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Abstract

With short term interest rates bounded at zero, monetary policy has aimed at affecting the yield curve at the longer end during the recent years. As the recent literature has shown, the quantitative easing programs conducted by the Federal reserve have significantly lowered long-term yields. This paper adds central bank reserves as a fourth factor to an affine term structure model to estimate the effect of quantitative easing on the yield curve. The cumulative effect on 10-year Treasury securities during the zero lower bound period is estimated to amount to 85 basis points. Of the total effect, one quarter is shown to be due to the liquidity effect and three quarters to the supply effect. To disentangle the two effects, the estimates for the US are compared to estimates for Swiss data because the Swiss national bank did not engage in any government bond purchases.

Keywords: Yield curve, zero lower bound, liquidity effect, reserves, Bayesian MCMC
JEL Classifications: E43, E52, C11, G12

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1 Introduction

Central banks around the globe have been flooding financial markets with liquidity since the breakdown of interbank markets in response to the collapse of Lehman Brothers in the fall of 2008. In its initial phase of unconventional monetary policy, also referred to as QE1, the US Federal Reserve performed large scale asset purchases consisting mainly of mortgage-backed securities. The Federal Reserve then gradually extended liquidity by outright purchases of Treasury securities. When the Treasury purchase program known as QE2 came to an end in mid 2011, non-borrowed reserves (NBR) held at the Federal Reserve exceeded USD 1.6 Trillion.

Aside from signalling effects about future policy, the literature has focussed mainly on portfolio balance effects of outright Treasury securities' purchases of the Federal Reserve[1], in particular the so-called supply effect and spill over effects from outright purchases of similar assets such as mortgage backed securities and agency debt[2]. Krogstrup et al. (2012) have shown that, in addition to the supply effect, there is also a liquidity effect of QE that has affected long-term Treasury bonds. While the supply effect depends on the composition of the central bank's balance sheet, it is an expansion of the balance sheet that triggers the liquidity effect.

This study proceeds in two steps: it first estimates the total effect of QE on term premia in an Affine Term Structure Model (ATSM) framework; in a second step it attempts to disentangle the supply from the liquidity effect. Our estimates show that QE has significantly lowered the term premium of Treasury yields. Estimates of the cumulative effect of QE on 10-year Treasury yields during the ZLB amounts to 85 basis points. This result is in line with the recent literature. Estimates of Swiss data suggest 10-year yields decreased 34 basis points in response to the expansion in liquidity. Because the Swiss National bank has not engaged in government bond purchases during the ZLB, the Swiss estimates can be fully ascribed to the liquidity effect. In the US, approximately 21 basis points can be attributed to the liquidity effect and 64 basis points to the supply effect.

QE may affect either the expectations' part of interest rates or the term premium. While signalling effects change expectations about future policy, the supply and the liquidity effect

[1]See e.g. Gagnon et al. (2011), Hamilton and Wu (2010b), Krishnamurthy and Vissing-Jorgensen (2011)

[2]See e.g. D'Amico and King (2010)

operate on the term premium. Term structure models are suitable to isolate the term premium. The three latent factors are typically shown to account for changes in expected future interest rates and thus capture any signalling effects and announcement effects that affect expectations about the future policy rate. Christensen and Rudebusch (2012) isolate the term premium in an affine term structure model and find that in US data, most of the effect of QE on yields was transmitted through policy expectations rather than the term premium. Hence, they find that the portfolio balance effect - or the supply effect - of QE on yields was much smaller than the signalling effect in the US. Information theory would suggest that the supply effect materializes immediately after announcements of the asset purchases. Therefore, many papers[3] apply event studies to estimate the effects. Instead, we add a fourth factor to the term structure model, a measure for liquidity, to explain changes in the term premium. The main analysis is computed for US data, divided into two subsamples: a pre-crisis sample that should account for "normal" times and a zero lower bound sample. To better separate the supply effect from the liquidity effect, we additionally run the analysis for Swiss data as a counterfactual case. The injected liquidity by the Swiss National Bank was entirely driven by foreign exchange purchases rather than government bonds. Hence, a supply effect can be excluded.

The rest of the paper is organized as follows. Section 2 lays out the different effects of quantitative easing on yields. Section 3 illustrates the US and Swiss data sets. Section 4 introduces the term-structure model used and Section 5 provides the details on the model estimation. Finally, Section 6 presents the results and Section 7 concludes.

2 The Different Effects of QE on Interest Rates

The recent literature has focused on two effects of the Federal Reserves' asset purchases. Fist, measures undertaken by the monetary authority are typically thought to have a signaling effect about future monetary policy. For example, Christensen and Rudebusch (2012) find that most of the declines in US interest rates can be attributed to lower policy expectations. Second, the purchase of long-term assets changes the maturity structure of the assets available to the public which might trigger a portfolio balance effect. Accordingly, Gagnon et al. (2011) attribute most of the effect of the LSAP on long-term yields to this portfolio balance effect, the so-called supply effect. Krogstrup et al. (2012) suggest a third mechanism at work: a

[3]See for example Gagnon et al. (2010) and Bauer and Rudebusch (2011).

liquidity effect on longer-term yields. Conversely to the first two effects, the liquidity effect hinges on the expansion rather than the composition of a central banks' balance sheet. By augmenting its balance sheet, the central bank creates additional funding and provides it to the economy, so there is more funding available to finance an unchanged amount of debt. Krogstrup et al. (2012) argue that some of the additional credit might have found its way into Treasury bonds and thereby brought down long-term yields.

The signalling effect can be clearly distinguished from the other two as it affects the expectations' part of yields rather than the term premium. It is not straightforward, however, to disentangle the liquidity effect from the supply effect. With an outright purchase of a Treasury bond, the Federal Reserve temporarily raises demand and therefore the price of that bond. Since the bond is drawn from the market, the supply of that bond decreases. Given unchanged market demand for this specific bond, the price increases and the yield decreases permanently. This effect is referred to as the supply effect in the literature.[4] The supply effect boils down to a portfolio balance effect that arises if there is some kind of "preferred-habitat" behavior of investors, e.g. if investors have a preference for the specific maturity of the asset purchased by the central bank. Some research suggests that this portfolio balance effect might even be at work across different asset classes. D'Amico and King (2010), for example, suggest that the purchases of MBS conducted by the Fed significantly affected long-term Treasury yields. In addition to the supply effect, the price change will spread to close-by maturities due to arbitrage. That said, an outright purchase of an asset conducted by the central bank raises the amount of credit in the form of reserves available in the aggregate. Assuming an unchanged amount of government debt, i.e. Treasury bonds, prices of these bonds will, all else equal, increase as there is more money to buy the same amount of assets. This is what Krogstrup et al. (2012) refer to as the liquidity effect.[5] The liquidity effect in the sense of Friedman (1968) refers to a fall in short-term nominal interest rates after an exogenous persistent increase in narrow measures of the money supply.[6] When the policy rate hits the zero lower bound (ZLB), the liquidity effect on the short-term yields is arguably absent, as the

[4]See e.g. Hamilton and Wu (2010a), Vayanos and Vila (2009), Greenwood and Vayanos (2010), Neely (2010), Kuttner (2006), D'Amico and King (2010) for recent research on the supply effect.

[5]The amount of money relative to the volume of assets increases, ceteris paribus, against the entire spectrum of assets. So the same price effect might be observed for other assets, too. The effect on Treasury bonds might be particularly large because these assets are very safe and liquid. Krishnamurthy and Vissing-Jorgensen (2011) argue that these characteristics make government bonds constitute a close substitute for money.

[6]See e.g.Thornton (2008)

short-term Treasury bills and money become equivalents. However, as the banks that sell assets to the central bank hold more money after the transaction, they might seek a positive return further out the yield curve.

The supply and the liquidity effect are difficult to distinguish empirically. The expansion of the Federal Reserve's balance sheet was largely due to the purchase of US Treasury bonds. Thus, the change in the supply of bonds, measured as the quantity of these bonds held by the Federal Reserve, is correlated with the amount of liquidity, measured as the deposits banks hold at the Federal Reserve. We address this difficulty by comparing estimates from US data with those from Swiss data. In the case of Switzerland, the supply effect cannot have been at play as the Swiss National bank increased liquidity through purchases of foreign exchange rather than government bonds.

3 Data

The following section briefly describes and illustrates both the US and the Swiss datasets used and explains how we conduct our analysis in two subsamples, one corresponding to "normal times" and other to the ZLB period.

3.1 US Data

The analysis is conducted on a weekly basis, so the data is converted into the same frequency. The sample spans from June 2005 to October 2011. Because NBR data are published on Wednesdays, all data are snapshots as of that day. Interest-rate data used stem from off-the-run US Treasury Bills and Bonds. In particular, we use 3-month and 6-month secondary market T-bills rates[7] and 1-year to 10-year off-the-run constant maturity yields from Gurkaynak et al. (2007)[8]. All the yields are continuously compounded, whereas quarterly and semi-annual compounding is assumed for the 3-month and the 6-month rate, respectively.[9]

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[7]Obtained from the Federal Reserve Economic Data base (FRED), under DTB3 and DTB6. The two are not constant-maturity yields which is acceptable due to their short maturity.

[8]The data can be downloaded from `http://www.federalreserve.gov/econresdata/researchdata.htm`. We use off-the-run treasuries to avoid the treatment of "repo-specialness" implicit in the on-the-run treasuries, see Duffie (1996) and Jordan and Jordan (1997).

[9]See Hull (2008).

In the literature, different measures of money have been used to analyze the liquidity effect. As concluded by Pagan and Robertson (1995), models defining money as narrow measures are more successful in giving evidence for a liquidity effect than those using broader measures of money. This is due to the fact that shocks to broader monetary aggregates are largely due to shocks in money demand rather than to shocks in money supply.[10] Hence, the expansion in liquidity conducted by the Federal Reserve is measured by the amount of reserves banks hold with the Federal Reserve. Total reserves are often found not to correlate with interest rates because the part of reserves which is borrowed is driven by the bank's demand for reserves. This makes borrowed reserves endogenous, as the demand for liquidity responds to monetary policy changes.[11] We therefore use non-borrowed reserves, i.e. total reserves minus borrowed reserves,[12] as the measure of liquidity.

The main hypothesis of the paper is that liquidity effects are present at the longer end of the yield curve only when interest rates are close to the ZLB. Therefore, the analysis is carried out in two subsamples. The first can be characterized as "normal times" with interest rates far away from the ZLB. It is defined as starting with the first week of June 2005 until the end of July 2007.[13] The second subsample corresponds to the "zero lower bound" period. It starts with the second week of December 2008, when the Federal funds rate reached the ZLB, and runs through the end of the sample (see Figure 1) i.e. to October 2011.

FIGURES 1 AND 2 ABOUT HERE

Figure 1 depicts the data. The NBR fluctuated around the level of 10 billion USD in the first subsample shown in the upper panel. NBR amounted to roughly 540 billion USD at the time the Federal funds rate reached the ZLB in late 2008. Then, during the ZLB period shown in the lower panel, NBR continued to grow to reach 1,560 billion USD by October 2011. In the first and the second subsample, the standard deviation of NBR equals 2.9 and 318.5 billion USD, respectively. In order to make the estimation results from the two periods comparable,

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[10] See e.g. Bernanke and Blinder (1992), Sims (1992), or Christiano and Eichenbaum (1995).

[11] See e.g. Carpenter and Demiralp (2008)

[12] Borrowed reserves are equal to the sum of credit extended through the Federal Reserve's regular discount window programs and credit extended through certain Federal Reserve liquidity facilities. The total reserves (WRESBAL) and the total borrowings (TOTBORR) are also available from FRED.

[13] We could have used a longer sample for the non-ZLB period. Trying different sample lengths and start/end points does not change the non-ZLB results.

the NBR factor is normalized so that it exhibits a standard deviation of one within each subsample.

3.2 Swiss Data

The data used in the analysis are the 3-month and the 6-month CHF Libor rates and 1-year to 10-year constant-maturity zero-coupon yields on Swiss Confederation Bonds. The data on Giro holdings of both domestic and foreign banks held at the SNB is used to measure liquidity. The SNB publishes this data on a weekly basis, so the analysis can also be conducted at the weekly frequency. In line with the US case, we choose the two subsamples so that one corresponds to "normal times" and the second to the ZLB period. The former subsample starts in the first week of January 2006 and ends on 12 September 2008.[14] The second subsample starts with the second week of December 2008, when the Swiss National Bank (SNB) decreased its targeted level for the 3-month Libor from 1% to 0.5%, and ends on 7 October 2011.

3.3 Simple Data Inspection

Figure 1 presents US data and Figure 2 Swiss data, where the upper panels show the earlier subsample and the lower panels depict the ZLB period. It is very difficult to see any comovement between interest rates and NBR in both samples in either dataset. To get a first idea on such a comovement, Table 1 depicts correlations between different yields and NBR for both subsamples. The numbers are reported for both the US and Switzerland. In the upper panel, the corrleation refers to levels, whereas the lower panel computes the correlations in first differences. The simple correlations draw a mixed picture.

TABLE 1 ABOUT HERE

Macroeconomic theory holds that, in normal times, the liquidity effect dominates at the short end of the yield curve, while at the longer end the Fisher effect is thought to obscure the liquidity effect. Hence, we would expect the correlation between NBR and short-term yields to be negative, while the correlation with long-term yields would turn positive with rising

[14]The exact choice of the "normal times" subsample is arbitrary and does not affect the results. Preliminary analysis of several different non-ZLB subsamples showed no liquidity effect irrespective of the exact period.

inflation expectations. As Table 1 shows in the first row, the data does provide evidence for the liquidity effect as correlations are negative in the US data. The negative correlation weakens with maturity, however, suggesting that the dominance of the Fisher effect increases as we move out on the yield curve. The Swiss data provides the exact opposite picture; there is no evidence for a liquidity effect, and the Fisher effect seems to decrease with maturity. Correlations decrease considerably during the ZLB period for both countries. In the US data the negative correlation is largest around the 1-year maturity while the Swiss data suggests longer maturities to be affected the most.

4 The Model

This Section defines the term structure model we use to estimate the liquidity effect. We start from the general asset pricing equation, define the pricing kernel and specify a Gaussian diffusion process of the underlying state factors. Then, we specify the one-period interest rate and formulate the bond prices across the maturity spectrum.

4.1 General Setting and State Dynamics

The general asset pricing equation[15] under the physical probability measure $\mathbb{P}$ reads:

$$P_{n,t} = E_t \left[M_{t+1} P_{n-1,t+1} | I_t \right], \tag{1}$$

where $P_{n,t}$ is the price of an n-periods to maturity zero-coupon bond at time t, M_{t+1} denotes the stochastic discount factor, and I_t represents the agents' current information set. In a risk-neutral world, where investors request no risk compensation, the price of the bond $P_{n,t}$ equals:

$$P_{n,t} = E_t^{\mathbb{Q}} \left[exp(-y_{1,t}) P_{n-1,t+1} | I_t \right], \tag{2}$$

where $\mathbb{Q}$ is the risk-neutral probability measure and $y_{1,t}$ is the short-term interest rate. The no-arbitrage argument assures that the two prices in (1) and (2) are equal. There exists an equivalent martingale measure $\mathbb{Q}$ according to which (2) holds[16] with the stochastic discount factor taking the form:

[15] See Campbell et al. (1997)

[16] See Harrison and Kreps (1979).

$$\begin{aligned} exp(-y_{1,t}) &= E_t\left[M_{t+1}|I_t\right] \\ &= exp(-y_{1,t})E_t\left[(d\mathbb{Q}/d\mathbb{P})_{t+1}|I_t\right]. \end{aligned}$$

$d\mathbb{Q}/d\mathbb{P}$ is the Radon-Nykodim derivative[17] which follows a log-normal process, so it reads:

$$(d\mathbb{Q}/d\mathbb{P})_{t+1} = exp\left(-\frac{1}{2}(\lambda_t)'\lambda_t - (\lambda_t)'\varepsilon_{t+1}\right) \tag{3}$$

where λ_t is the market price of risk associated with the sources of uncertainty ε_t[18]. Following Duffee (2002), the market price of risk is an "essentially affine" function of the state variables X_t, so it can be written as

$$\lambda_t = \lambda_0 + \lambda_1 X_t. \tag{4}$$

Equations (3) to (4) jointly define the pricing kernel of the model, where the essentially affine market price of risk constitutes the first fundamental building block of the Gaussian term structure model.

Another fundamental building block of the Gaussian term structure model is the multivariate state variable X_t. It follows a discrete version of the constant volatility Ornstein-Uhlenbeck process[19]. Under the physical probability measure $\mathbb{P}$, the process is

$$X_{t+1} = (I - \Psi)\mu + \Psi X_t + \Sigma\varepsilon_{t+1}. \tag{5}$$

The first term on the right-hand side of equation (5) is a vector of the factors' means. Ψ is the VAR matrix, Σ is the covariance matrix that normalizes the residuals ε_t which are assumed to be standard normal *i.i.d.* shocks.

4.2 Short Rate and Bond Prices

Following Duffie and Kan (1996), the one-period interest rate is an affine function of risk factors X_t as:

$$y_{1,t} = A_1 + B_1 X_t$$

[17] See Dai et al. (2007).

[18] See Ang and Bekaert (2002) and Ang and Piazzesi (2003).

[19] See Phillips (1972).

where the coefficient A_1 corresponds to the average one-period rate in the sample and B_1 is a vector of loadings of the risk factors on $y_{1,t}$. The risk factors are:

$$X_t = \begin{bmatrix} l_{1,t} \\ l_{2,t} \\ l_{3,t} \\ Reserves_t \end{bmatrix} \tag{6}$$

The first three factors l_t^i, $i = 1, 2, 3$ denote the latent factors backed out from yields. As commonly in the term structure literature, these factors can be interpreted as a level, a slope, and a curvature factor.[20] In order to estimate the liquidity effect on the longer-end of the yield curve, this paper adds central bank reserves, as the fourth factor to X_t.[21] As it can be noticed, the model assumes no ZLB for the short rate. Yet, as shown in the results Section, the model fits the data reasonably well and thus can be considered as a good approximation.

Assuming joint log-normality of bond prices and the pricing kernel in equation (1), the n-periods to maturity nominal bond price is an affine function of the state variables[22] and thus takes the form:

$$p_{n,t} = -A_n - B_n X_t, \tag{7}$$

with:

$$\begin{aligned} A_n &= A_{n-1} + B_{n-1}\left((I - \Psi)\mu - \Sigma\lambda_0\right) + \frac{1}{2} B_{n-1}\Sigma\Sigma' B_{n-1}' + A_1 \\ B_n &= B_{n-1}\left(\Psi - \Sigma\lambda_1\right) + B_1 \end{aligned} \tag{8}$$

5 Estimation

In this section the likelihood function used to construct the joint posterior of parameters and data is derived. In addition, the model is estimated with a simple version the Bayesian

[20]See Section 5 for more details on "backing out" of the yield-only factors.

[21]Similarly, Ang and Piazzesi (2003), for instance, are among many studies that use macro-economic variables as explicit factors.

[22]See for example: Cochrane and Piazzesi (2009).

Markov-Chain Monte-Carlo (MCMC) method and this section provides the general idea, rationale and the description of the algorithm.

5.1 Likelihood Function

Following Chen and Scott (1993), the 6-month, the 5-year and the 10-year yields are set to be observable, while the rest is measured with error. Let $y_{o,t}$ be a vector of observed yields i.e. yields perfectly priced by the model:

$$\begin{bmatrix} y_{o,t} \\ Reserves_t \end{bmatrix} = \begin{bmatrix} A_o \\ 0_{1\times 1} \end{bmatrix} + \begin{bmatrix} B_o & 0_{3\times 1} \\ 0_{1\times 3} & 1 \end{bmatrix} \begin{bmatrix} X_{o,t} \\ Reserves_t \end{bmatrix}$$

where A_o is a 3×1 vector and B_o a 3×3 matrix of factor loadings. $X_{o,t}$ are the three latent factors obtained by inverting the above equation as:

$$\begin{bmatrix} X_{o,t} \\ Reserves_t \end{bmatrix} = \begin{bmatrix} B_o & 0_{3\times 1} \\ 0_{1\times 3} & 1 \end{bmatrix}^{-1} \left(\begin{bmatrix} y_{o,t} \\ Reserves_t \end{bmatrix} - \begin{bmatrix} A_o \\ 0_{1\times 1} \end{bmatrix} \right)$$

The first part of the likelihood function refers to the evolution of the state variables X_t as given in equation (5). The assumption that ε_t is multivariate-Gaussian implies that the conditional probability density function of X_t is

$$f\left(X_t \mid y_{o,t-1}, Reserves_{t-1}\right) = \frac{exp\left(-\frac{1}{2}\varepsilon_t' \left(\Sigma'\Sigma\right)^{-1} \varepsilon_t\right)}{\sqrt{(2\pi)^T \mid \Sigma'\Sigma \mid}}. \tag{9}$$

Let the $y_{u,t}$ be a vector of remaining $N-3$ yields priced by the model with an error:

$$y_{u,t} = A_u + B_u X_t + \xi_t$$

The second part of the likelihood function refers to the pricing errors ξ_t. They are assumed to be distributed as *i.i.d.* $N\left(0, \omega^2 I\right)$, with the same variance ω and zero-correlations across yields, where I is an unity matrix. The conditional density of $y_{u,t}$ is thus given by:

$$f\left(y_{u,t} \mid X_t\right) = \frac{exp\left(-\frac{1}{2}\xi_t' \left(\left(\omega^2 I\right)' \omega^2 I\right)^{-1} \xi_t\right)}{\sqrt{(2\pi)^T \left|\left(\omega^2 I\right)' \omega^2 I\right|}} \tag{10}$$

The log-likelihood function is just the sum of logarithms of the "time-series part" in equation (9) and the "cross-sectional part" in equation (10) and thus takes the form:

$$\ln \mathcal{L}(\cdot) = \log(f\left(X_t \mid y_{o,t-1}, Reserves_{t-1}\right)) + \log(f\left(y_{u,t} \mid X_t\right)) \tag{11}$$

5.2 Econometric Identification

Solid identification of parameters is an essential part of dynamic term structure models estimation. The proposed identification scheme stems mainly from Dai and Singleton (2000) and Hamilton and Wu (2010a).

To begin with, the upper-left block of the VAR matrix Ψ, i.e. the one driving the dynamics of the yields-only factors in $X_{o,t}$, is set to be a power law structure,[23] with zero non-diagonal elements and the following power relation on the diagonal:

$$\psi_{zz} = \psi_{11}\alpha^{z-1}$$

where ψ_{11} is the largest eigenvalue and the AR coefficient of the first latent factor, α is a scaling parameter controlling the distance between the eigenvalues, and $z = 2, ...Z$, where Z is the number of latent factors. Preliminary estimation showed that the ψ_{11} parameter is near one. In line with the near co-integration assumption from previous studies[24], we simply set ψ_{11} to 1,[25] and estimate ψ_{zz}, where $z = \{2, 3\}$, together with the AR(1) coefficient of the reserves dynamics. The off-diagonal elements of the Ψ are set to zero, as well as vector the μ vector and the off-diagonal elements of matrix Σ in the transition equation (5).

We impose the usual boundary condition $A_0 = B_0 = 0$ on the parameters of the pricing equation given in (7). A_1 is normalized to average the one-period interest rate in the sample[26] while B_1 is normalized to $\begin{bmatrix} 1 & 1 & 1 & b_{Reserves} \end{bmatrix}'$.[27] Alternatively, one could set the covariance matrix of the transition equation (5) to a unity matrix and estimate all the elements of B_1.[28] Finally, the market price of risk dynamics is restricted to

$$\lambda_t^s = \begin{bmatrix} \lambda_{0,1} \\ \lambda_{0,2} \\ 0 \\ 0 \end{bmatrix} + \begin{bmatrix} 0 & 0 & 0 & \lambda_{1,14} \\ 0 & 0 & 0 & \lambda_{1,24} \\ 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 \end{bmatrix} X_t, \tag{12}$$

[23] See also Calvet et al. (2010) and Bauer and de los Rios (2011).

[24] See for instance Giese (2008) and Jardet et al. (2011).

[25] As in Diebold and Li (2006), Söderlind (2010) and Bauer (2011).

[26] Following Favero et al. (2007).

[27] As in Ang et al. (2008).

[28] See for instance Ang and Piazzesi (2003).

so that both "level" and "slope" risks are priced in yields.[29] This is the main market price of risk specification of the study. The results are also reported for the risk-neutral $\mathbb{Q}$ measure, where $\lambda_0 = \lambda_1 = 0$, and for the case where only the "level" shock is being a compensated risk, i.e. where $\lambda_{0,2}$ and $\lambda_{1,24}$ are set to 0.[30]

5.3 Bayesian Inference

The yield curve implied by the model is a complicated non-linear function of the underlying parameters. As this non-linearity tend to produce a multi-modal likelihood function[31], fitting a yield curve model with a standard maximum likelihood estimation is a daunting task. Bayesian Markov Chain Monte Carlo (MCMC) method seem to be a powerful alternative, providing both efficiency and tractability.

5.3.1 Setting

Let Θ be a vector of length K collecting all the parameters of the model to be estimated:

$$\Theta = \{\alpha, \psi_{Reserves}, \Sigma, \lambda_0, \lambda_1, \omega, b_{Reserves}\}$$

The key idea behind Bayesian estimation is to consider the vector as a multivariate random variable, and use the Bayes' rule to "learn" about the variable *given* the data:

$$p\left(\Theta \mid \text{data}\right) = \frac{p\left(\text{data} \mid \Theta\right) p\left(\Theta\right)}{p\left(\text{data}\right)} \tag{13}$$

where $p\left(\Theta \mid \text{data}\right)$ is the posterior density of Θ, $p\left(\text{data} \mid \Theta\right)$ is the likelihood function and $p\left(\Theta\right)$ denotes the prior density of the parameters. The term $p\left(\text{data}\right)$ is known as "normalising constant" and it is independent of Θ.[32] Consequently, the rule in (13) can be re-written as:

$$\ln p\left(\Theta \mid \text{data}\right) \propto \ln\mathcal{L} + \ln p\left(\Theta\right)$$

where $\ln\mathcal{L}$ is the logarithm of the likelihood function defined in (9) and (10):

[29]As in Duffee (2010) and Joslin et al. (2010).

[30]We follow exactly this specification from Cochrane and Piazzesi (2009) and add the slope factor. It is indeed a restricted set of models, yet previous studies show that many restrictions on the market price of risk are supported by the data, see Joslin et al. (2010) and Bauer (2011).

[31]See Chib and Ergashev (2009).

[32]In particular: $p\left(\text{data}\right) = \int p(\text{data} \mid \Theta)p(\Theta)d\Theta$. See Koop (2003).

$$\begin{aligned}\ln\mathcal{L} &= \sum_{t=0}^{T-1}\mathrm{lnpdf}\left(X_t \mid y_{o,t-1}, r_{t-1}\right) \\ &+ \sum_{t=0}^{T-1}\mathrm{lnpdf}\left(y_{u,t}|X_t, y_{o,t-1}, r_{t-1}\right)\end{aligned} \tag{14}$$

5.3.2 Priors

In the estimation, the priors $p(\Theta)$ are set to be non-informative or "flat", so that the posterior density of the model parameters is drawn with equal probability from the pre-defined support interval. Alternatively, we could derive the prior distributions for parameters Ψ and Σ, given the normality assumption of the state VAR process,[33] and for ω given the assumption of the Gaussian measurement error.[34] Chib (2001) propose a scaled beta distribution as an alternative to the uniform distribution. Nevertheless, we choose not to impose lower (or higher) probability areas from which the candidate values of parameters are drawn. In such a way, the estimation is almost completely data-driven and proves to be computationally efficient.

The parameters' suport intervals are specified by following the no-arbitrage condition and previous studies. In particular, the eigenvalues of the VAR matrix are set to be positive and less than one and the volatility parameters on the diagonal of Σ are set to be non-negative. The parameter b_{NBR} is constrained to be inside the unit-circle. The lower bound of the parameters in λ_0 vector and λ_1 matrix are set as in Chib and Ergashev (2009).

[33]For instance, see Ang et al. (2007).

[34]See for example Mikkelsen (2001).

5.3.3 Markov Chain Monte Carlo

We use a simple version of "Metropolis within Gibbs" algorithm[35] to draw the parameters from their posterior densities. The parameter candidates are drawn from continuous uniform distributions $U\left(\underline{\Theta},\overline{\Theta}\right)$ where the lower and the upper boundaries $\underline{\Theta}$ and $\overline{\Theta}$ for each parameter in Θ are specified in Table 2. The algorithm can be described in several steps:

Step 1: Set the initial values of parameters Θ^0. Two Markov chains with different starting values are set up. The initial values for the VAR parameters in first chain are obtained from OLS and the data descriptive statistics. The starting values of the market price of risk parameters and of the second chain are chosen arbitrarily.[36]

Step 2: Draw a candidate log-posterior density $\ln p\left(\Theta^* \mid \Theta^{mc-1}, \text{data}\right)$ conditional on previously drawn parameters' values Θ^{mc-1}. The number mc denotes current iteration. The draws are performed separately for every parameter in Θ. For instance, a proposal for the first element in the vector Θ is generated by the following Markov chain:

$$\theta_1^* = \theta_1^{mc-1} + \nu_1 U_1$$

where ν_k is a scaling factor and U_k is an uniformly distributed random number from interval [-1,1]. We initialise ν_k for parameters α and ψ_{NBR} to 0.01, for diagonal elements of Σ matrix and the market price of risk parameters to 0.1, and for the ω parameter to 0.00001[37]. The scaling factor is then automatically updated after every 5,000 sweeps[38] to obtain the accep-

[35] The "Metropolis within Gibbs" is a simple method and therefore often used in the literature, as for example in Gilks (1996), Koop (2003) and Lynch (2007). In particular, this algorithm features two favorable characteristics. First, in standard Metropolis new proposals for the parameter values are drawn all at once, whereas the Gibbs sampler draws only one proposal for only one parameter value at a time. Therefore, the Gibbs sampler results much more efficient, i.e. the estimator converges quicker. Second, the tuning of the estimator is simpler, where tuning refers to setting the scale factor in step 2 to different values. In standard Metropolis, one scale factor must "fit all". The Gibbs sampler is more flexible as it allows for a separate set of values for each parameter.

[36] For example, the volatility parameters' starting values are set to be 3 times larger in the second chain, the Ψ matrix parameters are set to 0.8 and the B_1 parameter is set to 0.5 and -0.5 in the first and the second Markov chain, respectively.

[37] Proposed in Ang et al. (2007) so that it roughly corresponds to a 30 basis points bid-ask spread on Treasuries. An average spread on the OTC plain vanilla swap market might be similar. See also Skarr (2010).

[38] The algorithm is ran for 100,000 times. The scaling factor is updated starting from the 10,000th iteration

tance ratio in step 4 of approximately 0.5.

Step 3: For every parameter in Θ^*, calculate the difference between the posterior density with the candidate value and the posterior density with the previously drawn parameter value, keeping the other parameter values unchanged. Using again the first element in Θ as an example, the difference reads:

$$\begin{aligned} \delta &= \ln p\left(\{\theta_1^*, \theta_2^{mc-1}, ..., \theta_K^{mc-1}\} \mid \Theta^{mc-1}, \text{data}\right) \\ &- \ln p\left(\{\theta_1^{mc-1}, \theta_2^{mc-1}, ..., \theta_K^{mc-1}\} \mid \Theta^{mc-1}, \text{data}\right) \end{aligned} \tag{15}$$

Step 4: Draw a random number $u \sim U(0,1)$ and accept the single parameter candidate from Step 2, whenever the following holds for the difference in Step 3:

$$min(0, \delta) > log(u)$$

Step 5: Repeat the Steps 2 to 5 until the joint posterior density of parameters converge in distribution.

The algorithm is ran 100,000 times and the first 40,000 are discarded as the burn-in period. The two Markov chains with different starting values for both joint posterior and the single parameters' posteriors converge to literally the same posterior distributions. Before estimating the entire model, the proposed parametrization is used to estimate the risk neutral specification. The model under $\mathbb{Q}$ converges even quicker[39] and thus the algorithm is ran for 50,000 times and the first 20,000 are discarded as burn-in.[40]

6 Results

QE may affect either the expectations' part of interest rates or the term premium - or both. Term structure models as applied in this paper are suitable to isolate the term premium. Therefore, the three factors should account for changes in expected future interest rates. Central bank reserves as a fourth factor are added to explain changes in the term premium.

to the 40,000th iteration.

[39] On average after 7,000 sweeps, between the non-ZLB sample and the ZLB sample.

[40] The scaling factor is also automatically adapted until the 20,000th iteration.

Our model performs well in that most of the variation in yields is explained by three latent factors. The fourth factor contributes only marginally and it is only slightly correlated with the latent factors. Our estimates show that QE has significantly lowered the term premium of long-term Treasury yields. Estimates of the cumulative effect of QE on 10-year Treasury yields during the ZLB amounts to 85 basis points. This result is in line with the recent literature. Estimates of Swiss data suggest 10-year yields decreased 34 basis points which can be fully ascribed to the liquidity effect because the Swiss National bank did not purchase government bonds to expand liquidity. In the following, this section lays out the model performance and then presents the estimation results and the economic interpretation. Finally, the supply and the liquidity effect are disentangled. As described above, the model is estimated under three different settings concerning the price of risk. Accordingly, the results are reported for the three different models.

6.1 Model Performance

Table 2 and 3 report the parameter estimates for the two subsamples for US and Swiss data, respectively. The tables report the estimated posterior modes of the parameters together with numerical standard errors (in brackets) for the subsample away from the ZLB (Θ^{nonZLB}) and the subsample at the ZLB (Θ^{ZLB}). The first two columns provide the support intervals, and the last two the average acceptance ratios and inefficiency factors (IF) for the two subsamples.[41]

Table 4 reports the cross-sectional fit for both countries and the two subsamples. By construction, the 3-month, the 5-year and the 10-year yields are explained perfectly by the model. The results are therefore not reported for these maturities. During the ZLB period, the shorter end of the yield curve approached zero. This lower bound brought down interest rate volatility at shorter maturities. Therefore, the pricing performance during the ZLB period clearly outperforms the pre-ZLB pricing performance at the shorter end of the yield curve. Comparing the different market price of risk settings, the time-varying level and slope risks clearly improve the fit above the $\mathbb{Q}$ model. This is true especially in the second subsample at mid to long maturities. Letting NBR affect the market price of risk considerably improves the fit only in the ZLB-subsample. Thus, the choice of market price of risk specification turns out to be relatively more important in the ZLB period than in the pre-crisis sample. The fit also improves in the earlier subsample, but the gain is relatively smaller. Allowing

[41] The two indicators are explained in the tables' headings.

NBR to additionally affect the slope risk improves the fit further at longer maturities only during the ZLB period. During normal times, as in the earlier subsample, the addition of the time-varying slope risk lowers the fit.

A broad consensus in the term structure models literature is that the first three yields-only factors account for most of the movements in the cross-section of yields.[42] Figure 4 shows the estimated latent factors in the two subsamples. Comparing the evolution of the latent factors with the yield curve depicted in Figure 1 shows that the first latent factor seems to drive the level movements of the yield curve in both periods, being relatively stable in the first period and substantially increasing in volatility in the second. The second latent factor l_2 seems to move the slope of the yield curve, but this is only clear in the non-ZLB period. The third latent factor l_3 co-moves to a great extent with the second factor l_2. The correlation in levels between the two factors amounts to -0.75 and -0.94 in the first and the second subsample, respectively. Ang and Piazzesi (2003) and Joslin et al. (2011) show that the entire yield curve can be explained fairly well with three latent factors, and that adding macro variables does not improve a model. The variance decomposition presented in Table 5 satisfies the conjecture.[43] It indicates the fraction of the total variance explained by each factor. The fourth factor, i.e. reserves, adds only marginal information to the model while most of the cross-variation in yields is explained by the latent factors.[44] Most of the cross-sectional variation in yields is given by the first two latent factors. In the earlier subsample, the first two latent factors explain 59% of the variance of yields at the 6-month horizon. This fraction increases with maturity and amounts to over 96% at the 9-year horizon. The model performs similarly well in the ZLB subsample: the fraction of variance explained by the fist two factors increases from 70% at the shortest to 94% at the longest maturity. With the third latent factor in addition, practically the entire variance is explained. The contribution of NBR does not exceed 0.1% at any of the maturities in the first subsample. At the ZLB, however, this contribution of the NBR factor increases considerably. Liquidity explains around 1% of total variation at the shorter end of the yield curve and increases to almost 4% at longer maturities. In the Swiss case, the marginal contribution of Giro holdings falls at short maturities but quadruples at

[42] One of the key contributions in this field is Litterman and Scheinkman (1991).

[43] Table 5 reports the variance decomposition for selected yields and an instructive derivation can be found in Dahlquist and Hasseltoft (2011).

[44] Our preliminary analysis also showed that the reserves factor contribution to pricing performance was minimal. We do not report this result.

long maturities when entering the ZLB period. Hence, at the ZLB, the size of Giro holdings has gained importance in explaining long-term yields.

FIGURE 4 AND TABLE 5 ABOUT HERE

The dynamics of the NBR factor is found to be independent from the yields-only factors. Table 1 reports the correlation of NBR with the three factors: in the first subsample, the correlations are -0.09, 0.19, and -0.22, respectively. In the second subsample, the correlations decrease in size to 0.07, 0.11, and 0.07. This shows that the NBR factor adds information to the analysis that is complementary to the information contained in the yield curve.

TABLE 4 ABOUT HERE

6.2 Parameters

The coefficient estimates are depicted in the third and fourth column of Table 2 and 3 for the US and Switzerland, respectively. To start at the bottom of Table 2, the estimated loading of the NBR factor on the short rate is significantly negative in both subsamples. This is in line with the hypothesis that expansionary monetary policy has a negative effect on yields. In the ZLB sample, the coefficient is much larger in size which suggests that a one standard deviation increase has had a larger impact on yields at the ZLB bound than in normal times. Opposed to the US estimates, the bottom of Table 3 shows that Swiss data does not provide a significant coefficient on the loading of banks Giro holdings.

λ_{01} and λ_{02} refer to the time-invariant market price of risk associated to the level and the slope of the yield curve, respectively. In both countries, the market price of level risk fell substantially when the ZLB was hit.[45] This reflects the fact that zero is a natural bound for interest rates where the risk becomes one sided. Conversely, the slope risk increased. There are many possible reasons for elevated uncertainty with respect to future rates: it may reflect uncertainty about the length of the financial crisis or it could reflect uncertainty not only about the future path of unconventional monetary policy but also about its consequences.

The time-varying prices of risk coefficients, λ_{14} and λ_{24}, do not significantly differ from zero in the pre-crisis subsample in either of the two countries. At the ZLB, however, they become

[45] Note that λ refers to the discount factor of a bond price. A higher coefficient thus means a lower market price of risk in the yield.

statistically significant. λ_{14} is associated with the effect of NBR (or Giro holdings) on the market price of risk in the interest rate level. The significantly negative λ_{14} in both countries suggests that during the ZLB period, the discount factor decreases with an expansion in NBR, and so the interest rate risk increases in the level. This could be evidence that QE elevated uncertainty with respect to future interest rates or inflation which in turn affects the term premium. For example, expansionary monetary policy will eventually lead to higher inflation rates which, anticipated in the form of increasing inflation expectations, drive up nominal interest rates (Fisher effect). The term premium increases with uncertainty associated with inflation expectations. Evidence of mechanisms such as the supply or the liquidity effect is found in the market price of risk associated with the slope of the yield curve. λ_{24} turns significantly positive at the ZLB which suggests that an increase in NBR would lead to a flattening of the yield curve. Hence, with short-term interest rates were already close to zero, the supply and the liquidity effects affected the longer end of the yield curve.

TABLE 2 ABOUT HERE

6.3 The Estimated Effect of Reserves on Interest Rates

The estimation results of the factor loadings are reported in Table 6 and 7 for the US and Switzerland, respectively. The upper tables refer to the pre-crisis subsample, the lower panels to the ZLB subsample. The factor loadings of reserves on yields are presented for selected maturities and the three different market price of risk settings as described above. The factor loadings indicate by how much yields change in response to a one standard deviation change in NBR. One standard deviation of NBR corresponds to \$2.3 billion in the pre-ZLB period and rises up to \$315.8 billion in the ZLB subsample. For the Swiss giro holdings, one standard deviation amounts to CHF 0.633 billion in the early sample and to CHF 43.524 billion during the ZLB period.

TABLES 6 AND 7 ABOUT HERE

6.3.1 Estimation Results

There is only weak evidence for an effect of reserves on yields in the pre-crisis subsample. The traditional view of the liquidity effect would call for a negative effect of an increase in liquidity on the short end of the yield curve. As the upper panel of Tabel 6 shows, a one standard deviation increase in NBR is followed by a decrease of the 3-month yield of 1 to 2 basis points, depending on the model. This estimate is not significantly different from zero according to the Z-test,[46] however. As expected, there is no effect of reserves on longer-term yields in the early sample in either of the two countries. This result is robust with respect to the different market price of risk settings. The remainder of this section therefore focusses on the ZLB period.

During the ZLB period, the effect of reserves on yields becomes significant in both countries at all maturities. As shown in the lower panel of Table 6, NBR have affected the yield curve differently across the models. The third model - which allows for time variation in the prices of both level and slope risk - indicates a particularly pronounced effect at mid to longer maturities. The other two models suggest a significant impact at all maturities. The first model - which allows for no time variation in the prices of risk - estimates a constant effect across maturities. The second model - which allows only the price of level risk to vary over time - exhibits an effect that is decreasing with maturity. The estimates for Switzerland give a very similar picture: During the ZLB period, giro holdings significantly affected the yield curve, including at the longer end. The relative effect across maturities in unclear as it varies depending on the model settings. In Swiss data, the effect seems to be fairly stable across maturities.

6.3.2 Size of the Effect Relative to GDP

At the 10-year horizon, a one standard deviation increase in NBR has lowered yields between 4 to 26 basis points according to our estimate. This amounts to 1 to 8 basis points per increase in NBR of USD 100 billion. With a nominal GDP of roughly $15.1 trillion, this translates into an increase of approximately 2 to 12 basis point due to an increase in NBR of the size of one percent of GDP. Accordingly, the Swiss data yields an effect of 3 to 7 basis points per standard deviation change in liquidity, or 7 to 16 basis points per CHF 100 billion.[47] With a Swiss GDP of approximately CHF 565 billion, an increase in liquidity in the range of one percent of nominal GDP lowers rates by 0.4 to 0.9 basis points. Hence, the total effect of the liquidity expansion on long-term yields is estimated to be considerably larger in the US

[46] For the Z-test, the null hypothesis is a zero-mean and the standard error is calculated from the estimation output.

[47] The average USDCHF exchange rate in the period of the Swiss ZLB subsample from mid December to the end of September 2011 was 1.0278, so CHF 100 billion is comparable to USD 97 billion.

compared to Switzerland.

6.3.3 Cumulative Effect During the ZLB

Figure 5 plots the estimated cumulative impact of an increase in NBR on Treasury yields during the ZLB period, Figure 6 depicts the equivalent for Switzerland. The graphs on the left refer to the 3-month yield, the graphs on the right to the 10-year rate. The upper row shows the estimates from the model under the $\mathbb{Q}$ measure, the middle row to the model allowing the level risk of interest rates to vary over time, and the bottom row to the model allowing both the level and slope risks of interest rates to be time varying. As the 90% confidence bands indicate the statistically significant effect of reserves on yields. The total increase in NBR of roughly 1,028 billion USD is associated with a fall in the 5-year yield of 98 basis points and a fall in the 10-year yield of 85 basis points.

FIGURE 5 ABOUT HERE

The other two market price of risk specifications estimate a smaller effect. The model under the $\mathbb{Q}$-measure estimates a total decrease of the 10-year Treasury yields of roughly 13 basis points due to the increase in NBR. The model accounting for a time-varying level risk estimates a fall in the long-term yields of approximately 52 basis points.

In Switzerland the cumulative effect of the expansion in Giro holdings from approximately 26 billion francs in December 2008 to 235 billion francs in October 2011 has led to an estimated fall in long-term yields between 14 to 34 basis points. The 90%-confidence bands are rather large as Figure 6 shows, however. But in all models, the cumulative impact is significantly different from zero.

FIGURE 6 ABOUT HERE

The size of the estimated impact for US data is in line with the existing literature. The 26 basis point fall in rates per standard deviation in NBR corresponds to an effect of 8 basis points per USD 100 billion. Gagnon et al. (2011) find a total supply effect 30 to 100 basis points which translates into 2 to 6 basis points per 100 billion USD. Hamilton and Wu (2010a) estimate a range of 17 to 48 basis points, equivalent to 12 to 17 basis points per USD 100 billion. D'Amico and King (2010) estimate an effect of 50 basis points in response to purchases of USD 300 billion (17 basis points per USD 100 billion), and Krishnamurthy and Vissing-Jorgensen (2011) estimate the total effect of the LSAP purchases to amount to roughly 100 basis points. Finally, Krogstrup et al. (2012) estimate an effect of 26 basis points per USD 100 billion of outright Treasury purchases on 10-year Treasury yields.

6.4 Disentangling the Different Effects

The signalling effects of monetary policy as well as expectations about the future path of interest rates should be captured by the first three factors of the model. Hence, the fourth factor contributes information on the term premium. This paper compares the estimates of US data with those of Swiss data to separate the supply effect from the liquidity effect. The former cannot be present in Swiss data because the SNB has not bought Swiss government bonds during the ZLB period. Up to our knowledge, the only paper that tries to disentangle the two effects is Krogstrup et al. (2012) who conduct the study using US data. According to their estimates, the cumulative liquidity effect exceeds the supply effect by a factor of three.

The fact that there is a significant effect of giro holdings on Swiss long-term yields provides evidence for the liquidity effect. Per standard deviation change in reserves, the effect in the US results up to almost four times the effect in Switzerland. In terms of a 1% increase of GDP, the estimated effect for the US exceeds the one for Switzerland by up to the twelvefold. This suggests that in the US, there were other effects at play on top of the liquidity effect. We conclude from these results that the liquidity effect is between a twelfth and a quarter the size of the supply effect. If three quarters of the estimated impact in the US are due to supply effects and one quarter to liquidity effects, the former would amount to and accumulated effect of 64 basis points and the latter to 21 basis points during the ZLB sample.

Figure 7 shows the development of non-borrowed reserves and the Federal Reserves holdings of Treasury securities with maturities of more than one year. It shows that the correlation is limited. The Treasury securities increased in two steps: in Q2 of 2009 and again in Q4 2011. Non-borrowed reserves increased the most in Q3 of 2009 and fluctuated much more than the Treasury holdings. Assuming that our estimates capture the liquidity effect and the supply effect, the difference in the change of non-borrowed reserves and the change in Treasury holdings would be a good measure of the "pure" liquidity increase. The lower panel of Figure 7 draws this "pure" cumulative liquidity effect. The bold black line depicts the liquidity effect with the smallest factor loading of 4 basis points per standard deviation which results from the model under the $\mathbb{Q}$-measure. The dashed red line depicts the liquidity effect with the largest factor loading of 26 basis points per standard deviation estimated by the third model.

These computations suggest that the cumulative "pure" liquidity effect amounted to 3 to 20 basis points. The rest of the effect is attributable to a mix between the supply and the liquidity effect. Thus, the 20 basis points most likely underestimate the cumulative liquidity effect.

Finally, information theory would suggest that supply effects are to a great extent priced in at the moment of the announcement (or even earlier if expected). If this is true, then NBR can only explain supply effects of Treasury securities that exceeded expectations. Conversely, the liquidity effect can only work once reserves have increased. Thus, our estimates for the liquidity effect are likely to underestimate the true liquidity effect of QE on the yield curve.

7 Conclusion

QE may affect either the expectations' part of interest rates or the term premium - or both. Term structure models as applied in this paper are suitable to isolate the term premium. Therefore, three latent factors should account for changes in expected future interest rates. Central bank reserves as a fourth factor are added to explain changes in the term premium. Our estimates show that QE has significantly lowered the term premium of long-term Treasury yields. The results suggest that an increase in liquidity has affected the entire yield curve at the ZLB in both US and Swiss data. The impact amounts up to 8 basis points per USD 100 billion in the US. This is line with the literature that estimates an effect of 2 to 26 basis points per USD 100 billion. On a cumulative basis, our estimates suggest that long-term yield have been lowered by up to 85 basis points in the US and 34 basis points in Switzerland.

The Swiss results provide evidence of liquidity effect at the longer end of the yield curve when the short-term interest rates are bounded by the ZLB because the Swiss national bank expanded liquidity by foreign exchange rather than government bond purchases. Therefore, a supply effect can be excluded. While the estimate is entirely attributable to the liquidity effect in Switzerland, the total effect in the US must be divided into the supply effect and the liquidity effect. According to our estimates, the estimated effect in the US exceeds the one in Switzerland by the threefold. If we deduce from this ratio that one quarter of the total effect in the US is due to the liquidity effect, it would amount to a cumulative 21 basis points in the ZLB sample. The second approach to isolate the liquidity effect yields the same result. We deduct the Treasury supply purchases from the changes in NBR and compute the liquidity effect by solely considering the increases in NBR that were not due to Treasury securities purchases. This yields a "pure" liquidity effect of 20 basis points. The remaining 64 basis points can be ascribed to a mix between the supply and the liquidity effect.

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8 Appendix - Tables and Figures

Table 1: *Sample Correlations.* The table reports contemporaneous correlations of selected interest rates and central bank reserves in levels (upper panel) and first differences (lower panel) for "normal times" and the ZLB subsamples.

Levels						
Non-ZLB						
	3M	1Y	3Y	6Y	8Y	10Y
USA	-0.39	-0.37	-0.31	-0.26	-0.25	-0.24
CH	0.19	0.19	0.14	0.16	0.16	0.15
ZLB						
	3M	1Y	3Y	6Y	8Y	10Y
USA	-0.54	-0.84	-0.61	-0.41	-0.38	-0.36
CH	-0.28	-0.27	-0.64	-0.69	-0.59	-0.53
First Differences						
Non-ZLB						
	3M	1Y	3Y	6Y	8Y	10Y
USA	0.05	-0.05	-0.21	-0.20	-0.19	-0.19
CH	0.29	0.10	0.03	-0.02	-0.06	-0.08
ZLB						
	3M	1Y	3Y	6Y	8Y	10Y
USA	-0.13	-0.17	-0.17	-0.15	-0.15	-0.15
CH	-0.40	-0.44	-0.28	-0.12	-0.10	-0.12

Table 2: *Parameter Estimates for the US.* The table reports the estimated posterior modes of the parameters for the US, together with numerical standard errors (in brackets) for the subsample away from the ZLB (Θ^{nonZLB}) and the subsample at the ZLB (Θ^{ZLB}). The first two columns provide the support intervals, and the last two the average acceptance ratios and inefficiency factors (IF) for the two subsamples. The acceptance ratio is the number of accepted parameters' proposals divided by the number of iterations after burn-in. The rate between 0.25 and 0.75 is often acceptable, see Lynch (2007) and Koop (2003). Inefficiency Factor (IF) of how well the sampler "mixes" is computed as $1+2\sum_{l=1}^{L}\rho(l)$, where $\rho(l)$ is the autocorrelation at lag l in the Markov chain sequence of a parameter, and L is the lag at which the autocorrelation function goes to zero. See Chib (2001) for details.

	$\underline{\Theta}$	$\overline{\Theta}$	Θ^{nonZLB}	Θ^{ZLB}	AccRatio	IF
α	0.00	0.99	0.92	0.93	0.70	811
			(0.00)	(0.00)		
ψ_{NBR}	0.00	1.00	0.84	1.00	0.52	61
			(0.04)	(0.00)		
σ_{11}	0.00	15.0	0.08	0.14	0.20	612
			(0.00)	(0.00)		
σ_{22}	0.00	15.0	0.22	0.33	0.41	78
			(0.01)	(0.01)		
σ_{33}	0.00	15.0	0.30	0.26	0.20	59
			(0.01)	(0.01)		
σ_{NBR}	0.00	15.0	1.07	0.09	0.33	11
			(0.05)	(0.00)		
λ_{01}	-100	100	1.87	5.10	0.57	1108
			(0.22)	(0.26)		
λ_{02}	-100	100	0.83	-2.00	0.26	759
			(0.23)	(0.24)		
λ_{14}	-100	100	-0.01	-0.08	0.80	166
			(0.01)	(0.02)		
λ_{24}	-100	100	0.00	0.09	0.68	291
			(0.01)	(0.05)		
ω	0.00	10.0	0.21	0.22	0.21	61
			(0.01)	(0.01)		
b_{NBR}	-1.00	1.00	-0.02	-0.24	0.35	335
			(0.02)	(0.20)		

Table 3: *Parameter Estimates for Switzerland.* The table reports the estimated posterior modes of the parameters for Switzerland, together with numerical standard errors (in brackets) for the subsample away from the ZLB (Θ^{nonZLB}) and the subsample at the ZLB (Θ^{ZLB}). The first two columns provide the support intervals, and the last two the average acceptance ratios and inefficiency factors (IF) for the two subsamples. The acceptance ratio is the number of accepted parameters' proposals divided by the number of iterations after burn-in. Inefficiency Factor (IF) of how well the sampler "mixes" is computed as $1 + 2\sum_{l=1}^{L} \rho(l)$, where $\rho(l)$ is the autocorrelation at lag l in the Markov chain sequence of a parameter, and L is the lag at which the autocorrelation function goes to zero. See Chib (2001) for details.

	$\underline{\Theta}$	$\overline{\Theta}$	Θ^{nonZLB}	Θ^{ZLB}	AccRatio	IF
α	0.00	0.99	0.90	0.92	0.66	195
			(0.00)	(0.00)		
$\psi_{SightDepo}$	0.00	1.00	0.99	1.00	0.62	33
			(0.01)	(0.00)		
σ_{11}	0.00	15.0	0.06	0.08	0.22	378
			(0.00)	(0.00)		
σ_{22}	0.00	15.0	0.21	0.18	0.38	63
			(0.01)	(0.01)		
σ_{33}	0.00	15.0	0.27	0.17	0.20	48
			(0.01)	(0.01)		
$\sigma_{SightDepo}$	0.00	15.0	0.83	0.17	0.25	10
			(0.04)	(0.01)		
λ_{01}	-100	100	1.58	1.95	0.60	1033
			(0.17)	(0.21)		
λ_{02}	-100	100	0.31	-1.06	0.27	1501
			(0.25)	(0.20)		
λ_{14}	-100	100	0.01	-0.04	0.81	184
			(0.01)	(0.02)		
λ_{24}	-100	100	-0.01	0.10	0.75	201
			(0.02)	(0.05)		
ω	0.00	10.0	0.20	0.20	0.21	27
			(0.01)	(0.01)		
$b_{SightDepo}$	-1.00	1.00	-0.01	0.03	0.23	181
			(0.03)	(0.08)		

Table 4: *Pricing Errors.* Mean absolute pricing errors in basis points across different modelling assumptions are reported for selected yields in the US (upper panel) and Switzerland (lower panel), and for the two subsamples.

USA

non-ZLB

	6M	1Y	3Y	6Y	8Y	9Y
$\mathbb{Q}$	11.6	12.5	3.5	1.9	4.3	3.4
l	9.7	7.9	3.1	1.1	1.7	1.3
l & s	10.1	8.9	3.3	1.3	2.3	1.8

ZLB

	6M	1Y	3Y	6Y	8Y	9Y
$\mathbb{Q}$	5.5	8.0	6.6	4.2	5.2	3.0
l	3.6	6.6	3.1	1.1	2.0	1.4
l & s	4.2	7.5	3.6	1.1	1.6	1.1

CH

non-ZLB

	6M	1Y	3Y	6Y	8Y	9Y
$\mathbb{Q}$	6.7	12.7	5.0	1.1	2.7	4.7
l	9.2	17.0	3.8	0.9	3.0	4.9
l & s	9.0	16.6	3.7	0.9	3.0	4.9

ZLB

	6M	1Y	3Y	6Y	8Y	9Y
$\mathbb{Q}$	14.2	10.0	6.2	1.5	7.6	3.9
l	14.9	11.0	5.2	1.5	7.9	4.0
l & s	15.1	11.0	5.3	1.4	7.1	3.4

Table 5: *Variance Decomposition.* The table reports variance decomposition of selected yields in % for the US (upper panel) and Switzerland (lower panel). The variance is decomposed by dividing each single state variable shock j to an n-periods yield: $MSE_n^j = B_n'\Sigma^j B_n + B_n'\Psi\Sigma^j\Psi B_n$, where Σ^j is a $K \times K$ matrix with zeros and a non-zero jj element corresponding to the volatility of state variable j; with the overall Mean Squared Error of forecasting the states 1 period ahead: $MSE_n = B_n'\Sigma B_n + B_n'\Psi\Sigma\Psi B_n$.

US

non-ZLB

	6M	1Y	3Y	6Y	8Y	9Y
l_1	17.4	22.1	40.5	62.3	72.3	76.1
l_2	41.6	44.8	44.6	31.5	23.6	20.5
l_3	40.9	33.0	14.8	6.2	4.1	3.4
NBR	0.1	0.1	0.1	0.0	0.0	0.0

ZLB

	6M	1Y	3Y	6Y	8Y	9Y
l_1	23.2	27.6	43.7	62.0	70.7	74.2
l_2	47.1	48.4	44.1	30.5	23.1	20.1
l_3	28.8	22.8	9.7	3.9	2.5	2.1
NBR	0.9	1.2	2.5	3.6	3.7	3.6

CH

non-ZLB

	6M	1Y	3Y	6Y	8Y	9Y
l_1	34.3	40.8	61.3	79.5	86.2	88.5
l_2	43.4	42.9	32.9	18.5	12.6	10.6
l_3	21.0	15.2	5.1	1.7	1.1	0.9
$SightDepo$	1.3	1.2	0.7	0.2	0.1	0.1

ZLB

	6M	1Y	3Y	6Y	8Y	9Y
l_1	43.8	49.8	67.9	82.9	88.2	90.1
l_2	41.8	40.0	28.9	15.8	10.7	9.0
l_3	14.0	10.0	3.2	1.1	0.6	0.5
$SightDepo$	0.3	0.2	0.0	0.3	0.4	0.4

Table 6: *Factor Loadings for the US.* The table reports the values of the NBR-factor loadings for selected yields, together with the Z-score from the Z-test (in parenthesis) and for the non-zero-lower-bound period (upper panel) and zero-lower-bound period (lower panel). The levels of significance of 0.1, 0.05 and 0.01 are denoted with *, ** and ***, respectively.

			non-ZLB			
			$\mathbb{Q}$			
	3M	1Y	3Y	6Y	8Y	10Y
B	-0.01	-0.01	-0.00	-0.00	-0.00	-0.00
Z-score	(-0.76)	(-0.77)	(-0.77)	(-0.76)	(-0.76)	(-0.75)
			level risk			
	3M	1Y	3Y	6Y	8Y	10Y
B	-0.01	-0.00	-0.00	0.00	0.00	0.00
Z-score	(-0.73)	(-0.61)	(-0.26)	(0.13)	(0.29)	(0.41)
			level & slope risk			
	3M	1Y	3Y	6Y	8Y	10Y
B	-0.02	-0.01	-0.01	-0.00	-0.00	-0.00
Z-score	(-1.34)	(-1.45)	(-1.43)	(-1.00)	(-0.79)	(-0.62)
			ZLB			
			$\mathbb{Q}$			
	3M	1Y	3Y	6Y	8Y	10Y
B	-0.04*	-0.04*	-0.04*	-0.04*	-0.04*	-0.04*
Z-score	(-1.85)	(-1.85)	(-1.85)	(-1.84)	(-1.84)	(-1.84)
			level risk			
	3M	1Y	3Y	6Y	8Y	10Y
B	-0.31***	-0.30***	-0.27***	-0.22***	-0.19***	-0.16***
Z-score	(-3.83)	(-3.82)	(-3.75)	(-3.52)	(-3.24)	(-3.06)
			level & slope risk			
	3M	1Y	3Y	6Y	8Y	10Y
B	-0.24	-0.26	-0.29**	-0.30***	-0.28***	-0.26***
Z-score	(-1.21)	(-1.47)	(-2.13)	(-2.84)	(-3.03)	(-3.04)

Table 7: *Factor Loadings for Switzerland.* The table reports the values of the reserves-factor loadings for the pre-ZLB period (upper panel) and ZLB-period (lower panel) for selected yields, together with the Z-score from the Z-test (in parenthesis). All the estimates regard the zero-lower-bound subsample. The levels of significance of 0.1, 0.05 and 0.01 are denoted with *, ** and ***, respectively.

pre-ZLB						
$\mathbb{Q}$						
	3M	1Y	3Y	6Y	8Y	10Y
B	-0.00	-0.00	-0.00	-0.00	-0.00	-0.00
Z-score	(-0.72)	(-0.72)	(-0.72)	(-0.72)	(-0.72)	(-0.72)
level risk						
	3M	1Y	3Y	6Y	8Y	10Y
B	0.01	0.01	0.01	0.00	0.00	-0.00
Z-score	(1.56)	(1.50)	(1.28)	(0.77)	(0.21)	(-0.67)
level & slope risk						
	3M	1Y	3Y	6Y	8Y	10Y
B	-0.01	-0.01	-0.01	-0.01	-0.01	-0.01
Z-score	(-0.24)	(-0.26)	(-0.36)	(-0.70)	(-1.01)	(-1.33)
ZLB						
$\mathbb{Q}$						
	3M	1Y	3Y	6Y	8Y	10Y
B	-0.04***	-0.04***	-0.04***	-0.03***	-0.03***	-0.03***
Z-score	-(2.64)	-(2.65)	-(2.66)	-(2.67)	-(2.68)	-(2.68)
level risk						
	3M	1Y	3Y	6Y	8Y	10Y
B	-0.08***	-0.08***	-0.08***	-0.08***	-0.07***	-0.07***
Z-score	-(2.59)	-(2.63)	-(2.75)	-(2.90)	-(2.96)	-(2.97)
level & slope risk						
	3M	1Y	3Y	6Y	8Y	10Y
B	0.03	0.00	-0.04	-0.06*	-0.07**	-0.07**
Z-score	(0.40)	(0.08)	-(0.82)	-(1.84)	-(2.16)	-(2.26)

Figure 1: *Interest Rates and Central Bank Reserves in the US.* The term structure of interest rates (left scale) is plotted for the non-ZLB (upper panel) and ZLB-period (lower panel), together with the NBR(right scale).

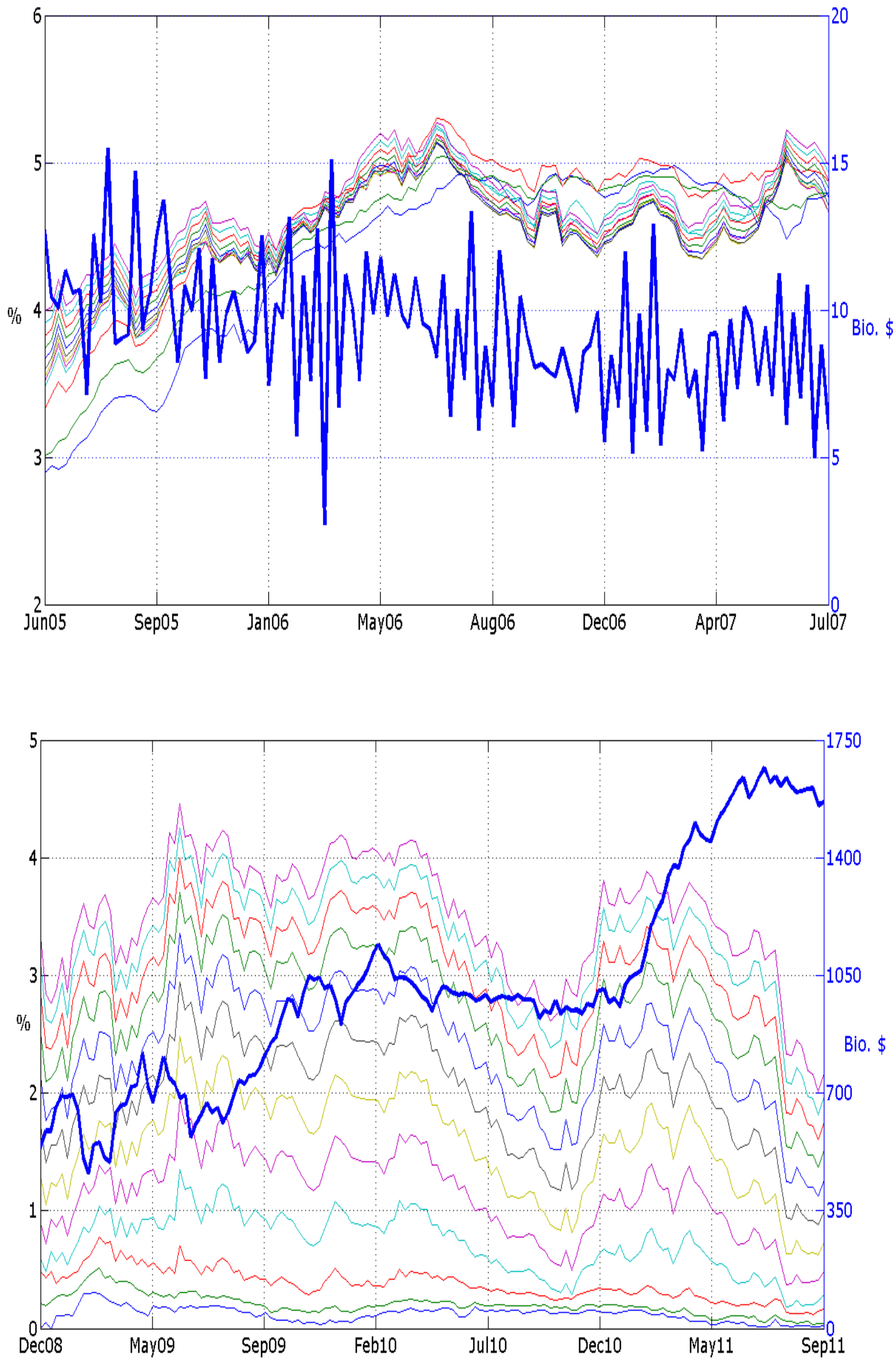


Figure 2: *Interest Rates and Central Bank Reserves in Switzerland.* The term structure of interest rates (left scale) is plotted for the non-ZLB (upper panel) and ZLB-period (lower panel), together with the total sight deposits (right scale).

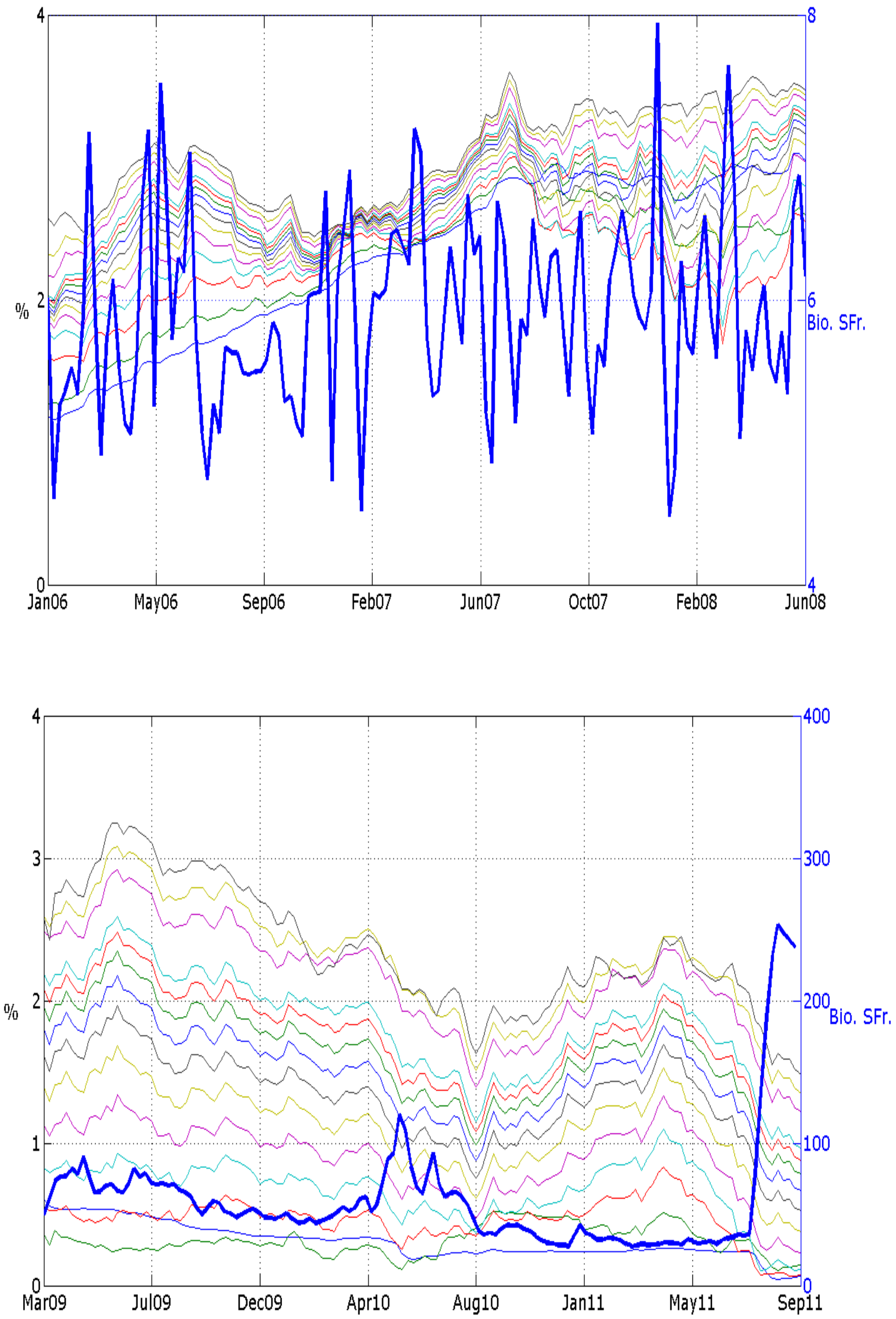


Figure 3: *Estimated Posteriors for the US.* The Figure illustrates the estimated posterior densities for all the parameters of the model used to estimate the liquidity effect in the US data.

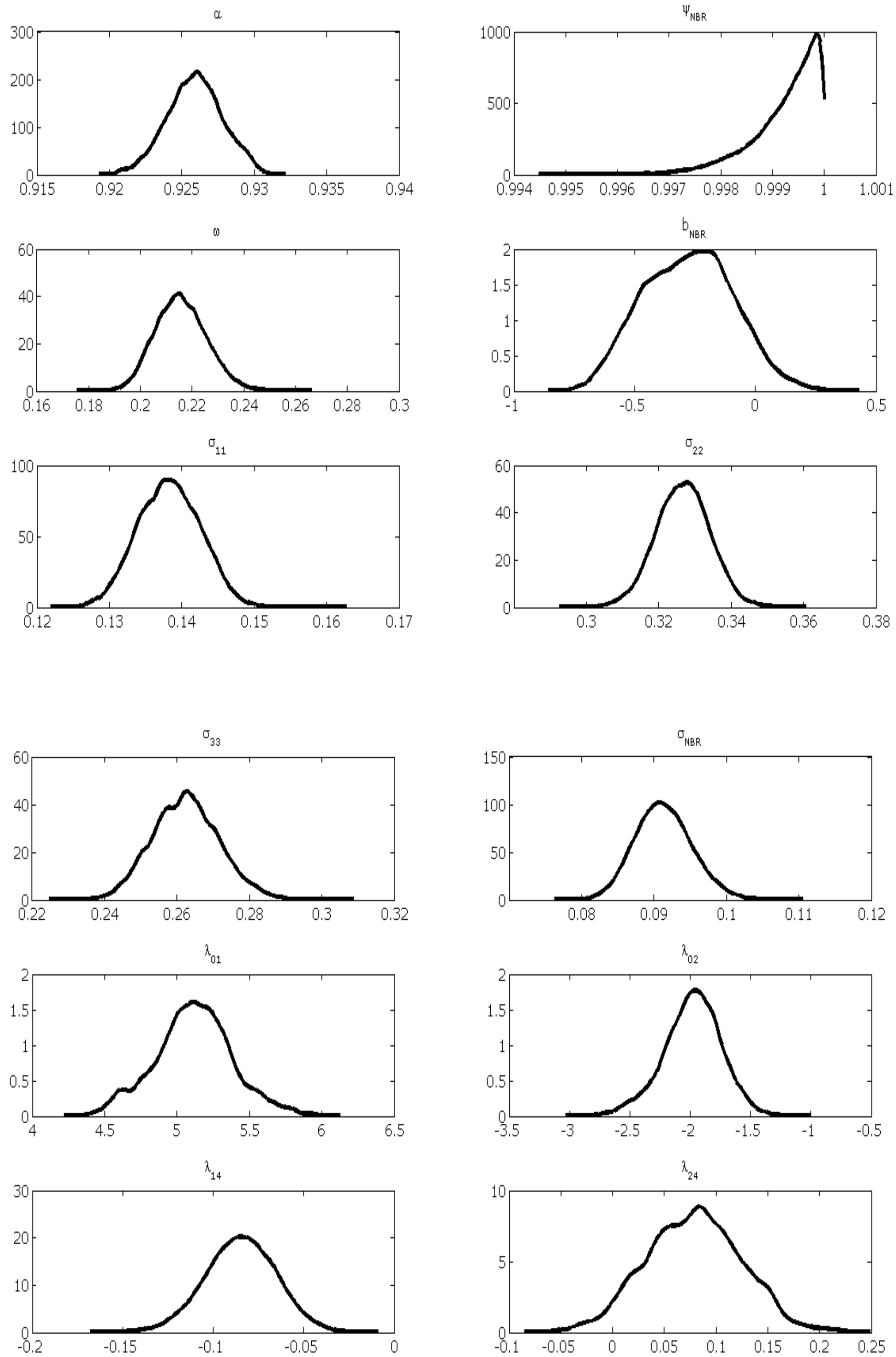
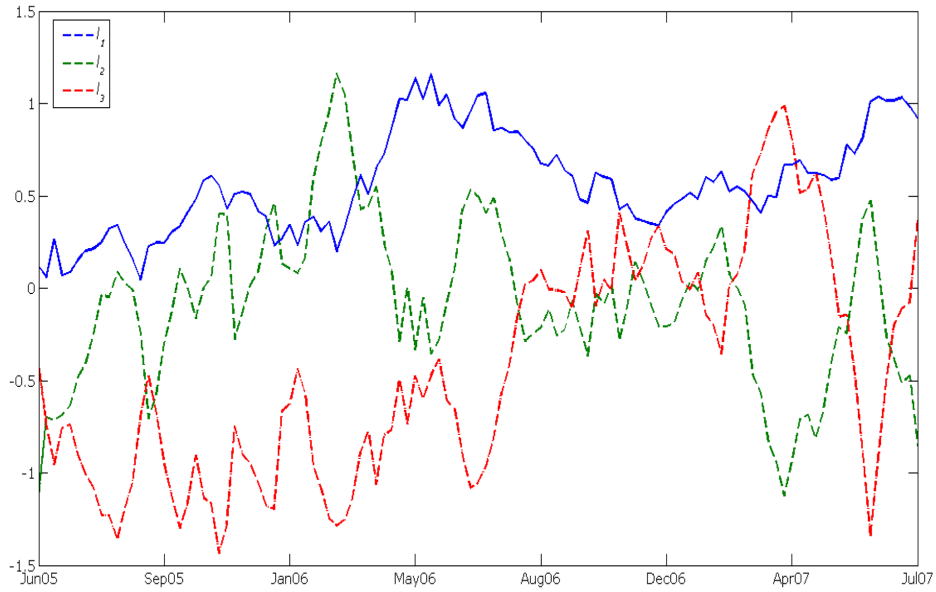


Figure 4: The Figure plots the evolution of the three estimated latent factors in the period away from the zero lower bound (upper panel) and at the zero lower bound (lower panel) for US data.



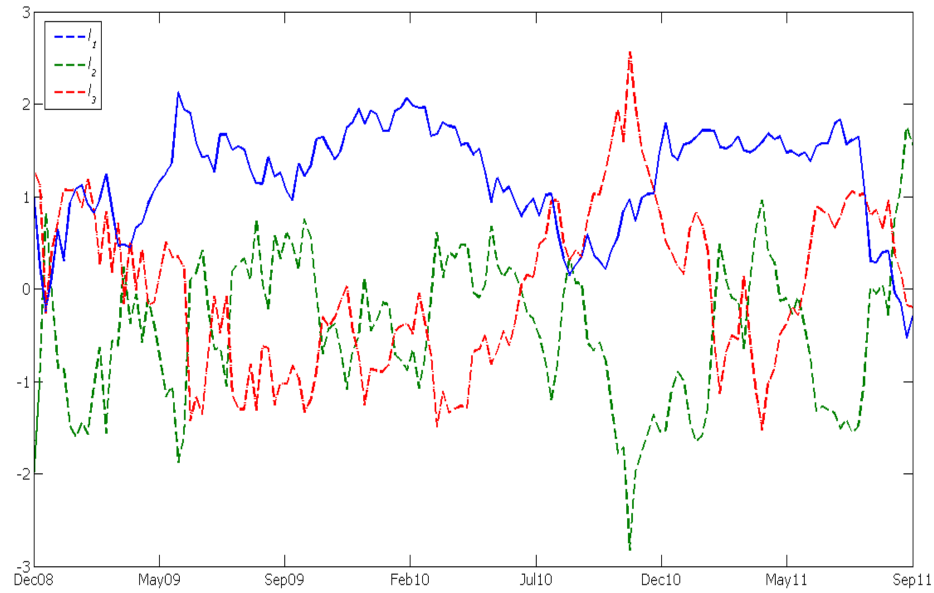


Figure 5: *Cumulative Effects for the US.* The figure plots cumulative effect of increase in NBR on the 3-month (left-hand side panels) and 10-year (right-hand side panels) interest rates according to the $\mathbb{Q}$-model (top row), the l-model (middle row) and $l\&s$-model (bottom row plots) together with 90% credible intervals. The credible interval is calculated using the posterior parameters' distributions, i.e. every 1000th set of parameters along the first Markov chain and after burn-in.

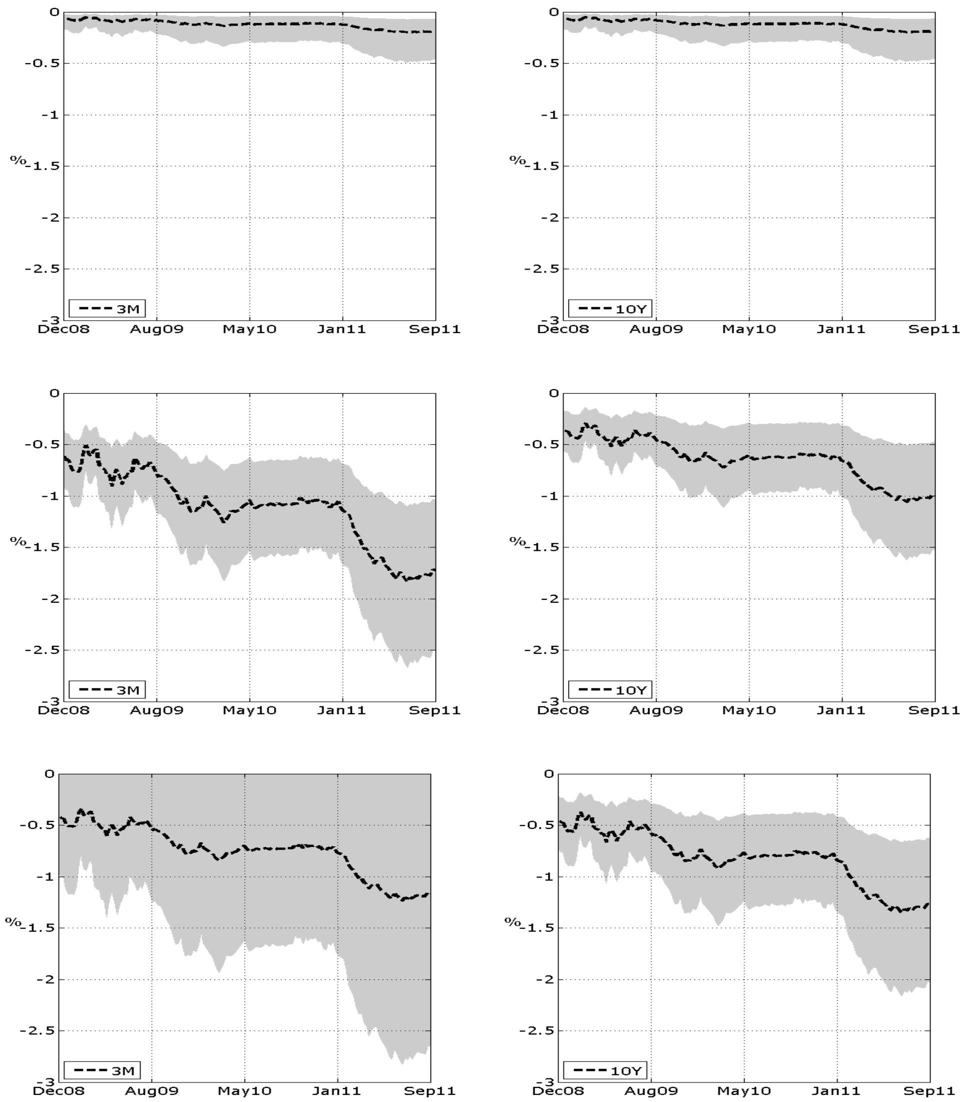


Figure 6: *Cumulative Effects for Switzerland.* The figure plots cumulative effect of increase total sight deposits at the SNB on the 3-month (left-hand side panels) and 10-year (right-hand side panels) interest rates according to the $\mathbb{Q}$-model (top row), the l-model (middle row) and $l\&s$-model (bottom row plots) together with 90% credible interval. The credible interval is calculated using the posterior parameters' distributions, specifically every 1000th set of parameters along the first Markov chain and after burn-in.

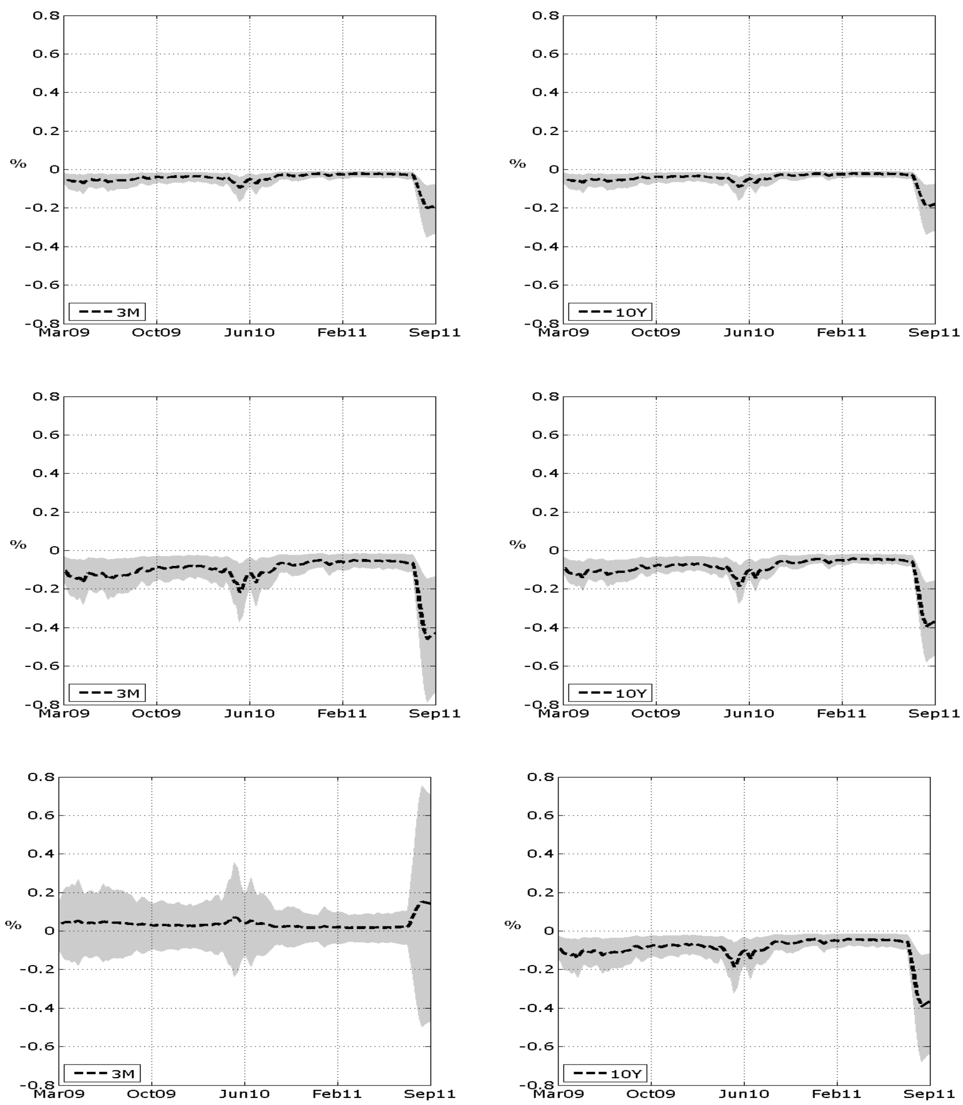


Figure 7: *Separating Supply and Liquidity Effect in US Data.* The upper panel depicts weekly data of non-borrowed Reserves and the Federal Reserves' holdings of Treasury securities with more than 1 year of maturity. The lower panel draws the cumulative liquidity effect by multiplying the difference in the changes of reserves and Treasury holdings by the factor loadings of 10-year maturity as estimated in Table 6. The solid black line refers to the most conservative estimate at the ZLB which results under the $\mathbb{Q}$-measure, the dashed red line to the most optimistic estimate that results from the third model.