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# Resolving Macroeconomic Uncertainty in Stock and Bond Markets*

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**Abstract.** We establish an empirical link between the ex-ante uncertainty about macroeconomic fundamentals and the ex-post resolution of this uncertainty in financial markets. We measure macroeconomic uncertainty using prices of economic derivatives and relate this measure to changes in implied volatilities of stock and bond options when the economic data is released. Higher macroeconomic uncertainty is associated with greater reduction in implied volatilities following the news release. It is also associated with increased volume and decreased open interest in option markets after the release, consistent with market participants using financial options to hedge or speculate on macroeconomic news.

## 1 Introduction

How important is uncertainty about macroeconomic fundamentals for financial markets?[1] The literature has tried to answer this question indirectly by measuring the response of asset prices, including those of derivatives, to macroeconomic

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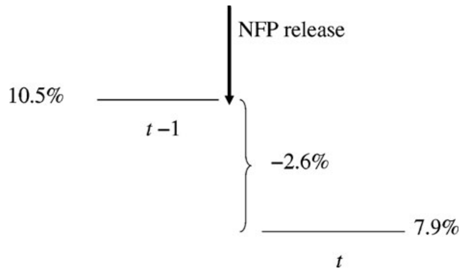
[1] We use the term uncertainty in the context of market participants being unsure about the current state of the economy. We do not take an explicit stance of whether the probability of outcomes can be quantified or not, as in the Knightian uncertainty sense.

announcements.[2] Evidence that new information about the economy matters for financial markets implies that uncertainty in these markets should be associated with uncertainty about the state of the economy. Consistent with this reasoning, Ederington and Lee (1996) and Beber and Brandt (2006) document that the uncertainty implicit in options written on Treasury securities drops substantially after the release of macroeconomic news. For example, the following diagram illustrates the unconditional one-day drop in annualized implied volatility of options on Treasury bond futures associated with the release of the non-farm payrolls in our sample:




This observation implies that as financial markets learn about the state of the economy through the employment report, uncertainty in financial markets is resolved. Motivated by this indicative evidence, the main goal of this paper is to examine how and to what extent the ex-post resolution of uncertainty in financial markets is related to the ex-ante uncertainty of market participants about the state of the economy.

Uncertainty about the state of the economy is unobserved and therefore inherently difficult to quantify. Previous studies have used measures of cross-sectional disagreement, such as the standard deviation of forecasts across economists, to proxy for uncertainty (e.g., Andersen et al., 2003), but such disagreement measures are in many cases only weakly correlated with measures of true uncertainty (see Zarnowitz and Lambros, 1987). We establish a more direct link between the ex-ante uncertainty of macroeconomic fundamentals and the ex-post resolution of uncertainty in financial markets. We measure macroeconomic uncertainty using prices of economic derivatives traded in a new auction-based market launched in 2002 jointly by Goldman Sachs and Deutsche Bank. The underlying is the initial release of a given macroeconomic statistic on the scheduled announcement date. These economic derivatives represent explicit bets on news about macroeconomic fundamentals and their option-implied second moments therefore provide a direct measure of how unsure market participants are about the news release.

[2] Several recent papers have investigated the response of bond, stock, and currency markets to scheduled macroeconomic announcements (e.g., Balduzzi, et al., 2001; Flannery and Protopapadakis, 2002; Andersen et al., 2003). The results reveal a significant impact of economic announcements on returns, their volatility, and market liquidity.

Although the economic derivatives market is relatively young, it is already closely watched by market participants and the media before scheduled macroeconomic announcements as a barometer of market views. For example, Bloomberg News reported the following before a recent announcement of U.S. payroll statistics:[3]

> . . . U.S. employers probably added 170,000 jobs last month, according to the median estimate of 75 economists surveyed by Bloomberg News. The Labor Department releases the figures at 8:30 a.m. in Washington. Traders expect 186,200 jobs new in April, an auction of economic derivatives showed. The derivatives, created by Deutsche Bank AG and Goldman Sachs Group Inc and marketed through ICAP Plc, were auctioned yesterday and today.

Besides this high degree of visibility, economic derivatives are likely to provide an accurate measure of ex-ante uncertainty about the state of the economy for at least two other reasons.[4] First, the market for economic derivatives is dominated by sophisticated investors, predominantly hedge funds and proprietary trading desks. Second, auctions take place only one day to a few hours prior to the announcement, leaving little room for market views to change before the subsequent news release.[5]

Our paper makes at least two contributions to the literature. First, we establish explicitly the link between the degree of ex-ante uncertainty about economic fundamentals implied by economic derivatives prices and the ex-post resolution of this uncertainty in financial markets. Based on our results, we can predict, for instance, how a temporary or structural change in the uncertainty about economic fundamentals will affect the volatility of bonds, stock indices, and individual stocks. There are several reasons for the uncertainty about economic fundamentals to change, including the economy transitioning into a different phase of the business cycle (a temporary change) or the Government changing its statistical reporting protocol (a structural change). Second, we examine how the ex-ante uncertainty about economic fundamentals is related to trading activity in financial markets, as measured by open interest and volume, before and after the announcement. Trading activity helps us understand how market participants deal with (meaning, do they hedge or speculate?) uncertainty about the state of the economy.

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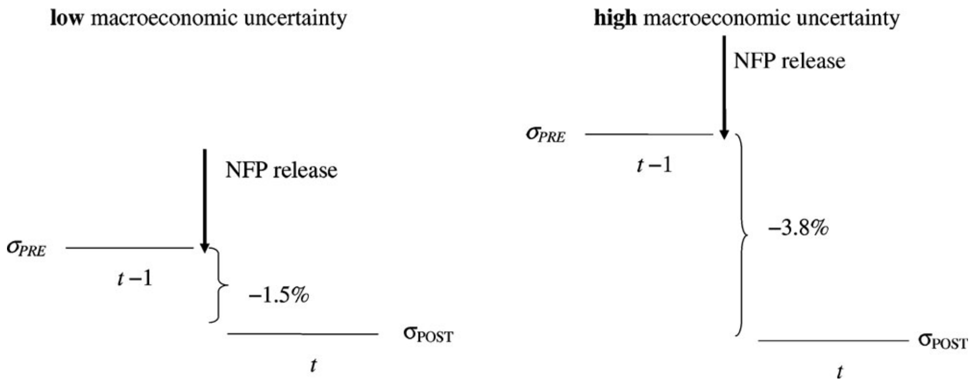
[3] Source: "Dollar advances on speculation job growth may exceed forecasts," Bloomberg News, May 7, 2004. Other headlines for this announcement include ". . . economic derivatives offered by Deutsche Bank and Goldman Sachs showed markets betting on Friday for February jobs growth of around 137,000. Markets had been going for about 145,000 on Thursday." ("Dollar firm while U.S. jobs data looms," Reuters News, March 5, 2004) and ". . . economists expected February payrolls to rise 130,000, according to a survey conducted by CBS MarketWatch. In the economic derivatives market run by Goldman Sachs, Deutsche Bank and ICAP were looking for a gain of about 140,000 new positions." (CBS/AP, March 5, 2004).

[4] In a concurrent paper, Gürkaynak and Wolfers (2005) show that expectations derived from economic derivatives are somewhat more accurate than survey-based forecasts.

[5] In contrast, survey expectations are often up to a week old.

The design of our empirical analysis is straightforward. We first confirm that stock and bond markets react to macroeconomic announcements in our sample period. We then extract a model-free measure of macroeconomic uncertainty from the observed prices of economic derivatives. Finally, we relate macroeconomic uncertainty to the changes in model-free implied variances of stock and bond options when the economic data is released as well as to the changes in transacted volume and open interest in stock and bond option markets.

We find that higher ex-ante uncertainty about macroeconomic fundamentals is associated with greater reduction in the implied volatility of stock and bond options when the economic data are released. The results are more pronounced for bonds than for stocks, and the effect appears to be permanent. Specifically, we observe that the degree of uncertainty about the non-farm payrolls explains about 40 percent of the variation in the change of the volatility of Treasury bond futures and about 15 percent of the change in the volatility of Eurodollar futures. Furthermore, this effect is also strong economically. For example, a high degree of uncertainty equal to one standard deviation above the average macroeconomic uncertainty in our sample predicts more than a one-third drop in the implied volatility of bond options, compared to less than a 15 percent drop when macroeconomic uncertainty is low (one standard deviation below the average). The following diagram illustrates this effect:




where $\sigma_{PRE}$ and $\sigma_{POST}$ denote the implied volatilities before and after the release. We also observe an association, although weaker, between macroeconomic uncertainty and the implied volatility of stock index options. Finally, we document that within the stock index, cyclical stocks are substantially more exposed to macroeconomic uncertainty than non-cyclical stocks.

Our results take advantage of the tight identification scheme related to precise measures of events (the announcements), macroeconomic uncertainty (economic derivatives) and financial market uncertainty (implied volatility). We can thus draw important implications for finance models that lower frequency empirical approaches lack (e.g., Schwert, 1989). For example, stochastic volatility models of bond returns should include uncertainty about macro fundamentals as a key state variable. Similarly, models of the volatility of stock returns should contain loadings on macro-uncertainty that depend on the sensitivity of cash flows to the business cycle.

Concerning the link between macroeconomic uncertainty and trading behavior, we find evidence that the degree of uncertainty affects significantly the trading strategies employed by market participants for both options written on Treasury bond futures and on cyclical stocks. In particular, higher macroeconomic uncertainty is associated with a greater increase in transacted volume after the news release. This observation is consistent with market participants entering hedging or speculative positions gradually before and unwind those positions shortly after the news release. We also find a negative relation between macroeconomic uncertainty and the reduction of open interest after the news release. This result is consistent with market participants closing out hedging or speculative positions and with the degree of these activities depending on the degree of macroeconomic uncertainty to be hedged or to profit from.

Besides contributing to the growing literature on the response of financial markets to macroeconomic news (e.g., Ederington and Lee, 1996; Beber and Brandt, 2006; and the references in footnote 2), our paper relates to the research examining the economic sources of return volatility. David and Veronesi (2002) show that an uncertainty measure obtained from a model of real earnings growth is related to the implied volatilities of equity options. Buraschi and Jiltsov (2006) find that a proxy for the differences in beliefs forecasts the implied volatility of index options and is related to index options volume. Patell and Wolfson (1981) and Dubinsky and Johannes (2006) study the evolution of uncertainty around earnings announcements using the implied volatilities of individual equity options. We add to this literature by employing an economic uncertainty measure that is based on prices of economic derivatives and is thus market-based, high-frequency, and intrinsically related to the macro economy. These features of our uncertainty measure allow us to uncover a strong relation between the uncertainty of economic fundamentals and financial markets volatility. We also demonstrate how the link between macroeconomic fundamentals and financial markets depends on the type of assets, with long-term bonds at one extreme and non-cyclical stocks at the other.

The paper proceeds as follows. In Section 2, we sketch out a theoretical framework to help guide our empirical investigation. In Section 3, we describe the economic derivatives and financial options data. Section 4 explains our methodology for estimating the uncertainty implicit in economic derivatives and how we relate this measure to the uncertainty in financial markets and to the trading strategies of options market participants. We present our empirical results in Section 5. Section 6 concludes with a summary of our findings.

## 2 Theory

In this Section, we sketch out a general asset pricing model for bonds and equities in the presence of macroeconomic news releases that allows us to illustrate the link between the uncertainty about macroeconomic fundamentals and the changes in bond and stock return volatilities. Our model combines the bond pricing with announcement effects of Piazzesi (2001) with the equities modeling framework of Mamaysky (2002). Given this straightforward combination of models, we keep the discussion relatively free of technical details and instead refer the interested reader to these two papers. Our discussion is organized into four parts. The first three parts specify the asset pricing model, characterize the state variables, and present the bond and stock pricing results. In the last part, we discuss the empirical predictions of the model (which we ultimately link to our empirical measures of uncertainty in Section 4.3 once we described the data).

### 2.1 ECONOMIC STATE VARIABLES AND INFORMATION FLOW

Consider an economy described by a $N + M$ dimensional vector of state variables $X(t)$ that drives the payoffs of all assets. We write $X(t)$ as:

$$X(t) = \begin{bmatrix} Y(t) \\ Z(t) \end{bmatrix},$$

for an $N$-dimensional vector $Y(t)$ and an $M$-dimensional vector $Z(t)$. As in Mamaysky (2002), we assume that bond prices depend only on the $Y$-type state variables, while stock prices depend on both $Y$-type and $Z$-type state variables. A natural interpretation of this decomposition is that the $Y$-type state variables reflect information about the macroeconomy while the $Z$-type state variables capture information orthogonal to the macroeconomy but still relevant for stock prices, such as non-macroeconomic related information on cashflows.

Bond prices can be obtained for quite a general $Y$ process (Duffie and Kan, 1996), but for simplicity we specify $Y$ to represent conveniently the pattern of macroeconomic news releases. Specifically, adopting the idea of Piazzesi (2001), the dynamics of the state variables $Y(t)$ are:

$$dY(t) = \mu(Y(t))dt + \sigma(Y(t))dW(t) + dJ(t), \tag{1}$$

where $W(t)$ is a vector of Brownian motions, $\mu$ is the drift, $\sigma$ is the diffusion matrix, and $J$ is a pure jump process. The jump process $J$ is activated at deterministic points in time, corresponding in our setting to scheduled macroeconomic announcements.[6]

[6] The idea that a deterministic jump process can capture scheduled macroeconomic releases is present also in Piazzesi (2001), Piazzesi (2005a), and Beber and Brandt (2006). Maheu and McCurdy (2004) and Dubinsky and Johannes (2006) model earning announcements with deterministic jumps.

These jump times are recorded by a deterministic counting process $N^{\mathrm{D}}$ starting at zero and counting up in increments of one. The drift and diffusion matrix of $Y(t)$ are given by:

$$\mu(Y(t)) = K_Y(\Theta - Y(t)) \tag{2}$$

$$\sigma(Y(t)) = \Sigma_Y S(Y(t)), \tag{3}$$

where $S(Y)$ is a diagonal $N \times N$ matrix with $i$-th diagonal element $S_i(Y) = \sqrt{S_{0i} + S_{1i}^{\top} Y}$ and where $S_{0i} \in \mathbb{R}$; $\Theta$, $S_{1i} \in \mathbb{R}^N$; $\Sigma_Y$, $K_Y \in \mathbb{R}^{N\times N}$ are constants. For any deterministic jump at time $\tau$ counted by the deterministic counting process, the jump size conditional on information right before the jump has a potentially multivariate Gaussian distribution $J \sim N(-1/2\sigma_J r, \sigma_J^2)$ uncorrelated with $W(t)$.[7]

Stocks are likely to depend on additional factors with respect to bonds, the $Z$-type state variables. The dynamics of $Z(t)$ are given by:

$$dZ(t) = \mu_Z dt - K_Z Y(t) dt + \Sigma_Z dW(t), \tag{4}$$

where $\mu_Z \in \mathbb{R}^M$, $K_Z \in \mathbb{R}^{M\times N}$, and $\Sigma_Z \in \mathbb{R}^{M\times(N+M)}$. $W(t)$ is a $N+M$ dimensional process of standard, independent Brownian motions.

### 2.2 BOND PRICES

For simplicity, we model bonds as zero-coupon securities. More specifically, the payoff of bonds is one unit at the terminal date $T$, so their price is given by:

$$P(t, T) = \mathbb{E}_t^Q \left[ \exp\left( -\int_t^T r(u) du \right) \right], \tag{5}$$

where $r(t)$ is the short term interest rate and whose expectation is taken under the risk-neutral probability measure $Q$.[8] The specification of the short rate process that allows us to solve for the expectation in (5) is given by

$$r(t) = r_0 + r_Y^{\top} Y(t) \tag{6}$$

The addition of jumps with stochastic intensities would make the theory more complex without meaningfully changing the model implications, because we take differences on the announcement days.

[7] Our assumption about the conditional distribution of the jump size is convenient for expositional reasons, but in principle it could be more general. In particular, Piazzesi (2001) shows that the jump size may have a conditional state-dependent distribution provided it has an exponential-affine Laplace transform.

[8] We assume the usual regularity conditions. Standard results show that if there exists a risk-neutral probability measure $Q$, the set of asset prices is arbitrage free. The converse is also true under reasonable restrictions on trading strategies. Duffie (2001), among others, provides details on and references for these results.

for $r_0 \in \mathbb{R}$ and $r_Y \in \mathbb{R}^N$. According to Piazzesi (2005a), the solution for bond prices in such a setting is an exponentially affine function in the state variables. The price of a bond at time $t$ that matures at time $T$ is given by:

$$P(Y, t, T) = \exp(a(t, T) + b(t, T)^\top Y). \tag{7}$$

The computation of $a(t,T)$ and $b(t,T)$ proceeds recursively backward through time, starting at the maturity date with boundary conditions $a(T, T) = 0$ and $b(T, T) = 0$.[9]

### 2.3 STOCK PRICES

Stocks are essentially priced analogously to bonds. The key ingredient for stocks is to specify a dividend process that leads to a tractable pricing model. Mamaysky (2002) shows that a particular choice of intermediate and terminal dividend processes leads to a very simple solution for the stock price.[10] More specifically, the stock price $S_i(t)$ is also affine and is given by:

$$S(Y, Z, t) = \exp\left(\alpha_i t - \beta_i^\top Y(t) - C_i^\top Z(t)\right), \tag{8}$$

where $\alpha_i$ and $\beta_i$ are solutions to a system of ODEs.

The subscript $i$ refers to different classes of stocks. This is an important feature of the model, because the stock prices of different types of stocks can have a different loading on the state variables. This characteristic is crucial in our empirical investigation of the class of cyclical versus non-cyclical stocks.

### 2.4 IMPLICATIONS FOR UNCERTAINTY

Within the asset pricing model described in the previous subsections and in particular equation (7), Ito's lemma implies that the variance of log bond returns is

$$\mathrm{Var}_{P(Y,t,T)} = b(t, T)^\top \mathrm{Var}(Y) b(t, T), \tag{9}$$

where

$$\mathrm{Var}(Y) = \sigma^2(Y) + \sum_{j=1}^{N^d} \sigma_j^2. \tag{10}$$

[9] The result of Piazzesi (2005a) is based on Piazzesi (2001), which basically shows that the bond price is exponentially affine just before the jump date and during the interim period between two deterministic jump dates.

[10] The crucial feature of the dividend process that leads to closed form solutions is that the intermediate dividend is equal to the stock price multiplied by an affine dividend yield.

For the purpose of pricing options with maturity $T^*$, the expected risk-neutral variance of continuous-time returns is thus going to be:

$$\mathrm{Var}_{P(Y,t,T,T^*)} = b(t,T)^\top \sigma^2(Y) b(t,T) + \frac{b(t,T)^\top \left(\sum_{j=1}^{N^d_{T^*}} \sigma_j^2\right) b(t,T)}{T^*} \tag{11}$$

where $N^d_{T^*}$ counts the number of deterministic jumps until the maturity of the option.

For simplicity, suppose there is only one jump at the time $\tau$ of a macroeconomic release, between $t$ and $T^*$. The expected risk neutral variance at $\tau-$ is:

$$\mathrm{Var}_{P(Y,\tau-,T,T^*)} = b(\tau-,T)^\top \sigma^2(Y) b(\tau-,T) + \frac{b(\tau-,T)^\top \sigma_{j1}^2 b(\tau-,T)}{T^* - \tau}. \tag{12}$$

The expected risk neutral variance after the jump is:

$$\mathrm{Var}_{P(Y,\tau,T,T^*)} = b(\tau,T)^\top \sigma^2(Y) b(\tau,T). \tag{13}$$

The comparison of equations (12) and (13) shows that the bulk of the difference in the bond risk-neutral variances between $\tau$ and $\tau-$ is explained by the variance of the jump sizes, $\sigma_{j1}^2$ scaled by the maturity of the option and by the outer product with the affine coefficient $b(\tau-,T)$. In general, the relation between the affine coefficient $b$ and the bond maturity depends on parameter estimates and is different for different state variables.[11] Therefore, the model does not have explicit implications for uncertainty at different bond maturities.

We can use the same logic for the stock price defined in equation (8). Using Ito's lemma, the variance of stock prices is:

$$\mathrm{Var}_{S(Y,Z,t)} = \beta_i(t,T)^\top \mathrm{Var}(Y) \beta_i(t,T) + C_i^\top \mathrm{Var}(Z) C_i, \tag{14}$$

where $\mathrm{Var}(Y)$ is as in equation (10) and $\mathrm{Var}(Z) = \sigma^2(Z)$. Similarly to the bond case, for the purpose of pricing options with maturity $T^*$, we can express the risk-neutral variance just before a deterministic jump at $\tau$ and after the jump:

$$\begin{aligned}\mathrm{Var}_{S(Y,Z,\tau-,T,T^*)} = {} & \beta_i(\tau-,T)^\top \sigma^2(Y) \beta_i(\tau-,T) \\ & + \frac{\beta_i(\tau-,T)^\top \sigma_{j1}^2 \beta_i(\tau-,T)}{T^* - \tau} + C_i^\top \sigma^2(Z) C_i.\end{aligned} \tag{15}$$

The expected risk neutral variance after the jump is:

$$\mathrm{Var}_{S(Y,Z,\tau,T,T^*)} = \beta_i(\tau,T)^\top \sigma^2(Y) \beta_i(\tau,T) + C_i^\top \sigma^2(Z) C_i. \tag{16}$$

[11] For example, Piazzesi (2005) estimates an affine bond pricing model with macroeconomic shocks that could be an appropriate benchmark in our setting. In her model estimated with U.S. data between 1994 and 1998, the relation between yield responses and bond maturities is decreasing, increasing, or hump-shaped, depending on the state variable that experiences a shock.

Again, the comparison of equations (15) and (16) shows that the bulk of the difference in the stock risk-neutral variances between $\tau$ and $\tau$- is explained by the variance of the jump size, scaled by the maturity of the option and by the outer product of the affine coefficient $\beta_i(\tau-, T)$. The subscript $i$ indicates that the resolution of uncertainty can be different for different classes of stocks (e.g., cyclical versus non-cyclical stocks).

The empirical analysis in the following sections of the paper tests the simple model predictions on whether the resolution of uncertainty in bond and stock markets is explained by the variance of the jump size in the state variables. Section 4 specifically formalizes the link between the model and the empirical specifications.

However, the model does not have explicit predictions on the relative magnitude of the coefficients (e.g., bonds versus stocks), except in very specific cases.[12] It is thus going to be an empirical question to establish if the resolution of uncertainty is stronger for bonds versus stocks or, within the stock asset class, for cyclical versus non-cyclical stocks. Still, simple economic intuition from present value type of models (e.g., Campbell and Mei, 1993) provides some guidance on the expected empirical results. If we think about macroeconomic announcements conveying both cash flow and discount rate news, we are likely to observe different patterns in bond versus stock prices (see Boyd, Hu, and Jagannathan, 2005). Risk-free government bond prices will not be affected by cash flow news and thus macroeconomic news will have unambiguous effects. In contrast, we could observe offsetting effects on stock prices whenever discount rate effects and cash flow effects are in the same direction (e.g., signals of overheating economy leading to higher discount rates and higher cash flows). In this case, the affine coefficient in the bond pricing equation $b$ will be higher than the affine coefficient in the stock pricing equation $\beta_i$, and we should thus observe a stronger effect of uncertainty resolution in bond versus stock markets. Within the stock market class, the cash flow effect is going to be more important for cyclical stocks. If this effect is strong enough to more than offset the discount rate effects, $\beta_i$ for cyclical stocks would be larger than $\beta_i$ for non-cyclical stocks and the resolution of uncertainty would play a more important role for cyclical stocks. Naturally, all these predictions hinge on the idea that discount rate effects and cash flow effects triggered by macroeconomic news are in the same direction. We could certainly envision business cycle phases where lower discount rates are associated with positive cash flow effects and in these cases our predictions would not be valid.

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[12] For example, Mamaysky (2002a) assumes that the state variables $Y$ and $Z$ follow Ohrnstein-Uhlenbeck processes without jumps. In this particular case, it can be shown that the affine coefficient $b$ on bond prices is generally larger than the coefficient $\beta$ on stock prices.

## 3 Data

### 3.1 ECONOMIC DERIVATIVES

Goldman Sachs and Deutsche Bank launched the market for derivative securities on scheduled macroeconomic announcements, covering initially non-farm payrolls (NFP), the Institute for Supply Management (ISM) manufacturing report, and U.S. retail sales ex-autos (RS) in October 2002.[13] The liquidity of this market was enhanced shortly thereafter through an agreement with ICAP (the largest interdealer broker for over-the-counter derivatives) to distribute the securities to the inter-dealer market.[14] Economic derivatives are priced in a parimutuel auction in which as much volume as possible is filled at the same clearing price.[15] This auction format is designed to maximize liquidity.

We collect data on auctions of economic derivatives completed between October 2002 and August 2006 from Goldman Sachs. Our sample consists of 233 auctions covering 47 releases of non-farm payroll, 43 releases of the Institute of Supply Management manufacturing report, 40 releases of retail sales ex-autos, and 103 releases of initial jobless claims. In the remainder of the paper, we focus exclusively on the release of non-farm payroll for a number of reasons. First, there is widespread evidence in the literature that this is the most influential macroeconomic announcement. Second, the time-series of economic derivative auctions on the non-farm payroll release is the longest available. Finally, we gather through informal conversations with Goldman Sachs that there was always active participation for auctions on non-farm payroll announcements during all the sample period.

Economic derivatives are structured as calls, puts, and digital options. These securities can be traded by themselves or in combinations such as spreads, straddles, strangles, and risk reversals. The strike prices are set to reflect the range of possible outcomes. The payoffs of the call (put) options are capped at the highest (lowest) available strike price. The underlying is the initial release of a given macroeconomic

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[13] Contracts on other releases were introduced later. Specifically, contracts on the Eurozone harmonized index of consumer prices (HICP) were introduced in May 2003, contracts on initial job claims started trading in February 2004, contracts on the U.S. gross domestic product started trading in January 2005, contracts on the U.S. international trade balance were introduced in February 2005, and contracts on the U.S. consumer price index started trading in June 2006.

[14] The Chicago Mercantile Exchange (CME) and Goldman Sachs announced a partnership in June 2005 to further enhance the liquidity of economic derivatives through integrated clearing and marketing agreements.

[15] All trades at a given strike are executed at the same price, as it is common in Dutch auctions. However, in the case of economic derivatives, it is possible to enter limit orders. The equilibrium price is determined by an auction-clearing algorithm that maximizes total trades. The parimutual auction mechanism and some of the operational details are described in Baron and Lange (2006).

statistic on the scheduled announcement date. Revisions of the data that typically follow the initial release do not matter for the payoffs of economic derivatives. The client base for economic derivatives is primarily hedge funds, proprietary trading desks, and inflation swap traders.

There are typically two auctions for each non-farm payroll announcement. At the beginning of the sample, the first auction was held three days before the release and the second auction was held one day before the release. After experimenting for a short period with only one auction the day before the release, there are then again two auctions for most of the sample. The first auction is held in the afternoon of the day before the release and the second auction is held in the early morning of the day of the release. Anecdotal evidence suggests that the participation in the first auction is mainly from U.S. customers, whereas the second auction gets slightly more than half U.K. and European participation.[16] The increasing interest for non-farm payroll derivatives led Goldman Sachs to add two more auctions in the morning of one and two days before the release in the last part of the sample.

We do not have figures on transacted volume. However, we gather through informal conversations with Goldman Sachs that digital options are about three quarters of the trades yet less than half of the transacted volume. Vanilla options are generally preferred by hedgers and are therefore transacted in greater size and volume. Figure 1 provides a snapshot of our data for one auction.

### 3.2 FIXED INCOME OPTIONS

We collect daily prices for options on the 30-year Treasury bond futures, options on the 10-year Treasury note futures, and options on the 5-year Treasury note futures, all traded at CBOT, for the sample period of the economic derivatives data. We collect options prices on the Eurodollar futures traded at CME for the same period. We also collect daily data on transacted volume and open interest for each option contract written on the 30-year Treasury bond futures, 10-year Treasury note futures, 5-year Treasury note futures, and Eurodollar futures from the CBOT and CME.

### 3.3 EQUITY OPTIONS

We also collect daily data on S&P 500 index options as well as on individual equity options from OptionMetrics. The data consists of midpoint quotes at the

[16] This geographical breakdown in auction participation is the outcome of the different time-zones, where the first auction is less convenient for European customers and the second is less convenient for U.S. customers.

Goldman Sachs US Non-Farm Payrolls Jun '04

Final Report
12:53

Auction Date: 2Jul04
Implied Forecast: 223.04

**Auction Prices for simple Digital and Vanilla Options**

| | Digital | | | | | | Vanilla | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | Calls | | | Puts | | | Calls | | | Puts | | |
| Strike | Clears At | Sell At | Buy At | Clears At | Sell At | Buy At | Clears At | Sell At | Buy At | Clears At | Sell At | Buy At |
| 75 | 0.946063 | 0.936063 | 0.956063 | 0.053937 | 0.043937 | 0.063937 | 149.817 | 146.317 | 153.317 | N/A | N/A | N/A |
| 100 | 0.92 | 0.91 | 0.93 | 0.08 | 0.07 | 0.09 | 126.504 | 123.254 | 129.754 | 1.6872 | 1.4372 | 1.9372 |
| 125 | 0.853 | 0.843 | 0.863 | 0.147 | 0.137 | 0.157 | 104.375 | 101.375 | 107.375 | 4.5582 | 4.0582 | 5.0582 |
| 150 | 0.81 | 0.8 | 0.82 | 0.19 | 0.18 | 0.2 | 83.609 | 80.859 | 86.359 | 8.7922 | 8.0422 | 9.5422 |
| 175 | 0.68 | 0.67 | 0.69 | 0.32 | 0.31 | 0.33 | 65.049 | 62.549 | 67.549 | 15.2322 | 14.2322 | 16.2322 |
| 200 | 0.58 | 0.57 | 0.59 | 0.42 | 0.41 | 0.43 | 49.349 | 47.099 | 51.599 | 24.532 | 23.282 | 25.782 |
| 225 | 0.508 | 0.498 | 0.518 | 0.492 | 0.482 | 0.502 | 35.785 | 33.785 | 37.785 | 35.968 | 34.468 | 37.468 |
| 250 | 0.358 | 0.348 | 0.368 | 0.642 | 0.632 | 0.652 | 25.035 | 23.285 | 26.785 | 50.218 | 48.468 | 51.968 |
| 275 | 0.286 | 0.276 | 0.296 | 0.714 | 0.704 | 0.724 | 17.021 | 15.521 | 18.521 | 67.204 | 65.204 | 69.204 |
| 300 | 0.22 | 0.21 | 0.23 | 0.78 | 0.77 | 0.79 | 10.729 | 9.479 | 11.979 | 85.912 | 83.662 | 88.162 |
| 325 | 0.152 | 0.142 | 0.162 | 0.848 | 0.838 | 0.858 | 6.1128 | 5.1128 | 7.1128 | 106.296 | 103.796 | 108.796 |
| 350 | 0.085804 | 0.075804 | 0.095804 | 0.914196 | 0.904196 | 0.924196 | 3.1733 | 2.4233 | 3.9233 | 128.357 | 125.607 | 131.107 |
| 375 | 0.051804 | 0.041804 | 0.061804 | 0.948196 | 0.938196 | 0.958196 | 1.4702 | 0.9702 | 1.9702 | 151.654 | 148.654 | 154.654 |
| 400 | 0.027804 | 0.017804 | 0.037804 | 0.972196 | 0.962196 | 0.982196 | 0.4871 | 0.2371 | 0.7371 | 175.671 | 172.421 | 178.921 |
| 425 | 0.011804 | 0.001804 | 0.021804 | 0.988196 | 0.978196 | 0.998196 | N/A | N/A | N/A | 200.183 | 196.683 | 203.683 |

*Figure 1.* Data.
This figure shows an example of the data source used to obtain economic derivatives prices. Goldman Sachs compiles a report at the end of each auction on U.S. Non-Farm Payrolls containing the final prices of vanilla and digital call and put options, for each level of strike prices. The clearing prices are the final outcome of the auction. In contrast, the ask and bid prices are automatically generated by adding and subtracting a fixed fee from the clearing prices.

close, implied volatilities and standard sensitivity measures. Due to the delay in the database update, this part of our data only goes through April 2006.

In contrast to index options, options markets for many individual firms are very illiquid. In order to assure a minimum level of liquidity and hence data integrity, we focus exclusively on individual equity options for firms that are part of the S&P 100 index, since an explicit requirement for membership in this index is to have a liquid options market.[17] We then sort stocks into industries and further into groups of industries that we anticipate to be cyclical or non-cyclical. Specifically, we base our industry classification on the results of Boudoukh et al. (1994), who sort industries by their correlation between industry level output growth and aggregate output growth. We label the five industries with the highest output growth beta as cyclical and the five industries with the lowest output growth beta as non-cyclical.[18] This sorting procedure results in a sample of individual equity options on 22 cyclical firms and on 14 non-cyclical firms.

[17] See the S&P 100 index fact sheet. Driessen et al. (2008) use the same selection criteria.
[18] We use the output growth betas from Table 1 of Boudoukh et al. (1994). The cyclical industries are primary metals, transportation equipment, rubber and plastics, metal products, and electrical machinery. The non-cyclical industries are food and beverage, tobacco, utilities, printing and publishing, and petroleum products.

## 4 Methodology

In this section, we explain the methodology to obtain empirical measures of macroeconomic uncertainty and financial markets uncertainty. We then relate these empirical proxies with the theoretical predictions of Section 2.

### 4.1 OBTAINING MACROECONOMIC UNCERTAINTY

We infer the implied uncertainty of the macroeconomic news release from the prices of vanilla options. We do not consider the prices of digital options, because the auction pricing system is designed such that the prices of digital options become redundant subject to minimal microstructure noise.[19] To compute the option-implied uncertainty, we use the model-free approach of Britten-Jones and Neuberger (2000). Unlike the traditional concept of implied volatility, this methodology does not rely on any specific option pricing model or distributional assumption and corresponds to the risk-neutral integrated variance.

We denote with $C(T,K)$ the price of a vanilla call option on the economic statistic $F$, with expiration date $T$ (the release date) and with strike price $K$. Britten-Jones and Neuberger (2000) show that the integrated variance between the current date and $T$ is fully specified by the set of call options expiring on $T$:

$$\mathrm{E}_0^Q\left[\int_0^T \left(\frac{dF_t}{F_t}\right)^2\right] = 2\int_0^\infty \frac{C(T,K) - \max(0, F_0 - K)}{K^2} dK, \tag{17}$$

where $F_0$ denotes the forward price. This approach can easily be adapted to include vanilla put options. We will refer to equation (17) as the model-free implied economic variance $\sigma^2_{ECO}$ and its square root as the model-free implied economic volatility $\sigma_{ECO}$.

Jiang and Tian (2005) show that the integral over a continuum of strike prices in (17) can be approximated accurately by a sum over a finite number of strike prices.[20] The implementation with economic derivatives requires three straightforward variations. First, we need to determine the expected outcome of the release, i.e. the forward price $F_0$ in equation (17). We use the forward price implied by the prices of vanilla call and put options and the put-call parity relation.[21] The

[19] The replication algorithm obtains vanilla options as combination of strings of knock-out calls and puts with digital options. We thank Ken Baron and Darrell Duffie for clarifications on this auction mechanism.

[20] When the underlying price process includes jumps, the strategy of approximating the variance of the underlying price process with log contracts fails to replicate perfectly. However, Carr and Wu (2004) and Jiang and Tian (2005) show that the approximation error in the model-free variance is small under realistic price processes and market settings.

[21] Alternatively, we obtain the forward price by solving a non-linear least squares problem with option market prices and model prices. We formulate an option pricing model imposing different

forward price easily inferred from economic derivatives prices is key information, because it represents a market-based expectation of the upcoming macroeconomic announcement to be confronted with analyst median expectations.

Second, we need to consider that the vanilla call and put options have caps and floors at the highest and lowest strike prices. If the probability in the tails of the distribution is very asymmetric, the estimate of $\sigma^2_{ECO}$ could be biased. We thus adjust all the call and put options prices by adding the value of a vanilla call written at the cap and a vanilla put written at the floor, respectively. We determine the value of a call with strike price at the cap using a Black-Scholes like option pricing formula, where we plug in the implied volatility at the highest strike price. The value of a put option with strike price at the floor is obtained analogously using the implied volatility at the lowest strike price. This adjustment is thus robust to the presence of implied volatility smiles in the economic derivatives market.

Finally, we take into account that the underlying asset, the release of non-farm payrolls, is measured in changes, but equation (17) is generally implemented in levels. We thus transform the unity of measure of option prices, strike prices, and forward prices from changes to levels. We obtain the model-free implied economic volatility from a discretized version of equation (17) and then transform it back into changes. At the end, we obtain an economically meaningful measure of uncertainty, in that $\sigma_{ECO}$ represents economic uncertainty expressed in thousands of jobs.

Table I, panel A, presents summary statistics for $\sigma_{ECO}$. During our sample period, economic uncertainty shows substantial time-variation, ranging from 59 to 134 thousand jobs, with a mean and median around 95 thousand jobs.

### 4.2 OBTAINING FINANCIAL MARKETS UNCERTAINTY

For each underlying fixed-income security, we obtain a measure of bond market uncertainty by computing the model-free implied variance described in equation (17). We obtain two series of model-free implied variance. The first series uses option settlement prices from the nearest expiry month. This series switches to the next available month on the first day of the expiry month. We also obtain a series of model-free implied variance for a constant time to maturity of 30 days, computed by interpolating the model-free variances of options maturing immediately before and after 30 days.[22] We construct both implied variance series using only out-of-the money call and put options, since they generally correspond to the most liquid contracts.

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distributional assumptions for the risk-neutral distribution of the underlying. The forward prices obtained with this methodology are very similar to the forward prices obtain through put-call parity.

22 The CBOE volatility index (VIX) is another example of a 30-days interpolated model-free implied variance. The CBOE website provides further details on the VIX construction methodology.

*Table I.* Summary statistics

| Panel A: Volatilities | | | | | |
|---|---|---|---|---|---|
| | min | median | mean | max | std.dev. |
| $\sigma_{ECO}$ | 58.81 | 94.97 | 95.84 | 133.52 | 17.65 |
| US | 0.0492 | 0.0921 | 0.0992 | 0.2025 | 0.0253 |
| TY | 0.0357 | 0.0619 | 0.0653 | 0.1201 | 0.0179 |
| FV | 0.0228 | 0.0462 | 0.0462 | 0.1044 | 0.0151 |
| ED | 0.0354 | 0.1968 | 0.2144 | 0.6832 | 0.1388 |
| S&P500 | 0.1023 | 0.1538 | 0.1723 | 0.4264 | 0.0615 |
| Cyclical | 0.2038 | 0.3463 | 0.3734 | 0.8508 | 0.1038 |
| Non-cyclical | 0.1791 | 0.2719 | 0.2855 | 0.6042 | 0.0815 |
| Panel B: Volume and Open Interest | | | | | |
| US volume | 1079 | 11920 | 13979 | 57803 | 8676 |
| US open interest | 26203 | 103683 | 147783 | 739402 | 98823 |
| TY volume | 1463 | 33278 | 39298 | 169874 | 24242 |
| TY open interest | 70107 | 303658 | 401093 | 1069171 | 239867 |
| FV volume | 505 | 9153 | 12308 | 80451 | 10794 |
| FV open interest | 8772 | 117599 | 148091 | 442619 | 100809 |
| ED volume | 125 | 13019 | 24298 | 391499 | 34695 |
| ED open interest | 7402 | 359475 | 839384 | 3876157 | 880209 |
| S&P500 volume | 595 | 2007 | 2269 | 6893 | 1150 |
| S&P500 open interest | 5451 | 16514 | 17290 | 38600 | 6826 |
| Cyclical volume | 289 | 947 | 998 | 3660 | 340 |
| Cyclical open interest | 4877 | 9646 | 10411 | 26457 | 3658 |
| Non-cyclical volume | 106 | 377 | 425 | 2145 | 233 |
| Non-cyclical open interest | 2154 | 5484 | 5868 | 17502 | 2251 |

This table shows summary statistics for economic uncertainty, $\sigma_{ECO}$, and options on the 30-year Treasury bond futures (US), the 10-year Treasury note futures (TY), the 5-year Treasury note futures (FV), or the Eurodollar futures (ED), the S&P500 index, cyclical and non-cyclical stocks. Panel A shows implied volatilities and Panel B shows trading volume and open interest for one-month to maturity options.

We apply the same methodology to obtain three time-series of model-free implied variances for equity options. The first series is an interpolated risk-neutral 30-day variance of S&P500 options, cyclical equity options, and non-cyclical equity options. We use daily interpolated implied volatilities at different deltas for each security from OptionMetrics and use Black-Scholes to infer the corresponding option prices.[23] The second and the third model-free implied variance series use the raw market prices of options with the shortest maturity and the next shortest maturity, respectively.

Table I, panel A, presents summary statistics for financial markets uncertainty. Consistent with intuition, bond volatilities increase for longer bond maturities,

[23] We use OptionMetrics data on forward prices, zero-coupon rates, and dividend rates to obtain the other inputs of the option pricing formula.

with the exception of Eurodollar options (ED).[24] Uncertainty in stock markets is larger, because of the combined effect of cash flow and discount rate uncertainty. More specifically, the volatility of cyclical stocks reflects the stronger relation with business cycle dynamics and is thus about one quarter larger than the volatility of non-cyclical stocks.

### 4.3 ECONOMIC AND FINANCIAL MARKETS UNCERTAINTY

We now link the theoretical relations derived in Section 2 with their empirical counterparts, setting the stage for the empirical analysis of the next section. Our proxy for the jump uncertainty $\sigma_j^2$ is the measure of macroeconomic uncertainty $\sigma_{ECO}^2$ extracted from economic derivatives according to the model-free methodology of equation (17). We assume a linear scaling relation between the jump uncertainty of the state variable $Y$ and our measure of macroeconomic uncertainty

$$\sigma_j^2 = \eta_0 + \eta_1 \sigma_{ECO}^2. \tag{18}$$

Similarly, we measure the expected risk-neutral variance of bond prices $\mathrm{Var}_P$ and stock prices $\mathrm{Var}_S$ – financial markets uncertainty – with the same model-free approach. We can now express the resolution of uncertainty in financial markets around macroeconomic announcement as a function of the *ex-ante* macroeconomic uncertainty. For bonds, the theoretical prediction:

$$\begin{aligned}\mathrm{Var}_{P(Y,\tau,T,T^*)} - \mathrm{Var}_{P(Y,\tau-,T,T^*)} &= b^\top(\tau, T)\sigma^2(Y)b(\tau, T) - \cdots \\ &\quad b^\top(\tau-, T)\sigma^2(Y)b(\tau-, T) - \cdots \\ &\quad \frac{b^\top(\tau-, T)\sigma_{j1}^2 b(\tau-, T)}{T^* - \tau-} \\ &\approx -\frac{b^\top(\tau-, T)\sigma_{j1}^2 b(\tau-, T)}{T^* - \tau-}\end{aligned} \tag{19}$$

is tested with the following empirical specification

$$\begin{aligned}\mathrm{Var}_t - \mathrm{Var}_{t-1} &= \alpha - \frac{b^\top\left(\eta_1\sigma_{ECO,t}^2\right)b}{T^* - \tau} + \epsilon_\tau \\ &= \alpha + \varphi_B \sigma_{ECO,t}^2 + \epsilon_t\end{aligned} \tag{20}$$

where $\mathrm{Var}_{t-1}$ is the model-free implied variance of a given bond the day prior to the news release, $\mathrm{Var}_t$ is the model-free implied variance of a given bond the day of the news release (after the announcement), $\varphi_B$ is an estimator of $-(b^\top)^2\eta_1/(T^* - \tau)$,

---

[24] The volatility of Eurodollar futures is conventionally quoted on the basis of the implied discount rate, given by 100 minus the prevailing futures price. As a result, it is comparable to the other bond volatilities only after a duration adjustment.

and we add a residual term. Note that the time-to-maturity does not introduce any bias in the empirical analysis, because we use options with constant time-to-maturity.[25] Analogously for stocks, the theoretical prediction

$$
\begin{aligned}
\mathrm{Var}_{S(Y,Z,\tau,T,T^*)} - \mathrm{Var}_{S(Y,Z,\tau-,T,T^*)} &= \beta_i^\top(\tau,T)\sigma^2(Y)\beta_i(\tau,T) + \cdots \\
&\quad C_i^\top(\tau,T)^2\sigma^2(Z)C_i(\tau,T) - \cdots \\
&\quad \beta_i^\top(\tau-,T)\sigma^2(Y)\beta_i(\tau-,T) - \cdots \\
&\quad \frac{\beta_i^\top(\tau-,T)\sigma_{j1}^2\beta_i(\tau-,T)}{T^* - \tau-} - \cdots \\
&\quad C_i^\top(\tau-,T)\sigma^2(Z)C_i^\top(\tau-,T) \\
&\approx -\frac{\beta_i^\top(\tau-,T)\sigma_{j1}^2\beta_i(\tau-,T)}{T^* - \tau-}
\end{aligned} \tag{21}
$$

is tested with the following empirical specification

$$
\mathrm{Var}_t - \mathrm{Var}_{t-1} = \alpha + \varphi_S\sigma^2_{ECO,t} + \gamma Ret_{t,t-1} + e_t, \tag{22}
$$

where $\mathrm{Var}_{t-1}$ is the model-free implied variance of a class of stocks the day prior to the news release, $\mathrm{Var}_t$ is the model-free implied variance the day of the news release (after the announcement), $\varphi_S$ is an estimator of $-(\beta_i^\top)^2\eta_1/(T^* - \tau)$, and we add a residual term. Furthermore, $Ret_{t,t-1}$ is the return of the underlying asset on the announcement day and represents a control for the leverage effect. The reason for including this term is that the resolution of uncertainty in the equity market can be overshadowed by the leverage effect, the empirical observation that a large negative return tends to be associated more with an increase in volatility than an equally large positive return. If, for some reason, stock prices drop on the announcement day, implied volatility increases due to the leverage effect and any resolution of uncertainty related to the macroeconomic release is harder to detect.[26]

We also estimate equation (20) and (22) using $\mathrm{Var}_{t+1}$ in place of $\mathrm{Var}_t$ to understand whether the resolution of uncertainty is transitory and whether the process of uncertainty resolution is instantaneous or takes some time. In addition, we also expand the specifications (20) and (22) to control for the magnitude of the surprise in the announcement, i.e. the standardized difference between the actual and the expected release. We thus make sure that our results do not depend on the degree to which the market is surprised by the announcement.

[25] This is clearly the case for the model-free implied variances interpolated at 30-days. It is generally the case also for the series using raw option prices, given that options have monthly expiration dates and macroeconomic announcements tend to be released with a recurrent monthly pattern.

[26] There is no leverage effect in the bond market in our sample period. As a result, we use the specification in equation (20) for the bond regressions.

4.4 ECONOMIC UNCERTAINTY AND TRADING

The previous section shows the model implications of the resolution of macroeconomic uncertainty for financial markets uncertainty. Although the model is silent about trading dynamics, it is interesting to investigate whether investors trade financial derivatives in ways that reflect macroeconomic uncertainty. We thus examine the relation between option trading volume, option open interest, and macroeconomic uncertainty. Since data is only available on exchange-traded options, and investors are likely to implement trading strategies also over-the-counter, any significant relation that we uncover in our sample is likely to be a conservative estimate.

Table I, panel B, presents summary statistics for volume and open interest on bond and stock options. We observe that contracts written on the 10-year (TY) and the Eurodollar are showing the largest trading activity and open positions among the bond underlyings. Options on cyclical stocks are more traded than options on non-cyclical stocks, with almost three times the traded volume and twice the open interest.

We investigate the trading behavior of market participants in the financial option markets by estimating the following two regressions:

$$\frac{(Volu_t - Volu_{t-1})}{Volu_{t-1}} = \alpha + \beta\sigma^2_{ECO,t} + e_t, \tag{23}$$

$$\frac{(Oint_t - Oint_{t-1})}{Oint_{t-1}} = \alpha + \beta\sigma^2_{ECO,t} + e_t, \tag{24}$$

where $Volu_t$ and $Oint_t$ are respectively the traded volume and open interest on the day of the announcement for the financial options contracts with closest quarterly maturity date.[27] We specify the dependent variable in percentage changes to control for the potential effects of trends in volume and open interest.

There are two general motives for why investors would trade financial options around macroeconomic announcements: hedging and speculation. An example of hedging would be an investor holding a bond portfolio and hedging against the adverse impact of macroeconomic news. An example of speculation would be an investor using options to exploit his information advantage (e.g., Amin and Lee, 1997; Pan and Poteshman, 2006, among others, show that options are used for information-based trading). Since private information about the macroeconomy has presumably a limited role, speculative trading stems from differences of opinion, where traders have common information, but differ in the way in which they interpret this information (e.g., Harris and Raviv, 1993; Kandel and Pearson, 1995). The

[27] The empirical results do not change if we consider trading volume and open interest of the options closest to maturity, regardless of whether it is a quarterly maturity date or not, or if we consider volume and open interest of all options with less than 60 days to maturity.

importance of both the hedging and speculation motive for trading activity is likely to be greater with higher macroeconomic uncertainty, since the benefits from hedging as well as the expected payoffs from speculation (or extent of difference in opinions) are increasing in uncertainty.

The hedging and speculative motives have similar implications for the expected signs in equations (23) and (24). Specifically, if hedging or speculative positions are opened the day before the announcement and closed right away after the release, we do not expect to find significant effects in equation (23). However, if hedging or speculative positions are opened gradually before the release and closed immediately after the release, we expect to find a positive relation in equation (23). One plausible reason for opening gradually a position over the days before the release is the "calm before the storm" effect (e.g., Jones et al., 1998), i.e., the documented reduced levels of liquidity preceding macroeconomic releases. The inventory-control models of Amihud and Mendelson (1980), Ho and Stoll (1983), O'Hara and Oldfield (1986), all emphasize the inventory risk to market-makers at times of high volatility. This fact evidently causes some dealers to widen or withdraw their quotes, inhibiting at the same time dealers from taking new positions. Since the risk to market-makers is likely to depend on the uncertainty about the upcoming macroeconomic release, we might expect the "calm before the storm" effect to be calmer when the storm is supposed to be more serious. Higher macroeconomic uncertainty is thus likely to be a stronger incentive to enter into hedging or speculative positions gradually.

The "calm before the storm" effect has been traditionally documented in spot markets, where market makers' inventory with directional exposure could be adversely affected by the high volatility induced by the macroeconomic announcement. We might observe the same effects on the option markets, if option market makers find it difficult to hedge their inventory in the imminence of a jump, that is, when the gamma risk of their option portfolios is high. Alternatively, the reduced liquidity on the spot market before the announcements could spill over to the option market for the same hedging reasons. Ultimately, the existence of a "calm before the storm" effect in option markets is an empirical question.

The "calm before the storm" effect could also be an incentive to delay trading for reasons unrelated to the macroeconomic release until after the uncertainty is resolved and liquidity is back to normal levels. In this case, we would also observe a positive relation in equation (23). We can distinguish between hedging or speculative trades and simple trading deferral by estimating equation (24). In the case of hedging/speculative trades, the relation between change in open interest and macroeconomic uncertainty should be negative, as a result of positions being closed out after the release. In the case of trading deferral, the relation should be positive, because new positions are opened after the uncertainty is resolved.

The discussion above does not suggest any empirical test to distinguish between hedging and speculation, since both trading motives have the same implications for volume and open interest. However, we can hope to gain some insight into which explanation likely dominates in two ways. First, by comparing the empirical results of estimating equations (23) and (24) for different underlying assets, in particular bonds versus stocks. While the model sketched out in Section 2 does not have specific predictions here, the intuitive argument from simple present value models suggests that the effect of macroeconomic news on bonds should be stronger than the effect on stocks, because there are no potential offsetting effects of cash flow news. Speculative trading on macroeconomic news would thus be naturally implemented with bond options. As a result, any significant result of equations (23) and (24) for options written on stocks is more likely an effect of investors trying to hedge a portfolio of cyclical stocks from the adverse effects of macroeconomic news.

A second way to disentangle the hedging and the speculative motive is to compare the results for different option types. In particular, hedging of long positions is most naturally implemented with put options. The aggregate position in stocks is evidently long and thus finding significant results for put to call ratios of volume and open interest would suggest a hedging motive. The same reasoning is hardly applicable to bonds, since it is not clear whether the aggregate position to hedge is long or short. Similarly, since speculative trades are not likely to be systematically in one direction, different options are used for different strategies. In summary, we expect to find significant results in put to call ratios only for stocks if there is an hedging motive.[28]

## 5 Empirical results

### 5.1 DO FINANCIAL MARKETS REACT TO MACROECONOMIC NEWS?

A precondition for macroeconomic uncertainty to be related to volatility in financial markets is that the new information released in the announcement moves asset prices. If financial markets do not react to macroeconomic news, there is no reason to expect that uncertainty about economic fundamentals is reflected in bond and stock market volatility. We therefore begin our empirical analysis by investigating whether bond and stock prices respond at the daily frequency to macroeconomic news during our sample period.[29]

[28] Since speculative trades are carried out sometimes with calls and sometimes with puts when macroeconomic uncertainty is higher, we could still obtain significant results by estimating equations (23) and (24) separately for calls and puts.

[29] While there is strong evidence in the literature that bond markets respond intra-daily to macroeconomic news (e.g., Fleming and Remolona, 1999a) and some evidence that stock markets also

To gauge the extent to which an announcement contains new information, we construct the following standardized measure of surprise:

$$S_t = \frac{A_t - F_{0,t}}{s}, \tag{25}$$

where $A_t$ is the value of the economic statistic released at time $t$, $F_{0,t}$ denotes the corresponding market forecast implied by the economic derivatives prices and obtained through put-call parity (see details in Section 4.1), and $s$ is the (unconditional) empirical standard deviation of the innovations $A_t - F_{0,t}$.[30] Standardizing the surprise by the empirical standard deviation helps in the interpretation of the results. We then estimate the following regression:

$$Ret_{t,t-1} = \alpha + \beta S_t + e_t, \tag{26}$$

where $Ret_{t,t-1}$ represents the percentage daily return on either the 30-year Treasury bond futures (US), the 10-year Treasury note futures (TY), the 5-year Treasury-note futures (FV), the Eurodollar futures (ED), the S&P 500 index, the portfolio of cyclical stocks, or the portfolio of non-cyclical stocks.

Table II presents the regression results for the case of non-farm payrolls (NFP), which is by far the most influential macroeconomic announcement during our sample period.[31] Panel A shows that a positive surprise in NFP is associated with strongly negative bond returns. In all cases, the $R^2$ indicates that more than 39 percent of the variance of announcement day bond returns is explained by the surprises in NFP. More specifically, a one-standard deviation positive surprise in NFP is associated with a negative 62, 46, 34, and three basis point daily return for the US, TY, FV, and ED futures, respectively. If we transform these bond returns into bond yield changes using the modified duration of the underlying, we naturally obtain the opposite ranking, i.e., the effect of NFP surprises on yields is greater for shorter maturity and smaller for longer maturity bonds (results not tabulated).[32]

respond intra-daily to the same news (e.g., Flannery and Protopapadakis, 2002), it is important for the empirical design of our paper to test whether this is true at the daily frequency and during our sample period.

[30] We also construct the standardized measure of surprise replacing the market forecast implied by economic derivatives $F_{0,t}$ with the median analyst forecast provided by Money Market Services. We obtain qualitatively and quantitatively similar results.

[31] During our sample period, NFP is always released concurrently with the civilian unemployment rate (CUR) in the employment report (ER). To ensure that the results in Table II are not driven by the CUR release, we also include as regressor the standardized surprise of the CUR release, using the median forecast provided by Money Market Services as proxy for market expectations. The results for the NFP release do not change in this expanded regression specification.

[32] We map returns into yield using a simple duration adjustment for each segment of the yield curve, using the standard formula for the modified duration: $dP/P = -D_m dY$, where $dP/P$ is the return on the underlying, $D_m$ is the modified MacCauley duration, and $dY$ is the change in yields. We calculate the duration of each Treasury futures contract as the modified duration of a deliverable

*Table II.* Macroeconomic news and financial markets returns

| | Panel A: Bond market | | | |
|---|---|---|---|---|
| | US | TY | FV | ED |
| constant | −0.00134 | −0.00135 | −0.00094 | −0.00008 |
| | (0.00121) | (0.00083) | (0.00057) | (0.00005) |
| std surprise | −0.00619*** | −0.00464*** | −0.00336*** | −0.00027*** |
| | (0.00113) | (0.00078) | (0.00054) | (0.00005) |
| $R^2$ | 0.3990 | 0.4432 | 0.4661 | 0.3950 |
| observations | 47 | 47 | 47 | 47 |

| | Panel B: Stock market | | |
|---|---|---|---|
| | S&P 500 | Cyclical | Non-cyclical |
| constant | −0.00006 | 0.00097 | 0.00023 |
| | (0.001354) | (0.00204) | (0.00166) |
| std surprise | 0.000886 | 0.00362** | −0.00072 |
| | (0.001283) | (0.00182) | (0.00156) |
| $R^2$ | 0.0105 | 0.0730 | 0.0046 |
| observations | 47 | 47 | 47 |

***, **, and * denote statistical significance at the 1, 5, and 10 percent levels, respectively.
This table shows the results of estimating the following regression:

$$Ret_{t,t-1} = \alpha + \beta S_{nfp,t} + e_{nfp,t},$$

where $Ret_{t,t-1}$ represents, in Panel A, the daily percentage return on the 30-year Treasury bond futures (US), the 10-year Treasury note futures (TY), the 5-year Treasury note futures (FV), or the Eurodollar futures (ED). In Panel B, $Ret_{t,t-1}$ represents the daily percentage returns on the S&P 500 index, the return on a portfolio of cyclical stocks, or the return on a portfolio of non-cyclical stocks. $S_{nfp,t}$ represents the standardized difference between the actual NFP release and the implied market forecast, as defined in equation (25).

Surprises in NFP have far less explanatory power for stock returns. Specifically, the announcement day return on the S&P 500 index is not significantly related to the surprises in NFP. Any effect of the macroeconomic news on aggregate cashflows and discount rates appears to be either offsetting (i.e., discount rates and expected cashflows changing in the same direction) or swamped by other news about individual firm cashflows. Consistent with this later explanation, we observe a significant, at the 5 percent level, relationship between the return on cyclical stocks and NFP surprises, explaining almost 10 percent of the variation in announcement day returns of cyclical stocks. The positive coefficient indicates that, for cyclical stocks, the cash flow effect completely dominates the discount rate effect. Specifically, a one standard deviation positive surprise in NFP

bond with the average maturity in the maturity range. We obtain the following modified durations for each Treasury futures: 11.82 for the 30-year Treasury bond futures, 6.09 for the 10-year Treasury note futures, 3.93 for the 5-year Treasury note futures, 0.24 for the Eurodollar futures.

*Table III. Ex-ante* macroeconomic uncertainty and *ex-post* surprises

| | (1) $\lvert S_{nfp,t}\rvert$ | (2) $S^2_{nfp,t}$ |
|---|---|---|
| constant | 0.8610 | 578.23 |
| | (0.5346) | (7365.48) |
| economic σ | −0.0003 | 70.34 |
| | (0.0055) | (75.61) |
| $R^2$ | 0.0001 | 0.0189 |
| observations | 47 | 47 |

This table shows the results of estimating the following regression:

$$f(S_{nfp,t}) = \alpha + \beta\sigma_{ECO,t} + e_t,$$

where $S_{nfp,t}$ represents the standardized difference between the actual NFP release and the implied market forecast, as defined in equation (25), and $\sigma_{ECO,t}$ is the model-free economic uncertainty estimated from economic derivatives prices before the release. In column (1), $f(.)$ is the absolute value. In column (2), $f(.)$ is the squared value.

is associated with a 36 basis points higher average daily return on the portfolio of cyclical stocks. Additional analysis, not reported in the table, reveals that if we estimate the regression (26) separately for each cyclical stock, we obtain positive slope coefficients for 20 of the 22 stocks, with 12 being significant at least at the 10 percent level. For two stocks the slope coefficient is negative but insignificant.

### 5.2 IMPLIED MACROECONOMIC UNCERTAINTY

We compute the macroeconomic uncertainty implied by the economic derivatives for all NFP announcements in our sample using data from the auctions directly preceding the news releases. We focus again on NFP because it is the most influential macroeconomic announcement, and we consider only the second auction because it reflects the most up-to-date information before the announcement. It turns out, however, that the informational contents of the auction held in the afternoon of the day before the announcement and of the auction held in the morning of the day of the announcement are very similar. Specifically, the implied forecasts $F_{0,t}$ of the two auctions have a correlation of 0.99 and the model-free implied variances $\sigma^2_{ECO,t}$ have a correlation of 0.87.

Before turning to a more detailed analysis relating macroeconomic uncertainty to changes in financial markets volatility, we first test whether the *ex-ante* measure of macroeconomic uncertainty $\sigma^2_{ECO,t}$ has predictive power for the unsigned *ex-post* surprise in the macroeconomic release. Table III shows that this is not the case, at least in our sample period. Ex-ante uncertainty in financial markets reflects ex-ante macroeconomic uncertainty, even when this is not significantly correlated with the realized surprise. This discrepancy reveals one of the main strengths of our approach. Ex-post realized measures of variance require a lot of observations to be estimated precisely, whereas option-implied measures are available at each point in time. Alternatively, our measure of economic uncertainty could include a risk-premium component that is not present in the realized measure. This negative result is very interesting for the upcoming empirical analysis. Macroeconomic shocks matter for bonds and cyclical stock returns, but *ex-ante* macroeconomic uncertainty does not predict the shock, do then financial markets reflect macroeconomic uncertainty?

### 5.3 DO FINANCIAL MARKETS REFLECT MACROECONOMIC UNCERTAINTY?

#### *5.3.1 Bonds*

As a first step in analyzing the resolution of macroeconomic uncertainty in bond markets, we compute the unconditional average implied volatility of at-the-money call and put options on the day before the NFP release and at the close of the announcement day. The average at-the-money implied volatility the day before the announcement is 10.15 percent for options written on the 30-year Treasury bond futures (US), 6.85 percent for options on the 10-year Treasury note futures (TY), 4.63 percent for options on the 5-year Treasury note futures (FV), and 24.02 percent for options on Eurodollar futures (ED), respectively. The average at-the-money implied volatility after the announcement is 9.26 percent for options on US, 6.12 percent for options on TY, 4.11 percent for options on FV, and 21.81 percent for options on ED, respectively. This represents a percentage decrease in implied volatility of 8.77 percent for options on US, 10.63 percent for options on TY, 11.25 percent for options on FV, and 9.20 percent for options on ED. We conclude from this evidence that bond market volatility drops unconditionally on macroeconomic announcement days.[33]

Table IV shows the results of estimating equation (20). We regress the change in the model-free implied variance of options written on US, TY, FV, or ED, on the uncertainty about the NFP announcement implied by the economic derivatives, expressed in million jobs. We correct for the effect of heteroscedasticity using standard two-stage weighted least squares. We observe that for the longer maturities (US, TY, and FV), between 25 and 40 percent of the across-announcement day variation in the variance drop is explained by the ex-ante uncertainty about the NFP

[33] The evidence provided here does not precisely match the figures presented in the introduction, because those figures are constructed using the weighted least squares estimates of Table IV.

*Table IV.* Resolution of bond market uncertainty

| | (1) release day | (2) release day | (3) day after | (4) day after |
|---|---|---|---|---|
| | | Panel A: US | | |
| option's time to maturity | closest | 30 | closest | 30 |
| constant | 0.0014** | 0.0018*** | 0.0015** | 0.0014*** |
| | (0.0006) | (0.0005) | (0.0006) | (0.0005) |
| economic $\sigma^2$ | −0.2873*** | −0.2891*** | −0.2677*** | −0.2169*** |
| | (0.0586) | (0.0547) | (0.0640) | (0.0565) |
| $R^2$ | 0.3586 | 0.3941 | 0.3043 | 0.2596 |
| observations | 47 | 47 | 47 | 47 |
| | | Panel B: TY | | |
| option's time to maturity | closest | 30 | closest | 30 |
| constant | 0.0008*** | 0.0005 | 0.0009** | 0.0008*** |
| | (0.0003) | (0.0003) | (0.004) | (0.0003) |
| economic $\sigma^2$ | −0.1406*** | −0.1309*** | −0.1841*** | −0.1243*** |
| | (0.0296) | (0.0424) | (0.0776) | (0.0289) |
| $R^2$ | 0.3545 | 0.2655 | 0.2879 | 0.3270 |
| observations | 47 | 47 | 47 | 47 |
| | | Panel C: FV | | |
| option's time to maturity | closest | 30 | closest | 30 |
| constant | 0.0004* | 0.0001 | 0.0003** | 0.0004** |
| | (0.0002) | (0.0001) | (0.0001) | (0.0002) |
| economic $\sigma^2$ | −0.0877*** | −0.0447*** | −0.0567*** | −0.0816*** |
| | (0.0261) | (0.0128) | (0.0126) | (0.0262) |
| $R^2$ | 0.2694 | 0.2424 | 0.3558 | 0.2956 |
| observations | 47 | 47 | 47 | 47 |
| | | Panel D: ED | | |
| option's time to maturity | closest | 30 | closest | 30 |
| constant | 0.0171 | 0.0127 | 0.0162 | 0.0241 |
| | (0.0254) | (0.0112) | (0.0175) | (0.0153) |
| economic $\sigma^2$ | −2.5951* | −2.1103** | −2.3296* | −2.6234* |
| | (1.4835) | (0.9696) | (1.5467) | (1.4516) |
| $R^2$ | 0.0512 | 0.1325 | 0.0750 | 0.0982 |
| observations | 47 | 47 | 47 | 47 |

***, **, and * denote statistical significance at the 1, 5, and 10 percent levels, respectively.
This table shows the results of estimating the following regression:

$$(\mathrm{Var}_t - \mathrm{Var}_{t-1}) = \alpha + \beta\sigma^2_{ECO,t} + e_t,$$

where $\mathrm{Var}_t$ is the model-free implied variance on the day of the release, obtained either from options with the closest maturity (column (1)), or interpolated at the 30 days horizon (column (2)). We also estimate the same regression substituting $\mathrm{Var}_t$ with $\mathrm{Var}_{t+1}$ (column (3) and (4)). Panel A, B, C, and D show the results for options on the 30-year Treasury bond futures (US), on the 10-year Treasury note futures (TY), on the 5-year Treasury note futures (FV), and on the Eurodollar futures (ED), respectively.

release. The higher the ex-ante uncertainty about NFP, the greater is the drop in implied variance following the announcement. This result is apparent in both the implied variances of the closest maturity options and the interpolated variances with a 30-day horizon. The relationship between macroeconomic uncertainty and changes in bond market uncertainty is not only highly significant statistically but also very important economically. For example, a high degree of uncertainty about NFP equal to the average uncertainty in our sample plus one standard deviation predicts a drop of about 4 percent in the 30-days model-free implied volatility of US options, which corresponds to a 40 percent negative change. This drop is obviously very large compared to the drop in implied volatility of only 1.5 percent on announcement days with low macroeconomic uncertainty (average uncertainty minus one standard deviation).

The comparison of the coefficients for different underlying bonds shows that the effect of macroeconomic uncertainty increases monotonically for longer maturities.[34] This result depends, however, on variances computed on bond returns and it would be informative to understand the effect on the variance of bond yields as well. Using a simple duration adjustment, we can change the unity of measure of the coefficients of Table IV. Specifically, we use the equivalence between variances in yields and variances in returns divided by the square of the modified duration of the underlying (see footnote 32 and earlier discussion). This transformation reverses the monotonic ordering, with short-term yield variances featuring the largest impact and long-term yield variances featuring the lowest impact of macroeconomic uncertainty (results not reported).

We can now resort to the simple model sketched out in Section 2 to *jointly* analyze the response of the underlying returns and the resolution of financial market uncertainty around macroeconomic releases. More specifically, the response of bond returns to shocks to the state variables described by equation (7) maps into the response of bond variances to macro uncertainty described by equation (20) up to a square product and a linear transformation. The theoretical model thus imposes restrictions on the parameters with respect to the unrestricted estimation carried out with the linear regressions of Table II and Table IV.

We cast this problem into a standard Generalized Method of Moments (GMM) estimation framework, with six moment conditions and four unknown parameters.[35] The J-test for the overidentifying restrictions equal to 4.89 with a p-value of 0.0865,

[34] As explained before, the volatility of Eurodollar futures is conventionally quoted on the basis of the implied discount rate. The coefficient of the regression in Table IV is thus comparable to the other coefficients only after a duration adjustment. The adjusted coefficient on the Eurodollar is the lowest of the fixed-income options.

[35] We exclude from this empirical exercise the options on the Eurodollar, given that their implied volatility is in yield-space rather than in return-space and it would thus imply duration adjustments that would unnecessarily complicate the analysis. We thus end up with six moment conditions (one

suggesting that the model is not strongly rejected by the data. The results of the GMM estimation are not reported but are very similar to the unrestricted estimates reported in Table II and IV.

To see if the resolution of uncertainty is transitory, we also test whether the effects on implied variance are still present at the close of the day following the announcement.[36] We observe that the effects of macroeconomic uncertainty on bond markets uncertainty persist, both in terms of magnitude and statistical significance of the coefficients. The explanatory power of macroeconomic uncertainty for the longer horizon drop in variance is somewhat lower, but still more than one third in general.

We extend the analysis above to the full week surrounding the Friday NFP release to gain further intuition on the evolution of uncertainty around macroeconomic announcements. We also aim to understand whether the documented resolution of uncertainty occurs because of an unusually high level of financial markets uncertainty the day before the announcement. Figure 2 presents the average pattern of financial market uncertainty in periods of high and low macroeconomic uncertainty for the four different bond underlyings. We define the releases with $\sigma_{ECO}$ in the top (bottom) quartile as the high (low) uncertainty announcements. We notice four stylized facts. First, the resolution of financial markets uncertainty with high ex-ante macroeconomic uncertainty is substantial. Second, the drop in uncertainty is not transitory. Third, financial market uncertainty on Thursdays is only moderately higher than on the previous days. Finally, all these patterns are valid across all the different bond underlyings.

We also try to understand whether there is a systematic role for good versus bad news about the economy in resolving the uncertainty in bond markets. For this, we add as a regressor the standardized surprise in the NFP release. The results (not tabulated) show that the new information embedded in the surprise does not have a significant impact on the change in bond implied variances. This also holds when we interact the standardized surprise in the release with our measure of macroeconomic uncertainty.

### *5.3.2 Aggregate stock index*

Table V shows the results for the S&P 500 index. In Panel A, we use changes in the model-free implied variances of S&P 500 index options interpolated to have a constant time to maturity of 30 days.[37] Panel B is for changes in the model-free

for each equation estimated in Table II and IV) and four parameters, i.e., the three sensitivities of bond returns to macroeconomic surprises and the linear transformation labeled $\eta_1$ in equation (20).

[36] Since the non-farm payrolls are generally released on Friday, we actually test if the relation between model-free implied variance and macroeconomic uncertainty is still present after the weekend.

[37] This measure of uncertainty corresponds basically to the square of the CBOE volatility index (VIX).

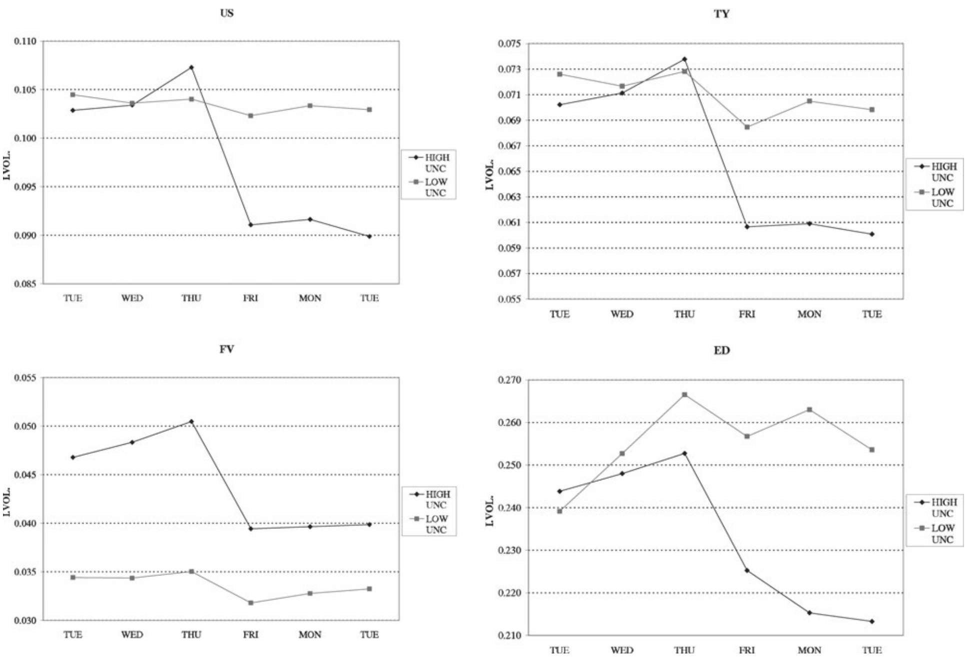


*Figure 2*. Bond Market Uncertainty around Macroeconomic Releases
This figure shows average financial market uncertainty for options on the 30-year Treasury futures (US), the 10-year Treasury futures (TY), the 5-year Treasury futures (FV), and the Eurodollar futures (ED) during the week surrounding the Friday NFP release, for high and low macroeconomic uncertainty releases.

implied variances of the closest to maturity options and next closest to maturity options, respectively. As before, the unity of measure of macroeconomic uncertainty is millions of jobs and we correct for the effect of heteroscedasticity using two-stage weighted least squares. All the coefficients relating changes in the implied index variances to our measure of macroeconomic uncertainty are negative, in line with the results in Table IV. In contrast to those bond market results, however, none of the coefficients for the stock index are statistically significant. This result is consistent with the observation from Table II that the stock index does not react strongly to macroeconomic news. The resolution of economic uncertainty therefore also has only little effect on the index volatility.

The coefficients on the index return are all negative and, as expected, highly significant. Recall from the discussion surrounding equation (20) that we add the index return to explicitly take into account the well-documented leverage effect. Judging by the $R^2$ of the regressions, the leverage effect explains a vast majority of the changes in index variance.

*Table V.* Resolution of aggregate stock market uncertainty

| | Panel A: S&P 500 interpolated options | |
|---|---|---|
| | release day | day after |
| constant | −0.0004 | 0.0001 |
| | (0.0004) | (0.0006) |
| S&P 500 return | −0.1889*** | −0.2427*** |
| | (0.0188) | (0.0226) |
| economic $\sigma^2$ | −0.0385 | −0.0037 |
| | (0.0290) | (0.0436) |
| $R^2$ | 0.7232 | 0.7948 |
| observations | 43 | 43 |

| | Panel B: S&P 500 raw options | | | |
|---|---|---|---|---|
| | release day | release day | day after | day after |
| options' average time to maturity | 14.82 | 45.14 | 14.82 | 45.14 |
| average no. of options | 42.13 | 35.63 | 39.24 | 35.38 |
| constant | −0.0006** | −0.0006 | 0.0001 | 0.0020 |
| | (0.0003) | (0.0005) | (0.0001) | (0.0013) |
| S&P 500 return | −0.1241*** | −0.0848*** | −0.0520 | −0.1953*** |
| | (0.0367) | (0.0261) | (0.0354) | (0.0269) |
| economic $\sigma^2$ | −0.0006 | −0.0707 | −0.0411 | −0.0387 |
| | (0.0507) | (0.0489) | (0.0685) | (0.0274) |
| $R^2$ | 0.1795 | 0.2120 | 0.0445 | 0.3921 |
| observations | 43 | 43 | 43 | 43 |

***, **, and * denote statistical significance at the 1, 5, and 10 percent levels, respectively.
This table shows the results of estimating the following regression:

$$(\mathrm{Var}_t - \mathrm{Var}_{t-1}) = \alpha + \beta\sigma^2_{ECO,t} + \gamma Ret_{t,t-1} + e_t,$$

where $\mathrm{Var}_t$ is the stock market model-free implied variance and $Ret_{t,t-1}$ is the S&P 500 index return on the day of the release. We also estimate the same regression substituting $\mathrm{Var}_t$ with $\mathrm{Var}_{t+1}$ and $Ret_{t,t-1}$ with $Ret_{t+1,t-1}$. Panel A shows the results for the model-free implied variance obtained from options on the S&P 500 Index interpolated at a 30-day time to maturity. Panel B shows the results for the model-free implied variance of options traded on both the pre-announcement and the announcement day for the traded maturities.

### *5.3.3 Cyclical and non-cyclical stocks*

Given the considerably weaker results for the S&P 500 index, relative to the results for long- and short-term bonds, we next consider a basket of options written on cyclical stocks. The goal is to examine how the effect of macroeconomic uncertainty differs for stocks that we know, from the regressions in Table II, react more strongly to macroeconomic announcements. Table VI shows the results for cyclical stocks using again either the change in interpolated 30-day model-free implied variances (Panel A) or the change in raw model-free implied variances (Panel B). The regression specification is identical to that for the stock index. However, we

*Table VI.* Resolution of cyclical stock uncertainty

| | Panel A: Interpolated options on cyclical stocks | |
|---|---|---|
| | release day | day after |
| constant | 0.0019 | 0.0080** |
| | (0.0017) | (0.0036) |
| cyclical stocks return | −0.3635*** | −0.4286*** |
| | (0.0678) | (0.0784) |
| economic $\sigma^2$ | −0.1858* | −0.1689* |
| | (0.1106) | (0.1010) |
| $R^2$ | 0.5766 | 0.7121 |
| observations | 43 | 43 |

| | Panel B: Raw options on cyclical stocks | | | |
|---|---|---|---|---|
| | release day | release day | day after | day after |
| options' average time to maturity | 14.82 | 45.14 | 14.82 | 45.14 |
| average no. of options | 66.97 | 98.09 | 97.44 | 60.74 |
| constant | −0.0003 | 0.0026** | 0.0286*** | 0.0081*** |
| | (0.0018) | (0.0010) | (0.0102) | (0.0017) |
| cyclical stocks return | −0.2105*** | −0.0411* | −0.1691 | −0.1799*** |
| | (0.0430) | (0.0222) | (0.0892) | (0.0766) |
| economic $\sigma^2$ | −0.1714* | −0.1779* | −0.1836* | −0.1548* |
| | (0.1013) | (0.1013) | (0.1089) | (0.0907) |
| $R^2$ | 0.4134 | 0.1837 | 0.1232 | 0.3944 |
| Observations | 43 | 43 | 43 | 43 |

***, **, and * denote statistical significance at the 1, 5, and 10 percent levels, respectively.
This table shows the results of estimating the following regression:

$$(\mathrm{Var}_t - \mathrm{Var}_{t-1}) = \alpha + \beta\sigma^2_{ECO,t} + \gamma Ret_{t,t-1} + e_t,$$

where $\mathrm{Var}_t$ is the model-free implied variance of a portfolio of cyclical stocks and $Ret_{t,t-1}$ is the return of a portfolio of cyclical stocks on the day of the release. We also estimate the same regression substituting $\mathrm{Var}_t$ with $\mathrm{Var}_{t+1}$ and $Ret_{t,t-1}$ with $Ret_{t+1,t-1}$. Panel A shows the results for the model-free implied variance obtained from options on cyclical stocks interpolated at a 30-day time to maturity, traded on both the pre-announcement and the announcement day. Panel B shows the results for the model-free implied variance of options traded on both the pre-announcement and the announcement day for the two traded maturities around the 30-day horizon.

apply an additional data filter in this case. We include in the analysis only options that are traded on both the pre-announcement day and on the announcement day. This is done to avoid any bias induced, for example, by options on volatile stocks being traded on only one of the two days causing a spurious difference between the average implied variances before and after the announcements.

Table VI shows that the implied variance of cyclical stocks is indeed related to the uncertainty about macroeconomic fundamentals. Higher ex-ante uncertainty about NFP is associated with a greater drop in implied variance when the uncertainty is resolved. This relation is statistically significant at the 10 percent level in all cases.

The coefficient is only slightly greater in magnitude for the 30-day interpolated variance in Panel A. Finally, we note that the drop in implied volatility appears to persist, given that the coefficients for the day after the release are similar in magnitude and statistical significance.

For comparison, we also repeat the analysis for non-cyclical stocks. The results, which are not tabulated to conserve space, show that the variance of non-cyclical stocks is unrelated to the uncertainty about macroeconomic fundamentals. The coefficients on the ex-ante uncertainty about NFP are never statistically significant, their magnitudes are small compared to the results for cyclical stocks, and the sign is even positive in a number of cases. These results confirm that non-cyclical stocks are relatively immune to changes in macroeconomic conditions, not only because contemporaneously their prices do not respond to announcements (see Table II) but also because ex-ante their implied volatilities do not reflect the degree of uncertainty about fundamentals.

### 5.4 HOW DO MARKETS DEAL WITH MACROECONOMIC UNCERTAINTY?

Our results thus far document a strong link between the ex-ante uncertainty about NFP releases and financial markets uncertainty. Clearly financial markets, and especially bond markets, are concerned about changes in macroeconomic conditions and the uncertainty associated with scheduled macroeconomic announcements. Given this empirical fact, it is reasonable to expect market participants to trade in a way that also reflects the degree of uncertainty about macroeconomic fundamentals. Specifically, we expect a greater degree of hedging or speculating before macroeconomic releases that are more uncertain, since the benefits of hedging or the gains from speculating increase with uncertainty. In this section, we explore the hypothesis developed in Section 4.4 by examining the relationship between the ex-ante uncertainty about NFP releases and changes in trading volume and open interest in bond and equity options on announcement days.

#### *5.4.1 Bonds*

Table VII presents the results of regressing the change in the trading volume for the closest quarterly maturity bond options on the ex-ante uncertainty about NFP releases. We observe that across the different option contracts a higher degree of macroeconomic uncertainty is associated with a greater increase in trading volume after the resolution of the uncertainty. We obtain the strongest statistical significance and explanatory power for the options written on the five-year Treasury note futures. This finding is consistent with the results of Table II, which show that the strongest explanatory power of macroeconomic news for bond returns is at the 5-year horizon. Accordingly, trading on macroeconomic uncertainty concentrates

*Table VII.* Bond market options volume

| | (1)<br>US | (2)<br>TY | (3)<br>FV | (4)<br>ED |
|---|---|---|---|---|
| | | Panel A: All options | | |
| constant | −0.3763 | −0.4070** | −0.7114*** | 0.1855 |
| | (0.2914) | (0.1543) | (0.2433) | (0.6464) |
| economic $\sigma^2$ | 65.3457** | 43.3397*** | 84.8365*** | 73.9421 |
| | (29.0683) | (15.3945) | (24.2662) | (64.4715) |
| $R^2$ | 0.1030 | 0.1526 | 0.2174 | 0.0290 |
| observations | 47 | 47 | 47 | 47 |
| | | Panel B: Call options | | |
| constant | −0.4066 | −0.1424 | 0.0811 | −0.0753 |
| | (0.3562) | (0.2683) | (0.4815) | (0.9368) |
| economic $\sigma^2$ | 84.1693** | 32.9626 | 31.1055 | 149.1184* |
| | (35.5249) | (26.7645) | (48.0255) | (93.4054) |
| $R^2$ | 0.1131 | 0.0333 | 0.0094 | 0.0547 |
| observations | 47 | 47 | 47 | 47 |
| | | Panel C: Put options | | |
| constant | −0.3312 | −0.4382*** | −0.6622*** | 0.1723 |
| | (0.3863) | (0.1558) | (0.2405) | (0.5354) |
| economic $\sigma^2$ | 70.7837* | 40.3792*** | 74.4085*** | 44.0173 |
| | (38.5325) | (15.5329) | (23.9881) | (52.5834) |
| $R^2$ | 0.0712 | 0.1331 | 0.1794 | 0.0157 |
| observations | 47 | 47 | 47 | 47 |

***, **, and * denote statistical significance at the 1, 5, and 10 percent levels, respectively.
This table shows the results of estimating the following regression:

$$(Volu_t - Volu_{t-1})/Volu_{t-1} = \alpha + \beta\sigma^2_{ECO,t} + e_t,$$

where $Volu_t$ is the trading volume for option contracts with the closest quarterly maturity traded on the release day $t$. Column (1) shows the results for options on the 30-year Treasury futures (US), column (2) for options written on the 10-year Treasury futures (TY), column (3) for options written on the 5-year Treasury futures (FV), and column (4) for options written on the Eurodollar futures (ED).

on the underlying asset with strongest ties with the release. This observation is also consistent with the evidence in the literature that price discovery is most pronounced for intermediate maturities (e.g., Fleming and Remolona, 1999b; Brandt and Kavajecz, 2004). Furthermore, the relation between macroeconomic uncertainty and trading volume is not only statistically significant but also economically important. For example, a high degree of uncertainty about NFP equal to one standard deviation above the mean uncertainty predicts a 38 percent increase in traded volume of FV call and put options, more than four times the unconditional increase in trading volume of nine percent on announcement days.

These results on trading volume are consistent with market participants using options to hedge against or speculate on the forthcoming NFP release. The greater the uncertainty about NFP, the greater are the benefits from hedging or gains from speculating. If a hedge or speculative position is entered into gradually over the days preceding the announcement and then unwound once the news is released, we would observe an increase in trading volume after the announcement. The magnitude of this increase in trading volume would depend on the size of the hedge or speculative position, which, in turn, is related on the degree of macroeconomic uncertainty. An alternative explanation is that market participants wait to trade for reasons unrelated to the announcement until the uncertainty about NFP is resolved. One reason to delay trades (or to enter into a position gradually) would be the reduced level of liquidity induced by the "calm before the storm" effect. If higher macroeconomic uncertainty implies more calm before a more serious storm, then the incentives to wait are greater.

The results on trading volume are qualitatively the same when we analyze separately the trading volume of call and put options (see panels B and C in Table VII). Although there is not a clear pattern in these disaggregated results, it appears that call option trading volume is most important for the long maturities and, to a lesser extent, for short bond maturities, whereas put option trading volume plays a larger role for intermediate bond maturities. We can only speculate on the reasons for these findings, but a story consistent with these disaggregated results is the following. The two largest fixed-income markets in the U.S. are the market for Treasuries and mortgage securities. The amount of outstanding marketable U.S. Treasury securities held by the public is largely at intermediate maturities.[38] Hedging macroeconomic uncertainty in the Treasury market would thus be presumably implemented mainly with protective put options at the 5-year (FV) and 10-year (TY) maturities. In contrast, prepayment risk hedging in the mortgage market is typically implemented with long maturities against lower yields or higher prices due to prepayment (e.g., Perli and Sack, 2003). Long-term call options would then be a suitable hedging tool. In addition, the two largest mortgage finance lenders, Fannie Mae and Freddie Mac, are known to be among the largest users of interest rate derivatives and to finance more than 40% of their assets with debt due in less than one-year (e.g., Jaffee, 2003). This last piece of evidence could explain the significant results on Eurodollar calls.

To gain further insight into which of the explanations for trading around macroeconomic releases is more plausible, we examine next the relationship between *ex-ante* macroeconomic uncertainty and the change in open interest on the announcement day. As explained in Section 4.4, the hedging/speculating explanation predicts a decrease in open interest while the waiting to trade explanation predicts an increase. In both cases, the rise or drop in open interest should in magnitude

[38] Statistics from the U.S. bureau of public debt show that notes represents about 57% of outstanding marketable U.S. Treasuries, bonds about 17% and bills about 26% (source: www.publicdebt.treas.gov).

*Table VIII.* Bond market options open interest

| | (1) US | (2) TY | (3) FV | (4) ED |
|---|---|---|---|---|
| | | Panel A: All options | | |
| constant | 0.0192* | 0.0241*** | 0.0356*** | 0.0050** |
| | (0.0107) | (0.0079) | (0.0120) | (0.0025) |
| economic $\sigma^2$ | −0.1604 | −1.6959** | −2.1696** | −0.6135** |
| | (1.0667) | (0.7936) | (1.1986) | (0.2473) |
| $R^2$ | 0.0005 | 0.0940 | 0.0692 | 0.1226 |
| observations | 47 | 47 | 47 | 47 |
| | | Panel B: Call options | | |
| constant | −0.0218 | 0.0206** | 0.0235 | 0.0017 |
| | (0.0172) | (0.0095) | (0.0194) | (0.0041) |
| economic $\sigma^2$ | −0.2382 | −1.6537** | −0.2993 | −0.5103 |
| | (1.7178) | (0.9466) | (1.9393) | (0.4112) |
| $R^2$ | 0.0044 | 0.0649 | 0.0054 | 0.0338 |
| observations | 47 | 47 | 47 | 47 |
| | | Panel C: Put options | | |
| constant | 0.0195 | 0.0284** | 0.0429*** | 0.0103** |
| | (0.0094) | (0.0127) | (0.0150) | (0.0044) |
| economic $\sigma^2$ | −0.3032 | −1.7438 | −3.0618** | −0.9396** |
| | (0.9422) | (1.2677) | (1.4990) | (0.4369) |
| $R^2$ | 0.0023 | 0.0412 | 0.0865 | 0.0951 |
| observations | 47 | 47 | 47 | 47 |

***, **, and * denote statistical significance at the 1, 5, and 10 percent levels, respectively.
This table shows the results of estimating the following regression:

$$(Oint_t - Oint_{t-1})/Oint_{t-1} = \alpha + \beta\sigma^2_{ECO,t} + e_t,$$

where $Oint_t$ is the open interest for option contracts with the closest quarterly maturity traded on the release day $t$. Column (1) shows the results for options on the 30-year Treasury futures (US), column (2) for options written on the 10-year Treasury futures (TY), column (3) for options written on the 5-year Treasury futures (FV), and column (4) for options written on the Eurodollar futures (ED).

be related to the degree of macroeconomic uncertainty. Table VIII presents the aggregated results for all near-term options in Panel A and the disaggregated results for near-term call and put options in panels B and C, respectively. Higher macroeconomic uncertainty is systematically associated with a greater decrease in open interest on the announcement day, which is consistent only with the hedging/speculating explanation. When we consider the open interest of all options or put options, the coefficients are more significant economically for options written on the five-year Treasury note futures, in line with the indications from the results of Table II and the results on trading volume of Table VII. The results for call options are more significant for options written on the ten-year Treasury note. Note, however, that the relation between macroeconomic uncertainty and changes in open interest is less important economically, compared to the results for the changes in trading volume. For example, a high degree of uncertainty about NFP equal to one standard deviation above the mean uncertainty predicts a 0.4 percent decrease in the open interest of put options written on the five-year Treasury note futures, compared to an unconditional increase of 1.4 percent on announcement days. This lower level of economic importance may be because the two explanations largely offset each other. As hedgers and speculators unwind their positions, other traders open new ones for reasons unrelated to the announcement, leaving open interest largely unchanged.

To further investigate trading strategies implemented with specific option types, we construct put-to-call ratios of volume and open interest and regress the difference from after to before the release on the level of macroeconomic uncertainty. This relation would be significant if calls or puts only are systematically used in conjunction with macroeconomic uncertainty. We do not find significant results, except in one case. For the open interest of put-to-call ratio of options on the 5-year Treasury note, the relation is significantly negative (results not reported). Higher macroeconomic uncertainty implies more put relative to call option positions unwound after the release. It is not clear why speculation would be systematically implemented with put options only. It seems more plausible that this finding is the outcome of hedging macroeconomic uncertainty of bond portfolios at intermediate maturities.

To have a better understanding of the relation between macroeconomic uncertainty and trading activity, we examine the patterns in option trading volume for several days surrounding the announcement. The aim of this analysis is to shed further light on whether the difference in trading volume from before to after the announcement is due to volume being unusually low before the announcement, a “calm before the storm” effect, or unusually high after the announcement, from the unwinding of hedge and speculative positions. Figure 3 depicts the evolution of trading volume in high and low macroeconomic uncertainty states corresponding to the top and bottom quartile of $\sigma_{ECO}$ in our sample. When uncertainty is high, trading volume generally dips moderately two days before the release (on Wednesdays) and then picks up somewhat before the release and very substantially so after the release. This graphical analysis suggests that in bond option markets we do not observe the “calm before the storm” effect documented for the underlying bonds, probably because option positions are opened before the release for hedging motives or speculative volatility-related motives that are greatest before macroeconomic announcements.

Finally, we examine the patterns of open interest for several days surrounding the announcement. Figure 4 depicts the evolution of open interest in high and low macroeconomic uncertainty states. In general, open interest increases after the

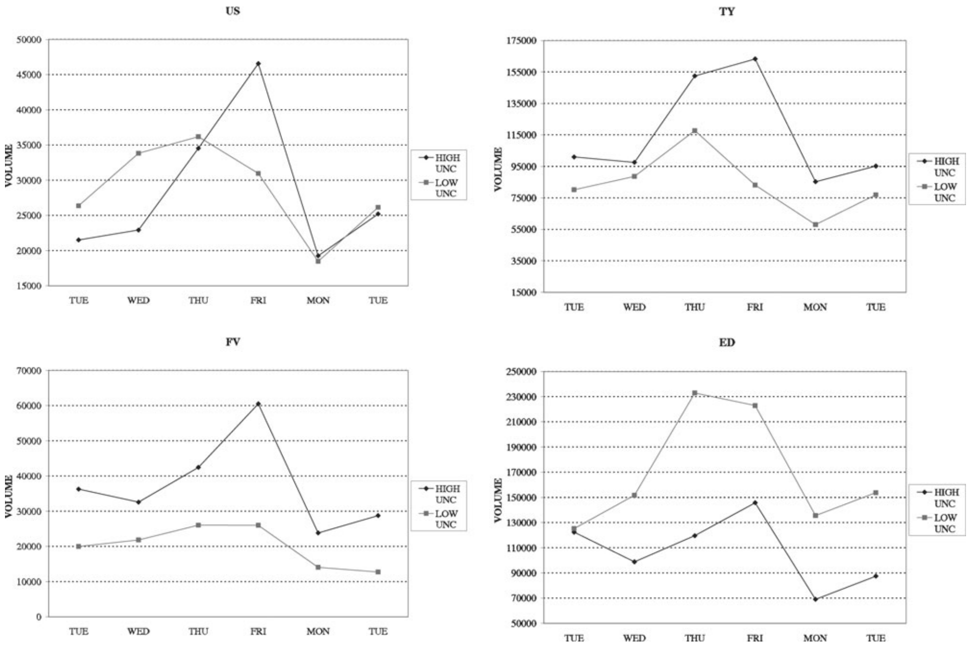


*Figure 3*. Bond Market Trading Volume around Macroeconomic Releases
This figure shows average trading volume for options on the 30-year Treasury futures (US), the 10-year Treasury futures (TY), the 5-year Treasury futures (FV), and the Eurodollar futures (ED) during the week surrounding the Friday NFP release, for high and low macroeconomic uncertainty releases.

release, suggesting that more positions are being opened when more is known about the state of the economy. When uncertainty is high, this increase in open interest is less substantial. This is an indication that some more positions are being closed with higher uncertainty, consistent with our earlier findings. The pattern during the week reveals a steady increase in opened positions until the release day, confirming our conjectures above that trading contracts, either for hedging or speculation, are opened gradually during the days preceding the announcement.

### *5.4.2 Aggregate stock index*

Turning next to the equity market, we first conduct the same analysis for trading volume and open interest of S&P 500 index options. Given the lack of an empirical relationship between macroeconomic uncertainty and changes in implied variance of index options, we do not expect to observe a relationship between macroeconomic uncertainty and trading activity. Nevertheless, the stock index results serve as

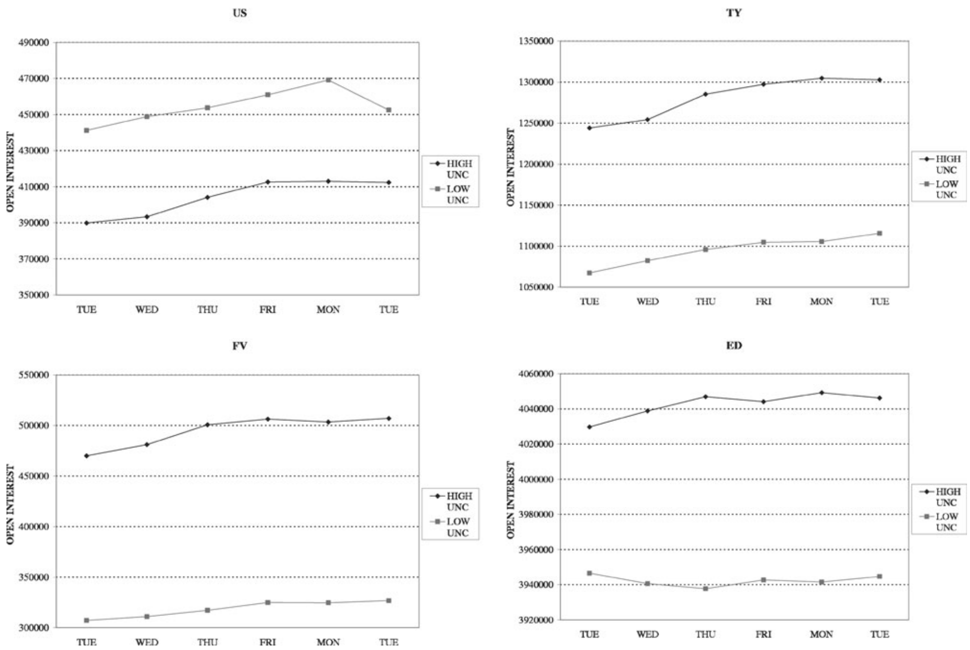


*Figure 4.* Bond Market Open Interest around Macroeconomic Releases
This figure shows average open interest for options on the 30-year Treasury futures (US), the 10-year Treasury futures (TY), the 5-year Treasury futures (FV), and the Eurodollar futures (ED) during the week surrounding the Friday NFP release, for high and low macroeconomic uncertainty releases.

natural benchmark for the more interesting findings involving cyclical stock options presented below.

Table IX shows the results for option trading volume. We do not report the results disaggregated for call and put options to conserve space. In all cases, there is a positive relationship between macroeconomic uncertainty and changes in trading volume of call and put options. However, none of the coefficients is statistically significant, and the explanatory power of the regressions is very modest. Table X presents the corresponding results for changes in open interest. In this table, the effect of macroeconomic uncertainty is generally negative. The coefficients for call and put options, which are not tabulated to conserve space, take different signs in some cases. Again, none of the coefficients is statistically significant, and the explanatory power of the regressions is very modest. We conclude that, as expected, there is no evidence of a relationship between macroeconomic uncertainty and trading activity in stock index options.

*Table IX.* S&P 500 index options volume

| | (1) short maturity | (2) medium maturity |
|---|---|---|
| average time to maturity (days) | 14.82 | 45.14 |
| average no. of traded options | 42.13 | 35.63 |
| constant | 0.0469 | 0.5871 |
| | (0.1829) | (0.2746) |
| economic $\sigma^2$ | 4.9998 | 28.8623 |
| | (17.6375) | (26.4783) |
| $R^2$ | 0.0282 | 0.0020 |
| observations | 43 | 43 |

***, **, and * denote statistical significance at the 1, 5, and 10 percent levels, respectively. This table shows the results of estimating the following regression:

$$(Volu_t - Volu_{t-1})/Volu_{t-1} = \alpha + \beta\sigma^2_{ECO,t} + e_t,$$

where $Volu_t$ is the trading volume for option contracts written on the S&P 500 index on the release day $t$. Column (1) shows the results for options written on the shortest traded maturity and column (2) for the next traded maturity.

*Table X.* S&P 500 index options open interest

| | (1) short maturity | (2) medium maturity |
|---|---|---|
| average time to maturity (days) | 14.82 | 45.14 |
| average no. of traded options | 42.13 | 35.63 |
| constant | 0.0465** | 0.0413 |
| | (0.0193) | (0.0339) |
| economic $\sigma^2$ | −2.9001 | −2.4503 |
| | (1.8673) | (3.2673) |
| $R^2$ | 0.0556 | 0.0135 |
| observations | 43 | 43 |

***, **, and * denote statistical significance at the 1, 5, and 10 percent levels, respectively. This table shows the results of estimating the following regression:

$$(Oint_t - Oint_{t-1})/Oint_{t-1} = \alpha + \beta\sigma^2_{ECO,t} + e_t,$$

where $Oint_t$ is the open interest for option contracts written on the S&P 500 index on the release day $t$. Column (1) shows the results for options written on the shortest traded maturity and column (2) for the next traded maturity.

### *5.4.3 Cyclical and non-cyclical stocks*

Since cyclical stocks react more strongly to macroeconomic news (see Table II) and their implied variances reflect the uncertainty about upcoming announcements (see Table VI), it seems logical to expect that trading activity in cyclical stock options is also more responsive to the degree of macroeconomic uncertainty. We therefore examine next the changes in trading volume and open interest of options written on

*Table XI.* Cyclical stocks options volume

| | (1) short maturity | (2) medium maturity |
|---|---|---|
| **Panel A: All options** | | |
| average time to maturity (days) | 14.82 | 45.14 |
| average no. of traded options | 66.97 | 98.09 |
| constant | −0.1340 | −0.2644 |
| | (0.1787) | (0.2091) |
| economic $\sigma^2$ | 31.3996** | 46.9940** |
| | (17.2260) | (20.1630) |
| $R^2$ | 0.0750 | 0.1170 |
| observations | 43 | 43 |
| **Panel B: Call options** | | |
| average time to maturity (days) | 14.82 | 45.14 |
| average no. of traded options | 34.33 | 52.30 |
| constant | −0.0864 | −0.2870 |
| | (0.2377) | (0.2504) |
| economic $\sigma^2$ | 29.3675 | 49.8051** |
| | (22.9229) | (24.1401) |
| $R^2$ | 0.0385 | 0.0941 |
| observations | 43 | 43 |
| **Panel C: Put options** | | |
| average time to maturity (days) | 14.82 | 45.14 |
| average no. of traded options | 32.65 | 45.79 |
| constant | −0.0904 | −0.2915 |
| | (0.2387) | (0.2391) |
| economic $\sigma^2$ | 34.1449* | 56.3910** |
| | (20.0152) | (23.0565) |
| $R^2$ | 0.0509 | 0.1273 |
| observations | 43 | 43 |

***, **, and * denote statistical significance at the 1, 5, and 10 percent levels, respectively. This table shows the results of estimating the following regression:

$$(Volu_t - Volu_{t-1})/Volu_{t-1} = \alpha + \beta\sigma^2_{ECO,t} + e_t,$$

where $Volu_t$ is the trading volume for option contracts written on cyclical stocks on the release day $t$. Column (1) shows the results for options written on the shortest traded maturity and column (2) for the next traded maturity.

cyclical stocks surrounding NFP announcements. Table XI documents the relation between our measure of macroeconomic uncertainty and changes in trading volume of all cyclical options (in Panel A) and of cyclical call and put options separately (in panels B and C). In all cases, greater macroeconomic uncertainty is associated with a larger increase in option trading volume immediately following the announcement, mirroring our results for bonds in Section 5.4.1. In aggregate as well as for calls and puts alone, this relation is strongest, both in economic magnitude and statistical

*Table XII.* Cyclical stocks options open interest

| | (1) short maturity | (2) medium maturity |
|---|---|---|
| | **Panel A: All options** | |
| average time to maturity (days) | 14.82 | 45.14 |
| average no. of traded options | 66.97 | 98.09 |
| constant | 0.0300*** | 0.0882*** |
| | (0.0071) | (0.0156) |
| economic $\sigma^2$ | −0.9487 | −4.0759*** |
| | (0.6871) | (1.5018) |
| $R^2$ | 0.0444 | 0.1523 |
| observations | 43 | 43 |
| | **Panel B: Call options** | |
| average time to maturity (days) | 14.82 | 45.14 |
| average no. of traded options | 34.33 | 52.30 |
| constant | 0.0227*** | 0.0933*** |
| | (0.0082) | (0.0202) |
| economic $\sigma^2$ | −0.5022 | −4.2669** |
| | (0.7868) | (1.9440) |
| $R^2$ | 0.0098 | 0.1051 |
| observations | 43 | 43 |
| | **Panel C: Put options** | |
| average time to maturity (days) | 14.82 | 45.14 |
| average no. of traded options | 32.65 | 45.79 |
| constant | 0.0515*** | 0.0650*** |
| | (0.0129) | (0.0258) |
| economic $\sigma^2$ | −2.4681** | −4.7236** |
| | (1.2480) | (2.4103) |
| $R^2$ | 0.0869 | 0.0801 |
| observations | 43 | 43 |

***, **, and * denote statistical significance at the 1, 5, and 10 percent levels, respectively. This table shows the results of estimating the following regression:

$$(Oint_t - Oint_{t-1})/Oint_{t-1} = \alpha + \beta\sigma^2_{ECO,t} + e_t,$$

where $Oint_t$ is the open interest for option contracts written on cyclical stocks on the release day $t$. Column (1) shows the results for options written on the shortest traded maturity and column (2) for the next traded maturity.

significance, for medium maturity options with an average of 46 days to maturity. For shorter maturities of 15 days on average, the results are almost as strong for all options and for put options. Table XII complements this analysis with corresponding results for changes in open interest. In all cases, greater macroeconomic uncertainty is associated with a larger drop in open interest following the announcement. In aggregate and for calls, this effect is significant only for medium maturities. For puts, the coefficients on macroeconomic uncertainty are significant, at the five-percent level, even for short maturities.

As argued in Section 4.4, these results are in principle again consistent with two different trading motives. Market participants could be using cyclical stock options to either hedge against or speculate on the economic release. Both activities predict that the announcement is followed by higher volume and a drop of open interest as hedge or speculative positions are unwound. Furthermore, in both cases the benefits from hedging or the potential gains from speculating increase with the degree of uncertainty about the announcement, leading to indistinguishable regression estimates. However, economic intuition makes hedging the more likely explanation of our results for cyclical stocks. This is because it seems far more likely that cyclical stock options are used to hedge against the effect of macroeconomic news on the cyclical stocks that underly these options, than that cyclical stock options are used to speculate broadly on announcements. After all, it is considerably cheaper to trade fixed income futures options than individual equity options, and the relation between bond returns and macroeconomic news is more than five times less noisy than that between cyclical stock returns and macroeconomic news (see Table II). This explanation is consistent also with the disaggregated results for call and put options. Taking for granted that the aggregate position in cyclical stocks is long, the most common hedging strategies would involve put options and the cheaper instruments would be at shorter maturities. In line with this prediction, the results of Table XI and XII for shorter maturity options are significant for put options. We conclude that, while our results for cyclical stocks are consistent with speculative trading, this explanation seems economically implausible.

To gain further insight into which of the hedging versus speculation explanations is more plausible, we regress again the difference in put-to-call ratios of volume and open interest on the level of macroeconomic uncertainty. We only find a significantly negative result for put-to-call ratios of the open interest of short-term options. This finding confirms our earlier conjecture of hedging portfolio of cyclical stocks with short-term put options when macroeconomic uncertainty is higher.

For completeness, we repeat this analysis one last time for non-cyclical stocks. The results, which are not tabulated to conserve space, confirm our earlier findings that non-cyclical stocks are even less related to macroeconomic announcements than the aggregate market index. Neither changes in option trading volume nor changes in open interest are significantly related to the uncertainty about NFP announcements. Furthermore, the regression coefficients on our measure of macroeconomic uncertainty take on different signs for different maturities and option types.

## 6 Conclusion

We established an empirical link between the ex-ante uncertainty about macroeconomic fundamentals and the ex-post resolution of this uncertainty in financial markets. We measured macroeconomic uncertainty using prices of economic derivatives and related this measure to changes in implied variances of stock and bond options when the economic data is released. Across the different assets we considered, we found that higher macroeconomic uncertainty is associated with greater reduction in implied variances of financial options. For bonds, the relationship between macroeconomic uncertainty and changes in implied variance is statistically and economically highly significant. Our uncertainty measure captures up to 40 percent of the variation of the change in implied variance across announcement days. Furthermore, a high degree of uncertainty equal to one standard deviation above the average uncertainty in our sample predicts more than a one-third drop in the implied variance of medium- to long-term bond options, compared to about a ten percent drop on announcement days with low uncertainty. The results are considerably weaker for the aggregate stock index. We showed, by further decomposing the aggregate stock index into cyclical and non-cyclical stocks, that these weaker results are largely due to non-cyclical stocks not responding to macroeconomic news. Cyclical stocks exhibit qualitatively the same significant relation between macroeconomic uncertainty and changes in implied variance as bonds, though quantitatively the results for cyclical stocks are still weaker.

We also examined the relationship between our measure of macroeconomic uncertainty and trading activity in stock and bond option markets surrounding the announcements. Higher macroeconomic uncertainty is associated with greater volume as well as larger drops in open interest of both bond and cyclical stock options following the announcements. These results are consistent with market participants using the option markets to either hedge against or speculate on the macroeconomic releases. We argued, however, that, at least in the case of cyclical stock options, the hedging explanation is economically more plausible.

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