News, No-News and Jumps in the US Treasury Market.[*]

Mardi Dungey▲, Michael McKenzie♦▲ and Vanessa Smith▲

▲CFAP, University of Cambridge
♦ Royal Melbourne Institute of Technology

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Abstract

Sufficiently fast and large disruptions to the continuous price process can be detected in high frequency data as jumps. Cojumping occurs when jumps occur contemporaneously across assets. This paper assesses cojumping in the US term structure using the Cantor-Fitzgerald tick dataset of 2002-2006, and finds that the middle of the curve is more likely to cojump and the ends have greater potential for idiosyncratic jumping. What is more, cojumping is strongly associated with responses to scheduled news announcements. In instances where cojumping occurs other than in response to scheduled news the price response is smaller than with the news announcements. The results are considered over a range of sampling frequencies.

Keywords: US Treasuries, high frequency, realized volatility, jumps, cojumping

JEL Classification:

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1 Introduction

The arrival of news to the market is an important event and a large literature has evolved that focuses on the issue of how information is incorporated into asset prices. In particular, macroeconomic news announcements have been the subject of a significant amount of research. This is not surprising given the importance of macroeconomic news announcements, which provide information that impacts on, and is closely monitored by participants in, all financial markets. Further, in contrast to most other forms of news, the timing of macroeconomic announcements is known well in advance. As such, traders frequently take positions in anticipation of the actual announcement and large price movements may manifest where those expectations are not met.

The potential importance of macroeconomic news announcements in understanding the price process is highlighted in the emerging literature on jumps in high frequency financial markets. In a recent paper that focuses on developing tests for discontinuities in pricing - so called 'jumps' - Barndorff-Neilsen and Shephard (2004a) provide an example which links a US trade balance announcement to a jump in the DM/USD exchange rate process.

The purpose of this paper is to provide further insights into the presence and causes of jumps in asset price data. Unlike Barndorff-Nielsen and Shephard (2006) however, we intend to focus on US bonds. Bond prices are important in their own right, and also as they are used as a reference rate for a myriad of heavily traded derivative products. Further, bond markets are of considerable interest because they are arguably the most important financial market for transmitting news on macroeconomic conditions (see Goodhart and O'Hara, 1997). Andersen, Bollerslev, Diebold and Vega (2005), find that bond prices produce the most pronounced response to macroeconomic news announcements, relative to other liquid US asset markets such as foreign exchange and equities. Our focus on the bond makets also provides an opportunity to consider disruptions to the price process for multiple traded assets. While similar studies have been undertaken in the context of financial contagion, the nature of the cross-asset, cross-country comparisons means that untangling the effects takes on a degree of complexity which econometrics is not fully equipped to deal with (see Dungey and Martin, 2006, for a discussion). The same cannot be said of bonds however, which are identical except for maturity and coupons. As such, the study of bond markets provides a unique opportunity to study the phenomena of co-jumps in asset prices.

To the best of the authors knowledge, there has not been a systematic investigation of the presence of jumps in bond markets, or their potential relationship with news announcements. This paper aims to fill that gap by investigating the association between immediate and measurably large disturbances to the pricing process, as identified by jumps testing, and their potential relationship with news events.

Most studies that focus on the effects of news announcements on bond markets, assess the impact of the unanticipated component of scheduled news announcements on daily price or yield changes (Fleming and Remolona, 1997, and Goldberg and Leonard, 2003). More recently, the focus has turned to high frequency datasets and the results show that the substantive response to news occurs within a very short period of the scheduled announcement time (see Balduzzi, Elton and Green, 2001, Gurkaynak and Wolfers, 2005, Andersen, Bollerslev, Diebold and Vega, 2006, and Campbell and Sharpe, 2006). In this paper, a new high frequency dataset obtained from Cantor-Fitzgerald is used to characterise the univariate jumps in bond prices. The results reveal that jumps occur most frequently at the short end of the maturity structure, which is consistent with the market segmentation theory. A relatively high number of jumps are also observed at the long end of the yield curve, which is consistent with elements of the liquidity premium hypothesis. The mid range maturities jumped less frequently than both the 2 and 30 year bonds.

As multivariate jump tests are not yet fully operational, we draw on the work of Bae, Karolyi and Stulz (2003) to develop a measure of the extent to which contemporaneous jumps (identified using the the univariate tests) across different maturities are observed on any given day. This 'cojump' analysis, reveals that the medium maturity 5 and 10 year bonds experience jumps that are almost entirely in conjunction with the long and short end of the curve. The 2 and 30 year bonds however, exhibit far more evidence of independent jump behaviour. A closer examination of the days on which cojumps are observed reveals that these events typically occur contemporaneously. Further, the time of these cojumps typically correspond to the times of scheduled US macroeconomic news announcements at 8:30am and 10:00am, the FOMC minutes release at 2:00pm or Treasury auctions. Confirming the results in Andersen, Bollerslev, Diebold and Vega (2006), the results of this paper reveal that bond prices react within 45 seconds of an announcement and most of the response by the market to the news is contained within the first 5 minutes.

The final part of this paper examines the source of these jumps more closely and we hypothesise that these disruptions to the continuous time process occur when the market is "surprised" by the content of the macroeconomic news announcements. An examination of the 5 minute returns around jump times shows that the unanticipated component of scheduled news announcements has a greater impact on prices than other forms of unanticipated news. Additionally, the impact on 5 minute returns is markedly greater at the 10 year maturity than for any of the others considered. The nature of this relationship would appear to be more complicated than is first thought however, as jumps are observed both with large surprises and when the median expections are met, ie. no surprise.[1]

The remainder of this paper proceeds as follows. Section 2 discusses the theoretical relationship between the term structure and the arrival of news to the market. Section 3 describes the price process and the econometric methods used in testing for univariate and multivariate jumps. The empirical application of jump tests to US Treasury bonds is considered in Section , with formal univariate tests and the application of a coexceedance measure of jumping. The cojumps are related to news using intradaily analysis in Section 5. This includes a consideration of the transmission and price impact of jumps following scheduled unanticipated news compared with unscheduled unanticipated news. Section 6 concludes.

2 Jumps in the Term Structure of Bonds

The expectations theory of the term structure of interest rates attributes the shape of the yield curve to a concensus forecast of future interest rates. In this context, any macroeconomic news that impacts on bond prices should affect all maturities and simultaneous jumps should be observed. This pure expectations theory of the term structure however, has long been discounted as a plausible explanation of interest rates as it assumes risk neutrality of investors. In response, economists have proposed a number of alternative theories to explain the shape of the yield curve.

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[1] One possible explanation for this result would be if the jumps generated by announcements with no surprise had a relatively large variation around the median compared to a no surprise announcement that does not generate a jump. Unreported empirical analysis fails to validate this hypothesis.

The liquidity preference theory of the term structure assumes that longer term rates are higher than the average of expected future rates by an amount equal to a liquidity risk premium. This premium reflects the relatively higher risk of long bonds, which possess a greater potential for capital loss before maturity. As such, rational risk averse investors demand a liquidity premium as compensation for bearing this additional risk. In the current context, the liquidity premium hypothesis highlights the fact that long bond prices are more responsive to the arrival of interest rate sensitive news compared to the shorter maturity issues. In this case, it is possible that information driven jumps may be more frequently observed at the long end of the yield curve.

One criticism of the liquidity preference theory is that it implies the risk premium would rise uniformly with maturity, which is unrealistic (albeit technically possible). The market segmentation theory also augments the expectations theory with a risk premium, however in this model it is not linked to maturity. Instead, investors are assumed to operate solely within particular segments of the yield curve and local supply and demand ultimately determine the equilibrium price for a bond at any given maturity. Investor preference for a particular maturity range may be a function of market characteristics (investors may prefer short-term instruments for reasons of liquidity) or reflect asset-liability management constraints. For example, insurance companies and pension funds typically have predictable long term liabilities, which they hedge by matching to long dated bonds. Commercial banks however, have a portfolio of short and medium term loans which prudent banking practice dictates should be funded by liabilities of a similar maturity. Thus, the segmented market theory assumes that bonds are not substitutable and the supply and demand for short-term and long-term instruments are independent. Modigliani and Sutch (1966) extended this model by removing the assumption of rigid market segmentation. Their preferred habitat theory argues that investors may be induced to move out of their chosen segment of the yield curve, where a risk premium is paid that reflects the marginal investors aversion to reinvestment risk.

The market segmentation/preferred habitat model suggests that speculators may be more active at the short end of the yield curve (where liquidity is higher) compared to the long maturity markets, which are dominated by institutional investors hedging long dated liabilities. In this case, news may generate a relatively greater response in short maturity bond prices as speculators alter their portfolio holdings whereas fund managers do not (unless that news happens to impact on the liability position of their portfolio). Thus, price jumps may be more common in short maturity bonds which contrasts to the prediction of the liquidity preference theory.

Turning to the empirical literature and a great deal of research has been undertaken in an attempt to identify which macroeconmoic news announcements are important for bond pricing (see inter alia Ederington and Lee, 1993, Beker, Finnerty and Kopecky, 1996, Fleming and Remolona, 1997, 1999a,b,c, Jones, Lamont and Lumsdaine, 1998). More recently, Arshanapalli, d'Ouville, Fabozzi and Switzer (2006) find that bonds have higher volatility on the day of macroeconomic announcements but it is a transitory effect. Pérignon and Villa (2006) observe that the switches in monetary policy play an important role in characterizing the time variation in the loadings on the common factors that drive interest rates.

In the current context, a relatively small number of papers have considered the responses of different bond maturities to the arrival of macroeconomic news. Barrett, Gosnell and Heuson (2004) found that the unexpected news component of four announcements had an impact on the term structure of zero-coupon yields that was felt equally across the entire maturity spectrum in most interest rate environments. In contrast, de Goeij and Marquering (2006) find that macroeconomic announcements seem to be especially influential at the intermediate and long end of the yield curve, while monetary policy seem to affect especially short-term bonds. Both of these papers consider daily data however, which may present a problem as most of the effect of news releases is observed in the period immediately following the announcement. In this case, intradaily data is required and papers that have considered this issue in the high frequency domain are Balduzzi, Elton and Green (2001), Gurkayanak and Wolfers (2005) and Campbell and Sharpe (2006). Campbell and Sharpe (2006) consider the 2 and 10 year bond and find relatively little difference in the news impact on the change in yield between the two maturities. This is in contrast with Balduzzi, Elton and Green (2001) who found substantially different effects by maturity, and that news impact was generally increasing with maturity for most macroeconomic announcements (I have looked in the paper and cant find this result. I can see that different macroeconomic announcements are important for different maturities and that the volatility response was greatest at the short end of the yield curve - see the volatility ratio discussion at the top of page 542). Gurkayanak and Wolfers (2005) use data from options contracts on future data announcements as an improved measure of anticipated macroeconomic news and also find that the news impact is broadly increasing from the short end to the longer end of the curve.

3 Identifying and measuring Jumps

Analysis of high frequency asset market data focuses on measures of the underlying volatility of the data generating process. The price of the asset is assumed to evolve as a continuous process of the form

$$p_t = \int_0^t a_s ds + \int_0^t \sigma_s dW_s \tag{1}$$

where p_t represents the price of the bond at time t, and the right hand side terms represent a continuous, locally bounded variation process, a_s, a strictly positive stochastic volatility process with well defined limits and right continuous, σ_s, and W_s is Brownian motion. Returns in this process are defined as $r_t = p_t - p_0$ and the associated quadratic variation is given by

$$[r,r]_t = \int\limits_0^t \sigma_s^2 ds \tag{2}$$

where the notation $[r,r]_t$ is taken to denote the equivalent of variance at time t (and commensurately $[r,q]_t$ represents a covariance between r and q). It is well known that asymptotically the quadratic variation in equation (2) can be approximated by realized variance, that is the sum of squared returns over the chosen sample period, often as here, a single trading day sampled at frequency δ, that is with n observations in a single day, where the quantity $n\delta = 1$. The subscript δ is used to identify the sampling frequency such that in expressing the realized variance,

$$RV_{t+1}(\delta) = \sum_{j=1}^{1/\delta} r_{t+j\delta,\delta}^2 \tag{3}$$

$r_{t+j\delta,\delta} = p_{t+j\delta} - p_{t+(j-1)\delta}$ are the δ period returns within the day.

Although realized variance has proved a useful concept in high frequency analysis it is also apparent that there are sometimes spikes in the daily realized variance potentially due to underyling events affecting the markets. The search for a means of identifying these spikes has led to a literature on jumps in realized volatility; see particularly Barndorff-Neilsen and Shephard (2004a) and Andersen and Bollerslev, Diebold (2006). This consists of augmenting the continuous process given in equation (1) with a potentially discontinous jump component as follows

$$p_t = \int_0^t a_u du + \int_0^t \sigma_u dW_u + \sum_{j=1}^{N_t} c_{jt} \tag{4}$$

where the final term is the jump process with c_{jt} a non-zero random number, and N is a count variable, representing the number of jumps.

The quadratic variation associated with this equation is given by

$$[r, r]_t = \int\limits_0^t \sigma_s^2 ds + \sum_{j=1}^{N} c_j^2. \tag{5}$$

Barndorff-Nielsen and Shephard (2004a) show how to separate the jumps using bi-power variation.[2] This technique for separating jumps relies on the observation that forms other than realized variance also converge to the true quadratic variation given in equation (2). In particular the Barndorff-Neilsen and Shephard (2004a) test exploits realized bi-power variation, which consists of the standardized sum of the product of consecutive returns given by

$$BV_{t+1}(\delta) = \mu_1^{-2} \sum_{j=2}^{1/\delta} \left| r_{t+j\delta,\delta} \right| \left| r_{t+(j-1)\delta,\delta} \right| .$$

The coefficient of standardization is the mean of the absolute value of the standard normally distributed random variable, $\mu_1 = \sqrt{2/\pi}$. Bi-power variation has the property that

$$BV_{t+1}(\delta) \to \int\limits_0^t \sigma_s^2 ds. \tag{6}$$

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[2] This turns out to be a special case of the more general testing framework for jumps recently proposed by Ait-Sahalia and Jaoud (2006), however, their new tests require extremely high numbers of observations to produce good sampling properties, the authors recommend less than one minute sampling, and are hence unsuited to the current data set.

However, it can be quickly seen that in the event of a large change in returns within a day (a jump) bi-power variation and realized volatility will not pick this up in the same manner. Hence the difference between realized volatility and the bi-power variation gives a consistent estimate of a jump. Asymptotically as $\delta \to 0$

$$RV_{t+1}(\delta) - BV_{t+1}(\delta) \to \sum_{0<s\leq t} \kappa_s^2.$$

In a finite sample it is possible that the sample bi-power variation may be negative, so it is convenient to truncate the measure of jumps at zero and define the jumps $J_{t+1}(\delta)$ as

$$J_{t+1}(\delta) = \max\left[RV_{t+1}(\delta) - BV_{t+1}(\delta), 0\right].$$

In order to select statistically significant jumps the jumps test statistic can be defined as

$$JS_{t+1}(\delta) = \frac{RV_{t+1}(\delta) - BV_{t+1}(\delta)}{\sqrt{(\mu_1^{-4} + 2\mu_1^{-2} - 5)\delta \int\limits_t^{t+1} \sigma^4(s)ds}} \to N(0,1) \qquad (7)$$

under the null hypothesis of no jump. An estimate of $\int\limits_t^{t+1} \sigma^4(s)ds$ is provided by the realized tri-power quarticity, $TQ_{t+1}(\delta)$, even in the presence of jumps. For $\delta \to 0$

$$TQ_{t+1}(\delta) = \delta^{-1}\mu_{4/3}^{-3} \sum_{j=3}^{1/\Delta} \left|r_{t+j\delta,\delta}\right|^{4/3} \left|r_{t+(j-1)\delta,\delta}\right|^{4/3} \left|r_{t+(j-2)\delta,\delta}\right|^{4/3} \to \int\limits_t^{t+1} \sigma^4(s)ds,$$

where $\mu_{4/3} = 2^{2/3}\Gamma(7/6)\Gamma(1/2)^{-1}$. Hang and Tauchen (2005) however, have shown that a statistic based on substituting $TQ_{t+1}(\delta)$ into equation (7) tends to over-reject the null. As such, the test statistic implemented in this paper contains a correction based on modifying the denominator of equation (7) (see also Andersen, Bollerslev and Diebold, 2006) as follows.

$$JS_{t+1}(\delta) = \frac{RV_{t+1}(\delta) - BV_{t+1}(\delta)}{\sqrt{(\mu_1^{-4} + 2\mu_1^{-2} - 5)\max\left\{1, TQ_{t+1}(\delta)BV_{t+1}(\delta)^{-2}\right\}}} \to N(0,1). \qquad (8)$$

The test is then implemented for chosen significance levels. In practice, the significance level chosen has to be quite high as the test tends to find rather a lot of jumps - for example see Beine et al (2006). A solution to this problem is an ongoing issue in the literature and so we choose a highly conservative significance level of 0.001.

3.1 Multivariate Extension

The current application has the novelty of a number of assets which can be regarded as homogeneous in all but maturity. We abstract from any possible complications relating to different coupon rates - in the analysis that follows, this is not considered to be an important issue. Hence, as well as identifying jumps in the prices of individual assets through the application of univariate jumps tests, we are also interested in events where bonds of differing maturity co-jump. To date, a formal multivariate test statistic has not been developed, although Barndorff-Neilsen and Shephard (2004b) have developed the main elements in the form of multivariate Quadratic Variation (denoted MQV) and multivariate bi-power covariation (BPCV).

In particular, Barndorff-Neilsen and Shephard (2004b) demonstrate that co-jumping series can be identified by a reduced rank matrix in the process describing the data. Consider the multivariate analog of equation (4), where P_t refers to a vector of prices.

$$P_t = \int_0^t a_s ds + \int_0^t \sigma_s dW_s + \sum_{j=1}^{N_t} C_j \tag{9}$$

where C_j represents the jump components in each of the bonds. If each of the assets in the set P_t jumps within a time period $[0,t]$ then a new process can be defined such that

$$\begin{aligned} X_t &= \int_0^t D_{s-} dP_s \\ &= \int_0^t D_s a_s ds + \int_0^t D_s \sigma_s dW_s + \sum_{j=1}^{N_t} D_{\tau j-} C_j \end{aligned} \tag{10}$$

where D is a non-zero matrix with the dimension k by number of rows in P_t whose elements are left bounded and right continuous and τ represents the arrival time of the information which causes the jump. Thus D may generate a process X_t which has continuous properties despite discontinuities in P_t during the same interval. This case is defined as co-jumping.

If we consider the variance of the X_t series, denoted $[X]_t$ this can be broken down as

$$[X]_t = \int_0^t D_s \sigma_s \sigma'_s D'_s dW_s + \sum_{j=1}^{N_t} D_{\tau j-} C_j C'_j D'_{\tau j-}. \tag{11}$$

Co-jumping will be identified in the last term of equation (11), where some of the elements of the diagonals (variances) are zero. If D is assumed to be time invariant then cojumping will occur whenever the matrix

$$\left[P_t^d\right] = \sum_{j=1}^{N_t} C_j C'_j \tag{12}$$

is of reduced rank. Further, if full rank is given by p then rank of $p-m$ should indicate m contemporaneous jumps.

In order to get a first look at how the co-jumping test tallies with the analysis of multiple jumps in univariate tests we examine the values of $\left[P_t^d\right]$ in the sample data set. Asymptotically the difference between the multivariate quadratic variation matrix, and the bi-power covariance matrix converges in probability to $\left[P_t^d\right]$ subject to a scaling factor - so that

$$[P_{\delta,t}] - \mu_1^{-2} \{P_{\delta,t}\} \xrightarrow{p} \left[P_t^d\right] \tag{13}$$

where $[P_{\delta,t}]$ denotes the multivariate quadratic variation of P_t, the analog of $RV_{t+1}(\delta)$ in the univariate case, and $\{P_{\delta,t}\}$ denotes the multivariate bi-power covariation, the analog of $BV_{t+1}(\delta)$ in the univariate case, with the scalar $\mu = \sqrt{2/\pi}$.

Individual elements of the realized quadratic variation $[P_{\delta,t}]$ are denoted $[P_t]_{l,k}$ and given for any day t as

$$[P_t]_{l,k} = p \lim_{n\to\infty} \sum_{j=1}^{n} \left(P_{l,tj} - P_{l,tj-1}\right)\left(P_{k,tj} - P_{k,tj-1}\right)$$

with clear implications when $l = k$ for the diagonal elements.

Practical implementation is undertaken using the bi-power realized covariation for the multivariate case developed in Barndorff-Neilsen and Shephard (2004b) where realized BPCV is defined as a square matrix with diagonal elements for asset P_l , where analagously to the univariate case, $r_{l,j} = P_{l,j} - P_{l,j-1}$ with j the time interval, so that

$$\{P_l; q\} = \gamma_{q,\delta} \sum_{j=q+1}^{n} |r_{l,j-q}| \, |r_{l,j}| \, , \qquad \gamma_{q,\delta} = \frac{n}{n-q} \tag{14}$$

where $\{P_l; q\}$ denotes the diagonal elements on the bi-power covariance matrix, q is the number of lags considered in the bi-power variation (we will default to $q = 1$ as in the univariate case above) and the parameter $\gamma_{q,\delta}$ is a correction factor taking into account the number of observations per day adjusted for the number of lags.

The corresponding off-diagonal terms for the BPCV matrix are given for P_l and P_k by

$$\{P_l, P_k; q\} = \frac{\gamma_{q,\delta}}{4} \left(\{P_l + P_k; q\} - \{P_l - P_k; q\} \right) \tag{15}$$

using polarisation results; see Barndorff-Neilsen and Shephard (2004b).

In the case of $q = 1$ which corresponds to the univariate tests above equations (14) and (15) correspond to

$$\{P_l\} = \frac{n}{n-1} \sum_{j=2}^{n} |r_{l,j-1}| \, |r_{l,j}| \, , \tag{16}$$

$$\{P_l, P_k\} = \frac{n^2}{4(n-1)^2} \left(\sum_{j=2}^{n} |r_{l,j-1} + r_{k,j-1}| \, |r_{l,j} + r_{k,j}| - \sum_{j=2}^{n} |r_{l,j-1} - r_{k,j-1}| \, |r_{l,j} - r_{k,j}| \right) \tag{17}$$

Analysis of this reduced rank property gives an indication of co-jumping on any day, despite the current lack of a formal test directly analagous to the univariate test given in equation (8).

4 Empirical Results

Previous research on US bond markets had typically focussed on the GovPX dataset. The use of GovPX data brings with it a number of issues related to identifying trades, matching the actual bid-ask spread to trades, and correctly calculating the volume of trade. These problems have meant that researchers have had to undertake complicated sample manipulation, as in Boni and Leach (2004), Brandt and Kavajecz (2004,) and

Dungey, Goodhart and Tambakis (2005), or make the decision to completely ignore these problems and sample from the entire database, as in Andersen and Benzoni (2005).

Since 2000, the US treasury market has undergone a significant number of changes (for details, see Mizrach and Neely, 2006), which have resulted in a severe drop in the coverage of the GovPX database and the emergence of two new US bond data vendors: Cantor Fitzgerald who provides the eSpeed database and ICAP who provide the BrokerTec database. Mizrach and Neely (2006) report that on-the-run trading is now almost completely electronic, with eSpeed (BrokerTec) capturing 40% (60%) of trading volume. They compare the two databases and find there is qualitatively few differences, which suggests that any empirical results are unlikely to be source dependent. In comparision to the GovPX data, the Cantor data uniquely identifies each individual transaction in the workup to an actual trade. There is some loss of richness in the data, in that the database does not cover the entire workup process and there is no information on prices posted where no transaction ensues.[3] The accurate identification of trades however, would seem to outweigh these disadvantages for most purposes.

In this paper, we sample data from the Cantor database begining with the first available observation on January, 2 2002 to September 29, 2006. These 1166 trading days provide over 13.5 million trades in on-the-run bonds, which averages to over 11.7 thousand trades each day. While the dataset covers the 2, 3, 5, 10 and 30 year maturities, we have excluded the 3 year market as data is only available from April 30, 2003. A trading day is defined as starting at 7:30am and finishing at 5:30pm New York time. After the data have been filtered to remove US public holidays, we include all days in our sample if there is trade prior to 3pm New York time on that day.

Figure 1 presents daily volumes and average trade size for each maturity. The average daily trading volume being highest in the 5 year bond, with the 10 year exhibiting a similar level of interest. The 30 year bond however, has a much lower trading volume. The average trade size for the 2 year bond is highest (averaging 10 trades per day) and this falls progressively as the bond maturity increases. A full discussion of the volume

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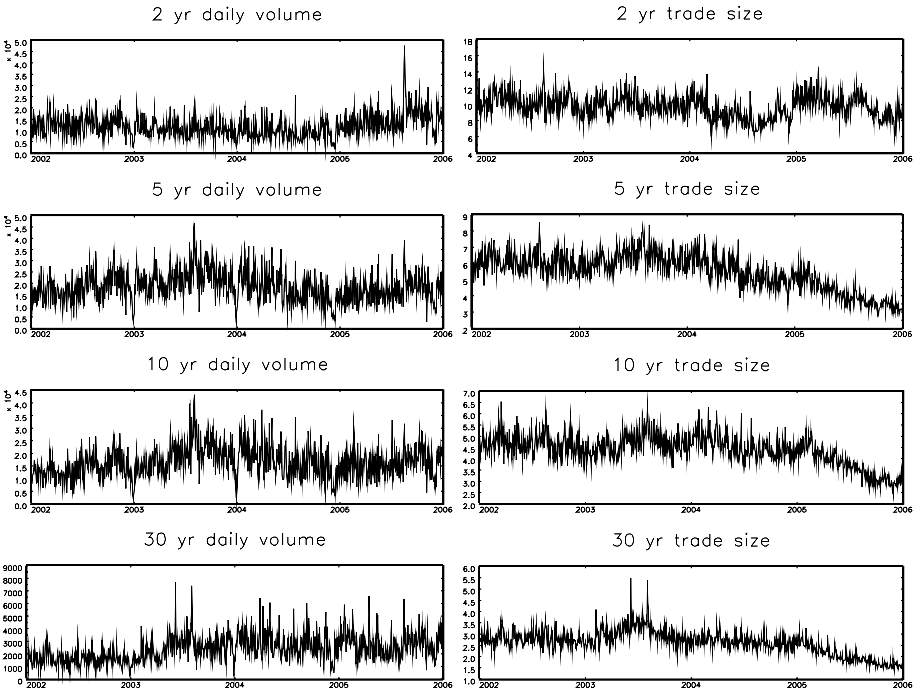
[3] The workup results from the expandable limit order book which operates in this market and refers to the process by which after a bid has been accepted by a counterparty the two negotiate on how much volume will actually be transacted at the agreed price. An analysis of workups in this market is provided in Boni and Leach (2004).

Figure 1: Daily volume and trade size by maturity 2002-2006

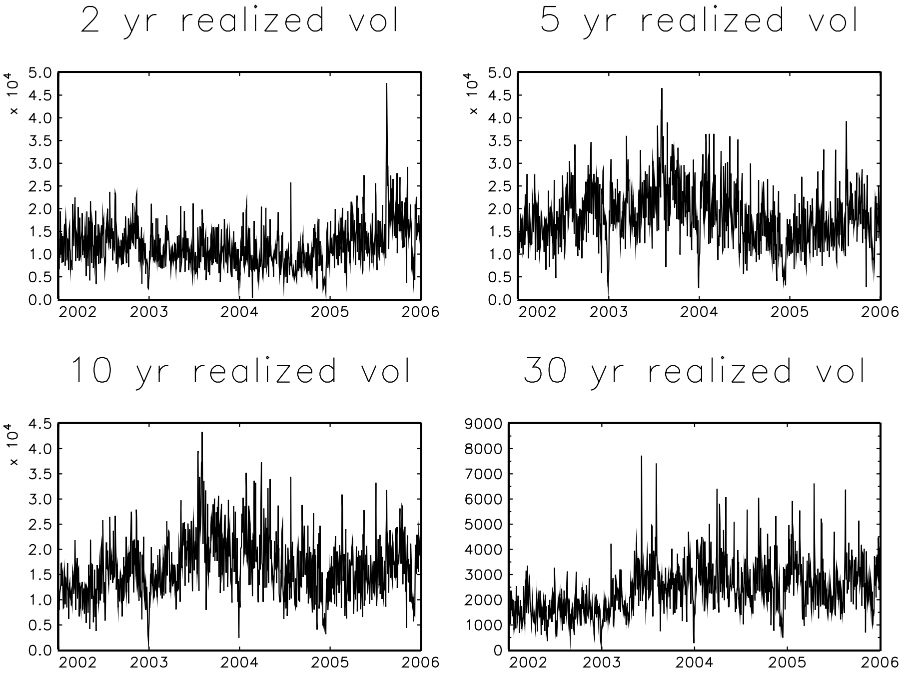


and trade flow properties of this data is beyond the scope of this paper and further details may be found in Dungey and Long (2007).

The daily realized variance for each of the four maturities sampled at a 15 minute frequency are presented in Figure 2. Realised volatility is lowest in the 30 year contract and highest in the 5 year maturity. A large number of outliers are visible, which may be indicative as to the presence of jumps in the data. We will return to consider this issue in the next section.

The jump testing procedures described in Section 3 may be applied to the data for each of the four bond maturities (2, 5, 10 and 30 year). As the data is trade by trade, it raises the important issue of sampling frequency. That is, a trade-off exists between sampling as frequently as possible to provide the greatest amount of information and sampling a noisy price signal. Bandi and Russell (2006), consider this issue and suggest that the optimal sampling frequency can be approximated by separately identifying microstructure noise from the true variance of the process based on the characteristics that average realized volatility over the sample days ($h = 1$ in their notation) consistently estimates the variance of the noise process, while the average

Figure 2: Realized Volatility calculated from 15 minute intervals by maturity



quartic volatility over the sample days consistently estimates the fourth moment of the noise. They show that the optimal sampling frequency δ^* is the minimimsation of the property

$$2\delta^3\widehat{\alpha} + \delta^2\widehat{\beta} - 2\widehat{Q_i}$$

where

$$\widehat{\alpha} = \left(\frac{\sum_{i=1}^{n} \sum_{j=1}^{1/\delta} \widetilde{r}_{j,i}^{2}}{nM} \right)^2$$

$$\widehat{\beta} = 2\frac{\sum_{i=1}^{n} \sum_{j=1}^{1/\delta} \widetilde{r}_{j,i}^{4}}{nM} - 3\left(\frac{\sum_{i=1}^{n} \sum_{j=1}^{1/\delta} \widetilde{r}_{j,i}^{2}}{nM} \right)^2$$

$$\widehat{Q_i} = \frac{M}{3} \sum_{j=1}^{1/\delta} \widetilde{r}_{j,i}^{4}$$

and $\sum_{j=1}^{1/\delta} \widetilde{r}_{j,i}^{2}$ is the estimate of realized volatility for a given day, and $\sum_{j=1}^{1/\delta} \widetilde{r}_{j,i}^{4}$ is the estimate of quartic volatility. Following Bandi and Russell (2006), we use 10 and 15 minute sampling intervals implement this test and and derive an optimal sampling length of 12 and 19 minutes respectively. Thus, the optimal interval increases with a lower sampling frequency and no real solution to the problem is provided.

As an alternative, we could simply choose to follow the sampling intervals specified in the previous literature. This is easier said than done however, as a wide range of intervals has been used. Fleming (1997) uses 30 minute samples in his study of the US bond spot market, while Andersen, Bollerslev, Diebold and Vega (2006) and Mizrach and Neely (2005) sample at 5 minute intervals in their study of bond futures data (although none of these authors apply an optimal sampling frequency test). A sampling interval of 15 minutes is common across much of the high frequency literature on foreign exchange markets, although equity market literature tends to sample far more frequently.

Rather than select one optimal frequency for analysis, in this paper we consider the robustness of the estimation results to the sampling interval. Thus, each of our testing procedures will be applied to data sampled at 5, 10, 15 and 30 minute intervals, where the last available observation in the sample interval is taken as the indicative price. Further, the sample volume for the period is taken as the total volume transacted within that time interval. This discretisation is subject to the potential problem of scrambling as described by Shephard (2006). In the interests of conserving space however, our focus will be on the 15 minute data where appropriate.

4.1 Univariate Jumping

Table 1 shows the rejection frequency of the jumps test for the different maturities at the 0.1% significance level, that is the proportion of total observations which are jumps. Where the data is sampled at a 5 mintue frequency, jumps are found in the 2 year bond price series on 914 days in a sample of 1166, or 78.4% of the time. The 30 year bond exhibits the second highest number of jumps (689), ie. a jump occurs on 59.1% of the days in the sample. The intermediate 5 and 10 year bonds jumps the least, generating a significant test score 42.7% and 44.8% of the time respectively. Turning to consider the other sampling intervals and the results qualitatively mirror the 5 minute outcome. That is to say, the 2 year bond consistently generates the highest number of jumps, well in excess of those observed in the other maturities. The 30 year bond provides the second highest number of jumps (except for the 30 minute interval), and the 5 and 10 year bond show the least number of jumps, but generate a similar number of signficant test scores. It is interesting to note that the proportion of jumps identified in each maturity increases with the sampling frequency. One possible explanation for this result could be due to increased prevalence of microstructure noise as opposed to the identification of actual jumps (see Bandi and Russell, 2006). It is not clear however, what is driving this result.

A graphical representation of these results is provided in Figure 3, which shows the jumps tests results for each of the maturities for the 15 minute sampling interval. Each observation represents the value of the test statistic over and above that of the critical value in each case. Thus, the height of each observation represents the significance of the estimated test score above the critical value. As is common in most jumps literature, the number of days with jumps is a relatively high proportion of the total number of days in the sample. But what is not evident from previous studies is how much this can vary across different price series that are governed by similar price dynamics. These results show that there are far more jumps in the shorter maturities than longer maturities. Further, this figure reinforces the earlier discussion on the greatest number of jumps occuring in the 2 year bond followed by the 30 year maturity, with the 5 and 10 year maturities generating the lowest number of price discontinuities. This accords with the behaviour of short and long bonds and their volatility characteristics as recorded in the previous section. An interesting question which arises from Figure 3 and Table 1 is the extent to which the jumps occur contemporaneously across maturities. The figures suggest a degree of coincidence in observed jumps insomuch as many of the large critical values appear contemporaneously across maturities and clustering in jump activity also appears common. In the next section, we consider this issue more closely.

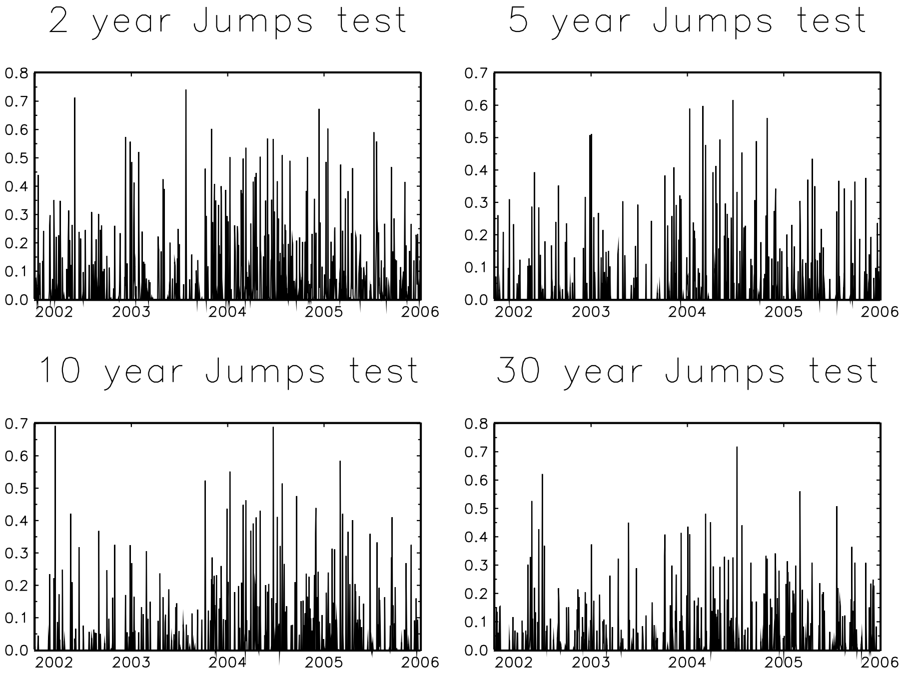
4.2 Cojumping

An area of considerable interest in the recent theoretical high frequency literature has been the development of a multivariate counterpart to the univariate jump test given in equation (8). This jump test would be used to identify when simultaneous discontinuities are observed in the price series for two or more assets. As with many cobreaking type tests, a fundamental difficulty exists in establishing whether the test

Table 1:
Rejection Frequency of Jumps Test for 7:30am to 5:30pm sample, number of jumps identified

Maturity	Total days	Rejection Frequency	No. of jump days	Rejection Frequency	No. of jump days
		30 minute sampling		*15 minute sampling*	
2 year	1166	0.279	325	0.389	453
5 year	1166	0.200	233	0.225	262
10 year	1166	0.204	238	0.241	281
30 year	1166	0.194	226	0.251	293
		10 minute sampling		*5 minute sampling*	
2 year	1166	0.521	607	0.784	914
5 year	1166	0.285	332	0.427	498
10 year	1166	0.272	317	0.448	522
30 year	1166	0.300	350	0.591	689

Figure 3: Barndorff-Neilsen and Shephard Univariate Jumps test by maturity at 0.001% significance.



statistic may reject a null of no jumps on the basis of a single large jump in one series or from combined multiple jumps. The rank of the matrix $\left[P^d\right]_t$ resulting from equation (13) then gives the indicative value for co-jumping. As this matrix is symmetric however, the result in the data is always for a full rank matrix, indicating no cojumping at odds with the strong results from the univariate test results. At this point, the development of a multivarite jumps tests remains an unresolved area of the literature that is subject to ongoing research.

An alternative avenue for examining the degree of cojumping in bond data is through the construction of co-exceedances, as applied by Bae, Karolyi and Stulz (2003) to extreme events in financial market returns. In the current context, the test is a simple count of the number of times the estimated jump test score simultaneously exceeds a pre-determined threshold across different maturities. The threshold is given by the critical value of the jump statistic, $JS_{t+1}(\delta)$, which will be determined independently for each series under consideration. More formally, denote $d_{i,t,\delta}$ as a binary variable indicating whether returns in bond of varying maturity subscripted i, $i = 1..n$, (sampled at frequency δ) contain a jump as indicated by the univariate jumps test,

$$d_{i,t} = \left\{ \begin{array}{ll} 1: & JS_{i,t}(\delta) > JS_{i,t}(\delta)_{critical} \\ 0: & \text{otherwise} \end{array} \right. . \tag{18}$$

The number of coexceedances for a jump in bond of maturity j recorded at time t can then be calculated as a simple sum of $d_{i,t}$ over all $i \neq j$,

$$E_{j,t} = \sum_{i=1, i\neq j}^{n} d_{i,t}$$

which in the current application of 4 maturities, $n = 4$, means that $E_{j,t}$ varies discretely between 0 and 3. Table 2 presents this information in terms of the number of co-exceedances associated with an observed jump in the maturity shown in the first column. For example, with 15 minute sampling, the 2 year bond is observed to jump independently (no. of co-exceedances is 0) on 176 occasions, and with a bond of an unspecified maturity 105 times, and contemporaneously with all the maturities in the sample 99 times (the column headed 3 coexceedances). The final column in the table gives the total number of jumps recorded in the maturity of that row. Table 3 gives the same information as in Table 2 with the figures in the columns expressed as a proportion of the total number of jumps in that maturity.

Table 2:
Number of coexceedances in jumps by maturity for 7:30am to 5:30 pm sample

Maturity	Co-exceedances 0	1	2	3	Total number of jumps
30 minute sampling					
2 year	124	81	53	67	325
5 year	32	68	66	67	233
10 year	36	66	69	67	238
30 year	68	49	43	67	227
15 minute sampling					
2 year	176	105	73	99	453
5 year	21	66	76	99	262
10 year	34	69	79	99	281
30 year	75	68	51	99	293
10 minute sampling					
2 year	217	169	110	111	607
5 year	32	97	92	111	332
10 year	23	86	97	111	317
30 year	71	98	70	111	350
5 minute sampling					
2 year	186	241	230	257	914
5 year	10	61	170	257	498
10 year	13	79	173	257	522
30 year	84	177	171	257	689

Table 3:
Proportion of coexceedances in jumps by maturity for 7:30am to 5:30pm sample

	Co-exceedances				Total number
Maturity	0	1	2	3	of jumps
30 minute sampling					
2 year	0.382	0.249	0.163	0.206	325
5 year	0.137	0.292	0.283	0.288	233
10 year	0.151	0.277	0.290	0.282	238
30 year	0.300	0.216	0.189	0.295	227
15 minute sampling					
2 year	0.389	0.232	0.161	0.219	453
5 year	0.080	0.252	0.290	0.378	262
10 year	0.121	0.246	0.281	0.352	281
30 year	0.256	0.232	0.174	0.338	293
10 minute sampling					
2 year	0.357	0.278	0.181	0.183	607
5 year	0.096	0.292	0.277	0.334	332
10 year	0.073	0.271	0.306	0.350	317
30 year	0.203	0.280	0.200	0.317	350
5 minute sampling					
2 year	0.204	0.264	0.252	0.281	914
5 year	0.020	0.122	0.341	0.516	498
10 year	0.025	0.151	0.331	0.492	522
30 year	0.122	0.257	0.248	0.373	689

The information provided in Tables 2 and 3 provides an interesting characterisation of the jumps. The 2 year bond displays a larger absolute number and also proportion of unique jumps (given in column headed 0 in Table 3) than the other maturities. For example, at the 15 minute frequency, 38.9% of the jumps in the two year bond price series were unique to that series. By way of comparison, only 8% (12.1%, 25.6%) of the observed jumps were unique for the 5 (10, 30) year bonds. The 5 and 10 year bonds however, the majority of observed jumps occur when at least the price of at least two other maturities are also observed to be jumping. The 15 minute data highlights this trend as around 70% of all price jumps for the 5 and 10 year bond occur during days on which at least two other price series also exhibit jumps. The 30 year bond series exhibits elements of both of these trends, ie. it exhibits a reasonable proportion of unique jumps, but also cojumps quite frequently. Overall, these results tend to suggest a stylized representation of the yield curve consistent with elements of both the liquidity preference and prefered habitat theory.

The choice of sampling frequency does not alter these basic observations about the results, however it does tend to magnify the pattern in the 5 and 10 year bonds. That is, in the most frequently sampled data, 5 and 10 year bond prices exhbiit unique jumps only infrequently, whereas around 85% of all price jumps occur during days on which at least two other price series also exhibit jumps. It is interesting to compare the impact of different sampling intervals on these cojumping results relative to their impact on the univariate jump results as discussed in the previous section. We find that greater sampling frequency generates not only more observed days of jumps, but also (disproportionately) more co-exceedances. That is, more frequent sampling of the data would appear to increase the probability of finding all maturities jumping on the same day with 5 minute sampling than 30 minute sampling. This is not to suggest however, that these jumps occur at the same time within that day, and hence they may not be truly contemporaneous. Section 5 returns to this aspect of intraday timing below.

To investigate these cojump results further, it is useful to consider whether common jumps are in general larger than those which are classified unique. To this end, Table 4 presents the average size of the jump statistic for each bond and all of the sampling intervals included in our paper. The average jump size for the 15 minute data ranges

Table 4:
Average Jump Size (size of Jstat) for 7:30 am to 5:30pm sample

	All Jumps Data		Unique Jumps		Co-exceedences = 3	
	Means	Variance	Mean	Variance	Mean	Variance
30 minute sampling						
2 year	0.193	0.013	0.041	0.014	0.249	0.027
5 year	0.152	0.007	0.006	0.005	0.258	0.022
10 year	0.149	0.007	0.008	0.002	0.238	0.022
30 year	0.140	0.006	0.026	0.009	0.196	0.017
15 minute sampling						
2 year	0.167	0.015	0.032	0.015	0.240	0.029
5 year	0.156	0.008	0.002	0.001	0.220	0.022
10 year	0.144	0.008	0.006	0.003	0.228	0.022
30 year	0.140	0.008	0.021	0.007	0.213	0.020
10 minute sampling						
2 year	0.175	0.018	0.032	0.016	0.240	0.033
5 year	0.135	0.008	0.003	0.002	0.206	0.026
10 year	0.133	0.008	0.006	0.032	0.204	0.020
30 year	0.136	0.009	0.020	0.023	0.173	0.017
5 minute sampling						
2 year	0.234	0.034	0.013	0.026	0.288	0.031
5 year	0.140	0.013	0.000	0.001	0.172	0.020
10 year	0.131	0.012	0.003	0.084	0.164	0.018
30 year	0.155	0.017	0.016	0.019	0.177	0.018

from 0.167 for the 2 year bond to 0.14 for the 30 year bond. The variance of these test scores is 0.008 for each of the 5, 10 and 20 year bonds, whereas the 2 year bond test scores have a variance of almost twice that. Where only the unique jumps are considered however, the average jump test scores are considerably lower for all of the bonds and the variance is also much smaller (except for the 2 year bond). Alternatively, the average test score and variance where all of the bond prices jump (the number of coexceedences is equal to 3) is markedly higher. Thus, a clear distinction exists between the size of the jump observed when one price series jumps on its own compared to when all bonds jump.

5 Sources and Timing of Jumps

The analysis of this paper shows that daily bond price dynamics are characterised by frequent discontinuities and that large jumps are often observed simultaneously across a number of maturities. In this section of the paper, we investigate what drives these episodes of cojumping and whether they can be attributed to the arrival of news to the market. In addressing this issue, it is necessary to identify the precise timing of the cojump within the day. Recall that the jump tests only identify the days on which jumps take place and it cannot be taken for granted that these jumps are contemporaneous. As such, we identify the precise timing of any given jump using an approach that is motivated by the method used in Beine et al (2006). Specifically, on a day in which a jump is known to have occured, the returns for each sampling interval in the day are ranked in terms of their absolute value for each maturity. This ranking is then compared across maturities to identify the time period during which the largest return is observed. To limit the scope of our results, we only formally consider where all 4 maturities cojump within the exact same interval[4] and Table 5 provides a detailed breakdown of the results. Consistent with our expecations, we find that the vast majority of jumps across maturities occur in the same time interval.[5] Focussing on the results obtained using the 15 minute data, we find that 40 of the 99 days on which jumps are observed across all four maturities occur in the time interval 8.30 to 9.00. A further 14 of these cojumping events occur after 10.00 and on one occassion a cojump across all four maturities is observed after 14.15. The remaining 22 cojumps occured at various other time intervals. Where the data is sampled at a different frequency, the basic tenor of these results is unchanged insomuch as most of the cojumps occur at 8:30am each day.

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[4] We tested the sensitivity of these results to other options and found they are qualitatively unchanged. For example, where we consider the case where 3 maturities experienced their largest intradaily return in the same interval, and the remaining maturity experienced its second largest intradaily return in that interval, the number of observations classified as cojumping increased and the number of times such cojumping was related to news events is reduced.

[5] A rejection of the null of no jumps in a day is not necessarily related to the largest jump in the day, and may be the result of multiple movements. We do not consider this problem here, and examination of the individual circumstances for the cojumping data suggests it is not an issue for the current application. However, in general the identification of the critical observations for jumps is more complex than this. Beine et al (2006) attempt to identify contributors to jump by stripping the largest observations from each day until the test statistic fails to reject the null, however they have failed to account for the bias such a process would create in the test statistic distribution. This is an area of future research.

In an attempt to explain these large price co-movements, we turn to the literature and the available empirical evidence providese clear evidence of a link between macroeconomic news announcements and movements in bond market prices. For examples of this literature see Fleming and Remolona (1997), Goldberg and Leonard (2003), and in the high frequency data domain, Bollerslev, Cai and Song (2000), Balduzzi, Elton and Green (2001), Gurkaynak and Wolfers (2005) and Campbell and Sharpe (2006).[6] Further, Andersen, Bollerslev, Diebold and Vega (2006) find that bond markets tend to respond more to macroeconomic announcements compared to foreign exchange or stock markets.

This literature clearly suggests that macroeconomic news announcements may manifest as simultaneous large price movements across all maturities in the US bond market. Recall that the majority of the cojumps across all four maturity markets are observed in the 8.30 - 9.00 trading interval. This is also the time at which most scheduled US macroeconomic news announcements are released to the market. The other important period of clustering in the jumps is around 10:00am, which is potentially associated with Treasury auction announcements. Further, there is also some price events occurring in the period from 14:15 to 14:45, which is the period where press statements are released concerning FOMC decisions.

Table 5 compiles the timing of the contemporaneous jumps across all maturities and their association with major macroeconomic news announcements, FOMC press announcements, auction news and those associated with no discernible news. The macroeconomic news announcements considered in compiling this list were announcements on non-farm payrolls, retail sales, CPI, PPI, GDP (advance, preliminary and final), housing starts, industrial production and durable goods numbers. For the 15 minute data, 78% (ie. 77 of 99 cojumps) of all days on which a cojumps occurs across all four maturities takes place contemporaneously. Further, of those 77 jumps, 71% (ie. 55) were uniquely identifiable as an event taking place in a window that coincides with some form of news announcement.

In a number of instances, even though the daily analysis reveals evidence of simultaneous jumps across all maturities, the intradaily ranking shows that the largest

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[6] The literature also contains papers linking high frequency data and macro news in other asset markets such as Andersen, Bollerslev, Diebold and Vega (2003) and Andersen and Bollerslev (1998) for foreign exchange.

absolute return is not contemporaneous. These occurrences are labelled 'unallocated' in Table 5. In terms of the 15 minute data, 22% (22 observations) of the total days where all four maturities cojump, are found to jump at different times during the day. Further, consistent with our expectations, the proportion of unallocated jumps increases with higher sampling frequency (ranging from 19% of the total observed jumps at 30 minute sampling to 67% at 5 minute sampling. In general, this is not due to an increase in the number of large returns in adjacent periods, but rather just due to increased number of days recording jumps at higher sampling.

To provide more insights into these results, Table 6 gives a more detailed breakdown of the time immediately following the major news announcements at 8:30 and 10:00 and the results show that there are some interesting differences in the way in which these jumps occur. For the 8:30 - 9.00 window, nearly all of the cojumps identified occurs in the first sampling window of that interval. For example, of the 68 cojumps identified in the 5 minute data in the 8:30 and 10:00 window, all of them occur in the first five minutes after the announcement, ie. 8.30 - 8.30. The same is true of the 15 and 30 minute data, which echoes the results of Andersen, Bollerslev, Diebold and Vega (2006). For the 10 mintue data however, 3 cases were observed in which all four maturies cojumped in the second 10 minute interval, ie. the 8.40 - 8.50 interval.

More pronounced evidence of these effects is the found in the 10am announcements, where the 10, 15 and 30 minute samples all indicate that a jump occurs in the periods beginning 10:00 am (and ending 10:10, 10:15 and 10:30 respectively). The cojump results for the 5 minute sample however, reveal that the jump occurs in the third sampling interval following 10.00, ie. the period 10:10-10:15. The combination of this information suggests that the reaction to the news released at 10:00 may be slightly delayed compared with the 8:30 announcements. It would be of great interest to compare these results to the 10:00 announcement results of Andersen, Bollerslev, Diebold and Vega (2006).

The preceeding analysis clearly indicates that contemporaneous jumps occur across each of the bond markets in the first five minutes after an announcement. To provide further insights into this result, it is necessary to transcend into the realm of transaction level data. This will allow us to determine whether any particular segment of the bond market may be systematically taking the lead in transmitting information, or whether

Table 5:
Jump Timings

time period	sampling frequency 5 min	 10 min	 15 min	 30 min
8:30 - 9:00	68	56	40	27
10:00 - 10:15	17	14	14	15
14:15 - 14:45	2	0	1	1
other	24	9	22	11
total	111	79	77	54
unallocated	146	32	22	13
total jump days	257	111	99	67
news jumps as propn of total (%)	79	88	71	80
unallocated as propn of jump days (%)	67	29	22	19
total as propn of total jump days (%)	43	71	78	81

Table 6:
8:30 and 10:00 announcement times: number of jumps observed in subperiods in greater detail

start of period	sampling frequency 5 min	 10 min	 15 min	 30 min
8:30	68	53	40	27
8:35	0	n.a.	n.a.	n.a.
8:40	0	3	n.a.	n.a.
10:00	0	14	14	15
10:05	0	n.a.	n.a.	n.a.
10:10	17	0	n.a.	n.a.

n.a. indicates not applicable

the different maturities do literally move together. As such, this analysis may be considered a direct test of the preferred location hypothesis. To this end, we gather together data on the precise timing of the price movements across all four markets following the news announcements. This identification is made difficult by the fact that there are multiple responses within a one second interval which are not distinctly identified by time of trade (although price and volume record distinct transactions) in the Cantor dataset. Keeping this difficulty in mind, we eschew a formal presentation of any results, instead opting to present casual observations on the nature of the data. Although it is quite difficult to generalise, there is some tendency for the 10 and 5 year maturities to respond prior to the other maturities. Further, all maturities respond to the news in the same direction within 45 seconds of the news announcements.

5.1 Scheduled versus unscheduled news

The timing of the cojumps is consistent with notion that there are large news effects on bonds stemming from macroeconomic announcements. This is not the entire story however, as there is also a number of cojumps that do not coincide with announcements. These are labelled 'other' in Table 5. The literature has examined the effects of unanticipated news by regressing a measure of the unanticipated component of news on the change in the price or yield of bonds of varying maturity over time periods around the news announcement. For example, Gurkaynak and Wolfers (2005) consider the change in bond yields from 5 minutes prior to the announcement to 25 minutes after, Campbell and Sharpe (2006) consider the change in bond yields from five minutes prior to ten minutes after the release and Balduzzi, Elton and Green (2001) use the change in bond price from 5 minutes before to 30 minutes after the announcement. Given that the impact of news is rapidly absorbed into prices, we concentrate on the returns to bonds in the 5 minutes immediately following the announcement, ie. 8.30 - 8.35. This return is calculated as the log price relative where the pre-announcement price is the last trade preceeding 8:30 and the the post-announcement price is the last trade in the interval. This appraoch is consistent with the yield and price change estimation technique used int the previous literature.

Figure 4a) illustrates the 5 minute returns occuring around each of the identified jumps across maturities. The horizontal axis is the bond maturity, and each point on a line indicates the observed 5 minute return after the news occurs. Figure 4b) gives the averages of the absolute 5 minute returns. The line labelled "total" is the average 5 minute return across all the observations for any given maturity in Figure 4a). Three other lines are also displayed: the lines labelled 8:30am and 10:00am are the average 5 minute returns associated with the news announcements at that time respectively, as identified in Table 5. Clearly the 8:30am news has relatively high impact, consistent with results of other research. The 10:00am news has a lower impact, lower than even the jumps at "other" times.

The biggest disruptions to the price process come with the release of macroeconomic news announcements at 8:30. Further analysis of this result can be undertaken by distinguishing between the particular types of news releases that occur at that time. Previous research has concluded that the most significant announcements for the bond markets are CPI, PPI, retail sales, housing starts and, in particular, non-farm payrolls. Thus, we focus on these announcements and Table 7 documents the number of news releases and summary information about the surprise content of each announcement type. The extent of the surprise here is measured as the difference between the average survey expectations data from Bloomberg and the actual vintage data release.

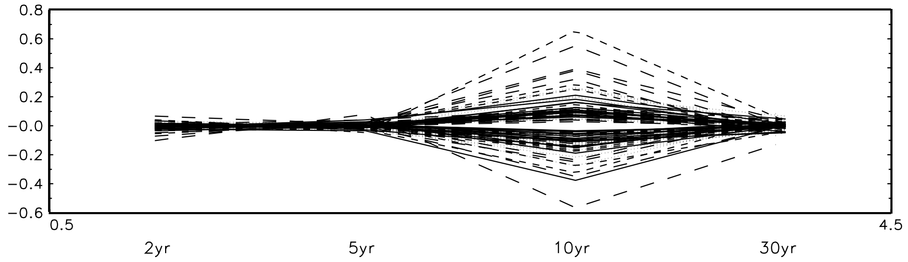
When we cross reference these news announcements to the cojumps across maturities, we find that the majority of jumps are associated with news releases, but also that there is still a large number of announcements that are not associated with jumps. Further, non-farm payrolls announcements produce the highest number of jumps across all four maturities, which is consistent with the previous literature.

It is intersting to note that the size of the surprise component in the announcement does not necessarily relate to the likelihood of a jump. Cross-referencing the unexpected component of the news release to the instances where there are jumps associated with the news releases, we find that the range of surprises in the announcments which do not have jumps exceeds those which do experience jumps for all but PPI. Further, the minimum surprise associated with a jump for each news release is less than the average of all the surprises for that announcement.

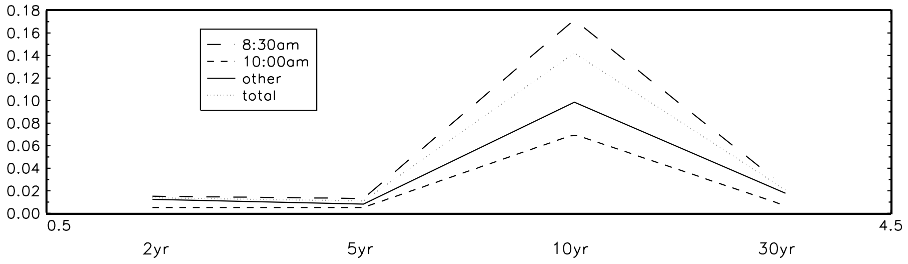
Thus, the scale of the unanticipated news does not clearly relate to disruption of the price process as evidenced by a jump. In fact, there are a number of occasions, particularly in the CPI and PPI releases, of zero surprise associated with jumps. Some

Figure 4: 5 minute returns in bonds across maturities in the period immediately associated with a detectable jump.

a) 5 minute returns in Jump interval



b) Abs(5 min return) Jump interval by time



of this may be related to problems with the measurement of surprise through potentially stale survey data. This issue is explored in Gurkaynak and Wolfers (2005), who suggest the use of derivative data as a better measure of expectations. This is scope for future work with this dataset, however we note that their results are typical of the previous literature and so we do not feel this is a likely explanation for our results. It is also clear from Table 7 that there is an asymmetry in the relationship between jumps and surprises, where we distinguish between positive and negative surprises in the announcement data. Jumps are more frequently associated with negative news for non-farm payrolls and PPI and more jumps are associated with positive news for CPI. The housing start and retail sales data however, are indistingushable.

Thus, the evidence here is not convincing that the extent of the surprise in the news event, or its direction, is solely responsible for significantly large disruptions in the price process. One possible avenue for further exploration of this issue would be to use the approach of Rigobon and Sack (2003) to more fully consider the differences in news announcements which do and do not result in jumps. This type of research would extend the work of Andersen, Bollerslev, Diebold and Vega (2006) who examine the impact of all news releases on bonds. We commend this as an area for future research.

6 Conclusion

High frequency data sets have confirmed that bond markets respond both strongly and quickly to the unanticipated component of macroeconomic news announcements. Evidence of this exists across the term structure, and generally finds that the impact increases with maturity, although there is some dissent. All the existing evidence proceeds on a univariate basis considering one maturity at a time.

This paper was concerned with significant disruptions to the frequently hypothesised continuous price process, as evidenced by statistically significant jumps in US Treasuries across maturities. Jumps were identified using univariate tests on each maturity. The multivariate analogue of these tests has not yet been developed, and we proposed a pragmatic and implementable alternative to identify cojumping across maturities using a coexceedance measure based on counting the number of contemporaneous jumps across maturities. Once days on which cojumping across the term structure

Table 7:
Surprises in scheduled news announcements and Jumps from 5 minute data sample, 2002-2006

	non-farm payrolls	CPI	PPI	retail sales	housing starts
number of announcements	59	58	59	59	58
surprise characteristics					
max abs(surprise)	318	0.3	1.2	1.5	256
min abs(surprise)	3	0	0	0	7.5
average abs(surprise)	67.83	0.10	0.37	0.39	90.56
jumps matching announcements	19	12	12	12	11
positive surprise	5	7	4	5	5
negative surprise	14	2	6	6	6
zero surprise	0	3	2	1	0
surprises with jumps characteristics					
max abs(surprise)	208	0.30	1.10	0.80	256
min abs(surprise)	9	0	0	0	13
average abs(surprise)	93.13	0.11	0.30	0.36	93.23

were identified we considered the evidence for those moves occuring simulatenously in the intraday data. A large proportion of cojumps in fact occurred contemporaneously within a day, and the vast majority of those occurred in the 5 minute period following a scheduled macroeconomic news announcement. This is consistent with the more general literature showing that bond markets respond to unanticipated components of news, or surpises. However, the jump response and news relationship displayed a number of interesting features. First, jumps were not necessarily associated with the larger news surprises, there were both larger surprises in the data not associated with jumps, and zero surprises which were associated with jumps. Although the average absolute surprise was likely to be larger with a jump than for the total sample, this was not true in all series, and the difference was relatively small in most cases. Further work is required to tease out the nature of the relationship of jumps with news. This may include obtaining better estimates of the surprise element of the news.

Overall the paper characterised jumps in the bond market as ocurring most frequently at the short end of the maturity structure, consistent with the market segmentation theory. However, the reduction in jumps with increasing maturity was not monotonic. The mid range maturities jumped less frequently than both the 2 and 30 year bonds. The somewhat higher jump rates in the 30 year bond are consistent with elements of the liquidity premium hypothesis. The size of jumps in response to news is also non monotonic. Although it generally increased over the 2 to 10 year maturities, with a pronounced maximum at 10 years, it declined again at 30 years. The relatively small impact at 30 years may reflect the high demand for this bond over the sample period.

There are a significant number of questions remaining to be addressed in this research agenda. These include the development of a truly multivariate jumps test, a more thorough examination of the sensitivity of the jumps to surprise news announcements through other measures of surprise and further work on the issue of optimal sampling frequency. In addition it is of considerable interest to explore further the cojumping instances not associated with discernible scheduled macroeconomic news announcements and instances of cojumping which do not extend across the entire term structure.

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