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# Data-driven structural BVAR analysis of unconventional monetary policy

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ABSTRACT

We apply a novel Bayesian structural vector autoregressive method to analyze the macroeconomic effects of unconventional monetary policy in Japan, the US and the euro area. The method exploits statistical properties of the data to uniquely identify the model without restrictions, and thus enables formal assessment of the plausibility of given sign restrictions. Unlike previous research, the data-based analysis reveals differences in the output and price effects of the Bank of Japan's, Federal Reserve's and European Central Bank's balance sheet operations.

## 1. Introduction

Many central banks undertook unconventional monetary policy (UMP) measures in the aftermath of the 2007-09 financial crisis to restore the normal functioning of the monetary transmission mechanism when the policy rates reached the zero lower bound of interest rates (ZLB), or to provide further stimulus to the economy. Each central bank adopted measures deemed most suitable to the circumstances of its currency area (See Fawley and Neely, 2013 and Ugai, 2007 for reviews) so that country-specific results can be thought to reflect the effectiveness of various measures (Gambacorta et al., 2014).

While conventional monetary policy targets low and stable inflation with a short-term interest rate as an instrument, UMP commonly consists of massive expansion of central banks' balance sheets and aims to influence longer term interest rates. In this paper UMP refers to the use of the central bank's balance sheet as a monetary policy instrument, also called 'balance sheet policies' by Borio and Disyatat (2010).[1]

The relatively limited literature analyzing the macroeconomic effects of central banks' balance sheet policies mostly uses structural vector autoregressions. In the few studies (Meinusch and Tillmann, 2016), (Weale and Wieladek, 2016), (Boeckx et al., 2017), (Gambacorta et al., 2014), (Schenkelberg and Watzka, 2013) focusing on the macroeconomic effects over a sample period during which central banks actually targeted macroeconomic conditions, no major differences between the countries arise. Specifically, an expansionary UMP shock is found to lead to a delayed significant temporary rise in output and prices in all countries, and the results are robust to alternative variables.

Sign restrictions are typically used for identification in the literature analyzing conventional or unconventional monetary policy. In the UMP literature, they are often combined with short-run zero restrictions to reduce the set of admissible impulse responses and hence to sharpen identification and, in some cases, to disentangle the UMP-shock from the business cycle or financial shocks (e.g. Gambacorta et al., 2014, Schenkelberg and Watzka, 2013). As the theoretical foundations of UMP are not well established, both the

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[1] This deliberate choice thus rules out those central bank's operations that leave the size of its balance sheet unaffected, for example the Federal Reserve's maturity extension program known as 'Operation Twist', and the central bank's use of communication about future policy decisions.

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signs and their restriction horizons are inevitably arbitrary. Obviously, if we are interested in the macroeconomic effects of certain policy, it is particularly desirable to leave the responses of macrovariables unrestricted.

To the best of our knowledge, the so-called statistical identification methods have not yet been employed in the UMP literature.[2] These methods facilitate statistical testing of exactly identifying short-run or long-run restrictions in SVAR models (see e.g. Lanne et al., 2017), whereas methods to assess the plausibility of sign restrictions have been either informal or difficult to generalize (see Lanne and Luoto, 2016, and the references therein). In this paper we employ the method recently put forth by Lanne and Luoto (2016) that exploits the statistical properties of the data to uniquely identify a SVAR model and enables the evaluation of the plausibility of sign restrictions by their probabilities of being compatible with the data. This is helpful in either labeling the statistically identified shocks, which do not carry any economic meaning as such, or in concluding that the sign restrictions imposed in the previous literature are not supported.

Apart from being able to assess the plausibility of sign restrictions, our approach has a number of additional benefits compared to the conventional approach to sign restrictions. First, it should yield more accurate impulse response functions. This follows from the fact that our impulse response analysis relies only on economic shocks that are found to plausibly satisfy the given restrictions. Second, since our model is uniquely identified, the uncertainty surrounding the impulse responses of sign and other set identified models – the so-called model identification problem (see e.g. Fry and Pagan, 2011) – disappears and reporting the results of the impulse response analysis is straightforward. Furthermore, as sign-identified SVAR models are set-identified, the posterior of the structural parameters is proportional to the prior and hence an uninformative prior cannot be used. In fact, Baumeister and Hamilton (2015) have recently shown that the results from sign-identified SVARs are driven by the (implicit) priors. In contrast, under our assumptions the impulse responses are point-identified so that their posterior distributions need not be driven by the priors and a genuinely uninformative prior can be used. This facilitates learning about the impulse responses from the data.

We find statistical support for the sign restrictions used in a number of previous studies in all three currency areas. This allows us to interpret the statistically identified shocks and impulse responses along the lines of our reference studies (Schenkelberg and Watzka, 2013), (Gambacorta et al., 2014), (Boeckx et al., 2017). However, our impulse responses of these shocks differ in interesting ways from those reported in these studies.

Importantly, unlike previous research, our analysis reveals differences in the macroeconomic impact of the three central banks' actions. Our unrestricted impulse response functions indicate that a UMP shock had an insignificant effect on the consumer price index (CPI) in Japan, while there is weak evidence of a lagged, positive impact on prices in the US and in the euro area, depending on the specification. Our results also point to an immediate positive output response in the euro area, to a more delayed and persistent impact in the US than previously found, and that the positive output effect in Japan was unlikely due to lower long-term interest rates. The differences in the effects of the balance sheet operations can be explained by the differences in the unconventional measures adopted by the three central banks.

Overall, there's more evidence of the effectiveness of UMP measures in stimulating economic activity than in stimulating inflation. In quantitative terms we find that unconventional monetary policy shocks had a somewhat stronger output effect in the US and in the euro area than reported in the benchmark studies. These results confirm the finding of the benchmark studies that unconventional monetary policy shocks lead to relatively larger output effects and weaker price effects than conventional monetary policy shocks.

The rest of the paper is organized as follows. Technical details of the econometric method are given in Section 2. Section 3 covers the empirical analysis and Section 4 concludes the paper.

## 2. Methodology

Structural vector autoregressions (SVARs) are a common tool to analyze monetary policy. Lanne et al. (2017) have shown that the SVAR model can be uniquely identified by statistical properties of the data. However, their model is only statistically, as opposed to economically, identified, and additional information is needed to give the shocks an economic interpretation. This information may come in the form of short-run or long-run restrictions that can also easily be tested in the framework of Lanne et al. (2017), and if not rejected, used for interpretation. However, as discussed in the Introduction, in the UMP literature identifying restrictions are typically sign restrictions that are not approached in a straightforward manner by classical methods, and to that end, we employ the Bayesian procedures recently devised by Lanne and Luoto (2016). In particular, they show how to assess the plausibility of a set of sign restrictions by their posterior probability, and we apply their approach to check the sign restrictions used in a number of previous empirical UMP studies.

Our empirical results are based on the following $n$-variate SVAR($p$) model

$$\mathbf{y}_t = \mathbf{a} + \mathbf{A}_1\mathbf{y}_{t-1} + \cdots + \mathbf{A}_p\mathbf{y}_{t-p} + \mathbf{B}\varepsilon_t, \tag{1}$$

where $\mathbf{y}_t$ is an $n \times 1$ vector of time series of interest, $\mathbf{a}$ ($n \times 1$) is an intercept term, $\mathbf{A}_1 \ldots, \mathbf{A}_p$ are $n \times n$ coefficient matrices and the $n \times n$ impact matrix $\mathbf{B}$, containing the contemporaneous relations of the structural errors $\boldsymbol{\varepsilon}_t$, is assumed nonsingular.

Following (Lanne et al., 2017), (Lanne and Luoto, 2016) assume that the error process $\boldsymbol{\varepsilon}_t = (\varepsilon_{1t}, \ldots, \varepsilon_{nt})$ is a sequence of independent and identically distributed random vectors with each component $\varepsilon_{it}$, $i = 1, \ldots, n$ having zero mean and finite positive variance $\sigma_i^2$. Furthermore, the components of $\boldsymbol{\varepsilon}_t = (\varepsilon_{1t}, \ldots, \varepsilon_{nt})$ are assumed mutually independent, and at most one of them has a Gaussian marginal

[2] As examples of using statistical information to identify conventional monetary policy shocks, see Bacchiocchi et al. (2017), Lanne et al. (2017), Lanne and Lütkepohl (2014), Normandin and Phaneuf (2004).

distribution. It should be noted that this allows one to specify a different distribution for each $\varepsilon_{it}$, of which at most one can be Gaussian. In fact, in Section 3 we assume that $\boldsymbol{\varepsilon}_i \sim t_{\lambda i}$, $i = 1, \dots, n$, allowing the degrees of freedom to differ across equations (therefore allowing even a normal distribution), which improves the overall fit of the model. Instead of imposing the distributions to be the same, we allow the data to define if the $\lambda_i$'s differ or are the same.

If the process $\mathbf{y}_t$ satisfies the stability condition

$$\det(\mathbf{I}_n - \mathbf{A}_1 z - \cdots - \mathbf{A}_p z^p) \neq 0,\ |z| \leqslant 1 (z \in \mathbb{C}),$$

then the SVAR($p$) model (1) has a moving average representation

$$\mathbf{y}_t = \boldsymbol{\mu} + \sum_{j=0}^{\infty} \boldsymbol{\Psi}_j \mathbf{B} \varepsilon_{t-j}, \tag{2}$$

where $\boldsymbol{\mu}$ is the unconditional expectation of $\mathbf{y}_t$, $\boldsymbol{\Psi}_0$ is the identity matrix and $\boldsymbol{\Psi}_j$, $j = 1, 2, \dots$ are obtained recursively as $\boldsymbol{\Psi}_j = \Sigma_{l=1}^{j} \boldsymbol{\Psi}_{j-l} \mathbf{A}_l$. Interest then lies in the matrices $\boldsymbol{\Psi}_j \mathbf{B} \equiv \Theta_j$, $j = 0, 1, \dots,$ the $k$th column of which contains the impulse responses of the $k$th structural shock $\varepsilon_{it}$, $i = 1, \dots, n$.[3]

To see how identification is obtained, consider first the case of Gaussian errors. Under the standard assumption of temporally uncorrelated and normally distributed error term $\boldsymbol{\varepsilon}_t$, the matrix **B** cannot be identified because, for any nonsingular $n \times n$ matrix **C,** we can always replace the matrix **B** with the product **BC**, and the error term $\boldsymbol{\varepsilon}_t$ with $\mathbf{C}^{-1}\boldsymbol{\varepsilon}_t$. The two factors of the product **BC** need to be made unique in order to obtain unique impulse responses $\psi_j$ ($j = 0, 1, \dots$). Assuming a diagonal covariance matrix of $\boldsymbol{\varepsilon}_t$ means that the transformation matrix **C** has to be of the form $\mathbf{C} = \mathbf{DO}$ with **O** orthogonal and **D** diagonal. In identification by sign restrictions, **BD** is typically taken as the Cholesky factor of the covariance of the residuals of the corresponding reduced-form VAR model and the **O** matrices are some orthogonal matrices satisfying the sign restrictions.

On the other hand, as Lanne et al. (2017) show, assuming non-Gaussianity and independence of the structural error term $\boldsymbol{\varepsilon}_t$, we can restrict the orthogonal matrix **O** in the product **DO** to a permutation matrix - meaning that the matrix **B** is uniquely identified up to permutation and scaling of its columns. Changing the order of the columns of **B** means a different ordering of the structural shocks $\varepsilon_{it}$. That is, if a process $y_t$ can be represented by two potentially different moving average representations

$$\mathbf{y}_t = \boldsymbol{\mu} + \sum_{j=0}^{\infty} \boldsymbol{\Psi}_j \mathbf{B} \varepsilon_{t-j} = \boldsymbol{\mu}^* + \sum_{j=0}^{\infty} \boldsymbol{\Psi}_j^* \mathbf{B}^* \boldsymbol{\varepsilon}_{t-j}^* \tag{3}$$

then for some diagonal matrix $\mathbf{D} = diag(d_1, \dots, d_n)$ with nonzero diagonal elements, for some permutation matrix $\mathbf{P}(n \times n)$, and for all $t$, $\mathbf{B}^* = \mathbf{BDP}$, $\varepsilon_t^* = \mathbf{P}'\mathbf{D}^{-1}\boldsymbol{\varepsilon}_t$ and $\boldsymbol{\mu}^* = \boldsymbol{\mu}$, $\boldsymbol{\Psi}_j^* = \boldsymbol{\Psi}_j (j = 0, 1, \dots)$. Therefore the SVAR (1) and its moving average representation (2) are unique apart from permutation and scaling of the columns of **B** (see Proposition 1 in Lanne et al. (2017) and the proof therein).

In this paper, we are only interested in the unconventional monetary policy shock. In other words, our goal is to find out whether there is a single shock among the $n$ statistically identified ones that satisfies the sign restrictions imposed in each of the previous studies that we consider. If such a shock can indeed be found, we compare its impulse responses to those of the original study. To that end, we employ the Bayesian procedure of Lanne and Luoto (2016). We start out by estimating the joint posterior distribution of the parameters of the unrestricted SVAR model (1), and then compute the posterior distribution of the reduced-form impulse response matrices $\boldsymbol{\Psi}_j$, $j \in L$, where $L$ consists of indices of the restricted impulse responses. For instance, if the sign restrictions are imposed on the first $q + 1$ impulse responses, $L = \{0, 1, \dots, q\}$.[4] Because any or none of the $n$ components of $\boldsymbol{\varepsilon}_t$ can satisfy the restrictions and hence be the structural shock of interest, we next compute the conditional probability of each shock $\varepsilon_{it,}$ $i = 1, \dots, n$ satisfying the restrictions, conditional on none of the others satisfying them. In practice this is done using the posterior distribution of the identified structural impulse responses $\boldsymbol{\Theta}_j = \boldsymbol{\Psi}_j \mathbf{B},\ j \in L$. For each $i \in \{1, \dots, n\}$, this probability can be interpreted as the posterior probability of the restricted SVAR model where the sign restrictions are imposed on the $i$th column of the $\boldsymbol{\Theta}_j$, $j \in L$ matrices only. Among the $n$ models, those satisfying the sign restrictions in the (true) data-generating process (DGP) are expected to have high posterior probabilities. Therefore, one can rank the SVAR models satisfying the restrictions by their posterior probabilities, and so find a shock that is most likely the shock of interest.[5] The economic shocks with the gratest probability can be given the economic interpretation related to the corresponding restrictions. On the other hand, if the sum of the posterior probabilities is small, i.e. all of the models take a negligible probability, we can conclude that the data does not lend support to the restrictions.

## 3. Empirical analysis of unconventional monetary policy

The Bank of Japan's (BoJ), the Federal Reserve's (Fed) and the European Central Bank's (ECB) unconventional actions mainly differ because of differences in the structures of the economies and financial markets in particular. While the euro area and Japan are bank-centric economies, bond markets play an important role in the United States. The respective central banks therefore provided

[3] Although the MA-representation (2) does not exist for integrated VAR($p$) processes, their impulse responses are given by the same recursion. A similar decomposition exists for I(1) variables and is known as the Beveridge-Nelson decomposition (see Lütkepohl, 2006, Section 6.1).

[4] Because different permutations of **B** produce the same shocks and impulse responses, the choice of the permutation does not matter. Just to ensure that the whole analysis is based on the same ordering of the shocks, the permutation of the columns of **B** is fixed (for details, see Lanne and Luoto 2016).

[5] The procedure described here can be generalized to the case of multiple structural shocks, see Lanne and Luoto (2016).

liquidity and support to different segments of the financial sector: the Fed concentrated on bond purchases, the ECB on lending directly to banks, and the BoJ's strategy involved both.

Most UMP measures consist of an active use of the central banks' balance sheet (Borio and Disyatat, 2010), which is therefore a natural gauge for UMP although other measures have also been used in the literature. In line with our reference studies, the policy instruments are central bank assets for the Fed and the ECB, and reserves for the BoJ given its explicit target for reserves in the early 2000s.

Although the major central banks' unconventional measures were only undertaken after the financial crisis, a few studies are based on longer samples (e.g. Lenza et al., 2010, Peersman, 2011).[6] Since UMP measures are only undertaken when the economy faces particularly difficult times, utilizing data far beyond such a period may not be adequate to assess the effects of those measures. Therefore our samples cover periods over which UMP was in use and the central banks had macroeconomic goals. A detailed description of the data is deferred to an appendix.

We now provide a few details concerning the practical implementation, and then present the results of the formal assessment of previously used identification schemes and analyze impulse response functions in each geographical area in turn.

### 3.1. The set-up

We first identify structural shocks statistically and, following (Lanne and Luoto, 2016) then proceed to formally assess the validity of the sign restrictions used by Schenkelberg and Watzka (2013) for Japan, Gambacorta et al. (2014) for the US and Boeckx et al. (2017) for the euro area. As the data turns out to lend support to the restrictions, we then move on to impulse response analysis of the economic shocks.

As discussed in Section 2, non-Gaussianity is required for identification. We assume that the $i$th independent component of the error vector $\varepsilon_{it}$ follows a univariate Student's $t$ distribution with $\lambda_i$ degrees of freedom. Given normality is the standard, generally accepted assumption, we choose the $t$ distribution, as it is close enough to the normal distribution but seems more suitable for macro data (see e.g. Ascari et al., 2015, Chib and Ramamurthy, 2014 and Gourieroux et al., 2017). The advantage of the $t$ distribution is its flexibility and, in our setting, the fact that from the degrees of freedom one can infer the strength of the identification. This is because a $t$-distributed random variable converges to a Gaussian as the number of degrees of freedom goes to infinity. Hence, small values indicate that the fat-tailed $t$ distribution is in fact a more suitable assumption for the errors than normality, implying (strong) identification.

To learn as much as possible about the impulse responses from the data, we use non-informative priors. We assume an exponential prior distribution with mean 5 and variance 25 for each degree of freedom parameter $\lambda_i$ and a Gaussian prior for the inverse of the error impact matrix $vec(\mathbf{B}^{-1}) \equiv \mathbf{b}$, $\mathbf{b} \sim \mathbf{N}(\underline{\mathbf{b}}, \underline{\mathbf{V}}_{\mathbf{b}})$ where $\underline{\mathbf{V}}_{\mathbf{b}}^{-1} = c_{\mathbf{b}}\mathbf{I}_{n^2}$ and $c_{\mathbf{b}} = 0$, which results in an uninformative (improper) prior for $\mathbf{B}^{-1}$, $p(\mathbf{B}^{-1}) \propto 1$. For the deterministic terms and coefficient matrices, collected in matrix $\mathbf{A} = [a, A_1^{'}, ..., A_p^{'}]'$, $vec(\mathbf{A}) \equiv \mathbf{a}$, we assume a normal prior distribution, i.e. $\mathbf{a} \sim \mathbf{N}(\underline{\mathbf{a}}, \underline{\mathbf{V}}_{\mathbf{a}})$ with $\underline{\mathbf{a}} = 0$ and $\underline{\mathbf{V}}_{\mathbf{a}} = 10000^2\mathbf{I}_{pn^2+n}$. For the US and the euro area we also present results based on a relatively more informative prior for $vec(\mathbf{A})$, which corresponds to the standard Minnesota/Litterman prior.[7]

### 3.2. Japan

The burst of the asset price bubble in the early 1990s in Japan led the Bank of Japan (BoJ) to be the first central bank to adopt the zero-interest rate policy. In March 2001 the BoJ changed its main operating target from the overnight call rate to bank reserves held at the BoJ, committed to maintain high reserves levels in the future and increased the outright purchases of long-term government bonds to attain its target (Ugai, 2007), (Borio and Disyatat, 2010).

We adopt the specification in Schenkelberg and Watzka (2013) who have analyzed the real effects of the Japanese unconventional monetary policy at the ZLB using post-1995 data in a sign-restricted BVAR. The Japanese data, plotted in Fig. 1, are analyzed with a five-variable structural BVAR model with an intercept and a trend.[8] Monthly data for Japan spans from March 1995 until September 2010.[9] The variables included are the core consumer price index (CPI), the Japanese industrial production index (IP), the bank reserves held at the Bank of Japan (RES), the 10-year yield of Japanese government bonds (LTY) and the real effective exchange rate of Yen against other currencies (EXR). Except for the long-term yield, all variables are expressed in logs. Given that we analyze the same variables and sample period as Schenkelberg and Watzka (2013), we follow them and include six lags in the VAR model. Maintaining the specification of the benchmark paper allows us to attribute any differences in results to the novel identification strategy we introduced.

[6] These also include nonlinear model specifications and policy instruments different from those discussed above (Darracq-Paries and De Santis, 2015), (Baumeister and Benati, 2012), (Kapetanios et al., 2012).

[7] Metropolis-within-Gibbs algorithm is used for the estimation of the posterior distribution of the parameters of the SVAR model given in (1). For details of the algorithm, see the appendix in Lanne and Luoto (2016). While the theoretical convergence of the algorithm is established in Geweke (2007), we verified the empirical convergence of the sampler with Geweke's MCMC diagnostics. However, we note that the MCMC diagnostics may not be reliable if the posterior distribution is bimodal or multimodal.

[8] The results are qualitatively the same with linearly detrended data and no trend in the model.

[9] The BoJ reintroduced QE measures – money market operations to increase the monetary base – in 2013 as part of the 'Abenomics' strategy. Since a linear model is not suitable to study a sample period which includes a change in the monetary policy regime, the sample cannot be extended to include the 'Abenomics' period.

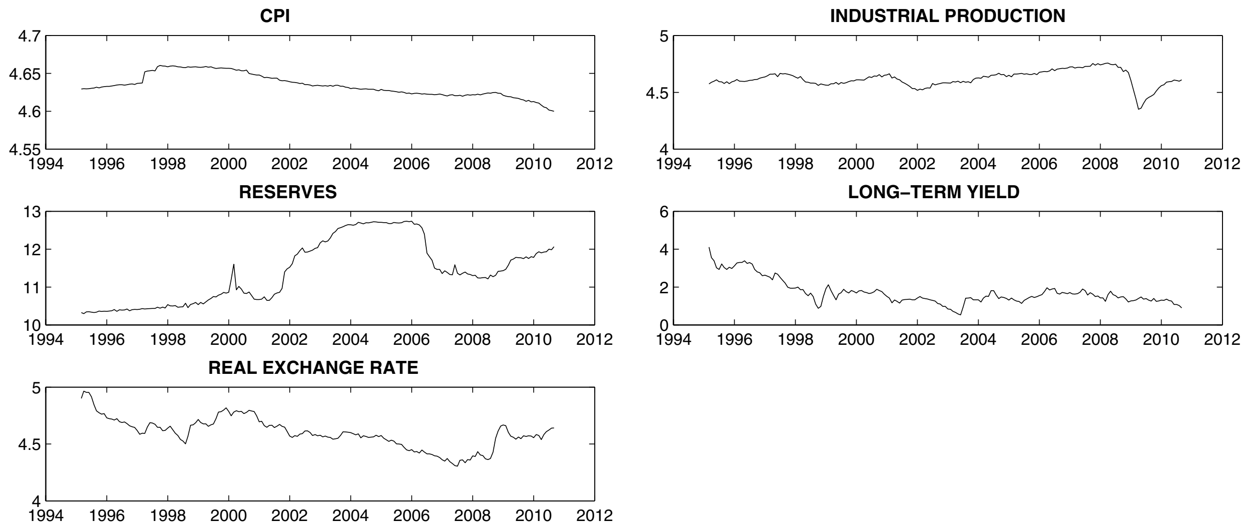


**Fig. 1.** Plot of logarithmic (excl. long-term yield) time series 1995M3–2010M9 for Japan.

In the present setup, the impact matrix **B** in (1) is uniquely identified under non-Gaussianity of at least four components of the error vector. The posterior means of the degree-of-freedom parameters of the univariate *t* distributions specified for the components of the error term lying between 2.2 and 4.6 thus provide evidence of non-normality and of successful identification.

To study the effects of UMP on output and price level, we need to pin down the right structural shock among the statistically identified ones. For that purpose we exploit the sign restrictions used by Schenkelberg and Watzka (2013) who assume that an expansionary UMP shock has a positive effect on the reserves held at the BoJ and a non-negative effect on consumer prices for 12 months.[10] Although (Schenkelberg and Watzka, 2013) use a 12-month restriction horizon, we first compute the posterior probability of each structural shock satisfying the restrictions on impact only ($h = 0$), and then for the cases $h = 0,\ 1$ and $h = 0,\ \ldots, 12$. The results are reported in the left panel of Table 1.

The sums of the posterior probabilities for these different cases range between 0.14 and 0.41, lending overall support to the restrictions irrespective of the horizon although the evidence is clearly weaker when the restrictions are required to hold for an entire year. Moreover, there is only one shock ($\varepsilon_{3t}$) with a high posterior probability when only the impact effect is restricted. It is found the likeliest candidate for the UMP shock also when the first two impulse responses are restricted although $\varepsilon_{1t}$ seems to be almost equally likely. Only in the case $h = 0,\ .,12$ the restrictions fail to pin down the shock. These results altogether speak in favor of a unique labeling of the UMP shock so that impulse responses can be analyzed. This labeling turns out to be robust to two alternative specifications, which we consider next.

Fig. 2 depicts the median impulse responses to a unit UMP shock along with the 16% and 84% percentiles of the posterior distribution. The UMP shock raises reserves approximately 3%, industrial production at most 0.15% after about two years but the impact on the price level and long-term government yield are insignificant in that the 68% highest posterior density credible sets contain the zero line. The effect on the real exchange rate is positive, increasing 0.25%, but barely significant.

In contrast to previous studies, these impulse response functions are obtained without restricting the effects on any of the variables and are solely based on the data. Therefore it is interesting to compare the results with those of Schenkelberg and Watzka (2013). It is worth noting that their response of reserves is of the same shape, persistence and magnitude as ours, and they also find a virtually insignificant effect of the UMP shock on the real exchange rate. On the other hand, their price response is weakly positive and temporary, while we find it to be insignificant also during the first year, when they restricted it non-negative. There is also a small difference in the negative impact response of industrial production, which only we find significant, but it is temporarily positive after 20 months in both studies. In both studies, the positive impact on industustrial production is mild. However the main difference is in the reaction of the long-term government bond yield, which (Schenkelberg and Watzka, 2013) report to be significantly *negative* for two years, whereas we observe a significantly *positive*, although very weak (one basis point), transient response of approximately six months. This finding is particularly interesting because asset purchases, which the BoJ engaged in to attain its target on reserves, are typically thought to work by lowering long-term rates. Although the positive response of long yields to an expansionary shock may seem counterintuitive, Schenkelberg and Watzka (2013) explicitly allow for this in their identification scheme and contrast it with previous studies. Our analysis only emphasizes the importance of remaining agnostic concerning the effects of UMP on the main variables of interest.

As a robustness check, we follow (Schenkelberg and Watzka, 2013) and consider a shorter sample period ranging from March

[10] The identification scheme in Schenkelberg and Watzka (2013) contains an additional contemporaneous zero restriction on consumer prices to disentangle the UMP-shock from demand and supply shocks. This is not required in our setup because identification is based on statistical properties of the data.

**Table 1**
Formal assessment of sign restrictions: Japan.

| | Benchmark model | | | Shorter sample | | |
|---|---|---|---|---|---|---|
| Shock | $h = 0$ | $h = 0, 1$ | $h = 0, .,12$ | $h = 0$ | $h = 0, 1$ | $h = 0, .,12$ |
| $\varepsilon_{1t}$ | 0.04 | 0.12 | 0.05 | 0.01 | 0.08 | 0.02 |
| $\varepsilon_{2t}$ | 0.02 | 0.04 | 0.01 | 0.03 | 0.12 | 0.05 |
| $\varepsilon_{3t}$ | 0.20 | 0.13 | 0.04 | 0.15 | 0.16 | 0.24 |
| $\varepsilon_{4t}$ | 0.02 | 0.05 | 0.03 | 0.00 | 0.02 | 0.01 |
| $\varepsilon_{5t}$ | 0.09 | 0.07 | 0.01 | 0.01 | 0.07 | 0.03 |
| Sum | 0.37 | 0.41 | 0.14 | 0.20 | 0.45 | 0.35 |

Notes: The figures in the top panel are the posterior probabilities of shock $\varepsilon_{it}$, $i = 1, \ldots, 5$ satisfying the sign restrictions that the reserves be positive and consumer prices be non-negative for various time horizons, and hence being the structural shock of interest. Benchmark model: reserves as the policy instrument. Shorter sample: sample period 2000M3–2007M3.

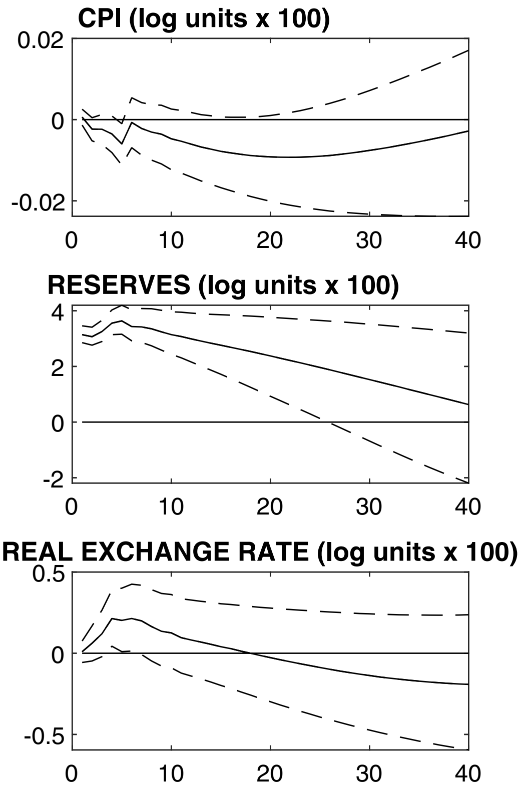


**Fig. 2.** Impulse responses to an expansionary UMP shock: Japan 1995M3–2010M9. Median responses (solid lines) together with 68% Bayesian credible sets (dashed lines).

2000 to March 2007.[11] The sample period covers approximately a year before and after the BoJ targeted current account balances. In fact, one could argue that although the BoJ's target rate was very close to zero since 1995, starting to target reserves marks the beginning of a different monetary policy regime. The posterior probabilities reported in the right panel of Table 1 show that, interestingly, the same shock ($\varepsilon_{3t}$) is uniquely identified as the UMP shock for all restrictions horizons.[12]

The impulse response functions are aligned with the short sample results in Schenkelberg and Watzka (2013). Their price response

[11] With the shorter sample lag length is set to $p = 2$.

[12] As another robustness check we analyzed a model with interpolated real GDP (instead of the industrial production), which has been used as a measure of aggregate output in Gambacorta et al. (2014) and Boeckx et al. (2017). A monthly measure of real GDP was constructed using the Chow-Lin interpolation method with monthly industrial production as a reference series. We observe that a similar pattern of probabilities emerges as in the previous specification: requiring reserves and the CPI to be non-negative on impact only uniquely identifies the UMP shock, while there are other shocks with positive probabilities in the case $h = 0, 1$, and no labeling is clearly supported for twelve months (posterior probabilities range from 0.02 to 0.05). There are also no major differences in the impulse responses compared to the benchmark case. The detailed results of the robustness analysis are not reported here to save space but are available upon request.

became insignificant as well, their response of real exchange rate turned from insignificant to positive, and in both studies the significant output effect occurs earlier than in the benchmark case. Interestingly, the main difference remains: we observe an insignificant effect on the long-term rate, while they documented an initial negative effect which then turns positive. As to the magnitude of the effects, Schenkelberg and Watzka (2013) reported quantitatively smaller effects on industrial production and CPI when the shorter sample was analyzed, while our results were quantitatively unaltered. We therefore conclude that our results are robust to the alternative output measure but shortening the sample period triggers sharper responses in output and real exchange rate, while the effect on the long term yield can be considered negligible in both cases.

Finally, as Weale and Wieladek (2016) have pointed out, unconventional monetary policies may reflect the reaction of the central bank to coincident developments such as domestic fiscal policy. To explore whether our results may suffer from the omission of important variables capturing fiscal policy, we have included the domestic government budget balance to GDP ratio in the VAR model. The results turn out to be robust to the addition of the variable.[13]

To summarize, the sign restrictions in Schenkelberg and Watzka (2013) are supported by the data on impact and after the first month following the shock, but they are not able to uniquely identify the UMP shock when imposed for an entire year, except for the shorter sample period. This allows us to pin down the right structural shock among the statistically identified ones and to conduct impulse response analysis.

Importantly, because we do not impose a positive price response, we are able to conclude that a UMP shock has no effect on the price level. This is in contrast to Schenkelberg and Watzka (2013) who forced the shock to have a positive effect for twelve months. They also documented a negative effect on the long-term government bond yield, whereas in our case positive, although very small values (one basis point) are included in the 68% posterior error bands. Our findings are robust to a different output measure but not entirely to a shorter sample period. The results indicate that the Japanese monetary policy with an explicit target for reserves had no effect on the core consumer price index. The policy managed to stimulate real economic activity slightly and with a delay but there is no strong evidence that it operated by lowering long-term interest rates.

### *3.3. United States*

In the aftermath of the 2007-09 financial crisis, the Fed and the ECB started to pursue less conventional monetary policies to restore financial and macroeconomic stability. Initially both focused on dysfunctional financial markets, while broader macroeconomic conditions soon became the targets. Due to the collapse of the housing price bubble and the related subprime crisis in the US, the Fed prioritized housing credit markets within its large scale asset purchase (LSAP) programs. In the first phase it pursued outright asset purchases of government-sponsored enterprise (GSE) debt, mortgage-backed securities (MBS) and long-term Treasury securities, the purchases of which several stages during the sample period. Although some of the operations were sterilized, leaving the monetary base unaffected, most of them were unsterilized (for details, see Fawley and Neely (2013)).

The existing literature on the macroeconomic effects of the Fed's balance sheet operations Gambacorta et al. (2014), Meinusch and Tillmann (2016), Weale and Wieladek (2016) uses different Bayesian VAR specifications (panel VAR, Qual VAR and SVAR, respectively), but obtains the same result for the key macroeconomic variables; an expansionary UMP shock leads to a temporary significant rise in output and prices.

According to Gambacorta et al. (2014) the main features of the crisis are captured by the following variables: the log of seasonally adjusted real GDP (GDP)[14], the log of seasonally adjusted consumer price index (CPI), the log of seasonally adjusted central bank assets (CBA) and the level of implied stock market volatility (VIX) to control for the central bank's balance sheet expansion resulting from financial market disturbances. We therefore adopt their specification but extend the sample to cover the period 2009M3-2014M5, plotted in Fig. 3. Using a different set of variables, Weale and Wieladek (2016) were the first to analyze this sample period, which does not span beyond the UMP period and is, hence, less susceptible to the Lucas Critique.

We specify a BVAR(2) with a constant consisting of the four variables.[15] With four variables, non-Gaussianity of at least three components of the error vector is crucial for identification. The posterior means of the degree-of-freedom parameters of the $t$ distributions of the error terms turned out to range from 2.8 to 4.2, lending support to fat-tailed error distributions and, hence, successful identification.

In order to find out whether any of the statistically identified shocks can be labeled as the monetary policy shock, we proceed with a formal assessment of the sign restrictions in Gambacorta et al. (2014) whereby an expansionary UMP shock increases central bank assets but does not increase stock market volatility on impact and one month after the shock.[16] We check the validity of the signs for

[13] Similarly, we have included the government budget balance to GDP ratio for the euro area and the public debt to GDP ratio for the US in the respective VAR models. Also these results turn out to be robust to the additional variables. Results are available on request.

[14] A monthly measure of real GDP is constructed using the Chow-Lin interpolation procedure with industrial production and retail sales as reference series.

[15] Both Gambacorta et al. (2014) and Weale and Wieladek (2016) include two lags in their models. As this choice is supported by the Hannan-Quinn information criteria and a subsequent Portmanteau test of (no) autocorrelation, we keep the lag length of two in our specification as well. Given the short time-series, increasing the number of lags would either lead to issues related to degrees of freedom problems or require the use of tight priors that do not allow the data to speak.

[16] Gambacorta et al. (2014) and Boeckx et al. (2017) impose additional contemporaneous zero restrictions on output and consumer prices to reduce the number of admissible impulse responses and so to sharpen identification. These are not required in our setup because the model is uniquely identified based on statistical properties of the data.

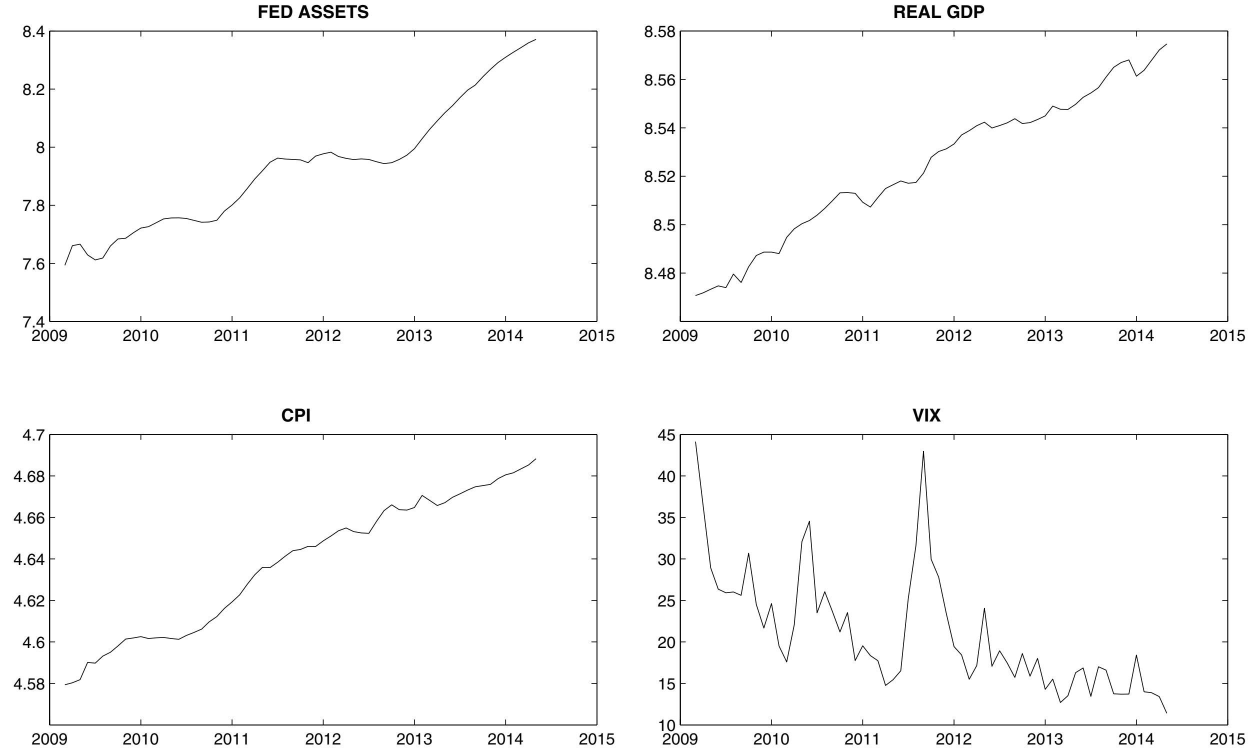


**Fig. 3.** Plot of logarithmic (excl. VIX) time series 2009M3–2014M5 for the US.

two restriction horizons: first on impact only ($h = 0$) and then for the case $h = 0, 1$. The results reported in the left panel of Table 2 show that there is not much difference between the posterior probabilities in the two cases. The sums of the posterior probabilities (0.12 and 0.16) lend overall support to the restrictions. Moreover, there is in both cases only one shock ($\varepsilon_{4t}$) with a high posterior probability, with the probability of the other shocks virtually zero, so that a UMP shock can be regarded as uniquely identified in probability.

The impulse responses, plotted in Fig. 4, show that a unit UMP shock increases the central bank assets on impact but the median peak response of 1% occurs after approximately eight months. While (Gambacorta et al., 2014) forced output and prices to respond with a lag and documented peak responses after six months, and Weale and Wieladek (2016) found output and prices to rise for 20–40 months after a UMP shock regardless of the identification scheme, our unrestricted impulse response functions indicate that the output response turns sigificantly positive only after ten months. The median response peaks at 0.08% as in Gambacorta et al. (2014), but in our case assets increase less, indicating more impactful QE. We also observe a more persistent output response, lasting up to 35 months. In contrast, the evidence for a positive CPI response is weaker, as the 68% Bayesian credible sets just include the zero line.

Taking into account the very small sample size, we also considered a more informative prior distribution, corresponding to the standard Minnesota/Litterman prior. Interestingly the relatively more informative prior resulted in a positive 0.08% price response after 30 months, with the rest of the responses unaltered. Moreover, further tightening the prior made the positive price response to

**Table 2**
Formal assessment of sign restrictions: United States.

| | Benchmark model | | Industrial production | | Monetary base | |
|---|---|---|---|---|---|---|
| Shock | $h = 0$ | $h = 0, 1$ | $h = 0$ | $h = 0, 1$ | $h = 0$ | $h = 0, 1$ |
| $\varepsilon_{1t}$ | 0.00 | 0.00 | 0.00 | 0.00 | 0.01 | 0.03 |
| $\varepsilon_{2t}$ | 0.00 | 0.00 | 0.00 | 0.01 | 0.01 | 0.02 |
| $\varepsilon_{3t}$ | 0.01 | 0.03 | 0.01 | 0.04 | 0.02 | 0.03 |
| $\varepsilon_{4t}$ | 0.11 | 0.16 | 0.05 | 0.11 | 0.39 | 0.42 |
| Sum | 0.12 | 0.19 | 0.06 | 0.16 | 0.43 | 0.50 |

Notes: The figures in the table are the posterior probabilities of shock $\varepsilon_{it}$, $i = 1, \ldots 4$ satisfying the sign restrictions that the central bank assets be nonnegative and the VIX be nonpositive for various time horizons, and hence being the structural shock of interest. The figures on the bottom line are the sums of the posterior probabilities. Benchmark model: central bank assets as policy instrument. Industrial production: Industrial production as a measure of aggregate output. Monetary base: monetary base as policy instrument.

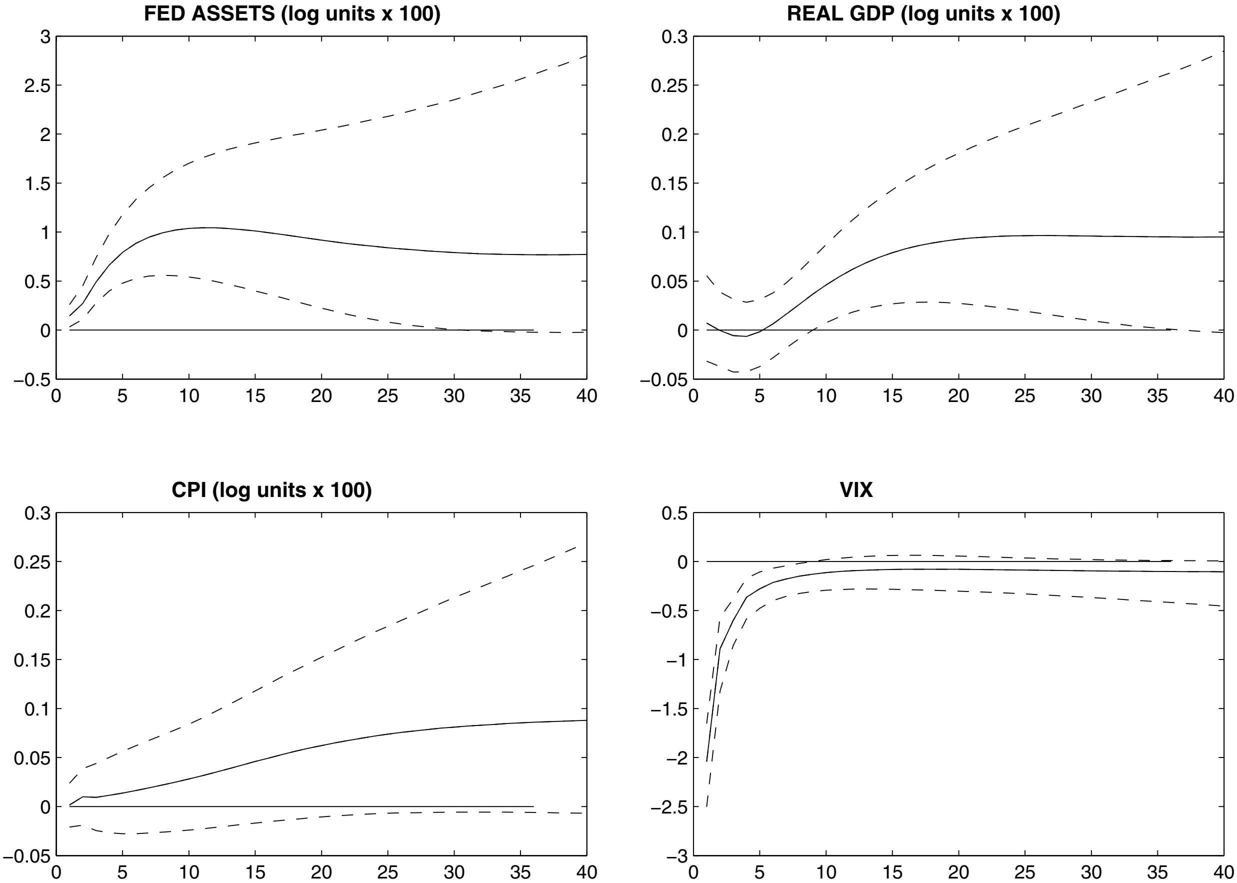


**Fig. 4.** Impulse responses to an expansionary UMP shock: US 2009M3–2014M5. Median responses (solid lines) together with 68% Bayesian credible sets (dashed line).

occur even earlier, but still much later than previously found. The effect is quantitatively similar to the benchmark study.

To check the robustness of our results, we considered industrial production as a measure of output and monetary base as the quantitative policy instrument. The middle and right panels of Table 2 show that the labeling is robust to both variables and the same shock ($\varepsilon_{4t}$) is uniquely identified in probability. There are, however, differences in the impulse response functions compared to the benchmark specification.[17] Interestingly, when industrial production is used, the positive CPI response becomes significant after 30 months even when a non-informative prior is used, while the rest of the responses remain the same. Quantitatively the responses remain unaltered. Again, tightening the prior has the same effect in that the CPI response becomes significantly positive earlier.

On the other hand, unlike documented by Gambacorta et al. (2014) and what we found for the euro area (see Section 3.4), the results from the impulse response analysis for the US are not robust to an alternative quantitative policy instrument. Although the posterior probabilities in Table 2 indicate that the sign restrictions are supported by the data, the impulse responses of the two macrovariables of interest are insignificant in that the 68% highest posterior density credible set includes the zero line. Furthermore, only a very tight prior triggers a significant positive output response similar to the previous specifications, while the price response remains insignificant. This finding is consistent with the fact that the effectiveness of balance sheet policies does not hinge on an accompanying change in the monetary base (Borio and Disyatat, 2010), and as already noted by Gambacorta et al. (2014), monetary base expanded less than central bank assets in the US over part of the sample period. It also indicates that differences between countries make panel methods less suitable to study the country-specific impact of unconventional monetary policies.

Overall our results indicate that the Fed's UMP measures were effective in supporting the macroeconomy in that they managed to stimulate economic growth, although the impact on infaltion remains unclear. We also confirm the finding in Gambacorta et al. (2014) that the effect of UMP on output is quantiatively larger and that on prices is smaller than those of conventional monetary policy.

Finally, given that one objective of the Fed for its QE policies was to reduce long-term rates, we have further analyzed the US system with the long-term interest rate variable included. Compared to the baseline set-up, the results remain unaltered. However, the impulse response of the long term rate, plotted in Fig. 5, turns out to be surprising. As in the case of Japan, the response is *positive*,

[17] These impulse response functions are not shown here to save space but are available upon request.

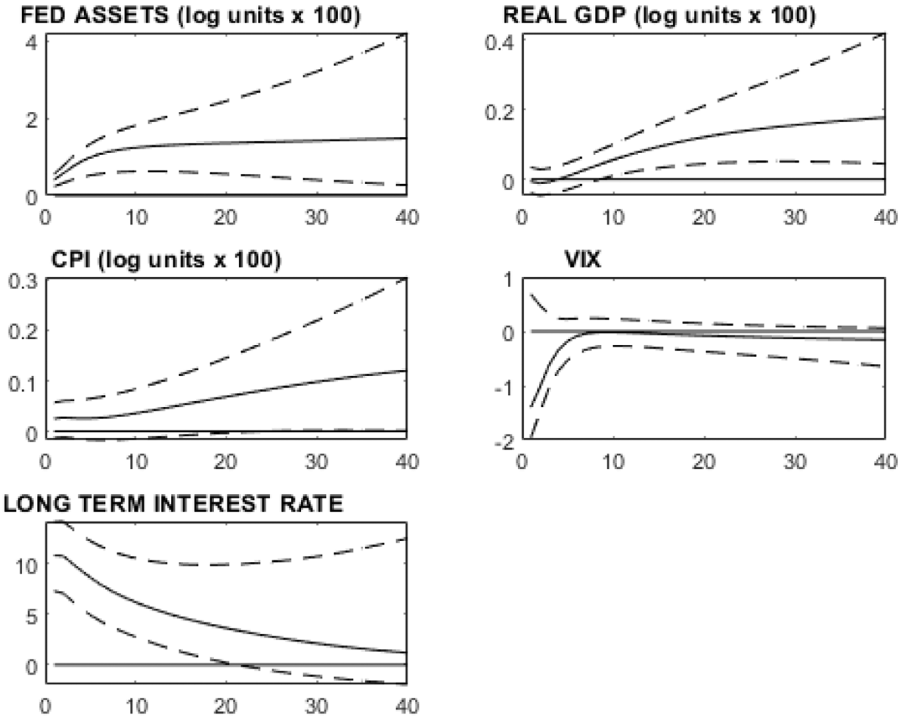


**Fig. 5.** Impulse responses for the US with long term interest rate. Median responses (solid lines) together with 68% Bayesian credible sets (dashed line).

approximately 10 basis points, and therefore does not seem to provide evidence for the interest rate channel being important for the effects of UMP. Some of the identification schemes in Weale and Wieladek (2016) also produce a similar effect for the UK, with the long rate increasing approximately 0.1% for assets purchases amounting to 1% of GDP. Given that there' s a broad literature studying the impact of UMP on financial variables finding evidence for the reduction of long rates, the result with macroeconomic variables is unexpected and should be investigated in more detail.

### 3.4. Euro area

Similarly to the Fed, the ECB's asset purchase programs aimed to improve the functioning of specific markets. The covered bond purchase program (CBPP) stimulated the issuance of covered bonds, and therefore eased funding conditions for banks (Beirne et al., 2011), whereas the objective of the Securities Markets Program (SMP) – later replaced by Outright Monetary Transactions (OMT) – was to address the malfunctioning of the securities markets caused by the sovereign debt crisis.[18] Apart from the SMP and its follower OMT, the majority of the ECB's operations during the sample period consisted of providing funding for banks. The ECB expanded both the availability and maturity of bank loans as well as eased the conditions for receiving funding on several occasions. Its asset purchases were modest in size and mostly sterilized, reversing their effects on the monetary base.

To investigate the effectiveness of the policy measures that expand the ECB's balance sheet, we adopt the VAR model specification of Boeckx et al. (2017)[19] The monthly ECB data, plotted in Fig. 6, spans from January 2007 until December 2014. Although the ECB has continued its unconventional policies beyond this date, we follow Boeckx et al. (2017) and end the sample period before the beginning of the Expanded Asset Purchase Program (EAPP).

The vector of endogenous variables comprises the log of seasonally adjusted real GDP (GDP), the log of seasonally adjusted consumer price index (CPI), the log of seasonally adjusted central bank assets (CBA) and the level of the Composite Indicator of Systemic Stress (CISS). Boeckx et al. (2017) also included in their model the main refinancing operations (MRO) policy rate and the spread between the EONIA and the MRO-rate. However with six variables the number of parameters to estimate increases considerably when no restrictions are imposed, and because of the short sample period this obviously creates problems in estimation.[20]

[18] See the 5.10.2010 ECB press release www.ecb.eu/press/pr/date/2010/html/pr100510.en.html

[19] As Boeckx et al. (2017) build on Gambacorta et al. (2014), also our study is related to theirs, with the difference of a longer sample period and the use of the CISS variable to measure overall financial stress in the euro area.

[20] In fact, with six variables the method adopted in this paper yielded results that did not allow us to make any conclusions even when using a very tight prior. Because one of the advantages of the method is the ability to check the compatibility with the data of the restrictions imposed in the conventional approach, we choose to stick to the 4-variable specification. Moreover, our conclusions turn out to be similar to those obtained by Boeckx et al. (2017) and most differences can be seen to follow from (the absense of) restrictions.

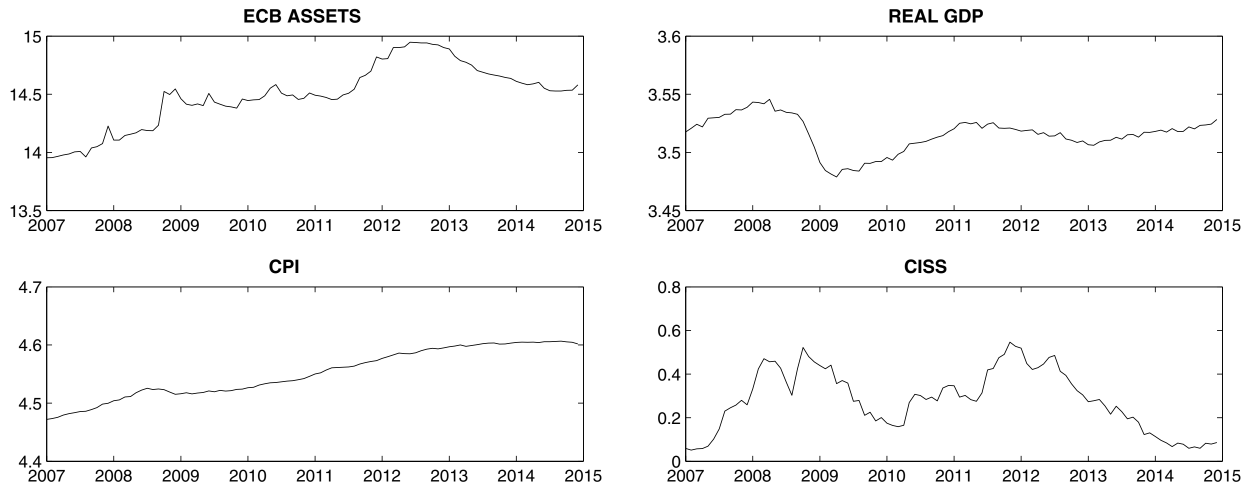


**Fig. 6.** Plot of logarithmic (excl. CISS) time series 2007M1–2014M12 for the euro area.

We include a constant and two lags in the VAR model.[21] The posterior means of the degree-of-freedom parameters of the *t* distributions specified for the components of the error term between 2.3 and 5.6 suggest that identification based on non-Gaussianity of the errors has once again been achieved. We therefore proceed with the formal assessment of the sign restrictions in Boeckx et al. (2017), who assume that a UMP shock increases the balance sheet of the ECB but does not increase financial stress. The restrictions are imposed on impact and in the first month after the shock.

The results reported in the left panel of Table 3 show that the restrictions are supported by the data and two of the shocks ($\varepsilon_{1t}$ and $\varepsilon_{2t}$) receive a relatively high probability (0.17 and 0.25, respectively). The results do not depend on the horizon over which the restrictions are imposed, and we regard $\varepsilon_{2t}$ maximizing the posterior probability as our UMP shock of interest.

An inspection of the impulse responses in Fig. 7 reveals that a unit UMP shock results in an increase in the ECB assets of approximately 0.4% on impact, leads to a significant increase in output and an (insignificant) initial decline in the CISS indicator.[22] The main difference with Boeckx et al. (2017) or the country-level results in Gambacorta et al. (2014) is the response of prices, which they found to be significantly positive persistently, while we find no significant effect. In contrast, the size of the output effect is similar to theirs, lasting less than a year. In quantitative terms we observe a similar output effect for smaller increase in assets. Given that our results are obtained without restrictions, it is interesting to note that also the timing of the output response differs from these studies. Specifically, when the impact response is not ruled out *ex ante*, a positive output response is found to occur earlier than reported in these previous studies. While Boeckx et al. (2017) found output to peak after eight months and Gambacorta et al. (2014) after three months, according to our results output peaks immediately. Taking into account the relatively small sample size and the implicit tight priors of conventional sign-identified SVARs, we also considered a more informative prior distribution. The relatively more informative prior resulted in a positive transient price response after 18 months, whereas the rest of the responses remained unaltered.

We checked the robustness of our results with respect to the monetary base instead of central bank assets as the monetary policy instrument. The right panel of Table 3 shows that the UMP shock is more sharply identified in that the posterior probability of the likeliest shock ($\varepsilon_{2t}$) is greater when monetary base is used, confirming that this shock indeed is our UMP shock of interest. The results from the impulse response analysis are qualitatively and quantitatively robust with respect to the alternative instrument save one interesting exception: a positive price response of 0.08% occurs already after one year even when a non-informative prior is used, i.e. the analysis is solely based on the data.[23]

The latter finding is in contrast to Boeckx et al. (2017), whose price response proved robust to the alternative policy instrument. Nonetheless, the authors point out an important difference between the two variables: the ECB's asset purchases were mostly sterilized and hence are not included in the monetary base. As a consequence, the evolution of the European Monetary Union's monetary base reflects extensions of the long term refinancing operations (LTROs) only (Fawley and Neely, 2013). This can explain our finding

[21] Again, our aim is to keep the model specification of Boeckx et al. (2017) who explain on p.12 that "most criteria suggest a lag length of 3 or even shorter". In line with this, our results are robust to $p = 3$ although the IRFs are somewhat smoother with $p = 2$.

[22] The normalization rule used to compute the posterior probabilities reported in Table 3 generates bimodal posterior distributions for the impulse response functions, resulting in error bands that do not properly reflect parameter uncertainty (see Waggoner and Zha, 2003). For the error bands to be informative about the reliability of the estimates, we report impulse responses computed with a different normalization rule which, however, does not affect the posterior probabilities.

[23] Following (Rogers et al., 2014), we also used the spread between Italian and German 10-yr government bonds instead of central bank assets as a proxy for UMP in the euro area. Even with only four variables and an informative prior, the Bayesian credible sets are wide. However, the median responses are basically the same as in the baseline, although not as clearly as with monetary base as a proxy varaible. These impulse response functions are not shown here to save space but are available upon request.

**Table 3**
Formal assessment of sign restrictions: Euro area.

| | Benchmark model | | Monetary base | |
|---|---|---|---|---|
| Shock | $h = 0$ | $h = 0, 1$ | $h = 0$ | $h = 0, 1$ |
| $\varepsilon_{1t}$ | 0.17 | 0.19 | 0.07 | 0.08 |
| $\varepsilon_{2t}$ | 0.25 | 0.28 | 0.34 | 0.38 |
| $\varepsilon_{3t}$ | 0.06 | 0.05 | 0.04 | 0.05 |
| $\varepsilon_{4t}$ | 0.06 | 0.04 | 0.16 | 0.16 |
| Sum | 0.54 | 0.56 | 0.61 | 0.67 |

Notes: The figures in the table are the posterior probabilities of shock $\varepsilon_{it}$, $i = 1, ...4$ satisfying the sign restrictions that the central bank assets be nonnegative and the CISS be nonpositive for various time horizons, and hence being the structural shock of interest. The figures on the bottom line are the sums of the posterior probabilities. Benchmark model: central bank assets as policy instrument. Monetary base: monetary base as policy instrument.

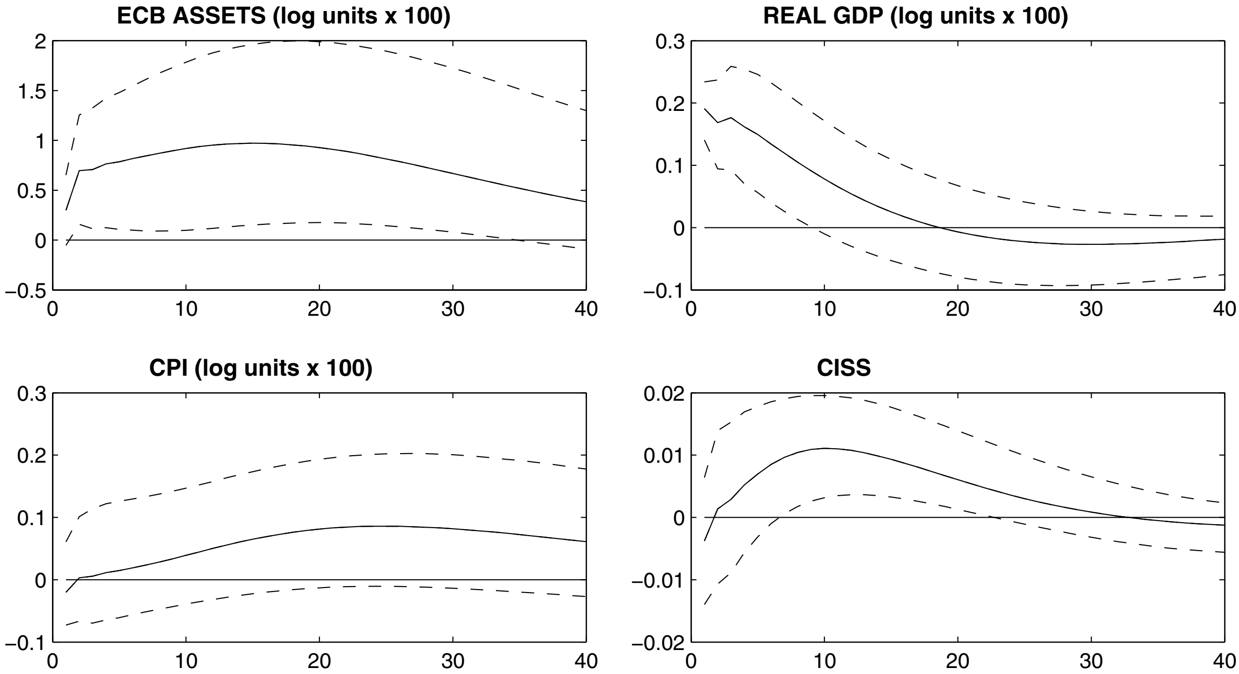


**Fig. 7.** Impulse responses to an expansionary UMP shock: Euro area 2007M1–2014M12. Median responses (solid lines) together with 68% Bayesian credible sets (dashed lines).

that central bank assets and monetary base had a different impact on the price level and suggests that extending the maturity of the longer bank loans showed up sooner in the euro area consumer prices than purchases of private assets or government bonds.

## 4. Conclusions

We have applied a novel Bayesian SVAR identification method due to Lanne and Luoto (2016) to estimate the macroeconomic effects of the Bank of Japan's, the Federal Reserve's and the European Central Bank's balance sheet operations. The procedure exploits non-Gaussianity and independene of the structural error terms to uniquely identify the shocks as in Lanne et al. (2017). Compared to the conventional approach to sign restrictions, the main advantage of our approach is that we are able to formally assess the plausibility of the restrictions used in the literature against the data.

According to our results, the sign restrictions used in the previous literature were mostly supported by the data. However, unlike previous literature, we found an expansionary unconventional monetary policy shock to have different macroeconomic effects in the three geographical areas. Not only the timing, persistence and the significance of the output and price responses varied between currency areas but also the robustness of the results to alternative variables used in the literature. Although we looked at policies that expand each central bank's balance sheet, the policy instrument encompassess different operations for each central bank, which therefore turned out to have different economy-wide effects. As a future research topic, this identification strategy could be used to study the impact of UMP on the stock market. We thank an anonymous referee for the suggestion.

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## Appendix

The data have been retrieved from the FRED database provided by the Federal Reserve Bank of St Louis (https://research.stlouisfed.org/fred2/), from the Bank for International Settlements' (BIS) website (www.bis.org), Bank of Japan's statistics (BOJ) website (http://www.boj.or.jp/en/statistics/index.htm/), CBOE (www.cboe.com) and the ECB Statistical Data Warehouse (ECB) (http://sdw.ecb.europa.eu/).

Series employed in the empirical analysis for Japan:

- Real effective exchange rate (RNJP), BIS
- Core consumer price index (JPNCPICORMINMEI), FRED
- Industrial production (JPNPROINDMISMEI), FRED
- Average outstanding current account balances (BJ'MABS1AN113), BOJ
- 10-year government bond yield (IRLTLT01JPM156N), FRED
- Real GDP (NAEXKP01JPQ661S), FRED
- Government primary balance as % of GDP, OECD (quarterly frequency linearly interpolated to monthly frequency)

Series for the USA:

- Total Federal Reseve bank's assets (WALCL), FRED
- Consumer price index (CPALTT01USM661S), FRED
- CBOE volatitily index (VIX), CBOE
- Industrial production (INDPRO), FRED
- Retail sales (RSXFS), FRED
- Monetary base (AMBSL), FRED
- Total Public Debt as Percent of GDP (GFDEGDQ188S), FRED (quarterly frequency linearly interpolated to monthly frequency)

Series for the euro area:

- Central bank assets for the euro area (ECBASSETS), ECB
- Composite indicator of sovereign stress (CISS.M.U2.Z0Z.4F.EC.SOV_GDPW.IDX), ECB
- Harmonized index of consumer prices (ICP.M.U2.Y.000000.3.INX), ECB
- Real GDP (NAEXKP01EZQ661S), FRED
- Euro area government budget balance as % of GDP, ECB statistical data warehouse (quarterly frequency linearly interpolated to monthly frequency)

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