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# Real Effects of Quantitative Easing at the Zero-Lower Bound: Structural VAR-based Evidence from Japan

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## Abstract

Using post-1995 Japanese data we propose a theory-based sign-restriction SVAR approach to identify monetary policy shocks when the economy is at the zero-lower bound. The identifying restrictions accord with predictions of corresponding DSGE models. Our results show that while a quantitative easing shock leads to a significant but temporary rise in output, the effect on inflation is not significantly different from zero. This suggests that while the Japanese Quantitative Easing experiment was successful in stimulating real activity in the shortrun, it did not lead to any increase in inflation. These results are interesting not only for Japan, but also for other advanced economies where monetary policy is currently constrained by the ZLB.


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# 1 Introduction

We study the real effects of Quantitative Easing (QE) in a structural VAR (SVAR) when the shortterm interest rate is constrained by the Zero-Lower-Bound (ZLB). Using monthly Japanese data since 1995 - a period during which the Bank of Japan's target rate, the overnight call rate, has been very close to zero - and sign restrictions based on corresponding DSGE models, we find that an increase in reserves leads to a significant 0.5 percent rise in industrial production after about two years. This rise lasts for about two years. On the other hand, our results indicate that the same shock has no effect on inflation. Thus our results provide mixed evidence on the successfulness of QE in Japan. Whilst real economic activity does seem to pick up after a QE-shock, this does not seem to affect inflation in such a way that Japan could exit its deflationary period through such a policy shock. However, this conclusion strictly holds only under the usual caveat in SVAR- analysis that the monetary policy shock we consider must be a small one - one that is not allowed to change the policy regime or any other of the structural relations we estimate. Whilst we argue this is precisely the kind of shock that central banks currently inflict on our economies, we should be careful not to conclude that any more aggressive policy changes by central banks to escape the deflationary period of the liquidity trap[1] are doomed to fail.

Our study adds to the existing literature in various important ways. First, focusing specifically on post-1995 Japanese data where the policy rate of the Bank of Japan, the call rate, was virtually zero,[2] allows us to identify a monetary policy shock under liquidity trap conditions. We call such a shock unconventional monetary policy shock or QE-shock for short. Second, including standard macro variables in our VAR allows us to study the effects of such a QE-shock on a broader set of variables than usually studied in the literature on unconventional monetary policy effects. In particular, our approach allows us to study the effects of a QE-shock on real economic activity and on the inflation rate - the two variables of ultimate interest to the central bank. Third, using a sign-restriction approach to identify our QE-shock allows us to remain *agnostic* about whether, how, and when real activity and inflation respond to the QE-shock. Fourth, because our restrictions are firmly grounded in liquidity trap theory we believe they are *credible* in the sense of Sims (1980) and that what we measure in our SVAR is indeed the structural QE-shock we are aiming at. Finally, because shortterm policy rates in the US, the Euro Area, the UK and other economies around the world are currently very close to zero and therefore possibly also constrained by the ZLB, our results shed light on the effects of the currently implemented nonstandard policy measures adopted by the leading central banks in the world.

The rest of the paper is organised as follows: The next section discusses the key findings in the literature on monetary policy effects at the ZLB. Section 3 then briefly discusses key features of the main monetary policy decisions implemented by the Bank of Japan since the stock and

[1]For instance along the lines of Krugman (1998) or Svensson (2003).

[2]See Figure 1.

housing market crashes in the early 90s. Our SVAR and its key identification strategies using sign-restrictions based on liquidity trap theory are explained in Section 4. Results are then presented in Sections 5. Finally, Section 6 concludes.

## 2 Effects of monetary policy shocks at the ZLB in the empirical literature

The effects of monetary policy shocks when monetary policy is *not* constrained by the ZLB has been well documented in the literature. By and large there is a broad consensus that expansionary monetary policy, by lowering the policy interest rate, affects inflation and output positively, but only very sluggishly and only temporarily.[3] This of course is roughly in line with our macroeconomic theories on how monetary shocks affect the real economy under normal times when the interest rate is not constrained by the ZLB. It is somewhat surprising therefore that there is much less empirical evidence on the real effects of monetary policy shocks when monetary policy *is* in fact constrained by the ZLB. One obvious reason might be that most economies until very recently have not been in such a situation and that sample periods to use in estimation would thus be notoriously short. However, it is also true that at least since 2000, when the Fed was fast to lower the Federal Funds rate to very low levels in response to the bursting of the IT-bubble, there has been an important theoretical discussion amongst central bankers as how to avoid liquidity traps and how to escape them once an economy found itself in the trap.[4]
The recent financial crisis has led to renewed interest in the empirical effects of the so-called unconventional monetary policies implemented by the leading central banks. However, most of these studies focus on the effect unconventional policies have on various interest rates or interest rate spreads. They do not study the effects of those policies on other standard macro variables like output or inflation. But these variables of course are the key variables of interest to the central bank and the public and of course important for welfare considerations. The growing body of literature that studies the effects of unconventional monetary policy on such financial market variables include Bernanke et al. (2004), Gagnon et al. (2010), Hamilton and Wu (2010) and Stroebel and Taylor (2009) for the US, Meier (2009) for the UK, ECB (2010) for the Euro Area, and Oda and Ueda (2007) and Ueda (2010) for Japan. Broadly speaking, these studies do find negative effects on yield spreads of unconventional policies, or more precisely of announcements of such policies, in the sense that the yields of various assets do tend to decline thereby narrowing the spread to the corresponding riskless rate. However, these effects are generally found to be rather small. For instance, Hamilton and Wu (2010) find that a purchase of 400 billion US dollars in 10-year US Treasury Bonds would lead to a 14 basis points fall in the 10-year yield.

[3] Compare Christiano et al. (1998). But note different identifying restrictions do in fact lead to slightly different results, compare Uhlig (2005) and Lanne and Lütkepohl (2008).

[4] See e.g. Bernanke (2002), Bernanke and Reinhart (2004), Krugman (1998) or Svensson (2003).

Gagnon et al. (2010) find that the same policy measure would lower longterm yields by 20 basis points. Meier (2009) estimates that the Bank of England's QE-related asset purchases lowered gilt yields by around 40-100 basis points.

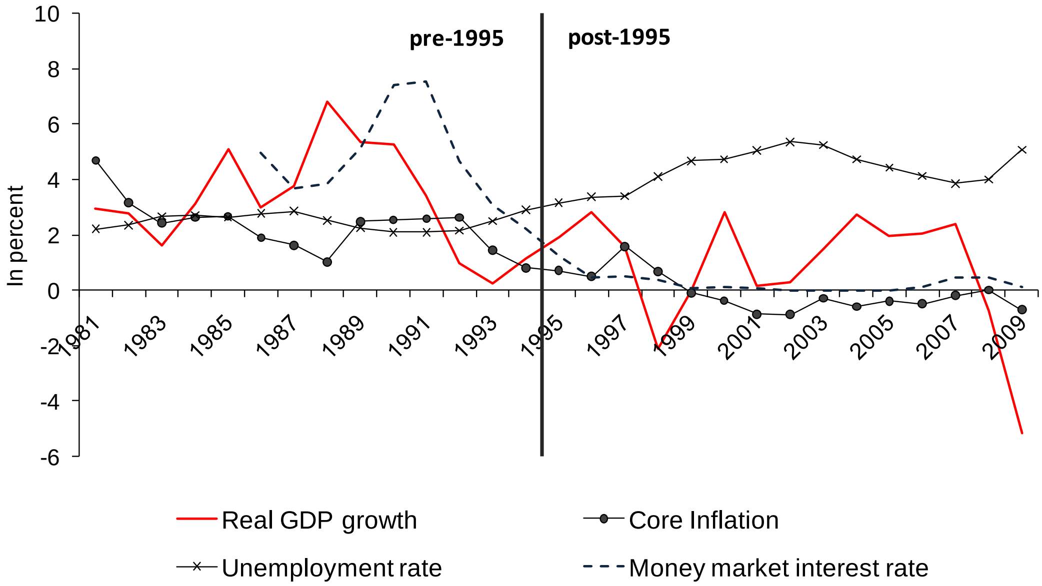
It is important to note that the theoretical impact of such a policy announcement on longterm yields is far from clear. Most theoretical studies, such as e.g. Doh (2010), refer to some kind of imperfect substitutability between assets and explain the expansionary effect of such a policy decision by arguing that the purchase of longterm bonds by the central bank will naturally lower longterm bond yields, because the central bank buys those bonds at a higher price than the market. This lower longterm yield on government bonds then feeds through - via portfolio shifts (Meltzer, 1995) - to other asset markets, like the corporate bond market and the stock market making longterm financing for investment and durable goods cheaper thereby stimulating aggregate demand. Being constrained by the ZLB the traditional interest channel which normally lowers longterm yields through the expectations hypothesis does not function anymore and the central bank circumvents this by directly intervening in the market for longterm bonds. This argument is partly supported by the empirical evidence of the above mentioned studies. However, theoretically it is not clear that longterm yields are indeed supposed to fall after such a policy announcement. Indeed, if market participants believe the central bank intervention is successful in stimulating the economy by increasing aggregate demand, inflation and real rates are likely to rise in the future. Thus inflationary expectations as of today should rise and longterm nominal yields should in fact rise. In other words, the effectiveness of such a policy move might instead be seen by rising longterm yields, not by falling yields.

We therefore argue that it is important to focus on the effects on the real economy when analyzing unconventional monetary policy. So far, the correpsonding empirical evidence is rather scarce and a consensus on the effectiveness of these measures in terms of the real economy has not yet been reached. Studies using sign restrictions to identify unconventional monetary shocks include Kamada and Sugo (2006), Baumeister and Benati (2010) and Peersman (2010). While Baumeister and Benati (2010) find some significant real effects of quantitative easing in different countries including Japan, results reported by Kamada and Sugo (2006) are less optimistic. Both studies rely, however, on relatively restrictive identification schemes. Peersman (2010) finds that unconventional shocks can in principle affect macroeconomic variables in the Euro area; the responses of output and prices are, however, much more delayed compared to standard policy measures. Using a large-scale Bayesian VAR model Lenza et al. (2010) report some significant effects of unconventional monetary shocks on macroeconomic variables in the Euro area but focus on policy measures that actually affected the interest rate spread. Finally, Chung et al. (2011), using a set of structural and time series statistical models, find that asset purchases by the Fed have been successful at mitigating the macroeconomic costs of the ZLB in the US.

# 3 Monetary Policy in Japan since the late 1980s

This section briefly summarises key monetary policy developments in Japan since the late 1980s/early 1990s - in other words since the bursting of the Japanese stock and real estate bubbles. We only sketch key developments, for a thorough discussion please consult Mikitani and Posen (2000), Ugai (2007) and Ueda (2010). We divide this period into pre- and post-1995 based on the behaviour of the Bank of Japan's target interest rate, the call rate, which has been lowered by the Bank of Japan to virtually zero during 1995. Figure 1 shows key macroeconomic variables for Japan since 1981.

Figure 1: Key macroeconomic variables in Japan since 1981. The thick line indicates 1995.

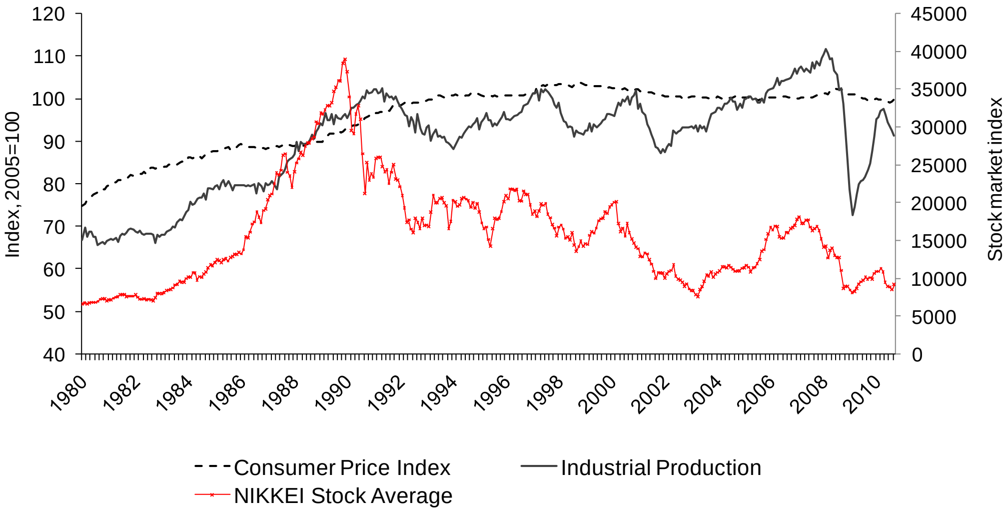



The figure reveals the widely documented behaviour of the dramatic fall in real GDP growth rates after the bursting of the asset prices bubbles in 1990/91. Whilst GDP grew in the pre-1991 period by an average rate of 3.9 percent per year, it slowed down to only 0.8 percent post-1991. This of course is the numerical basis for the well-known label "Japan's lost decade." Meanwhile the notoriously low Japanese unemployment rate has more than doubled while the core inflation rate has steadily trended below zero since 2000. The following subsections give some more details to Japanese monetary policy pre- and post-1995.

### 3.1 The bursting of the bubbles and delayed monetary policy reaction

The bursting of the stock market bubble can be seen in Figure 2. The stock market was rising dramatically until around 1990. The figure shows that this went together with rapid increase in industrial production under fairly low and constant rates of inflation (compare also Figure 1).

Figure 2: Industrial production, Consumer Price Index and NIKKEI stock index since 1980. The stock market bubble burst in 1990.

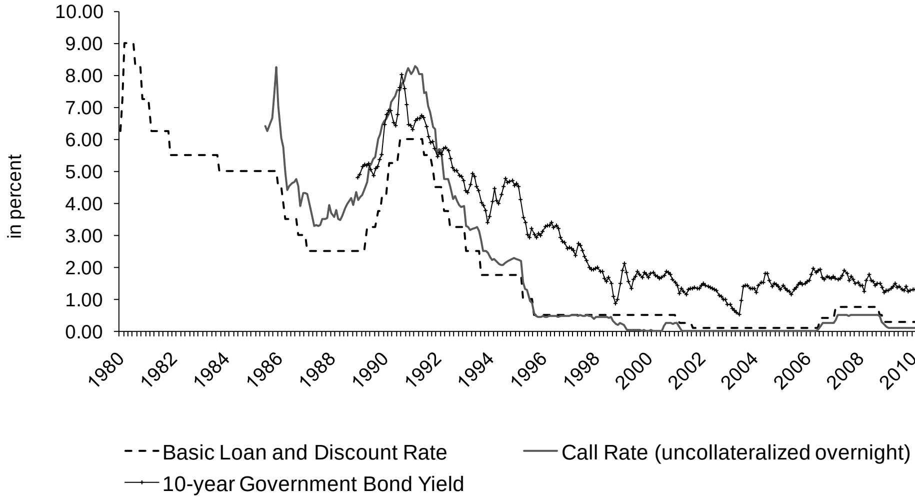



Realising that the elevated stock and land prices seemed out of touch with fundamentals the Bank of Japan did in fact continuously increase the call rate (compare again Figure 1). Optimism turned into pessimism around 1990/91 and both stock and land prices started falling rapidly. It is nowadays widely agreed that the initial response of the Bank of Japan to the bursting of the asset price bubbles was too slow and not aggressive enough (Jinushi et al., 2000). In fact, Figure 1 shows that the call rate was high until 1992/3 and then only lowered very gradually until it reached 0.5 percent in the last quarter of 1995.

### 3.2 Post-1995

But even with its key policy rate at 0.5 percent and therewith close to zero, the Bank of Japan was very slow in implementing unconventional expansionary policy measures. In fact, only in 1999 did the Bank of Japan officially introduce its so-called zero interest rate policy (ZIRP) when it lowered the call rate to 0.03 percent (see Figure 3). It also tried to steer market expectations by adding commitments to its policy statements indicating that it would keep the call rate low for longer time.

Figure 3: Key Bank of Japan interest rates since 1980. The call rate has been the Bank of Japan's main policy rate between the mid-80s and early 2000.



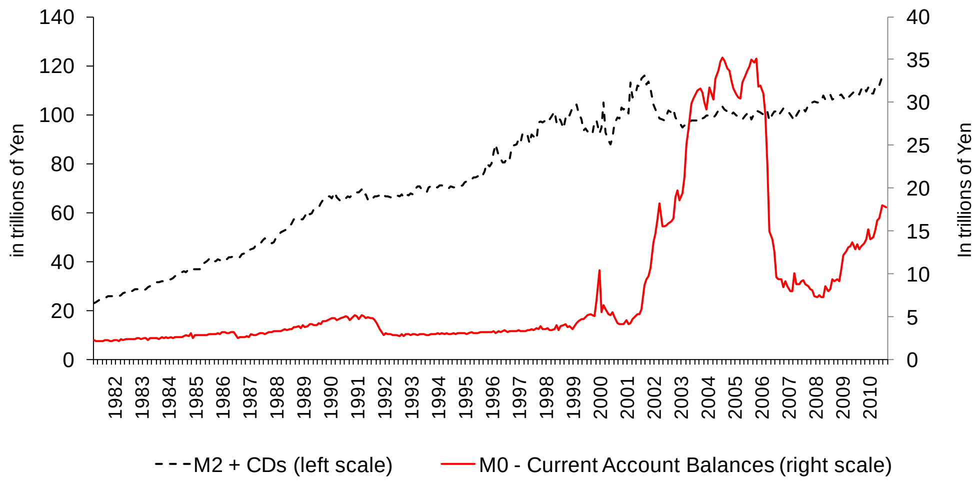

When the Japanese economy started to recover slightly in 2000 with real GDP growth of 2.8 percent, the call rate was raised to 0.25 percent and ZIRP was officially ended. However, the worldwide economic recession following the bursting of stock market bubbles in response to the IT-bubble led to renewed macroeconomic problems in Japan. This time the Bank of Japan introduced a more aggressive policy programme. From March 2001 until March 2006 it implemented the so-called "Quantitative Easing Policy" (QEP) which consisted basically of three elements: (i) the operating target was changed from the call rate to the outstanding current account balances held by banks at the Bank of Japan, (ii) to commit itself to continue providing ample liquidity to banks until inflation stabilised at zero percent or a slight increase, and (iii) to increase the amount of outright purchases of longterm Japanese government bonds.[5]

The monetary development and the effect of the Bank of Japan's QEP measures can be seen in Figure 4. We plot that part of the monetary base that is the current account holdings of banks at the Bank of Japan, in other words these are bank reserves held at the central bank. The figure clearly shows the enormous increase in reserves during the QEP period and later again when the recent financial crisis hit. At the same time we see that the growth rate of M2 and Certificates of Deposits (CDs) steadily slowed since 1980.

Having these macroeconomic and monetary developments in mind we next want to present our identification strategy based on the reasonable assumption that the Bank of Japan since 1995 did not conduct its monetary policy through the call rate anymore - which was constrained by the ZLB - but by changing the reserve holdings of banks at the Bank of Japan.

[5]See the authorative survey by Ugai (2007) for more details.

Figure 4: Monetary aggregates in Japan. Quantitative Easing was implemented between 2002 and 2006.




# 4 Identification of structural shocks in a sign-restriction VAR

## 4.1 Basic VAR model

### 4.1.1 Benchmark specification

To analyse the effects of monetary policy on economic activity and the price level at the ZLB, the following reduced-form VAR model is estimated:

$$Y_t = c + A(L)Y_{t-1} + u_t, \tag{1}$$

where $c$ is a vector of intercepts,[6] $Y_t$ is a vector of endogenous variables, $A(L)$ is a matrix of autoregressive coefficients of the lagged values of $Y_t$ and $u_t$ is a vector of residuals. In this model, the reduced-form error terms are related to the uncorrelated structural errors $\epsilon_t$ according to:

$$u_t = B^{-1}\epsilon_t. \tag{2}$$

[6]Following Uhlig (2005) and Peersman (2010) we do not include a time trend in our benchmark VAR model. In fact, including a trend would require an adjustment of the prior used in the Bayesian estimation (Uhlig, 1994). However, our main results are insensitive to linearly detrending the variables prior to estimation. Results are available upon request.

In our benchmark regression we include the following four macroeconomic variables in the VAR-system:

$$Y_t = [P_t,\ IP_t,\ RES_t,\ LTY_t], \tag{3}$$

where $P_t$ denotes the core consumer price index and $IP_t$ indicates the Japanese industrial production index. Moreover, we include reserves ($RES_t$)[7] and the 10-year yield of Japanese government bonds ($LTY_t$) in the set of regressors. The VAR model is estimated by means of Bayesian methods using monthly data over the period January 1995 to September 2010. In the benchmark case, six lags of the endogenous variables are included in the estimation, which seems to be sufficient to capture the dynamics of the model[8]. Except for the longterm yield, all variables are seasonally adjusted and incl uded as log-levels.[9] We do not include a measure of Japanese money supply such as M2 + CDs since a number of empirical studies conclude that the relationship between money supply and economic activity or prices disappeared in the course of the 1990s (see e.g. Miyao (2005)). This is in line with the descriptive evidence given in Figure 4 showing that during the period of massive quantitative easing in the early 2000's M2 + CDs did in fact not increase with the monetary base. Thus, the money stock is not likely to be an important variable with respect to the transmission of monetary policy during the 1990's and 2000's. A detailed description of the data is given in Appendix A.

#### 4.1.2 Additional specifications

Within the theoretical literature on monetary policy at the ZLB the role of the exchange rate in the transmission of unconventional policy has been stressed by a number of studies (Orphanides and Wieland, 2000; Coenen and Wieland, 2003; McCallum, 2000). These models usually imply a real depreciation of the domestic currency following a base money injection due to portfolio rebalancing effects. In order to shed more light on the role of the exchange rate at the ZLB we estimate an additional specification including the real effective exchange rate of the Yen against other currencies ($EX_t$):

$$Y_t = [P_t,\ IP_t,\ RES_t,\ LTY_t,\ EX_t]. \tag{4}$$

We argue above that Japan has been at the zero lower bound during the whole sample period under consideration. In the course of 1995, the call rate has been reduced to 0.5% severely mitigating its importance as a policy instrument. However, officially, the reserve target has

[7]As a measure of reserves we consider outstanding current account balances held at the Bank of Japan, which is a base money component.

[8]While different lag length criteria lead to different suggestions concerning the number of lags to include, all of them tend to propose an even shorter lag length. Our main results are, however, robust to varying the lag length.

[9]According to Sims et al. (1990) this leads to consistent parameter estimates even in the presence of unit roots.

replaced the interest rate as the main policy instrument in 2001 only with the introduction of QEP. Therefore, we additionally estimate a further specification that augments the model given by equation (4) by adding the call rate, denoted by $R_t$. If indeed this variable has been unimportant as a policy instrument even prior to 2001 we should not observe any significant reaction of the nominal interest rate to the monetary shock and the remaining results should remain unchanged. Thus, we additionally estimate the following specification:

$$Y_t = [P_t,\ IP_t,\ RES_t,\ LTY_t,\ EX_t,\ R_t]. \tag{5}$$

## 4.2 Identification of structural shocks

As in Uhlig (2005), Canova and Nicolo (2002) and Peersman (2005) we impose sign restrictions on the impulse response functions to identify an unconventional monetary shock. In order to prevent that other disturbances enter the identified unconventional monetary shock we additionally identify two traditional shocks; a positive demand and a positive supply shock. To be able to distinguish between the responses to the respective shocks, we require these disturbances to be orthogonal to the monetary shock.[10] Using this specification we make sure that the expansionary monetary shock is not confused with disturbances related to business cycle fluctuations. In contrast to identification strategies based on Cholesky or Blanchard-Quah decompositions, the sign restriction approach explicitly incorporates assumptions that are often used implicitly allowing a more transparent procedure. Moreover, imposing zero restrictions on contemporaneous or long-run impulse responses is avoided. While zero-restrictions on contemporaneous interactions may not hold in reality (Faust, 1998), long-run restrictions may be biased in small samples (Faust and Leeper, 1997). The sign restriction approach is implemented by taking draws for the VAR parameters from the Normal-Wishart posterior[11], constructing an impulse vector for each draw and calculating the corresponding impulse responses for all variables over the specified horizon.[12] In particular, the reduced-form residuals $u_t$ are related to the structural shocks according to equation (2) above with $B = W\Sigma_\epsilon^{1/2}Q$, where $W\Sigma_\epsilon^{1/2}$ is the Cholesky factor obtained from the Bayesian estimation of the VAR model for each of the 1000 draws, and Q is an orthogonal matrix with $QQ' = I$. To generate $Q$, we draw a random matrix U from an N(0,1) density and decompose this matrix using a QR decomposition. For each of the 1000 Cholesky factors we

[10] Mountford and Uhlig (2009) show how the identification setup in Uhlig (2005) can be extended to control for additional shocks. Our estimation strategy closely follows their approach.

[11] Uhlig (1994) offers a detailed discussion on the Normal-Wishart prior. His observation 5 states that in the case of nonexplosive roots, i.e. $|\rho| \leq 1$, using a flat Normal-Wishart prior is equivalent to using a *critics prior* in practical applications. While the Normal-Wishart prior puts equal weights on all values of $\rho$, the critics prior emphasizes larger values of $\rho$ and is thus consistent with the prior belief that unit roots are unlikely.

[12] Estimation was performed on the basis of Fabio Canova's SVAR Matlab codes, which can be downloaded from his website http://www.crei.cat/people/canova/.

repeat drawing U until we find a matrix generating responses to the respective shocks that are in line with the sign restrictions we impose. Additionally, as in Peersman (2010) we employ exact zero restrictions on the contemporaneous impact matrix B. Using a mixture of sign restrictions and zero restrictions on selected impact responses allows us to improve identification of the structural shocks and thus to enhance the interpretation of the respective impulse response functions by exploiting additional economic information (Kilian, 2009).[13] The impulse response functions $r^k_{ijt}$ of variable $j = 1, ..., 5$ to shock $i = 1, 2, 3$ at horizon $t = 1, ..., 60$ constructed using model $k = 1, ...1000$ (where k indexes the different values of Q) are then summarised by computing the median over $k$ of $r^k_{ijt}$.

It is important to note, however, that solely reporting the median of all admissible impulse responses may be problematic, especially if several shocks are identified at the same time (Fry and Pagan, 2007). First, since the median *over k* summarises information obtained from different models, the reported structural impulse response functions may be hard to interpret. Second, and related, since two shocks may be generated from two different models, the structural disturbances are not necessarily orthogonal. We account for these issues by following Fry and Pagan (2007) and additionally reporting impulse responses generated by *one* model Q; the model that leads to impulse responses that are as close to the median over $k$ of $r^k_{ijt}$ as possible. This model is found by first standardizing the impulse responses $r^k_{ijt}$ by subtracting off their median and divide by their standard deviation over the 1000 models satisfying the sign restrictions. The standardized impulse responses are then grouped into a vector $\phi^k$ for each value $Q^k$. We subsequently choose the model that minimizes $\phi^{k'}\phi^k$ and report the corresponding impulse responses; see Section 5.5.

### 4.3 Demand, supply and monetary shocks at the ZLB in the theoretical literature

As has been stressed above, the existing empirical VAR literature on the transmission of unconventional monetary policy is rather scarce and thus a broad consensus about the identification of a QE-shock at the ZLB is yet to be reached. Moreover, it is not clear ex ante whether the usual identifying restrictions for aggregate demand and supply shocks are still valid if the interest rate is close to zero. In particular, the main impediment to disentangling the monetary shock from business cycle disturbances at the ZLB is the fact that the interest rate cannot move following either shock. Nevertheless, we show below that it is still possible to derive a clear identification

[13]There have been alternative approaches to improve the interpretation of structural impulse response functions generated by using a sign restriction approach by using additional information to identify the shocks. One proposition has been to impose restrictions on cross correlations of impulse responses, thereby narrowing down the set of admissible responses (Canova and Nicolo, 2002; Canova and Paustian, 2010). Furthermore, Uhlig (2005) employs a penalty function in order to use more identifying information. Finally, identifying an oil market VAR, Kilian (2009) show how to exploit additional information concerning the short-run oil supply elasticity to improve a pure sign restriction identification scheme. Checking sensitivity of our results to such alterations is beyond the scope of our analysis, so we leave this for future work.

setup using a mix of exact zero- and sign-restrictions that are implied by theoretical models. Thus, as a first step, we take a closer look at the theoretical DSGE literature concerned with the modelling of the ZLB before deriving our identifying restrictions.

One approach within the theoretical literature concerned with modelling the ZLB has been to calibrate (McCallum, 2000; Orphanides and Wieland, 2000) or estimate (Coenen and Wieland, 2003) open-economy macromodels allowing for zero interest rates. Allowing the quantity of base money to affect output and inflation even if the interest rate is zero these models imply that liquidity injections lead to an increase in output and inflation, respectively, given that these policy measures are sufficiently aggressive. The particular channel that these models rely on is the portfolio balance effect along the lines of Meltzer (1995, 2001) and Mishkin (2001) implying a rebalancing of investors' portfolios towards, for instance, foreign assets following a base money injection. The resulting real exchange rate depreciation in turn helps to increase output and prices. Relative to this class of macromodels, more microfoundation is provided by a growing DSGE literature aiming at a characterization of optimal monetary policy in a situation of zero interest rates including Eggertsson and Woodford (2003), Jung et al. (2005), Eggertsson (2006) and Nakov (2008). This stream of literature stresses changing expectations of future monetary policy as the main channel of transmission of base money injections instead of a direct quantity effect. Thus, if a base money injection is successful in that it leads to lower expected interest rates in the future and increases inflationary expectations as of today, it may increase output and inflation. While these different approaches focus on diverging channels underlying the effect of quantitative easing, the outcome is similar: a rise in the reserve component of the monetary base in a situation of zero interest rates should lead to a non-negative effect of output and prices.

Yano (2009) presents a New Keynesian DSGE model under liquidity trap conditions that is estimated using Japanese data and thus offers more insights on the reaction of output, inflation and the interest rate following different business cycle shocks at the ZLB. In particular, the model implies that prices and output move in the same direction following a demand shock and in opposite directions after a supply shock. The interest rate stays fixed at zero after both shocks.

Finally, Eggertsson (2010) provides a DSGE model in which the ZLB is the outcome of an exogenous negative shock moving the economy away the from the zero-inflation natural rate steady state and into the ZLB. Again, in this model a positive aggregate demand shock increases output and inflation. However, in contrast to the responses implied by the model of Yano (2009) an aggregate supply shock also leads output and inflation to move in the same direction. More specifically, a positive supply shocky further enhances deflation at the ZLB, which further increases the real rate of interest and hence depresses aggregate demand.

## 4.4 Identifying sign-restrictions

Using the implications of these theoretical models we will now derive our identifying set of sign-restrictions. As far as the identification of the business cycle shocks are concerned we will take into account the diverging predictions of the DSGE models of Yano (2009) and Eggertsson (2010), respectively, by implementing restrictions implied by the former in our benchmark identification, while the restrictions in line with the latter model are used in an alternative identification scheme.

### 4.4.1 Benchmark identification

We first describe our benchmark identification scheme for our benchmark specification. Restrictions are binding for twelve months following the shock[14], while the zero restrictions are imposed on impact only. A summary of the restrictions considered for the benchmark model is provided in table 1.

Table 1: Identifying sign restrictions - benchmark identification

| | Response to | | | Restriction horizon |
|---|---|---|---|---|
| Variable | Demand shock | Supply shock | QE shock | |
| CPI | $> 0$ | $< 0$ | $0, \geq 0$ | $k = 0, ..., 11$ |
| Ind. production | $> 0$ | $> 0$ | $0$ | $k = 0, ..., 11$ |
| Reserves | | | $\geq 0$ | $k = 0, ..., 11$ |
| Longterm yield | | | | |

As table 1 shows, to identify an aggregate demand shock we restrict output and prices to move in the same direction; both variables are assumed to increase following a positive demand disturbance. For an aggregate supply shock we impose that output and prices move in the opposite direction. It can easily be seen that these assumptions are sufficient to disentangle these respective shocks. As has been explained above, these restrictions are in line with the predictions of DSGE models explicitly modeling the zero lower bound, such as Yano (2009). Moreover, similar restrictions are implied by standard DSGE models (Straub and Peersman, 2006; Canova and Paustian, 2010) and are also imposed in more traditional VAR studies (Peersman, 2005; Canova et al., 2007). The unconventional monetary shock is identified by restricting reserves not to decrease following the shock. Furthermore, we follow the usual approach in the VAR literature assuming a lagged impact of a monetary shoc k on output and prices; the contemporaneous impact on these variables is restrained to zero. Similar zero restrictions have also been used by Peersman (2010). Additionally, however, we assume a non-negative response of the price level to the QE-shock. As outlined above, this is in line with a wide range of theoretical models assuming

[14] A similar restriction horizon is used by e.g. Scholl and Uhlig (2005).

a zero lower bound (Coenen and Wieland, 2003; Eggertsson and Woodford, 2003; Eggertsson, 2006).

Because the central question assessed in this paper is concerned with the effectiveness of unconventional monetary policy measures on the real economy at the zero-lower bound, which is the ultimate concern of central banks facing a liquidity trap situation, we leave the response of industrial production to a QE-shock unrestricted. Moreover, we abstain from restricting the 10-year government yield. As discussed in the last section, the effects of quantitative easing on longterm yields are theoretically not clear; observing *rising* yields following a base money expansion may be possible as a consequence of increasing inflation expectations or increasing risk premia. In this sense our identification scheme can be considered "agnostic" in that we let the data speak concerning the effects of an unconventional monetary shock on the real economy and longterm interest rates. Crucially, the contemporaneous zero restrictions following a QE-shock imposed on the real variables ar e sufficient to disentangle the unconventional monetary shock from the business cycle disturbances (Peersman, 2010).

The set of identifying restrictions for our alternative specifications given in equations (4) and (5) can be found in table 2. The restrictions imposed on the CPI, industrial production and reserves are the same as those used for the benchmark specification above. In contrast to Kamada and Sugo (2006) we abstain from restricting the exchange rate. Leaving the response of the exchange rate unrestricted allows us to let the data speak concerning the effect of the QE-shock on this variable and thus its role in the transmission of unconventional policy.

Table 2: Identifying sign restrictions - benchmark identification (extended specifications)

| | Response to | | | Restriction horizon |
|---|---|---|---|---|
| Variable | Demand shock | Supply shock | QE shock | |
| CPI | $> 0$ | $< 0$ | $0, \geq 0$ | $k = 0, ..., 11$ |
| Ind. production | $> 0$ | $> 0$ | 0 | $k = 0, ..., 11$ |
| Reserves | | | $\geq 0$ | $k = 0, ..., 11$ |
| Longterm yield | | | | |
| Exchange rate | | | | |
| Call Rate | 0 | 0 | 0 | |

As far as the shortterm interest rate is concerned, because we identify an expansionary shock, the call rate should not increase following the QE-shock. Moreover, since we estimate the model for a zero-lower bound situation, the nominal interest rate cannot fall further. We implement this simply by assuming a zero reaction to the QE-shock on impact and let the data speak concerning the response in the remaining periods. If we correctly specified the model for the zero lower bound, we should not observe any significant reaction of the call rate following the shock.[15] Moreover, we assume the call rate not to react on impact following a demand and a supply shock, respectively, and leave the sign of the response open thereafter. Again, these restrictions are in line with recent DSGE models of Yano (2009) and Christiano et al. (2009).

### 4.4.2 Alternative identification scheme

In order to check whether our results concerning the QE-shock are still valid when we account for the somewhat diverging effects of a positive supply shock at the ZLB predicted by Eggertsson (2010) we try to implement these restrictions in an alternative setup, summarised in table 3. Since both the demand and supply shocks should now induce output and prices to move in the same direction, we cannot easily disentangle the two shocks. To deal with this problem we propose another way to disentangle shocks using the different slope properties of the aggregate supply and demand equations in the model. In particular, in the model of Eggertsson (2010) the AD-curve will always be steeper than the AS-curve and thus a positive demand shock will lead to a proportionately larger impact on the value of output versus inflation than a positive supply shock. Thus we restrict the response of this ratio to be larger than one in absolute value for the demand shock, and less than one for the supply shock. At the same time, a positive demand shock is assumed to lead to a positive reaction of both output and prices, while a positive supply shock is restricted to lower these variables. The QE-shock is identified as before and can again be disentangled from the other shocks by imposing the exact zero-restrictions.

Table 3: Identifying sign restrictions - alternative identification

| | Response to | | | Restriction horizon |
|---|---|---|---|---|
| Variable | Demand shock | Supply shock | QE shock | |
| CPI | $>0$ | $<0$ | $0, \geq 0$ | $k=0,...,11$ |
| Ind. production | $>0$ | $<0$ | $0$ | $k=0,...,11$ |
| Reserves | | | $\geq 0$ | $k=0,...,11$ |
| $\lvert\frac{\Delta\hat{y}}{\Delta\pi}\rvert$ | $>1$ | $<1$ | | $k=0,...,11$ |

[15] Alternatively, we could implement this by means of a "near-zero" restriction specifying the response of the call rate to stay "reasonably close" to zero following the unconventional monetary shock: $-\epsilon \leq r_{R,QE,t} \leq \epsilon$, where $r_{R,QE,t}$ denotes the impulse response of the call rate following a quantitative easing shock at horizon $t = 1, ...K$ and $\epsilon$ denotes a threshold set close to zero. Our main results are robust to using this identification scheme. However, since it is much more restrictive we prefer to use zero restrictions on impact only.

# 5 Results

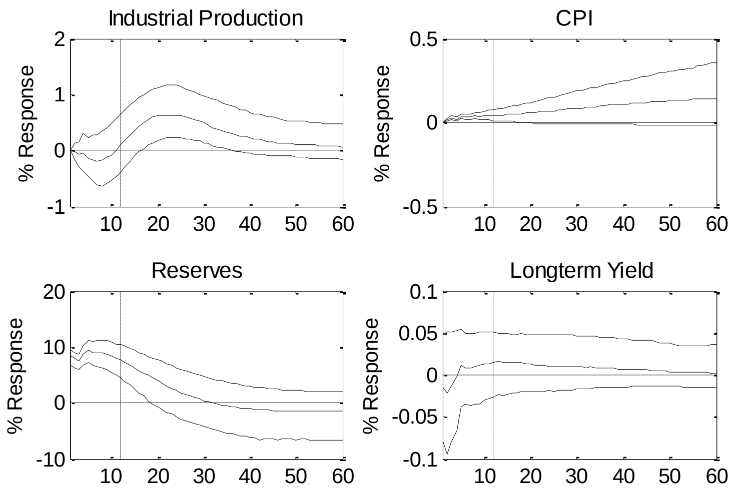
## 5.1 Impulse response analysis - benchmark regression

Figures 5, 6 and 7 show the impulse responses to the three shocks based on the benchmark identification and specification schemes explained above. Figure 5 shows the responses to our unconventional monetary policy shock. In the figure, the inner lines denote the median impulse responses from a Bayesian vector autoregression with 1000 draws, while the outer lines indicate one-standard error confidence bands. The response of reserves has been restricted not to decrease following the shock, so the immediate positive response is not surprising by construction. In particular, reserves rise by up to 10% and stay significantly above the zero line for quite a bit longer than preset. As restricted, CPI does not react on impact and responds positively thereafter. It can be seen that the response of the price level is rather weak staying below 0.1%. Moreover, the response becomes insignificant soon aft er the end of the restriction horizon. Hence, all in all, the response of the price level to the QE shock is temporary and rather weak implying that the rate of inflation also reacts only temporarily and weakly.

Crucially, the main variable of interest, industrial production, has been left unrestricted except for the contemporaneous zero restriction. It can be seen in the figure that an expansionary QE-shock leads to a significant increase of industrial production by about 0.5% after about 18 months. This response is temporary and fades after about three years. Thus, our VAR-based evidence indicates that unconventional monetary policy by increasing reserves can in fact increase economic activity temporarily. Finally, in contrast to some previous studies we did not restrict the response of the longterm government bond yield since its reaction following a QE-shock shock is theoretically unclear. In fact figure 5 shows an initial negative but insignificant reaction of this variable. Hence, our result does not support the view that QE works by lowering longterm rates. This finding confirms the validity of our agnostic approach with respect to the longterm yield; ide ntification of an unconventional shock by an explicit negative restriction on this variable may lead to misleading results.

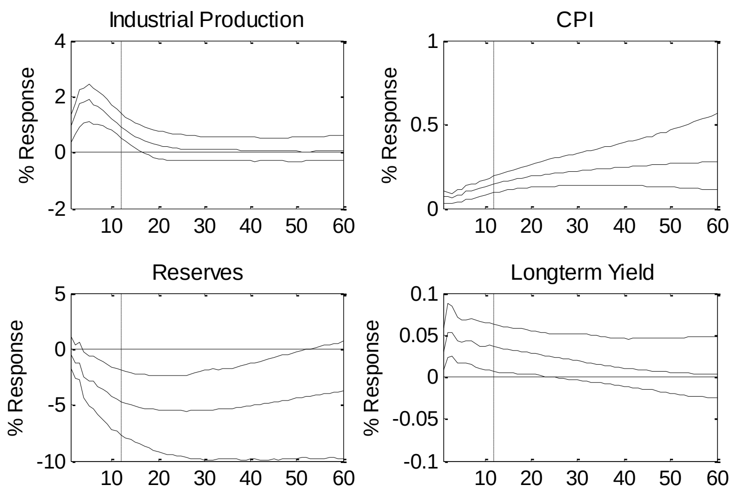
All in all, the results presented in figure 5 suggest that the quantitative easing strategy adopted by the Bank of Japan in the early 2000's in a situation of near-zero interest rates has been successful in stimulating real economic activity, at least in the short run. However, our results also show that the Bank of Japan's second main goal motivating this policy, namely to permanently raise inflation and to eventually bring an end to Japan's deflationary episode, does not seem to have been achieved by the QE-shock. In fact, core CPI only rises to the extent that we restrict it and is just on the verge of insignificance thereafter. More importantly, whilst it may still be argued that core CPI rises somewhat after a QE-shock, the same cannot be said for its rate of change, the core inflation rate. Hence, our benchmark results provide mixed evidence for the effectiveness of the Bank of Japan's QE-policy.

Figure 5: Impulse responses to a QE-shock - benchmark identification and model



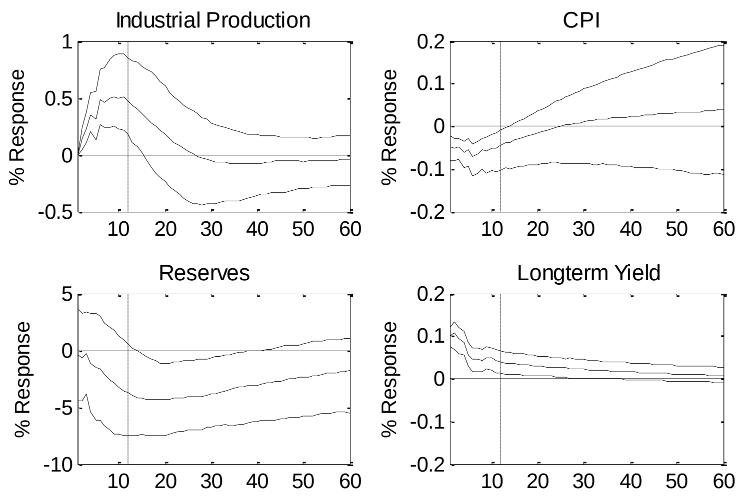

The figure displays responses to a QE-shock as identified above. The inner lines denote the median impulse responses from a Bayesian vector autoregression with 1000 draws, while the outer lines indicate one-standard error confidence bands. The vertical dotted lines indicate the restriction horizon.

Figure 6: Impulse responses to a demand shock - benchmark identification and model




The figure displays responses to a demand shock as identified above. The inner lines denote the median impulse responses from a Bayesian vector autoregression with 1000 draws, while the outer lines indicate one-standard error confidence bands. The vertical dotted lines indicate the restriction horizon.

Figure 7: Impulse responses to a supply shock - benchmark identification and model




The figure displays responses to a supply shock as identified above. The inner lines denote the median impulse responses from a Bayesian vector autoregression with 1000 draws, while the outer lines indicate one-standard error confidence bands. The vertical dotted lines indicate the restriction horizon.

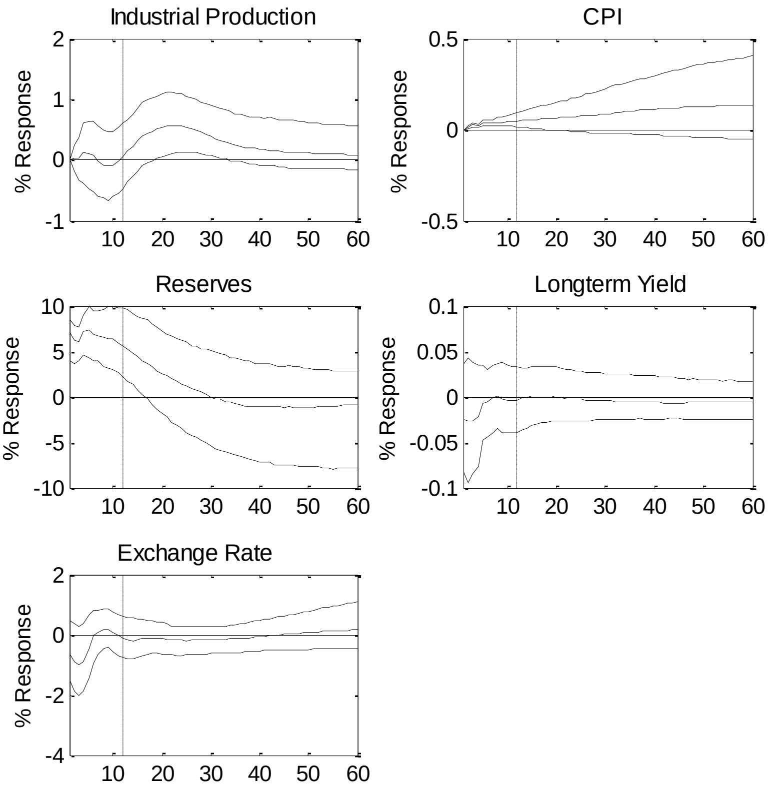
The impulse response functions for the demand and supply shocks are shown in figures 6 and 7 respectively. These two shocks are mainly identified for the purpose of controlling for other business cycle disturbances with which the QE-shock might be confused. Because most variables have been restricted we only briefly discuss the results here. Following a demand shock, industrial production and the CPI are restricted to rise. Hence the initial increase in these variables is not surprising. However, note that CPI rises significantly over the entire response horizon we study. And also industrial production does stay significantly positive for somewhat longer than the restriction horizon. Turning to the responses of reserves and longterm yields to the demand shock, we find reserves falling significantly and in a hump-shaped pattern by around 5% and the longterm yield rising significantly on impact by 0.05%. These results can be explained by reserves being run down by banks needing to increase lending in response to the positive demand shock and by inflationary pressures bidding up longterm yields. Figure 7 finally shows the impulse response functions following a supply shock. Again, the initial increase in industrial production and decrease in CPI are by construction. It is interesting to see that the impulse responses become insignificant soon after the restriction horizon. At the same time the responses of reserves and yields are quite similar to those to the demand shock.

### 5.2 Alternative specifications

We now focus on the responses to the QE-shock only and discuss the results of the other two specifications. Figure 8 shows the impulse responses to the QE-shock resulting from the specification including the exchange rate. This serves both as a robustness test and because we want to study whether the exchange rate follows an interesting pattern that might help explaining the transmission mechanism of the QE-shock. As can be seen in the figure, the qualitative results do not change after controlling for business cycle disturbances. Industrial production still rises by up to 0.5%; however, error bands are somewhat wider. As in the benchmark case, the response of the consumer price index becomes significant after a while, however, the delay is somewhat longer. The responses of the other variables are very similar to those in the benchmark case. Thus, our extended identification scheme does not change our main conclusion that while production and prices could be increased temporarily by quantitative easing measures, the longterm inflation environment has not be en affected by this policy. Moreover, the response of the longterm yield is robust to this extended specification confirming that longterm rates do in fact not fall after such an unconventional shock. However, adding the exchange rate does not help us in shedding more light on the specifics of the transmission mechanism. In fact, the real effective exchange rate is insignificant over the entire horizon. Its dynamics does however point to a gradual depreciation.[16]

[16]The exchange rate first seems to appreciate (an appreciation is a fall in the impulse response), but then to gradually depreciate.

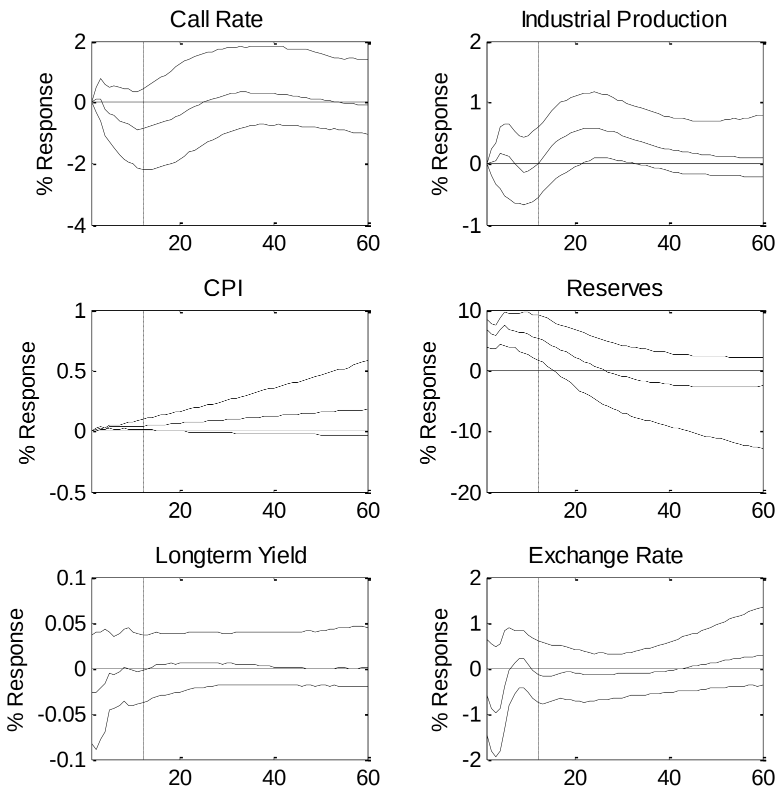
Figure 8: Impulse responses to a QE-shock - including the exchange rate




The figure displays responses to a QE-shock as identified above. The inner lines denote the median impulse responses from a Bayesian vector autoregression with 1000 draws, while the outer lines indicate one-standard error confidence bands. The vertical dotted lines indicate the restriction horizon.

Turning to the specification including the call rate of interest, figure 9 shows the impulse responses to the QE-shock. Our key results are unchanged in this specification; again, the confidence intervals are somewhat wider. Interestingly the call rate, which is restricted not to change on impact, does not react significantly to the QE-shock. In other words, our assumption that the Bank of Japan's key monetary policy instrument after 1995 was indeed the amount of current account balances, i.e. reserves, is supported by the data.

Figure 9: Impulse responses to a QE-shock - including the call rate

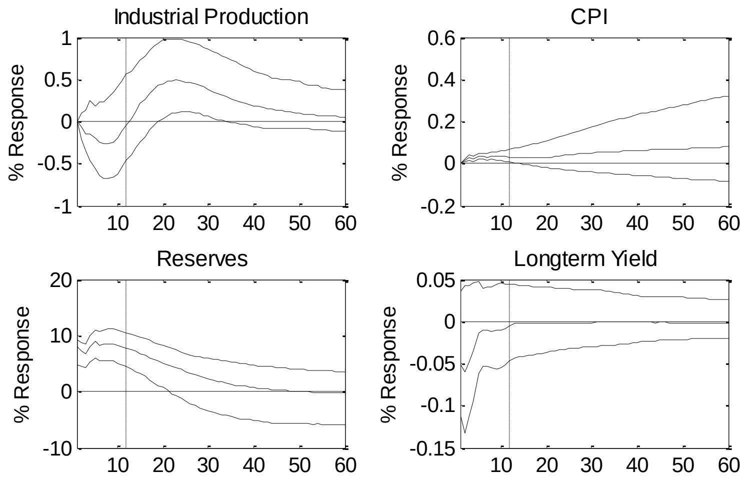



The figure displays responses to a QE-shock as identified above. The inner lines denote the median impulse responses from a Bayesian vector autoregression with 1000 draws, while the outer lines indicate one-standard error confidence bands. The vertical dotted lines indicate the restriction horizon.

## 5.3 Alternative identification scheme

We next turn to our benchmark specification results when we identify our three shocks according to the alternative sign restrictions given in table 3. These restrictions differ from the benchmark restrictions only in the identification of the demand and supply shocks. Because the theoretical predictions from the DSGE-model are the same for a positive demand and a negative supply shock, we need to impose the additional restriction on the relative magnitudes of the output and price responses. Results for the three shocks are given in figures 10, 11 and 12.

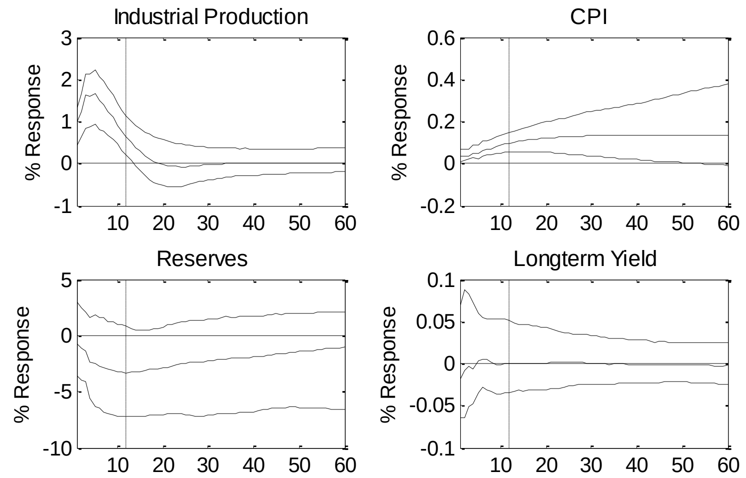
Figure 10: Impulse responses to a QE-shock - alternative identification scheme




The figure displays responses to a QE-shock as identified above. The inner lines denote the median impulse responses from a Bayesian vector autoregression with 1000 draws, while the outer lines indicate one-standard error confidence bands. The vertical dotted lines indicate the restriction horizon.

Figure 10 shows the impulse responses to the QE-shock using our alternative identification scheme. As expected the results are very similar to those from our benchmark identification. More interestingly, turning to the impulse responses to the demand and supply shocks under the alternative identification scheme, we do find some differences. Figure 11 shows the impulse responses to the demand shock. Again, the initial increase in industrial production and CPI is by construction. Again, the price level remains significantly positive for much longer than the restricted horizon. But overall the response is less strong compared to our benchmark identification scheme. Similarly, longterm yields do not react significantly at all this time.

Figure 11: Impulse responses to a demand shock - alternative identification scheme



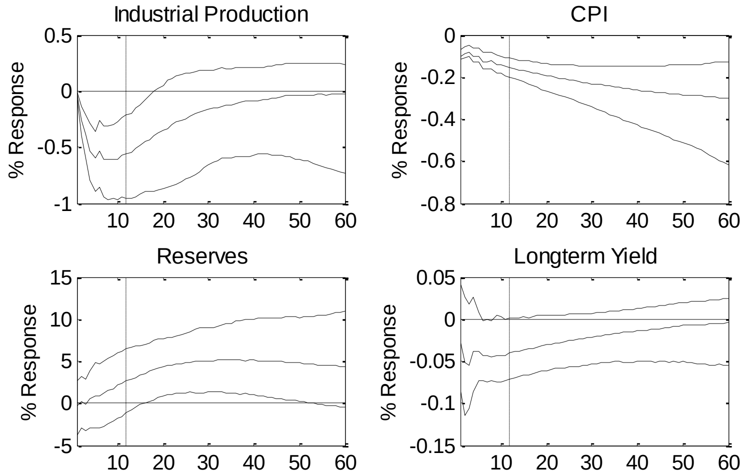

The figure displays responses to a demand shock as identified above. The inner lines denote the median impulse responses from a Bayesian vector autoregression with 1000 draws, while the outer lines indicate one-standard error confidence bands. The vertical dotted lines indicate the restriction horizon.

Finally, figure 12 shows the impulse responses to the supply shock under this identification: industrial production and CPI are restricted to fall following a positive supply shock. The crucial identification restriction regarding the relative magnitudes of the responses of production and prices can be seen by comparing the absolute size of the response of production to the demand and supply shocks. Industrial production is restricted to respond stronger than CPI following the demand shock, but less strong than CPI following the supply shock. This is confirmed by figures 11 and 12. But note now the difference in the responses to the supply shock as shown in figure 12 from the benchmark model shown in figure 7. CPI remains significantly negative for the entire period, reserves significantly rise after around 1-2 years, and the longterm yield is on the verge of being significantly negative after half a year.

## 5.4 Forecast error variance decomposition

In order to get a better understanding of the relative importance of our identified QE-shock for the variables of interest we calculate the forecast error variance decomposition which gives the estimated shares of the variability of each variable due to the respective shocks. Our main interest is of course focused on the variance shares of the QE-shock because they can be interpreted as measures of the quantitative effect of unconventional policy shocks on the real economy. Table 4 displays the median of the forecast error variance shares of the endogenous variables for each of the three identified shocks at the one to five-year forecast horizon. The last column of the

Figure 12: Impulse responses to a supply shock - alternative identification scheme




The figure displays responses to a supply shock as identified above. The inner lines denote the median impulse responses from a Bayesian vector autoregression with 1000 draws, while the outer lines indicate one-standard error confidence bands. The vertical dotted lines indicate the restriction horizon.

left panel shows the sum of the variance shares of the respective variables due to all identified shocks. It can be seen that together the structural shocks explain between 55% and 84% of the variations in the endogenous variables, which is a relatively large share. Moreover, it can be seen that the QE-shock explains some of the variations in the CPI and industrial production, our main variables of interest; however, these shares are rather small.While the unconventional shock explains fluctuations in output by up to 13%, the share is only 3 - 8% for the CPI. For both variables, the demand shock is the dominant source of variation with variance shares of around 40 - 50%. The right panel of Table 4 shows some interesting findings for the alternative specification including the exchange rate as well as the call rate. While the QE-shock still has a relatively minor role in explaining variations in output and prices, it seems to be relatively more important for the fluctuations in the exchange rate explaining around 15%. This suggests that while the response of this variable to the unconventional shock is found to be insignificant, this shock still has some explanatory power with regard to exchange rate fluctuations pointing to a non-negligible role of the exchange rate in the transmission of such shocks at the ZLB. Finally, as one would expect for a liquidity trap situation, our structural shocks only explain a relatively small share of variations in the call rate; the sum of the variance shares only range from 12% to 35%. As expected, the demand shock is the dominant source of variations in the shortterm interest rate.

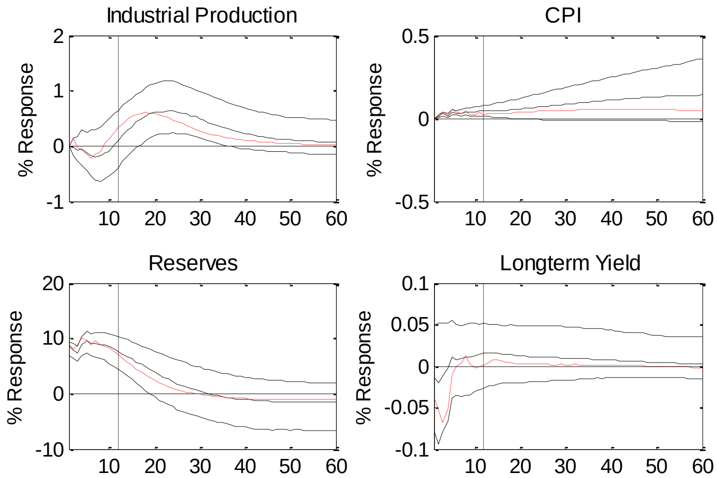
Table 4: Forecast error variance decomposition

| | | Benchmark specification - shock to: | | | | Alternative specification - shock to: | | | |
|---|---|---|---|---|---|---|---|---|---|
| Variable | Horizon | QE | Supply | Demand | Total | QE | Supply | Demand | Total |
| Ind. production | 1st | 1 | 2 | 54 | 57 | 2 | 21 | 27 | 50 |
| | 2nd | 6 | 6 | 51 | 63 | 6 | 15 | 27 | 48 |
| | 3rd | 12 | 7 | 45 | 64 | 10 | 14 | 23 | 47 |
| | 4th | 13 | 8 | 43 | 64 | 11 | 14 | 22 | 47 |
| | 5th | 13 | 8 | 43 | 64 | 11 | 14 | 22 | 47 |
| CPI | 1st | 3 | 17 | 38 | 58 | 4 | 21 | 25 | 50 |
| | 2nd | 5 | 8 | 49 | 62 | 5 | 10 | 29 | 44 |
| | 3rd | 6 | 5 | 49 | 60 | 6 | 6 | 27 | 39 |
| | 4th | 7 | 5 | 46 | 58 | 7 | 5 | 25 | 37 |
| | 5th | 8 | 5 | 42 | 55 | 7 | 5 | 23 | 35 |
| Reserves | 1st | 70 | 9 | 5 | 84 | 45 | 7 | 5 | 57 |
| | 2nd | 50 | 13 | 14 | 77 | 27 | 8 | 10 | 45 |
| | 3rd | 38 | 16 | 21 | 75 | 21 | 7 | 15 | 43 |
| | 4th | 34 | 17 | 24 | 75 | 18 | 7 | 16 | 41 |
| | 5th | 30 | 17 | 25 | 72 | 17 | 7 | 17 | 41 |
| LT yield | 1st | 12 | 49 | 13 | 74 | 13 | 7 | 12 | 32 |
| | 2nd | 14 | 37 | 17 | 68 | 13 | 7 | 14 | 34 |
| | 3rd | 17 | 32 | 17 | 66 | 14 | 7 | 14 | 35 |
| | 4th | 18 | 31 | 18 | 67 | 15 | 7 | 15 | 37 |
| | 5th | 18 | 30 | 19 | 67 | 15 | 7 | 15 | 37 |
| Exchange rate | 1st | | | | | 16 | 10 | 8 | 34 |
| | 2nd | | | | | 15 | 13 | 9 | 37 |
| | 3rd | | | | | 15 | 14 | 11 | 40 |
| | 4th | | | | | 15 | 13 | 14 | 42 |
| | 5th | | | | | 14 | 12 | 17 | 43 |
| Call rate | 1st | | | | | 3 | 3 | 6 | 12 |
| | 2nd | | | | | 7 | 7 | 12 | 26 |
| | 3rd | | | | | 10 | 8 | 13 | 31 |
| | 4th | | | | | 12 | 9 | 13 | 34 |
| | 5th | | | | | 12 | 9 | 14 | 35 |

## 5.5 Robustness

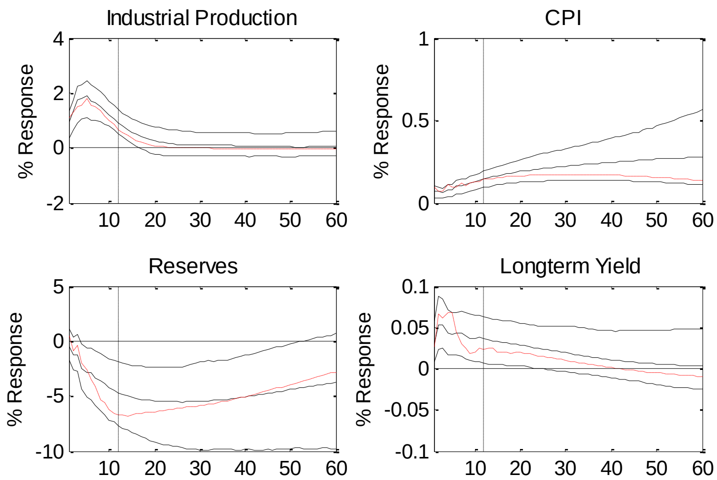
The first robustness check is concerned with the median as a way to summarise the information obtained from the Bayesian approach to calculating impulse responses. Figures 13 to 15 replicate the median impulse responses along with the 68% confidence intervals to the respective shocks. The red dashed lines additionally show the impulse responses generated by the one model that is closest to the median over all 500 models. It can be seen that generally, the impulse responses generated by this "close-to-median" model are quite similar to the median over all models.

Figure 13: Impulse responses to a QE-shock - close to median model



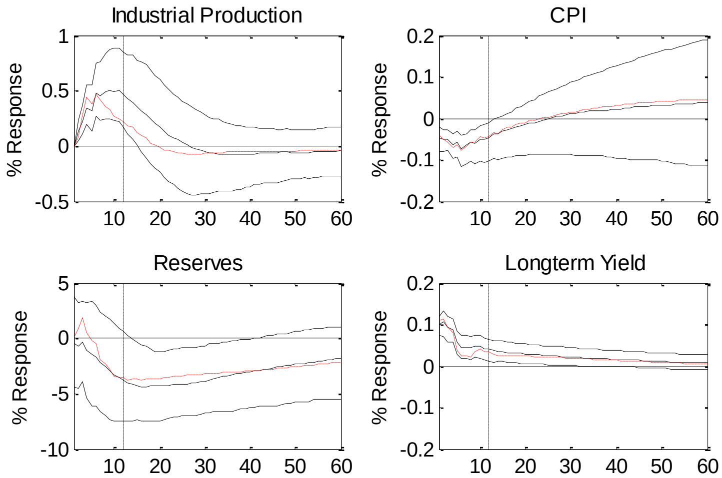

Responses to a QE-shock. The inner lines denote the median impulse responses from a Bayesian vector autoregression with 1000 draws, while the outer lines indicate one-standard error confidence bands. The red dotted lines display the response generated by the close to median model.

Figure 14: Impulse responses to a demand shock - close to median model




Responses to a demand shock. The inner lines denote the median impulse responses from a Bayesian vector autoregression with 1000 draws, while the outer lines indicate one-standard error confidence bands. The red dotted lines display the response generated by the close to median model.

Figure 15: Impulse responses to a supply shock - close to median model



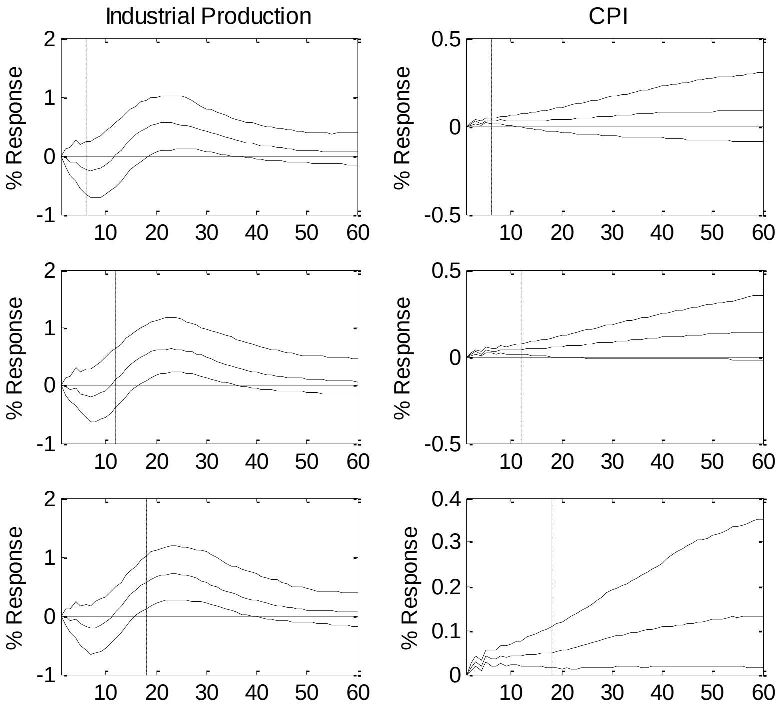

Responses to a demand shock. The inner lines denote the median impulse responses from a Bayesian vector autoregression with 1000 draws, while the outer lines indicate one-standard error confidence bands. The red dotted lines display the response generated by the close to median model.

The second robustness check involves specifying the restriction horizon k. As noted by, for instance, Uhlig (2005), it is difficult to base the choice of the appropriate restriction horizon on economic theory resulting in some degree of arbitrariness in specifying this parameter. We therefore check sensitivity of our results to this choice by estimating the benchmark model (for the 1-shock case) for different restriction horizons. Figure 16 shows the impulse response functions for our variables of interest, CPI and industrial production, for a lower restriction horizon compared to the benchmark model $k = 6$ and for a longer horizon $k = 18$ (displayed in the first and the third row, respectively). The benchmark case, $k = 12$ is given in the second row. It can be seen in the figure that our main results are qualitatively insensitive to variations in k; industrial production shows a significant and positive response at least over several months. However, the magnitude of this increase differs among the respective cases. While for $k = 6$ the positive impact on economic activity is significant only with a delay of about two years and vanishes rather fast following the shock[17], the response is stronger and lasts somewhat longer for $k = 18$.[18]

[17]Similar results are obtained for a restriction horizon of nine months or eight months.

[18]Again, results are very similar for even longer restriction horizons of, say, 24 months.

Figure 16: Impulse responses to a QE-shock - varying the restriction horizon




The figure displays responses of industrial production and the CPI to a QE-shock shock as identified above. The inner lines denote the median impulse responses from a Bayesian vector autoregression with 500 draws, while the outer lines indicate one-standard error confidence bands. The vertical dotted lines indicate the restriction horizon.

# 6 Discussion and Conclusion

The primary objective of this paper has been to agnostically assess the real effects of QE measures adopted by the Bank of Japan for a liquidity trap episode. We suggest to use results from the theoretical literature to derive our identifying restrictions for our SVAR. In particular, we propose a set of sign restrictions based on predictions of DSGE models explicitly taking into account the ZLB that clearly identify an unconventional shock without imposing restrictions on interest rates or spreads. Given that a broad consensus is still missing as to how to identify monetary shocks at the ZLB, we used two different identification strategies and various specifications. Our results show that a QE-shock does positively and significantly affect industrial production. After around two years industrial production has risen by about 0.5% following an increase in reserves by 10%. On the other hand, this shock has virtually no effect on core CPI and thus the rate of inflation. Concerning longterm government yields we do not find any significant reduction in yields following the QE-shock. Moreover, while the exchange rate seems to depreciate after an expanisonary shock, the effect is not significant.

Overall, our empirical results show that unconventional monetary policy can positively affect real economic activity even when the economy is in the liquidity trap. However, the QE-shocks we identify do not significantly affect inflation. Our results concerning the longterm yield and the exchange rate therefore suggest that a direct quantity effect such as a portfolio rebalancing channel in the spirit of Meltzer (1995) has not been at work following the QE policies in Japan. We believe these results are interesting not only for the Japanese economy, but also for other advanced economies where monetary policy is constrained by the ZLB.

# A Data

In the benchmark case we include four variables reflecting the macroeconomic and monetary environment of the Japanese economy. We use monthly observations for the period of January 1995 to September 2010. The start of the sample period is motivated by the fact that the Bank of Japan first decreased nominal interest rates to below 1% during the course of 1995 and we are mainly interested in the effectiveness of monetary policy at near-zero interest rates.

*Monetary variables*

As far as the monetary variables are concerned, we include the shortterm interest rate as well as a measure of reserves; both series have been obtained from the Bank of Japan's statistics website. As the shortterm interest rate we include the monthly average of the uncollateralized overnight call rate, which has been the target rate for the Japanese Central Bank from July 1985 (Miyao, 2002). As far as reserves are concerned, to be able to identify the QE-shock we include the average outstanding current account balances held by financial institutions at the Bank of Japan. This is the part of the monetary base that can be referred to as reserves held at the central bank. Under the QE policy this variable has gained importance as the main operating target for the Bank of Japan.

*Prices*

We include the core consumer price index, which measures the development of consumer prices excluding energy and food. Base year is 2005. The core CPI has been obtained from Datastream. Moreover, we include a narrow index of the real effective exchange rate of the Yen against other currencies as published on the Bank for International Settlements' (BIS) website. Both series are seasonally adjusted by X12-ARIMA.

*Industrial Production*

We include a measure of the Japanese industrial production as a generally used indicator of economic activity. Base year is 2005. The series has been obtained from Datastream and is seasonally adjusted by X12-ARIMA.

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