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Monetary Policy at the Zero Lower Bound: Information in the Federal Reserve's Balance Sheet ⋆

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Abstract

I examine the impact of the actual purchases of Treasury securities by the Federal Reserve on Treasury yields. Using structural stability tests I find significant breaks in the relation between these variables. I find that in the zero lower bound period following the first phase of quantitative easing, May 2010 to December 2015, the actual purchases of Treasury securities by the Federal Reserve are positively related to changes in Treasury yields. This effect is driven primarily by the positive relation of the Treasury purchases with the bond risk premium, but they are also positively related to the expected inflation rate and the real rate of interest. The evidence is consistent with the liquidity channel hypothesis as put forward by Krishnamurthy and Vissing-Jorgensen (2011), since the Federal Reserve's Treasury purchases also strongly predict a lower corporate yield spread. Using a macro-finance term structure model I provide counterfactual estimates of the Treasury yields in the zero lower bound period. The counterfactual 10-year yield and the term premium are considerably smaller during QE2 but close to the actual time series for most of the rest of the lower bound period.

Keywords: quantitative easing, zero lower bound, unconventional monetary policy, Treasury yields, liquidity channel.
JEL classification: E4, E5, G1.

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1. Introduction

In ‘normal’ times the federal funds rate is considerably positive and is the main instrument used by the Federal Reserve to conduct monetary policy. Faced with the greatest economic crisis since the great depression, the Federal Reserve set this rate to zero in December 2008. Since the federal funds rate could no longer be used to conduct monetary policy, the Fed adopted unconventional monetary policy (hence UMP) measures to boost real activity, prevent deflation and improve the functioning of financial markets. One of these measures was large-scale asset purchases, also commonly known as Quantitative Easing (QE), which is based on expanding the size of the central bank’s balance sheet, in the hope of stimulating the economy (Bernanke, Reinhart and Sack, 2004). In this paper, using the information on the Federal Reserve’s balance sheet, I provide new evidence on the effects of the large-scale asset purchases at the zero lower bound (ZLB).

One of the stated goals of the large-scale asset purchases was to reduce the long-term interest rates (Dudley, 2010). Lower interest rates were supposed to incite consumption and investment spending, and overcome the threat of deflation looming at the outset of the global financial crisis (Joyce, Tong and Woods, 2011). The successful realization of this objective has been confirmed by a number of studies (see e.g. Gagnon, Raskin, Remache and Sack, 2011, Krishnamurthy and Vissing-Jorgensen, 2011, Hamilton and Wu, 2012,Neely, 2010 among others), mostly relying on the event studies of the QE announcements. However, as noted in a Greenlaw, Hamilton, Harris and West (2018), by selecting the event dates arbitrarily, the results reported in these papers can suffer from a number of potential biases. By taking a systematic approach to the FOMC announcement dates and Fed news days, Greenlaw et al. show that the overall news effect of QE on interest rates was positive rather than negative. This paper further contributes to the sceptical assessment of the effectiveness of QE as a tool to lower interest rates. I examine the effects of the actual purchases of Treasury securities by the Federal Reserve using regression analysis. I control for the FOMC announcements by including the dummy variables for all FOMC meeting weeks.

The Efficient Market Hypothesis predicts that, with rational, forward-looking agents, the actual purchases of the securities should not affect asset prices if they are perfectly anticipated. As such, the findings that the (presumably anticipated) Treasury securities purchases by the Federal Reserve tend to *increase* the interest rates seem to contradict the efficient markets hypothesis. For example, the fact that anticipated purchases of Treasury securities positively correlate with bond yields could be explained by market overreaction at the announcement times.[1] However, if the purchases contain some information beyond that

[1]Lou, Yan and Zhang (2013) document bond price reaction around anticipated Treasury auctions.

already included in the announcements, the price effect can be consistent with the Efficient Market Hypothesis. An alternative explanation would be that the relation is spurious, in the sense that once the QE was implemented, the economic conditions were improving, which increased the probability of the interest rate lift-off and consequently moved up the term structure of interest rates. I try to rule out the latter possibility by controlling for a spectrum of variables measuring the economic and financial market conditions.

To eliminate the possible endogeneity bias, I examine the relation between the Federal Reserve purchases and Treasury prices in a predictive manner. This allows me to investigate the implications of the Federal Reserve's purchases for the term premium and expected future interest rates. The modern finance theory sees the yields as a risk-neutral expectation of the average short rate and a term and liquidity premium. I find that the strong price reaction of Treasury securities is primarily driven by a rise in the term premium. This result is consistent with other studies that found that monetary policy shocks have a large impact on the risk premium, which, in turn, moves the long-term rates (Gilchrist, López-Salido and Zakrajšek, 2015, Hanson and Stein, 2015). Gertler and Karadi (2015) argue that credit conditions are very sensitive to monetary policy shocks through the term premium channel, which in turn affects the real economy in the long run and thus the long-term rates.

Although my findings are consistent with these papers, I find that the relationship has not been stable. In particular, in a comprehensive analysis of parameter stability I find three significant structural breaks. The first break occurs in October 2007 and can be attributed to the turmoil around the solvency problems of BNP Paribas and Bear Sterns. The second break is estimated at October 2008, the peak of the financial crisis following the collapse of Lehman Brothers and multiple banks bailouts. Finally, I found a third break estimated at March 2010, which coincides with the end of the QE1. Thus, to analyse the impact of the QE, I focus on the ZLB period when the structural model parameters were stable: May 2010 to December 2015.[2]

One of the important implications of these structural breaks is to challenge the validity of the approach to the counterfactual analysis of the impact of QE on yields (or other macro variables) that is prevalent in the literature, based on the estimates from the normal period (see the literature review in Section 2). To find the effect of Treasury purchases, I employ the Joslin, Priebsch and Singleton (2014) macro-finance term structure model and eliminate the effect of Federal Reserve purchases on other variables in the ZLB period, which allows us to obtain counterfactual Treasury yields and term premium. My analysis indicates that

[2]In a related paper Swanson and Williams (2014) find the reduced sensitivity of the bond yields to the macroeconomic news in the ZLB period. My results are different, as I find a change in the sign of the relation between the variables of interest.

there were periods when the impact of Treasury purchases by the Federal Reserve on the long-term nominal rates was significant, particularly during QE2 when the observed yields were higher than counterfactual by about 0.7%, but they converged together towards the end of the ZLB period.

Although this paper is predominantly empirical, it can be noted that the findings are consistent with the liquidity channel as put forward by Krishnamurthy and Vissing-Jorgensen (2011). The liquidity channel hypothesis states that as the Federal Reserve swaps Treasury bonds for more liquid reserve balances, investors have more liquidity at hand and, therefore, the (negative) liquidity premium in Treasury bonds decreases, which increases their yields relative to other, less liquid assets. Indeed, in the aftermath of the subprime mortgage crisis culminating in the bankruptcy of Lehman Brothers in September 2008, the financial world witnessed an unprecedented decline in the liquidity of financial markets and a collapse of the shadow banking system. The consequence of this economic turmoil was the flight-to-liquidity, which is a rapid transfer of capital from more risky assets (the stock market) to assets that are considered to be safer and more liquid (the Treasury market). In consequence, the price of bonds increased more than predicted by traditional asset pricing models, thus creating a negative liquidity premium in bond prices (see Figure 1 in Krishnamurthy and Vissing-Jorgensen, 2011). This mechanism is consistent with the theoretical model of flight-to-liquidity proposed by Vayanos (2004).[3]

An important prediction of the liquidity transmission channel is that, after a flight-to-liquidity crisis, an increase in QE should relieve the liquidity pressure leading to an *increase* in yields on Treasury securities in relation to less liquid assets. I find confirmation of these predictions in the data. In particular, I find that in the ZLB period the purchases of Treasury securities by the Federal Reserve are associated not only with higher nominal and real rates, but also with lower corporate yields relative to Treasury yields, and lower yield spread in corporate bonds for different safety, and thus liquidity, classes. Also, I find that during the ZLB period the Treasury security purchases predict stock market returns. These effects have strong statistical significance and are economically meaningful, but only in the ZLB period; in the normal period, increasing the balance sheet of the Federal Reserve had a negligible impact on interest rates and other variables.

[3] Gagnon et al. (2011) consider another type of liquidity channel that is associated with market functioning. They conjecture that after the outbreak of the financial crisis the Treasury securities had a liquidity premium due to market strains. By providing constant demand for these instruments Federal Reserve decreased the liquidity premium that dealers and other market participants required for holding long-term assets, which should lead to higher bond prices and lower interest rates.

2. Related literature

The literature on the impact of QE has been burgeoning since the beginning of the financial crisis of 2007-2008 and I do not intend to mention all the related papers, as this would be virtually impossible. Instead, I limit my attention to the papers that are most relevant to my research.

In a comprehensive study, Krishnamurthy and Vissing-Jorgensen (2011) consider the effectiveness of QE through several transmission mechanisms: signalling, duration risk, liquidity, safety, prepayment risk premium, default risk and inflation channel. They employ a range of tools and methods to separately assess each of them. My objective is more narrow; I evaluate the impact of the security purchases by the Federal Reserve on the interest rates. Together with additional evidence from the corporate bond market, I find support for the liquidity transmission channel.

To assess the effectiveness of QE, Gagnon et al. (2011) use two investigation methods: event studies and time series regressions. In time series regressions they regress Kim and Wright (2005) term premium on publicly held debt (excluding the Federal Reserve's holdings and foreign agencies) to examine the portfolio substitution channel. They use the sample January 1985 to June 2008 to draw an inference about the Large Scale Asset Purchases. I argue that due to the break at the beginning of the financial crisis the inference about the impact of QE based on earlier data is questionable. In contrast to Gagnon et al. (2011), I run the regression of changes (not levels) in the term premium on the balance sheet variables before and after the federal funds rate hit the ZLB. A similar methodology is used by Kozicki, Santor and Suchanek (2012), who run time series regressions of long-term forward rates on central bank claims and assets. They too draw their inference from regressions on the pre-crisis period, 1980 to 2007, and conclude that expansion of the Federal Reserve's balance sheet is associated with a negative effect on the interest rates. Their results are consistent with my findings for the normal period. The structural break at the ZLB means, however, that this conclusion cannot be extended to the ZLB period. D'Amico, English, López-Salido and Nelson (2012) regress nominal yields and nominal, real and inflation risk premia on privately held Treasury debt (thus excluding the Federal Reserve holdings) and the duration gap. They use weekly data from December 2002 to October 2008, thus from the pre-crisis period. Using a market microstructure approach D'Amico and King (2013) use cross-sectional (stock effect) and panel data (flow effect) regressions to examine the effect of purchases of individual Treasury bonds on their returns in the period 18 March to 30 October 2009. They find negative stock and flow effects on Treasury yields, but they also note that the flow effect for notes is fully reversed within a few days, while the effect for bonds is more persistent. Nonetheless, this might be the reason for the difference between the results reported by them and the results in this paper.

A popular approach to the assessment of the impact of QE on the interest rates is to use an equilibrium term structure model and examine the change in particular interest rates components around the important announcement days. This method was implemented by Christensen and Rudebusch (2012) and Bauer and Rudebusch (2014), among others. In this paper I use a no-arbitrage decomposition estimated by three different methods: Kim and Wright (2005), Joslin, Singleton and Zhu (2011) and Bauer, Rudebusch and Wu (2012)), and I use it in the time series regression on the balance sheet variables in a similar vein to Gagnon et al. (2011) and Kozicki et al. (2012). Hamilton and Wu (2012) construct the unconditional monetary policy variables from the ratio of public debt with particular maturities to the total outstanding publicly held debt. They estimate a term structure model using weekly data before (January 1990 - July 2007) and during the ZLB period (March 2009 - Augugust 2010). They estimate the dynamics of the term structure variables (level, slope and curvature) with the UMP variables in the spirit of unspanned variables of Joslin et al. (2014), but to draw the inference about the impact of QE, they mix the forecasting regressions (the real world dynamics) estimated on the normal period with the term structure model estimated on the ZLB period. Another interesting approach was proposed by Li and Wei (2013), who estimate the no-arbitrage term structure model that includes debt supply variables: total public debt excluding Federal Reserve holdings and the total supply and average duration of private MBS holdings. However, they also estimate their model on the basis of the normal period, from March 1994 to July 2007.

This paper is also related to the fast growing literature on the relation between interest rates and the quantity of public debt (Bernanke et al., 2004, Hubbard and Engen, 2004, Han, Longstaff and Merrill, 2007, Krishnamurthy and Vissing-Jorgensen, 2012, Lou et al., 2013) as well as to the empirical stream of the flight-to-liquidity literature (Longstaff, 2004, Beber, Brandt and Kavajecz, 2009, Goyenko and Sarkissian, 2010).

3. Effects of Treasury security purchases on Treasury yields

3.1. Unconventional monetary policy variable

The main subject of my interest is the informational content of the stock of the U.S. Treasury securities held by the Federal Reserve. Thus, I define a measure of the UMP Federal Reserve activity as the total value of Treasury debt held by the Federal Reserve scaled by the nominal GDP interpolated to the weekly frequency:

$$FTT \equiv \log\left(\frac{Total\ Treasury\ securities\ held\ by\ Fed}{Nominal\ GDP}\right). \tag{1}$$

The time series of the FTT variable, in levels and first-differences, is plotted in Figure 1.

[Insert Figure 1 near here]

The differenced FTT series is the only constructed variable in my analysis.[4] Its descriptive statistics are presented in Table 1, where, based on the analysis from Section 3.2, they are reported separately for two periods: 31 December 2003 to 3 October 2007 (Normal period) and 7 April 2010 to 16 December 2015 (ZLB period). The mean of the changes in FTT is close to zero in the normal period and positive, 0.003, in the ZLB period. In the ZLB period the standard deviation of the changes in Treasury holdings increases significantly, from 0.002 to 0.007. The changes in Treasury holdings display small negative skewness in the normal period and positive skewness in the ZLB period. The variable exhibits excess kurtosis in both periods. The changes in FTT are only weakly persistent in the normal period, but become moderately persistent in the ZLB period with the first order autocorrelation equal to 0.73.

[Insert Table 1 near here]

3.2. Sup Wald test

In the predictive regression analysis I focus on the Federal Reserve's Treasury purchases while controlling for a spectrum of macro-finance variables that are likely (on theoretical or empirical grounds) to affect bond returns. In particular, I consider the predictive regressions of weekly changes of zero coupon yields, $\Delta y_{n,t+1}$:

$$\begin{aligned}\Delta y_{n,t+1} \quad = \quad & \alpha_n + \beta_{1,n} level_t + \beta_{2,n} slope_t + \beta_{3,n} VXTYN_t + \beta_{4,n} BAspr_t \\ & + \beta_{5,n} InitCl_t + \beta_{6,n} ExpInf_t + \beta_{7,n} \Delta PubDebt_t + \beta_{8,n} \Delta FTT_t \\ & + \sum\nolimits_j \delta_{j,n} \mathbf{1}\left(t+1 = T_j^{FOMC}\right) + u_{n,t+1}, \qquad (2)\end{aligned}$$

where level is the cross-sectional average of yields with maturities $n = \{1, 3, 5, 7, 10, 15\}$ years and slope is the difference between the 15−year yield and the 1−year yield. Intuitively, including the level of yields on the right hand side of the regression stands for capturing the mean-reversion of interest rates. If interest rates are mean-reverting, I should expect

[4] In the QE literature it is customary to control for the average duration of the debt securities held by the public (e.g. Bernanke et al., 2004, Gagnon et al., 2011, Swanson, 2011, Hamilton and Wu, 2012, Li and Wei, 2013, Greenwood and Vayanos, 2014). In the earlier version of this paper I included the dollar weighted average maturity of the Treasury securities held by the Federal Reserve. Since I did not find it significant in any regression, I decided to drop it to simplify the discussion. The results remain virtually unchanged.

a negative coefficient here. Alternatively I could include a lagged yield in each regression, which leads to almost the same results. The slope is commonly believed to be related to the time varying term premium (e.g. Fama and Bliss, 1987, Campbell and Shiller, 1991). The modern term structure literature typically assumes that the cross-section of interest rates is a linear combination of three factors (the level, slope and curvature, see Litterman and Scheinkman, 1991). Adopting this approach, the dynamics of these factors determine (at least partially) changes in yields. I do not include the curvature factor since it does not exhibit any power in the regressions. The time series of the 3, 5, 10 and 15−year yields are plotted in Figure 2. As can be seen in the figure, at the end of each phase of QE the Treasury yields were higher than they were at the beginning.

[**Insert Figure 2 near here**]

In (2) $VXTYN$ is a logarithm of the 10−year Treasury note volatility index; *BAspr* is a logarithm of the liquidity index spread measuring changes in the bid and ask prices on 3−Month U.S. Treasuries; *InitCl* is the deviation of the log of initial claims reported by the U.S. Employment and Training Administration from its 3−month moving average; and *ExpInf* is the in-5-for-5-year forward inflation expectation rate obtained from the nominal and inflation adjusted Treasury securities.[5] The volatility and liquidity indexes are meant to control for a fast-moving, forward-looking financial environment, while initial claims, expected inflation and public debt control for the broad macroeconomic conditions.

Furthermore, $PubDebt$ is a logarithm of the outstanding debt held by public, i.e. the total U.S. national debt excluding intragovernmental holdings. Figure 3 shows the Federal Reserve holdings of Treasury securities (×10) plotted together with the public debt.[6] Comparing it to Figure 1, it is evident that the FTT variable is a close image of the Federal Reserve holdings. It can be seen that although the public debt was growing steadily throughout the whole period, it was increasing faster in the period following the financial crisis. In particular, in the normal period (December 2003 to October 2007) the public debt increased from 4 to 5.1 trillion dollars or by 6.1% on annual basis, while in the ZLB period (April 2010 to December 2015) it increased from 8.3 to 13.6 trillion dollars or by 13.9% annually. In

[5] In my sample correlation of $ExpInf$ with level and slope is mild, amounting to 0.12 and 0.29 in the normal period, and 0.06 and 0.43 in the ZLB period, respectively. This suggests that expected inflation is an unspanned factor in interest rates as in Joslin et al. (2014). I estimate the term structure model of interest rates with expected inflation as one of the unspanned factors in Section 7.

[6] The daily data on the public debt is available at `https://treasurydirect.gov/govt/reports/pd/pd.htm` since 31 March 2005. Before this date, I interpolated the public debt by scaling total debt, which is available daily, by the average of the ratio of public debt to total debt at the beginning and the end of the month (generally close to 58%).

the ZLB period, the the Treasury holdings by the Fed also grew rapidly, from 0.78 to 2.46 trillion dollars (an increase by 36% on annual basis). Although the correlation of both series in levels is self-evident, the correlation of first-differenced series, as used in (2), is close to zero and amounts to -0.004 and 0.032 in the normal and the ZLB period, respectively.

[**Insert Figure 3 near here**]

I include dummy variables for all weeks with FOMC meetings, which is otherwise equivalent to removing these observations. To match the reporting convention of the Federal Reserve holdings, all variables are constructed on a Wednesday-to-Wednesday basis, except initial claims that are reported weekly on a Saturday-to-Saturday basis. The log transformation of the variables is commonly used to reduce heteroskedasticity. All the macro-finance variables, except public debt, are available from the website of the Federal Reserve Bank of St. Louis.

Figure 4 shows the Sup Wald test for parameter instability with an unknown change point, proposed by Andrews (1993), for the predictive regression (2) for the 3, 5, 10, and 15− year yields. The yields come from the dataset constructed by Gürkaynak, Sack and Wright (2007) and are available from the Federal Reserve Board website.[7] The horizontal lines denote the 10%, 5% and 1% significance levels. The full data sample spans the period from 31 December 2003 to 16 December 2015. The end of the sample is chosen to mark the lift-off of the federal funds rate from the zero lower bound.

[**Insert Figure 4 near here**]

The full-sample test results indicate the existence of three significant breaks. The first break, estimated at 10 October 2007, can be attributed to the collapse of three investment funds belonging to BNP Paribas and two hedge funds belonging to Bear Sterns. This break is particularly significant for the 3−year yield (Panel (a)), but is virtually non-existent in the test of the 10 and 15−year yields (Panels (c) and (d)). The second break is estimated at 15 October 2008 and it is relevant for all yields. It occurs in the time of the biggest market turmoil following the collapse of Lehman Brothers in September 2008, but also amid the Icelandic crisis and multiple banks bailouts. The last break, estimated at 31 March 2010, is particularly visible in the tests for the regression of the 10 and 15−year yields and it coincides exactly with the end of the QE1.

The results of the test provide the motivation to exclude the 'crisis' period between the first and third break from the regressions (the shaded area in Figure 4), and to focus on two

[7] `http://www.federalreserve.gov/pubs/feds/2006/200628/200628abs.html`.

subperiods: 31 December 2003 to 3 October 2007 and 7 April 2010 to 16 December 2015, which I dub a 'normal' and a 'ZLB' period, respectively. The sup Wald test calculated for each subsample shows that, except for the 3−year yield, the null hypothesis of the stability of the parameters is not rejected at the 5% level.

3.3. Predictive regression results

Table 2 reports the results of the predictive regressions (2) of changes in yields for 3, 5, 10 and 15−year yields. The numbers reported in round brackets are Newey-West standard errors calculated with 6 lags.

[Insert Table 2 near here]

In the normal period there is a notable lack of any predictive power in any of the considered variables, with the exception of the bis-ask spread, which is significant at the 10% level for the 3−year and 5−year yields. In the last two rows of the table I report the number of FOMC meeting-week dummies that are significant at the 5% level. In the normal period there were 30 FOMC meetings. In this period the number of significant dummies ranges from 16 in the regression of the 15−year yield to 19 for the 3−year yield.

The results for the ZLB period are very different from those for the normal period. In particular, the coefficient on the level factor is significant and negative, which indicates mean-reversion tendency in this sub-sample. The bid-ask spread is statistically significant for short and medium yields with a negative coefficient, which is consistent with the liquidity premium in Treasury bonds (Fontaine and Garcia, 2012). The expected inflation has significant predictive power for all considered yields. As should be expected, the effect is stronger for long maturities, since it measures break-even inflation in the in-5-for-5 year interest rate futures. The association of expected inflation with bond returns has been well-established in the asset pricing literature (see e.g. Fama and Schwert, 1977, Ang, Bekaert and Wei, 2008, Chernov and Mueller, 2012, Goliński and Zaffaroni, 2016).[8] The coefficients on public debt are positive and statistically significant at least at the 5% level. These results indicate that a rise in public debt by 1% predicts an increase in Treasury yields by about 3 basis point.

The most remarkable predictive effect in the ZLB period is obtained by changes in the total Treasury securities holdings by the Federal Reserve. The estimated coefficients are positive and their t−statistics are between 3.5 and 4.3 standard deviations from zero. This effect is stronger for longer maturities. Increasing the total holdings of the Federal Reserve by 1% (in a ratio to the GDP) predicts increases of medium and long maturity yields by

[8] In similar regressions with contemporaneous variables I confirm that expected inflation has a strong positive association with Treasury yields (results not reported, available upon request).

about 4 basis points, which is economically meaningful, especially in a low interest rate environment.

Finally, there are 45 FOMC weeks in the ZLB period. A great majority of the dummy variables for those weeks are significant at the 5% level, ranging from 34 for the 3−year yield to 38 for the 15−year yield.

The predictive power of the Federal Reserve purchases of Treasury securities in the ZLB period can be seen from the marginal increase in the adjusted R−squared coefficient reported in Table 3.[9] Without ΔFTT variable the adjusted R−squared ranges between 3% for the 15−year yield to 3.7% for the 5−year yield. Although the fit of the predictive regression with only one variable, ΔFTT, and a constant is rather modest with the adjusted R−squared generally below 1%, the addition of this variable to other control variables greatly improves the goodness of fit as evidenced by the adjusted R−squared statistic ranging from 6.9% for the 3−year yield to 8.9% for the 5−year yield. These results contrast with the fact that in the normal period the adjusted R−squared coefficient decreases when adding the Federal Reserve security purchases to the predictive regression, which is typical for statistically insignificant variables.

[Insert Table 3 near here]

4. Transimission channel of UMP

4.1. Term premium

To understand the transmission channel of the monetary policy, it is important to learn whether the changes in the total Treasury holdings of the Fed affect interest rates through expectation of future interest rates or through the risk premium:

$$y_{n,t} = y^{\mathcal{Q}}_{n,t} + TermPr_{n,t}, \tag{3}$$

where $y^{\mathcal{Q}}_{n,t}$ is the risk-neutral yield defined as the average expected path of the short rate over the maturity of the bond and $TermPr_{n,t}$ stands for the yield term premium.

To this end, I adopt a methodology similar to that in Gagnon et al. (2011) and examine the effect of Federal Reserve security purchases on the term premium and risk neutral yields in a time series regression. These quantities are generally unobserved and, as such, they need to be estimated from a structural no-arbitrage model (see e.g. Hamilton and Wu, 2012,

[9] Since the FOMC meeting week dummies set the residuals to zero, the statistic is calculated based on the residuals corresponding to non-meeting weeks.

Christensen and Rudebusch, 2012). Since in the literature there is no consensus on what the estimates of term premium should look like (see e.g. the discussion in Bauer, Rudebusch and Wu, 2014 and Wright, 2014), I consider three different term premium estimates. The first term premium comes from the Kim and Wright (2005) model, which is based on the algorithm proposed by Kim and Orphanides (2012) and incorporates information from interest rate surveys. The estimates of the term premium reported on a daily basis are available from the website of the Federal Reserve Board.[10] The second term premium estimates are based on the method proposed by Joslin et al. (2011). They assume that under the assumption that some combinations of yields are observed without error, the forecast of the observable state vector can be found by estimating the $VAR(1)$ system by the OLS. The term premium is then found by subtracting the mean expected short rate from the fitted yield (subject to a constant convexity adjustment). This approach was used by Wright (2011) to evaluate term premia for a number of countries. Following Joslin et al. (2011) I assume that the term structure of interest rates is spanned by three factors and the first three principal components are observable without error. Finally, Bauer et al. (2012) propose a simulation-based bias correction to the estimates of the VAR system. This is the third method that I use to find the estimates of the term premium.[11] The term premia are found by estimating the model on the basis of the whole sample.

Hence, I regress the term premium for the $10-$year bond on the Federal Reserve Treasury purchases, controlling for level, slope and dummy variables for announcement weeks in a predictive manner:

$$\Delta TermPr_{10,t+1} = \alpha + \beta_1 level_t + \beta_2 slope_t + \beta_3 \Delta FTT_t + \sum\nolimits_j \delta_j \mathbf{1}\left(t+1 = T_j^{FOMC}\right) + u_{t+1}. \quad (4)$$

In a complementary regression, I regress the risk neutral yield:

$$y_{10,t}^{Q} \equiv \frac{1}{10 \times 52} \sum_{j=1}^{10\times 52} E_t\left[y_{1/52,t+j-1}\right] \quad (5)$$

on the same set of independent variables. The results of both sets of regressions are presented in Table 4. The Newey-West standard errors with 6 lags are reported in parentheses.

[Insert Table 4 near here]

The results are generally dependent on the model used to decompose the $10-$year yield,

[10] `https://www.federalreserve.gov/econresdata/researchdata/feds200533.html`.

[11] I use part of the code available on the Cynthia Wu's website: `http://faculty.chicagobooth.edu/jing.wu/`, for which I am grateful.

but nonetheless we can make a few systematic observations. First of all, the term premium obtained by the bias correction method proposed by Bauer et al. (2012) exhibits little predictability in any of the periods considered. As for the other two decomposition methods, in the normal period the level can predict the risk-neutral yields (and the term premium for the Kim and Wright (2005) model) and slope can predict term premium, but there is no evidence of the predictability of either the term premium or the risk-neutral yield by the Federal Reserve's purchases. On the other hand, in the ZLB period, the monetary policy variable (FTT) is highly significant in the regression of the term premium obtained from the Kim and Wright (2005) model and Joslin et al. (2011) model. The results for the Bauer et al. (2012) method are not significant, but we can note that the point estimates have the same sign and the order of magnitude as for the other models, but the lack of significance is due to larger standard errors. The Federal Reserve purchases do not seem to have any predictive power for the changes in the risk-neutral yield, with the notable exception of that obtained by the Kim and Wright model.

4.2. Excess returns

To further corroborate the relation between the Federal Reserve's Treasury purchases and term premium, I turn to the economic meaning of the term premium, which is the expected excess return on a long maturity bond over the risk-free rate. The results of the term premium regressions suggest that in the ZLB period increasing the Treasury holdings by the Federal Reserve increases the expected excess return on long maturity bonds.[12] To evaluate this hypothesis I construct a monthly series of annual excess returns on the zero-coupon bonds:

$$\begin{aligned} xr_{n,t+12} =& p_{n-1,t+12} - p_{n,t} + p_{1/12,t} \\ =& ny_{n,t} - (n-1)y_{n-1,t+12} - y_{1/12,t}. \end{aligned} \tag{6}$$

I then regress the excess returns on a constant and monthly changes in the total Treasury holdings by the Federal Reserve:

$$xr_{n,t+12} = \alpha + \beta \Delta FTT_t + u_{n,t+12}. \tag{7}$$

[12]Strictly speaking, since the regressions of the term premium (4) are run in a predictive manner, the increase in the term premium is accompanied by a contemporaneous decline in realized returns *ceteris paribus*, followed by higher expected future returns. Thus, the multi-period returns predictability likely reflects the serial correlation of the predictor, ΔFTT, reported in Table 1.

The two subperiods match the samples selected for the regressions with the weekly data, i.e. the normal period is from December 2003 to October 2007, and the ZLB period is from May 2010 to December 2015.

The results are reported in Table 5. The Newey-West standard errors with 12 lags are reported in parentheses. Note that since for different maturities the volatility of yields is of the same order of magnitude, the volatility of excess returns increases almost linearly with maturity and so do the regression coefficients. In the normal period ΔFTT does not have power to forecast excess bond returns on the 3−year and 5−year bonds but has a positive and statistically significant coefficient at the 5% level for 10 and 15−year bond returns. The adjusted R−squared is low and does not exceed 2% in any regression. In the ZLB period, however, ΔFTT is significant at the 1% level for each bond with the adjusted R−squared coefficient ranging from 39% for the 10−year bond to 45% for the 3−year bond.

[Insert Table 5 near here]

Furthermore, I check whether the predictive content of the information about the Federal Reserve's holdings is already included in the current term structure of interest rates. Cochrane and Piazzesi (2005) argue that excess returns on different bonds are driven by one return forecasting factor, which can be constructed from the cross-section of current forward rates. Thus, including the Treasury purchases in the return forecasting regressions together with the return forecasting factor can be interpreted as testing whether ΔFTT is an unspanned factor in interest rates.

I construct the return forecasting factor adopting the Cochrane and Piazzesi method to my maturities. The Cochrane-Piazzesi-like factor (CP) is the fitted part from the regression

$$\overline{xr}_{t+12} = \gamma_0 + \gamma' \mathbf{f}_t + \varepsilon_{t+12}, \tag{8}$$

where $\overline{xr}_{t+12}$ is the mean excess return, $\mathbf{f}_t$ is a vector with 1−year forward rates with maturities 2, 4, 6, 9 and 14 years, estimated over the whole sample. As shown in Table 5, the predictive power of CP is substantial in both subperiods. In the normal period the predictability is strong for bond excess returns with maturity of 5 years and longer, with the adjusted R−squared reaching 50% for the 10−year bond. The forecasting power of the CP factor is even stronger in the ZLB period, with the adjusted R−squared ranging from 52% for the 3−year bond to 71% for the 15−year bond.

In a joint regression with the Treasury purchases variable I find that in the normal period CP renders the Treasury purchases insignificant in all regressions. In the ZLB period ΔFTT has forecasting power beyond that already incorporated in forward rates; in all excess return regressions the monetary policy variable is always positive and statistically significant, at least at the 5% level. The adjusted $R-$squared increases by about $4\% - 10\%$ above those in simple regressions with CP variable. In general, the excess bond returns monthly regression results are consistent with the weekly term premium regression results.

The results from Table 5 are remarkable and difficult to dismiss. First of all, ΔFTT is an exogenous variable that includes only the information on the Federal Reserve's holdings, but is not related in any way (at least directly) to economic or financial indicators. Second, by design it does not have the look-ahead bias, unlike the CP factor, which results from the way the factor is constructed. Finally, it is moderately persistent - the first-order autocorrelation coefficient in the ZLB period is 0.73 and is mainly due to systematic purchases of Treasury securities during QE2. Nonetheless, as opposed to many well-known return predictors, the persistence is not large enough to raise the suspicion of spurious predictability.[13]

5. Robustness checks

Admittedly, the vital question is whether the relation between the Federal Reserve's Treasury purchases and the bond yields is not spurious. Based on the presented evidence I argue that a spurious correlation between the Federal Reserve's purchases and Treasury yields is unlikely. To validate the results presented in the previous sections I performed multiple robustness checks. In all of these exercises I consistently found that increasing the total Treasury holdings by the Federal Reserve was associated with a rise in the yields in the ZLB period. I briefly describe them here. They are available upon request.

First of all, I tried different accounting of the Fed balance sheet variable. I found that including the Agency debt does not change the results in any significant way - the estimated FTT coefficients are actually slightly larger and the $t-$values exceed 3.5. On the other hand, when we include MBS in the FTT variable, the coefficients in the 3 and $5-$year yield regressions are still significant at the 1% level, but for the 10 and $15-$year yields they are significant only at the 5% and 10% level, respectively. I attribute these results to two factors: 1) MBS have very different risk characteristics, since they contain a significant prepayment risk; 2) the scale of MBS purchases was comparable to that of Treasury securities, while Agency debt constituted a small fraction of total Federal Reserve purchases. Also, as a scaling factor in the FTT variable I use nominal GDP, following the Greenwood and Vayanos (2014) and Krishnamurthy and Vissing-Jorgensen (2012). I tried alternatively to scale it by the total national debt (i.e. including intragovernmental holdings) or by the outstanding

[13] Using simulation method, Ferson, Sarkissian and Simin, 2003 establish that spurious regression bias does not arise to any serious degree provided the autocorrelation of predictors is 0.9 or less.

debt held by the public (public debt) - the results were almost the same.

Second, the advantage of the weekly data is that it provides us with a relatively high number of observations, which increases the power of statistical inference, but brings forth the problem of price reactions to market expectations. I repeated the regressions using the monthly data and excluding the FOMC meeting dummies. Despite the obvious drawback of a smaller number of observations when working with lower frequency data, the monthly regressions give very similar results, both in terms of a sign and magnitude of the coefficients, and their statistical significance.

Third, it could be hypothesized that some FOMC meetings were more important than other and we should specifically control only for these. For instance, Christensen and Rudebusch (2012) and Krishnamurthy and Vissing-Jorgensen (2011) select 'key' announcement dates and examine yield reaction around these times. I checked that the choice of the FOMC meeting dummies, as well as skipping all the dummies altogether, had a negligible effect upon the regression results. Furthermore, I checked the hypothesis of a 'delayed' reaction to FOMC announcement by including lagged dummies. Although some of them were statistically significant, they had little effect on other regression coefficients. Also, I tried regressions with a permanent effect of the FOMC announcements:

$$\Delta y_{n,t+1} = \alpha_n + \boldsymbol{\beta}_n' \mathbf{x}_t + \sum\nolimits_j \delta_{j,n} \mathbf{1}\left(t+1 \geq T_j^{FOMC}\right) + u_{n,t+1}, \tag{9}$$

but again it did not change the main results; if anything, the results were even stronger than with single point dummy variables.

One might be also concerned whether the results could be driven by some particular event or episode during the ZLB period. For example, one such episode could be a rapid increase in the total security holdings by the Fed during QE2, which was accompanied by the increase in the medium and long maturity rates at the end of 2010 and the first half of 2011. The rise in the government bond yields at that time was caused by the increase in interest rates in Europe and dissipating market concerns in the U.S. resulting in declining safe heaven premium (OECD, 2011). To check this possibility I run the baseline regressions in the ZLB period starting after QE2 (6 July 2011 to 16 December 2015). The general pattern of results is very similar to those reported in Table 2. The point estimates of the FTT coefficients are actually slightly larger, but the standard errors are larger as well (the coefficient for the 15$-$year yield is not statistically significant and for the 10$-$year yield is significant at the 5 percent level, but not at the 1 percent). When we run the regressions, however, starting from January 2012, all coefficients are significant at the 1 percent level.

More generally, one could be concerned that the results are influenced by some outliers or extreme observations. In Figure 5 I plot the fitted vs. observed changes in yields during the ZLB period. If the model gives a perfect fit, the points would lie along the 45−degree line. On the other hand, if the model does not have any explanatory power, then the observations would be scattered vertically around 0. Consistent with the moderate values of the adjusted $R-$squared coefficients reported in Table 3, the scatterplot shows a mild correlation between the fitted and observed values. More importantly, however, we can see that the correlation is not driven by any outliers: the overall picture does not changes if we eliminate any particular observation(s) from the sample.

[**Insert Figure 5 near here**]

Finally, I also tried other control variables, but have not found any meaningful change in the results. My volatility and liquidity proxies are weekly averages. Using other proxies as well as using the end-of-period values does not make any difference. I tried different specifications of the control variables, e.g. changes instead of levels or simple levels instead of logs, but generally I found the specification used in Section 3.3 to be most statistically significant, while other variables have little effect on the ΔFTT estimates.

6. Effects of UMP on expected inflation and real interest rate

In this section I consider an alternative decomposition of a nominal interest rate, into a real interest rate and expected inflation. As with the term premium, generally such decomposition would require specifying and estimating a no-arbitrage model, such as in Campbell, Sunderam and Viceira (2009) or Chernov and Mueller (2012), which would also enable estimation of the inflation risk premium component. I take a simpler route and analyse the relation between observable quantities. As the expected inflation I use the 5−year forward inflation rate, as in Section 3.3. As a real rate I use the 10−year real yield estimated from the index-linked bonds by Gürkaynak, Sack and Wright (2010).[14]

For both subperiods I run regressions on weekly data:

$$\Delta x_{t+1} = \alpha_n + \beta_1 x_t + \beta_2 \Delta FTT_t + \sum\nolimits_j \delta_{j,n} \mathbf{1}\left(t+1 = T_j^{FOMC}\right) + u_{t+1}, \qquad (10)$$

where x_t is either the expected inflation or the real yield. The results are presented in Table 6.

[**Insert Table 6 near here**]

[14]Available at: `https://www.federalreserve.gov/pubs/feds/2008/200805/200805abs.html`.

Since both the expected inflation and the 10−year real rate are persistent variables, the estimated coefficient on the lagged term is close to zero for both variables in both periods. The Federal Reserve Treasury purchases do not seem to predict any of these variables in the normal period. However, the results for the ZLB periods suggest a positive relation between Treasury purchases by the Federal Reserve and both the expected inflation and the real rate. The significance is particularly strong, at the 1% level, in the expected inflation regression, while in the real rate regression it is significant at the 10% level and close to the 5% significance level.

To summarise, in the ZLB period we find a strong and positive association between Treasury security purchases and expected inflation and a marginally positive relation between Treasury purchases and the real interest rate. On the other hand, we do not find any statistically significant effect between these variables in the normal period.

7. Counterfactual analysis: effects of UMP on yields

7.1. Term structure model

In this section I analyse the cumulative effect of the Treasury securities purchases by the Federal Reserve on the term structure of Treasury yields. To this end, I estimate the no-arbitrage term structure model with unspanned variables proposed by Joslin et al. (2014) (hence JPS). The model is specified in terms of risk-neutral and physical dynamics. The principal components of yields are the only factors that span the term structure of interest rates contemporaneously, while other (unspanned) variables can affect the future realizations of the term structure factors in a dynamic relation.

A convenient feature of the JPS model is that it allows us to estimate the physical dynamics of the system by OLS regression, which greatly reduces the computational problem. Consistent with the structural break analysis in the previous sections, the model is estimated over the ZLB period, 7 April 2010 to 16 December 2015, using the weekly data on yields with maturities of 1, 3, 5, 7, 10 and 15 years. Since at this time the short end of the yield curve was effectively fixed at the lower bound, the variation in the level factor was restricted. Indeed, Figure 6 shows the loadings on the first two principal components estimated on the ZLB period. Clearly, the usual level factor disappears in this period and the first two principal components can be identified as the slope and curvature factors. Hence, I estimate the model with two yield curve factors spanning all interest rates.[15]

[15] Although the principal components have different interpretation from the level and slope factors that I used in previous sections, using these two sets of factors is consistent with each other since the model is invariant to the factor rotation (see Joslin et al., 2011).

[Insert Figure 6 near here]

Following JPS as unspanned macro factors, I include variables that represent real activity (initial claims) and expected inflation, both as defined in Section 3.2, but I also add the Treasury purchases by the Federal Reserve. Thus, denote by $\mathbf{x}_t$ a 5×1 vector containing the first two principal components of yields, initial claims, expected inflation and ΔFTT, i.e. $\mathbf{x}_t = [PC_{1,t}, PC_{2,t}, InitCl_t, ExpInf_t, \Delta FTT_t]'$. To stay consistent with the methodology in the previous sections, I also include the dummy variables for FOMC meeting weeks. Keeping the dummies in the model allows us to avoid the mechanical effect of expanding the Federal Reserve's balance sheet on yields in the counterfactual analysis. The correlations between the variables are showed in Table 7. As before, these calculations are based on the non-meeting weeks. In consequence, the correlation between the principal components is not zero but slightly negative, -0.001. Initial claims do not exhibit strong correlation with principal components and expected inflation, but expected inflation is strongly correlated with the second principal component (0.65). The changes in the total Fed's holdings, on the other hand, are strongly positively correlated with both principal components (0.50 and 0.32, respectively) and expected inflation (0.36).

[Insert Table 7 near here]

The physical dynamics of the state vector are then given by the augmented $VAR(1)$ process:

$$\mathbf{x}_t = \boldsymbol{\mu} + \boldsymbol{\Phi}'\mathbf{x}_{t-1} + \sum\nolimits_j \boldsymbol{\delta}_j \mathbf{1}\left(t = T_j^{FOMC}\right) + \mathbf{u}_t, \tag{11}$$

where $\mathbf{u}_t \sim N(\mathbf{0}, \boldsymbol{\Sigma})$. The estimates of $\boldsymbol{\Phi}$ are reported in Table 8. The Newey-West standard errors with 6 lags are reported in parentheses.

The results indicate that Treasury purchases by the Federal Reserve are an important determinant of the principal components and expected inflation. Although the initial claims variable do not seem to be an important driver of the term structure factors, it is weakly statistically significant in the expected inflation equation, which, in turn, strongly Granger causes the first principal component of yields. On the other hand, ΔFTT seems to be weakly Granger caused by the principal components of yields.

[Insert Table 8 near here]

I also estimated the system without the dummy variables for the FOMC meeeting weeks (unreported, available upon request). The results differ very little from the results presented in Table 8, either in terms of the size of the coefficients or their statistical significance. This is because although most of the FOMC meetings had a significant impact on the interest rates, their sign was not systematic, which means that to a large extent they were cancelling each other.

7.2. Effect on yields

Keeping the original parameters of the model and the estimated residuals, I plug in the new series to obtain the hypothesised term structure quantities. I compute two versions of the counterfactual time series. To find the first counterfactual series ('Counterfact., with FOMC'), I re-evaluate the system by assuming that in the ZLB period the holdings of Treasury securities by the Federal Reserve remain unchanged, i.e. $\Delta FTT = 0$. To this end, given the estimates of (11) I find a counterfactual time series of $\mathbf{x}_t$ by re-evaluating the system with the last element of the estimated residual vector $\mathbf{u}$ and the last element of the $\boldsymbol{\delta}$'s in (11) set to zero. I also set to zero the last element of $\boldsymbol{\mu}$. To eliminate the effect of other variables on ΔFTT I also replace the last column of $\mathbf{\Phi}$ by a vector of zeros. This counterfactual series corrects only for the effect of the actual changes in the Federal Reserve's balance sheet, but allows for the effect of the FOMC meetings on other variables. In the second version ('Counterfact., no FOMC') I eliminate the effect of shocks of the FOMC meetings on the interest rates and expected inflation, but not on initial claims. This is achieved by setting the coefficients on the dummy variables to $\boldsymbol{\delta}_j = [0, 0, \delta_{3,j}, 0, 0]'$.[16] The distance between the actual yield and the counterfactual series with the FOMC shocks accounts purely for the change in the Federal Reserve's balance sheet, while the difference between the two counterfactual series measures the effect of the FOMC meetings.

In Figure 7(a) I plot the actual and the two counterfactual series of the $10-$year yield, while the figures for other maturity yields display a very similar pattern. As could be expected from predictive regressions, since the biggest increase in the Federal Reserve's holdings took place during QE2, removing the effect of the Treasury purchases results in considerably lower interest rates during this phase (by about 70 basis points), regardless whether we account for the FOMC meeting effects or not. After QE2 the total holdings of the Federal Reserve were slowly subsiding, which results in closing the gap between the actual and the counterfactual yield (with FOMC), although probably most of this re-alignment is due to the FOMC shocks that pushed the actual yields down. From the second quarter of 2012 to the beginning to 2015 both the counterfactual series follow quite closely the actual $10-$year yield. In 2015, however, we can see that the changes in the total holdings of the Federal Reserve did not have a material impact on the level of interest rates, but the FOMC shocks resulted in the reduction of the level of the $10-$year yield by about 30 basis points.

[16] I checked that the setting the FOMC shocks to zero makes virtually no difference to the results.

[Insert Figure 7 near here]

Using the estimated dynamics of (11) for the counterfactual series we find the term premium for the 10−year yield, which I plot in Figure 7(b). The term premium is calculated as a difference between the fitted (actual or counterfactual) yield and the expectation of the average short rate over the next ten years. The counterfactual expectation of the short rate is found by assuming that the Federal Reserve's holdings will not change in the future, which is obtained by setting particular coefficients of the model to zero, as described previously.

The general pattern of the counterfactual term premium is similar to that of the 10−year yield level. The model suggests that the rapid expansion of the Federal Reserve's balance sheet during QE2 increased the term premium by about 80 basis points. This gap was subsequently reduced and closed in the first half of 2012. The purchases of the Treasury securities during QE3 increased the term premium by about 40 basis points and this effect persisted until the end of 2014. From 2015, if we allow for the FOMC shocks, the counterfactual term premium is lower by about 30 basis points, but absent these shocks the counterfactual term premium closely follows the actual term premium.

To summarize, I find that the changes in the Federal Reserve holdings of Treasury securities in the period April 2010 to December 2015 had a non-negligible effect both on interest rates and on the term premium. This effect was particularly pronounced during QE2 in 2011 and, in consequence of QE3, in the second half of 2013 and the first half of 2014.

8. Testing predictions of the liquidity channel

So far I have documented the impact of the actual purchases of Treasury securities on Treasury yields arguing that the main transmission channel is the risk premium. As I outlined in the introduction, the positive effect of QE on Treasury yields, and on the Treasury term premium in particular, can be explained by the liquidity channel as postulated by Krishnamurthy and Vissing-Jorgensen (2011). If Treasury securities purchases by the Federal Reserve relieve the liquidity tension in financial markets, the perception of riskiness of Treasury bonds relative to other asset classes should change. Put another way, the risk premium for holding Treasury bonds as a 'safe haven' asset class declines in comparison to other assets. Although the liquidity channel seems more relevant in the early stages of the financial crisis, in this section I present the evidence showing that the expansion of the Federal Reserve's balance sheet had a substantial effect on the risk premium in particular asset classes.

8.1. Corporate yield spread

To test this prediction of the liquidity channel hypothesis in a predictive regression I examine the impact of Federal Reserve purchases on corporate yield spread. To this end, I perform regressions similar to that in Section 6 in Eq. (12), where x_t represents one of three options: the spread between Moody's seasoned Baa corporate yield and 10−year Treasury constant maturity, the spread between Aaa yield and 10−year Treasury constant maturity, and the spread between Baa and Aaa corporate yields. The series, plotted in Figure 8, are reported on a daily basis and available on the website of the Federal Reserve Bank of St. Louis. As before, I run regressions using the weekly Wednesday-to-Wednesday data:

$$\Delta x_{t+1} = \alpha + \beta_1 x_t + \beta_2 \Delta FTT_t + \sum\nolimits_j \delta_j \mathbf{1}\left(t+1 = T_j^{FOMC}\right) + u_{t+1}, \tag{12}$$

where x_t represents different credit spreads.

[**Insert Figure 8 near here**]

The results are presented in Table 9. The Newey-West standard errors with 6 lags are reported in parentheses. In the normal period none of the spreads seems predictable either by the Treasury purchases or by itself. In the ZLB period, however, ΔFTT is negative and statistically significant at the 5% level in the regression of the Baa-Treasury spread and the Baa-Aaa spread. As can be seen in Figure 8, these spreads were systematically falling in each phase of QE. The coefficient on the monetary policy variable in the Aaa-Treasury spread regression is negative but not statistically significant.

[**Insert Table 9 near here**]

8.2. Stock market returns

Using similar logic, we can hypothesise that stock returns, as the more risky asset class than Treasury securities, are positively related to the Treasury purchases by the Federal Reserve as the liquidity premium narrows and equity becomes more expensive. For instance, by eyeballing Figure 9, we can see that the stock market index $S\&P$ 500 was systematically growing in each phase of QE.

[**Insert Figure 9 near here**]

In Table 10 I report the results of the regression of weekly log returns on the $S\&P$ 500 index on the Treasury purchases variable with one and two lags and lagged level and slope of the interest rates as the control variables. Thus, the regression is:

$$\begin{aligned} r_{t+1} &= \alpha + \beta_1 level_t + \beta_2 slope_t + \beta_3 \Delta FTT_t + \beta_4 \Delta FTT_{t-1} \\ &\quad + \sum\nolimits_j \delta_j \mathbf{1}\left(t+1 = T_j^{FOMC}\right) + u_{t+1}, \end{aligned} \tag{13}$$

where r_{t+1} is the log return on the $S\&P$ 500 index. As previously, I add the dummy variable for the announcement weeks.

[Insert Table 10 near here]

Consistent with the safety channel hypothesis, in the ZLB period the Treasury security purchases by the Federal Reserve, ΔFTT, appear to be a strong predictor of stock returns with a positive coefficient. However, in comparison to the corporate bond market, this liquidity effect seems to permeate slower to the equity market, since it is significant only in the second lag. The 1% increase in Treasury holding to GDP ratio predicts increases in the stock returns by about 38 basis points on a weekly basis (based on the cumulative sum of the two lags). In the normal period the coefficient estimates are larger in magnitude but only marginally significant at the 10% level.

To summarize, the corporate spread and equity return regressions provide support for the liquidity channel hypothesis of the QE. The yield on less liquid asset classes declines relative to safer and more liquid securities. This is evident both in the effect of Treasury security purchases on different asset classes, but also within a relatively homogeneous asset class, like corporate bonds.

9. Conclusions

In this paper I analyse the effect of Treasury security purchases by the Federal Reserve. First of all, I document a significant structural change in the relation between the Treasury holdings of the Federal Reserve and Treasury yields. In the ZLB period from May 2010 to December 2015, Treasury purchases by the Federal Reserve are associated with a rise in Treasury yields. This effect is significant after controlling for the FOMC meeting dates as well as a number of macroeconomic and financial variables. I argue that the positive relation between the Federal Reserve's purchases and the bond risk premium in the ZLB period is the main driver of this result, but a similar relation also exists with expected inflation. All of these effects are statistically significant.

As noted by Wu and Zhang (2016), in the ZLB period the Federal Reserve's holdings increase and long-term rates decline. It should be noted, however, that by running a multiple regression I look at the conditional rather than unconditional correlation, which allows controlling for other factors. Insofar as the launch of QE Treasury purchases were known in advance, my findings are puzzling from the perspective of the efficient market hypothesis. However, the evidence is consistent with the liquidity premium in Treasury bonds following the bankruptcy of Lehman Brothers in September 2008 and the subsequent flight-to-quality in financial markets. This transmission channel was postulated by Krishnamurthy and Vissing-Jorgensen (2011). Accordingly, I find support for the liquidity channel hypothesis by finding the significant relation between Treasury purchases by the Federal Reserve and corporate bond spread between bonds with different safety classes. My counterfactual analysis suggests that there were periods when long-maturity Treasury yields would have been lower without the Federal Reserve's interventions, but at the end of the ZLB period the counterfactual yield would have been close to the actual rates.

The collected evidence allows us to conjecture that the effect of the UMP in the ZLB period has been multidimensional. The presented results should not be interpreted in a prescriptive manner for monetary policy; it is inappropriate to conclude that the Federal Reserve can lower interest rates by selling rather than buying Treasury securities. First of all, as documented in Krishnamurthy and Vissing-Jorgensen (2011) and Christensen and Rudebusch (2012), in the first phase of the crisis the QE announcements had a strong negative impact on the interest rates, which could have been critical for the stability of financial markets and the macroeconomy. Thus, it is possible that the following actual Treasury purchases were just gradually reversing this effect.

Moreover, despite the fact that the effect of actual purchases runs counter to the narrow objective of the QE of lowering the long-term rates, its economic significance should be considered in a broader context. For instance, I have found evidence that the UMP variables tend to be positively related to the expected inflation, which was also one of the QE objectives (see Joyce et al., 2011). This is an important outcome of the QE as the menace of deflation was looming at the outset of the financial crisis. Second, the Federal Reserve purchases are negatively correlated with the corporate yield spread, which must have been favourable for the firms by lowering the real cost of capital.

Finally, it could be conjectured that the Treasury securities purchases by the Federal Reserve had a strong positive signalling effect in the ZLB period, as indicated by higher real rates, higher term premium on nominal bonds and higher inflation expectations. Put another way, relieving the liquidity pressure by actual purchases of long-maturity bonds corresponds with expectations of 'good times' for the economy. This, in turn, can be expected to decrease aggregate risk aversion and increase optimism about the future state of the economy. A rise in economic optimism is hard to overvalue when the economy is fragile, since it increases the propensity to consume and increases investors' confidence (see e.g. Kozicki et al., 2012). All in all, despite the fact that QE did not lower the Treasury yields significantly, it can be contended that it importantly contributed to the revival of the economy.

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Tables

ΔFTT	Normal period	ZLB period
Obs.	197	297
Mean	-0.0003	0.0031
Std.dev.	0.0015	0.0069
Min.	-0.0072	-0.0153
Max.	0.0047	0.0393
Skewness	-0.1249	2.0846
Ex.kurtosis	4.1122	6.1087
Autocorr.(1)	0.2168	0.7284

Table 1: Descriptive statistics of the logarithm of the first differences of the ratio of the total holdings of Treasury securities by the Federal Reserve to the nominal GDP, ΔFTT. The normal period is 31 December 2003 to 3 October 2007 and the ZLB period is 7 April 2010 to 16 December 2015.

	Normal period				ZLB period			
	3y	5y	10y	15y	3y	5y	10y	15y
level	-0.0667	-0.0514	-0.0305	-0.0168	-0.0583***	-0.0920***	-0.1091***	-0.1006***
	(0.0444)	(0.0465)	(0.0474)	(0.0466)	(0.0200)	(0.0273)	(0.0352)	(0.0365)
slope	-0.0367	-0.0331	-0.0307	-0.0301	0.0127	0.0250	0.0310	0.0257
	(0.0285)	(0.0302)	(0.0282)	(0.0271)	(0.0172)	(0.0234)	(0.0272)	(0.0279)
$\log(VXTYN)$	-0.1053	-0.0952	-0.0372	-0.0061	0.0027	-0.0105	-0.0261	-0.0159
	(0.1079)	(0.1124)	(0.0999)	(0.0883)	(0.0260)	(0.0351)	(0.0402)	(0.0408)
$\log(BAspr)$	0.2595*	0.2444*	0.1934	0.1824	-0.0218***	-0.0213**	-0.0092	-0.0070
	(0.1392)	(0.1433)	(0.1337)	(0.1310)	(0.0084)	(0.0106)	(0.0105)	(0.0105)
$\log(InitCl)$	0.0385	0.0596	0.0536	0.0366	-0.1464	-0.2148	-0.2200	-0.2063
	(0.1281)	(0.1259)	(0.1359)	(0.1457)	(0.1215)	(0.1580)	(0.1713)	(0.1694)
$ExpInf$	0.0657	-0.0006	-0.0641	-0.0887	-0.0453**	-0.0770***	-0.1097***	-0.1024***
	(0.1214)	(0.1230)	(0.1161)	(0.1105)	(0.0218)	(0.0291)	(0.0341)	(0.0343)
$\Delta PubDebt$	-2.4433	-1.9736	-1.0521	-0.2091	2.4908**	3.3000**	3.6449**	3.5600***
	(2.4329)	(2.5354)	(2.4306)	(2.2789)	(1.1959)	(1.4774)	(1.7618)	(1.8807)
ΔFTT	-2.6539	-2.9162	-1.9814	-0.6150	2.4203***	3.9450***	4.8250***	4.3243***
	(5.1683)	(5.2234)	(4.8391)	(4.7595)	(0.6600)	(0.9144)	(1.1716)	(1.2120)
# sign. dummies	19	17	17	16	34	36	37	38
% of all dummies	(63.33%)	(56.67%)	(56.67%)	(53.33%)	(75.56)	(80.00)	(82.22)	(84.44)

Table 2: Regression of changes in Treasury yields. The Newey-West standard errors with 6 lags are reported in parentheses. *, **, and *** denote significance at the 10, 5, and 1 percent level, respectively. The normal period is 31 December 2003 to 3 October 2007 and the ZLB period is 7 April 2010 to 16 December 2015.

	3y	5y	10y	15y
	Normal period			
Control variables only	0.0335	0.0199	0.0029	-0.0035
ΔFTT only	-0.0059	-0.0061	-0.0060	-0.0061
Control variables + ΔFTT	0.0287	0.0152	-0.0026	-0.0098
	ZLB period			
Control variables only	0.0348	0.0369	0.0326	0.0295
ΔFTT only	0.0042	0.0108	0.0081	0.0036
Control variables + ΔFTT	0.0685	0.0885	0.0866	0.0713

Table 3: Adjusted $R-$squared coefficient for predictive regressions for the normal period (top panel, 31 December 2003 to 3 October 2007) and the ZLB period (bottom panel, 7 April 2010 to 16 December 2015).

	Kim and Wright		Joslin, Singleton, Zhu		Bauer, Rudebusch and Wu	
	Normal period	ZLB period	Normal period	ZLB period	Normal period	ZLB period
			$\Delta TermPr$			
level	-0.0437***	-0.0340*	-0.0320	-0.0275	0.0310	0.0250
	(0.0169)	(0.0196)	(0.0242)	(0.0315)	(0.0525)	(0.0730)
slope	-0.0176**	-0.0150	-0.0193***	-0.0316	-0.0119	-0.0744
	(0.0077)	(0.0141)	(0.0076)	(0.0221)	(0.0185)	(0.0537)
ΔFTT	-0.8898	3.3524***	-0.4106	4.0140***	1.3587	3.5883
	(2.9850)	(0.8673)	(3.6671)	(1.4046)	(7.9918)	(3.4919)
$\overline{R}^2$	0.0144	0.0594	0.0090	0.0333	0.0086	-0.0012
			Δy^Q			
level	-0.0239**	-0.0035	-0.0334**	-0.0076	-0.0986*	-0.0625
	(0.0116)	(0.0084)	(0.0165)	(0.0133)	(0.0523)	(0.0629)
slope	-0.0067	-0.0117*	-0.0043	0.0045	-0.0124	0.0477
	(0.0048)	(0.0060)	(0.0076)	(0.0102)	(0.0241)	(0.0485)
ΔFTT	1.0483	0.9559***	0.6577	0.1638	-1.2002	0.7200
	(2.7048)	(0.3703)	(3.0532)	(0.5736)	(9.5133)	(3.0443)
$\overline{R}^2$	0.0116	0.0371	0.0347	-0.0104	0.0202	-0.0071

Table 4: Regression of the changes in the 10-year yield term premium obtained from the Kim and Wright (2005) (left panel), Joslin et al. (2011) (middle panel) and Bauer et al. (2012) (right panel) decomposition. The Newey-West standard errors with 6 lags are reported in parentheses. *, **, and *** denote significance at the 10, 5, and 1 percent level, respectively. The normal period is 31 December 2003 to 3 October 2007 and the ZLB period is 7 April 2010 to 16 December 2015.

	$xr_{3,t+12}$			$xr_{5,t+12}$			$xr_{10,t+12}$			$xr_{15,t+12}$		
						Normal period						
ΔFTT_t	1.30		-43.38	57.69		-66.75	264.38**		-84.09	476.19**		-73.83
	(25.59)		(36.61)	(40.66)		(69.63)	(106.34)		(153.80)	(199.04)		(247.62)
CP_t		0.19*	0.21*		0.57***	0.61***		1.66***	1.71***		2.65***	2.70***
		(0.12)	(0.13)		(0.15)	(0.19)		(0.28)	(0.36)		(0.57)	(0.66)
$\overline{R}^2$	-0.0303	0.0805	0.0667	-0.0212	0.2705	0.2587	0.0091	0.4960	0.4838	0.0164	0.4623	0.4465
						ZLB period						
ΔFTT_t	16.56***		9.52***	57.36***		27.87**	179.26***		67.90**	277.50***		104.56**
	(2.78)		(3.09)	(7.84)		(10.43)	(22.78)		(26.63)	(39.91)		(42.17)
CP_t		0.12***	0.08***		0.46***	0.35***		1.58***	1.32***		2.45***	2.05***
		(0.03)	(0.03)		(0.10)	(0.11)		(0.19)	(0.28)		(0.22)	(0.39)
$\overline{R}^2$	0.4519	0.5227	0.6202	0.4124	0.5789	0.6407	0.3931	0.6810	0.7155	0.4070	0.7073	0.7431

Table 5: Regression of excess bond returns. The Newey-West standard errors with 12 lags are reported in parentheses. *, **, and *** denote significance at the 10, 5, and 1 percent level, respectively. The normal period is December 2003 to October 2007 and the ZLB period is May 2010 to December 2015.

Δx_{t+1} :	Expected inflation		10-year real yield	
	Normal period	ZLB period	Normal period	ZLB period
x_t	-0.0816	-0.0411	-0.0378	-0.0237
	(0.0355)	(0.0152)	(0.0239)	(0.0102)
ΔFTT_t	-1.3165	1.8240***	0.6722	1.4850*
	(1.7456)	(0.6279)	(5.0773)	(0.7759)
$\overline{R}^2$	0.0270	0.0266	0.0076	0.0150

Table 6: Regression of the changes in the expected inflation and changes in the 10−year real yield. The Newey-West standard errors with 6 lags are reported in parentheses. *, **, and *** denote significance at the 10, 5, and 1 percent level, respectively. The significance of the lagged regressand in level is not indicated. The normal period is 31 December 2003 to 3 October 2007 and the ZLB period is 7 April 2010 to 16 December 2015.

	$PC1$	$PC2$	$InitCl$	$ExpInf$	ΔFTT
$PC1$	1				
$PC2$	-0.0009	1			
$InitCl$	-0.0897	-0.0592	1		
$ExpInf$	0.1207*	0.6488***	-0.0683	1	
ΔFTT	0.4972***	0.3238***	-0.0834	0.3560***	1

Table 7: Correlation coefficients of the first two principal components of yields, initial claims, expected inflation and Treasury security purchases by the Federal Reserve on the FOMC non-meeting weeks. * and *** denote significance at the 10 and 1 percent level, respectively. The sample period is from 7 April 2010 to 16 December 2015.

	$PC1$	$PC2$	$InitCl$	$ExpInf$	ΔFTT
$PC1_{t-1}$	0.9495***	0.0061*	-0.1215	-0.0033	0.1182**
	(0.0110)	(0.0032)	(0.2061)	(0.0056)	(0.0576)
$PC2_{t-1}$	0.0476	0.9946***	-0.3054	0.0285*	0.1948*
	(0.0385)	(0.0132)	(0.6684)	(0.0166)	(0.1074)
$InitCl_{t-1}$	-0.0049	0.0009	0.3857***	-0.0022*	-0.0027
	(0.0034)	(0.0011)	(0.0451)	(0.0012)	(0.0085)
$ExpInf_{t-1}$	-0.1762***	0.0103	0.5832	0.9331***	0.1704
	(0.0403)	(0.0169)	(0.9779)	(0.0202)	(0.1258)
ΔFTT_{t-1}	0.0903***	-0.0110*	-0.0966	0.0186**	0.5833***
	(0.0218)	(0.0063)	(0.2536)	(0.0084)	(0.0977)
$\overline{R}^2$	0.9672	0.9756	0.1297	0.9363	0.5826

Table 8: Regression of the $VAR(1)$ with first two principal components of Treasury yields, initial claims, expected inflation and Treasury security purchases by the Federal Reserve. The regression includes dummy variables for the FOMC meeting weeks. The Newey-West standard errors with 6 lags are reported in parentheses. The $p-$values for the $F-$test are reported in square brackets. *, **, and *** denote significance at the 10, 5, and 1 percent level, respectively. The sample period is from 7 April 2010 to 16 December 2015.

Δx_{t+1} :	Baa-Treasury		Aaa-Treasury		Baa-Aaa	
	Normal period	ZLB period	Normal period	ZLB period	Normal period	ZLB period
x_t	-0.0315	-0.0114	-0.0159	-0.0677	-0.0279	-0.0040
	(0.0163)	(0.0116)	(0.0116)	(0.0210)	(0.0188)	(0.0117)
ΔFTT_t	-0.3185	-1.3326**	-0.9991	-0.4030	0.6399	-0.7817**
	(2.4107)	(0.5513)	(1.5890)	(0.5873)	(1.5325)	(0.3078)
$\overline{R}^2$	0.0121	0.0200	0.0011	0.0381	-0.0009	0.0089

Table 9: Regression of the changes in the corporate spread. The Newey-West standard errors with 6 lags are reported in parentheses. *, **, and *** denote significance at the 10, 5, and 1 percent level, respectively. The significance of the lagged regressand in level is not indicated. The normal period is 31 December 2003 to 3 October 2007 and the ZLB period is 7 April 2010 to 16 December 2015.

	r_{t+1}	
	Normal period	ZLB period
$level_t$	-0.0061*	-0.0083**
	(0.0032)	(0.0040)
$slope_t$	-0.0018	0.0012
	(0.0012)	(0.0029)
ΔFTT_t	-1.0723*	-0.1074
	(0.5573)	(0.1682)
ΔFTT_{t-1}	0.9190*	0.4881***
	(0.4936)	(0.1796)
$\overline{R}^2$	0.0114	0.0096

Table 10: Regression of the stock market returns. The Newey-West standard errors with 6 lags are reported in parentheses. *, **, and *** denote significance at the 10, 5, and 1 percent level, respectively. The normal period is 31 December 2003 to 3 October 2007 and the ZLB period is 7 April 2010 to 16 December 2015.

Figures

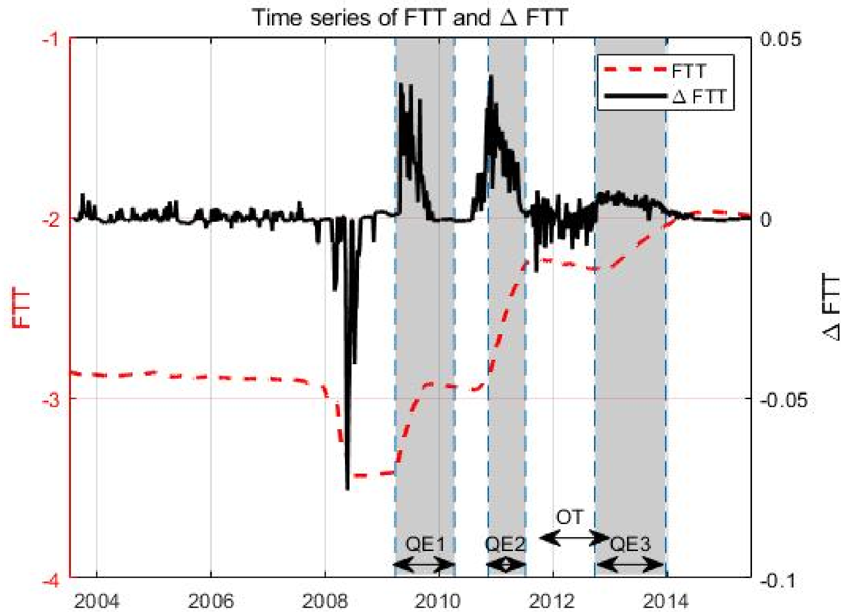


Figure 1: Time series of the log ratio of the total Federal Reserve holdings of Treasury securities to the nominal GDP (FTT, left scale), and its first differences (ΔFTT, right scale). The grey areas denote the three phases of QE.

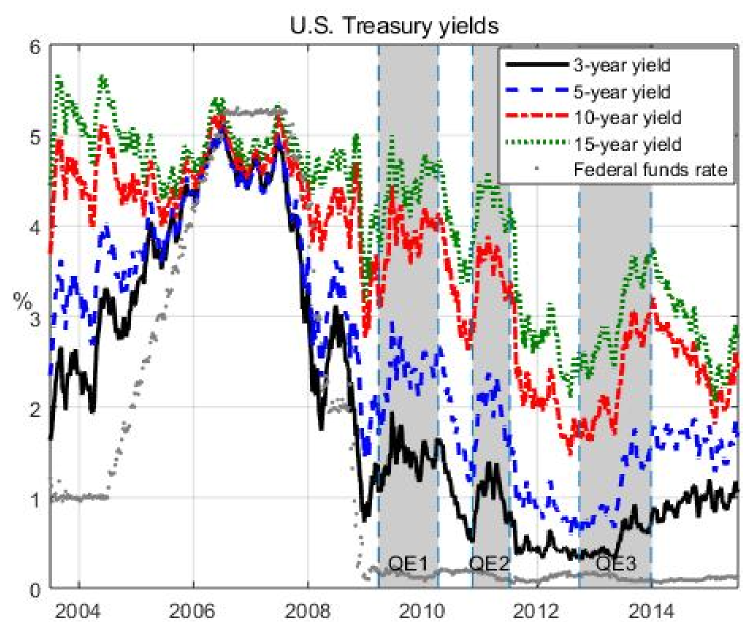


Figure 2: Time series of the 3-, 5-, 10- and 15-year Treasury yields and the federal funds rate. The grey areas denote the three phases of QE.

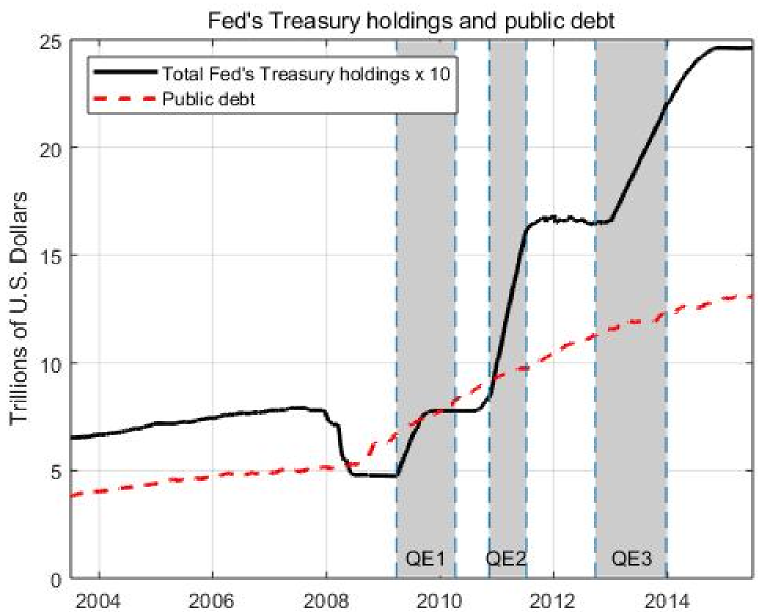


Figure 3: Time series of the total holdings of Treasury securities by the Federal Reserve (×10) and the total public debt. The grey areas denote the three phases of QE.

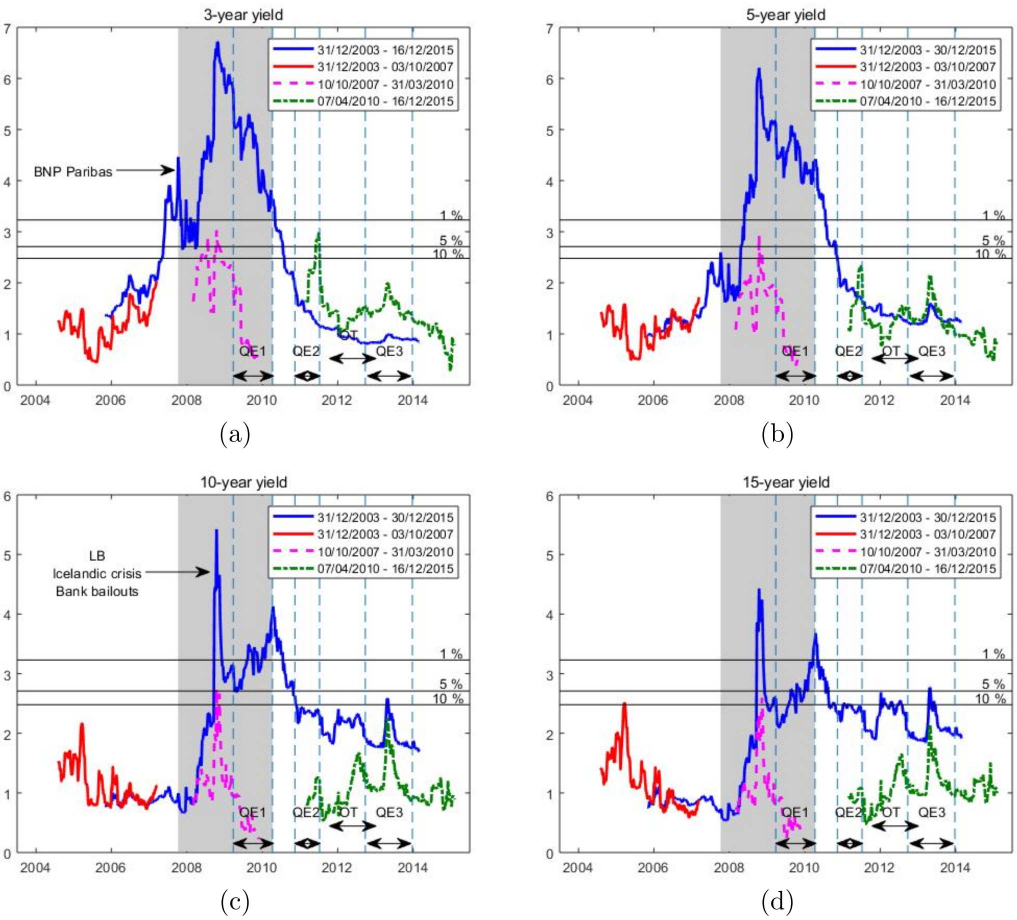


Figure 4: Sup Wald test for the sample 31 December 2003 to 30 December 2015 and three subsamples. The grey area denotes the period of instability in the regression of changes in Treasury yields on monetary policy and macro variables. The horizontal lines denote the significance level of the test.

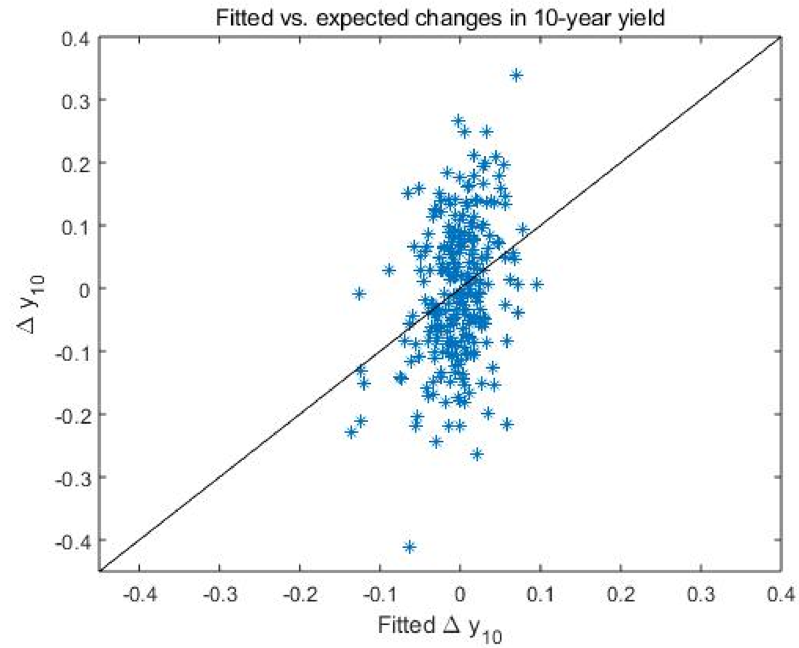


Figure 5: Actual vs. fitted observations from the regression (2) for the 10−year yield for the sample period 7 April 2010 to 16 December 2015.

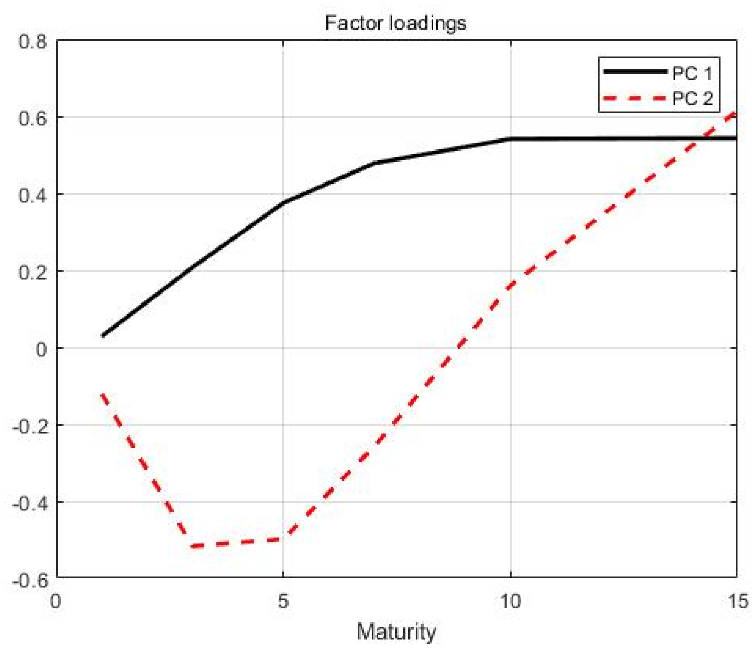
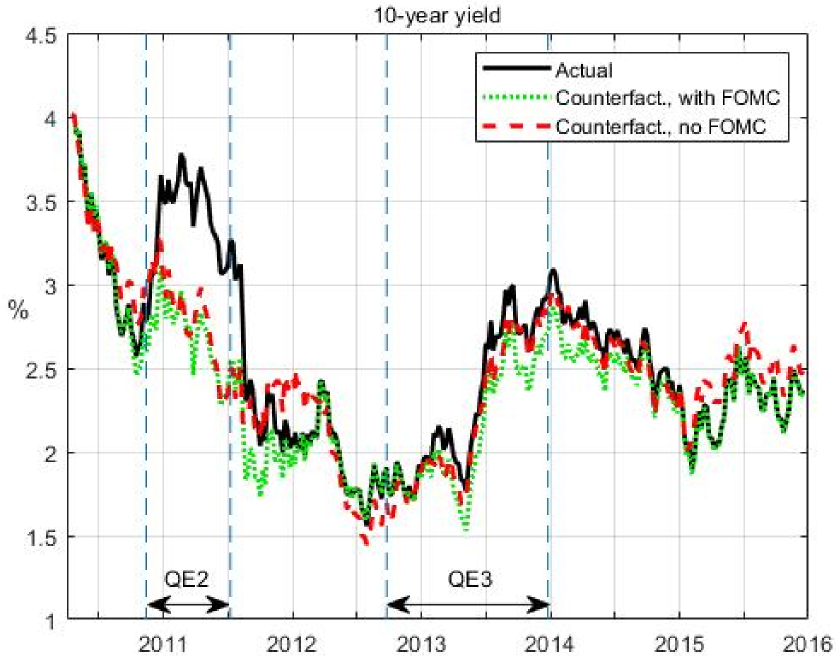
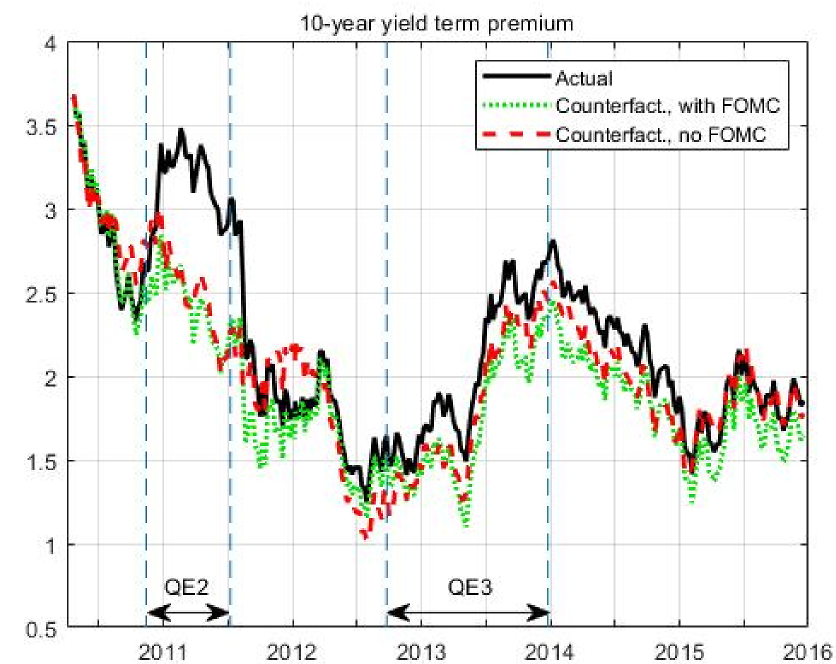


Figure 6: Factor loadings for the first two principal components of yields estimated on the sample period 7 April 2010 to 16 December 2015.



(a)



(b)

Figure 7: Actual and counterfactual 10−year yield and yield term premium. The sample period is 7 April 2010 to 16 December 2015.

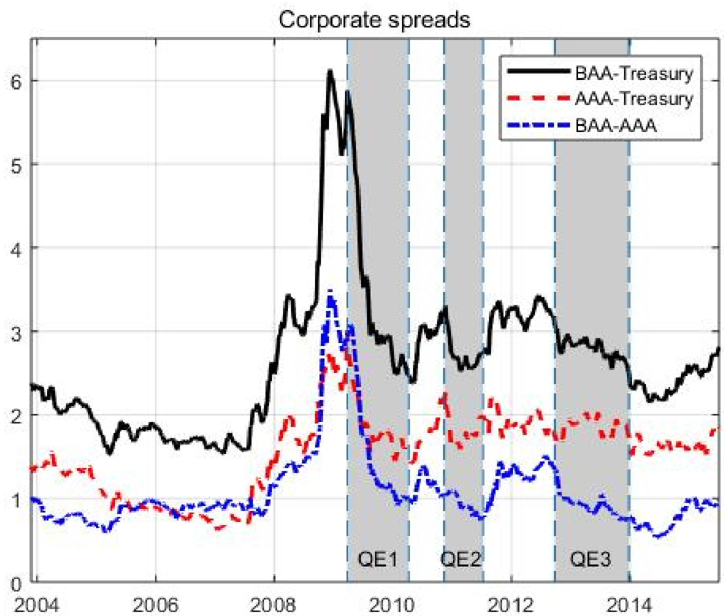


Figure 8: Time series of three corporate spreads: the spread between Moody's seasoned Baa corporate yields and the 10−year Treasury yield; the spread between Aaa yield and 10−year Treasury yield and the spread between Baa and Aaa yields. The grey areas denote the three phases of QE.

Figure 9: Time series of $S\&P$ 500 index. The grey areas denote the three phases of QE.