BANK FOR INTERNATIONAL SETTLEMENTS

BIS Working Papers
No 855

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April 2020

JEL classification: E32, E44, E52.

Keywords: liquidity trap, effective lower bound, monetary transmission.

This publication is available on the BIS website (www.bis.org).

ISSN 1020-0959 (print)
ISSN 1682-7678 (online)

DOES THE LIQUIDITY TRAP EXIST?*

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Abstract. The liquidity trap is synonymous with ineffective monetary policy. The common wisdom is that, as the short-term interest rate nears its effective lower bound, monetary policy cannot do much to stimulate the economy. However, central banks have resorted to alternative instruments, such as QE, credit easing and forward guidance. Using state-of-the-art estimates of the effects of monetary policy, we show that monetary easing stimulates output and inflation, also during the period when short-term interest rates are near their lower bound. These results are consistent across the United States, the euro area and Japan.

I. Introduction

In this article, we compare the effects of monetary policy shocks in situations where the short-term interest rate is near its effective lower bound (ELB) with their effects when this is not the case, using a Bayesian structural vector autoregression (SVAR) framework. We define "ELB times" as periods when both the level and the standard deviation of short-term interest rates are close to zero. The level of the nominal short-term interest rate may even be negative, as has been the case in the euro area, Japan and several other countries. The ELB times extend from January 2009 to December 2015 for the United States economy, since July 2012 for the euro area and since January 1996 for Japan. Estimates of the effect of monetary

Date: March 28, 2020.

Key words and phrases: liquidity trap, effective lower bound, monetary transmission.
JEL Classification: E44, E52.

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policy shocks are compared with their estimates in periods when short-term interest rates can "normally" fluctuate in positive territory. We call these later times the "normal times."

Monetary policy shocks are identified through a combination of standard sign and zero restrictions on the impulse response functions (IRFs) and the systematic component of the monetary policy equation following Arias, Caldara, and Rubio-Ramírez (2019). Our models include industrial production, the CPI, a short-term interest rate, a monetary aggregate and a measure of financial market sentiment (corporate credit spread for the euro area and the United States and stock prices for Japan) at monthly frequency. Because the monetary policy rate is bounded below during the ELB times, for the short-term interest rate, we chose a two-year interest rate to be consistent across periods. Our sign and zero restrictions on the IRFs imply that the monetary policy shock contemporaneously impacts the short-term rate the credit spread in a similar direction and the money measure in the opposite direction. At the same time, industrial production and prices do not react simultaneously to monetary policy shocks. The restrictions on the systematic component of the monetary policy equation directly follow Arias, Caldara, and Rubio-Ramírez (2019) and will be described below.

To simplify our analysis as much a possible, for each country we will split our sample into ELB and normal times and run two different SVARs. Our main result is that monetary policy shocks have remained effective during ELB times. The result is robust across countries. We find that the signs of the IRFs of the model's variables to a monetary policy (MP) shock do not change when the economy moves from normal to ELB times. We also find that the contribution of monetary policy shocks to unexpected movements of output and prices is still relevant during ELB times. For the United States, the contribution of MP shocks to the variance of output is even more important during the periods when the movements of the short-term rate were restricted by the ELB. That is not the case for the other two economies.

The analysis of the effects of monetary policy when rates are low is important because interest rates may remain at or near the ELB in the decade ahead of us. As we write this paper, the Fed has reduced its interest rate by 75 basis points from the 2018 peak in the interest rate cycle. Neither the ECB nor the Band of Japan (BoJ) are likely to lift the short-term rate into positive territory for several quarters. It is therefore necessary to measure the effect of monetary policy shocks when rates are near the ELB.

In many DSGE models used for the analysis of monetary policy, the ELB prevents the economy from adjusting. The logic of the argument is straightforward. If the reaction function of the central bank is the main stabilisation mechanism of the economy and a sequence of negative demand shocks brings interest rates to their lower bound, the economy loses its main adjustment mechanism and it may diverge into a deflation spiral (see Krugman, 1998; Eggertsson et al., 2003; Adam and Billi, 2006, 2007). Among recent empirical analyses implied by DSGE models, Gust, Herbst, López-Salido, and Smith (2017) estimate that the ELB accounts for 30 percent of the 2008-2009 contraction in United States output. Using a DSGE model Evans, Fisher, Gourio, and Krane (2016) also show that monetary policy is less effective at the ELB. Beyond monetary policy, DSGE models also imply that the ELB has important implications for the effects of fiscal policy, which would be much larger if, unlike in normal times, interest rates fail to increase following a stimulus (see Christiano, Trabandt, and Walentin, 2011; Woodford, 2011; Coenen, Straub, and Trabandt, 2013).

But this argument takes models with rational expectations a bit too literally. In the real world, monetary policy can still affect output and prices when the policy rate is at the ELB since several other rates can still adjust up or down. In particular, the central bank can lift asset prices. The set of measures focused on affecting asset prices have been labelled as unconventional monetary policies. Although the yields on these bonds are tied to the short-term rate of the central bank via arbitrage, they also depend on expectations for future short-term rates, term-, liquidity- and credit risk-premia, all of which can be influenced by the central bank's asset purchases, lending policy, and commitment to future short-term rates. Hence, monetary policy shocks may spur output and inflation during ELB times. They would be expected to do so if unconventional monetary policies work effectively.

Our results show that these policies have stimulated economic activity and inflation during ELB times. We show that a short-term rate stuck near zero does not prevent a central bank from spurring credit growth if it seeks to do so. In particular, our results accord with the view of the monetary transmission in Bernanke and Gertler (1995). Promoting credit and releasing liquidity constraints matter more than lowering the level of interest rates (see Swanson and Williams, 2014; Gertler and Karadi, 2015, for similar results). Our results also support Friedman and Schwartz; during a deep recession, credit maybe unbounded and, hence, unconventional monetary measures have indeed sustained the effectiveness of monetary policy.

This paper proceeds as follows. Section II relates our contributions to the literature. Section III introduces the methodology. Section IV presents the main results. Section V concludes.

II. Literature review

Since Hicks (1937), the liquidity trap has been synonymous with the inability of monetary policy to stimulate activity when interest rates are near zero. In the words of Krugman (1998) "John Hicks, in introducing both the IS-LM model and the liquidity trap, identified the assumption that monetary policy is ineffective, rather than the assumed downward inflexibility of prices, as the central difference between Mr. Keynes and the classics." Most of the New Keynesian model-based analysis under the ELB confirms Hicks' conclusions. However, Eggertsson (2008) has challenged this view. In his view, the central bank can in principle stimulate activity at the ELB by committing itself to keep lowering rates in the future when the economy has recovered (see Eggertsson et al., 2003; Adam and Billi, 2006, 2007, for similar arguments). This future stimulus would lift the economy and inflation today under relatively specific conditions. In particular, if the central bank can credibly commit its future interest rate moves, the expectations channel is powerful enough to prevent a liquidity trap. As predicted by Hicks (1937), short of such commitment, given that the interest rate cannot decline far below zero, the New Keynesian model implies that economy may diverge into a liquidity and deflationary trap; real interest rates keep increasing as prices decline and nominal interest rates cannot adjust to derail this spiral. Also challenging the view implied by simple New Keynesian models, DelNegro, Eggertsson, Ferrero, and Kiyotaki (2017) introduce liquidity frictions into an otherwise standard New Keynesian model and show that, once the nominal interest rate reaches the lower bound, the central bank can halt the deflationary spiral by increasing its liquidity provision.

Empirical studies are more polarized. A comprehensive review of the expanding literature that assesses the effects of non-conventional monetary policy goes much beyond the ambition of this paper. But the interested reader can find a comparison of the estimated effects of non-conventional monetary policies, as available in 2016, in Table 7 of Borio and Zabai (2018), and a review of no less than 150 papers on the topic, most of which were published in the last five years, in Neri and Siviero (2019).

Some papers stress that the undermining of monetary policy by the ELB is very costly. Chung, Laforte, Reifschneider, and Williams (2012) claim that the cost of the ELB has been underestimated because most assessments consider the mainly benign demand shocks that characterized the Great Moderation period. They also stress that non-conventional monetary policies are only a poor substitute for conventional monetary policies; they are much less effective than the short-term interest rate in stimulating the economy. Boeckx, Dossche, and Peersman (2014), Filardo and Nakajima (2018), Borio and Gambacorta (2017), Borio and Hofmann (2017), Hesse, Hofmann, and Weber (2018) and Burriel and Galesi (2018) find similar results. Campbell, Evans, Fisher, and Justiniano (2012) argue that the ELB is costly because, although unconventional monetary policy instruments may work, they do not work as well as conventional monetary policy, and they may exhibit decreasing returns as we approach the ELB also for long-term interest rates. Borio and Gambacorta (2017) and Borio and Hofmann (2017) also stress that low rates for long and extending easing via unconventional monetary policies may inflict costly side effects on financial stability.

Others stress instead that monetary policy has remained effective at the ELB. Using macroeconomic data from eight advanced economies, Gambacorta, Hofmann, and Peersman (2014) show that expanding the balance sheet of the central bank has stimulated output and inflation. Using a panel VAR for eight advanced economies, they show that expanding the assets held by central banks brings down the VIX and stimulates output and inflation. In a paper more closely related to ours, Panizza and Wyplosz (2018) estimate the effects of unconventional monetary policy in the euro area, Japan the United Kingdom and the United States. They compare the effects of shocks with shadow rates and the effects of QE announcements using both SVARs and local projection methods. The results are mixed. While asset purchase programs have stimulated output and inflation consistently, the effects of monetary easing as measured by identified shocks of a shadow rate vary across countries and estimation approaches. Their results for the United Kingdom and the United States are consistent with the estimates of Weale and Wieladek (2016). Finally, Koeda (2019), Kimura and Nakajima (2016) confirm previous estimates showing that changes in the stance of monetary policy via QE have stimulated output in Japan.

Other papers have estimated a shadow rate to translate the effects of non-conventional monetary policy measures into an “equivalent” short-term negative rate. Wu and Xia (2016) showed, for instance, that the US shadow rate seems to have reached as low as -4 percent. Such a low level is synonymous with powerful non-conventional stimuli. Estimates for other countries indicate that shadow rates have reached similar low values. Related Mouabbi and Sahuc (2019) challenge results in Gust, Herbst, López-Salido, and Smith (2017) since, using a DSGE New Keynesian model and, using the shadow rate as a proxy for unconventional monetary policies, they show that monetary policy has remained a valid instrument during ELB periods. Likewise, Swanson and Williams (2014) stress in particular that long-term rates can still adjust, and the fiscal multiplier need not increase as much as in models where all rates are stuck at zero. Gertler and Karadi (2015) use high-frequency data to highlight that the mechanisms in Eggertsson (2008) may be empirically relevant. Altavilla, Giannone, and Lenza (2014) use the OMT announcements to show that unconventional monetary policy can be effective at the ELB. Finally, Debortoli, Gali, and Gambetti (2018) show both indirect and direct evidence for the stabilizing effect of monetary policy in the United States. The time series properties of US macroeconomic variables have hardly changed with the ZLB. And shocks identified as impacting long-term rates negatively and GDP and prices positively, as would a monetary stimulus whether the economy is at the ZLB or not, have similar effects before and after 2008.

Our approach complements these previous studies. We estimate the effects of shocks to the money supply by applying an agnostic identification that combines a minimal set of sign and zero restrictions in a model that captures transmission through a financial accelerator as in Bernanke and Gertler (1995). A monetary easing shock raises money, it reduces the short-term interest rate and credit spreads (or stimulates stock prices in the case of Japan). We make no assumption on its effects on output and inflation. We then estimate this model separately for the G3 so that we do not impose homogeneous adjustments across the board. Finally, our estimates infer changes in the stance of monetary policy directly from the application of identifying restrictions to the joint dynamics of observable variables without relying on a shadow rate, which would require a separate estimation.

III. The Methodology

III.1. **Baseline Specification.** Consider the SVAR with the general form, as in Rubio-Ramírez, Waggoner, and Zha (2010)

$$\mathbf{y}_t'\mathbf{A}_0 = \sum_{\ell=1}^{p} \mathbf{y}_{t-\ell}'\mathbf{A}_\ell + \mathbf{c} + \boldsymbol{\varepsilon}_t' \text{ for } 1 \leq t \leq T, \tag{1}$$

where $\mathbf{y}_t$ is an $n \times 1$ vector of endogenous variables, $\boldsymbol{\varepsilon}_t$ is an $n \times 1$ vector of exogenous structural shocks, $\mathbf{A}_\ell$ is an $n \times n$ matrix of parameters for $0 \leq \ell \leq p$ with $\mathbf{A}_0$ invertible, $\mathbf{c}$ is a $1 \times n$ vector of parameters, p is the lag length, and T is the sample size. The vector $\boldsymbol{\varepsilon}_t$, conditional on past information and the initial conditions $\mathbf{y}_0, ..., \mathbf{y}_{1-p}$, is Gaussian with mean zero and covariance matrix $\mathbf{I}_n$, the $n \times n$ identity matrix. The model described in Equation (1) can be compactly written as

$$\mathbf{y}_t'\mathbf{A}_0 = \mathbf{x}_t'\mathbf{A}_+ + \boldsymbol{\varepsilon}_t' \text{ for } 1 \leq t \leq T, \tag{2}$$

where $\mathbf{A}_+' = \begin{bmatrix}\mathbf{A}_1' & \cdots & \mathbf{A}_p' & \mathbf{c}'\end{bmatrix}$ and $\mathbf{x}_t' = \begin{bmatrix}\mathbf{y}_{t-1}' & \cdots & \mathbf{y}_{t-p}' & 1\end{bmatrix}$ for $1 \leq t \leq T$. The dimension of $\mathbf{A}_+$ is $m \times n$, where $m = np + 1$.

III.2. **Data.** We will use data for three countries; the United States, the euro area and Japan. We use a monthly frequency from January 1990 to December 2015 for the United States, from January 1999 to December 2018 for the euro area, and from January 1980 to December 2018 for Japan. We employ five endogenous variables: $y_t = [ip_t, p_t, m_t, d_t, r_t]'$; ip_t is an index of industrial production; p_t is an index of consumer price; r_t is a monetary policy instrument; m_t is a monetary aggregate. We also include a financial variable: for the United States and the euro area, we employ a credit spread, sp_t following the recent work by Gertler and Karadi (2015). For Japan, we use the Nikkei stock market index, $stock_t$. Data sources are presented in Appendix A. For estimation, we measure all variables in log units, except for variables expressed in rates. Thus, industrial production, prices, and money are measured in log-levels, while the spread and the policy instrument are measured in levels.

As a measure for the monetary policy instrument, we consider a two-year nominal interest rate. The choice of a medium-term interest rate is justified by the fact that typical short-term interest rates do not move during ELB times. Of course, we could have chosen a different measure for the policy stance between the two samples, one shorter-term for the normal times and one longer-term for liquidity trap times, but this would make the comparison more difficult. As shown in Figure 1, the strong correlation between the three-month and two-year rates allows us to employ medium-term rates while keeping a very good approximation of the monetary policy stance. Hanson and Stein (2015), Gertler and Karadi (2015), and Gilchrist, López-Salido, and Zakrajšek (2015) also consider changes in the two-year nominal interest rate to reflect the policy stance. As a robustness check, we have also estimated VARs with a short-term interest rate for the normal times and show that the main results remain unchanged. Alternatively, we could have approximated the short-term nominal interest rates by the shadow rate, which would incorporate the non-standard monetary policies implemented during the Great Recession. However, such a rate is subject to a wide range of uncertainty, which would complicate the interpretation of results.

III.3. **Effective Lower Bound.** We will split our sample into two sub-periods and compare the IRFs across both periods. The first period will be called normal times while the second will be called ELB times. For this reason, we need now to define what we mean by ELB in the data. ELB is when the short-term interest rates are flat and close to zero and their standard deviation is at a historical minimum.

Figure 1 reports the historical path of the level (top panel) and the standard deviation (bottom panel) of the three-month and the two-year interest rate for the United States, euro area, and Japan. The vertical bar indicates when, according to this definition, the ELB begins: January 2009 for the United States, July 2012 for the euro area, and January 1996 for Japan. Table 1 summarizes these sub-periods. Given that the three-month interest rate becomes nearly constant, by definition, during ELB times, we use a two-year interest rate to track changes in the stance of monetary policy. Indeed, central banks engage in forward guidance and asset purchase programs to modify the stance of their policies. Note that for the United States, our sample ends in December 2015, as thereafter the Federal Reserve raised its fed funds rate by 0.25 percentage points. So the ELB time was then over.

III.4. **The Reduced-form Representation and the Identification Problem.** The reduced-form representation implied by Equation (2) is

$$\mathbf{y}_t' = \mathbf{x}_t'\mathbf{B} + \mathbf{u}_t' \text{ for } 1 \leq t \leq T, \tag{3}$$

where $\mathbf{B} = \mathbf{A}_+\mathbf{A}_0^{-1}$, $\mathbf{u}_t' = \boldsymbol{\varepsilon}_t'\mathbf{A}_0^{-1}$, and $\mathbb{E}\left[\mathbf{u}_t\mathbf{u}_t'\right] = \boldsymbol{\Sigma} = (\mathbf{A}_0\mathbf{A}_0')^{-1}$. The matrices $\mathbf{B}$ and $\boldsymbol{\Sigma}$ are the reduced-form parameters, while $\mathbf{A}_0$ and $\mathbf{A}_+$ are the structural parameters. Following Rubio-Ramírez, Waggoner, and Zha (2010), two of the parameters $(\mathbf{A}_0, \mathbf{A}_+)$ and $(\tilde{\mathbf{A}}_0, \tilde{\mathbf{A}}_+)$

are observationally equivalent if and only if they have the same reduced-form representation. Hence, we have an identification problem. To solve the identification problem, one often imposes sign and/or zero restrictions on either the structural parameters or some function of the structural parameters, like the IRFs. We describe our identification restrictions in the next section.

III.5. **Identification Scheme.** To identify the monetary policy shock in the model described in Equation (2), we impose sign and zero restrictions on IRFs of monetary policy shocks, as well as on the systematic component of monetary policy. The former follows works by Faust (1999), Uhlig (1999), Canova and Nicolo (2002), Rubio-Ramírez, Waggoner, and Zha (2010), and Arias, Rubio-Ramírez, and Waggoner (2018), and the latter was recently proposed by Arias, Caldara, and Rubio-Ramírez (2019).

First, we discuss our sign and zero restrictions on IRFs. Table 2 summarizes our sign and zero restrictions on the contemporaneous IRFs. Regarding the sign restrictions, we assume that a contractionary (i.e., negative) monetary policy shock lowers the policy rate but increases the monetary aggregate. Such a restriction allows us to capture the so-called liquidity effect. In addition, we impose that a monetary policy shock also leads to a decline in the credit spread. This restriction is mostly motivated by the theory of the financial accelerator in Bernanke and Gertler (1995) and Bernanke, Gertler, and Gilchrist (1999). The decline of interest rates results in a strengthening of borrower balance sheets (via the "balance sheet" channel) as their interest expenses fall and asset values rise. As a consequence, the capacity of firms to raise funds externally increases, which reduces the external finance premium. Regarding the zero restrictions, we impose that the production sector, namely ip_t and p, does not respond contemporaneously to monetary policy shocks. In other words, monetary policy shocks have only lagged effects on the production sector. These restrictions can be justified by the work of Leeper, Sims, and Zha (1996), who explains that "there are planning processes involved in changing output or in changing the prices of final goods". This set of restrictions is defined as Restriction I. Thus, in the language of Rubio-Ramírez, Waggoner, and Zha (2010), we have that

$$\mathbf{S}_1 = \begin{pmatrix} 0 & 0 & 1 & 0 & 0 \\ 0 & 0 & 0 & -1 & 0 \\ 0 & 0 & 0 & 0 & 1 \end{pmatrix}, \mathbf{Z}_1 = \begin{pmatrix} 1 & 0 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 & 0 \end{pmatrix} \text{ and } \mathbf{F}(\mathbf{A}_0, \mathbf{A}_+) = (\mathbf{A}_0^{-1})'$$

and $s_1 = 3$ and $z_1 = 2$ are the rank of $\mathbf{S}_1$ and $\mathbf{Z}_1$ respectively. Clearly $\mathbf{S}_j$ and $\mathbf{Z}_j$ for $j \geq 2$ are empty because we only identify the monetary policy shock.

Second, we now discuss the restrictions on the systematic part of monetary policy. Without loss of generality, the first equation of the SVAR,

$$\mathbf{y}_t'\mathbf{a}_{0,1} = \sum_{\ell=1}^{p} \mathbf{y}_{t-\ell}'\mathbf{a}_{\ell,1} + \boldsymbol{\varepsilon}_{1,t} \text{ for } 1 \leq t \leq T, \tag{4}$$

is the monetary policy equation, where $\boldsymbol{\varepsilon}_{1,t}$ denotes the first entry of $\boldsymbol{\varepsilon}$, $\mathbf{a}_{\ell,1}$ denotes the first column of $\mathbf{A}_\ell$ for $0 \leq l \leq \nu$, and $\mathbf{a}_{\ell,ij}$ denotes the (i,j) entry of $\mathbf{A}_\ell$ and describes the systematic component of monetary policy. Let us re-write monetary policy Equation (4), abstracting from lag variables, as

$$r_t = \psi_{ip} ip_t + \psi_p p_t + \psi_m m_t + \psi_{sp} sp_t + \sigma \varepsilon_{1,t},$$

where $\psi_{ip} = -\mathbf{a}_{0,51}^{-1}\mathbf{a}_{0,11}$, $\psi_p = -\mathbf{a}_{0,51}^{-1}\mathbf{a}_{0,21}$, $\psi_m = -\mathbf{a}_{0,31}^{-1}\mathbf{a}_{0,41}$, $\psi_{sp} = -\mathbf{a}_{0,51}^{-1}\mathbf{a}_{0,11}$, and $\sigma = \mathbf{a}_{0,51}^{-1}$. Equipped with this representation of the monetary policy equation, we now impose restrictions directly on some structural parameters. More specifically, Restrictions II and III are given as follows.

Restriction II. *The contemporaneous reaction of the policy rate to output, prices, money and stock market is positive, i.e.,* $\psi_y, \psi_p, \psi_m > 0, \psi_{stock} > 0$.

Restriction III. *The contemporaneous reaction of the policy rate to credit spread is negative, i.e.,* $\psi_{sp} < 0$.

To sum up, Restrictions I, II, and III form the set of all restrictions that identify monetary policy shocks.

III.6. **The Prior and Posterior Distributions.** We employ Bayesian tools to make inferences. We will be interested in making independent draws from the normal-generalized-normal posterior distribution over the structural parameterization conditional on the sign and zero restrictions described in Section III.5. Equation (2) represents the SVAR in terms of this structural parameterization, which is characterized by $(\mathbf{A}_0, \mathbf{A}_+)$. Arias, Rubio-Ramírez, and Waggoner (2018) denote the normal-generalized-normal distribution by $NGN(\nu, \mathbf{\Phi}, \mathbf{\Psi}, \mathbf{\Omega})$, where $\nu \leq n$ is a scalar, $\mathbf{\Phi}$ is an $n \times n$ symmetric and positive definite matrix, $\mathbf{\Psi}$ is an $m \times n$ matrix, and $\mathbf{\Omega}$ is an $m \times m$ symmetric and positive definite matrix.

As explained in Arias, Rubio-Ramírez, and Waggoner (2018), it would be easier to independently draw from uniform-normal-inverse-Wishart distribution over the orthogonal reduced-form parameterization conditional on the sign and zero restrictions and then transform the draws to the structural parameterization.[1]

The orthogonal reduced-form parameterization is characterized by the reduced-form parameters $\mathbf{B}$ and $\mathbf{\Sigma}$ together with an orthogonal matrix $\mathbf{Q}$ and it is given by the following equation

$$\mathbf{y}_t' = \mathbf{x}_t'\mathbf{B} + \boldsymbol{\varepsilon}_t'\mathbf{Q}'h(\mathbf{\Sigma}) \text{ for } 1 \leq t \leq T,$$

where the $n \times n$ matrix $h(\mathbf{\Sigma})$ is any decomposition of the covariance matrix $\mathbf{\Sigma}$ satisfying $h(\mathbf{\Sigma})'h(\mathbf{\Sigma}) = \mathbf{\Sigma}$. We will take h to be the Cholesky decomposition, although any differentiable decomposition would do. Arias, Rubio-Ramírez, and Waggoner (2018) denote uniform-normal-inverse-Wishart distribution over the orthogonal reduced-form parameterization by $UNIW(\nu, \mathbf{\Phi}, \mathbf{\Psi}, \mathbf{\Omega})$. This distribution implies a normal-inverse-Wishart, denoted by $NIW(\nu, \mathbf{\Phi}, \mathbf{\Psi}, \mathbf{\Omega})$, over the reduced-form parameters and an uniform distribution over the orthogonal matrix conditional on the reduced-form parameters.

[1] Baumeister and Hamilton (2015) directly draws in the structural parameterization. This is a very interesting and novel approach since the rest of the literature (including us) works in the orthogonal reduced-form parameterization. While working in the structural parameterization has clear advantages, mainly being able to define priors densities directly on economically interpretable structural parameters, this approach uses a Metropolis-Hastings algorithm to make the draws. Hence, this approach could become computationally inefficient, as compared with ours, in larger models.

The mapping between $(\mathbf{A}_0, \mathbf{A}_+)$ and $(\mathbf{B}, \mathbf{\Sigma}, \mathbf{Q})$ by

$$f_h(\mathbf{A}_0, \mathbf{A}_+) = (\underbrace{\mathbf{A}_+\mathbf{A}_0^{-1}}_{\mathbf{B}}, \underbrace{(\mathbf{A}_0\mathbf{A}_0')^{-1}}_{\mathbf{\Sigma}}, \underbrace{h((\mathbf{A}_0\mathbf{A}_0')^{-1})\mathbf{A}_0}_{\mathbf{Q}}).$$

By a direct computation, it is easy to see that $h((\mathbf{A}_0\mathbf{A}_0')^{-1})\mathbf{A}_0$ is an orthogonal matrix. The function f_h is invertible, with the inverse defined by

$$f_h^{-1}(\mathbf{B}, \mathbf{\Sigma}, \mathbf{Q}) = (\underbrace{h(\mathbf{\Sigma})^{-1}\mathbf{Q}}_{\mathbf{A}_0}, \underbrace{\mathbf{B}h(\mathbf{\Sigma})^{-1}\mathbf{Q}}_{\mathbf{A}_+}).$$

If the prior distribution over the orthogonal reduced-form parameters is $UNIW(\bar{\nu}, \bar{\mathbf{\Phi}}, \bar{\mathbf{\Psi}}, \bar{\mathbf{\Omega}})$ and then the posterior distribution over the orthogonal reduced-form parameters is $UNIW(\tilde{\nu}, \tilde{\mathbf{\Phi}}, \tilde{\mathbf{\Psi}}, \tilde{\mathbf{\Omega}})$, where

$$\tilde{\nu} = T + \bar{\nu},$$

$$\tilde{\mathbf{\Omega}} = (\mathbf{X}'\mathbf{X} + \bar{\mathbf{\Omega}}^{-1})^{-1},$$

$$\tilde{\mathbf{\Psi}} = \tilde{\mathbf{\Omega}}(\mathbf{X}'\mathbf{Y} + \bar{\mathbf{\Omega}}^{-1}\bar{\mathbf{\Psi}}),$$

$$\tilde{\mathbf{\Phi}} = \mathbf{Y}'\mathbf{Y} + \bar{\mathbf{\Phi}} + \bar{\mathbf{\Psi}}'\bar{\mathbf{\Omega}}^{-1}\bar{\mathbf{\Psi}} - \tilde{\mathbf{\Psi}}'\tilde{\mathbf{\Omega}}^{-1}\tilde{\mathbf{\Psi}},$$

for $\mathbf{Y} = [\mathbf{y}_1 \ \cdots \ \mathbf{y}_T]'$ and $\mathbf{X} = [\mathbf{x}_1 \ \cdots \ \mathbf{x}_T]'$.

Arias, Rubio-Ramírez, and Waggoner (2018) use the following algorithm to make independent draws from the $NGN(\tilde{\nu}, \tilde{\mathbf{\Phi}}, \tilde{\mathbf{\Psi}}, \tilde{\mathbf{\Omega}})$ posterior over the structural parameterization conditional on the sign and zero restrictions.[2]

Algorithm 1. *The following algorithm independently draws from the $NGN(\tilde{\nu}, \tilde{\mathbf{\Phi}}, \tilde{\mathbf{\Psi}}, \tilde{\mathbf{\Omega}})$ distribution over the structural parameterization conditional on the sign and zero restrictions described in Section III.5.*

(1) *Draw $(\mathbf{B}, \mathbf{\Sigma})$ independently from the $NIW(\tilde{\nu}, \tilde{\mathbf{\Phi}}, \tilde{\mathbf{\Psi}}, \tilde{\mathbf{\Omega}})$ distribution.*

(2) *For $1 \leq j \leq n$ draw $\mathbf{x}_j \in \mathbb{R}^{n+1-j-z_j}$ independently from a standard normal distribution and set $\mathbf{w}_j = \mathbf{x}_j / \parallel \mathbf{x}_j \parallel$.*

[2]The reader should notice that the parameters of the posterior over the structural parameterization are equal to the parameters of the posterior over the orthogonal reduced-form parameterization. This is not necessary. In the case that the researcher wants to use different parameterization across posterior, the importance weight can be modified (see Arias, Rubio-Ramírez, and Waggoner, 2018, for details).

(3) *Define* $\mathbf{Q} = [\mathbf{q}_1 \ \cdots \ \mathbf{q}_n]$ *recursively by* $\mathbf{q}_j = \mathbf{K}_j \mathbf{w}_j$ *for any matrix* $\mathbf{K}_j$ *whose columns form an orthonormal basis for the null space of the* $(j-1+z_j) \times n$ *matrix*[3]

$$\mathbf{M}_j = \begin{bmatrix} \mathbf{q}_1 & \cdots & \mathbf{q}_{j-1} & (\mathbf{Z}_j \mathbf{F}(f_h^{-1}(\mathbf{B}, \mathbf{\Sigma}, \mathbf{I}_n)))' \end{bmatrix}'.$$

(4) *Set* $(\mathbf{A}_0, \mathbf{A}_+) = f_h^{-1}(\mathbf{B}, \mathbf{\Sigma}, \mathbf{Q})$.

(5) *If* $(\mathbf{A}_0, \mathbf{A}_+)$ *satisfies the sign restrictions,*

$$\mathbf{S}_1 \mathbf{F}(\mathbf{A}_0, \mathbf{A}_+) > \mathbf{0},$$

then set its importance weight to

$$\frac{|\det(\mathbf{A}_0)|^{-(2n+m+1)}}{v_{(g \circ f_h)|\mathcal{Z}}(\mathbf{A}_0, \mathbf{A}_+)},$$

where $(\mathbf{B}, \mathbf{\Sigma}, \mathbf{Q}) = f_h(\mathbf{A}_0, \mathbf{A}_+)$ *and* $\mathcal{Z}$ *denotes the set of all structural parameters that satisfy the zero restrictions. Otherwise, set its importance weight to zero.*[4]

(6) *Return to Step 1 until the required number of draws has been obtained.*

(7) *Re-sample with replacement using the importance weights.*

Finally, Algorithm 1 is very efficient if one wants to draw from a conjugate prior distribution over the structural parameterization that can be induced by a uniform-normal-inverse-Wishart distribution over the orthogonal reduced-form parameterization. It turns out that the normal-generalized-normal over the structural parameterization that can be induced by a uniform-normal-inverse-Wishart distribution over the orthogonal reduced-form parameterization.

Algorithm 1 follows the steps highlighted above: it draws from a distribution over the orthogonal reduced-form parameterization conditional on the sign and zero restrictions and then transforms the draws into the structural parameterization. Step 5 is the crucial one. The re-sampling Step 7 allows us to have unweighted and independent draws from the desired posterior distribution. Given the desired number of independent draws, the researcher should require enough draws so that the effective sample size is larger than the desired number of independent draws.

[3]Because $\mathbf{Z}_j$ is empty for $j \geq 2$, $\mathbf{M}_j = \begin{bmatrix} \mathbf{q}_1 & \cdots & \mathbf{q}_{j-1} \end{bmatrix}'$ for $j \geq 2$.

[4]See Arias, Rubio-Ramírez, and Waggoner (2018) for definition of the volume element $v_{(g \circ f_h)|\mathcal{Z}}(\mathbf{A}_0, \mathbf{A}_+)$.

It is also important to notice that computing the volume element $v_{(g\circ f_h)|\mathcal{Z}}(\mathbf{A}_0, \mathbf{A}_+)$ in Step 5 is the most expensive part in implementing Algorithm 1. The rest of Algorithm 1 is quite fast. But the reader should note that we do not need to compute $v_{(g\circ f_h)|\mathcal{Z}}(\mathbf{A}_0, \mathbf{A}_+)$ for all the draws, only for those that satisfy the sign restrictions.

IV. Results

For each economy, we consider an SVAR with 6 lags and a constant term, and we use a normal-inverse-Wishart prior distribution, $NIW(\bar{\nu}, \bar{\mathbf{\Phi}}, \bar{\mathbf{\Psi}}, \bar{\mathbf{\Omega}})$, for the reduced-form parameters. We set $\bar{\nu} = 0$, $\bar{\mathbf{\Phi}} = \mathbf{0}$, and $\bar{\mathbf{\Psi}} = \mathbf{0}$ and $\bar{\mathbf{\Omega}}^{-1} = \mathbf{0}$. We apply the identification procedure developed above to each economy as a way to examine the role of monetary policy as a source of business cycle fluctuations.

First, we present the time series of the estimated monetary policy shocks in the ELB periods in IV.1. Then we report and discuss the IRFs of variables to monetary policy shocks in Section IV.2. For each country, we compare the IRFs of monetary policy in normal times with those during the ELB times. Third, we describe how the coefficients of a quasi-reaction function of the central bank respond, i.e. how the interest rate responds to the other variables of the model across sub-periods. Finally, we assess the quantitative importance of these shocks using the variance decomposition of forecast errors in Section IV.4.

The results shown in the following sub-sections are based on 10,000 draws that satisfy the sign and zero restrictions in Restrictions I-III using the Gibbs-sampling procedure presented by Algorithm 1. Table 6 shows the effective sample size for each of set of draws. As the reader can see, all the effective sample sizes are reasonably high.

IV.1. **The estimated monetary policy shocks.** In this section, we report the estimated monetary policy shocks for normal and ELB times. We focus our analysis on the monetary policy shocks during ELB times since we think these are of more interest to the reader. The bottom panels of Figures 2(b), 3(b) and 4(b) report the time series of estimated monetary policy shocks when the short-term interest rate is at the ELB for each economy.[5] These time series, which are a measure of exogenous changes of monetary policy, can be used for future analyses of the effects of monetary policy. In Figures 5(b), 6(b) and 7(b), we report a centered MA(3) of the same shocks. This smoothing procedure helps in spotting periods of consistent

[5]We report posterior means for each period.

monetary policy easing. Strikingly, the most important periods of monetary policy easing correspond to the implementation of major steps in non-conventional monetary policies, in particular, the periods when non-conventional monetary policies were either pre-announced, announced or implemented.

For the United States, these include the early months of QE1, from April 2009 to December 2009, periods in the neighborhood of the beginning of QE2 in September and October 2010, the first three months of QE3 in August 2012, and the period between October 2013 and August 2014. During this last period, the Fed kept a highly accommodatory monetary policy stance, after the taper tantrum of early 2013.

Turning to the euro area, the most important monetary policy easings as estimated by the model are the OMT in the summer of 2012, negative rates in June 2014 and the run-up to the asset purchase program in late 2014 and early 2012. Interestingly, a couple of studies, such as Andrade, Breckenfelder, De Fiore, Karadi, and Tristani (2016) and Altavilla, Carboni, and Motto (2015), have shown that the pre-announcements of QE by Mario Draghi between September 2014 and January 2015 had larger effects on asset prices than the purchases themselves after March 2015.

Finally, in the case of Japan, according to the model, the QQE program in March 2013 and the yield curve control in 2016 correspond to the Bank of Japan's most important easing periods.

IV.2. **Dynamic effects of monetary policy shocks.** In this section, we compare the effects of monetary policy during normal and ELB times. Specifically, for each economic area, we consider the effects of monetary policy shocks during both normal and ELB times on endogenous macroeconomic variables, namely industrial production, prices, monetary policy indicator, monetary aggregate, and credit spread (or stock market for Japan).

Figures 8(a), 9(a), and 10(a) report the IRFs of endogenous variables for the United States, the euro area, and Japan respectively, in normal and ELB times. More specifically, we report in red the IRFs in normal times, while blue indicates those for ELB times. For each panel, the median is reported in a solid line and the 68% probability intervals in colored areas. In addition, for series in log-levels (industrial production, prices, and money), we report the deviation in percent, whereas other panels (policy indicator and spread) display the deviation in percentage points. For comparability across periods, the monetary policy shock is scaled to induce an immediate reduction of 10 basis points in the two-year interest rate. All the IRFs will be to an expansionary monetary policy shock.

Regarding the results for the US economy (see Figure 8(a)) both in normal and ELB times the credit spread falls and the monetary aggregate rises after an expansionary monetary policy shock. The effects persist during both sub-periods although the effects seem to last longer during ELB times. What are the effects on output and prices? For both sub-periods, output rises slowly, reaching its maximum after more than 20 months in normal times and around 10 months in ELB times. While the long-run effects on output turn out to be much stronger in normal than in ELB times, output reacts much faster during ELB times. Moreover, the 68% probability intervals lie within the positive region, indicating that monetary policy shocks have positive effects on output. The effect on prices is also qualitatively consistent across sub-periods, although there are some important differences. In normal times, the increase in price levels appears to be very persistent and relatively modest, with a short-run price puzzle. The response is much larger and faster during ELB times; prices rise immediately and they start decaying six months after the shock. What is the message from this figure? Overall, most 68% IRF intervals show that the shape of the IRFs does not change across sub-periods. This clearly indicates that monetary policy has remained effective during the ELB period in the United States. These results are broadly consistent with previous studies that estimated the effects of US monetary policy, although on somewhat different samples or estimation approaches, including Weale and Wieladek (2016), Debortoli, Gali, and Gambetti (2018) and Gertler and Karadi (2015) and Panizza and Wyplosz (2018).

Regarding the results for the euro area economy (see Figure 9(a)), an expansionary monetary policy shock is also estimated to have positive effects on both output and prices, although the levels of uncertainty differ substantially. Hence, the transmission of output and prices is qualitatively similar across sub-periods but uncertainty around the median response is much larger during ELB times. This may reflect that, as the sample is much shorter in this case, the estimates are less precise. Regarding the response of money and credit spreads, the response is also similar across sub-periods. As was the case with the United States, the results show qualitative similarities across sub-periods, indicating that monetary policy has remained effective during the ELB period in the euro area. Boeckx, Dossche, and Peersman (2014) found some effects the non-conventional policies of the ECB on activity, although their sample at the ELB is shorter than ours. The counterfactual simulations reported in Figure 12 of Neri and Siviero (2019) are consistent with our estimates and the ones of several recent papers which they review. The ECB's unconventional policy has stimulated economic activity even at the ELB.

Turning to Japan, the impulse response (see Figure 10(a)), of output is somewhat different across regimes. While in both periods, a monetary policy expansion produces an increase in output, the magnitude of this response is about four times larger in ELB times. Furthermore, the shock propagates faster in ELB times. This differs from the United States and Japan, where output behaves similarly in both sub-periods. Regarding prices, the behaviour is very similar to the one in the United States: stronger and faster effects in ELB times. Interestingly, the Nikkei index rises immediately after the easing of monetary policy, with larger effects on the stock market index during ELB times. Our results on the effectiveness of monetary policy at the ELB concur with the estimates reported in Koeda (2019), Panizza and Wyplosz (2018) and Kimura and Nakajima (2016).

In summary, our most relevant result is that monetary policy easing shocks are estimated to spur activity and prices in both normal and ELB times in the United States, the euro area, and Japan.

IV.3. **Restrictions II and III.** As described in Section III.5, following Arias, Caldara, and Rubio-Ramírez (2019), we impose restrictions on the systematic component of monetary policy. Arias, Caldara, and Rubio-Ramírez (2019) show how important is to restrict these coefficients in order to induce a contraction in output after a surprise increase in the federal funds rate. They also show that, if these restrictions are not imposed, one obtains systematic components of monetary policy that have no economic meaning with high posterior probability.

Tables 3, 4, and 5 show the posterior of the monetary policy coefficients for the three economies. It is interesting to see that, for all three economies, the systematic part of monetary policy has become less responsive to inflation during the ELB periods. The parameter ψ_p is smaller during the ELB times for all three economies. The change in the coefficient of output is similar; the parameter ψ_y is smaller during the ELB times for all three economies. Our results indicate that monetary policy authorities are less responsive to both output and prices during ELB times. The monetary authority also induces lower responses of the two-year interest rate to changes in the money aggregate and the indicator of credit sentiment during ELB times (the only exception is the euro area, where the response of the two-year interest rate to money is stable across regimes). In any case, the largest drops in the coefficients are observed in the United States. For example, the two-year interest rate response to changes in output is almost 20 times smaller during ELB times. This lack of response could be interpreted as a deliberate move on the part of the Federal Reserve to keep liquidity ample and interest rates low until well into the recovery. Considered together with the very large contribution of MP shocks to the variance of US output during the ELB period, our estimates highlight that monetary policy had a considerable effect on the business cycle during the ELB period.

IV.4. **Relative importance of monetary policy shocks.** Using variance decompositions, we now evaluate the relative importance of monetary policy shocks in driving fluctuations in endogenous variables. Figures 8(b), 9(b), and 10(b) report, for each economic area, the percentage of the variance of the error made in forecasting each endogenous variable due to monetary policy shocks across the two sub-periods at forecasting horizons between the first and the 36th months after the initial shock. As before, red areas represent the contribution of money supply shocks in normal times, while blue areas are those in ELB times.

The variance decompositions for the US economy (see Figure 8(b)) indicate that the contribution of disturbances to monetary policy equation to business cycle fluctuations differs substantially across the two sub-periods. In normal times, these shocks do not explain much of the variations in output and prices, although they represent a non-negligible source for long-run output variations (about 20 percent). Interestingly, these results are consistent with Caldara and Herbst (2019), who show that the contribution of policy shocks to variation in industrial production is higher in VAR models that allow an immediate endogenous response of monetary policy to credit spreads. Specifically, the authors report that monetary shocks explain about 20 percent of output variability. As Caldara and Herbst (2019), our results differ from the VAR literature that reports a much lower contribution from monetary policy shocks. Interestingly, when looking at ELB times, the fraction of those shocks explaining macroeconomic variables increases dramatically. Indeed, they explain about 40 percent of the volatility of output and prices at a three-year horizon. Overall, the contribution is substantially larger for all endogenous variables in comparison with normal periods, although only error bands for output, prices and money across the two sub-periods do not overlap, reinforcing our finding.

In the euro area (see Figure 9(b)), the contribution of disturbances to monetary policy equation to endogenous variables differs substantially across regimes as well. In normal times, money supply shocks account for about 40 percent of variability in output and prices, while their contributions are more modest, although non-negligible, in ELB times (about 15 percent). The important contribution of money supply shocks to output variability is consistent with the monetary literature in the euro area. For example, Peersman and Smets (2001) and Smets and Wouters (2003) report similar results, where money supply shocks explain significant shares of the long-run variability of output. Finally, the fact that the contribution of MP shocks is lower in ELB times contrasts with the US result. One likely difference is the role of the sovereign debt crisis, which dragged on from for several years after 2010 and has contributed to the 2012-2013 recession in the euro area. Hence, the beginning of our ELB sample for the euro area may see a larger role for shifts in credit sentiment shocks, which may have had larger effects, thus reducing the weight of monetary shocks in fluctuations of output. This conjecture may deserve further investigation in future work.

Regarding variance decompositions for the Japanese economy (see Figure 10(b)), like the United States and euro area, our model shows a non-negligible role in output and prices movements for monetary policy shocks, although their contribution differs across regimes. More specifically, MP shocks explain about 40 percent of output variability in normal times, and only about 10 percent in ELB times. Such patterns are qualitatively similar to ones observed in the euro area. Regarding the level of prices, the contribution of money supply shocks is estimated less precisely. The 68 percent error bands of the contributions overlap across regimes that are therefore not clearly different from one another.

In summary, although its contribution to the fluctuation of output and price differs across countries and periods, monetary policy has maintained its impact even when interest rates were at the ELB.

V. Conclusion

To answer the question of whether the liquidity trap exists, our results clearly show that monetary policy has remained effective in a low interest rate environment. In the case of the euro area, Japan and the United States, a short-term interest rate of zero or below has not, at least up till now, translated into a ”Hicks-type liquidity trap”. Our results reinforce a view of monetary transmission that works predominantly through credit quantities and where the interest rate channel has only a marginal role. In such a view, the central bank's inability to lower the short-term interest rate is irrelevant, provided that it can ramp up credit supply and if at least some non-financial economic agents are credit-constrained. Of course, the effectiveness of non-conventional monetary policy cannot be taken for granted. Our estimates remain subject to the Lucas critique. However, so far, the liquidity trap has been more of a theoretical concept than an empirical reality.

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Appendix A. Data

All data are organized monthly, but with a different sample period for each country due to data availability. For each economy, data sources are reported below.

A.1. **The United States.** Data comes from the Federal Reserve Economic Data (FRED) developed by St. Louis Fed., except for the credit spread.

(1) ip_t: The industrial production index (INDPRO);
(2) p_t: The consumer price index for all urban consumers: all items (CPIAUCSL);
(3) r_t: The 2-year government bond yield;
(4) m_t: Monetary aggregate M1 (M1SL);
(5) sp_t: Non-financial corporate credit spread by Gilchrist and Zakrajsek (2012).

A.2. **Euro Area.** Data comes from the ECB - Statistical Data Warehouse, with the exception of the credit spread, which comes from Gilchrist and Mojon (2018).

(1) ip_t: The industrial production index (Euro Area 19, fixed composition, working day and seasonally adjusted);
(2) p_t: The harmonized index of consumer price (HICP). The series has been seasonally adjusted with the Jdemetra+ software;
(3) r_t: The 2-year government bond yield;
(4) m_t: Monetary aggregate M1;
(5) sp_t: Non-financial corporate credit spread by Gilchrist and Mojon (2018).

A.3. **Japan.** Most of the data comes from FRED, which proposes Japanese macroeconomic time series initially provided OECD, with the exception of the credit spread.

(1) ip_t: Production of Total Industry (JPNPROINDMISMEI);
(2) p_t: Consumer Price Index of All Items (CPALTT01JPM661S);
(3) r_t: The 2-year government bond yield;
(4) m_t: Monetary aggregate M1 (MYAGM1JPM189S);
(5) $stock_t$: Nikkei stock market index.

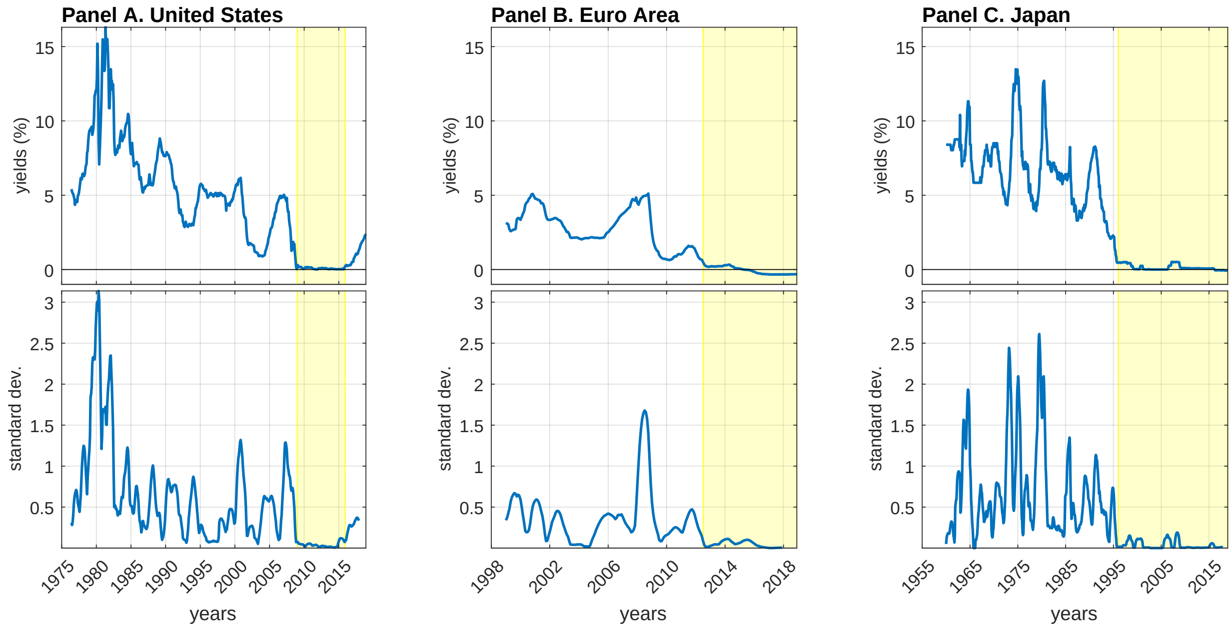
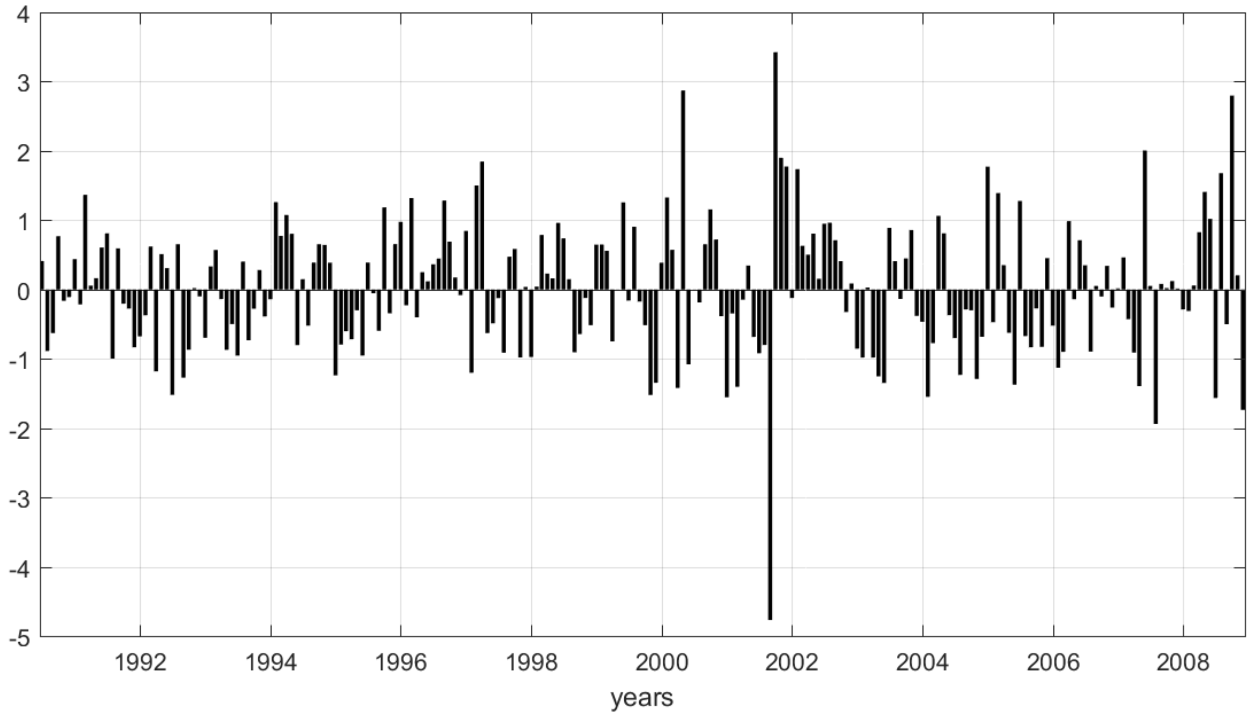
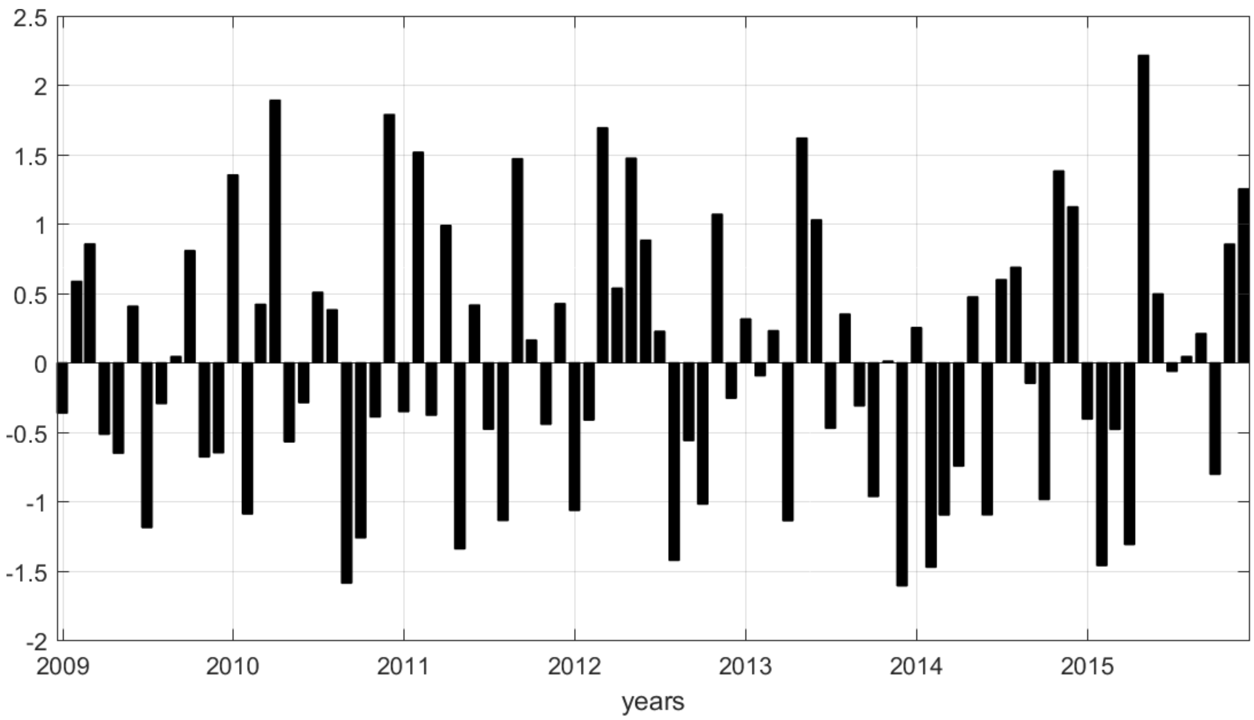


FIGURE 1. Time series of short-term policy rates. The yellow areas represent the ELB times: 2009.M1—2015.M12 (United States), 2012.M07—2018.M12 (Euro Area), 1996.M01—2018.M12 (Japan).

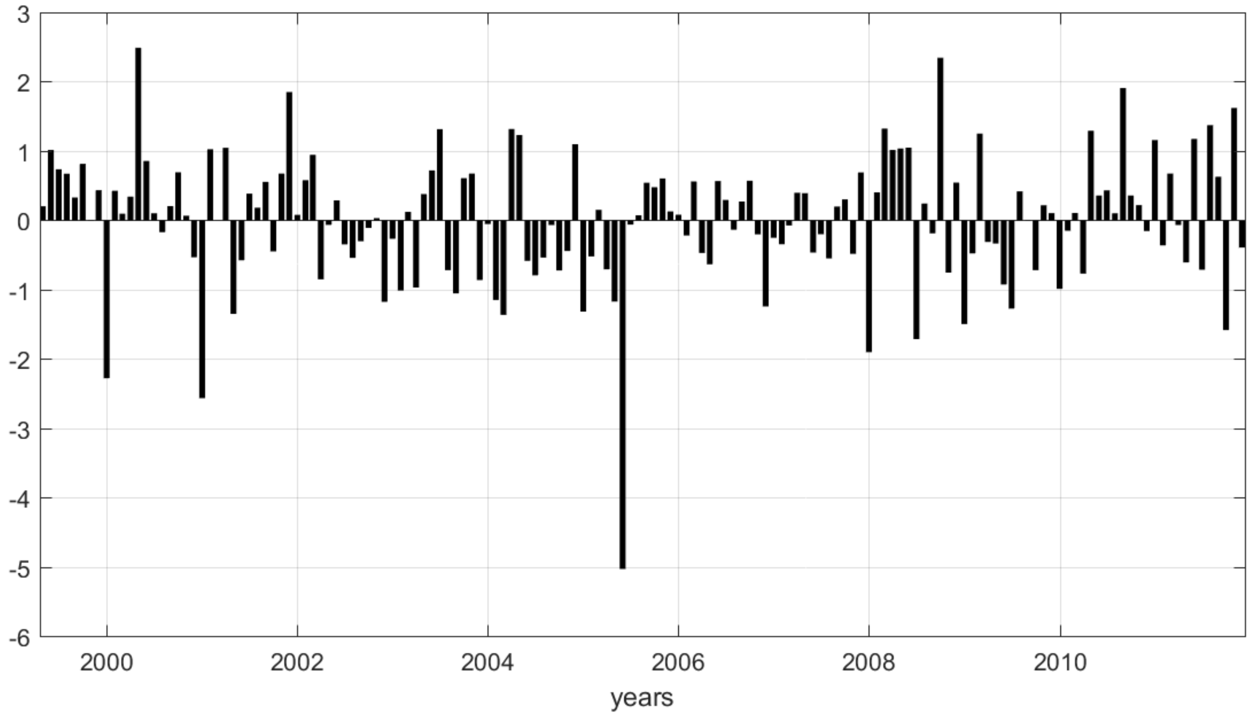


(a) Normal

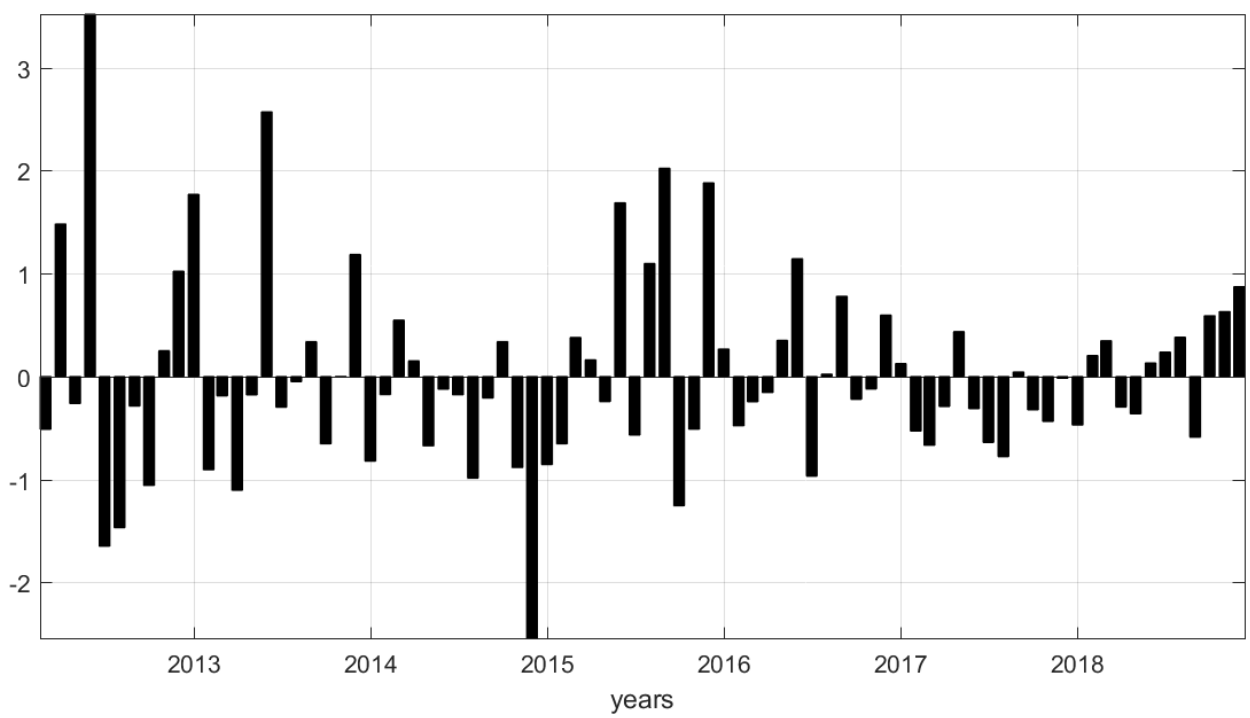


(b) ELB

FIGURE 2. Rough monetary policy shocks — United States

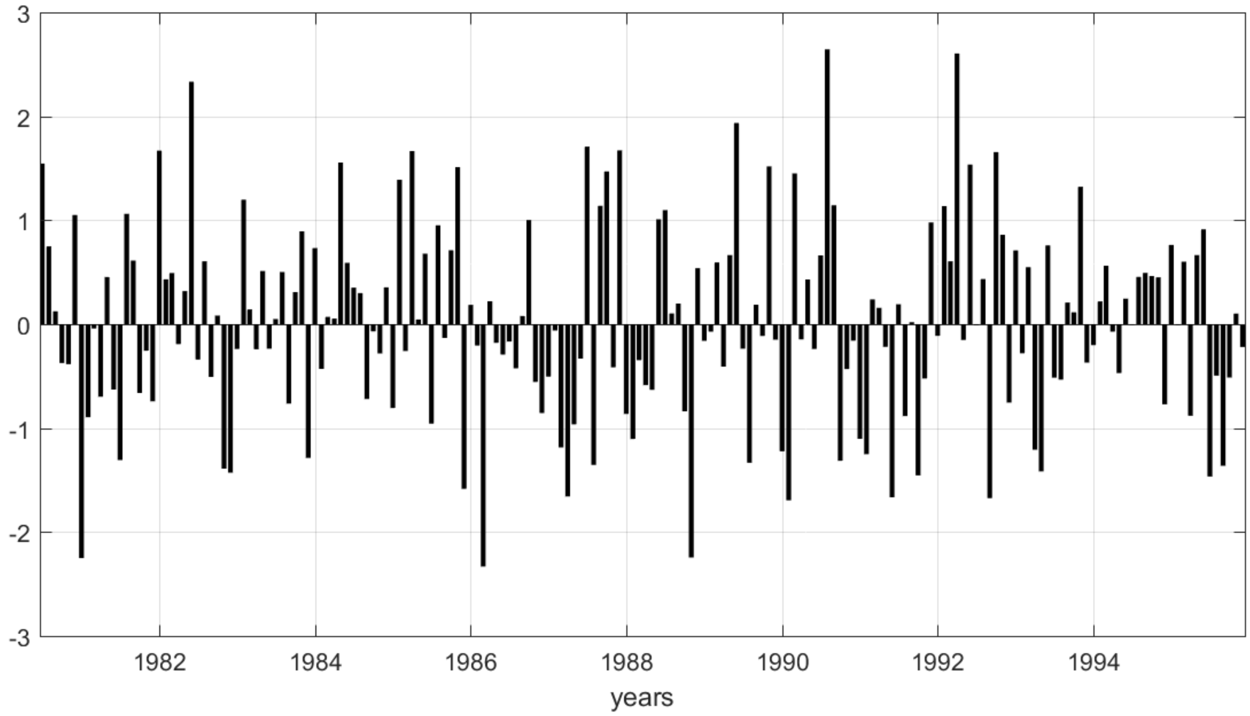


(a) Normal

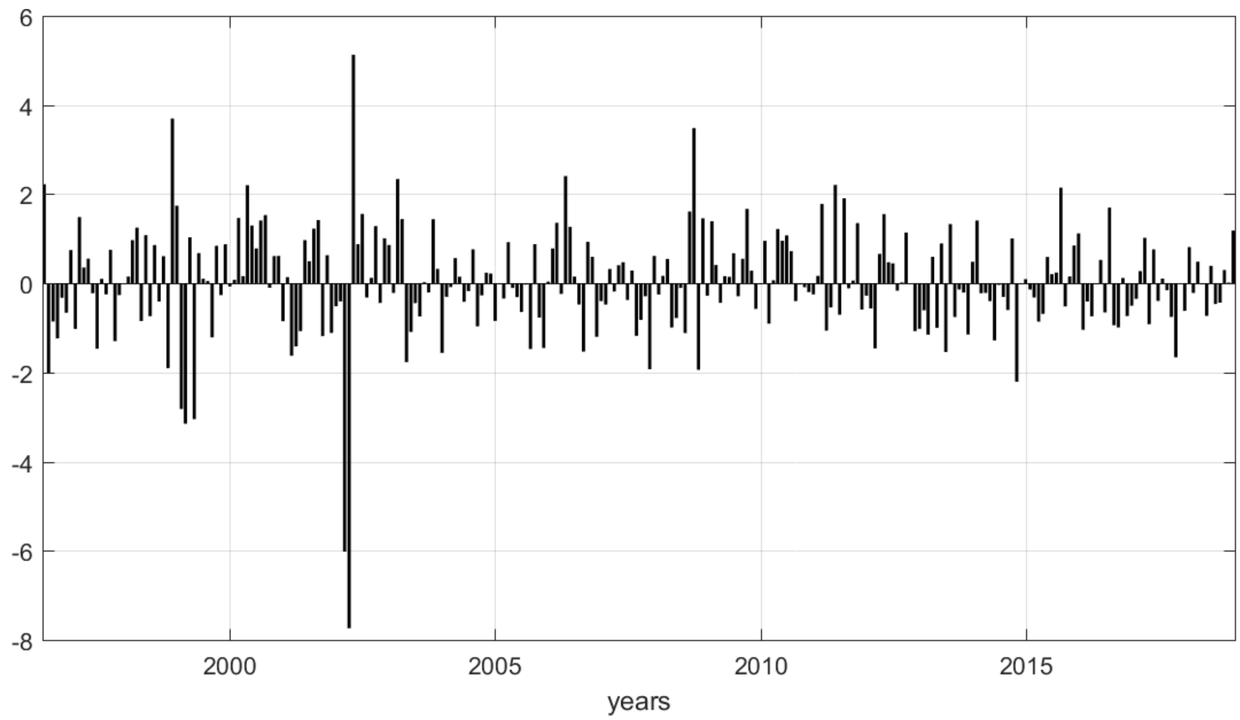


(b) ELB

FIGURE 3. Rough monetary policy shocks — Euro Area

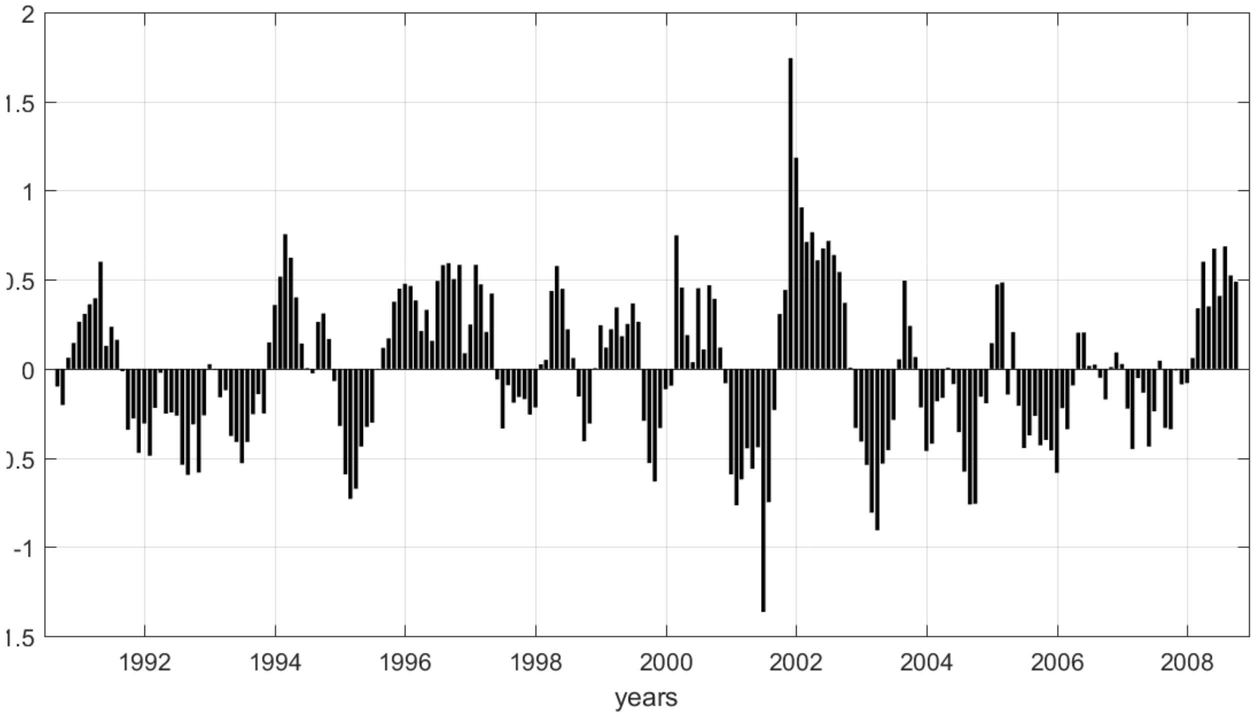


(a) Normal

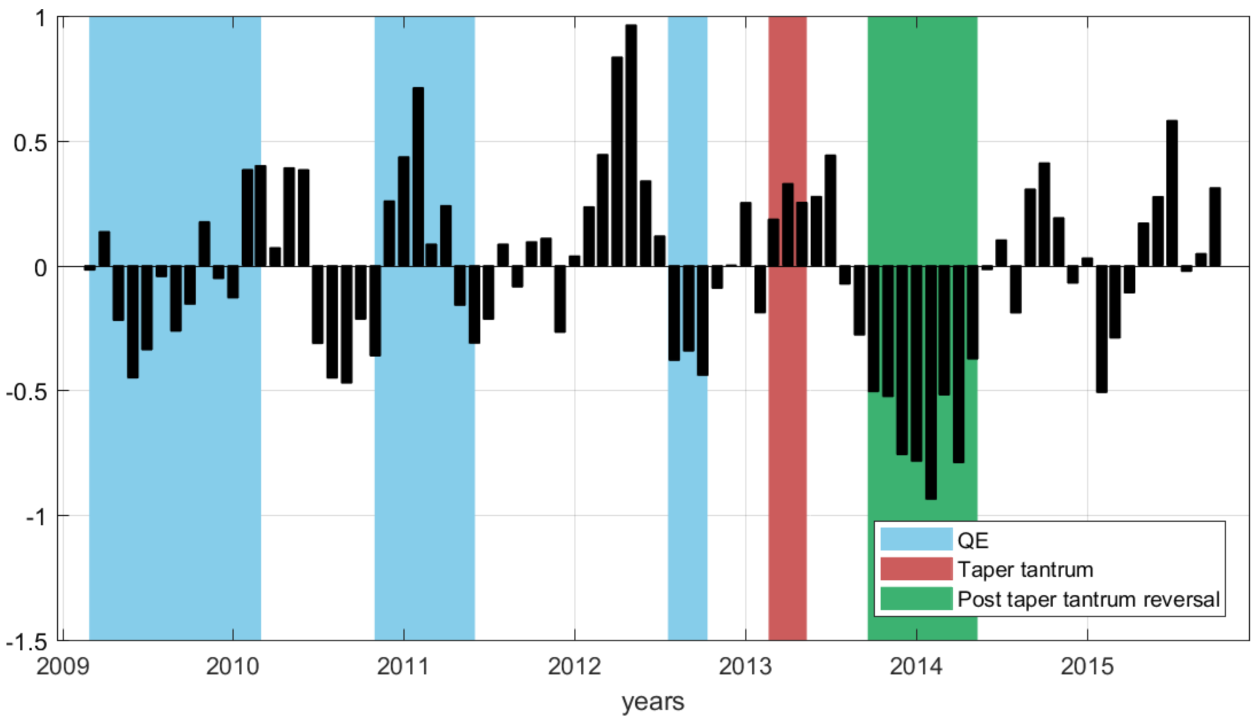


(b) ELB

FIGURE 4. Rough monetary policy shocks — Japan

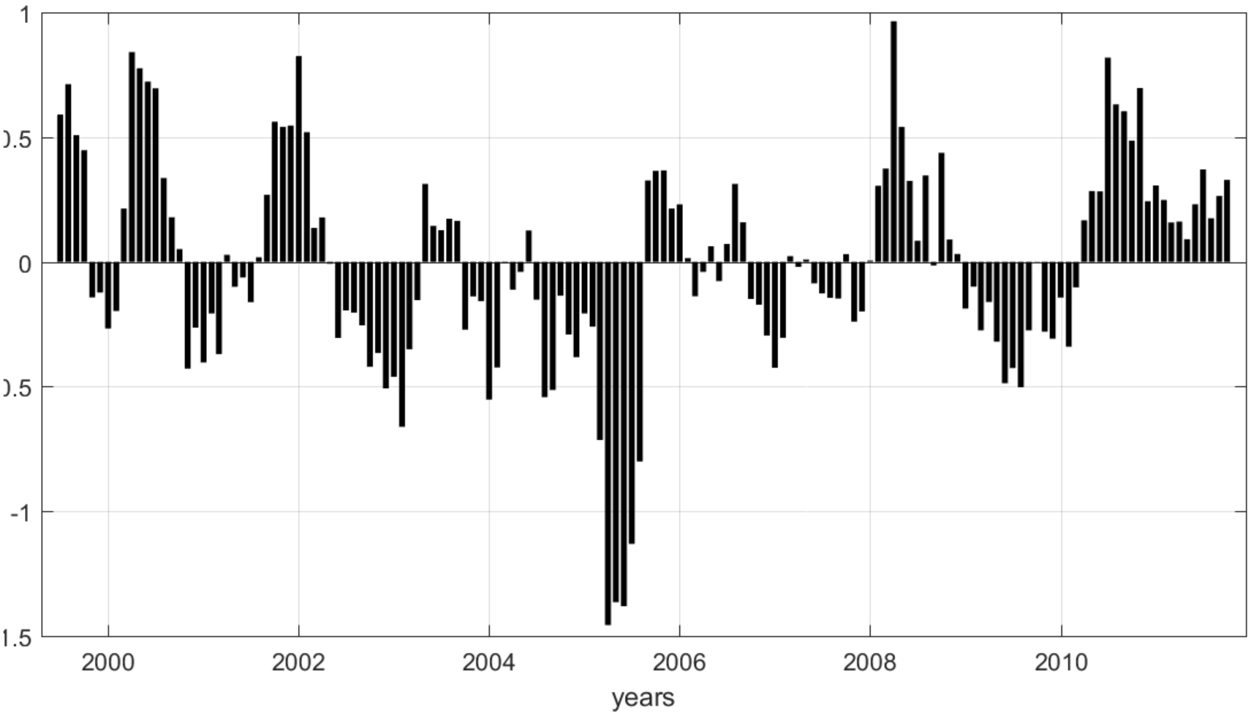


(a) Normal

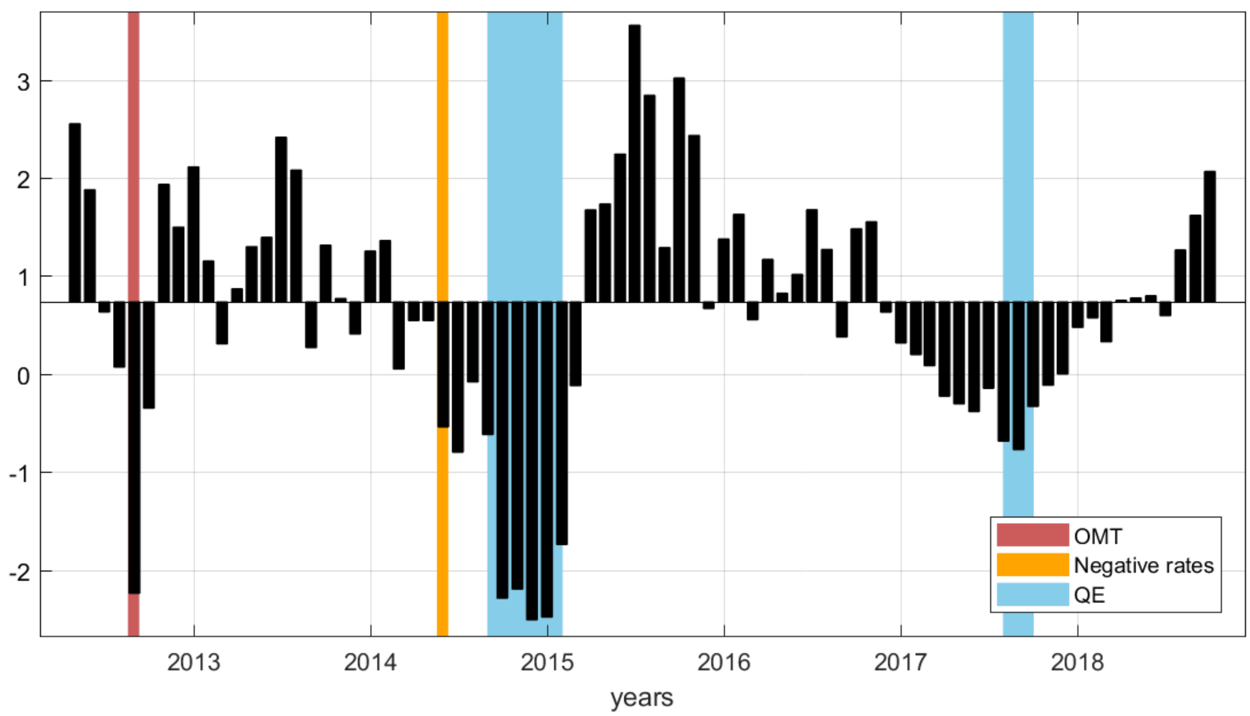


(b) ELB

FIGURE 5. Smoothed monetary policy shocks — United States

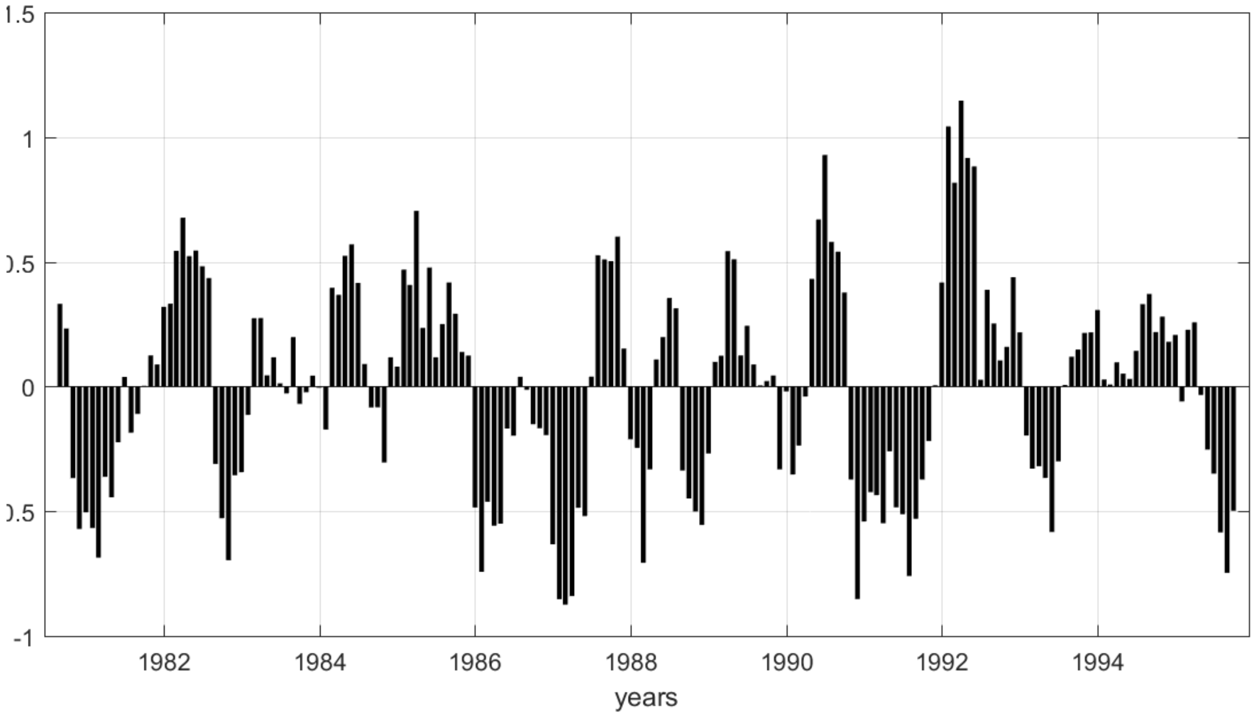


(a) Normal

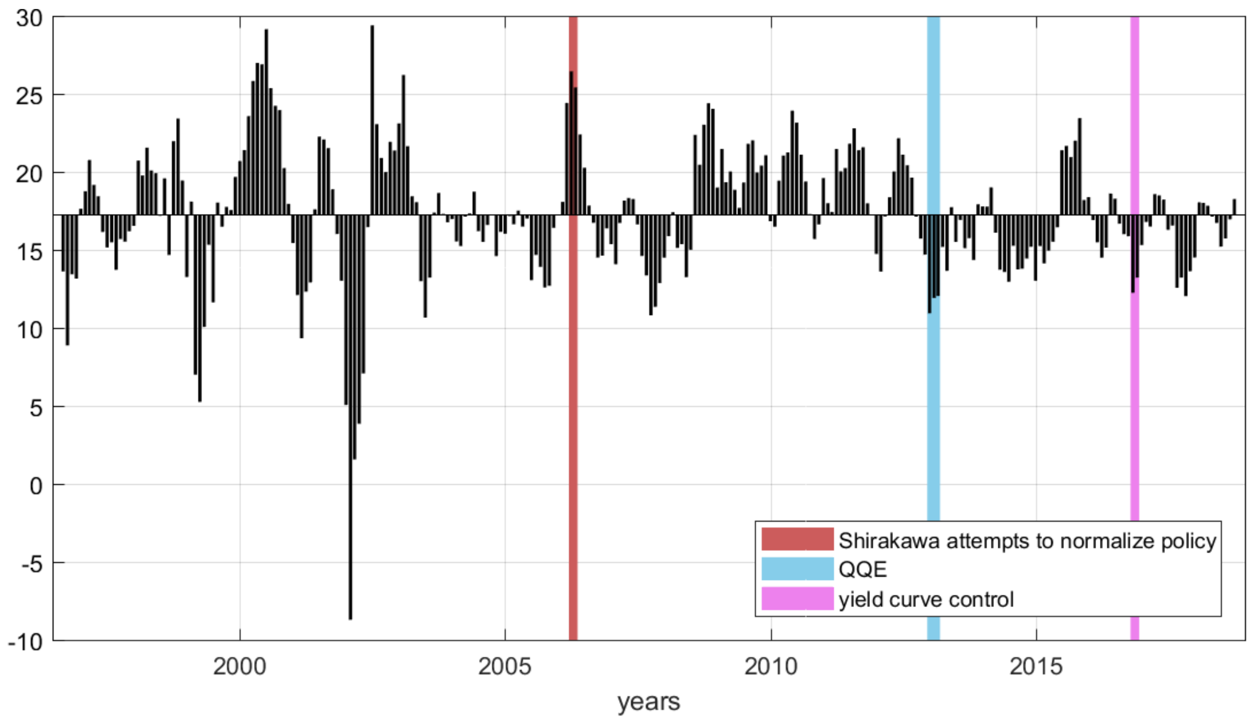


(b) ELB

FIGURE 6. Smoothed monetary policy shocks — Euro Area

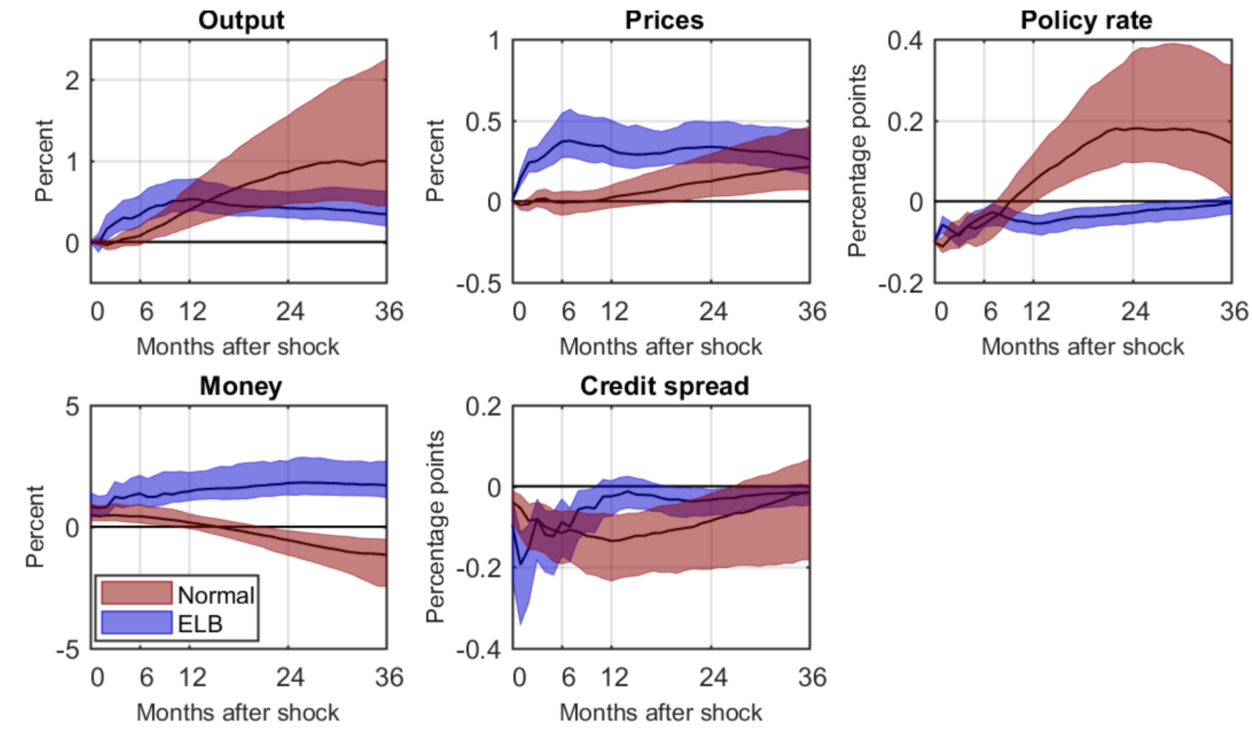


(a) Normal

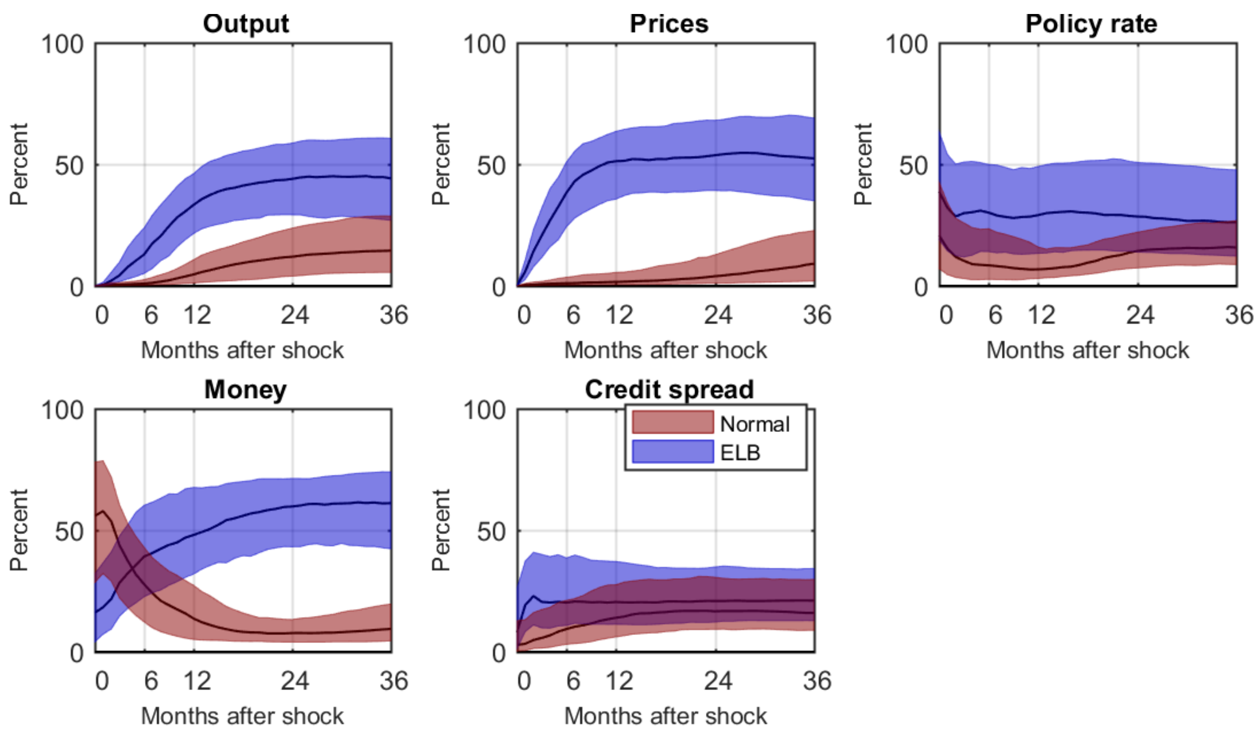


(b) ELB

FIGURE 7. Smoothed monetary policy shocks — Japan

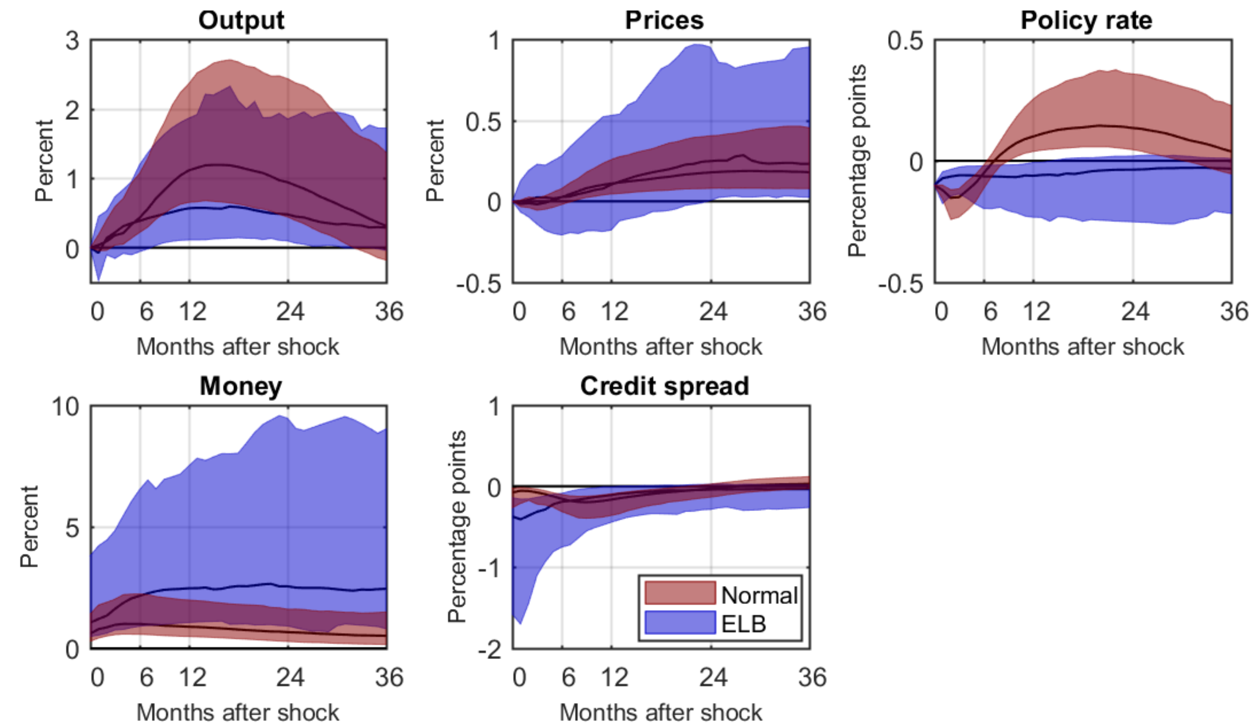


(a) IRFs

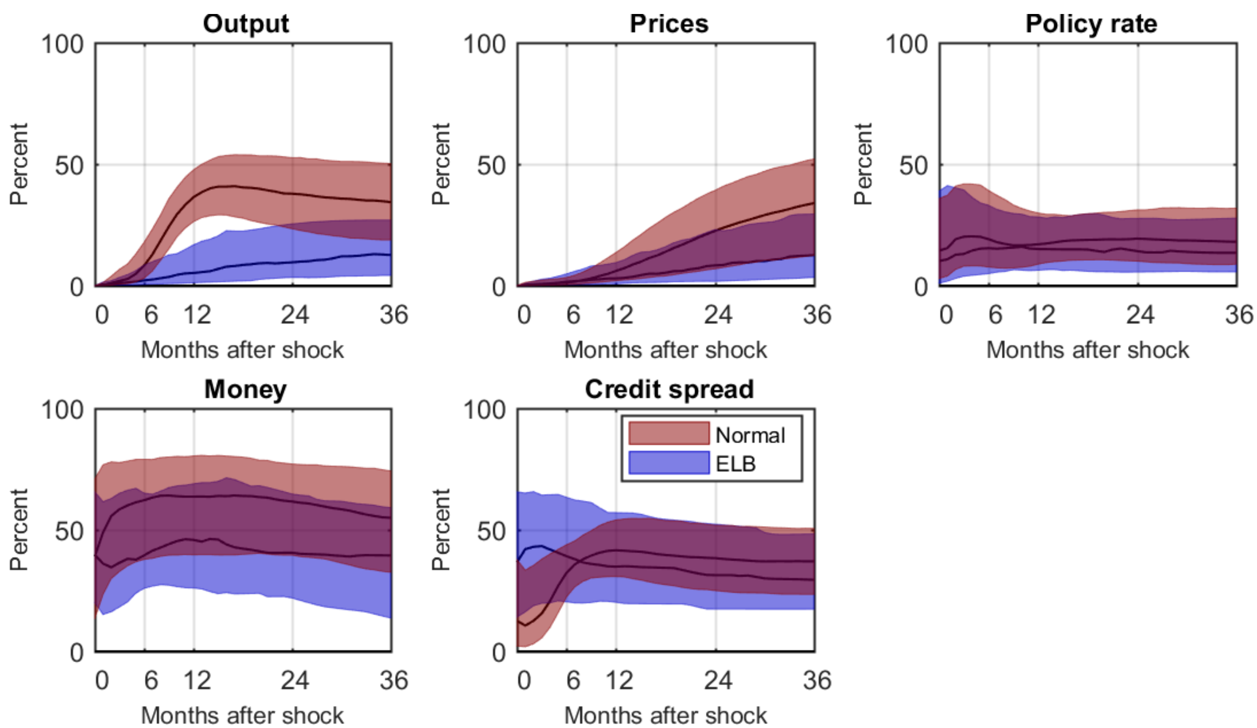


(b) VDs

FIGURE 8. United States

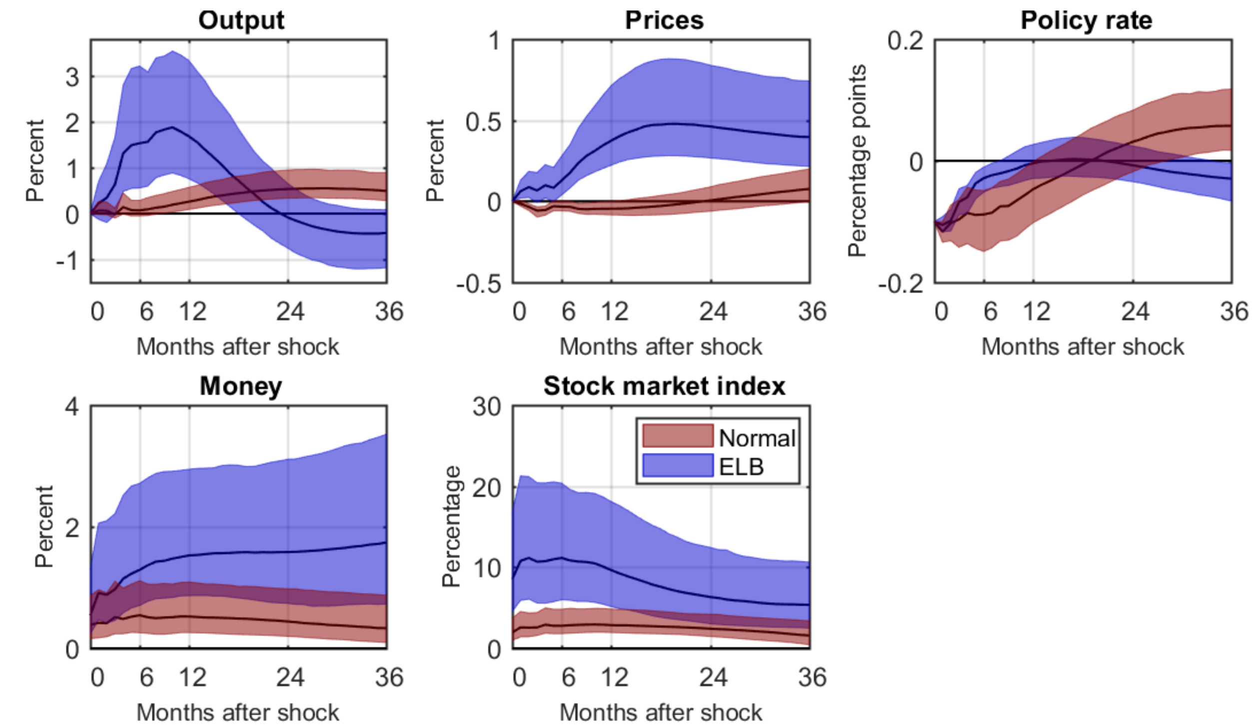


(a) IRFs

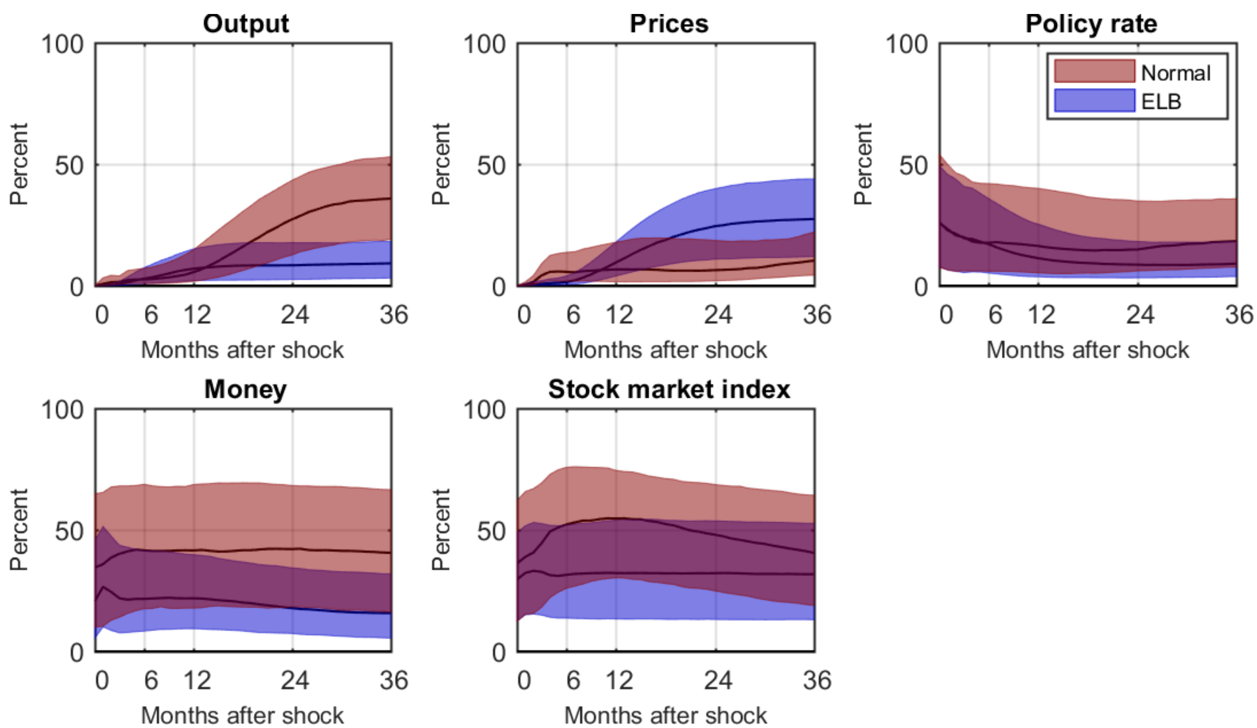


(b) VDs

FIGURE 9. Euro Area



(a) IRFs



(b) VDs

FIGURE 10. Japan

Table 1. Monetary regimes.

	Normal Times	**ELB Times**
United States	1990.M01—2008.M12	2009.M01—2015.M12
Euro Area	1999.M01—2012.M06	2012.M07—2018.M12
Japan	1980.M01—1995.M12	1996.M01—2018.M12

Note: ELB times are defined as periods where short-term interest rate reaches zero and its standard deviation is flat at its minimum.

Table 2. Restrictions I.

	(ip_t)	(p_t)	(r_t)	(m_t)	(sp_t)
Monetary policy shock	0	0	-	+	-

Note: Sign and zero restrictions. "0" stands for no contemporaneous response, "+" positive response, and "-" negative response. Sign restrictions are imposed on impact and over the next five months after the initial shock.

Table 3. Coefficients in the Monetary Policy Equations for United States

Coefficient	ψ_y	ψ_p	ψ_m	ψ_{sp}
[Normal period]				
Median	0.1714	0.0387	0.5258	-0.8668
68 percent error bands	[0.1086; 0.2835]	[0.0105; 0.1086]	[0.2806; 0.9606]	[-1.3895; -0.5548]
[ELB times]				
Median	0.0104	0.0226	0.0477	-0.3748
68 percent error bands	[0.0034; 0.0280]	[0.0049; 0.0511]	[0.0242; 0.0779]	[-0.5695; -0.2456]

Table 4. Coefficients in the Monetary Policy Equations for Euro Area

Coefficient	ψ_y	ψ_p	ψ_m	ψ_{sp}
[Normal period]				
Median	0.0253	0.1505	0.2193	-1.1618
68 percent error bands	[0.0078; 0.0565]	[0.0473; 0.2929]	[0.0979; 0.4609]	[-1.8266; -0.8040]
[ELB times]				
Median	0.0134	0.1114	0.2249	-0.8131
68 percent error bands	[0.0033; 0.0504]	[0.0233; 0.3332]	[0.0842; 0.7831]	[-2.6486; -0.3250]

Table 5. Coefficients in the Monetary Policy Equations for Japan

Coefficient	ψ_y	ψ_p	ψ_m	ψ_{stock}
[Normal period]				
Median	0.0389	0.1289	0.4076	0.0796
68 percent error bands	[0.0106; 0.1222]	[0.0372; 0.3621]	[0.1629; 1.1632]	[0.0345; 0.1919]
[ELB times]				
Median	0.0071	0.0846	0.1415	0.0148
68 percent error bands	[0.0025; 0.0173]	[0.0255; 0.2293]	[0.0527; 0.3706]	[0.0093; 0.0251]

Table 6. Diagnostic Analysis

Country	Euro Area	United States	Japan
[Normal period]			
Draws satisfying signs	1554	1162	1976
Effective sample size	0.59	0.78	0.56
[ELB times]			
Draws satisfying signs	2183	1112	2936
Effective sample size	0.76	0.38	0.50

Note: Effective sample size as a share of the draws satisfying the sign and zero restrictions.

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