coverage bands at 68% and 90% levels. The shock is identified using a high-frequency external instrument.

E Robustness: Global Growth and Global Risk Aversion – For Online Publication

E.1 Global Growth

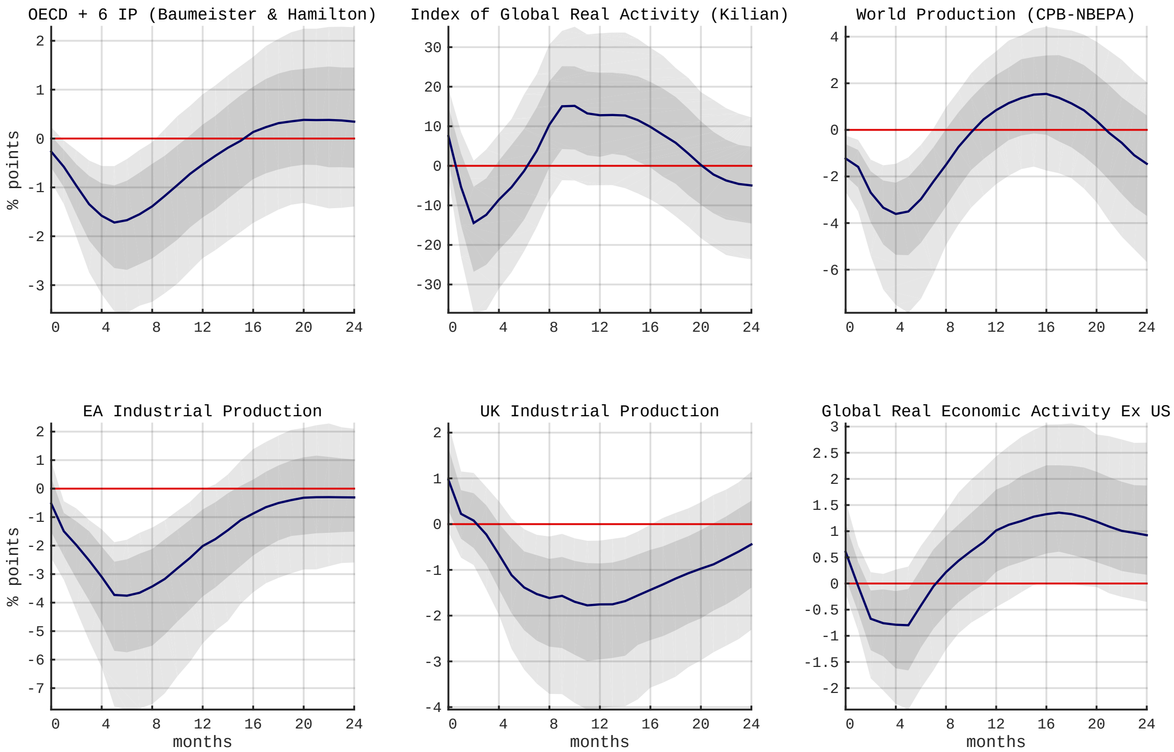
Figure E.1 reports IRFs to a contractionary US monetary policy shock of a variety of global growth measures. The IRFs are obtained by sequentially substituting a different growth variable in the same VAR, the composition of which is equal to that of Figure D.1 (our benchmark global VAR). All VARs are estimated over the sample 1980:01-2010:12 with the exception of the one that includes the World Production (CPB-NBEPA) which is only available from 1990:01.

The measures reported in the figure are:

- World industrial production measured as the weighted average of the IP of OECD countries + 6 EMEs (originally distributed by the OECD, updated in Baumeister and Hamilton, 2019) available at `https://sites.google.com/site/cjsbaumeister/research`;
- Global real activity as measured in Kilian (2009) updated and corrected in Kilian (2019) and available at `https://sites.google.com/site/lkilian2019/research/data-sets`;
- World industrial production as weighted average of IP indices of both AEs and EMEs, measured by the CPB Netherlands Bureau for Economic Policy Analysis and distributed in their World Trade Monitor publication, available at `https://www.cpb.nl/en/worldtrademonitor`;
- IP for the EA, fixed composition. The variable is NSA at source (ECB data series ID: STS.M.I8.W.PROD.NS0020.3.000), we have seasonally adjusted it prior to its inclusion in the VAR;
- UK Index of Production as distributed by the UK Office for National Statistics;
- Global real economic activity Ex US constructed as the component of the Baumeister and Hamilton (2019) series that is orthogonal to the US cycle. This is our benchmark variable used in Section 3.

Results indicate that global production seems to contract following a US monetary tightening mostly as a result of the US slowdown. While the two global IP measures decline significantly, the component that is orthogonal to the US cycle (last panel) is not affected in a significant way. Kilian's measure that is based on nominal shipping costs

Figure E.1: Alternative Measures of Global Growth



Note: Responses of alternative measures of global growth to a US contractionary monetary policy shock that induces a 1% increase in the policy interest rate. Median IRFs with posterior coverage bands at 68% and 90% levels. The shock is identified using a high-frequency external instrument. Combines IRFs estimated from different VARs all estimated over the sample 1980-2010 with the exception of the one that includes the World Production (CPB-NBEPA) which is only available from 1990. With the exception of the global growth variable the composition is identical to that of Figure D.1.

is only marginally affected. On the other hand, production indices in the UK and the EA do contract with some delay. Figure 10 in the paper however also shows that the US monetary contraction significantly affects financial conditions in both the UK and EA, which in turn may be detrimental for real activity.

E.2 Global Risk Aversion

In general asset pricing models, the residuals of the projection of asset prices on realized variance can be decomposed into news about future expected GDP growth, changes in risk-free discount rates, and changes in risk aversion (see e.g. Zhou, 2018).[46]

[46]Some papers which posit a *constant* level of risk aversion add another term, time-varying macroeconomic uncertainty. Instead, we allow for time-varying risk aversion as in Bekaert et al. (2013); Bekaert and Hoerova (2014). We rationalize the time variation using a model where fluctuations in risk aversion reflect the time-varying importance of different financial intermediaries in international markets (see

Our benchmark risk aversion proxy ignores fluctuations in contemporaneous discount rates and future expected GDP growth, in line with the simple model of Section 4.[47] It is computed from the centered residuals of the following regression:

$$GFAC_t = \alpha + \beta \ln(GRVAR_t) + \varepsilon_t,$$

where $GFAC_t$ is our global factor (in log units) and $GRVAR_t$ is the measure of realized variance obtained from the returns of the Global MSCI Index.

In order to account for these more general models, we construct an alternative aggregate global risk aversion proxy by regressing the global factor on global realized variance, and also on discount rates and survey forecasts of GDP growth for four countries which we take to be representative of our global set: the US, the UK, Germany and Japan. Specifically, we recover a refined proxy for risk aversion as the centered residual of the following regression:

$$GFAC_t = \alpha + \beta \ln(GRVAR_t) + \sum_j \gamma_j DR_{j,t} + \sum_j \delta_j \mathbb{E}_t \left[GDP_{j,t+12}\right] + \varepsilon_t,$$

where j denotes the following 4 countries: US, GER, UK, JP; DR_t are discount rates (averages of daily figures) obtained from the St Louis Fed FRED database, and $\mathbb{E}_t \left[GDP_{j,t+12}\right]$ are monthly survey forecasts for GDP growth a year (12 months) ahead, constructed from the calendar-year forecasts distributed by Consensus Economics. Both regressions are estimated at monthly frequency.

The two proxies are plotted in the bottom panel of Figure E.2 as a solid and dashed line respectively. The new series shares many similarities with our original one, with some small differences; there is for example a quicker reversal to a low-perceived-risk regime following the global financial crisis, and higher risk-aversion in the early 2000s.

Figure E.3 reports IRFs to a contractionary US monetary policy shock of a variety of risk aversion measures. The IRFs are obtained by sequentially substituting a different risk aversion variable in the same VAR, the composition of which is equal to that of Figure D.1 (our benchmark global VAR). All VARs are estimated over the sample 1990:01-2010:12.

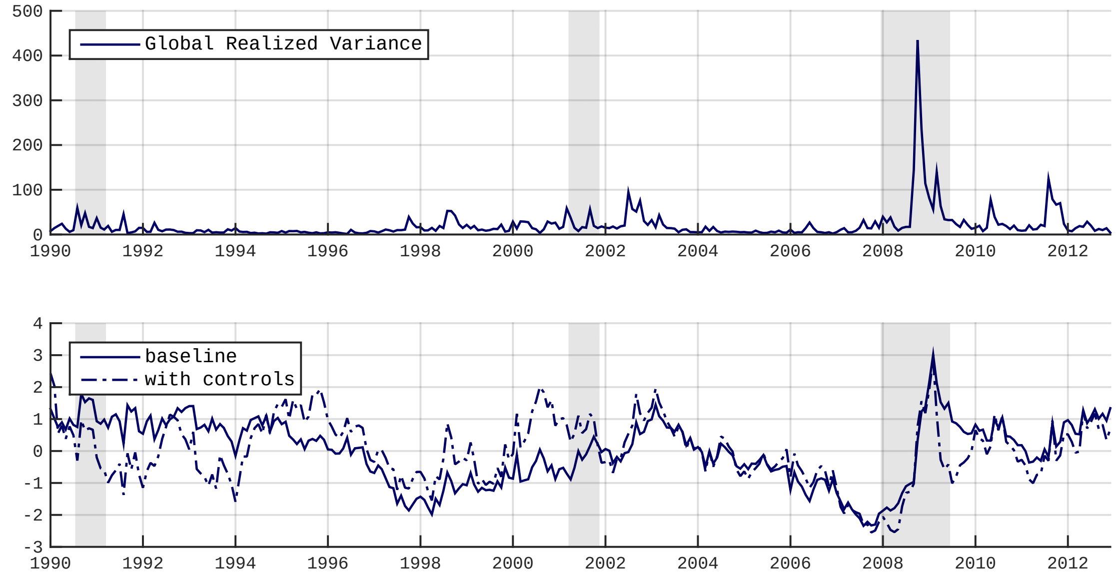
The measures reported in the figure are:

- Our benchmark proxy for aggregate risk aversion;
- A more general proxy for aggregate risk aversion that also controls for discount rates and GDP growth forecasts in the 4 countries listed;

Section 4).

[47] We note that our BVAR controls for past discount rates and past GDP growth.

Figure E.2: global factor decomposition (including controls)

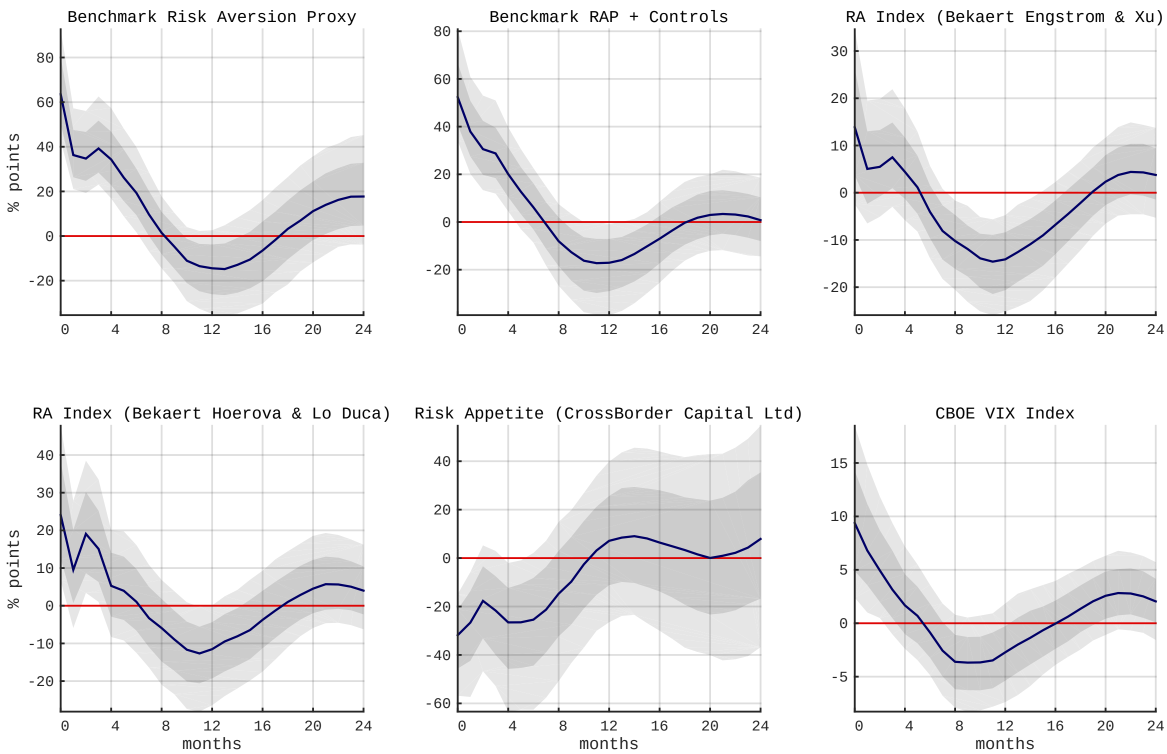


Note: [top panel] Monthly global realized variance measured using daily returns of the MSCI Index. [bottom panel] Index of aggregate risk aversion calculated as (the inverse of) the residual of the projection of the global factor onto the realized variance. Solid line: baseline estimate of Section 2; Dashed line: includes controls for discount rates and expected growth. Shaded grey areas highlight NBER recession times. *Source:* Global Financial Data and authors calculations.

- The risk aversion index for the US market constructed in Bekaert et al. (2019) available at `https://www.nancyxu.net/risk-aversion-index`;
- The risk aversion index for the US market constructed in Bekaert et al. (2013) as a variance risk premium along the lines of our benchmark aggregate risk aversion proxy, available at `http://mariehoerova.net/RA_UC_series.xls`;
- A measure of risk appetite distributed by CrossBorder Capital Ltd and calculated as the difference between the Equity Exposure Index (based on the balance sheet exposure of all investors (by type) in the asset class) and the Bond Exposure Index (equivalent of the equity exposure index for bonds);
- The CBOE VIX Index (`http://www.cboe.com/products/vix-index-volatility/vix-options-and-futures/vix-index/vix-historical-data`).

The chart confirms that a US monetary contraction triggers an increase in risk aversion (decline in risk appetite), regardless of the specific measure that is used.

Figure E.3: Alternative Measures of Risk Aversion



Note: Responses of alternative measures of risk aversion to a US contractionary monetary policy shock that induces a 1% increase in the policy interest rate. Median IRFs with posterior coverage bands at 68% and 90% levels. The shock is identified using a high-frequency external instrument. Combines IRFs estimated from different VARs all estimated over the sample 1990-2010: with the exception of the global growth variable the composition is identical to that of Figure D.1.