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Macroeconomic announcements and their effects on bond market volatility

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Abstract

This paper examines the effects of macroeconomic news announcements on bond market volatility for the U.S., Germany and the U.K. Using an ARCH-type specification, volatility effects and their persistence is modeled for employment, PPI and target interest rate announcements. It is found that all types of announcements have significant effects on U.S. Treasury volatility, while this is only the case for target interest rate announcements in Europe. Furthermore, it is found that announcement day volatility sometimes persist, seemingly inconsistent with the semi-strong form of the Efficient Market Hypothesis. Moreover, it is shown that FOMC announcements are more important than ECB and BoE announcements for longer-term maturity bonds in Europe, although domestic announcements are found to be more important in the short end of the yield curve.

Contents

1 Introduction

One of the most well known characteristics of financial markets is the relation between risk and return. By investing money in stocks, bonds, derivatives or other assets, investors take upon them risk. Economic theory states that, on average, the return from investing is proportional to the risk taken, leading to the risk/return trade off. In this trade off, 'risk' is defined as a measure of the uncertainty in returns, often described as the degree by which returns vary over time. This 'volatility' of returns can be predicted and hence a risk analysis of a financial asset can be made.

'What drives asset price volatility?' is a question that has been extensively studied in economics, although the exact sources of observed volatility remain unknown. From theory we know that market prices (should) reflect all information currently available regarding the financial asset that is priced. Hence, 'new' information such as earning statements or dividend announcements are a possible source for asset price volatility (Aharony & Swary, 1980). However, for government bonds such information is not available, as the 'dividends' usually consist of fixed coupon payments. Consequently, other information sources are of interest, such as the release of macroeconomic news.

In fact, this topic has received quite some attention in research already and was the subject of the paper by Jones, Lamont & Lumsdaine (1998). They studied how regularly scheduled U.S. government releases of the Producer Price Index (PPI) and the Employment Situation Report affected Treasury bond prices, in the period between October 9, 1979 and December 31, 1995. By studying periodic, preannounced macroeconomic news announcements they investigated whether non-autocorrelated news give rise to autocorrelated volatility. Their results confirm that the release of macroeconomic information has a significant positive effect on bond market volatility, but that these shocks to volatility do not persist in the days following an announcement. Furthermore, they find that a risk premium can be earned on announcement days, consistent with the idea of a higher exposure to macroeconomic risk on those days. This paper will extend the works of Jones et al. (1998) by studying more countries and more announcement types, leading to the research question of this paper:

What is the effect of macroeconomic news announcements on bond market volatility, and do these effects differ for various countries and announcement types?

This question will be answered by studying excess return volatility for government bonds in the U.S., Germany and the U.K., for the period January 1, 1999 to December 31, 2016. The paper by Jones et al. (1998) will be extended by using PPI and employment announcements, as well as target interest rate announcements done by the Federal Open Market Committee (FOMC), the European Central Bank (ECB) and the Bank of England (BoE).

In order to analyze differing effects for the U.S., Germany and the U.K., three thing will be compared. First, the effects of macroeconomic announcements on the first and second moments of government bond returns will be studied. Secondly, it will be investigated whether new macroeconomic information is immediately reflected in bond prices or that this takes longer, with shocks to volatility persisting in the following days. By doing so, results obtained by Jones et al. (1998) will be replicated for a more recent sample period. The third point of my analysis focuses on the relative importance of central bank announcements. For Germany and the U.K., announcements done by the ECB and BoE will be compared to announcements done by the FOMC, thereby extending earlier academic literature (e.g., Nikkinen & Sahlström, 2004; Andersson, Overby & Sebestyén, 2009).

Next to academic relevance, the results of this research will also be of social relevance. Gaining insights into the effects of macroeconomic news on bond markets could benefit policy makers, such as the Federal Reserve and the European Central Bank. Especially if notable differences are found in how financial markets react, potential improvements could be identified towards more effective monetary policy making. Besides policy makers, investors could also profit from further insights into the forces at play in bond markets, potentially leading to profitable trading strategies. Moreover, it could provide investors with helpful insights into the differences in risk between the European and the U.S. bond market. This would further enable investors to invest according to their individual risk-preference, resulting in a more optimal risk allocation.

This paper is organized as follows: Section 2 presents an overview of applicable economic theory as well as current academic literature related to this work. This will provide a theoretical framework for this research and explain how my hypotheses are embedded in existing literature. In Section 3 a description and overview of the data will be given, followed by a preliminary analysis using ordinary-least squares (OLS) regressions in Section 4. This section shows that both PPI and employment announcements have significant positive effects on bond market volatility in the U.S., confirming earlier findings by Jones et al. (1998). However, for similar announcements done in Germany and the U.K., no such effect is found. It is also shown that significant announcement day effects on bond market volatility are present on days when central banks make announcements. However, for none of the countries or announcements studied, significant higher excess returns are found on announcement days. In Section 5 I continue by modeling conditional variance using a mixture GARCH(1,1) specification developed by Jones et al. (1998), in order to study volatility persistence. It is found that volatility shocks related to PPI and employment announcements do not persist at all, but that shocks related to announcements done by central banks do persist. Furthermore, evidence is found that for Germany and the U.K, FOMC announcements have more impact on the volatility of longer maturity government bond returns than announcements made by the ECB and BoE. Finally, in Section 6, a conclusion is given, limitations of this work are discussed and suggestions for further research are presented.

2 Literature review

The first thing that will be studied in this paper are the possible effects of macroeconomic announcements on the first and second moments of government bond returns. However, the fact that financial markets would react to macroeconomic announcements at all, is evidence against the 'strong' form of the Efficient Market Hypothesis (EMH). The EMH states that if markets are 'efficient', asset prices should always incorporate and reflect all relevant information concerning an asset. The hypothesis, as described by Malkiel & Fama (1970), is presented in three forms: in the 'weak' form, where the information set consists of only historical information. In the 'semi-strong' form, all historical information is incorporated into prices and 'new' public information is immediately reflected in prices, hence the information set consists of all publicly known information. In the 'strong' form of the hypothesis, not only public information is incorporated into prices, but also private information, effectively making 'insider' trading impossible. At the same time, this would imply that merely announcing macroeconomic figures should not have any effect on asset prices. This follows from the fact that the PPI is simply the result of a survey of prices taken by the Bureau of Labor Statistics. Hence, announcing the PPI is merely making private information publicly known.

Although the exact extent to which markets are efficient is often disputed in Finance, the EMH forms a theoretical background for the idea that information affects asset prices. This idea led to a rich variety of academic literature regarding the relation between information and asset prices. For example, Roll (1988) examined the extent to which stock prices of large U.S. firms can be explained *in hindsight*, using information about economic risk factors, industry specific information and information about events specific to the firm itself. However, even when including all explanatory variables available, the percentage of the variation of stock prices that he was able to explain (as seen from the R^2 of the regression), was on average only

35% for monthly data and 25% for daily data. This work, and other research as done by for example by Mitchell & Mulherin (1994), shows that linking asset price movements to measures of news is difficult, although from a theoretical perspective, they are certainly related.

Another way to study the link between information and asset prices is by taking a subset of all information, and relating that to observed asset price volatility. That is what was done by Ederington & Lee (1993), when they studied the impact of scheduled macroeconomic news announcements on interest rate and foreign exchange futures markets. Using 5-minute intraday data, they showed that macroeconomic announcements explain the largest part of observed volatility patterns, explaining both time-of-the-day and the day-of-the-week effects. They found that the largest part of the price adjustments (following new announcements) is within the first minute, followed by substantially higher volatility in the first 15 minutes, and slightly higher volatility in the following hours.

Similar research was done by Kuttner (2001), who studied the effect of monetary policy actions taken by the FOMC on Treasury bonds yields. However, instead of using the Federal Reserve's constant maturity interest rate series (as did Jones et al. 1998; and as this paper does), Kuttner used data from the futures market for federal funds. By doing so, he was able to separate announcements into an anticipated (and hence priced) component and an unanticipated component. By regressing only on the latter component, he found large and highly significant responses of interest rates.

Together with the results obtained by Jones et al. (1998), this leads to the first hypothesis: "Macroeconomic news announcements have a positive effect on government bond volatility and a significant higher excess returns can be earned on announcement days". The second part of this hypothesis is not only based on finding by Jones et al. (1998), but can also be motivated from theory. In financial economics, the link between risk and return is evident. As pointed out by Engle, Lilien & Robins (1987) if "the degree of uncertainty in asset returns varies over time, the compensation required by risk averse economic agents for holding these assets, must also be varying" (p. 391). Hence, if macroeconomic announcements have a significant positive effect on bond market volatility, we expect that this coincides with higher expected returns.

The second aim of this paper is to see whether new macroeconomic information is immediately reflected in bond prices, or that this does not happen instantaneous. In that case, shocks to volatility could persist in the days following an announcement, leading to autocorrelated returns. As documented by, for example Bollerslev, Chou & Kroner (1992), empirical evidence shows that volatility in financial markets is correlated and clustered over time. This has led to the autoregressive conditional heteroskedasticity (ARCH) framework of Engle (1982), later followed by the generalized autoregressive conditional heteroskedasticity (GARCH) framework of Bollerslev (1986). For an overview of previous literature on these types of models, please refer to Bollerslev et al. (1992) or Pagan (1996).

Besides modeling of (empirically observed) persistent volatility, it is also of interest to study possible sources of autocorrelated volatility. In their paper, Jones et al. (1998) mention several possible explanations for autocorrelated volatility. As it is known that volatility is caused by the arrival of new information, the first possible explanation they give for persistent volatility is that the arrival of news itself is serial correlated. Using an index of news events (as proxied by *New York Times* headlines), Jones et al. (1998) show that this is indeed the case. Another explanation for autocorrelated volatility could be that private information diffuses gradually among investors, as they 'learn' from each other by observing market volumes and prices. As formalized by Hong & Stein (1999), this could lead to an under reaction of prices in the short run, while prizes would overreact in the long run due to momentum trading strategies. Research by Hong, Lim & Stein (2000) proxies the speed of information flow by the number of analysts covering a stock. They find evidence for this behavioral explanation of autocorrelated volatility, although they mention that evidence should be tempered with caution. Other possible explanations mentioned by Jones et al. (1998) are that prices respond instantly to news, but incorrectly due to not fully rational behavior by investors. Daniel, Hirshleifer & Subrahmanyam (1998) show that investor overconfidence about private information and biased self-attribution (asymmetric shifts in investors confidence due to investment outcomes) lead to negative autocorrelation in returns and excess volatility.

In this paper, earlier work of Jones et al. (1998) is be followed with respect to possible sources of autocorrelated volatility. By studying periodic, preannounced macroeconomic news releases (that is by assumption, non-autocorrelated news), it is tried to find evidence for the first explanation of autocorrelated volatility. If it is found that macroeconomic news announcements have a significant effect on bond market volatility (as the first hypothesis states), and if these shocks do not persist, this would affirm the theory that markets are efficient in the way information is incorporated into prices. However, if it is found that shocks to volatility do persist, this would provide evidence for alternative explanations, as given above. Of course, a critical assumption that is made here, is that the announcements studied are in fact non-autocorrelated; I assume they each consist of a one-time piece of new information. Based on earlier finding by Jones et al. (1998), this leads to the following hypothesis: "Macroeconomic news is immediately incorporated in bond prices and announcement day shocks to volatility do not persist". This hypothesis will be tested using a mixture GARCH specification, as proposed by Jones et al. (1998).

The third point of our analysis focuses on the relative importance of central bank announcements done by the ECB and BoE versus announcements done by the FOMC. Nikkinen & Sahlström (2004) study the importance of both domestic and U.S. macroeconomic announcements as information sources for German and Finnish stocks markets, in the period January 1996 to December 1999. They regress changes in implied volatility indices on daily dummy variables for different types of announcements, being the Employment Situation Report, CPI and PPI announcements, as well as dummy variables for meeting days of central banks. Their results show that only the U.S. employment report and FOMC meeting days have a significant impact on implied volatility measures. Hence, they conclude that U.S. macroeconomic announcements are a valuable source of information for European stock markets, while domestic announcements appear to be unimportant. This paper will use similar announcements to study volatility effects on government bond markets, using a larger and more recent time period. By combining central bank announcements done by the ECB, BoE and FOMC in a mixture GARCH specification, the relative importance of each announcement can be studied. This will be done for different government bond maturities, in order to see whether the relative importance of different announcements is related to the term to maturity of the bonds studied.

Similar research was done by Andersson et al. (2009), who studied market response to macroeconomic news and ECB monetary policy releases, using intra-day prices on German long-term bond futures. They find that U.S. announcements are the most influential among the announcements studied and give three possible explanations for this finding. First, they perceive the U.S. as the engine of global growth, explaining its importance in international financial markets. Secondly, they mention the ongoing integration of business cycles, leading to a higher degrees of interdependence between economies. Thirdly, they note that U.S. macroeconomic information is usually released earlier than equivalent European data, hence decreasing the informational value of the latter.

Based on the results by Nikkinen & Sahlström (2004) and (Andersson et al., 2009), the third hypothesis is formulated as follows: "Central bank announcements done by the FOMC have more impact on bond market volatility than similar announcements done by the ECB or BoE". This hypothesis will be tested using daily government bond returns of German Bunds and English Gilts, as it is expected that U.S. announcements also effect European financial markets. As it was earlier found by Nikkinen & Sahlström (2004) that European announcements do not significantly effect implied volatility on European markets, the effects of European announcements on U.S. Treasury markets are not separately investigated.

3 Data

For this research daily returns on U.S. and European bonds will be used, with maturities of one, ten and 30 years[1]. These maturities are chosen in order to differentiate between the impact of macroeconomic news announcements on short-term, mid-term and long-term government bonds. Furthermore, these assets are chosen because earlier research shows that the macroeconomic news announcements that are of interest for this paper, have material impact on (primarily) the bond market (Jones et al., 1998). As there are no bonds issued by the European Central Bank, national bonds of two countries in Europe will be used as proxy, being Germany and the U.K. The choice for these two countries is based on their GDP output, as percentage of the output of the total Euro area. With Germany being the largest economy within the Euro area, and the U.K. running second, these countries are considered appropriate benchmarks for the European market.

Returns of U.S. government bonds are calculated using the Federal Reserve's constant maturity interest rate series (FRB H15 release), obtained from the website of the Federal Reserve. For Germany and the U.K. their respective counterparts are obtained via Bloomberg[2]. To calculate returns from the published (daily) yields, a similar method is used as Jones et al. did in their 1998 paper, please refer to appendix A for the exact calculations. The excess returns r^e_{t+1} (for U.S. Treasuries) are calculated over the risk free rate, assumed equal to the rate on three-month Treasury bills. For calculating the excess returns of holding German Bunds and English Gilts, the procedure described in appendix A is used as well, but then using the three-month Euribor and LIBOR as risk free rate. For the Euribor and LIBOR, on some bank holidays no rate was published, although yield information from German and English government bonds was available. In order to be able to calculate excess returns on those days, the Euribor and LIBOR rates were linearly interpolated if data was missing. This was done for not more than two consecutive days (Friday and Monday). Furthermore, it was observed that sample sizes differ slightly between countries in the same time period, as the exact amount of trading days/bank holidays differs. This is not considered an issue, as analyses are done per country. Besides that, it was observed that the U.K. government one-year maturity Gilt has nine observations less than the ten-year and 30-year Gilts, due to a missing week of data in October 2007. Consequently, one employment and one PPI announcement less is included in the analysis for the one-year maturity Gilt. The other four missing data points were all non-announcement dates.

The dates of employment report announcements and PPI announcements (as made by the Bureau of Labor Statistics) were obtained through Bloomberg. For both European countries, their respective counterparts for the U.S. employment report and PPI announcements were used, as published by the DESTATIS Statistisches Bundesambt and the Office for National Statistics. For consistency reasons, there was decided not to change the type of announcements selected for Germany and the U.K., in order to aid comparison between countries. Moreover, it was checked that PPI announcements were not directly preceded by (similar) CPI announcements, as this would effect the informational value of the announcement published latest. Just like in the U.S., it was found that the Producer Price Index is generally published a few days before the Consumer Price Index, hence making the latter less informative from the market's perspective. The dates of German and English Employment and PPI announcements were obtained through Bloomberg as well[3].

[1] The original work of Jones et al. (1998) used five year maturity Treasuries instead of one year maturity Treasuries. In order to be able to focus on short term effects, I have chosen to 'switch' from five year to one year maturity Treasury bonds

[2] Bloomberg tickers: GDBR1, GDBR10 and GDBR30 for Germany. GUKG1, GUKG10 and GUKG30 for the U.K.

[3] Bloomberg tickers: GRUEPR Index and GRPFIMOM Index for Germany. UKUER Index and UKPPIOC

Data regarding meetings of, and interest rate decisions taken by, the Federal Open Market Committee (FOMC), the Governing Council of the European Central Bank (ECB) and the Monetary Policy Committee of the Bank of England (BoE) were obtained from their respective websites[4] and Bloomberg (using the UKBRBASE Index). However, not all data obtained regarding the FOMC meetings was publicly known at the time decisions were taken, as meeting outcomes were only immediately announced from 1995 onwards. Before then, announcements were only published several weeks/months after the FOMC meeting had taken place. In order to include only publicly known information, for which the announcement date was known in advance, the analysis of announcements made by central banks extends only from 1999 to 2016, excluding earlier observations for the U.S. Moreover, due to time differences, FOMC announcements in general took place when European markets were already closed. Therefore, the first trading day after FOMC announcements is taken as announcement day for Germany and the U.K. when considering FOMC announcement effects.

The first part of the analysis of this paper will be focusing on the situation in the U.S., for the period 1982 - 2016 (8751 observations). By doing so, the earlier work of Jones et al. (1998) will be replicated with a larger data sample. The second part of the analysis will be focusing on the comparison of the situations in the U.S. and Europe, for which a smaller time period is investigated, running from 1999 to 2016. The start of the second sample is set later in time, as besides the structural break described above, another structural break in the European time series is observed, following the formation of the European Central Bank and the introduction of the Euro as currency.

Several aspects are important to keep in mind when using the aforementioned sample time frame. First of all, the complete sample used for the U.S. runs for 34 years, which is a considerable period of time. On the one hand, this is a positive aspect, as more data in general results in better statistical inferences. On the other hand, when using the complete sample, the implicit assumption is made that the topic of interest, in this case the reaction of the bond market to macroeconomic news, did not change during the time period studied. However, the longer the sample time frame is, the stricter this assumption becomes and the more likely it becomes that this assumption is violated. Furthermore, the sample time frame includes the period 2007-2008, where we've seen a global financial crisis, and the period 2010 onwards, where several European countries faced a sovereign debt crisis. Both these crises had major implications for financial markets and led to several policy implications, which could severely affect the analysis that is to follow. However, within the scope of this paper, no explicit corrections are made for these crises, although this could certainly be a good start for further research.

3.1 U.S. Government bonds

In Table 1, descriptive statistics are shown for U.S. Treasury returns over the the time period of January 5, 1982 to December 30, 2016, for both announcement and non-announcement days. Table 2 and 3 show similar statistics for Germany and U.K. Treasury returns, which will be discussed in subsection 3.2.

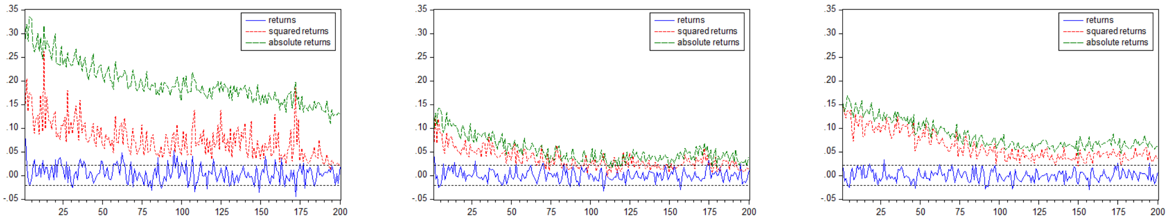
As can be seen in Table 1, excess returns vary from on average 0.003% per trading day for one-year securities, to 0.026% per trading day for 30-year securities, roughly equal to 0.72% and 6.75% per year respectively (continuously compounded, using 250 trading days per year). However, daily returns are quite variable, as can be seen from the standard deviation, with daily excess returns as high as 4.8% for ten-year Treasury returns (on October 20, 1987, the day after the stock market crash) and as low as -5.0% for 30-year Treasury returns on November 9, 2016, following the election of Donald J. Trump as the 45th president of the United States. Neither of the two dates is an announcement date, illustrating that only part of the variation will

Index for the U.K.

[4]for the FOMC refer to: `https://www.federalreserve.gov/monetarypolicy/fomc_historical_year.htm` and for the ECB refer to: `https://www.ecb.europa.eu/press/govcdec/mopo/2017/html/index.en.html`

be captured by the announcements of interest in this study. Furthermore, it can be seen that excess returns are positively skewed (although less so for long-term maturities) and significantly fat-tailed (especially the one-year Treasury returns), far exceeding the theoretical value of three. Hence, it is clear that the excess returns are not normally distributed.

Moreover, in Table 1 first-order autocorrelations (ρ) are shown. In the full sample, for the squared value of excess returns, ρ ranges 0.07 to 0.15, while ranging from 0.10 to 0.29 for the absolute value of excess returns. However, for the first moment of the returns first-order autocorrelation is only in the range of 0.02 to 0.08. From these fact, graphically illustrated in Figure 1, the use of ARCH class models can be motivated. Interesting to see is that the empirical autocorrelation function (EAF) of the one-year maturity Treasury differs from the EAF's of the ten- and 30-year maturity Treasuries, showing significant higher autocorrelation in the absolute returns. However, in general the bond returns show similar autocorrelated volatility as stock returns and foreign exchange returns (Cont, 2001).



1-year 10-year 30-year

Figure 1: Empirical Autocorrelation Functions for U.S. Treasury bond returns

The second half of Table 1 focuses on (various types of) announcements dates. Analogous to Jones et al. (1998), Employment and PPI announcements are pooled, as they are relatively similar in their effects on the first and second moment of daily excess returns (as seen in Table 1). However, it can be seen that volatility is in general higher on employment announcement dates than on PPI announcement days, although the latter announcements show significant higher excess returns than non-announcement dates, which is not the case for employment announcement dates.

For announcements done by the Federal Open Market Committee (FOMC), regarding the target interest rate, significant higher volatility is observed on announcement dates for ten- and 30-year maturity treasuries, while this is to a smaller degree the case for one-year maturities. Furthermore, significant excess returns are obtained for one- and ten-year maturity bonds on FOMC announcement dates, although this is not the case for longer maturities in this sample.

In general it is interesting to see that the 12% of the trading days with announcements (1069 in total, ten times the FOMC announcements coincided with employment or PPI-announcements), saw significant higher excess returns averaging 0.050% per trading day (for the ten-year bond, 13.3% on an annual base) than the non-announcement dates, where the average excess returns averaged only 0.013% for the same maturity (3.3% on an annual base). Moreover, we generally see increasing excess returns with increasing maturities, which is consistent with the idea that a higher exposure to macroeconomic risk should be accompanied by higher expected returns. Although theoretically a risk premium over the risk-free rate is expected for holding longer-term government bonds, this does not necessarily have to hold, as shown by, for example by Campbell (1995), who finds negative excess returns over the period 1952-1991 for ten-year Treasury bonds.

Table 1: Summary statistics: U.S. Treasury bond daily excess returns

XR_t is the daily continuously compounded excess return of the relevant constant maturity Treasury security over the three-month Treasury bill. Returns are expressed in percent, i.e., multiplied by 100. The sample extends from January 5, 1982 to December 30, 2016.

	1-yr			10-yr			30-yr		
	XR_t	XR_t^2	$\|XR_t\|$	XR_t	XR_t^2	$\|XR_t\|$	XR_t	XR_t^2	$\|XR_t\|$
Full Sample ($N = 8751$)									
Mean	0.003[c]	0.004[c]	0.036[c]	0.017[c]	0.240[c]	0.362[c]	0.026[c]	0.680[c]	0.607[c]
Std. Dev.	0.062	0.020	0.050	0.490	0.574	0.330	0.824	1.563	0.558
Min.	−0.925			−2.716			−4.993		
Max.	0.791			4.804			7.522		
ρ (autocorr.)	0.077	0.152	0.290	0.039	0.072	0.101	0.016	0.148	0.151
Kurtosis	27.012			6.695			6.280		
Skewness	0.610			0.124			0.077		
Employment report announcement dates ($N = 416$)									
Mean	0.008[x]	0.008[cz]	0.063[cz]	0.000	0.481[cz]	0.548[cz]	−0.024	1.162[cz]	0.856[cz]
Std. Dev.	0.092	0.017	0.067	0.695	0.741	0.426	1.079	1.735	0.657
ρ (t to t+1)	0.044	0.282	0.278	0.012	0.071	0.062	−0.024	0.037	0.041
PPI announcement dates ($N = 415$)									
Mean	0.011[cz]	0.005[cx]	0.042[cz]	0.073[cy]	0.284[cz]	0.411[cz]	0.136[cz]	0.738[cx]	0.679[cz]
Std. Dev.	0.068	0.016	0.054	0.529	0.496	0.340	0.849	1.109	0.527
ρ (t to t+1)	0.021	0.142	0.278	−0.016	0.037	0.105	−0.055	0.180	0.209
Pooled announcement dates ($N = 830$)									
Mean	0.009[cz]	0.007[cz]	0.053[cz]	0.036[a]	0.383[cz]	0.480[cz]	0.055	0.951[cz]	0.768[cz]
Std. Dev.	0.081	0.017	0.062	0.619	0.638	0.392	0.974	1.472	0.602
ρ (t to t+1)	0.036	0.216	0.269	0.001	0.032	0.066	−0.038	0.087	0.101
FOMC announcement dates ($N = 249$)									
Mean	0.014[cz]	0.004[c]	0.039[cx]	0.075[ax]	0.376[cy]	0.424[cz]	0.072	0.881[cy]	0.670[cy]
Std. Dev.	0.059	0.012	0.046	0.610	1.365	0.444	0.938	2.053	0.658
ρ (t to t+1)	−0.078	0.012	0.162	0.013	0.072	0.068	0.180	0.359	0.262
Non-announcement dates ($N = 7682$)									
Mean	0.002[c]	0.004[c]	0.035[c]	0.013[b]	0.221[c]	0.348[c]	0.021[b]	0.644[c]	0.587[c]
Std. Dev.	0.060	0.020	0.048	0.470	0.518	0.316	0.802	1.551	0.547
ρ (t to t+1)	0.087	0.148	0.297	0.044	0.083	0.109	0.016	0.141	0.153

[a]Different from zero at the 10% significance level (two-tailed for returns, one-tailed for variances)
[b]Different from zero at the 5% significance level (two-tailed for returns, one-tailed for variances)
[c]Different from zero at the 1% significance level (two-tailed for returns, one-tailed for variances)
[x]Different from non-announcement mean value at the 10% significance level (one-tailed)
[y]Different from non-announcement mean value at the 5% significance level (one-tailed)
[z]Different from non-announcement mean value at the 1% significance level (one-tailed)

3.2 European bonds

Next we turn to the European bonds, as proxied by German government Bunds and U.K. government Gilts. For these government bonds summary statistics are shown in Tables 2 and 3. In the sample period studied for the German and U.K. government bonds, extending from January 1, 1999 to December 30, 2016, excess returns appear to be a bit lower than previously seen in the larger U.S. sample. For German Bunds, excess returns averaged between 0.000% and 0.025% per trading day, while averaging between 0.000% and 0.015% per trading day for English Gilts.

From Table 2 and 3 it can be seen that minimum and maximum daily excess returns appear to be the same order of magnitude as previously seen in for U.S. Treasuries. Furthermore, extreme returns can sometimes be 'linked' to specific events in time. For example, the maximum one-day return for the German ten-year government Bund, equaling 2.347%, was on November 1, 2011. This coincided with the Greek proposed economy referendum concerning acceptance of borrowing conditions set out by the International Monetary Fund (IMF) and European Central Bank (ECB). Another example would be the minimum one-day return for the German 30-year government Bund, equaling -4.608%, on June 2, 2015, following the release of the fourth economic bulletin by the ECB that year. The latter is considered an announcement date (due to the employment report that was published on the same day), the former is not. Furthermore, just like U.S. Treasury returns, German and U.K. government bond returns are significantly fat tailed, although not always positively skewed. Autocorrelations (not shown in Tables 2 and 3), were checked and showed similar patterns as autocorrelations for U.S. returns did.

However, when turning to the announcement dates, shown in the bottom halves of Tables 2 and 3, the results shown are rather dissimilar. Shown in Table 2 is that the average excess return on non-announcement dates (0.014%) is actually higher than on announcement dates (0.011%). Employment report and PPI announcements dates show quite similar excess returns, that are on average appear to be lower than on non-announcement dates, albeit not significantly so. Moreover, German bond market volatility, as measured by squared returns (XR_t^2) and absolute returns ($|XR_t|$), appears not to be significantly different (higher) on these dates, than on non-announcement dates. The situation in the U.K., as shown in Table 3, appears to be similar, although it is seen that on Employment report announcements dates, positive excess returns are obtained for one- and ten-year maturity Gilts. For further analysis, Employment report and PPI announcements will be pooled for both countries, as they appear to have rather similar effects on the first and second moments of Germany government Bund and U.K. government Gilt returns.

More interesting effects are found when looking at the ECB announcement, BoE announcement and FOMC announcement dates. It can be seen that German bond market volatility is significantly higher on announcement dates, both when the ECB and FOMC release new information. However, while this (on average) leads to significant lower excess returns for ten- and 30-year maturity government bonds in the case of ECB announcements, no such effect appears to be the case for FOMC announcements. In fact, the opposite appears to be the case, with significantly higher excess returns than on non-announcement days. Excess returns of 0.005% are observed for one-year maturity Bunds and even higher announcement day excess returns of 0.188% for 30-year maturity Bunds. For the U.K., similar increased bond market volatility is observed on days BoE and FOMC announcement days, although significant positive excess on announcement days returns are only observed for the one-year maturity Gilt.

Although the summary statistics discussed before give some indication regarding the data used, it is by no means a sound statistical analysis. One of the things that is overlooked when assessing the summary statistics, is the day of the week on which macroeconomic news announcements are released. For example, U.S. employment statistics are released on Friday in 96.6% of the cases (in general the first Friday of the month), while PPI statistics are released around the middle of the month, with 50.6% being released on Friday, followed by 20.7% on Thursdays and 16.6% on Tuesdays. More complete information regarding the day-of-the-week on which macroeconomic announcements were released, please refer to Appendix B, Table 23-25. In order to explore the relation between risk and return more in depth, while taking into account day-of-the-week effects, there is now turned to simple OLS regressions.

Table 2: Summary statistics: German government Bund daily excess returns

XR_t is the daily continuously compounded excess return of the relevant constant maturity German government Bund security over the three-month EURIBOR rate. Returns are expressed in percent, i.e., multiplied by 100. The sample extends from January 1, 1999 to December 30, 2016.

	1-yr			10-yr			30-yr		
	XR_t	XR_t^2	$\|XR_t\|$	XR_t	XR_t^2	$\|XR_t\|$	XR_t	XR_t^2	$\|XR_t\|$
Full Sample ($N = 4696$)									
Mean	0.000	0.001[c]	0.019[c]	0.011[b]	0.140[c]	0.278[c]	0.025[b]	0.706[c]	0.604[c]
Std. Dev.	0.030	0.004	0.023	0.375	0.280	0.252	0.840	1.650	0.584
Min.	−0.257			−2.050			−4.608		
Max.	0.439			2.347			6.462		
Kurtosis	20.483			4.988			6.470		
Skewness	0.327			−0.176			−0.068		
Employment report announcement dates ($N = 217$)									
Mean	−0.001	0.001[c]	0.020[cy]	−0.012	0.140[c]	0.280[c]	0.014	0.728[c]	0.597[c]
Std. Dev.	0.031	0.003	0.023	0.375	0.294	0.248	0.855	1.831	0.611
PPI announcement dates ($N = 216$)									
Mean	0.000	0.001[c]	0.018[c]	−0.006	0.137[c]	0.281[c]	0.010	0.637[c]	0.616[c]
Std. Dev.	0.028	0.002	0.021	0.370	0.254	0.241	0.800	1.016	0.509
Pooled announcement dates ($N = 433$)									
Mean	0.000	0.001[c]	0.019[cx]	−0.009	0.138[c]	0.280[c]	0.012	0.683[c]	0.606[c]
Std. Dev.	0.029	0.002	0.022	0.372	0.275	0.244	0.827	1.481	0.562
ECB announcement dates ($N = 241$)									
Mean	−0.001	0.003[cz]	0.034[cz]	−0.027[x]	0.208[cz]	0.351[cz]	−0.072[x]	0.905[cx]	0.679[cy]
Std. Dev.	0.050	0.007	0.037	0.457	0.388	0.292	0.951	2.128	0.668
FOMC ($N = 149$)									
Mean	0.005[x]	0.001[cy]	0.025[cz]	0.065	0.291[cz]	0.418[cz]	0.188[by]	1.364[cz]	0.899[cz]
Std. Dev.	0.037	0.003	0.027	0.538	0.450	0.343	1.157	2.217	0.748
Non-announcement dates ($N = 3941$)									
Mean	0.000	0.001[c]	0.018[c]	0.014[b]	0.132[c]	0.269[c]	0.028[b]	0.676[c]	0.590[c]
Std. Dev.	0.028	0.004	0.021	0.363	0.261	0.244	0.822	1.603	0.572

[a]Different from zero at the 10% significance level (two-tailed for returns, one-tailed for variances)
[b]Different from zero at the 5% significance level (two-tailed for returns, one-tailed for variances)
[c]Different from zero at the 1% significance level (two-tailed for returns, one-tailed for variances)
[x]Different from non-announcement mean value at the 10% significance level (one-tailed)
[y]Different from non-announcement mean value at the 5% significance level (one-tailed)
[z]Different from non-announcement mean value at the 1% significance level (one-tailed)

Table 3: Summary statistics: U.K. government Gilt daily excess returns

XR_t is the daily continuously compounded excess return of the relevant constant maturity U.K. government Gilt security over the three-month LIBOR rate. Returns are expressed in percent, i.e., multiplied by 100. The sample extends from January 1, 1999 to December 30, 2016. The smaller number of observations is for the one-year maturity securities, where some values are missing.

	1-yr			10-yr			30-yr		
	XR_t	XR_t^2	$\|XR_t\|$	XR_t	XR_t^2	$\|XR_t\|$	XR_t	XR_t^2	$\|XR_t\|$
Full Sample ($N = 4682/4691$)									
Mean	0.000	0.002^{c}	0.026^{c}	0.009	0.174^{c}	0.310^{c}	0.015	0.575^{c}	0.562^{c}
Std. Dev.	0.042	0.012	0.033	0.418	0.360	0.280	0.759	1.239	0.510
Min.	−0.762			−2.036			−3.605		
Max.	0.641			2.715			5.905		
Kurtosis	46.011			5.274			5.620		
Skewness	−0.946			−0.021			0.175		
Employment report announcement dates ($N = 215/216$)									
Mean	0.005^{x}	0.002^{c}	0.031^{cz}	0.006	0.210^{cx}	0.333^{cx}	−0.008	0.594^{c}	0.578^{c}
Std. Dev.	0.048	0.008	0.037	0.459	0.438	0.315	0.773	1.127	0.511
PPI announcement dates ($N = 215/216$)									
Mean	−0.004	0.002^{c}	0.025^{c}	−0.017	0.192^{c}	0.316^{c}	−0.016	0.643^{c}	0.550^{c}
Std. Dev.	0.040	0.007	0.032	0.439	0.514	0.305	0.803	2.195	0.585
Pooled announcement dates ($N = 430/432$)									
Mean	0.001	0.002^{c}	0.028^{cy}	−0.005	0.201^{cx}	0.325^{cx}	−0.012	0.618^{c}	0.564^{c}
Std. Dev.	0.044	0.007	0.034	0.449	0.477	0.310	0.787	1.743	0.549
BoE announcement dates ($N = 215$)									
Mean	0.006^{ay}	0.002^{cy}	0.033^{cz}	0.006	0.228^{cy}	0.364^{cz}	−0.021	0.728^{cx}	0.636^{cy}
Std. Dev.	0.050	0.006	0.038	0.479	0.499	0.309	0.855	1.739	0.571
FOMC announcement dates ($N = 149$)									
Mean	0.004^{x}	0.002^{c}	0.032^{cz}	0.015	0.327^{cz}	0.448^{cz}	0.036	1.015^{cz}	0.770^{cz}
Std. Dev.	0.045	0.004	0.033	0.573	0.529	0.356	1.010	1.641	0.652
Non-announcement dates ($N = 3927/3934$)									
Mean	−0.001	0.002^{c}	0.025^{c}	0.010	0.165^{c}	0.302^{c}	0.019	0.552^{c}	0.552^{c}
Std. Dev.	0.041	0.013	0.033	0.406	0.328	0.272	0.743	1.123	0.497

[a]Different from zero at the 10% significance level (two-tailed for returns, one-tailed for variances)
[b]Different from zero at the 5% significance level (two-tailed for returns, one-tailed for variances)
[c]Different from zero at the 1% significance level (two-tailed for returns, one-tailed for variances)
[x]Different from non-announcement mean value at the 10% significance level (one-tailed)
[y]Different from non-announcement mean value at the 5% significance level (one-tailed)
[z]Different from non-announcement mean value at the 1% significance level (one-tailed)

4 Preliminary Analysis

In Table 4 the results of the OLS analysis are shown for the complete sample period for the U.S. Just as Jones et al. (1998) did, day-of-the-week and announcement indicator variables are used to document the volatility of daily excess returns. By doing so, day-of-the-week effects can be separated from announcement effects, as for example 57% of the announcements took place on Fridays, while only 35% of the Fridays had announcements (for the U.S.). Although only results for absolute excess returns will be discussed, similar conclusions apply to the analysis using squared excess returns.

In Table 4 it can be seen that there are moderate day-of-the-week effects, for U.S. Treasury return volatility. The lowest volatility is in general observed on Mondays, followed by an increase on Tuesday, slight dip on Wednesday followed by an increase on Thursday and Friday. Bond market volatility is generally highest on Friday and lowest on Monday, leading to an increasing trend over the week. However, a different trend is observed in the one-year maturity treasuries, where a U-curve in the volatility can be seen over the course of the week. This is in line with what Jones et al. (1998) find in their analysis covering 1979-1995. Bond market volatility patterns with respect to day-of-the-week effects for Germany and the U.K. (as shown in Tables 5 and 6) are similar to the pattern observed in the ten- and 30-year maturity Treasuries. These patterns concur somewhat with patterns observed in international stock markets, as shown by Kiymaz & Berument (2003).

Table 4 also shows that, after controlling for day-of-the-week effects, bond market volatility is significantly higher on announcement dates (both for pooled and FOMC announcements). Moreover, announcement day effects seem to be increasing with the Treasury maturity, with the absolute value of the excess returns increasing 0.016 for one-year Treasuries, 0.105 for ten-year maturities and 0.143 for 30-year maturities on pooled announcement dates. A similar increased announcement day volatility is observed for FOMC announcements. Using a weighted average bond market volatility (calculated using the release day percentages shown in Tables 23), it can be seen that announcement day volatility is on average between 13% and 44% higher (FOMC announcements, 30-year maturity and pooled announcements, one-year maturity respectively) than on non-announcement days for the U.S. These differences are highly statistically significant and corroborate earlier findings by Jones et al. (1998), although they did find larger magnitudes of volatility increases (more than 33%) across all maturities studied.

Next, the same OLS analysis is done using German and U.K. bond market data. When compared to the situation in the U.S., this yields somewhat different results regarding announcement day volatility. As can be seen in Tables 5 and 6, employment and PPI announcements (pooled for this analysis, separate analyses showed similar results) do not lead to a significant increased bond market volatility for any of the maturities studied. This is an interesting first finding, as it seems that the reaction of European bond markets differs significantly from the reaction of U.S. bond markets to comparable macroeconomic news announcements. In order to check whether this was not caused by the different sample periods used, the OLS regression analysis was repeated for the U.S., using only data from January 1, 1999 onwards. The results of this sensitivity analysis showed that the increase in announcement day volatility became less significant (which could also be partly due to the smaller sample size), but that the overall pattern of announcement day volatility did not change. This indicates that the reaction of the bond market to macroeconomic news announcements indeed differs between the U.S. and Europe.

A second interesting finding is that bond market volatility *does* significantly increase for the other type of announcements studied, being the target interest rate announcements made by the ECB and BoE. For the ten-year German government Bund, the absolute value of the excess returns was on average 0.068 higher on ECB announcement days, while a comparable announcement by the BoE on average increases the absolute value of the excess returns by 0.039 for the 10-year U.K. government Gilt. In general, the increase in central bank announcement day volatility appears to be more pronounced for the German Bund market than for the U.K. Gilt market, although still less pronounced than seen in the U.S. market.

A third interesting finding, related to the second finding, is that U.S. central bank announcements done by the FOMC appear to have a significant effect on European bond markets as well. Especially for long-term maturity government bonds, the FOMC announcements appear to have a larger effect than announcements made by the ECB and BoE. It can be seen that the absolute value of daily excess returns increases on average 0.292 for the 30-year German Bund (an increase of about 50% over the normal volatility levels) on days when the FOMC makes an announcement, while an increase of only 0.064 is observed when the ECB does so. Moreover, the volatility effect of an ECB announcement cannot be deemed significant at a 10% significance level, while the volatility increase seen on days when the FOMC makes an announcement is highly significant ($p < 0.0001$). A similar effect is seen in the U.K. market, where a BoE announcement leads to a (not significant) increase of volatility by 0.055 for the 30-year government Gilt, while a comparable FOMC announcement leads to an increase in the absolute value of excess returns of 0.196 ($p = 0.0003$).

Just as Jones et al. (1998) did, a one-day leading announcement dummy is included in the OLS regressions, in order to investigate the so called 'calm before the storm' effect. This effect, which is frequently heard of in the financial press, states that financial markets see less trade volume on days prior to important macroeconomic announcements. A typical headline seen on *Reuters* would be 'TREASURIES-Prices narrowly weaker as traders await Fed' (11/4/2009), implying that both returns and volatility measures are down. Moreover, this is also seen internationally, as for example this headline of *USA Today* reads 'World stock markets muted as investors await U.S. jobs report' (11/2/2016). If such an effect is present, we should be able to capture this effect using the one-day ahead dummy for FOMC announcements in the OLS regression using Bund or Gilt returns. However, a 'calm before the storm' effect is not found in the datasets used. As far as coefficient estimates for the one day ahead dummies are significant, they are positive, indicating increased trading leading up to the announcement date. However, this is rarely the case, as most of the time the one day ahead dummies coefficient estimates are not significantly different from zero.

Lastly, a one-day lagging announcement dummy is included, in order to see whether hypothesized shocks to volatility persist. If they do so, we would expect to see this in a lower than usual volatility on the day after the announcements, as captured by said dummy. Similar to what Jones et al. (1998) find, we see *lower* than average volatility in Tables 4, 5 and 6. However, none of the coefficient estimates for the one-day lagging announcement dummies is significant, so no (formal) inferences can be made from them. However, a preliminary inferences could be that there is no evidence for an increase in autocorrelated volatility following an one-time release of news. That being said, it is important to realize that a simple OLS regression analysis is not a sufficient method to model conditional heteroskedasticity, so no conclusions can be drawn from these Tables. More elaborate and statistically sound modeling of the conditional variance process will be done in Section 5, using an ARCH type of model specification.

A second question that Jones et al. considered in their 1998 paper was whether the theoretical link between risk and return could be shown in practice. In other words, they explored whether the increase in announcement day volatility, and hence increase in exposure to macroeconomic risk, coincided with an increase in expected excess returns. They found that announcement day risk premiums could be earned, as shown by a significant positive dummy variable when regressing the first moment of returns on days-of-the-week and an announcement day dummy. The same analysis using an OLS regression was done for the dataset used in this paper, however no significant risk premiums were found for either the U.S., Germany or the U.K., hence the results of these regressions are not shown.

In summary, the simple OLS regressions presented in this section confirmed earlier finding by Jones et al. (1998) regarding significant increased bond market volatility on announcement days for the U.S. market, although the effects seem to be less pronounced nowadays. Moreover, it was found that announcement day volatility effects are also present when studying announcements made by the FOMC, both in the U.S. and European markets. Interestingly is that FOMC announcement in the European market appear to have more impact than announcements made by the ECB and BoE, especially on the long-term securities (ten-year and 30-year). However, for short term German government Bunds and U.K. government Gilts, domestic central bank announcement appear to be more important.

Besides that, the release of one time information does not appear to give rise to autocorrelated volatility and no evidence for higher than average excess returns is found on announcement days. This is contrary to what one would expect from theory, as risk and (expected) return are in general related. However, it is seen that relation between volatility and government bond maturity is as expected. Tables 4-6 show that the returns of longer-maturity bonds show higher announcement day volatility, as well as higher average volatility on non-announcement days, when compared to shorter-maturity bonds. Next, I will turn to more statistically sound ways of modeling (announcement day) volatility. This enables jointly studying the second and first moments of excess returns, while explicitly modeling conditional heteroskedasticity.

Table 4: Treasury bond volatility by the day of week and event day

Mean values of the volatility of the daily continuously compounded excess return of the relevant constant maturity Treasury security over the three-month Treasury bill, estimated using an OLS regression with dummy variables for weekdays and announcement days. Returns are expressed in percent, i.e., multiplied by 100. Announcement is a dummy variable which equals one on the respective announcement dates. The sample extends from January 5, 1982 to December 30, 2016. Heteroskedasticity-consistent standard errors are given in parentheses (White, 1980).

	1-yr		10-yr		30-yr	
	Emp. & PPI	FOMC	Emp. & PPI	FOMC	Emp. & PPI	FOMC
Panel A: Absolute value of excess returns						
Monday	0.036	0.023	0.321	0.320	0.546	0.608
	(0.002)	(0.001)	(0.010)	(0.011)	(0.016)	(0.022)
Tuesday	0.033	0.022	0.346	0.365	0.589	0.691
	(0.001)	(0.001)	(0.008)	(0.011)	(0.013)	(0.021)
Wednesday	0.031	0.022	0.335	0.358	0.570	0.675
	(0.001)	(0.001)	(0.007)	(0.011)	(0.013)	(0.021)
Thursday	0.036	0.024	0.366	0.383	0.619	0.715
	(0.001)	(0.001)	(0.009)	(0.011)	(0.016)	(0.021)
Friday	0.038	0.026	0.383	0.416	0.634	0.756
	(0.001)	(0.001)	(0.008)	(0.011)	(0.015)	(0.020)
Announcement (t+1)	0.002	0.005	0.018	0.121	0.017	0.211
	(0.002)	(0.003)	(0.014)	(0.028)	(0.023)	(0.052)
Announcement	0.016	0.011	0.105	0.109	0.143	0.095
	(0.002)	(0.003)	(0.014)	(0.028)	(0.023)	(0.052)
Announcement (t-1)	−0.001	0.003	−0.003	−0.015	−0.011	−0.017
	(0.002)	(0.003)	(0.013)	(0.028)	(0.023)	(0.052)
Panel B: Squared excess returns						
Monday	0.005	0.002	0.196	0.187	0.574	0.696
	(0.001)	(0.000)	(0.018)	(0.019)	(0.044)	(0.063)
Tuesday	0.003	0.001	0.227	0.223	0.651	0.798
	(0.000)	(0.000)	(0.017)	(0.018)	(0.042)	(0.061)
Wednesday	0.002	0.001	0.205	0.223	0.625	0.828
	(0.000)	(0.000)	(0.013)	(0.018)	(0.036)	(0.061)
Thursday	0.004	0.002	0.236	0.250	0.716	0.904
	(0.000)	(0.000)	(0.013)	(0.018)	(0.051)	(0.061)
Friday	0.004	0.002	0.249	0.284	0.706	0.952
	(0.000)	(0.000)	(0.011)	(0.018)	(0.034)	(0.059)
Announcement (t+1)	0.000	0.001	0.034	0.148	0.035	0.554
	(0.001)	(0.001)	(0.036)	(0.045)	(0.063)	(0.152)
Announcement	0.003	0.002	0.140	0.258	0.254	0.327
	(0.001)	(0.001)	(0.022)	(0.045)	(0.055)	(0.152)
Announcement (t-1)	0.000	0.001	0.004	0.011	−0.038	−0.003
	(0.001)	(0.001)	(0.021)	(0.045)	(0.059)	(0.152)

5 Modeling time-varying variance

One of the most frequently used specifications to model financial asset return volatility is probably the GARCH(1,1) model originally described by Bollerslev (1986). Although the GARCH(1,1) model is not necessarily the best specification to model a return-generating process, it is able to replicate three stylized facts frequently observed in (daily) assets returns. These stylized facts, that were also observed in Section 3, are: **i)** the distribution of returns is not normal, **ii)** (almost) no significant autocorrelation in returns and **iii)** small, but very slowly declining autocorrelation in squared and absolute returns. Moreover, GARCH(1,1) models are able to adequately model volatility clustering (periods of relatively small and relatively large changes in prices alternate). Furthermore, theoretical results are available for quasi-maximum likelihood estimators of this model (e.g., Lumsdaine, 1996).

Table 5: Germany government bond return volatility by day of the week and event days

Mean values of the volatility of the daily continuously compounded excess return of the relevant constant maturity government bond over the three-month EURIBOR rate, estimated using and OLS regression with dummy variables for weekdays and announcement days. Returns are expressed in percent, i.e., multiplied by 100. Announcement is a dummy variable which equals one respectively on the PPI and employment report, ECB and FOMC announcement dates. The sample extends from January 1, 1999 to December 30, 2016. Heteroskedasticity-consistent standard errors are given in parentheses (White, 1980). Furthermore, resulting parameter estimates and standard errors for the 1-yr maturity using squared excess returns, are multiplied by 10 for readability.

	1-yr			10-yr			30-yr		
	Emp. & PPI	ECB	FOMC	Emp. & PPI	ECB	FOMC	Emp. & PPI	ECB	FOMC
Panel A: Absolute value of excess returns									
Monday	0.018	0.018	0.018	0.247	0.247	0.247	0.513	0.517	0.517
	(0.001)	(0.001)	(0.001)	(0.008)	(0.008)	(0.008)	(0.018)	(0.017)	(0.017)
Tuesday	0.017	0.017	0.017	0.265	0.265	0.266	0.603	0.606	0.606
	(0.001)	(0.001)	(0.001)	(0.008)	(0.008)	(0.008)	(0.020)	(0.020)	(0.020)
Wednesday	0.019	0.019	0.018	0.283	0.287	0.276	0.623	0.635	0.612
	(0.001)	(0.001)	(0.001)	(0.009)	(0.010)	(0.009)	(0.021)	(0.023)	(0.020)
Thursday	0.022	0.018	0.021	0.301	0.284	0.289	0.627	0.613	0.600
	(0.001)	(0.001)	(0.001)	(0.009)	(0.009)	(0.009)	(0.020)	(0.022)	(0.019)
Friday	0.018	0.016	0.018	0.293	0.277	0.292	0.636	0.620	0.632
	(0.001)	(0.001)	(0.001)	(0.009)	(0.009)	(0.009)	(0.021)	(0.023)	(0.020)
Announcement (t+1)	−0.001	0.007	0.001	0.002	0.066	0.006	0.012	0.079	0.062
	(0.001)	(0.002)	(0.002)	(0.013)	(0.021)	(0.023)	(0.029)	(0.044)	(0.057)
Announcement	0.000	0.016	0.005	−0.002	0.068	0.135	−0.006	0.064	0.292
	(0.001)	(0.003)	(0.002)	(0.012)	(0.021)	(0.029)	(0.029)	(0.050)	(0.063)
Announcement (t-1)	−0.002	−0.002	0.002	−0.001	−0.021	−0.016	0.031	−0.030	−0.017
	(0.001)	(0.001)	(0.002)	(0.014)	(0.019)	(0.019)	(0.033)	(0.047)	(0.048)
Panel B: Squared excess returns									
Monday	0.009	0.009	0.009	0.115	0.116	0.116	0.526	0.538	0.537
	(0.002)	(0.002)	(0.002)	(0.008)	(0.008)	(0.008)	(0.044)	(0.042)	(0.042)
Tuesday	0.007	0.007	0.007	0.129	0.129	0.130	0.722	0.727	0.731
	(0.001)	(0.001)	(0.001)	(0.009)	(0.009)	(0.010)	(0.056)	(0.058)	(0.060)
Wednesday	0.009	0.009	0.008	0.146	0.147	0.139	0.745	0.758	0.720
	(0.001)	(0.001)	(0.001)	(0.010)	(0.012)	(0.010)	(0.055)	(0.067)	(0.054)
Thursday	0.012	0.007	0.011	0.157	0.139	0.143	0.724	0.671	0.654
	(0.001)	(0.001)	(0.001)	(0.010)	(0.010)	(0.009)	(0.050)	(0.058)	(0.047)
Friday	0.008	0.006	0.007	0.155	0.140	0.153	0.778	0.755	0.757
	(0.001)	(0.001)	(0.001)	(0.010)	(0.010)	(0.009)	(0.071)	(0.080)	(0.065)
Announcement (t+1)	−0.002	0.005	0.001	−0.003	0.061	0.010	−0.008	0.098	0.210
	(0.001)	(0.002)	(0.003)	(0.013)	(0.024)	(0.029)	(0.083)	(0.123)	(0.176)
Announcement	0.000	0.018	0.004	−0.006	0.069	0.150	−0.037	0.230	0.678
	(0.001)	(0.004)	(0.003)	(0.014)	(0.028)	(0.038)	(0.076)	(0.162)	(0.186)
Announcement (t-1)	0.001	−0.002	0.001	0.010	−0.010	−0.022	0.120	−0.015	−0.062
	(0.004)	(0.002)	(0.002)	(0.019)	(0.023)	(0.017)	(0.099)	(0.147)	(0.130)

Table 6: U.K. government bond return volatility by day of the week and event days

Mean values of the volatility of the daily continuously compounded excess return of the relevant constant maturity government bond over the three-month LIBOR rate, estimated using and OLS regression with dummy variables for weekdays and announcement days. Returns are expressed in percent, i.e., multiplied by 100. Announcement is a dummy variable which equals one respectively on the PPI and employment report, ECB and FOMC announcement dates. The sample extends from January 1, 1999 to December 30, 2016. Heteroskedasticity-consistent standard errors are given in parentheses (White, 1980). Furthermore, resulting parameter estimates and standard errors for the 1-yr maturity using squared excess returns, are multiplied by 10 for readability.

	1-yr			10-yr			30-yr		
	Emp. & PPI	BoE	FOMC	Emp. & PPI	BoE	FOMC	Emp. & PPI	BoE	FOMC
Panel A: Absolute value of excess returns									
Monday	0.023	0.024	0.024	0.257	0.262	0.262	0.479	0.482	0.482
	(0.001)	(0.001)	(0.001)	(0.009)	(0.009)	(0.009)	(0.016)	(0.015)	(0.015)
Tuesday	0.025	0.026	0.026	0.290	0.298	0.300	0.552	0.556	0.559
	(0.001)	(0.001)	(0.001)	(0.010)	(0.009)	(0.009)	(0.018)	(0.016)	(0.016)
Wednesday	0.025	0.027	0.026	0.316	0.332	0.317	0.575	0.589	0.571
	(0.001)	(0.001)	(0.001)	(0.010)	(0.011)	(0.009)	(0.018)	(0.019)	(0.017)
Thursday	0.027	0.026	0.027	0.332	0.327	0.326	0.588	0.582	0.576
	(0.001)	(0.001)	(0.001)	(0.010)	(0.010)	(0.010)	(0.018)	(0.018)	(0.017)
Friday	0.025	0.025	0.025	0.326	0.320	0.332	0.597	0.590	0.595
	(0.001)	(0.001)	(0.001)	(0.010)	(0.010)	(0.010)	(0.019)	(0.020)	(0.019)
Announcement (t+1)	0.003	0.002	0.002	0.010	0.052	−0.019	0.027	0.039	0.019
	(0.002)	(0.002)	(0.002)	(0.015)	(0.023)	(0.021)	(0.025)	(0.044)	(0.040)
Announcement	0.003	0.007	0.005	0.024	0.039	0.126	0.014	0.055	0.196
	(0.002)	(0.003)	(0.003)	(0.016)	(0.023)	(0.030)	(0.029)	(0.043)	(0.055)
Announcement (t-1)	0.002	−0.002	−0.002	0.024	−0.046	−0.038	0.002	−0.044	−0.041
	(0.002)	(0.004)	(0.002)	(0.015)	(0.019)	(0.020)	(0.026)	(0.039)	(0.040)
Panel B: Squared excess returns									
Monday	0.015	0.016	0.016	0.132	0.138	0.138	0.442	0.449	0.450
	(0.003)	(0.002)	(0.002)	(0.010)	(0.010)	(0.010)	(0.032)	(0.031)	(0.031)
Tuesday	0.021	0.021	0.022	0.158	0.164	0.166	0.548	0.544	0.549
	(0.006)	(0.005)	(0.005)	(0.012)	(0.011)	(0.011)	(0.037)	(0.031)	(0.032)
Wednesday	0.021	0.017	0.022	0.176	0.198	0.179	0.578	0.603	0.573
	(0.007)	(0.002)	(0.007)	(0.012)	(0.014)	(0.011)	(0.042)	(0.040)	(0.036)
Thursday	0.017	0.015	0.017	0.193	0.184	0.183	0.618	0.582	0.579
	(0.002)	(0.002)	(0.002)	(0.013)	(0.012)	(0.013)	(0.044)	(0.038)	(0.042)
Friday	0.013	0.013	0.013	0.189	0.183	0.195	0.675	0.658	0.677
	(0.001)	(0.001)	(0.001)	(0.014)	(0.014)	(0.014)	(0.059)	(0.060)	(0.060)
Announcement (t+1)	0.003	−0.001	0.001	0.002	0.046	−0.031	−0.011	0.078	−0.049
	(0.004)	(0.002)	(0.003)	(0.019)	(0.036)	(0.022)	(0.050)	(0.153)	(0.077)
Announcement	0.001	0.010	0.002	0.037	0.045	0.145	0.074	0.148	0.435
	(0.005)	(0.005)	(0.005)	(0.025)	(0.036)	(0.045)	(0.093)	(0.123)	(0.138)
Announcement (t-1)	0.001	0.019	−0.009	0.015	−0.063	−0.047	−0.028	−0.048	−0.073
	(0.004)	(0.027)	(0.005)	(0.023)	(0.020)	(0.019)	(0.069)	(0.087)	(0.082)

5.1 Announcement day effects

For these reasons, an adjusted GARCH(1,1) specification is used as starting point to model daily bond returns, while incorporating (possible) announcement day effects. The baseline GARCH(1,1) model that will be used is based on the model of Jones et al. (1998), who in turn based their model on the procedure outlined in Andersen & Bollerslev (1997). That is, returns are assumed to follow the specification shown below (there will be adhered to the notation used by Jones et al. (1998)):

$$R_t = \mu + \theta I_t^A + \phi_1 R_{t-1} + s_t^{1/2}\epsilon_t \tag{1}$$

$$s_t = 1 + \delta_0 I_t^A \tag{2}$$

$$h_t = \omega + \alpha\epsilon_{t-1}^2 + \beta h_{t-1} \tag{3}$$

where I_t^A is the announcement indicator dummy variable, s_t is the volatility seasonal for time t, δ_0 measures the volatility effect of an announcement day on t and ϵ_t is a random variable with conditional mean zero and conditional variance h_t, independent of s_t. In other words, on non-announcement days the observed excess return innovations are distributed as a random variable with mean 0 and conditional variance h_t, while on announcement days, the observed excess return innovations are distributed as a random variable with mean 0 and conditional variance $(1 + \delta_0)h_t$. Intuitively, this means that the model described above can be seen as a mixture of two GARCH processes, resulting in two conditional variances processes, h_t and $g_t = (1 + \delta_0)h_t$, where g_t is always a scaled version of h_t. Non-announcement day errors are drawn from a distribution with h_t as conditional variance, while announcement day errors are drawn from a distribution with the scaled conditional variance g_t. However, using equation (3), it can be seen that h_t is modeled using the (partly non-observed) non-announcement day errors ϵ_t. This means that on announcement days, the observed error $s_t^{1/2}\epsilon_t$, first needs to be scaled in order to model h_{t+1}.

Just like Jones et al. (1998) did, we choose not to include day-of-the-week effects in the return specification or in the volatility seasonal, although this can be easily implemented. Nonetheless, it is found that day-of-the-week effects contribute little to the model (as measured by the increase in log likelihood value) and hence they have been left out. Moreover, we follow Jones et al. (1998) by including an autoregressive term ϕ for the first moment of excess returns, as we found in section 3 that there is a significant autocorrelation in the different types of government bond returns.

Estimation procedures used by Jones et al. (1998) will be followed, obtaining parameter estimates for equation (1)-(3) using quasi-maximum likelihood estimation. Starting values for the conditional variance process h_t are set equal to the unconditional variance, as suggested in Engle & Bollerslev (1986). The maximization algorithm is ran a minimum of 50 times, using randomized starting values for the parameter estimates, in order to minimize the chance of finding a local maximum for the log likelihood function. Normal standard errors will be used throughout the whole paper, contrary to the robust standard errors used by Jones et al. (1998). The normal standard errors are calculated by taking the square root of the diagonal elements of the covariance matrix, as estimated by taking the inverse of the negative Hessian, calculated in the final iteration of the optimization algorithm.

First, the results for the first moment of the returns will be discussed. The coefficient θ in equation (1) captures the change in mean excess return on announcement dates. Thus, a significant positive coefficient estimate of θ would implicate that a risk premium can be earned on announcement dates, as would be expected if increased announcement day volatility is found. Effects for the pooled announcement types studied by Jones et al. (1998) shall be discussed first. In Table 7, it can be seen that returns are on average 0.0015, 0.0373 and 0.0639 percent higher on announcement days, which is in line with what was seen in the summary statistics of Table 1.

When comparing these figures to the results found in the paper by Jones et al. (1998), it is seen that the magnitude of the risk premium has decreased by a factor two, as the original paper found risk premiums of 0.09 and 0.012 percentage points for the ten- and 30-year Treasuries. In addition, a small positive autoregressive term ϕ is observed, that is significant for the ten-year and 30-year maturity series at the 5% level.

When turning to the second moment of the returns, the coefficient estimates of δ_0 are of interest, as they capture a (possible) volatility increase on announcement days. It is seen that employment report and PPI announcements have a highly statistically significant impact on Treasury volatility on the day of announcement, with volatility increasing between 77% and 136% for the different Treasury maturities. These results are in line with the results found by Jones et al. (1998), although slightly lower figures are found for the longer-term maturities than they did.

In order to compare the situation in the U.S. with the situation in Germany and the U.K., the same model (using German and English employment and PPI announcements respectively) has been estimated for those countries, for which the results can be found in Tables 9 and 11. Contrary to the situation in the U.S., *lower* average excess returns are found on announcement days in Germany and the U.K. Parameter estimates for θ of -0.0169 and -0.0159 percentage points are seen for German and English ten-year maturity government bonds, although these are not significantly different from zero. Another contrast with the situation in the U.S. is the (non-existing) effect of German and English pooled macroeconomic announcements on the volatility of excess returns. As can be seen in Tables 9 and 11, there is no statistical evidence that such an effect is present. Small positive coefficient estimates for δ_0 are observed for one- and ten-year Bunds and Gilts, while small negative coefficient estimates are observed for 30-year maturities. Yet, standard errors are relatively large, making estimates statistically indistinguishable from zero. At the same time, estimates for α and β are comparable to what was seen in the U.S. market, indicating that the conditional variance process does not differ between countries.

Next, there is turned to the other types of announcements studied, being the central bank announcements made by the FOMC, the ECB and the BoE. In Table 8 results are shown for the U.S. using FOMC announcements, in Table 10 results are shown for Germany using ECB announcements and in Table 12 results are shown for the U.K. using BoE announcements. As for the U.S., parameter estimates for θ are insignificant for all maturities studied, as can be inferred from the standard errors used.

However, it is found that announcements have a significant effect on the volatility of the excess returns, increasing volatility for the one-year Treasury by on average 85%, as can be seen in Table 8. For Germany, similar announcements made by the ECB appear to have an even more pronounced impact on excess returns volatility of short term Bunds, as can be seen from the parameter estimate of 1.6851 (meaning a 169% increase in volatility) for δ_0 in Table 10. For ten-year maturity Bunds this effect is smaller and more in line with the situation in the U.S. However, for the longest maturity studied, the 30-year government Bunds, an average announcement day volatility increase of 25% is seen. This is larger than the effect of a FOMC announcement on the 30-year Treasuries, where we only find a much smaller and statistically insignificant effect. For the UK, announcements made by the BoE have a less distinct effect on government Gilt excess return volatility, with δ_0 estimates showing that volatility increases are in the range of 24-32% for the different maturities studied. Furthermore, estimates for θ are all insignificant, indicating that average expected excess returns on days when the ECB and BoE make announcements do not differ from expected returns on non-announcement days.

5.2 Volatility persistence

In order to further investigate the persistence of the announcement day volatility shocks, as observed in our first GARCH model, I follow the original paper of Jones et al. (1998) and continue by extending the benchmark GARCH(1,1) model. This is done to study whether

Table 7: Benchmark GARCH(1,1) model of daily U.S. bond returns with pooled announcements

Quasi-maximum likelihood estimates of the model

1. $R_t = \mu + \theta I_t^A + \phi_1 R_{t-1} + s_t^{1/2} \epsilon_t$
2. $s_t = 1 + \delta_0 I_t^A$
3. $h_t = \omega + \alpha \epsilon_{t-1}^2 + \beta h_{t-1}$

where R_t is the daily continuously compounded excess return of the relevant constant maturity Treasury security over the three-month Treasury bill, ϵ_t is an independent random variable with conditional mean zero and conditional variance h_t, and I_t^A is an indicator variable equal to one on employment or PPI announcement days (Table 7) and FOMC announcements respectively (Table 8). Returns are expressed in percent, i.e., multiplied by 100. The sample extends from January 5, 1982 to December 31, 2016. Standard errors are given in parentheses.

	First moment parameters				Second moment parameters		
	1-yr	10-yr	30-yr		1-yr	10-yr	30-yr
μ	0.0001	0.0115	0.0203	ω	0.0000	0.0026	0.0047
	(0.0002)	(0.0047)	(0.0077)		(0.0000)	(0.0006)	(0.0010)
θ	0.0015	0.0373	0.0639	α	0.0779	0.0468	0.0397
	(0.0012)	(0.0207)	(0.0327)		(0.0025)	(0.0044)	(0.0035)
ϕ	0.0153	0.0456	0.0215	β	0.9300	0.9414	0.9529
	(0.0113)	(0.0109)	(0.0109)		(0.0014)	(0.0061)	(0.0043)
				δ_0	1.3681	0.9576	0.7720
					(0.1157)	(0.0991)	(0.0901)
Log L	15788.73	−5563.07	−10002.95				

Table 8: Benchmark GARCH(1,1) model of daily U.S. bond returns with FOMC announcements

	First moment parameters				Second moment parameters		
	1-yr	10-yr	30-yr		1-yr	10-yr	30-yr
μ	0.0001	0.0093	0.0169	ω	0.0000	0.0015	0.0069
	(0.0002)	(0.0066)	(0.0122)		(0.0000)	(0.0007)	(0.0019)
θ	0.0011	−0.0016	−0.0272	α	0.0795	0.0351	0.0382
	(0.0019)	(0.0463)	(0.0710)		(0.0033)	(0.0054)	(0.0047)
ϕ	−0.0398	0.0043	−0.0096	β	0.9273	0.9585	0.9536
	(0.0162)	(0.0152)	(0.0154)		(0.0019)	(0.0075)	(0.0059)
				δ_0	0.8505	0.6157	0.0941
					(0.2381)	(0.1855)	(0.1251)
Log L	10016.88	−2955.04	−5738.75				

Table 9: GARCH(1,1) model of daily German bond returns with pooled announcements

Quasi-maximum likelihood estimates of the model

1. $R_t = \mu + \theta I_t^A + \phi_1 R_{t-1} + s_t^{1/2} \epsilon_t$
2. $s_t = 1 + \delta_0 I_t^A$
3. $h_t = \omega + \alpha \epsilon_{t-1}^2 + \beta h_{t-1}$

where R_t is the daily continuously compounded excess return of the relevant constant maturity bond security over the three-month EURIBOR rate, ϵ_t is an independent random variable with conditional mean zero and conditional variance h_t, and I_t^A is an indicator variable equal to one on employment or PPI announcement days (Table 9) and FOMC announcements respectively (Table 15). Returns are expressed in percent, i.e., multiplied by 100. The sample extends from January 4, 1999 to December 31, 2016. Standard errors are given in parentheses.

	First moment parameters				Second moment parameters		
	1-yr	10-yr	30-yr		1-yr	10-yr	30-yr
μ	−0.0007	0.0119	0.0190	ω	0.0000	0.0007	0.0032
	(0.0003)	(0.0050)	(0.0099)		(0.0000)	(0.0006)	(0.0011)
θ	0.0008	−0.0169	−0.0243	α	0.0703	0.0323	0.0521
	(0.0011)	(0.0167)	(0.0325)		(0.0031)	(0.0069)	(0.0065)
ϕ	0.0714	0.0198	0.0397	β	0.9269	0.9629	0.9441
	(0.0166)	(0.0150)	(0.0152)		(0.0016)	(0.0106)	(0.0071)
				δ_0	0.1493	0.0192	−0.0060
					(0.0995)	(0.0713)	(0.0677)
Log L	10730.50	−1741.39	−5208.01				

Table 10: GARCH(1,1) model of daily German bond returns with ECB announcements

	First moment parameters				Second moment parameters		
	1-yr	10-yr	30-yr		1-yr	10-yr	30-yr
μ	−0.0006	0.0127	0.0213	ω	0.0000	0.0007	0.0032
	(0.0003)	(0.0049)	(0.0097)		(0.0000)	(0.0006)	(0.0011)
θ	−0.0020	−0.0423	−0.0808	α	0.0712	0.0329	0.0526
	(0.0022)	(0.0253)	(0.0454)		(0.0031)	(0.0076)	(0.0066)
ϕ	0.0728	0.0163	0.0380	β	0.9246	0.9621	0.9435
	(0.0160)	(0.0149)	(0.0152)		(0.0016)	(0.0117)	(0.0072)
				δ_0	1.6851	0.4816	0.2500
					(0.2340)	(0.1351)	(0.1120)
Log L	10827.24	−1729.37	−5203.02				

Table 11: GARCH(1,1) model of daily U.K. Gilt returns with pooled announcements

Quasi-maximum likelihood estimates of the model

1. $R_t = \mu + \theta I_t^A + \phi_1 R_{t-1} + s_t^{1/2}\epsilon_t$
2. $s_t = 1 + \delta_0 I_t^A$
3. $h_t = \omega + \alpha\epsilon_{t-1}^2 + \beta h_{t-1}$

where R_t is the daily continuously compounded excess return of the relevant constant maturity U.K. government bond security over the three-month LIBOR rate, ϵ_t is an independent random variable with conditional mean zero and conditional variance h_t, and I_t^A is an indicator variable equal to one on employment or PPI announcement days (Table 11), BoE announcements days (Table 12 or FOMC announcement days respectively (Table 16). Returns are expressed in percent, i.e., multiplied by 100. The sample extends from January 4, 1999 to December 31, 2016. Standard errors are given in parentheses.

	First moment parameters				Second moment parameters		
	1-yr	10-yr	30-yr		1-yr	10-yr	30-yr
μ	−0.0009	0.0061	0.0121	ω	0.0000	0.0005	0.0027
	(0.0004)	(0.0056)	(0.0100)		(0.0000)	(0.0005)	(0.0010)
θ	0.0003	−0.0159	−0.0383	α	0.0285	0.0256	0.0362
	(0.0014)	(0.0190)	(0.0330)		(0.0011)	(0.0060)	(0.0048)
ϕ	0.0602	0.0165	0.0383	β	0.9748	0.9720	0.9594
	(0.0168)	(0.0150)	(0.0150)		(0.0006)	(0.0084)	(0.0057)
				δ_0	0.0027	0.0506	−0.0162
					(0.0904)	(0.0744)	(0.0679)
Log L	9018.25	−2282.33	−5020.74				

Table 12: GARCH(1,1) model of daily U.K Gilt returns with BoE announcements

	First moment parameters				Second moment parameters		
	1-yr	10-yr	30-yr		1-yr	10-yr	30-yr
μ	−0.0011	0.0048	0.0105	ω	0.0000	0.0005	0.0027
	(0.0004)	(0.0055)	(0.0097)		(0.0000)	(0.0005)	(0.0010)
θ	0.0035	−0.0050	−0.0434	α	0.0287	0.0256	0.0360
	(0.0022)	(0.0286)	(0.0505)		(0.0011)	(0.0061)	(0.0048)
ϕ	−0.0179	0.0182	0.0395	β	0.9749	0.9720	0.9595
	(0.0167)	(0.0150)	(0.0150)		(0.0006)	(0.0086)	(0.0057)
				δ_0	0.3213	0.2656	0.2436
					(0.1259)	(0.1233)	(0.1190)
Log L	9035.55	−2279.73	−5018.17				

announcement day shocks are more or less persistent than non-announcement day shocks. Under the assumption that markets immediately incorporate new information into prices (the Efficient Market Hypothesis), one would expect to find that announcement day shocks are less persistent than non-announcement day shocks (as was the result of the original paper by Jones et al. (1998)). This leads to the following model, consisting of equation (1) and the following adjusted specification for the volatility seasonal and the conditional variance:

$$s_t = (1 + \delta_0 I_t^A)(1 + \delta_1 I_{t-1}^A) \tag{2'}$$

$$h_t = \omega + (\alpha_0 + \alpha_A I_{t-1}^A)\epsilon_{t-1}^2 + (\beta_0 + \beta_A I_{t-1}^A)h_{t-1} \tag{3'}$$

In the specification above the conditional variance is a regime-switching GARCH process, in which regime shifts occurs (deterministically) at dates on which macroeconomic news announcements are made. When comparing to the original specification, three more parameters are added, being δ_1, α_A and β_A. Using δ_1, a deterministic change in volatility on the day after an announcement can be modeled. Under the assumption that markets are efficient, this coefficient is expected to be statistically indistinguishable from zero. Furthermore, using α_A, announcement day shocks are allowed to influence the conditional variance process differently than non-announcement day shocks (that feed into the conditional variance process through α_0). If announcement day shocks are truly one-time events, containing a one-time piece of information, α_A is expected to be equal to $-\alpha_0$, indicating that announcement day shocks do not persist in the conditional variance process. Last of all, the model enables lagged conditional volatility to influence current conditional volatility differently on the day after an announcement, through β_A. This 'extended' specification is more flexible with respect to the effects of announcement and non-announcement shocks on volatility. However, if $\delta_1 = \alpha_A = \beta_A = 0$, equation 2' and 3', reduce to equation 2 and 3, thereby nesting the original model.

In order to conserve space, results for the extended model specification are shown in the Appendix, in Tables 17 to 22. First of all, it is tested whether the new model specification is better than the benchmark model. Because the benchmark model is nested in the new model specification, this can be done using the Wald test with the joint null hypothesis that $\delta_1 = 0$, $\alpha_A = 0$ and $\beta_A = 0$. As the joint null hypothesis consists of three restrictions, the 5% critical value is taken from the $\chi^2(3)$ distribution, resulting in a 5% critical value of 7.81.

For the U.S., using pooled announcements (results shown in Table 17), the null hypothesis is rejected, with Wald test statistic values of 21.90, 13.95 and 12.57 for the one-, ten- and 30-year Treasuries respectively. The same result is found when looking at the U.S. market, but then using FOMC announcements (results shown in Table 18, with Wald test statistic values of 12.08, 12.78 and 21.39. However, it can be seen that the rejection of the joint null hypothesis for both types of announcements has different reasons. For the pooled announcements, results are quite similar to what Jones et al. (1998) found in their paper, at least for the ten- and 30-year Treasuries. In Table 17, negative coefficient estimates for α_A can be observed, not significantly different from $-\alpha_0$. The Wald test for the hypothesis $\alpha_A = -\alpha_0$ shows test statistic values of 1.21 and 1.04 for the ten- and 30-year Treasuries, well below the 5% critical value of 3.84 from the $\chi^2(1)$ distribution. Furthermore, it can be seen in Table 17 that coefficient estimates for β_A and δ_1 are both statistically indistinguishable from zero.

This implies that announcement day shocks to volatility do not persist. Moreover, on days following an announcement, conditional variance decays on a 'normal' rate β_0 and the average unconditional variance is not higher than on non-announcement days. However, for the one-year maturity Treasuries, it is observed that announcement day shocks do persist ($\alpha_A = -\alpha_0$ is strongly rejected with a test value of 36.83), conditional variance actually decays somewhat faster than normally the case following an announcement day ($\beta_A = -0.1160$) and the day after an announcement sees an increased unconditional variance, as can be seen from the fact that $\delta_1 = 0.3156$.

For the other type of announcements studied for the U.S., being the FOMC announcements, results are shown in Table 18. From this Table it follows that macroeconomic news announcements done by the FOMC have similar effects as the pooled employment and PPI announcements have on the excess returns of the shortest maturity, the one-year Treasury bond. However, for the ten- and 30-year Treasury bonds, it is seen that FOMC announcements actually lead to volatility *persistence* as significant positive coefficient estimates are obtained for δ_1. This effect appears to be stronger for longer maturities, as it is observed in Table 18 that for the 30-year bond $\delta_1 > \delta_0$ and β_A is significant with a positive value of 0.2053. One possible explanation for this finding is that longer-term maturity government bonds are primarily influenced by expectations of future interest rates, the so called 'path' surprise of the central bank announcements.

Next I'll turn to the extended model results for Germany and the U.K, as shown in Tables 19 to 22. For the pooled employment and PPI announcements, it is observed that for both Germany and the U.K., the new model adds little. In fact, for the ten- and 30-year maturity Bunds and Gilts, the Wald test shows that the null-hypothesis $\delta_1 = 0$, $\alpha_A = 0$ and $\beta_A = 0$ is not rejected, with test statistic values of 0.52 and 0.06 for Germany's ten- and 30-year Bunds and values of 5.49 and 4.36 for U.K.'s ten- and 30-year Gilts. For the shorter-term one-year maturity Bund and Gilt, the Wald test shows that the null-hypothesis is not rejected, indicating that there might be some degree of conditional volatility persistence on the days following an announcement. However, as for both models the impact of announcements on volatility is statistically insignificant (as seen from the standard errors of δ_0 in Tables 19 and 21), no further model interpretation will be done.

The last extended models to interpret are the extended models using the central bank announcements, as made by the ECB and BoE and shown in Table 20 and 22. For Germany it is seen that a similar Wald test as done before, for the null hypothesis that $\delta_1 = 0$, $\alpha_A = 0$ and $\beta_A = 0$, yield test statistics of 0.63, 22.42 and 21.32 (with a 5% critical $\chi^2(3)$ value of 7.81) for the one-, ten- and 30-year government Bund securities. Further inspection as to why the null hypothesis is not rejected for the shortest term maturity, shows that this is due to large standard errors. For the longer-term maturities significant positive coefficient estimates for δ_1 are found, indicating a higher than usual unconditional variance on the day following an ECB announcement in Germany. Estimates for both α_A and β_A are insignificant and hence cannot be interpreted.

For the U.K., the null hypothesis that $\delta_1 = 0$, $\alpha_A = 0$ and $\beta_A = 0$ is also tested using a Wald-test, yielding test statistic values of 33.33, 64.91 and 5.71 for the one-, ten- and 30-year maturity Gilts. As was the case with the one-year maturity Bund, large standard errors are the main reason for not rejecting the null hypothesis for the 30-year maturity Gilt. Furthermore, in Table 22, it is observed that there is some kind of shock persistence following BoE announcements, as shown from the significant positive coefficient for α_A for the one-year maturity Gilt and the significant positive coefficient for δ_1 for the ten-year maturity Gilt. However, for the one-year maturity Gilt the coefficient for δ_1 is not significantly different from zero and for the ten-year maturity Gilt the coefficient for $alpha_a$ is negative. Hence, the results are interpreted as showing some kind of volatility persistence, but more research to as the exact ways this process takes place is needed.

In summary, the research as done by Jones et al. (1998) into volatility persistence is replicated, and extended by including different countries and also central bank announcements. Similar results are found as Jones et al. (1998) did for the pooled employment and PPI announcements on US Treasuries, showing significant positive excess return premiums on announcement days and a significant announcement day effect on the volatility of said returns. However, using the same type of announcements, I do not find significant effects on the first or second moment of excess returns for Germany and the U.K. Turning to announcements made by the central banks (the FOMC, ECB and BoE respectively), I find significant announcement day effects on the second moment of the excess returns for all countries studied.

Next to that, volatility persistence is studied on the days following a macroeconomic announcement, by implementing a more flexible model specification, as proposed by Jones et al. (1998). Again I am able to replicate earlier results for pooled announcements in the U.S., showing that there is no persistence of announcement day shocks to volatility at all. Yet, for the other type of announcements studied (as done by the central banks), I *do* find that announcement day shocks to volatility persist, although results are mixed for different maturities studied and hence further research is needed.

5.3 Comparing central bank announcements

The third point of my analysis focuses on the relative importance of central bank announcements. The question central questions that I would like to answer is: "Which central bank's policy is seen as most influential by the financial markets?" Preliminary analysis (shown in the appendix in Table 15 and 16) showed that FOMC announcements do not only have significant effects on U.S. Treasury excess return volatility, but also on German Bund and U.K. Gilt excess return volatility.

This question will be investigated more in-depth by estimating a third model, obtained by adjusting the original model as proposed by Jones et al. (1998), resulting in the following model specification:

$$R_t = \mu + \theta_1 I_t^{A1} + \theta_2 I_t^{A2} + \phi_1 R_{t-1} + s_t^{1/2}\epsilon_t \tag{4}$$

$$s_t = (1 + \delta_1 I_t^{A1})(1 + \delta_2 I_t^{A2}) \tag{5}$$

$$h_t = \omega + \alpha\epsilon_{t-1}^2 + \beta h_{t-1} \tag{3}$$

where I_t^{A1} and I_t^{A2} are announcement indicator dummy variables, s_t is the volatility seasonal for time t, and δ_1 and δ_2 measure the volatility effect of announcements on day on t, and ϵ_t is a random variable with conditional mean zero and conditional variance h_t.

As literature shows that historically the FOMC was the most influential central bank for financial markets, it is studied whether this is still the case for German government Bunds and U.K. government Gilts. This is done by taking both the countries own central bank's announcements (the ECB for Germany and the BoE for the U.K.) and the FOMC announcements and modeling excess returns using the specification described above. In the model domestic central bank's announcements are taken as the first announcement (with possible announcement effects being captured by the obtained coefficients for θ_1 and δ_1), and the FOMC announcements as the second announcements (captured by θ_2 and δ_2).

The results for this analysis are shown in Tables 13 and 14. It can be seen that for both countries, the domestic central bank's announcements have more impact on excess return volatility than FOMC announcements do, for the one-year maturity government bonds. This effect is slightly more distinct for the U.K. than for Germany, as seen from the ratio $\frac{\delta_1}{\delta_2}$ for both countries. Furthermore, it can be seen that for longer-term maturities, FOMC announcements have more impact on excess return volatility than ECB or BoE announcements. For Germany the distinction between the influence of ECB and FOMC announcements becomes stronger for longer maturities (as seen from the ratio $\frac{\delta_1}{\delta_2}$), for the U.K. this does not appear to be the case, as can be seen when comparing results for ten- and 30-year maturity Gilts.

However, these results have to be examined with some consideration. One ground to do so is the timing of the different central bank's announcements. This timing is important for two reasons. First of all the informational value of the 'first' central bank's announcement to change the target interest rate can be considered larger than the informational value of subsequent announcements done by other central banks, given the ongoing globalization of trade and investment. However, when analyzing our sample, no systematic lead/lag relations were found between announcements made by different central banks. Secondly, the frequency of central bank's announcements influences the informational value of said announcements, hence (possibly) influencing the magnitude of the announcement day effect on return volatility. In my sample it can be seen that the ECB and BoE make more frequent announcements, which might decrease the informational value of each independent announcement. Thirdly, if the timing of central bank's announcements coincides (or differs only a few days), it could be that volatility persistence effects interfere with announcement day effects. When analyzing the sample used, we find that ECB (BoE respectively) and FOMC announcements coincide on about 6-10% of the announcement days. However, the ECB and BoE announcement coincide on about 55% of the announcement days, hence rendering comparison between those announcements more difficult.

6 Conclusion

Understanding the relation between asset prices and risk has been at the center of interest of financial economics for quite some time. Predicting asset prices using an asset's sensitivity to (non-diversifiable) systematic risk has seen a lot of interest, following the formulation of the Capital Asset Pricing Model by Sharpe (1964). This has been the basis for a lot of academic work concerning asset prices and how these can be predicted, based on the asset's exposure to different types of economic risk (Fama & French, 1992). At the center of this research has been the question of how 'news' can be linked to risk. However, even linking asset returns in hindsight to news events has proven difficult, especially for volatility measures (Mitchell & Mulherin, 1994; Berry & Howe, 2004, e.g.,).

However, linking 'news' to volatility is exactly what I tried to do in this paper. But, instead of taking all asset price movements and explaining them by some kind of 'news' measure, I took a subset of periodic, preannounced macroeconomic news announcements and studied what their effect is on asset price volatility. More specifically, the effect of PPI, employment and central bank announcements on government bonds returns in the U.S., Germany and the U.K are examined. In order to answer my research question *"What is the effect of macroeconomic news announcements on bond market volatility, and do these effects differ for various countries and announcement types?"*, three hypotheses were formulated.

The first hypothesis stated that "Macroeconomic news announcements have a positive effect on government bond volatility and a significant higher excess returns can be earned on announcement days". Looking at the data, this first hypothesis must be rejected. Although it is found that macroeconomic news announcements (in general) have a positive effect on the volatility of excess returns, I am not able to statically show that this coincides with significant higher excess returns in any of the countries studied. Furthermore, while it is seen that all types of announcements have a positive effect on Treasury bond volatility, a positive effect on German Bund and English Gilt volatility is only found for announcements done by the ECB and BoE.

The second hypothesis stated that "Macroeconomic news is immediately incorporated in bond prices and announcement day shocks to volatility do not persist". From the data it is found that this is indeed the case for PPI and employment announcements in the U.S. When studying FOMC announcements it was seen that results differ over the maturities studied. While information appears to be immediately incorporated in short term Treasuries, this is not the case for longer-term maturities, as can be seen from an increasing degree of volatility persistence over the maturities studied. Furthermore, evidence is found that announcement day shocks, and subsequent volatility effects, tend to persist the days following an announcement, when studying central bank's announcements in Europe.

The last hypothesis stated that "Central bank announcements done by the FOMC have more impact on bond market volatility than similar announcements done by the ECB or BoE".

Table 13: GARCH(1,1) model of daily German Bund returns with ECB and FOMC announcements

Quasi-maximum likelihood estimates of the model

1. $R_t = \mu + \theta_1 I_t^{A1} + \theta_2 I_t^{A2} + \phi_1 R_{t-1} + s_t^{1/2}\epsilon_t$
2. $s_t = (1 + \delta_1 I_t^{A1})(1 + \delta_2 I_t^{A2})$
3. $h_t = \omega + \alpha\epsilon_{t-1}^2 + \beta h_{t-1}$

where R_t is the daily continuously compounded excess return of the relevant constant maturity German government bond security over the three-month EURIBOR rate, θ_1, I_t^{A1} and δ_1 correspond to ECB (BoE respectively) announcements, θ_2, I_t^{A2} and δ_2 to FOMC announcements. Returns are expressed in percent, i.e., multiplied by 100. The sample extends from January 4, 1999 to December 31, 2016. Standard errors are given in parentheses.

	First moment parameters				Second moment parameters		
	1-yr	10-yr	30-yr		1-yr	10-yr	30-yr
μ	−0.0006	0.0126	0.0229	ω	0.0000	0.0007	0.0031
	(0.0003)	(0.0049)	(0.0096)		(0.0000)	(0.0008)	(0.0011)
θ_1	−0.0027	−0.0505	−0.0902	α	0.0696	0.0313	0.0487
	(0.0016)	(0.0255)	(0.0458)		(0.0031)	(0.0093)	(0.0063)
θ_2	0.0009	−0.0116	−0.1111	β	0.9220	0.9612	0.9443
	(0.0021)	(0.0363)	(0.0515)		(0.0017)	(0.0156)	(0.0072)
ϕ	0.0566	0.0242	0.0513	δ_1	0.6087	0.5207	0.2766
	(0.0167)	(0.0151)	(0.0154)		(0.0945)	(0.1437)	(0.1201)
				δ_2	0.5447	0.9714	0.8504
					(0.1918)	(0.2338)	(0.2214)
Log L	10813.02	−1707.95	−5182.64				

Table 14: GARCH(1,1) model of daily U.K Gilt returns with BoE and FOMC announcements

	First moment parameters				Second moment parameters		
	1-yr	10-yr	30-yr		1-yr	10-yr	30-yr
μ	−0.0010	0.0059	0.0120	ω	0.0000	0.0004	0.0027
	(0.0004)	(0.0055)	(0.0097)		(0.0000)	(0.0006)	(0.0010)
θ_1	0.0037	−0.0052	−0.0489	α	0.0278	0.0239	0.0341
	(0.0022)	(0.0279)	(0.0501)		(0.0011)	(0.0074)	(0.0046)
θ_2	−0.0006	−0.0379	−0.0720	β	0.9751	0.9728	0.9598
	(0.0026)	(0.0386)	(0.0513)		(0.0006)	(0.0111)	(0.0057)
ϕ	0.0070	0.0241	0.0481	δ_1	0.3023	0.2066	0.2354
	(0.0168)	(0.0151)	(0.0151)		(0.1182)	(0.1196)	(0.1222)
				δ_2	0.1425	0.9190	0.7630
					(0.1483)	(0.2295)	(0.2085)
Log L	9036.90	−2261.10	−5002.51				

I show that this is indeed the case for longer-term maturities studied in Germany and the U.K. However, it is found that for shorter term-maturities, announcements done by the ECB and BoE have a more pronounced effect on volatility and hence I conclude that those announcements are more important in the short end of the yield curve.

Using my dataset, I am able to replicate the findings by Jones et al. (1998), except for their finding that "the predictable risk that bonds bear on announcement days is compensated with higher expected excess returns" (p. 334). Several possible explanations can be given for this. First of all, it could be that the increasing speed of the trading process makes it more difficult to statically show higher expected returns while studying daily frequency data. This warrants further research, using higher frequency data. Secondly, theory states that higher expected returns are only expected when information is not already priced in. This implies that one should look at the 'surprise' component of macroeconomic announcements, which was not done in this paper. This is another limitation of this research and provides one more starting point for further research (e.g., Kuttner, 2001). The third limitation is on a technical level. One can only reliably test higher expected returns if the return process is modeled correctly. It could be that this is not the case for the models I use, thereby influencing my findings. For this research, models as proposed by Jones et al. (1998) have been used, however it could be that those are not sufficient for other types of announcements studied than the ones included in the original paper. Moreover, while the original paper by Jones et al. (1998) used robust standard errors, as proposed by Bollerslev & Wooldridge (1992), I ran into scaling problems when trying to do so. Therefore, I was forced to use 'normal' standard errors. While these are consistent, they are not efficient when heteroskedasticity is present, therefore biasing my results against finding significance. Next to those limitations, the limitations applicable to the original paper of Jones et al. (1998) also apply to this research. Only a small subset of available news events is taken into account in this paper, exclusively consisting of public information. However, the announcement types included in this paper were found to be especially influential (de Goeij & Marquering, 2006).

The main implications of my paper are twofold. The first implication, as also mentioned by Jones et al. (1998), has to do with the modeling of conditional volatility in asset returns. This paper confirms earlier findings that the persistence of returns shocks can be related to their sources. In order to model conditional variance correctly, this finding should be incorporated, otherwise inferior estimates may be obtained. The second implication has to do with policy making, especially as done by central bank's. The finding that announcements done by the FOMC are more influential than announcements done by the ECB, could be reason for the latter to review announcement procedures or the transparency in the decision making process. However, it could also be that influence effects are simply due to timing, which could be reason for the ECB and BoE to review the pace in which decisions are made and consequently communicated to financial markets (Andersson et al., 2009).

Consequently, much remains for future research. This could be done along a multitude of dimensions, including but not limited to: studying higher frequency data, including more types of announcements, correcting for the 'surprise' effect of announcements and further investigating integration and interdependence between different financial markets and ultimately different economies.

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Appendix A: Calculations of Returns

In this appendix further elaboration is given upon the calculation of the bond returns, using the methodology and notation described by Jones et al. (1998) and de Goeij & Marquering (2006).

The U.S. government Treasury bonds have begin-of-period prices equal to its face value, as the bonds have semi-annual coupon payments equal to half the stated coupon yield. The end-of-period price is calculated using the next day's yield, from which the $\#hd$-return can be calculated, as shown in (6). Total returns are obtained by summing the capital appreciation and the excess income over the short rate, accruing over the holding period (varying from one to five days due to weekends and (bank) holidays).

$$r_{t+1} = \sum_{i=1}^{2n-1} \frac{\frac{1}{2}y_{nt}}{(1+\frac{1}{2}y_{n,t+1})^i} + \frac{1+\frac{1}{2}y_{nt}}{(1+\frac{1}{2}y_{n,t+1})^{2n}} + \frac{\#\text{ holding days}}{365} y_{nt} - 1 \tag{6}$$

where r_{t+1} is the $\#hd$-return at time $t+1$, n is the maturity of the bond in years, t is the time and y_{nt} is the yield of an n-year bond at time t.

Afterwards, the excess returns r^e_{t+1} are calculated over the risk free rate (assumed equal to the rat eon three-month Treasury bills), once again taking into account the holding period, as shown in (7).

$$r^e_{t+1} = r_{t+1} - \frac{\#\text{ holding days}}{365} y_{3m,t} \tag{7}$$

Appendix B: Tables

Table 15: GARCH(1,1) model of daily German bond returns with FOMC announcements

	First moment parameters				Second moment parameters		
	1-yr	10-yr	30-yr		1-yr	10-yr	30-yr
μ	−0.0007	0.0098	0.0169	ω	0.0000	0.0007	0.0032
	(0.0003)	(0.0048)	(0.0095)		(0.0000)	(0.0008)	(0.0011)
θ	0.0020	0.0180	−0.0043	α	0.0753	0.0320	0.0529
	(0.0023)	(0.0366)	(0.0700)		(0.0033)	(0.0094)	(0.0066)
ϕ	0.0660	0.0218	0.0404	β	0.9227	0.9631	0.9432
	(0.0166)	(0.0148)	(0.0151)		(0.0017)	(0.0149)	(0.0073)
				δ_0	0.8130	0.8765	0.7727
					(0.2724)	(0.2163)	(0.2016)
Log L	10738.98	−1722.86	−5192.67				

Table 16: GARCH(1,1) model of daily U.K. bond returns with FOMC announcements

	First moment parameters				Second moment parameters		
	1-yr	10-yr	30-yr		1-yr	10-yr	30-yr
μ	−0.0009	0.0056	0.0097	ω	0.0000	0.0005	0.0027
	(0.0004)	(0.0054)	(0.0096)		(0.0000)	(0.0006)	(0.0010)
θ	−0.0007	−0.0256	−0.0345	α	0.0283	0.0254	0.0360
	(0.0025)	(0.0409)	(0.0715)		(0.0011)	(0.0069)	(0.0048)
ϕ	−0.0176	0.0177	0.0393	β	0.9751	0.9721	0.9595
	(0.0167)	(0.0149)	(0.0149)		(0.0006)	(0.0098)	(0.0057)
				δ_0	0.1260	0.8510	0.7678
					(0.1374)	(0.2134)	(0.2027)
Log L	9029.47	−2264.68	−5006.11				

Table 17: Ext. GARCH(1,1) model of daily U.S. Treasury bond returns with pooled announcements

Quasi-maximum likelihood estimates of the model

1. $R_t = \mu + \theta I_t^A + \phi_1 R_{t-1} + s_t^{1/2}\epsilon_t$
2. $s_t = (1 + \delta_0 I_t^A)(1 + \delta_1 I_{t-1}^A)$
3. $h_t = \omega + (\alpha_0 + \alpha_A I_{t-1}^A)\epsilon_{t-1}^2 + (\beta_0 + \beta_A I_{t-1}^A)h_{t-1}$

where R_t is the daily continuously compounded excess return of the relevant constant maturity Treasury security over the three-month Treasury bill rate, ϵ_t is an independent random variable with conditional mean zero and conditional variance h_t, and I_t^A is an indicator variable equal to one on employment or PPI announcement days (Table 17) and FOMC announcements respectively (Table 18). Returns are expressed in percent, i.e., multiplied by 100. The sample extends from January 5, 1982 to December 31, 2016. Standard errors are given in parentheses.

	First moment parameters				Second moment parameters		
	1-yr	10-yr	30-yr		1-yr	10-yr	30-yr
μ	0.0001	0.0115	0.0197	ω	0.0000	0.0029	0.0055
	(0.0002)	(0.0047)	(0.0077)		(0.0000)	(0.0006)	(0.0010)
θ	0.0015	0.0339	0.0619	α_0	0.0800	0.0499	0.0422
	(0.0011)	(0.0208)	(0.0328)		(0.0029)	(0.0048)	(0.0039)
ϕ	0.0172	0.0453	0.0212	α_A	−0.0139	−0.0391	−0.0329
	(0.0115)	(0.0109)	(0.0109)		(0.0119)	(0.0112)	(0.0100)
				β_0	0.9411	0.9410	0.9543
					(0.0041)	(0.0075)	(0.0059)
				β_A	−0.1160	−0.0065	−0.0227
					(0.0367)	(0.0414)	(0.0405)
Log L	10790.31	-5556.47	-9997.30	δ_0	1.0905	0.8854	0.6951
					(0.1094)	(0.1038)	(0.0932)
				δ_1	0.3156	0.0775	0.0348
					(0.0735)	(0.0591)	(0.0566)

Table 18: Ext. GARCH(1,1) model of daily U.S. Treasury bond returns with FOMC announcements

	First moment parameters				Second moment parameters		
	1-yr	10-yr	30-yr		1-yr	10-yr	30-yr
μ	0.0001	0.0089	0.0171	ω	0.0000	0.0015	0.0064
	(0.0002)	(0.0066)	(0.0122)		(0.0000)	(0.0007)	(0.0018)
θ	0.0014	0.0002	−0.0232	α_0	0.0837	0.0365	0.0350
	(0.0018)	(0.0461)	(0.0704)		(0.0036)	(0.0063)	(0.0047)
ϕ	−0.0394	0.0033	−0.0136	α_A	−0.0722	−0.0267	0.0347
	(0.0161)	(0.0152)	(0.0153)		(0.0232)	(0.0227)	(0.0363)
				β_0	0.9276	0.9554	0.9499
					(0.0030)	(0.0081)	(0.0059)
				β_A	−0.0524	0.0780	0.2053
					(0.0671)	(0.0790)	(0.0938)
Log L	10021.74	-2944.72	-5722.01	δ_0	0.6237	0.6531	0.2547
					(0.2215)	(0.2018)	(0.1519)
				δ_1	0.0634	0.5381	0.5088
					(0.1189)	(0.1853)	(0.1827)

Table 19: Ext. GARCH(1,1) model of daily German Bund returns with pooled announcements

Quasi-maximum likelihood estimates of the model

1. $R_t = \mu + \theta I_t^A + \phi_1 R_{t-1} + s_t^{1/2}\epsilon_t$
2. $s_t = (1 + \delta_0 I_t^A)(1 + \delta_1 I_{t-1}^A)$
3. $h_t = \omega + (\alpha_0 + \alpha_A I_{t-1}^A)\epsilon_{t-1}^2 + (\beta_0 + \beta_A I_{t-1}^A)h_{t-1}$

where R_t is the daily continuously compounded excess return of the relevant constant maturity German bond security over the three-month EURIBOR rate, ϵ_t is an independent random variable with conditional mean zero and conditional variance h_t, and I_t^A is an indicator variable equal to one on employment or PPI announcement days (Table 19) and ECB announcements respectively (Table 20). Returns are expressed in percent, i.e., multiplied by 100. The sample extends from January 4, 1999 to December 31, 2016. Standard errors are given in parentheses.

	First moment parameters				Second moment parameters		
	1-yr	10-yr	30-yr		1-yr	10-yr	30-yr
μ	−0.0007	0.0120	0.0190	ω	0.0000	0.0007	0.0032
	(0.0003)	(0.0050)	(0.0099)		(0.0000)	(0.0007)	(0.0011)
θ	0.0009	−0.0170	−0.0248	α_0	0.0626	0.0313	0.0520
	(0.0010)	(0.0167)	(0.0328)		(0.0035)	(0.0084)	(0.0066)
ϕ	0.0738	0.0198	0.0396	α_A	0.0817	0.0089	0.0021
	(0.0165)	(0.0150)	(0.0152)		(0.0330)	(0.0147)	(0.0205)
				β_0	0.9091	0.9605	0.9455
					(0.0066)	(0.0118)	(0.0093)
				β_A	0.2350	0.0282	−0.0162
					(0.0841)	(0.0803)	(0.0720)
Log L	10746.87	−1741.07	−5207.98	δ_0	0.0947	0.0467	−0.0101
					(0.0885)	(0.0893)	(0.0769)
				δ_1	−0.2553	0.0006	0.0107
					(0.0559)	(0.0784)	(0.0781)

Table 20: Ext. GARCH(1,1) model of daily German Bund returns with ECB announcements

	First moment parameters				Second moment parameters		
	1-yr	10-yr	30-yr		1-yr	10-yr	30-yr
μ	−0.0006	0.0120	0.0197	ω	0.0000	0.0007	0.0035
	(0.0003)	(0.0048)	(0.0096)		(0.0000)	(0.0008)	(0.0011)
θ	−0.0018	−0.0442	−0.0985	α_0	0.0709	0.0341	0.0530
	(0.0015)	(0.0271)	(0.0459)		(0.0035)	(0.0094)	(0.0067)
ϕ	0.0832	0.0160	0.0376	α_A	−0.0068	0.0050	0.0411
	(0.0164)	(0.0149)	(0.0151)		(0.0194)	(0.0285)	(0.0301)
				β_0	0.9243	0.9620	0.9444
					(0.0030)	(0.0141)	(0.0072)
				β_A	0.0369	−0.0287	−0.0781
					(0.0550)	(0.0612)	(0.0544)
Log L	10782.05	−1705.86	−5182.76	δ_0	0.7273	−0.1085	−0.0464
					(0.6747)	(0.0999)	(0.1149)
				δ_1	0.1899	0.9755	0.8875
					(0.1228)	(0.2354)	(0.2256)

Table 21: Ext. GARCH(1,1) model of daily U.K. Gilt returns with pooled announcements

Quasi-maximum likelihood estimates of the model

1. $R_t = \mu + \theta I_t^A + \phi_1 R_{t-1} + s_t^{1/2}\epsilon_t$
2. $s_t = (1 + \delta_0 I_t^A)(1 + \delta_1 I_{t-1}^A)$
3. $h_t = \omega + (\alpha_0 + \alpha_A I_{t-1}^A)\epsilon_{t-1}^2 + (\beta_0 + \beta_A I_{t-1}^A)h_{t-1}$

where R_t is the daily continuously compounded excess return of the relevant constant maturity U.K. government bond security over the three-month LIBOR rate, ϵ_t is an independent random variable with conditional mean zero and conditional variance h_t, and I_t^A is an indicator variable equal to one on employment or PPI announcement days (Table 21) and BoE announcements respectively (Table 22). Returns are expressed in percent, i.e., multiplied by 100. The sample extends from January 4, 1999 to December 31, 2016. Standard errors are given in parentheses.

	First moment parameters				Second moment parameters		
	1-yr	10-yr	30-yr		1-yr	10-yr	30-yr
μ	−0.0009	0.0072	0.0131	ω	0.0000	0.0005	0.0028
	(0.0004)	(0.0056)	(0.0100)		(0.0000)	(0.0008)	(0.0010)
θ	−0.0013	−0.0221	−0.0380	α_0	0.0227	0.0235	0.0373
	(0.0015)	(0.0190)	(0.0333)		(0.0017)	(0.0091)	(0.0054)
ϕ	−0.0188	0.0183	0.0381	α_A	0.0582	0.0235	−0.0060
	(0.0166)	(0.0151)	(0.0151)		(0.0197)	(0.0148)	(0.0139)
				β_0	0.9962	0.9792	0.9667
					(0.0044)	(0.0097)	(0.0069)
				β_A	−0.2101	−0.0801	−0.0822
					(0.0393)	(0.0726)	(0.0436)
Log L	9038.91	−2277.88	−5018.58	δ_0	0.1314	0.0525	−0.0551
					(0.1336)	(0.0881)	(0.0698)
				δ_1	0.1772	0.1629	0.1045
					(0.0725)	(0.0885)	(0.0810)

Table 22: Ext. GARCH(1,1) model of daily U.K Gilt returns with BoE announcements

	First moment parameters				Second moment parameters		
	1-yr	10-yr	30-yr		1-yr	10-yr	30-yr
μ	−0.0010	0.0043	0.0100	ω	0.0000	0.0003	0.0024
	(0.0004)	(0.0054)	(0.0097)		(0.0000)	(0.0004)	(0.0009)
θ	0.0044	−0.0047	−0.0417	α_0	0.0245	0.0244	0.0351
	(0.0022)	(0.0270)	(0.0505)		(0.0013)	(0.0054)	(0.0047)
ϕ	−0.0118	0.0177	0.0391	α_A	0.0579	−0.0633	−0.0220
	(0.0163)	(0.0147)	(0.0150)		(0.0187)	(0.0138)	(0.0199)
				β_0	0.9854	0.9686	0.9574
					(0.0034)	(0.0041)	(0.0060)
				β_A	−0.2065	0.1940	0.1047
					(0.0614)	(0.0409)	(0.0819)
Log L	9050.46	−2265.56	−5014.76	δ_0	0.2950	0.2891	0.2768
					(0.1358)	(0.1256)	(0.1316)
				δ_1	−0.0983	0.2727	0.1737
					(0.0930)	(0.2354)	(0.2256)

Table 23: Summary statistics for U.S. macroeconomic announcements

	Day of the week					Announcement type			
U.S.	Mon.	Tue.	Wed.	Thu.	Fri.	PPI	Emp.	Pooled	FOMC
All	0.184	0.206	0.206	0.202	0.201	0.047	0.048	0.095	0.028
PPI	0.002	0.166	0.118	0.207	0.506	1	0.002	1	0.022
Emp.	0	0.002	0.005	0.026	0.966	0.002	1	1	0.002
Pool	0.001	0.084	0.061	0.117	0.736	0.500	0.501	1	0.012
FOMC	0.004	0.538	0.418	0.036	0.004	0.036	0.004	0.040	1

Table 24: Summary statistics for German macroeconomic announcements

	Day of the week					Announcement type				
Germany	Mon.	Tue.	Wed.	Thu.	Fri.	PPI	Emp.	Pooled	ECB	FOMC
All	0.200	0.200	0.200	0.200	0.200	0.046	0.046	0.092	0.051	0.032
PPI	0.176	0.171	0.153	0.157	0.343	1	0	1	0.019	0.023
Emp.	0.000	0.327	0.221	0.438	0.014	0	1	1	0.106	0.097
Pool	0.088	0.249	0.187	0.298	0.178	0.499	0.501	1	0.062	0.060
ECB	0.004	0.000	0.058	0.938	0.000	0.017	0.095	0.112	1	0.066
FOMC	0.000	0.007	0.396	0.564	0.034	0.034	0.141	0.174	0.107	1

Table 25: Summary statistics for U.K. macroeconomic announcements

	Day of the week					Announcement type				
U.K.	Mon.	Tue.	Wed.	Thu.	Fri.	PPI	Emp.	Pooled	BoE	FOMC
All	0.199	0.200	0.200	0.200	0.200	0.046	0.046	0.092	0.046	0.032
PPI	0.556	0.236	0	0.009	0.199	1	0	1	0.009	0
Emp.	0	0.005	0.981	0	0.014	0	1	1	0	0.116
Pool	0.278	0.120	0.491	0.005	0.106	0.500	0.500	1	0.005	0.058
BoE	0.014	0.005	0.023	0.958	0	0.009	0	0.009	1	0.056
FOMC	0	0.007	0.396	0.564	0.034	0	0.168	0.168	0.081	1