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Credit constraints and the international propagation of US financial shocks☆

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ABSTRACT

This paper investigates whether credit constraints in the US economy amplify the international propagation of US financial shocks. We model the dynamics of the US economy jointly with global macroeconomic and financial variables using a threshold vector autoregression. This model captures regime-specific dynamics conditional on the severity of credit constraints in the US economy. We identify three main episodes of tight credit in US financial history over the past thirty years. These occur in the late-1980s, in the early 2000s, and during the 2007–09 financial crisis. We find that US financial shocks are associated with a significant contraction in global economic activity in times of tight credit. By contrast, there is little impact of US financial shocks on the global economy in normal times. This asymmetry highlights an international dimension of the US financial accelerator mechanism.

1. Introduction

The 2007–08 turmoil in US financial markets gave rise to a credit crunch with widespread effects on the global economy. Considering the pivotal role of the United States in international financial markets, a key question is how financial shocks that originate from the US propagate across the globe. In particular, we ask whether US financial shocks generate global spillovers that depend nonlinearly on the severity of credit constraints in the US economy. The novelty of this paper is to assess the impact of US financial shocks in an empirical open-economy model that features nonlinear dynamics. Our results reveal that credit constraints amplify the international propagation of US financial shocks.

We model the dynamics of the US economy jointly with global macroeconomic and financial variables in a structural threshold vector autoregression (TVAR). This model distinguishes between "normal" and "tight" credit regimes in the US economy (see also Balke, 2000). In contrast to models in which regime switching is governed by a latent Markov-process, transition across regimes in the TVAR is determined by an endogenous variable that measures the tightness of credit constraints in the US economy. If this variable crosses a threshold, the economy shifts from a regime characterized by unconstrained access to credit to a regime in which borrowers face stringent credit constraints. The model parameters vary across these two regimes, which allows us to assess the regime-specific effects of US financial shocks.

The excess bond premium (EBP) proposed by Gilchrist and Zakrajsek (2012) enters the TVAR model as the threshold variable. The EBP reflects a risk premium demanded by investors for bearing exposure to credit risk across the entire maturity spectrum (from 1- to 30-years) and the range of credit quality (from D to AAA) in the US corporate bond market, beyond the compensation for the usual counter-cyclical movements in expected corporate default.

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The EBP is thus a potentially useful measure of credit supply conditions in the US economy. Using a DSGE model, Gilchrist and Zakrajsek (2011) show that an adverse financial shock, calibrated to match fluctuations in the EBP, is associated with a reduction in the risk-bearing capacity of the financial sector that raises the cost of external finance for non-financial borrowers, leading to a decline in aggregate spending and production.

Our objective is to investigate the regime-specific effects of US financial shocks, associated with unexpected changes in the EBP. We begin with formulating a TVAR for the US economy which comprises output, prices, bank loans, the federal funds rate, and the EBP. This model specification is augmented with aggregate global output and with the US real effective exchange rate (REER) to study the propagation of EBP shocks to the global economy. The seven-variate TVAR is subsequently enlarged by including global realized stock market volatility, trade between the US and the rest of the world, global consumer prices, the world price of oil, and the global interest rate in order to analyze potential shock transmission channels. We identify EBP shocks recursively, postulating that the EBP may react contemporaneously to all shocks that affect the economy, while the remaining variables respond with a delay to EBP shocks.

Using data for the period from January 1984 to December 2012, we detect three major tight-credit episodes in US financial history. The first episode takes place during the savings and loan crisis of the 1980s, the second occurs in the early 2000s, and the third is associated with the 2007–09 financial crisis. The US economy responds asymmetrically to an unexpected rise in the EBP when distinguishing between normal and tight credit regimes. In the normal credit regime, the US economy is resilient in the face of financial disturbances, and the impact of EBP shocks on the macroeconomy is not statistically significant. In contrast, EBP shocks have statistically significant real effects in the tight credit regime. Specifically, a 10 basis point positive shock to the EBP has a negative impact of about 0.6 percentage points on US output one year after the shock. The surprise increase in the EBP is also associated with a significant reduction in the volume of bank loans and a fall in consumer prices. Moreover, the federal funds rate drops by about 10 basis points one year after the EBP shock, which suggests that the Federal Reserve eases monetary policy amid deteriorating macroeconomic conditions.

Crucially, we find that global economic activity contracts significantly following EBP shocks when credit is scarce in the US, while EBP shocks have little impact in normal times. Global output declines in the tight credit regime by about 0.6 percentage points one year after a 10 basis point positive shock to the EBP, which is comparable to the effect of the shock on economic activity in the US. A historical decomposition shows that EBP shocks contributed negatively to global output growth particularly during the 1980s, in the early 2000s, and during the 2007–09 global financial crisis. The contribution of EBP shocks to global output growth as a fraction of the contributions of all structural shocks is 45% on average in the tight credit regime, while it equals 18% on average in the normal credit regime. EBP shocks are thus a relatively more important driver of global business cycles during periods of tight credit.

EBP shocks propagate to the global economy through various channels in the tight credit regime, while all transmission channels remain muted in normal times. First, we find that during a credit crunch the US dollar appreciates significantly in response to an unexpected rise in the EBP, which is in line with the exchange rate movements during the 2007–09 crisis. This result is consistent with the view advanced by Fratzscher (2009), Gourinchas et al. (2012), and Prasad (2014) that the US acts as a “global insurer” during financial crisis periods, as US dollar denominated assets satisfy increased demand from flight-to-liquidity and safe-haven flows. Second, the EBP shock increases global stock market volatility in the tight credit regime. This suggests that international stock markets serve as a conduit for the EBP shock in the spirit of an “international finance multiplier” mechanism advanced by Krugman (2008) and featured in the structural models of Devereux and Yetman (2010), Devereux and Sutherland (2011), Dedola and Lombardo (2012), and van Wincoop (2013). In addition, the EBP shock is followed by a significant reduction in US trade with the rest of the world, in global consumer prices, in the world price of oil, and in the global interest rate. In sum, a global economic downturn unfolds in the tight credit regime after an US financial shock.

Our key results are robust to a variety of robustness checks. First, we implement an alternative approach to identify US financial shocks, which is based on a combination of zero and sign restrictions on the estimated impulse responses in line with Peersman (2012). Second, to control for the possibility that EBP shocks may be confounded with uncertainty shocks, we include into the baseline model specification the CBOE VXO implied volatility index which is a popular proxy for uncertainty (see Bloom, 2009). Third, we replace the real GDP weights used to obtain global variables in the baseline model specification with trade weights and with financial weights. Fourth, we estimate our model on data up to November 2007 in order to verify whether our results are dominated by the recent Great Recession period. Finally, since the re-estimation of our various TVAR model specifications may deliver model estimates inconsistent with one another, we also estimate the full system jointly with all variables. The asymmetric impact of EBP shocks on the global economy proves to be a robust feature across all different settings considered.

The 2007–08 global financial crisis has revived the literature on the international transmission of financial shocks (see, e.g., Bagliano and Morana, 2012; Bekaert et al., 2014; Fry-McKibbin et al., 2014; Blatt et al., 2015). Several transmission channels through which the turmoil emanating from US financial markets spread to the global economy have been documented, for instance, cross-border holdings of asset-backed securities (e.g., Longstaff, 2010; Manconi et al., 2012) and bank credit default swaps (e.g., Eichengreen et al., 2012), balance-sheet rebalancing by globalized banking conglomerates (e.g., Cettorelli and Goldberg, 2012; Giannetti and Laeven, 2012; De Haas and Van Horen, 2013), equity market contagion (e.g., Bekaert et al., 2014), and the collapse of global trade (e.g., Bems et al., 2011). Thus far, however, the empirical literature on the link between credit market frictions and international financial spillovers is scarce. The papers by Helbling et al. (2011) and Eickmeier and Ng (2015) on the international transmission of US credit shocks are perhaps closest to ours, but use constant-parameter VAR models.

Our paper belongs to a growing group of empirical studies on macro-financial linkages (e.g., Lown and Morgan, 2006; Gilchrist and Zakrajsek, 2011, 2012; Fornari and Stracca, 2012; Meeks, 2012; Prieto et al., 2016). Macro-financial models often fail to capture nonlinear amplification effects. Balke (2000) and Hubrich and Tetlow (2015) constitute important exceptions, but focus on a closed economy setup. We add to existing studies by accounting for nonlinearities in a macro-financial model that incorporates global spillovers.

From a methodological perspective, the abrupt transition VAR model proposed by Artis et al. (2007) to estimate the effects of US output shocks on European economies constitutes the approach most closely related to ours. Nevertheless, time variation has been introduced in different ways into VAR models in order to study the international transmission of financial shocks. For example, Favero and Giavazzi (2002) formulate a VAR with variance regime shifts, Blatt et al. (2015) employ a VAR with structural breaks in the VAR parameters, and Abbate et al. (2016) propose a time-varying parameter factor augmented VAR. In this paper we adopt the TVAR approach because it captures explicitly the role of credit constraints in the nonlinear transmission of financial shocks. Hence, it enables us to provide new insights into the mechanisms through which US financial shocks propagate to the global economy.

Our econometric approach constitutes an empirical counterpart to the recently developed macroeconomic models that feature "occasionally binding" financial constraints. A consensus seems to emerge from structural models that occasionally binding constraints are central to understanding the nonlinearities observed during financial crisis episodes (see, e.g., Mendoza, 2010; Bianchi, 2011; Brunnermeier and Sannikov, 2014; Perri and Quadrini, 2014). Specifically, this strand of the literature predicts that economies are resilient to shocks as long as the flow of credit is unconstrained. Binding credit constraints can, however, give rise to aggregate economic contraction. Moreover, recent theoretical studies have shown that financial frictions lead to an amplification of cross-border shocks, and structural models featuring such frictions provide a more realistic picture of international macroeconomic fluctuations (see, e.g., Krugman, 2008; Devereux and Yetman, 2010; Devereux and Sutherland, 2011; Olivero, 2010; Kollmann et al., 2011; Dedola and Lombardo, 2012; van Wincoop, 2013). Empirical models that ignore nonlinear amplification mechanisms may therefore deliver biased estimates of cross-country spillovers. We complement the theoretical literature with empirical evidence on regime-specific spillover effects that arise from financial frictions in the US economy.

The remainder of the paper is organized as follows. We present our econometric approach in Section 2. Section 3 offers a brief description of the data, and it outlines our empirical results. Finally, Section 4 summarizes our findings, and it concludes the paper.

2. Econometric framework

At the core of our analysis is a dynamic system that comprises five variables for the US economy: the first difference of the log of industrial production (Δq_t), the first difference of the log of consumer prices (π_t), the first difference of the log of real bank loans (Δl_t), the effective federal funds rate (i_t), and the EBP (ebp_t). This system is augmented with a weighted average that captures the growth rate of aggregate global output (Δq_t^*) and the first difference of the log of the REER (Δe_t). Hence, our baseline model specification labeled as TVAR-7 is given by $Y_t = [\Delta q_t^*,\ \Delta q_t,\ \pi_t,\ \Delta l_t,\ i_t,\ \Delta e_t,\ ebp_t]$.

We assume that the $N \times 1$ vector Y_t follows a structural threshold vector autoregressive model given by:

$$Y_t = \begin{cases} \mathrm{A}^1 Y_t + \Theta^1(L)Y_t + \varepsilon_t^1 & \text{if } \ ebp_{t-d} < \gamma, \\ \mathrm{A}^2 Y_t + \Theta^2(L)Y_t + \varepsilon_t^2 & \text{if } \ ebp_{t-d} \geq \gamma, \end{cases} \tag{1}$$

for $t \in \{1,\dots, T\}$. This model constitutes a regime-switching VAR according to which Y_t follows regime-specific dynamics determined by the EBP which acts as a threshold variable with a delay of d months (ebp_{t-d}). If the EBP crosses a threshold level γ, the economy shifts from a regime in which access to credit is unconstrained ("normal" credit regime) to a regime in which borrowers face stringent credit constraints ("tight" credit regime). The regime-specific parameter matrices A^r reflect the contemporaneous relationships between the endogenous variables contained in Y_t, and the lag polynomial matrices $\Theta^r(L) = \Theta_1^r L^1 + \dots + \Theta_{p^r}^r L^{p^r}$ describe their dynamic interaction in regime r ($r = 1, 2$). The vectors of orthogonal regime-specific shocks ε_t^r are assumed to be normally distributed with zero mean and regime-dependent positive definite covariance matrices $\Sigma_\varepsilon^r = E(\varepsilon_t^r \varepsilon_t^{r\prime})$. The model described in Eq. (1) constitutes a variant of the TVAR models proposed by Tsay (1998) and Balke (2000) with regime-switching volatility in line with Galvao (2006) and Artis et al. (2007).

The reduced form of the TVAR model is given by:

$$Y_t = \begin{cases} B^1(L)Y_t + u_t^1 & \text{if } \ ebp_{t-d} < \gamma, \\ B^2(L)Y_t + u_t^2 & \text{if } \ ebp_{t-d} \geq \gamma, \end{cases} \tag{2}$$

where $B^r(L) = (I - \mathrm{A}^r)^{-1}\Theta^r(L)$ are p^r-order lag-polynomial matrices of the reduced form coefficients (where $p^r \in \mathbb{N}$), and where $u_t^r \sim (0, \Sigma_u^r)$ are vectors of reduced form Gaussian white noise forecast errors with $\Sigma_u^r = E(u_t^r u_t^{r\prime})$ positive definite. The reduced form parameters are estimated using the maximum likelihood estimator (MLE) described in Galvao (2006), and the autoregressive lag length is selected using information criteria proposed by Tsay (1998) and Wong and Li (1998). A detailed description of the estimation procedure is presented in the Appendix.

Our objective is to investigate the regime-specific effects of US financial shocks, associated with unexpected changes in the EBP. Conditional on the threshold γ, the TVAR constitutes a piecewise linear VAR model, which allows us to calculate regime-specific structural impulse response functions that describe the effects of EBP shocks within each regime. The regime-specific structural shocks relate to the reduced-form forecast errors according to $\varepsilon_t^r = \mathrm{A}^r u_t^r$. Identification of regime-specific shocks can be achieved by imposing orthogonality restrictions on the contemporaneous relationships A^r.

We identify EBP shocks via a recursive scheme which is implemented by performing a Cholesky decomposition of the regime-specific reduced-form covariance matrices with the EBP ordered last. The identifying assumption entails that the EBP responds contemporaneously to all shocks hitting the economy, while all remaining variables react with a delay to EBP shocks. Our identification scheme is in line with related VAR studies in which financial shocks are identified recursively by ordering the financial variable of interest below other macroeconomic and financial variables (see, e.g., Lown and Morgan, 2006; Gilchrist and Zakrajsek, 2011; 2012; Bagliano and Morana, 2012; Hubrich and Tetlow, 2015). This identifying assumption is motivated by the high-frequency nature of financial markets, whereby asset prices and risk premia promptly reflect economic developments. We attach an economic interpretation solely to the EBP shock, while we do not interpret the remaining orthogonal shocks from a structural perspective, i.e., these may reflect a mixture of the true underlying structural disturbances.

When analyzing the international transmission of financial shocks, Dungey and Martin (2007) and Blatt et al. (2015) make a distinction between common shocks and cross-border shock spillovers based on the timing of the initial impact of the shock. Common shocks are thought to have contemporaneous effects, while transmission across borders may require a certain time lag to materialize, such that spillovers refer to transmission in period t or later of shocks which occurred in period t–1. Conforming to this idea, the recursive ordering might help us to distinguish US-specific shocks from global common shocks. Nevertheless, there might still be global financial shocks, or financial shocks that originate from another country, which co-move with US financial shocks and also affect the global economy with a delay. Ideally, one would control for this possibility by using a global EBP which, however, is not available.[1] This constitutes an important caveat that should be kept in mind when interpreting our findings.

To study the diverse shock transmission channels, we expand the TVAR-7 to include the log of global realized stock market volatility (vol_t^*), the first difference of the log of global trade (Δtrd_t^*), the first difference of the log of global consumer prices (π_t^*), the first difference of the log of the world real price of oil

[1] Gilchrist and Mojon (2016) construct credit risk indicators for euro area banks and non-financial corporations following Gilchrist and Zakrajsek (2012). However, their time series date back only to 1999.

(Δoil_t^*), and the global interest rate (i_t^*) into the baseline model specification. The TVAR is a richly parameterized model, and with the inclusion of additional variables into the model the degrees of freedom necessary to provide reliable parameter estimates are quickly consumed. Therefore, in order to keep the estimated model specifications as parsimonious as possible, the additional global variables are included one-at-a-time into the baseline TVAR-7 specification. Nevertheless, we also consider joint estimation of the full system for the sake of robustness.

3. Empirical results

3.1. Data

We use monthly data for the period from January 1984 to December 2012. The first observation is cancelled due to first differencing, which leaves us with $T = 347$ observations. The sample begins with the Great Moderation and its end is constrained by the availability of the EBP. We obtain the US industrial production index, the US consumer price index, the volume of commercial and industrial loans issued by all US commercial banks, and the effective federal funds rate from the FRED database of the Federal Reserve Bank of St. Louis. The REER comes from the OECD Main Economic Indicators database, and it is based on competitiveness-weighted manufacturing consumer price indices for the overall economy in dollar terms. The weights take into account the structure of competition in both export and import markets of the goods sector of 34 OECD countries and 15 non-OECD countries. An increase in the REER indicates a real effective appreciation.

Global variables are obtained as weighted averages of the time series for 18 major economies.[2] The weights reflect the average overall size of the economy over the estimation period, measured by average PPP-adjusted real GDP from the Penn World Tables. Global output is measured by industrial production data obtained from the OECD. Realized stock market volatility is obtained as the sum of squared daily stock market returns within each month using the MSCI price index of the total national stock market. We proxy US trade by the total sum of bilateral imports and exports between the US and its 18 counterparts (deflated by US CPI), obtained from the IMF Direction of Trade statistics. We retrieve national CPI indices from Datastream, and we use short-term interest rates that constitute the main monetary policy instrument in each country. Finally, we use the West Texas Intermediate spot oil price measured in dollars per barrel from the FRED database, and deflate it by US CPI to obtain real values.

3.2. The excess bond premium and US credit supply

The EBP proposed by Gilchrist and Zakrajsek (2012) arguably constitutes a comprehensive measure of credit supply conditions in the US economy. To obtain the EBP, Gilchrist and Zakrajsek (2012) construct a composite credit spread index as an arithmetic average of credit spreads on senior unsecured corporate bonds issued by 1,112 nonfinancial firms that encompass the entire maturity spectrum (from 1- to 30-years) and the range of credit quality (from D to AAA) in the US corporate bond market. For each firm, the credit spread for a corporate bond of a given maturity is obtained as the difference between the corporate bond yield and the yield of a corresponding synthetic risk-free security from the Treasury yield curve. Gilchrist and Zakrajsek (2012) decompose the credit spread index into various components using a Black–Scholes–Merton option-pricing model. This model captures (i.) the systematic counter-cyclical movements in firm-specific distance-to-default, (ii.) the level, slope and curvature of the Treasury yield curve, and (iii.) the realized volatility of ten-year Treasury bonds. The EBP is the residual component unexplained by these factors, it thus reflects systematic deviations in the pricing of US corporate bonds from the expected default risk of the underlying issuers.

Fig. 1 depicts the quarterly average of the EBP together with the net percentage of US banks tightening loan underwriting standards for commercial and industrial loans to large and middle-market firms, obtained from the Senior Loan Officer Opinion Survey on Bank Lending Practices (SLOOS), which constitutes another popular proxy for US credit supply. The SLOOS is conducted four times per year by the Federal Reserve among major US banks that account for around 60% of all US bank loans and about 70% of all US bank business loans (see Lown and Morgan, 2006). The dashed dotted line in the figure represents the net percent of banks reporting tightening bank lending standards, measured by the number of loan officers reporting tightening less the number reporting easing divided by the total number reporting. The correlation coefficient between this survey-based measure and the EBP is 0.76 over the full sample period. This co-movement supports the argument that the EBP might provide an adequate proxy for credit supply conditions in the US economy.

3.3. Model selection

Following Altissimo and Corradi (2002), Galvao (2006), and Artis et al. (2007), we employ two model selection criteria to choose between a linear VAR model under the null hypothesis and a TVAR model under the alternative. The threshold γ is not identified and constitutes a nuisance parameter under the null. We therefore use the bounded supWald (BW) and bounded supLM (BLM) statistics, which constitute consistent model selection criteria when a nuisance parameter is present only under the nonlinear alternative.[3] The TVAR model is preferred over the linear VAR if the statistics exceed unity (BW > 1 and, similarly, BLM > 1). This model selection rule ensures that type I and type II errors are asymptotically zero.

Table 1 shows the BW and BLM statistics that guide our model selection between a constant-parameter linear VAR against the threshold-VAR alternative. The table shows the test statistics for each individual equation in the US TVAR-5, the baseline open-economy TVAR-7, and the TVAR-12 comprising all variables. The equation-wise supremum statistics speak unequivocally in favor of the nonlinear TVAR model, which suggests that US credit constraints give rise to significant macro-financial nonlinearities.

[2] The countries included are Argentina, Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, India, Italy, Japan, Korea, Mexico, the Netherlands, Spain, Sweden, and the United Kingdom. We treat the data as an unbalanced panel and aggregate accordingly, as some time series do not stretch back to 1984.

[3] The BW statistic is given by:

$$BW = \frac{1}{2\log(\log(T))}\left(\sup_{\gamma_L \le \gamma \le \gamma_U} T\left(\frac{SSR^{lin} - SSR^{nlin}(\gamma)}{SSR^{nlin}(\gamma)}\right)\right)^{\frac{1}{2}},$$

and the BLM is given by:

$$BLM = \frac{1}{2\log(\log(T))}\left(\sup_{\gamma_L \le \gamma \le \gamma_U} T\left(\frac{SSR^{lin} - SSR^{nlin}(\gamma)}{SSR^{lin}}\right)\right)^{\frac{1}{2}},$$

where SSR^{lin} is the the sum of squared residuals under the linear VAR null, and $SSR^{nlin}(.)$ is the sum of squared residuals under the TVAR alternative hypothesis. The statistics BW and BLM provide the asymptotic bounds on the supremum of the Wald and LM statistics computed over a grid $\gamma_L \le \gamma \le \gamma_U$ of possible values for the threshold γ.

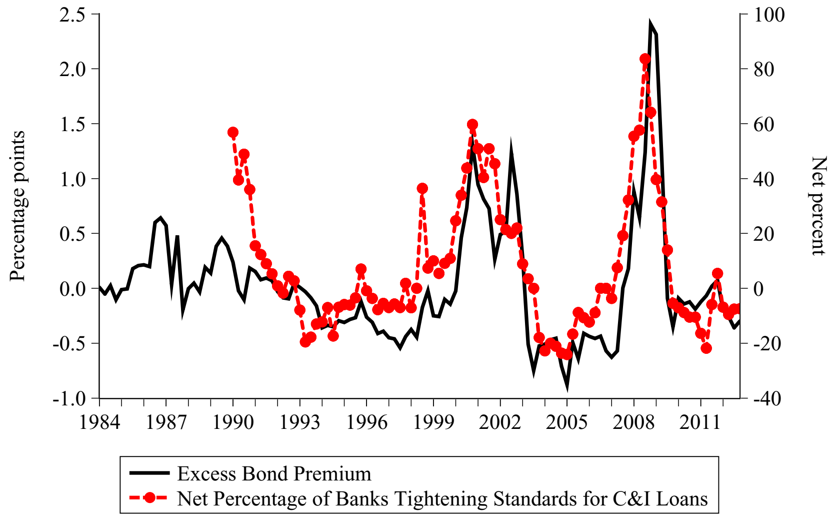

Fig. 1. The EBP and the net percentage of US banks tightening credit standards. *Note*: The solid line depicts the quarterly average of the excess bond premium proposed by Gilchrist and Zakrajsek (2012) (measured on the left axis in percentage points). The dashed dotted line represents the net percentage of US banks tightening loan underwriting standards for commercial and industrial loans to large and middle-market firms, obtained from the Senior Loan Officer Opinion Survey on Bank Lending Practices (measured on the right axis in the net % of banks tightening standards). The sample ranges from 1984Q1 to 2012Q4. The survey-based measure starts in 1990Q1.

Table 1
Model selection criteria.

Equation	US VAR-5		Spillover VAR-7		Spillover VAR-12	
	BW	BLM	BW	BLM	BW	BLM
q_t^*			4.15	3.86	5.54	4.89
π_t^*					5.48	4.86
tra_t					5.60	4.94
q_t	4.17	3.88	4.69	4.28	5.72	5.02
π_t	3.84	3.60	4.13	3.85	5.32	4.75
l_t	4.43	4.08	4.61	4.22	5.88	5.12
i_t^*					6.66	5.62
i_t	4.90	4.44	3.63	3.43	5.02	4.53
oil_t					5.00	4.51
e_t			3.90	3.65	4.89	4.43
vol_t^*					5.73	5.03
ebp_t	4.74	4.32	4.58	4.20	6.00	5.21

Note: The table shows the BW and BLM statistics for each equation of the estimated models. The model specifications are: TVAR-5: $Y_t = [\Delta q_t, \pi_t, \Delta l_t, i_t, ebp_t]$; TVAR-7: $Y_t = [\Delta q_t^*, \Delta q_t, \pi_t, \Delta l_t, i_t, \Delta e_t, ebp_t]$; and TVAR-12: $Y_t = [\Delta q_t^*, \pi_t^*, \Delta trd_t^*, \Delta q_t, \pi_t, \Delta l_t, i_t^*, i_t, \Delta oil_t^*, \Delta e_t, vol_t^*, ebp_t]$. The variables in the different model specifications are: global output growth (Δq_t^*), global inflation (π_t^*), trade (trd_t^*), US output growth (Δq_t), US inflation (π_t), real bank loans (Δl_t), global interest rate (i_t^*), federal funds rate (i_t), oil price (Δoil_t^*), real effective exchange rate (Δe_t), global financial volatility (vol_t^*), and the excess bond premium (ebp_t). The nonlinear TVAR model is chosen over the linear VAR if BW > 1 and, similarly, if BLM > 1.

3.4. Identified credit regimes

Using the EBP as the threshold variable in the TVAR, we estimate an US-specific threshold value $\hat{\gamma}_{US}$ endogenously from the 5-variate model specification for the US economy, estimated with $p_1 = 4$ lags in the normal credit regime and $p_2 = 3$ lags in the tight credit regime. The estimated threshold value equals $\hat{\gamma}_{US} = 0.1004$ percentage points with a delay of $\hat{d} = 1$ month, which amounts to $T^1 = 235$ observations in the normal credit regime and $T^2 = 112$ observations in the tight credit regime. In order to facilitate comparability across different model specifications, $\hat{\gamma}_{US}$ is exogenously held constant across all estimated TVARs. That is, whenever global variables are added to Y_t, the TVAR is re-estimated with $\hat{\gamma}_{US}$. This approach ensures that the identified regimes reflect distressed credit conditions in the US economy, and it amounts to studying the international transmission of EBP shocks in times when constraints bind in US credit markets.

Fig. 2 illustrates the lagged EBP (solid line) together with the estimated threshold (dashed line). The shaded areas correspond

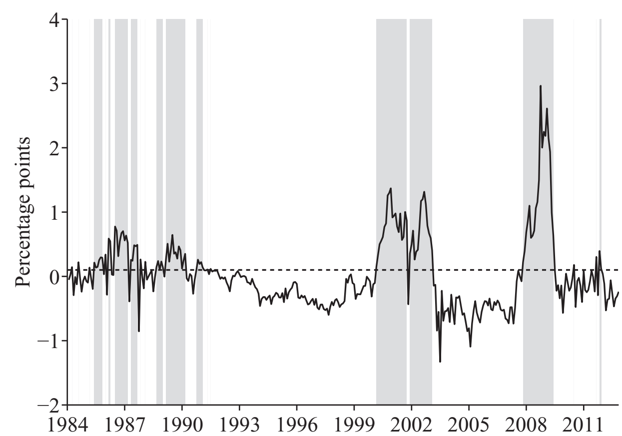

Fig. 2. Credit regimes. *Note*: The solid line depicts the lagged excess bond premium and the dashed line corresponds to the threshold value estimated endogenously from the TVAR-5 model for the US economy ($\hat{\gamma}_{US} = 0.1004$). Tight credit regimes that correspond to periods when the EBP exceeds the threshold are shaded in grey. The sample ranges from January 1984 to December 2012.

to periods when the EBP exceeds the threshold. At a first glance, three major episodes of distress in US banking and credit markets stand out. The first wave of tight credit coincides with the savings and loan crisis of the 1980s and early 1990s (see Federal Deposit Insurance Corporation, 1997, for a historical account of the banking crises in that period). Following a period of relative financial stability during the 1990s, the US economy is again characterized by stringent credit supply conditions at the wake of the new millennium, around the Enron, Y2K, and 9/11 debacles, and the burst of the dotcom bubble. Finally, credit constraints are tight throughout the recent global financial crisis. The tight credit regime associated with the recent crisis covers the 20 months long period from December 2007 until July 2009, which broadly corresponds with the business cycle peak (December 2007) and trough (June 2009) dates published by the NBER for the Great Recession period.

3.5. The regime-specific effects of EBP shocks

Our aim is to investigate the regime-specific effects of EBP shocks on the global economy, conditional on whether the US economy resides in a normal or tight credit regime. Before

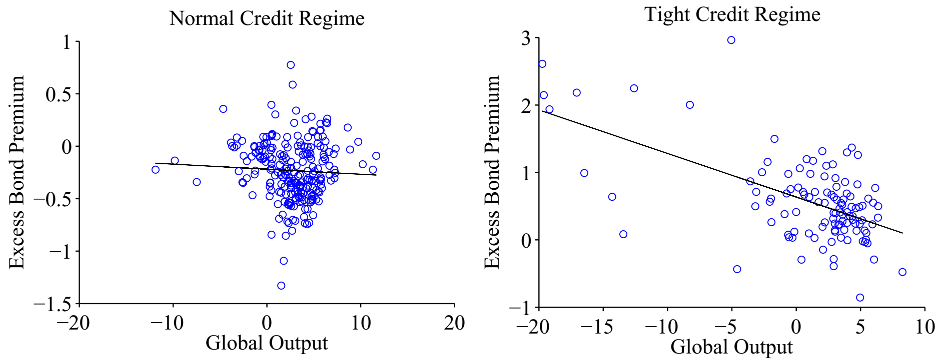

Fig. 3. EBP vs. global output. *Note*: The sample from January 1984 to December 2012 is split according to the threshold $\hat{\gamma}_{US}$. The figure on the left depicts the scatter plot of the EBP and year-on-year growth of global output together with a linear regression line in the normal credit regime. The figure on the right depicts the scatter plot of the EBP and year-on-year growth of global output together with a linear regression line in the tight credit regime.

turning to the structural analysis, we consider a simple graphical assessment of the relationship between the EBP and global output growth. We split the sample according to the threshold $\hat{\gamma}_{US}$, and Fig. 3 shows a scatter plot of the EBP vs. year-on-year changes in global output in each of the two regimes. The relationship between the EBP and global economic activity seems to differ across US credit regimes: high values of the EBP are associated with very low negative growth rates of global output in the tight credit regime, while the link between EBP and global output is much weaker in the normal credit regime.

We now turn to the structural analysis. All open-economy TVAR models are estimated using four lags in each regime, selected according to the criteria proposed by Tsay (1998) and Wong and Li (1998). We compute regime-specific impulse response functions because they enable structural identification of EBP shocks in each regime. This constitutes an important advantage compared to "generalized" impulse response functions occasionally used in conjunction with TVARs. Even though generalized impulse responses enable explicitly modeling switches from one regime to another, they reflect the responses to shocks that have not been orthogonalized, which makes economic interpretation of the shocks difficult. In contrast, regime-specific impulse response functions have a meaningful economic interpretation, they are therefore very informative regarding the effects of EBP shocks, at least before a possible regime switch is likely to change the economy's dynamics. For this reason we only report impulse responses up to 20 months, i.e., the longest identified period of tight credit (the 2007–09 crisis).[4]

Fig. 4 shows the regime-specific impulse responses to a 10 basis point rise in the EBP identified via Cholesky decomposition from the baseline TVAR-7. The first column of Fig. 5 depicts the impulse responses of additional variables added to the TVAR-7 one-at-a-time. We only show the responses of the additional variables, as the impulse responses of the baseline variables are nearly identical across the different model specifications.

Our results indicate that the US economy responds in an asymmetric fashion to an unexpected increase in the EBP when distinguishing between normal and tight credit regimes. In times when credit is readily available, the US economy does not respond significantly to the EBP shock. In contrast, the EBP shock has significant real effects when credit is scarce and the non-financial sector faces difficulties in raising external funds. In the tight credit regime a 10 basis point positive shock to the EBP induces banks to curtail lending, and the volume of bank loans decreases by about 0.7 percentage points 20 months after the shock.[5] The EBP shock is also associated with a downturn in aggregate economic activity. US output contracts by about 0.6 percentage points and consumer prices decline by about 0.15 percentage points one year after the shock, suggesting that binding credit constraints force firms and households to postpone investment and consumption plans until credit conditions improve. The federal funds rate falls by about 10 basis points one year after the EBP shock, implying that the Federal Reserve eases monetary policy amid deteriorating macroeconomic conditions in the US economy.

Crucially, we find that global output contracts significantly by about 0.6 percentage points one year after the EBP shock when credit is scarce in the US, while it does not respond significantly in normal times (see Fig. 4). To assess the economic significance of this result, we decompose the historical values of global output growth into the contributions of EBP shocks and all other orthogonalized shocks. The analytical details of the historical decomposition are presented in the Appendix. Fig. 6(a) depicts the contributions of EBP shocks to (demeaned) global output growth together with the sum of the contributions of all other shocks. In addition, in order to gauge the relative importance of EBP shocks, Fig. 6(b) shows the contribution of EBP shocks to global output growth as a fraction of the contributions of all shocks. The first thing to note is that EBP shocks explain a non-negligible part of movements in global output growth over the entire sample period. A large part of the negative movements in global output growth can be attributed to EBP shocks particularly during the three sub-periods associated with tight US credit market conditions (the mid-1980s, the early 2000s, and between 2007–09). The contribution of EBP shocks to global output growth as a fraction of the contributions of all structural shocks is 45% on average in the tight credit regime, while it equals 18% on average in the normal credit regime, which suggests that EBP shocks are a relatively more important driver of global business cycles during periods of tight credit.

[4] Previous studies that consider regime-specific impulse response functions within regime-switching VARs include Ehrmann et al. (2003), Candelon and Lieb (2013), and Hubrich and Tetlow (2015).

[5] Loans increase slightly in the first three months after the shock, and they decline significantly only ten months after impact. This temporary rise might be puzzling at first sight, however, it does not preclude *new* loan issuances from decreasing on impact, and instead it may rather reflect drawdowns of credit lines that had been granted prior to the shock, as argued by Ivashina and Scharfstein (2010) and Adrian et al. (2012). According to Ivashina and Scharfstein (2010): "commercial and industrial (C&I) loans reported on the aggregate balance sheet of the U.S. banking sector actually rose by about $100 billion from September to mid-October 2008, from a base of about $1.5 trillion [...] However, we show that this increase was not driven by an increase in new loans, but rather by an increase in drawdowns by corporate borrowers on *existing* credit lines (prior commitments by banks to lend to corporations at prespecified rates and up to prespecified limits)" (see page 320). As a result, credit may lag rather than lead the business cycle, as already pointed out by Bernanke et al. (1996).

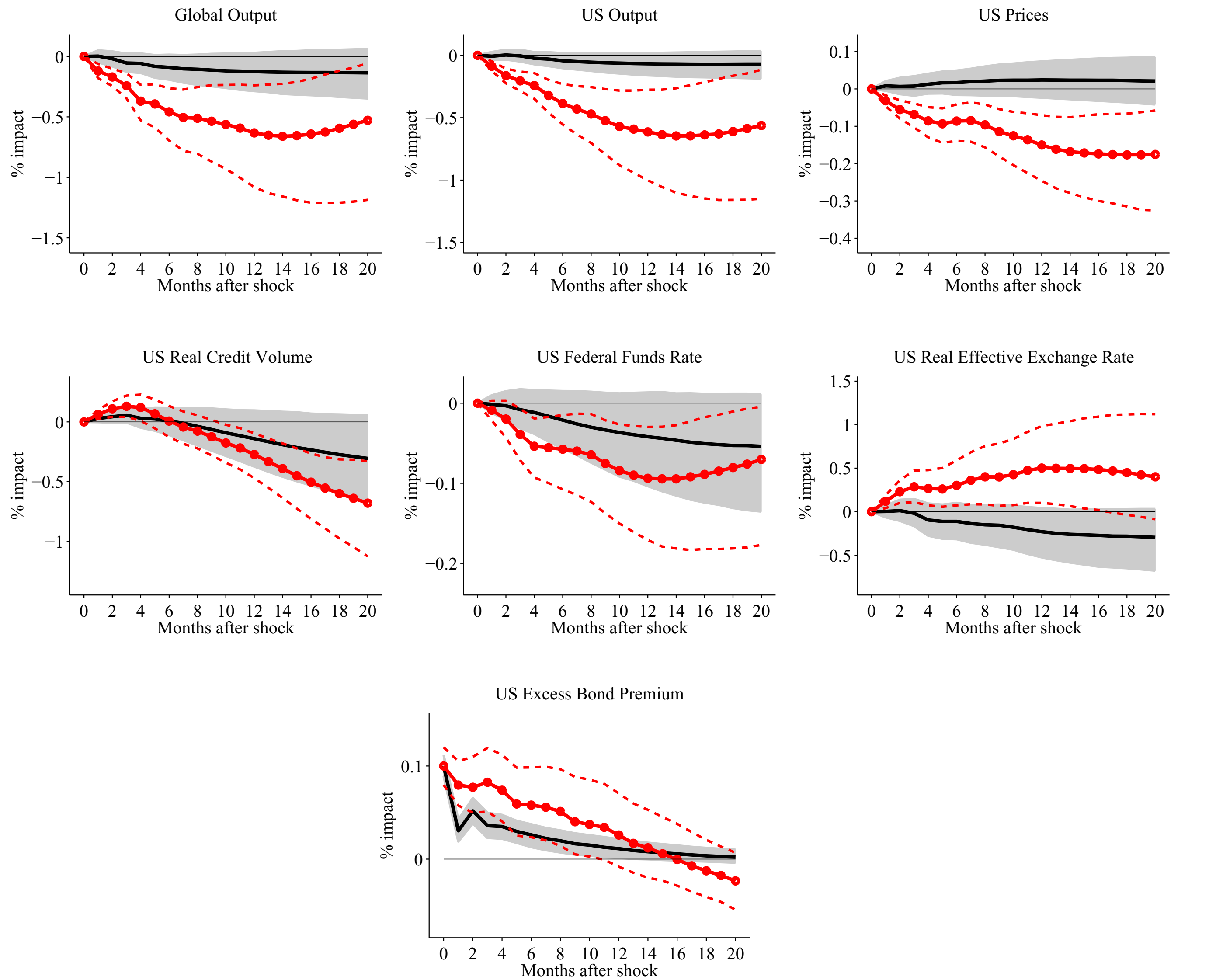

Fig. 4. Responses to an EBP shock – Baseline model estimates. *Note*: Regime-specific impulse responses to an unexpected 10 basis point rise in the EBP from the TVAR-7 depicted for 20 months. The estimated TVAR model is given in Eq. (1). The TVAR-7 includes in the estimation order: $Y_t = [\Delta q_t^*, \Delta q_t, \pi_t, \Delta l_t, i_t, \Delta e_t, ebp_t]$. EBP shocks are identified via a Cholesky decomposition of the regime-specific reduced form variance-covariance matrices. The black solid lines are the median impulse responses from the TVAR model in the normal credit regime with shaded areas representing 90% confidence bands based on 1000 draws. The red dotted lines are the median impulse responses from the TVAR model in the tight credit regime with dashed lines representing 90% confidence bands. (For interpretation of the references to color in this figure legend, the reader is referred to the web version of this article.)

The EBP shock does not have statistically significant effects on global variables in the normal credit regime. On the contrary, it has a significant adverse impact on all remaining global variables in the tight credit regime. Specifically, the EBP shock is followed by a real appreciation of the US dollar, which is in line with exchange rate movements in the 2007–08 global financial crisis, during which the US dollar appreciated against numerous currencies. Fratzscher (2009) attributes the lion's share of the US dollar appreciation during the crisis to US-specific shocks and a flight-to-safety by investors. In the same vein, Gourinchas et al. (2012) ascribe the exchange rate dynamics in the 2007–08 crisis to the role played by the US as a "global insurer" during financial crises.[6] This safe haven property of the US dollar is labeled as the "dollar trap" by Prasad (2014).[7] The real appreciation of the US dollar is associated with a deterioration in the competitive position, the EBP shock is therefore followed by a 2 percentage point reduction in US trade with the rest of the world one year after the EBP shock. Moreover, global consumer prices decline by 0.1 percentage points one year after impact, which together with the contraction in global production point toward a reversal in global demand for industrial commodities. Consequently, the world price of oil falls significantly, reaching its trough at 2.75 percentage points one year after the EBP shock. Finally, the EBP shock is followed by a significant increase in worldwide stock market volatility and by a reduction of 7.5 basis points in the global interest rate within 1.5 years after the shock.

[6] In a TVAR-7 augmented with the yields on 3-month US Treasury Bills we have found that T-Bill yields decline following an EBP shock in the tight credit regime, which provides independent corroborating evidence for the global insurer/flight-to-safety argument.

[7] We have performed a robustness check by estimating the TVAR with data that end in November 2007 (not shown here), which confirms the tendency of the REER to appreciate after an EBP shock in the tight credit regime, suggesting that this is not a particular feature of the Great Recession period.

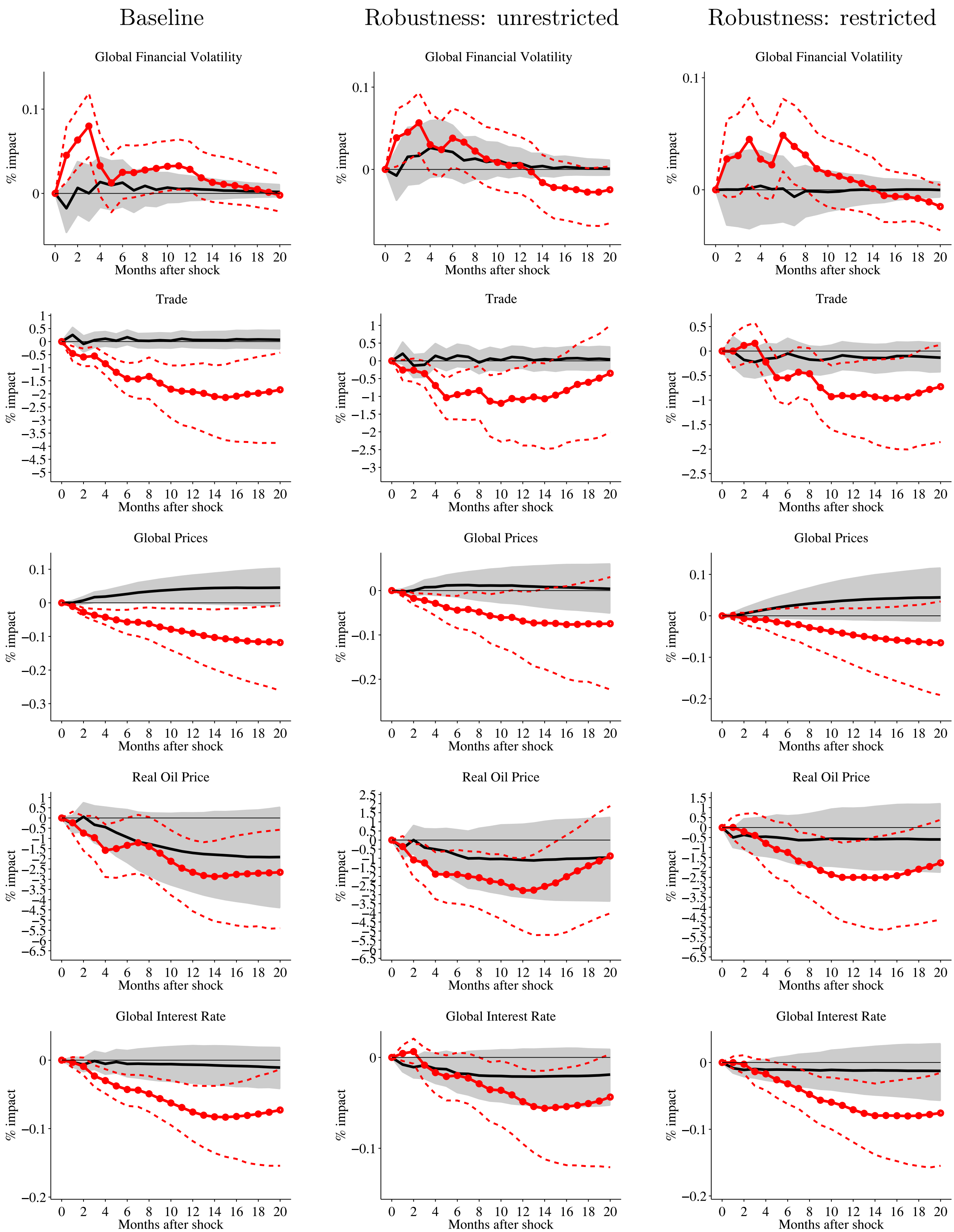

Fig. 5. Responses to an EBP shock – Additional variables. *Note*: Impulse responses to an EBP shock of additional variables added to the baseline TVAR-7 specification. The first column depicts the impulse responses of additional variables included into the TVAR-7 one-at-a-time. The second column depicts the impulse responses of additional variables included into the TVAR-7 jointly. The third column depicts the impulse responses of additional variables included into the TVAR-7 jointly, and estimated from a model specification with VAR coefficient restrictions imposed according to a subset VAR modeling strategy developed by Brueggemann (2004) to find all entries in the VAR coefficient matrices that do not differ significantly from zero.

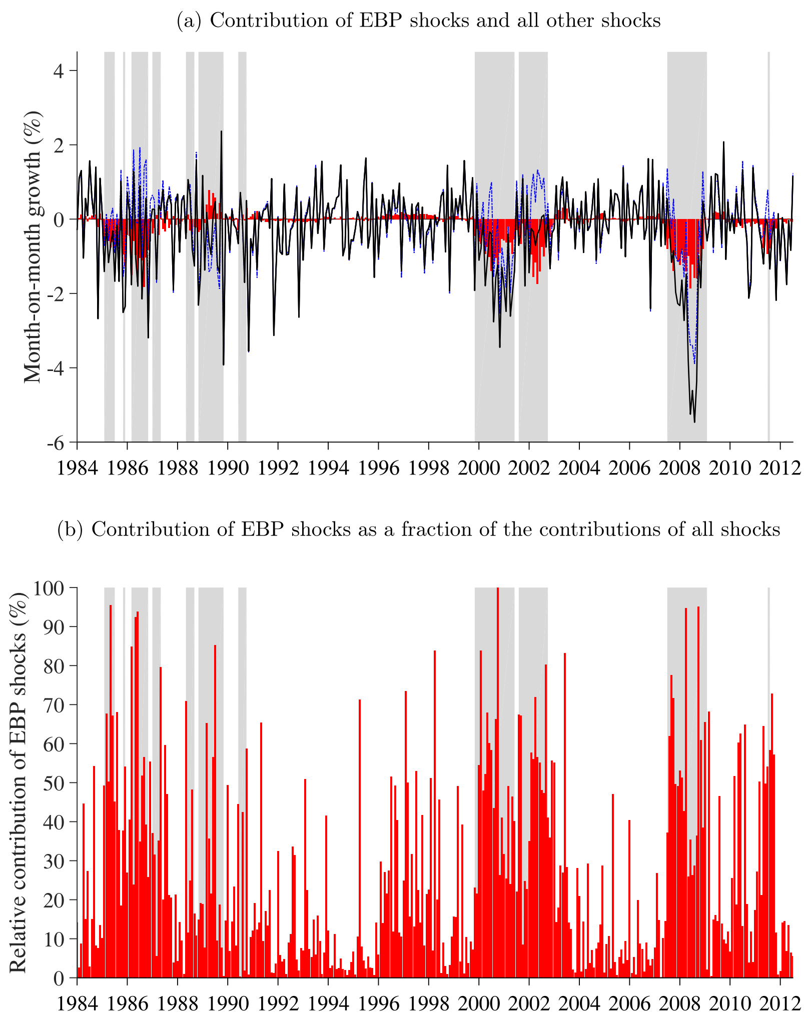

Fig. 6. Historical decomposition of global output growth. *Note*: Panel (a): Red bars: historical contribution of EBP shocks to global output growth. Blue dashed line: contribution of all other shocks (excluding EBP shocks). Black solid line: deviation of global output growth from its deterministic component (historical contribution of all shocks). Grey shaded areas indicate the identified tight credit regimes. Panel (b): Historical contribution of EBP shocks to global output growth relative to the contributions of all shocks. Sample: January 1984–December 2012. (For interpretation of the references to color in this figure legend, the reader is referred to the web version of this article.)

To conclude, the crucial role of the US in international financial markets implies that US financial shocks generate global spillovers that depend nonlinearly on the severity of credit constraints in the US economy. US financial shocks reverberate through the global economy during an US credit crunch, while they have little impact in normal times. This key empirical result highlights the importance of accounting for financial frictions when modeling international macroeconomic fluctuations, as suggested by, e.g., Krugman (2008), Devereux and Yetman (2010), Devereux and Sutherland (2011), Olivero (2010), Kollmann et al. (2011), Dedola and Lombardo (2012), and van Wincoop (2013). Moreover, it complements the theoretical works by Mendoza (2010), Bianchi (2011), Brunnermeier and Sannikov (2014), and Perri and Quadrini (2014) with empirical evidence on regime-dependent dynamics that arise due to occasionally binding credit constraints.

The results presented thus far suggest that a surprise increase in the EBP is followed by macroeconomic dynamics consistent with an adverse shock to the supply of credit: a rise in the EBP is associated with a decline in the risk-bearing capacity of the financial sector, which induces banks to cut back lending, leading to a downturn in aggregate economic activity. However, our results may in fact capture something different than a credit supply channel. We explore this possibility in the next section.

3.6. Do EBP shocks capture a credit supply channel?

One possibility is that part of what we assign to be a credit supply shock might actually be a monetary policy shock that works through a risk-taking channel. This issue arises as empirical evidence suggests that financial risk appetite moves in tandem with monetary policy. In particular, Bekaert et al. (2013) and Bekaert and Hoerova (2014) show that option-implied stock market volatility – measured by the VIX index – can be decomposed into a component that reflects uncertainty and another that captures a variance risk premium associated with risk taking in financial markets. Moreover, Bekaert et al. (2013) document that this variance risk premium is causally affected by monetary policy. Consequently, the question arises whether the asymmetric effects of US financial shocks on the global economy hold up when the effects of monetary policy are explicitly purged from credit supply disturbances traced by the EBP.

To address this question, we consider an alternative identification scheme of US financial shocks, which isolates changes in the supply of credit from other macro and financial shocks and in particular from monetary policy induced changes in credit supply. We impose a combination of zero and sign restrictions on the estimated impulse responses proposed by Peersman (2012). He distinguishes between three different types of bank lending shocks:

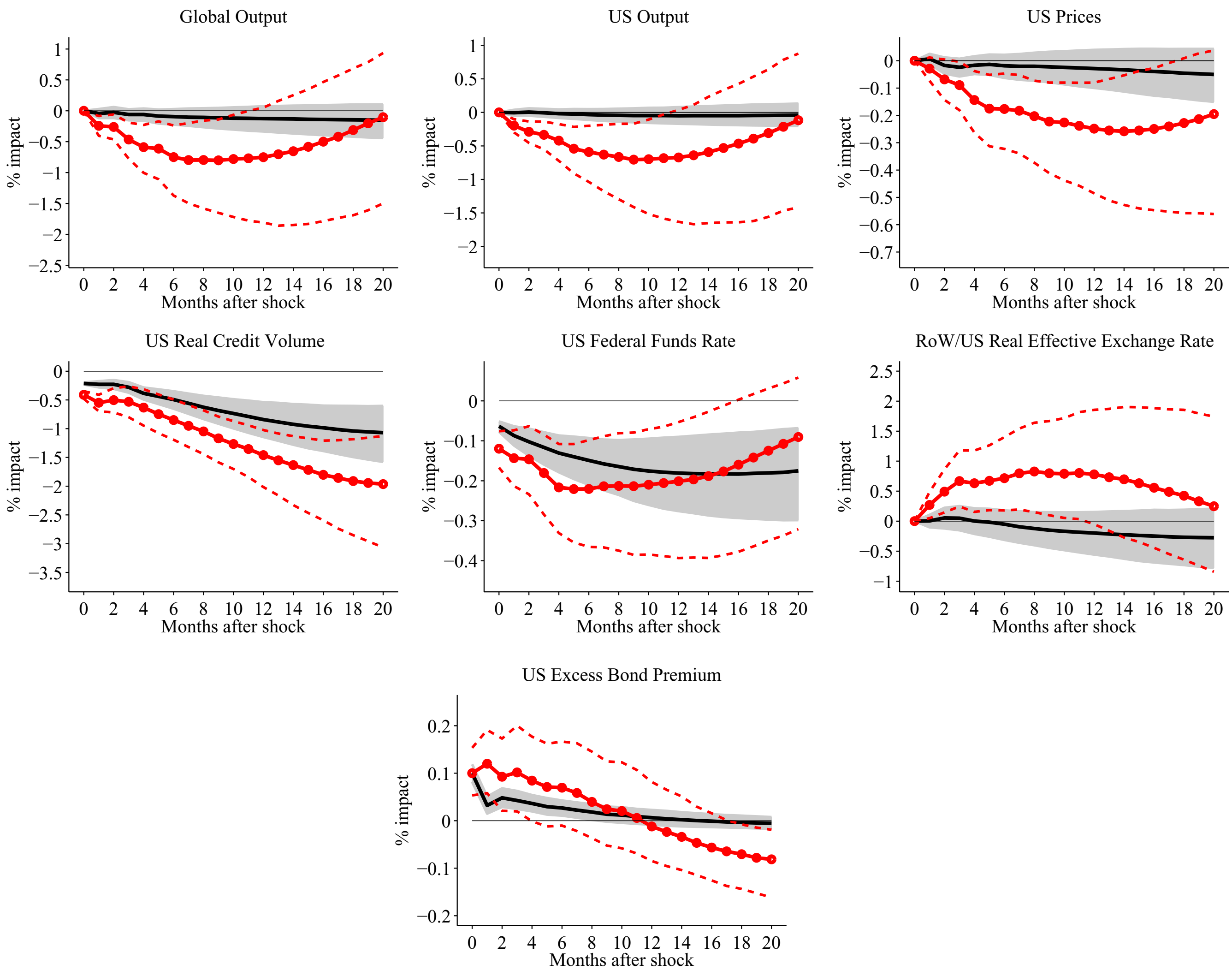

Fig. 7. Responses to an EBP shock – TVAR-7 identified with sign restrictions. *Note*: Regime-specific impulse responses to an unexpected 10 basis point rise in the EBP from the TVAR-7 depicted for 20 months. The estimated TVAR model is given in Eq. (1). The TVAR-7 includes in the estimation order: $Y_t = [\Delta q_t^*, \Delta q_t, \pi_t, \Delta l_t, i_t, \Delta e_t, ebp_t]$. EBP shocks are identified via zero and sign restrictions on the estimated impulse responses. The black solid lines are the median impulse responses from the TVAR model in the normal credit regime with shaded areas representing 90% confidence bands based on 1000 draws. The red dotted lines are the median impulse responses from the TVAR model in the tight credit regime with dashed lines representing 90% confidence bands. (For interpretation of the references to color in this figure legend, the reader is referred to the web version of this article.)

a loan demand shock; a loan supply shock which is caused by changes in monetary policy; and a loan supply shock that is due to changes in the credit supplied by banks independently of a shift in monetary policy, labeled as a “lending multiplier shock”. We impose the zero and sign restrictions associated with this latter shock in order to isolate independent credit supply disturbances. In particular, we assume that an unexpected tightening in the supply of credit is associated with a rise in the EBP ($ebp_t \geq 0$) accompanied by a decline in loan volume ($l_t \leq 0$) and a drop in the federal funds rate ($i_t \leq 0$). These sign restrictions are assumed to hold for at least six months following the shock. The contemporaneous impact on US and global output and on consumer prices is restricted to zero. Finally, other shocks hitting the economy are assumed to display a different pattern.

The sign restrictions approach is implemented by using the method by Rubio-Ramirez et al. (2010). It is well known that sign restrictions do not allow us to achieve unique identification of shocks. Hence, we draw rotation matrices until 500 of them yield shocks consistent with our sign restrictions. We adopt the median target approach to pick among the 500 rotations the one which yields impulse responses that are closest to the median response (see Fry and Pagan, 2011). Fig. 7 depicts the structural impulse response functions obtained with this particular rotation matrix. The results paint a broadly similar picture to those obtained with the baseline Cholesky scheme. In the tight credit regime the EBP shock is followed by a significant decline in both US and global output, a fall in prices, and an appreciation of the REER. Bank loans and the federal funds rate decline on impact, consistent with the sign restrictions imposed. Remarkably, the median effects are in general stronger compared to the baseline results. In line with the baseline specification, in the normal credit regime the shock does not have a significant impact on output, prices, and the REER. The only notable difference comes from the response of credit and the federal funds rate which – again, in line with the sign restrictions – respond with a significant decline on impact also in normal times, albeit the effects are weaker than in the tight credit regime. On balance, we may thus conclude that the asymmetric effect of EBP shocks on the global economy does not hinge upon the structural identification scheme.

Another potential concern is that the EBP may fluctuate with changes in the level of uncertainty in financial markets, and as uncertainty is known to have real effects (see, e.g., Bloom, 2009), we may capture an uncertainty shock instead of a credit supply shock. To refute this interpretation of the data, we augment our baseline

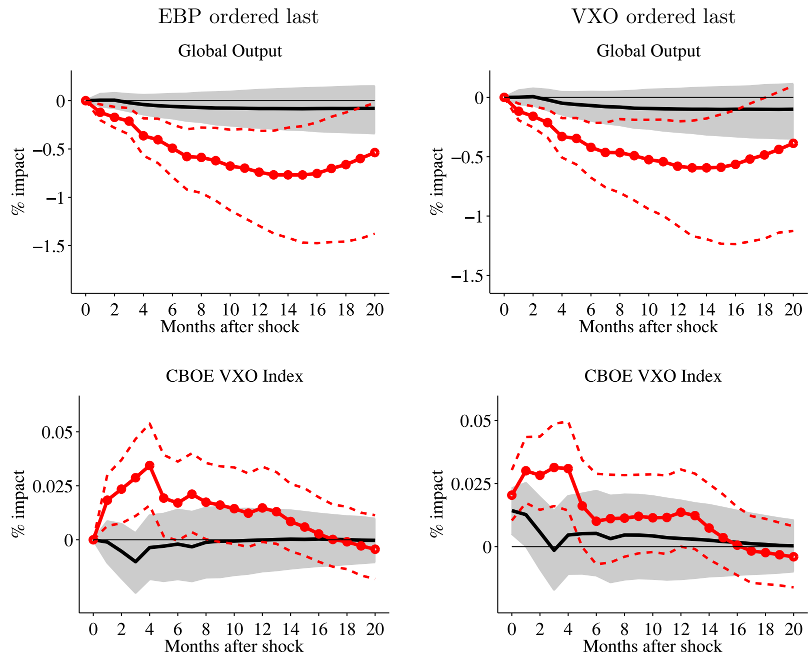

Fig. 8. Responses to an EBP shock – Selected variables from TVAR-8 with VXO index. *Note*: Regime-specific impulse responses to an unexpected 10 basis point rise in the EBP from the TVAR-7 augmented with the CBOE VXO index depicted for 20 months. The estimated TVAR model is given in Eq. (1). The model includes in the estimation order: $Y_t = [\Delta q_t^*, \Delta q_t, \pi_t, \Delta l_t, i_t, \Delta e_t, vxo_t, ebp_t]$ and $Y_t = [\Delta q_t^*, \Delta q_t, \pi_t, \Delta l_t, i_t, \Delta e_t, ebp_t, vxo_t]$, respectively. EBP shocks are identified via a Cholesky decomposition of the regime-specific reduced form variance-covariance matrices. The black solid lines are the median impulse responses from the TVAR model in the normal credit regime with shaded areas representing 90% confidence bands based on 1000 draws. The red dotted lines are the median impulse responses from the TVAR model in the tight credit regime with dashed lines representing 90% confidence bands. (For interpretation of the references to color in this figure legend, the reader is referred to the web version of this article.)

TVAR-7 with the Chicago Board of Options Exchange VXO index of percentage implied volatility based on trading of S&P 100 options – a proxy for uncertainty also used by Bloom (2009). Fig. 8 shows the regime-specific responses of global output and of the VXO index to an EBP shock from the model specification that includes the VXO. We estimate this model from January 1986 onward because the VXO has a price history that dates back to that date. Since both the VXO and the EBP are financial variables, their ordering is somewhat ambiguous. Therefore, we first place the VXO above the EBP, and subsequently this order is reversed. In the normal credit regime the VXO index barely moves after the EBP shock. In contrast, in the tight credit regime the EBP shock is associated with a significant rise in the VXO. By construction, the VXO responds to the EBP shock only with a lag when ordered above the EBP, whereas, not surprisingly, it jumps on impact when ordered last. Yet, the reaction of our key variable of interest – global output – to the EBP shock remains qualitatively unchanged in spite of the inclusion of the VXO index, even when the VXO is allowed to respond contemporaneously to the EBP shock. This suggests that the effects of EBP shocks on global output are unlikely to stem from uncertainty shocks (in as far as the VXO adequately captures uncertainty; see Bekaert and Hoerova, 2014).

3.7. Robustness exercises

We perform several additional robustness exercises. Fig. 9 depicts the regime-specific responses of global output to an EBP shock obtained from the baseline model specification together with the impulse response functions from five different robustness checks.[8] Even though the estimates may somewhat vary quantitatively across these robustness checks, the asymmetric impact of EBP shocks on the global economy turns out to be a salient feature across all different settings considered.

In the baseline model specification global variables are obtained as GDP-weighted averages of the country-specific variables. We control for robustness to the weights used in the aggregation by considering two alternative weighting schemes, based on bilateral trade and financial positions.[9] The second and third plot of Fig. 9 depict the impulse responses from the TVAR with global output weighted according to the trade and financial weighting scheme, respectively. Our key result remains unchanged irrespective of the aggregation method employed.

To ensure that our findings are not dominated by the recent Great Recession period, we re-estimate the baseline TVAR-7 using data that end in November 2007. Thus, we exclude the tight credit episode associated with the 2007–09 crisis and the period thereafter. The fourth plot of Fig. 9 depicts the impulse responses of global output from this subsample robustness exercise. Not surprisingly, the confidence bands around the sub-sample estimates are wider as less observations are available for estimation. Yet, our main conclusions continue to hold.

The re-estimation of each TVAR-8 may deliver model estimates inconsistent with one another. Thus, for the sake of robustness, we jointly estimate the full system with 12 endogenous variables. The TVAR-12 requires the estimation of a large number of parameters, which quickly erodes the degrees of freedom for estimation. Therefore, besides unrestricted model estimation, we also employ

[8] We do not show the responses of variables other than global output to save space, however, these are in line with the baseline results. Detailed results are available from the authors upon request.

[9] Trade weights are constructed as $w_i^{Tra} = (EX^{\mathrm{US\,to}\,i} + IM^{\mathrm{US\,from}\,i})/(\Sigma_i EX^{\mathrm{US\,to}\,i} + \Sigma_i IM^{\mathrm{US\,from}\,i})$ following Frankel and Rose (1998), where $EX^{\mathrm{US\,to}\,i}$ denotes US exports to country i and $IM^{\mathrm{US\,from}\,i}$ denotes US imports from country i. Following Imbs (2004), financial weights are constructed as $w_i^{Fin} = |(NFA_i/GDP_i) - (NFA_{US}/GDP_{US})|$, using the data from Lane and Milesi-Ferretti (2007). NFA_i denotes the net foreign asset position in country i. The weight w_i^{Fin} take high values for countries that have diverging external positions with respect to the US, as such countries are more likely to lend and borrow from the US according to Imbs (2004). Trade as well as financial weights are normalized and sum to 1.

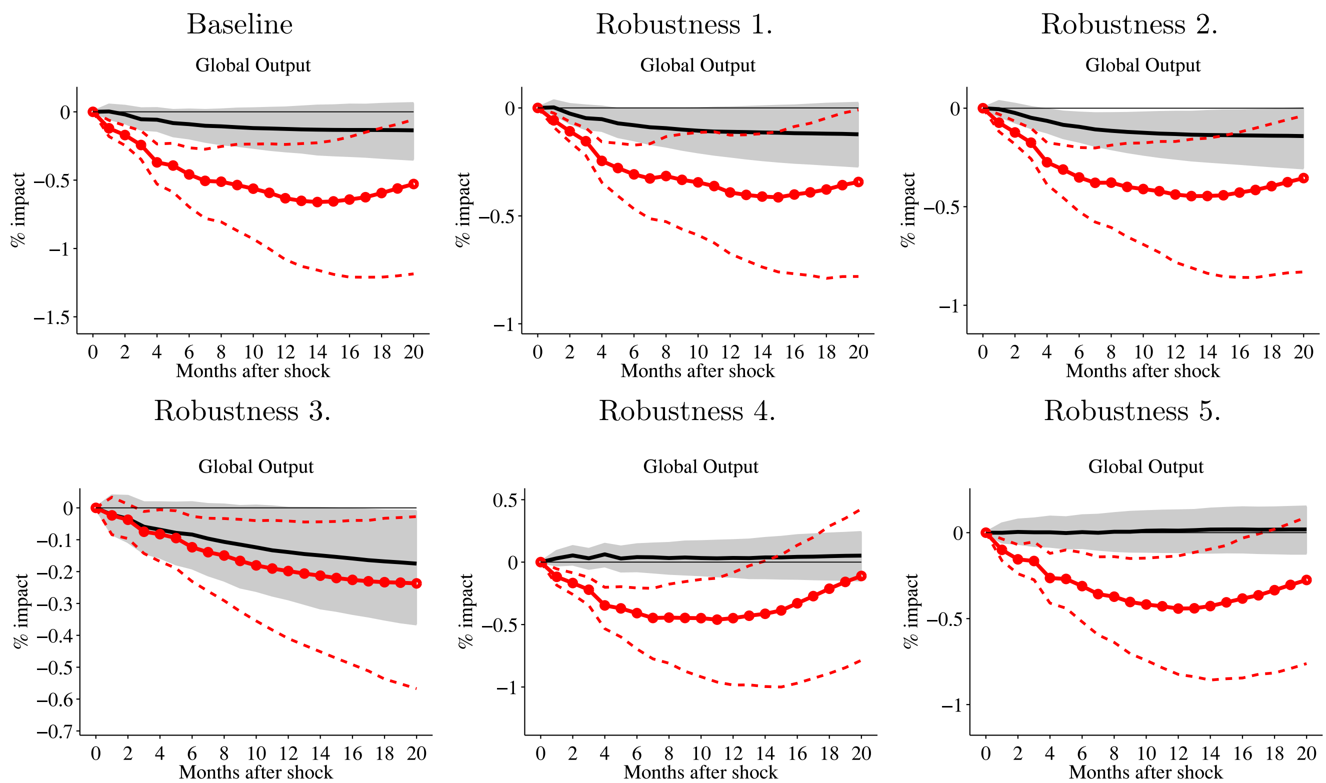

Fig. 9. Robustness checks. *Note*: Baseline: impulse response function of global output to an EBP shock from the baseline TVAR-7 model specification. Robustness 1.: global variables are obtained as a trade-weighted average of country-specific variables. Robustness 2.: global variables are obtained as a finance-weighted average of country-specific variables. Robustness 3.: TVAR-7 estimated on data that end in Nov. 2007. Robustness 4.: estimation results from the unrestricted full model with 12 endogenous variables (TVAR-12). Robustness 5.: estimates of the TVAR-12 model with zero coefficient restrictions.

a subset VAR modeling strategy designed to find all entries in the VAR coefficient matrices that do not differ significantly from zero, along the lines of the general-to-specific model reduction approach advocated by Hendry (1995). In particular, we reduce the dimension of each regime-specific VAR model separately using the system test procedure developed by Brueggemann (2004).[10] This procedure helps us to determine which variables play a less significant role in the VAR dynamics, and it enables us to constrain the model along those dimensions. The fifth (sixth) plot of Fig. 9 shows the impulse response of global output to an EBP shock from the unrestricted (restricted) TVAR-12. Furthermore, the second (third) column of Fig. 5 shows the impulse responses of the remaining global variables in the unrestricted (restricted) TVAR-12. The joint model estimates qualitatively resemble those obtained from the baseline model specification, and our main conclusions remain broadly unchanged.[11] Crucially, the asymmetric impact of EBP shocks on the global economy is confirmed across all model specifications.

4. Conclusion

A consensus seems to emerge from theoretical models that financial frictions are central to understanding the nonlinearities observed during financial crisis episodes. Moreover, theoretical studies have shown that financial frictions lead to an amplification of cross-border shocks, and structural models featuring such frictions provide a more realistic picture of international macroeconomic fluctuations. However, most empirical studies on macro-financial linkages resort to linear models that fail to account for the nonlinear amplification mechanisms implied by these theoretical considerations. There is an equally limited empirical literature that investigates the relation between financial frictions and global spillovers. This paper aims to fill these gaps.

We model the dynamics of the US economy jointly with global macroeconomic and financial variables using a threshold vector autoregressive model. This model captures regime-dependent dynamics conditional on the tightness of credit supply conditions in the US economy, measured by the excess bond premium proposed by Gilchrist and Zakrajsek (2012) which enters our model as the threshold variable. Transition from a state of unconstrained access to credit to a regime characterized by binding financial constraints arises endogenously in this framework whenever the excess bond premium crosses an estimated threshold value.

Using data for the period from 1984 to 2012, we identify three main periods of distress in US banking and credit markets. The first tight credit episode takes place during the savings and loan crisis of the 1980s and early 1990s. The second episode occurs in the early 2000s, around the Enron, Y2K, and 9/11 debacles and following the burst of the dotcom bubble. Finally, the 2007–09 financial crisis is identified as the most recent credit crunch.

We study the nonlinear effects of an US financial shock on the global economy in the TVAR model using regime-specific impulse response functions. We find that credit constraints amplify business cycle fluctuations within as well as across economies. Upon distinguishing between normal and tight credit regimes in the US economy, we uncover a clear asymmetry in the impulse responses to an US financial shock. The US financial sector absorbs the shock when borrowers have unconstrained access to credit, and there are

[10] In the system test procedure, the VAR is estimated jointly, and in each model reduction step one variable is deleted if its corresponding t-ratio is less than a certain critical value τ. The VAR is then re-estimated with this new zero restriction imposed. If all absolute t-ratios exceed the critical value, the algorithm stops. We set $\tau = \sqrt{2}$ according to the AIC criterion.

[11] The specification with coefficient restrictions delivers insignificant global price responses. Financial factors might thus play a relatively less important role in global price dynamics.

no aggregate economic consequences. In contrast, the US financial shock is followed by a significant contraction both in the US and in the global economy when the US resides in a tight credit regime. These results are consistent with an international dimension of the US financial accelerator mechanism. However, when interpreting our results it is important to keep in mind that there might be global financial shocks which co-move with US financial shocks and also affect the global economy. Therefore, an extension of our analysis to other parts of the world, such as the euro area, may constitute a promising avenue of future research.

Appendix

This appendix provides descriptions of the ML estimation of the TVAR model and computation of regime-specific impulse responses and historical decompositions within this framework.

To estimate the reduced form parameters of the TVAR model, we follow the approach described in Galvao (2006). This entails computing the constrained MLE for $B^r(L)$ and Σ_u^r for all possible values of delay d and threshold value γ on an equally spaced grid of ebp_{t-d}. For a given d and γ, the MLE are the OLS estimators given by:

$$\begin{bmatrix} B_1^r \\ B_2^r \\ \vdots \\ B_{p^r}^r \end{bmatrix}' = \left(\left(\left(\begin{bmatrix} Y_{t-1} \\ Y_{t-2} \\ \vdots \\ Y_{t-p^r} \end{bmatrix}' D_t^r \right)' \left(\begin{bmatrix} Y_{t-1} \\ Y_{t-2} \\ \vdots \\ Y_{t-p^r} \end{bmatrix}' D_t^r \right) \right)^{-1} \left(\begin{bmatrix} Y_{t-1} \\ Y_{t-2} \\ \vdots \\ Y_{t-p^r} \end{bmatrix}' D_t^r \right)' Y_t \right)'$$

where $D_t^1 = I(ebp_{t-d} < \gamma)$ and $D_t^2 = I(ebp_{t-d} \geq \gamma)$ are indicator functions in regime $r = 1, 2$. The estimated residuals are obtained as: $\hat{u}_t^r = Y_t D_t^r$ - $([Y'_{t-1}, Y'_{t-2}, ..., Y'_{t-p^r}]D_t^r)[\hat{B}_1^r, \hat{B}_2^r, ..., \hat{B}_{p^r}^r]$. Finally, the MLEs for the covariance matrices are $\hat{\Sigma}_u^r = 1/T^r \sum_{t=1}^{T^r} \hat{u}_t^r \hat{u}_t^{r'}$, where $\sum_{r=1}^{2} T^r = T$.

The unconstrained parameter estimates are then obtained by solving the following optimization problem:

$$(\hat{\gamma}, \hat{d}) = \min_{\substack{\gamma_L \leq \gamma \leq \gamma_U \\ 1 \leq d \leq d_{max}}} \left(\sum_{r=1}^{2} \frac{T^r}{2} \log(|\Sigma_u^r|) \right).$$

We consider a maximum delay of six months ($d_{max} = 6$), and the search region is restricted to a minimum of 15% of the observations in each regime, such that γ_L is the 15%th percentile and γ_U is the 85%th percentile of the empirical distribution of ebp_{t-d}.

Conditional on the regime r, the TVAR model reduces to a piecewise linear VAR. The regime-specific structural shocks relate to the reduced-form errors according to $\varepsilon_t^r = A^r u_t^r \sim (0, \Sigma_\varepsilon^r = A^r \Sigma_u^r A^{r'})$, and they can be recovered using the Cholesky decomposition of the reduced form error covariance matrix Σ_u^r. The moving average representation of this model is given by:

$$Y_t = \sum_{i=1}^{\infty} \Psi_i^r \varepsilon_{t-i}^r,$$

where the $\Psi_i^r = \Phi_i^r (A^r)^{-1}$ contain the regime-specific structural impulse responses, and the Φ_i^r matrices can be obtained recursively as

$$\Phi_i^r = \sum_{j=1}^{i} \Phi_{i-j}^r B_j^r, \quad i = 1, 2, \ldots,$$

where B_j^r are regime-specific coefficient matrices ($j = 1, \ldots, p^r$, $\Phi_0^r = I_N$, and $B_j^r = 0$ for $j > p^r$).

We examine the contribution of EBP shocks to global output growth by adopting the historical decomposition proposed by Burbidge and Harrison (1985) to the TVAR setup. Let global output growth constitute the jth element of Y_t and let the EBP shock constitute the nth structural shock ($n = 1, \ldots N$). The contribution of the EBP shock to global output growth in period t, given the starting values $Y_0, \ldots, Y_{-p+1}$, is then obtained as:

$$Y_{j,t}^{(n)} = \sum_{i=0}^{t-1} \psi_{j,n,i}^{r_i} \varepsilon_{n,t-i}^{r_i} + \beta_{j,1}^{(t)} Y_0 + \cdots + \beta_{j,p}^{(t)} Y_{-p+1}$$

where $\psi_{j,n,i}^{r_i}$ is the (j, n)th element of $\Psi_i^{r_i}$ and r_i denotes the regime r in period i. Furthermore, $\beta_{j,l}^{(t)}$ is the jth row of $B_l^{(t)}$ $(l = 1, \ldots, p)$, where $[B_1^{(t)}, \ldots, B_{p-1}^{(t)}, B_p^{(t)}]$ consist of the first N rows of the companion matrix $\mathbf{B}^t$ raised to the power t computed as:

$$\mathbf{B}^t = \begin{bmatrix} B_1^{r_t} & \cdots & B_{p-1}^{r_t} & B_p^{r_t} \\ I_N & & 0 & 0 \\ & \ddots & & \vdots \\ 0 & & I_N & 0 \end{bmatrix} \times \begin{bmatrix} B_1^{r_{t-1}} & \cdots & B_{p-1}^{r_{t-1}} & B_p^{r_{t-1}} \\ I_N & & 0 & 0 \\ & \ddots & & \vdots \\ 0 & & I_N & 0 \end{bmatrix} \times \cdots \times \begin{bmatrix} B_1^{r_1} & \cdots & B_{p-1}^{r_1} & B_p^{r_1} \\ I_N & & 0 & 0 \\ & \ddots & & \vdots \\ 0 & & I_N & 0 \end{bmatrix}.$$

Finally, the relative contribution of EBP shocks is obtained as $|Y_{j,t}^{(n)}|/|(\sum_{k=1;k\neq n}^{N} Y_{j,t}^{(k)})|$.

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