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BERNANKE-KUTTNER REDUX

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Movements in Yields, not the Equity Premium: Bernanke-Kuttner Redux
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ABSTRACT

We show that the stock market price reaction to monetary policy surprises upon announcements of the Federal Open Market Committee (FOMC) is explained mostly by changes in the default-free term structure of yields, not by changes in the equity premium. We reach this conclusion based on a new model-free method that uses dividend futures prices to obtain the counterfactual stock market index price change that results purely from the change in the default-free yield curve induced by the monetary policy surprise. The yield curve change in turn partly reflects a change in expected future short-term interest rates, as measured by changes in professional forecasts, and partly a change in the term premium. We further find that the even/odd week FOMC cycle in stock index returns is also largely due to an FOMC cycle in the yield curve rather than the equity premium.

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1 Introduction

In their seminal study of the stock market reactions to monetary policy announcements of the Federal Open Market Committee (FOMC), [Bernanke and Kuttner \(2005\)](#) (BK) find that much of the stock market reaction to monetary policy surprises is explained by changes of expected excess returns and very little by changes in expected real interest rates. They interpret this result as evidence that monetary policy has effects on stock market risk, investor risk aversion, or investor sentiment. In this paper, we use a different and arguably better methodological approach based on data that was not available at the time of BK. We reach a very different conclusion: The bulk of the stock market reaction to monetary policy surprises in FOMC announcement windows is explained by changes in the default-free term structure of bond yields, without equity cash-flow risk premium effects.

Decomposing the stock market index movements in FOMC announcement windows into components due to equity premium changes and other sources is a challenging task. [Bernanke and Kuttner \(2005\)](#) do this by first estimating a vector autoregression (VAR) as in [Campbell and Ammer \(1993\)](#) that includes stock market excess returns and the dividend-price ratios, among other variables, with monthly data. In a second step, they regress the monthly VAR innovations on Federal funds rate surprises from within-month FOMC announcement windows. By iterating on the VAR, they then obtain longer-run impulse responses with the result that future expected excess returns respond positively to Federal funds rate surprises. What drives these VAR results is that when stock prices fall in the case of an unexpected monetary tightening, the level of dividends is sticky in the short run, and hence the dividend-price ratio rises. Based on the fact that movements in the dividend-price ratio are generally associated with movements in expected excess returns, this then leads to the conclusion that expected excess returns must have moved in response to the monetary surprise. However, the fact that the dividend-price ratio is *generally* associated with movements in expected excess returns, as captured by the VAR estimated on a long sample of monthly data, does not necessarily mean that high-frequency movements in the dividend-price ratio around a selected

set of *specific events*, such as FOMC announcements, are also associated with movements in expected excess returns. It could well be the case that the dividend-price ratio movements around these specific events are instead associated with, say, changing expectations of future real interest rates, even though this is not generally true for dividend-price ratio movements outside of FOMC announcement windows.

To make progress on this question, we use a different approach that does not rely on VAR estimates. Our starting point is the fact that the price of a stock market index share can be expressed as

$$P_t = \sum_{n=1}^{\infty} B_{n,t} G_{n,t}, \quad (1)$$

where $G_{n,t}$ is the price, at time t , of dividend futures maturing in period $t + n$, and $B_{n,t}$ is the price, at time t , of a zero-coupon bond maturing, with a one dollar payoff, in period $t + n$. In other words, no arbitrage implies that the current value of the index is equal to the sum of dividend futures prices discounted at the default-free zero-coupon yield curve. Equipped with data on futures prices, we can then ask: How much of price movements in FOMC announcement windows associated with monetary policy surprise measures can be explained just by movements in the yield curve, without any change in risk premia (other than those embedded in the default-free term structure)? We keep $G_{n,t}$ fixed at the pre-announcement day futures prices and we compute the implied change in P_t just based on the observed change in zero-coupon bond prices around the announcement. This approach gives us a model-free assessment of the contribution of changes in default-free yields to changes in stock prices. Unlike in BK, we do not have to estimate a VAR, we do not have to make an assumption that dividend-price ratio changes in FOMC announcement windows reflect the same components as dividend-price ratio changes in general, and we can fully account for any shape that the yield curve response to the monetary policy surprise might take.

What we find is very different from BK. Our estimates suggest that essentially all movements in the stock market index in FOMC announcement windows related to the monetary policy surprise can be explained just with movements in the yield curve. There is very

little room for additional effects from equity premium changes and the small component unaccounted for by yield curve changes is not statistically significant. This result holds for different definitions of the monetary policy surprise. Our baseline analysis uses the first principal component of unexpected changes in yields at five maturities less than one year as in [Nakamura and Steinsson \(2018\)](#), with expectations based on pre-announcement Federal funds and Eurodollar futures prices. As an alternative measure, we also use unexpected changes in the target Federal funds rate measured relative to pre-announcement Federal funds futures as in BK.

While essentially model-free, our approach based on dividend futures still requires some assumptions. Specifically, we only observe prices of dividend futures with maturity up to 7 years. In our baseline analysis, we assume that bond forward rates at maturities beyond 7 years do not change in response to the monetary surprise. We show that if this assumption holds, only dividend futures with maturity up to 7 years are needed to compute stock market index price changes implied by changes in the yield curve. Empirically, the bulk of forward rate reactions to FOMC announcements happens at maturities shorter than 7 years, so this assumption is reasonable. Nevertheless, to check robustness, we use the approach of [Knox and Vissing-Jørgensen \(2022\)](#) to estimate futures prices at maturities beyond 7 years under a Gordon growth model assumption. Our results continue to hold with this alternative measure. We also consider an approach that does not use dividend futures prices. In the [Campbell and Shiller \(1988\)](#) log-linear present value framework, the discount factor ρ is a function of the average dividend-price ratio and it captures the duration of the stock market index. With ρ calibrated to the observed average dividend-price ratio, and using observed changes in the forward rate curve in the FOMC announcement window, this framework allows us to calculate another measure of stock market index price changes implied by changes in the yield curve. We obtain results very similar to our baseline analysis using dividend futures. Moreover, both the dividend futures approach and the Campbell-Shiller framework can be applied to assess stock price changes within 30-minute windows around FOMC announcements, using high-

frequency yield responses. The conclusions align closely with those from the daily analysis.

We then use interest rate forecasts from the BlueChip survey of professional forecasters to decompose the yield curve reaction around FOMC announcements into changes in expected future short-term interest rates and changes in the term premium. The results differ depending on the monetary policy surprise measure. For the Nakamura-Steinsson measure, we find that each component accounts for about half of the reaction. For the Federal funds rate surprise measure, we attribute almost the entire reaction to the term premium. That the term premium component plays a substantial role is consistent with the finding in [Cieslak and Pang \(2021\)](#), based on structural vector autoregressions, that monetary policy surprises have a large effect on the common term premium in stocks and bonds.

Finally, the dominant role of yield curve changes in explaining the stock price reaction to monetary policy surprises motivates us to examine another intriguing regularity in stock prices related to FOMC: the stylized fact documented in [Cieslak et al. \(2019\)](#) that average returns are much higher in even weeks than odd weeks in FOMC cycle time. We find that stock index price changes implied by yield curve changes over the FOMC cycle account for a substantial share of the FOMC cycle in stock returns. Hence, the puzzling FOMC cycle effect in stock prices is, to a substantial extent, really the puzzle that yield curve changes are predictable with FOMC cycle time.

Our finding that yield curve changes mostly explain why stock prices comove with monetary policy surprises induced by FOMC announcements leaves little room for an equity risk premium effect of monetary policy surprises. This also implies that the comovement of risk appetite measures with monetary policy surprises observed by [Bauer et al. \(2023\)](#) and [Bekaert et al. \(2013\)](#) does not appear to be an equity risk premium effect.

This is relevant for the theoretical literature that has been spawned by BK's findings. Researchers have aimed for models of asset prices and the macroeconomy in which monetary policy surprises move the equity premium. [Pflueger and Rinaldi \(2022\)](#) focus on time-varying risk aversion. In [Kekre and Lenel \(2022\)](#), monetary policy surprises have a time-varying effect

on aggregate risk-bearing capacity by changing the wealth distribution. In [Drechsler et al. \(2017\)](#), monetary policy affects the liquidity premium and hence cost of capital of levered agents and their willingness to bear risk. Our findings are better aligned with models in which monetary policy surprises move stock prices without changes in the equity risk premium, such as [Bianchi et al. \(2022a\)](#), [Caballero and Simsek \(2023\)](#), and [Nakamura and Steinsson \(2018\)](#).

That said, we focus, like BK, on regressions of the stock market announcement window returns on monetary policy surprise measures constructed from unexpected changes in interest rates. Since monetary policy surprise measures are far from perfectly correlated with stock market index returns within announcement windows, there must be other FOMC-related news that moves stock prices but is not captured by conventional monetary policy surprise measures. [Kroencke et al. \(2021\)](#) and [Boehm and Kroner \(2024\)](#) study these components of asset price reactions orthogonal to default-free interest rates. Furthermore, evidence in [Bianchi et al. \(2022b\)](#) and [Cieslak and McMahon \(2023\)](#) indicates that monetary policy news may affect the equity premium in intermeeting time periods between FOMC announcements. Our findings suggest that if monetary policy affects the equity risk premium, it is likely through such alternative channels unrelated to default-free interest rate movements in FOMC announcement windows. Theoretically, it is also possible that the effects of equity premium changes cancel with valuation effects of changes in cash flow expectations, but in most existing theories of a time-varying equity premium the effects would work in the same direction.

Methodologically, our decomposition of announcement window stock market index returns is related to, but different from, the decomposition of [Knox and Vissing-Jørgensen \(2022\)](#). They also use dividend futures data, but they are interested in a decomposition into stock price changes due to risk-free rate news, news about expected excess returns, and news about cash flows. Their decomposition requires a number of assumptions. As we show, if one is interested only in stock price changes due to yield curve changes, such assumptions are not necessary and the yield curve effect on stock prices can be obtained in an essentially model-free way.

Finally, our finding that yield curve movements matter a lot more for the stock market reaction to monetary policy surprises than previously thought is also related to the finding in [van Binsbergen \(2020\)](#) that large changes in real interest rates during the past few decades seem to account for a lot of the returns that stock market indices earned during these periods.

2 Decomposition of announcement returns

We start by developing our decomposition of announcement returns based on dividend futures prices. In what follows, all quantities are quoted in nominal terms unless otherwise specified.

Let M_{t+j} be the stochastic discount factor (SDF) and consider the dividend stream of the stock market broken up into its individual pieces (strips). For example, the price at t of a dividend strip that pays the stock market index dividend D_{t+2} at $t + 2$ is

$$\begin{aligned} P_{2,t} &= \mathbb{E}_t[M_{t+1}M_{t+2}D_{t+2}] \\ &= \mathbb{E}_t[M_{t+1}M_{t+2}]\mathbb{E}_t[D_{t+2}] + \text{cov}_t(M_{t+1}M_{t+2}, D_{t+2}). \end{aligned} \quad (2)$$

Denoting with $B_{2,t} = \mathbb{E}_t[M_{t+1}M_{t+2}]$ the price of a default-free zero-coupon bond that pays \$1 in two periods, we can write this as

$$P_{2,t} = B_{2,t} \underbrace{\left[\mathbb{E}_t[D_{t+2}] + \text{cov}_t \left(\frac{M_{t+1}M_{t+2}}{\mathbb{E}_t[M_{t+1}M_{t+2}]}, D_{t+2} \right) \right]}_{G_{2,t}}. \quad (3)$$

The strip price therefore depends on the default-free yield curve, via $B_{2,t}$, expected dividends $\mathbb{E}_t[D_{t+2}]$ and the equity risk premium at two-period horizon captured by the covariance term.¹ According to spot-future parity, the sum of the two term in brackets in (3) represents the dividend futures price $G_{2,t}$.

¹ Note that if shocks to the SDF are multiplicative, the ratio $\frac{M_{t+1}M_{t+2}}{\mathbb{E}_t[M_{t+1}M_{t+2}]}$ within the covariance extracts innovations to the SDF and its conditional mean. So when the strip price changes, it could be due to changes in the yield curve (and hence $B_{2,t}$), changes in expected dividends, or changes in the equity risk premium.

Going through the same calculation at other horizons allows us to write the price of the entire dividend stream of a share of the stock market index as

$$P_t = \sum_{n=1}^{\infty} P_{n,t} = \sum_{n=1}^{\infty} B_{n,t} G_{n,t}. \quad (4)$$

Now consider a change in zero-coupon bond prices from $B_{n,t-}$ to $B_{n,t+}$ with expected dividends and the equity premium and hence dividend futures prices remaining unchanged. The implied percentage price change of the stock market index is

$$\Delta P_{B,t} \equiv \sum_{n=1}^{\infty} \frac{G_{n,t-}}{P_{t-}} (B_{n,t+} - B_{n,t-}). \quad (5)$$

This is the key variable for our baseline analysis. We look at the change $B_{n,t+} - B_{n,t-}$ around FOMC announcements, holding the futures price fixed at the observed pre-announcement value $G_{n,t-}$.

In practice, we only have futures data up to a maturity of 7 years. We deal with this issue by making the following assumption:

$$\frac{B_{n,t+}}{B_{n,t-}} = \frac{B_{m+1,t+}}{B_{m+1,t-}}, \quad \forall n > m+1, \forall t, \quad (6)$$

where m denotes the observable maximum maturity on the dividend futures prices. Given that $B_{n,t} = \exp(-ny_{n,t})$, this amounts to the assumption that

$$f_{n,m+1,t+} - f_{n,m+1,t-} = 0, \quad \forall n > m+1, \forall t, \quad (7)$$

where $f_{n,m,t} \equiv \frac{1}{n-m}(ny_{n,t} - my_{m,t})$ denotes the log forward rate m periods into the future and paying off at n . In other words, we assume forward rates far out in the future do not change on FOMC announcement days. This is plausible, as monetary policy is typically not thought to have such extremely long-horizon effects. Moreover, we present evidence below consistent with this assumption and conduct formal tests to validate it. We also show that

this difference does not comove with monetary shocks.

Now with (6), we have

$$\begin{aligned}
\Delta P_{B,t} &= \sum_{n=1}^m \frac{G_{n,t-}}{P_{t-}} (B_{n,t+} - B_{n,t-}) + \sum_{n=m+1}^{\infty} \frac{G_{n,t-}}{P_{t-}} (B_{n,t+} - B_{n,t-}) \\
&= \sum_{n=1}^m \frac{G_{n,t-} B_{n,t-}}{P_{t-}} \left(\frac{B_{n,t+}}{B_{n,t-}} - 1 \right) + \sum_{n=m+1}^{\infty} \frac{G_{n,t-} B_{n,t-}}{P_{t-}} \left(\frac{B_{m+1,t+}}{B_{m+1,t-}} - 1 \right) \\
&= \sum_{n=1}^m \frac{G_{n,t-} B_{n,t-}}{P_{t-}} \left(\frac{B_{n,t+}}{B_{n,t-}} - 1 \right) + \left(1 - \sum_{n=1}^m \frac{G_{n,t-} B_{n,t-}}{P_{t-}} \right) \left(\frac{B_{m+1,t+}}{B_{m+1,t-}} - 1 \right). \quad (8)
\end{aligned}$$

Note that we can observe all the right-hand side components using prices on the market index, dividend futures, and zero-coupon bonds. Essentially, the assumption (6) implies that the value of all the dividend strips at maturities $n > m$ changes around the announcement by the same factor. Since we can figure out the total value of these strips by subtracting the value of the first m strips from the total index value P_t , we can also figure how much their value changes when they are all multiplied by the same factor.

As the market for dividend futures may be quite illiquid, it is advantageous that accurate measurement of dividend futures prices is not critical for the calculation in (8). The dividend futures prices act only as weights, so if they happen to be mismeasured by a few percent, it has minimal impact on the outcome of the calculation. In contrast, accurate measurement of zero-coupon bond prices is important, as the differences $B_{n,t+} - B_{n,t-}$ are much more sensitive to measurement errors than the weights are.

To check the joint contribution of changes in expected dividends and the cash flow risk premium, we can compute the implied percentage price change by holding bond prices fixed at $B_{n,t-}$ and looking at the changes in futures prices, i.e.,

$$\Delta P_{G,t} \equiv \sum_{n=1}^{\infty} \frac{B_{n,t-}}{P_{t-}} (G_{n,t+} - G_{n,t-}). \quad (9)$$

In practice, we simply use the available dividend futures to arrive at

$$\Delta P_{G,t} = \sum_{n=1}^m \frac{B_{n,t-}}{P_{t-}} (G_{n,t+} - G_{n,t-}). \quad (10)$$

This calculation implicitly assumes that expectations about dividends beyond maturity m do not change in response to the news conveyed by the FOMC announcement. As a caveat, this calculation is more sensitive to dividend futures price mismeasurement than the calculation in (8) that is of main interest in our analysis.

For robustness, we also calculate two alternative measures of $\Delta P_{G,t}$. One of the alternative measures extends the available maturities of $G_{n,t}$ by estimating dividend futures prices for $n > 7$ under a Gordon growth model following [Knox and Vissing-Jørgensen \(2022\)](#). [Appendix A.2](#) provides more details.

2.1 Alternative approach based on Campbell-Shiller approximate present-value identity

Let p_t and d_t denote the (per-share) log value and dividends of the stock market index and r_t the log return. The Campbell-Shiller approximate present-value identity is

$$p_t \approx \frac{\kappa}{1 - \rho} + \mathbb{E}_t \sum_{n=1}^{\infty} \rho^{n-1} [(1 - \rho)d_{t+n} - r_{t+n}], \quad (11)$$

where $\rho = 1/(1 + \exp(\overline{d - p}))$ and $\kappa = -\log \rho - (1 - \rho) \log(1/\rho - 1)$. The key parameter here is the log-linearization parameter ρ . It is a function of the mean log dividend-price ratio $\overline{d - p}$, which captures the duration of the stock market index. For example, if expected dividend growth is higher, $d - p$ is lower, and duration is higher. In our analysis, we use data for the log dividend-price ratio of the CRSP value-weighted index from 1926 to 2023 to estimate $\overline{d - p} = -3.41$. We use this value in our calculations.

We can now use the present-value identity to obtain stock price changes implied by changes in the yield curve. For this purpose, we decompose the return r_{t+n} into an excess return and

a forward rate. Let $f_{n,t}$ be the time- t log forward rate for the period between $t + n - 1$ to $t + n$,

$$f_{n,t} = ny_{n,t} - (n - 1)y_{n-1,t}, \quad (12)$$

where $y_{n,t}$ is the continuously compounded zero-coupon yield on a n -maturity bond at time t . Define $x_{n,t+n} \equiv r_{t+n} - f_{n,t}$ as the excess return of the stock index in $t + n$ relative to the time- t forward rate for that future period. Then

$$p_t = \frac{\kappa}{1 - \rho} + \mathbb{E}_t \sum_{n=1}^{\infty} \rho^{n-1} [(1 - \rho)d_{t+n} - x_{n,t+n}] - \sum_{n=1}^{\infty} \rho^{n-1} f_{n,t}. \quad (13)$$

We can calculate the log price change, from p_{t-} just before the FOMC announcement to p_{t+} thereafter, as:

$$\Delta p_t \equiv p_{t+} - p_{t-} = (\mathbb{E}_{t+} - \mathbb{E}_{t-}) \sum_{n=1}^{\infty} \rho^{n-1} [(1 - \rho)d_{t+n} - x_{n,t+n}] + \sum_{n=1}^{\infty} \rho^{n-1} (f_{n,t-} - f_{n,t+}). \quad (14)$$

The last term in the above equation represents the log price change implied by a change in the yield curve, without changes in expectations of the stream of d_{t+n} or $x_{n,t+n}$:

$$\Delta p_{F,t} \equiv \sum_{n=1}^{\infty} \rho^{n-1} (f_{n,t-} - f_{n,t+}). \quad (15)$$

In practice, we only have forward rates up to a 30-year horizon. We therefore calculate

$$\Delta p_{F,t} = \sum_{n=1}^{30} \rho^{n-1} (f_{n,t-} - f_{n,t+}). \quad (16)$$

This amounts to assuming that forward rates beyond the 30-year horizon do not change in response to FOMC announcements.

3 Data

This section introduces the data we use in the empirical analysis.

3.1 FOMC announcements and monetary surprise measures

Our main measure of monetary surprises, POLICY, follows [Nakamura and Steinsson \(2018\)](#) and is the first principal component of price changes, in a 30-minute window around the FOMC announcement, of five interest rate futures contracts with maturities of less than 1 year (Federal funds futures expiring at the end of the month in which the FOMC meeting takes place and those expiring after the next FOMC meeting; 3-month Eurodollar futures with maturities of 1, 2, and 3 quarters ahead). We also consider an alternative measure, FFR, that uses just the change in the Federal funds futures price of the contract expiring at the end of the month after the FOMC meeting. The latter measure is the same as the one used by BK, but here with intraday instead of daily data.

For both surprise measures, we use the updated series from [Acosta \(2023\)](#) which are available between February 1995 and December 2023 for scheduled announcements.² This means most of our analysis involving monetary surprises is focused on the period after 1994 when the Fed started to publicly announce changes in the Federal funds rate target following each meeting.

Panel A, Table 1 reports the summary statistics of these two monetary surprise measures. The POLICY shock series is already standardized to unit standard deviation. We also standardize the FFR series before we use it in our empirical analysis.

3.2 Asset prices

Daily nominal yields on Treasury zero-coupon bonds with maturities up to 30 years are obtained from [Filipović et al. \(2022\)](#) and are available starting from June 1961. We also obtain the intraday responses of Treasury yields in the 30-minute window around FOMC announcements as measured from futures prices from [Bauer and Swanson \(2023a\)](#).³ The available maturities include 2-, 5-, 10-, and 30-year.

² We thank the author for providing data on the website.

³ We thank the authors for providing data on the website.

TABLE 1
Summary Statistics

In Panel A, POLICY is the first principal component of 30-minute price changes of five interest-rate futures prices with maturities of less than 1 year as in Nakamura and Steinsson (2018) and as updated by Acosta (2023). FFR denotes the price changes, in percentage points, of Federal funds futures expiring at the end of the month after the FOMC meeting. In Panel B, ΔP and Δp are the actual percentage and log price changes of the S&P 500 index around scheduled FOMC announcements, respectively. ΔP_B and Δp_F are the corresponding counterfactual percentage and log price changes implied by yield changes using the dividend futures method or the Campbell-Shiller present-value identity, respectively. ΔP_G denotes the counterfactual percentage price changes using only changes in dividend futures prices while holding the yield curve fixed. In Panel C, ΔP^H , Δp^H , ΔP_B^H , and Δp_F^H are the high-frequency counterparts of actual and counterfactual percentage and log price changes of the S&P 500 index around FOMC announcements in 30-minute windows using data from Bauer and Swanson (2023a). All price changes are quoted in basis points.

	Mean	S.D.	Median	5th	95th
A. Monetary Surprise Measures					
POLICY	0	1	0.08	-1.72	1.54
FFR	-0.40	3.87	0	-6.35	5.64
B. Announcement-Day Price Changes					
Oct 2002–Dec 2023:					
ΔP	31.79	123.04	11.28	-136.10	261.56
ΔP_B	7.82	62.51	6.57	-84.97	131.08
ΔP_G	-0.72	8.68	0.16	-13.59	8.28
Feb 1994–Dec 2023:					
Δp	26.55	115.70	11.28	-155.15	213.80
Δp_F	14.43	117.65	2.49	-146.29	198.32
C. High-Frequency Price Changes					
Nov 2002–Dec 2019:					
ΔP^H	2.88	59.34	1.76	-95.48	91.42
ΔP_B^H	12.27	85.19	3.17	-33.77	59.73
Feb 1994–Dec 2019:					
Δp^H	-2.17	56.58	-4.71	-93.37	82.45
Δp_F^H	2.28	73.89	0.85	-97.08	101.77

For the aggregate stock market prices, we use daily prices on the S&P 500 index from CRSP. The intraday responses of the stock market in the 30-minute window around FOMC announcements as measured from futures prices come from [Bauer and Swanson \(2023a\)](#). Daily prices of dividend futures on the S&P 500 index between October 2002 and August 2014 come from [van Binsbergen and Koijen \(2017\)](#).⁴ We supplement the series with daily prices of S&P 500 Annual Dividend Index Futures since November 2015 from Bloomberg (mnemonics “ASDZXXIndex” where “XX” denotes the maturing year). These dividend futures contracts are listed for the nearest 11 years and expire on the third Friday of December. We linearly interpolate the log futures prices between two maturities to obtain constant-maturity futures prices up to a maturity of 7 years.⁵

Panel B, Table 1 reports the summary statistics for the actual and counterfactual stock market index price changes on FOMC announcement days that we construct from these data. In the shorter sample from 2002 to 2023 in which we have dividend futures data available, the counterfactual percentage price changes just based on yield curve changes using the dividend futures method, ΔP_B , also have high volatility, but only about half of the volatility of actual percentage price changes. In contrast, the counterfactual price changes ΔP_G that keep the yield curve fixed and use only changes in dividend futures prices have a much smaller standard deviation. Based on the long sample from 1994 to 2023, the counterfactual log price changes implied by yield curve changes using the Campbell-Shiller method, Δp_F , have almost the same volatility as the actual log price changes Δp . These summary statistics are therefore already a hint that yield curve changes could potentially explain a lot of the stock index price reaction to FOMC announcements.

⁴ We thank the authors for kindly sharing the data.

⁵ We follow [van Binsbergen and Koijen \(2017\)](#) to look at futures with maturities up to 7 years as longer maturity futures are illiquid. S&P 500 Annual Dividend Index Futures also have very low open interests beyond year 7.

3.3 Survey forecasts

Monthly forecasts of macroeconomic variables and Treasury bill rates come from two surveys published by Wolters Kluwer, the *Blue Chip Economic Indicators* (BCEI) and the *Blue Chip Financial Forecasts* (BCFF). The BCEI survey is typically released on the 10th day of each calendar month with responses collected during the first week of the same month. The BCFF survey is typically released on the first day of each calendar month with responses collected during the last week of the previous month. From BCEI, we collect forecasts of real GDP, GDP price index, unemployment rate, and 3-month Treasury bill rates. From BCFF, we collect forecasts of 3-month Treasury bill rates. Both surveys report forecasts of quarterly averages in future quarters. Forecast revisions are calculated as monthly changes in quarterly forecasts.⁶

We also obtain long-range forecasts of 3-month Treasury bill rates from both surveys. These forecasts are reported bi-annually by survey participants and are available for the next year and up to 6 years ahead. The forecasters also report a long-term 5-year average for horizons between 7 to 11 years ahead. The long-range surveys are typically conducted in March and October. BCFF switched the survey months to June and December in December 1996.

4 Asset price responses to monetary policy surprises

Before looking at the stock market, we start by examining the response of Treasury yields to monetary policy surprises on FOMC announcement days. In our method based on dividend futures, we only have dividend futures data up to maturities of 7 years. For this reason, the dividend futures method would not capture the effects of changes in forward rates at horizons beyond 7 years.

⁶ For example, to compute the revisions in the 1-quarter-ahead forecasts from March (last month of the quarter) to April (first month of the next quarter), we subtract the 2-quarter-ahead forecasts in March from the 1-quarter-ahead forecasts in April. For the other two months within a calendar quarter, we simply use the difference in 1-quarter-ahead forecasts.

4.1 Treasury yield response

Panel A in Table 2 shows that monetary policy surprises have effects on the yield curve that stretch quite far out in the term structure of zero-coupon yields. In terms of point estimates, the POLICY surprise measure is associated with a yield change at the 20-year maturity that is still about a third as big as the yield change at the 1-year maturity.

However, as Panel B shows, in terms of forward rates, the response is clearly concentrated at maturities up to 5 years.⁷ For our dividend futures method, the crucial assumption is that forward rates beyond a horizon of 7 years do not move in response to monetary policy surprise news in FOMC announcements. The results here support this assumption. Table A.1 in Appendix A.1 provides formal statistical tests of this assumption.

4.2 Stock market response

We now turn to our main analysis, revisiting the conclusions of BK about the effects of monetary policy surprises on the equity premium. We arrive at a very different conclusion.

BK regress the CRSP value-weighted index returns on unexpected changes in the Federal funds rate target and find that an unanticipated 25-bp cut in the Federal fund rates target is associated with a 1% increase in the stock market index. Here we have a sample that covers a much longer sample and, to match with our dividend futures data, we use the S&P 500 index instead of the CRSP value weighted index, but we find a broadly similar result. In the long sample in Panel B of Table 3 we find that an unanticipated 25-bp cut in the Federal fund rates target, which represents a 6.46 s.d. movement from the sample mean, is associated with a $6.46 \times 20.6 \approx 133$ bp increase in the stock market index. Using the POLICY surprise

⁷ Using a shorter sample from 1999 to 2012, [Hanson and Stein \(2015\)](#) find a somewhat stronger, but still statistically insignificant response at the 20-year horizon. [Kekre et al. \(2024\)](#) find that nominal forward rates respond negatively to monetary shocks at a horizon of more than 20 years. We replicate their finding in Appendix B for their sample period and with their monetary surprise measure based on [Jarociński and Karadi \(2020\)](#), but we show this result appears only with forward rates in the [Gürkaynak et al. \(2007\)](#) (GSW) dataset, and not with the [Filipović et al. \(2022\)](#) forward rates that we use in our analysis. [Filipović et al. \(2022\)](#) show that the forward rates obtained with their method are more precise and robust than those based on the parametric GSW method. Relatedly, [Liu and Wu \(2021\)](#) point out that the the long end of the yield curve in the parametric GSW method is subject to extrapolation errors.

TABLE 2
Bond Market Response to Monetary Policy Surprises

This table reports the response of zero-coupon yields and forward rates extracted from nominal Treasury securities to monetary policy surprises on scheduled FOMC announcement days. In Panel A, the dependent variables are 1-day changes in zero-coupon Treasury yields. In Panel B, the dependent variables are 1-day changes in 1-year forward rates. All dependent variables are quoted in basis points. The monetary policy surprise measures include the policy news shock from [Nakamura and Steinsson \(2018\)](#) as updated by [Acosta \(2023\)](#) (POLICY) and the unexpected changes in the target Federal funds rates (FFR). Both surprises are standardized over the full sample period to have unit standard deviations. All regressions include a constant term that is not reported. We report the t -statistics calculated using heteroskedasticity-robust standard errors in brackets. The sample period runs from February 1995 to December 2023.

Panel A: Δy_n												
	1Y		2Y		5Y		10Y		20Y		30Y	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
POLICY	3.72 [11.04]		4.42 [9.29]		3.98 [8.27]		2.69 [5.01]		1.45 [2.92]		1.01 [1.94]	
FFR		2.17 [5.26]		1.82 [3.21]		1.56 [2.91]		1.09 [1.97]		0.61 [1.23]		0.54 [1.02]
N	230	230	230	230	230	230	230	230	230	230	230	230
Adj. R^2	0.45	0.15	0.39	0.06	0.25	0.03	0.13	0.02	0.05	0.01	0.02	0.00
Panel B: Δf_n												
	1Y		2Y		5Y		10Y		20Y		30Y	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
POLICY	3.72 [11.04]		5.11 [7.24]		2.90 [4.66]		1.30 [1.89]		0.00 [0.00]		0.45 [0.65]	
FFR		2.17 [5.26]		1.47 [1.92]		1.25 [2.13]		0.77 [1.20]		0.82 [1.00]		0.30 [0.41]
N	230	230	230	230	230	230	230	230	230	230	230	230
Adj. R^2	0.45	0.15	0.30	0.02	0.10	0.01	0.03	0.01	-0.00	0.01	-0.00	-0.00

measure instead, the regression produces a statistically significant coefficient of -23.28 (t -stat -2.22). Combined with column (1) in Panel A of Table 2, this point estimate implies that a POLICY shock that decreases the 1-year nominal Treasury yield by 25 bp is associated with a $23.28 \times (25/3.72) \approx 156$ bp increase in the stock market index.

To find out how much of the stock market response can be attributed to changes in the default-free yield curve without changes in the equity premium, we repeat the same regression, but now using our counterfactual price change measures ΔP_B , based on dividend futures, in Panel A, and Δp_F , based on the Campbell-Shiller present-value identity method, in Panel B.

Columns (3) and (4) in Panel A show that the magnitude of coefficients are very similar to the coefficients in columns (1) and (2). Specifically, when using the POLICY surprise measure, we have an estimate of -25.47 , which is close in magnitude to the estimate of -36.84 in column (1). Furthermore, as column (7) shows, the difference is not statistically significant at conventional levels. The same is true for the FFR surprise measure.

With the longer sample and the Campbell-Shiller present-value identity method in Panel B, the counterfactual price changes implied by yield curve changes are even closer to the actual price changes. For the POLICY surprise measure, the difference in slope coefficient point estimates is only about 3 bp.

Columns (5) and (6) in Panel A show how prices would have changed if only dividend futures prices had changed, with no change in the yield curve. For this analysis we hold $B_{n,t}$ fixed at pre-announcement values and look at changes in implied prices due to changes in $G_{n,t}$. The results show that the price change implied by dividend futures price changes is basically zero.⁸ This further confirms our central result that yield curve movements, not movements in the equity premium, explain stock prices changes in FOMC announcement windows that are correlated with monetary policy surprise measures. However, this analysis based on dividend futures price changes is more sensitive to measurement error in dividend

⁸ Golez and Matthies (2023) find that very short-horizon dividend claim prices do seem to move in response to monetary policy surprises. However, our calculation here combines dividend futures over the whole term structure many years out.

TABLE 3
Stock Market Response to Monetary Policy Surprises

The first two columns report the OLS estimates from regressing percentage price changes in Panel A and log changes in Panel B of the S&P 500 market index on monetary policy surprises on scheduled FOMC announcement days. The monetary policy surprise measures include the policy news shock from [Nakamura and Steinsson \(2018\)](#) as updated by [Acosta \(2023\)](#) (POLICY) and the unexpected changes in the target Federal funds rates (FFR). Both surprises are standardized to have unit standard deviations. Columns (3) and (4) use ΔP_B as the dependent variable in Panel A and Δp_F in Panel B. Columns (5) and (6) in Panel A use ΔP_G as the dependent variable. The rest of columns test the difference between slope coefficients. All dependent variables are quoted in basis points. We report the t -statistics calculated using heteroskedasticity-robust standard errors in brackets. The sample period runs from November 2002 to December 2023 in Panel A and from February 1995 to December 2023 in Panel B.

Panel A: Dividend Futures Method

	ΔP		ΔP_B		ΔP_G		$\Delta P - \Delta P_B$		$\Delta P - \Delta P_G$	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
POLICY	-36.84		-25.47		0.52		-11.37		-37.36	
	[-2.53]		[-4.50]		[0.72]		[-0.81]		[-2.51]	
FFR		-24.33		-10.48		-0.95		-13.84		-23.37
		[-1.37]		[-1.54]		[-0.75]		[-0.77]		[-1.25]
Constant	32.91	32.03	8.56	7.74	-0.78	-0.72				
	[3.39]	[3.25]	[1.84]	[1.57]	[-1.12]	[-1.05]				
N	158	158	158	158	158	158				
Adj. R^2	0.07	0.02	0.14	0.02	-0.00	0.00				

Panel B: Campbell-Shiller PV Identity Method

	Δp		Δp_F		$\Delta p - \Delta p_F$	
	(1)	(2)	(3)	(4)	(5)	(6)
POLICY	-23.28		-26.48		3.20	
	[-2.22]		[-2.69]		[0.28]	
FFR		-20.60		-12.45		-8.15
		[-1.94]		[-1.25]		[-0.62]
Constant	27.33	27.35	11.75	11.78		
	[3.63]	[3.62]	[1.58]	[1.55]		
N	230	230	230	230		
Adj. R^2	0.04	0.03	0.05	0.01		

futures prices than our main analysis, which uses the dividend futures prices as weights for bond price changes, and should therefore be interpreted with more caution.

In summary, our findings are very different from BK. Based on their VAR method, BK attribute most of the stock market response to changes in the equity premium. However, their approach is based on the strong assumption that dividend-price ratio movements around FOMC days are associated with movements in expected excess returns to exactly the same extent as on all other non-FOMC days. Our methods do not require such a strong assumption; they are model-free in that they do not require estimation of a VAR to back out changes in the equity premium. Based on this model-free approach, we find that almost all of the stock market response to monetary policy surprises on FOMC announcement days can be attributed to changes in the default-free yield curve.

4.3 Attribution to short- and long-horizon yield curve changes

Since the implied price changes ΔP_B and Δp_F are weighted averages of zero-coupon yield and forward rate changes at different maturities, we can further study whether short- or long-maturity components mainly contribute to the implied stock market response. We do so by regressing components of the summations that yield ΔP_B and Δp_F in equations (5) and (15) on the monetary policy surprise measures. This allows us to conduct a further check of the assumption that our dividend futures method is based on.

Specifically, for ΔP_B , we regress the $\frac{G_{n,t-}}{P_{t-}}(B_{n,t+} - B_{n,t-})$ terms at different maturities n on POLICY and FFR. For horizons beyond 7 years, we approximate $\frac{G_{n,t-}}{P_{t-}}(B_{n,t+} - B_{n,t-})$ using the [Knox and Vissing-Jørgensen \(2022\)](#) method as outlined in [Appendix A.2](#). [Figure 1](#) shows the result. In both panels, the negative coefficients flatten out at maturities close to 7 years (in fact, the magnitudes even revert a bit for POLICY in panel A, but the estimates at very long maturities come with big standard errors). Our dividend futures method assumes this flattening out to get to expression (8). This requires that forward rates beyond 7 years do not change in response to FOMC announcements and, as a consequence, the $\frac{G_{n,t-}}{P_{t-}}(B_{n,t+} - B_{n,t-})$

all change by the same factor beyond the 7-year maturity.

Figure 2 shows similar results for the Campbell-Shiller present-value identity method. We regress $\rho^{n-1}(f_{n,t-} - f_{n,t+})$ on POLICY and FFR and here the coefficients should be close to zero for maturities beyond 30 years to be consistent with the assumption we make in the Campbell-Shiller present-value identity method. As the figure shows, the estimates are supportive of this assumption. Overall, with both methods, changes in the forward rate curve are sufficiently near-term to allow us to fully capture implied price changes with the observable yield and forward rate data.

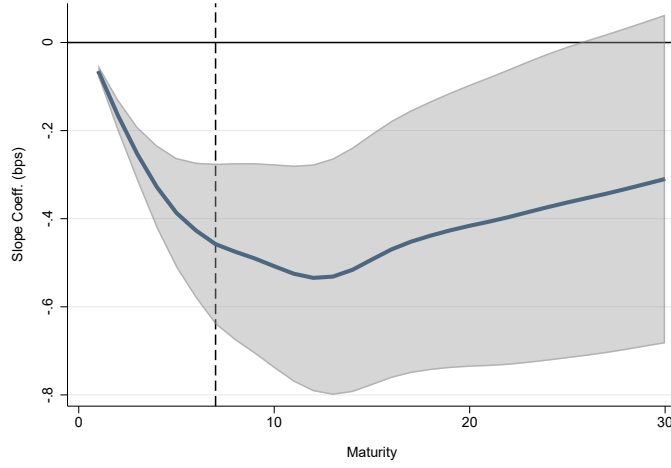
4.4 High-frequency responses

So far, we have used daily yield changes on FOMC announcement days. To further eliminate noise in yield changes that are not related to monetary surprises, we use intraday yield responses in the 30-minute window around FOMC announcements to calculate high-frequency counterfactual price changes. In the following analysis, we also use high-frequency S&P 500 index futures price changes, a high-frequency monetary policy surprise measure (BS), and the residuals from regressing BS on six macro and financial variables ($BS^\perp$) from [Bauer and Swanson \(2023a\)](#).

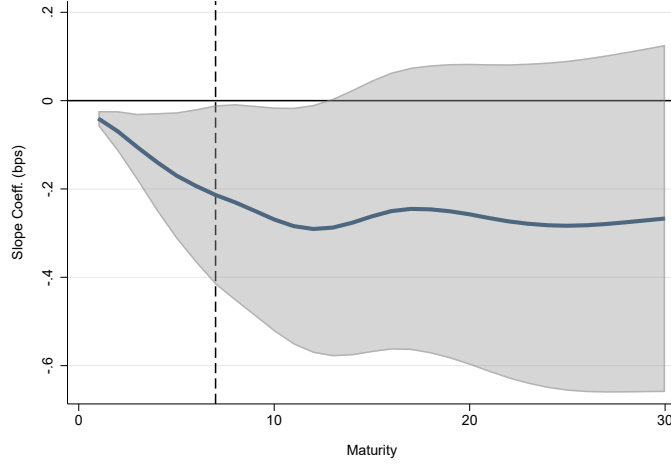
We first construct the high-frequency counterfactual percentage price changes using the dividend futures method. As we do not have data on dividend futures prices 10 minutes before each FOMC announcement, we use the dividend futures prices from the end of the previous day to calculate weights on high-frequency yield responses:

$$\Delta P_{B,t}^H = \sum_{n=1}^m \frac{G_{n,t-} B_{n,t-}}{P_{t-}} \left(\frac{B_{n,t+}^H}{B_{n,t-}^H} - 1 \right) + \left(1 - \sum_{n=1}^m \frac{G_{n,t-} B_{n,t-}}{P_{t-}} \right) \left(\frac{B_{m+1,t+}^H}{B_{m+1,t-}^H} - 1 \right). \quad (17)$$

Since dividend futures prices only serve as weights and are not used to extract asset price changes, the error introduced by using previous day futures prices instead of intraday pre-announcement prices is minor. In contrast, it is crucial that $B_{n,t+}^H/B_{n,t-}^H = \exp(n(y_{n,t-}^H - y_{n,t+}^H))$ can be calculated with high-frequency yield responses $y_{n,t+}^H - y_{n,t-}^H$, which we observe



(A) POLICY

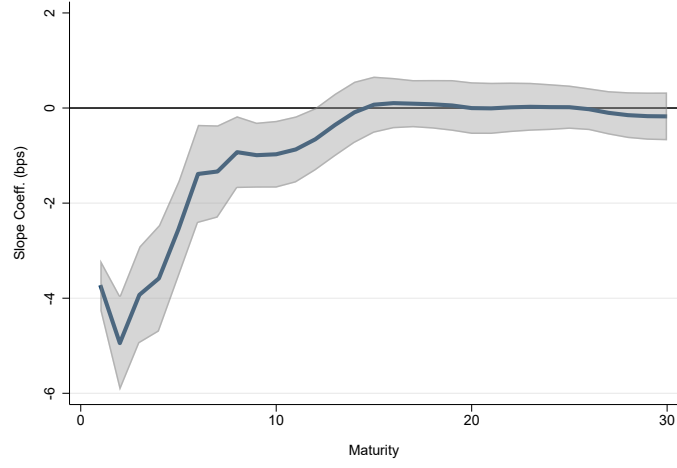


(B) FFR

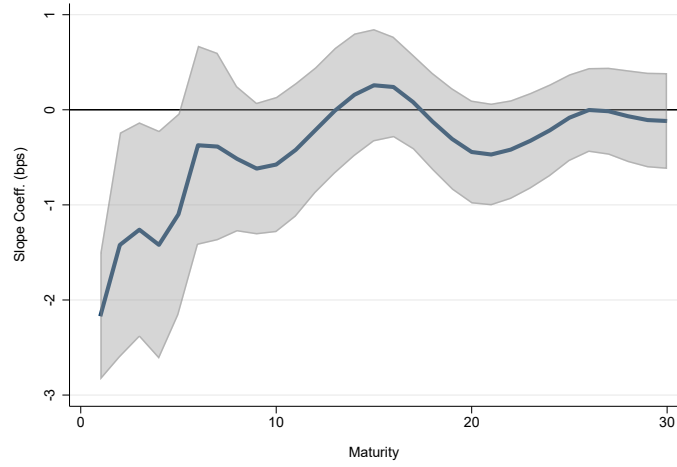
FIGURE 1

Regressing Weighted Zero-Coupon Yield Changes on Monetary Policy Surprises

The blue line plots the slope coefficients from regressing the weighted zero-coupon yield changes $\frac{G_{n,t-}}{P_{t-}}(B_{n,t+} - B_{n,t-})$ on monetary policy surprises for maturities up to 30 years. For $n \leq 7$, we use directly observed dividend futures prices to calculate $G_{n,t-}$; for $n > 7$, we use estimated dividend futures prices under a Gordon growth model following [Knox and Vissing-Jørgensen \(2022\)](#). The monetary policy surprise measures include the policy news shock from [Nakamura and Steinsson \(2018\)](#) as updated by [Acosta \(2023\)](#) (POLICY) and the unexpected changes in the target Federal funds rates (FFR). The black dashed line indicates the 7-year maturity. The gray area represents the 95% confidence interval.



(A) POLICY



(B) FFR

FIGURE 2

Regressing Weighted Forward Rate Changes on Monetary Policy Surprises

The blue line plots the slope coefficients from regressing the weighted forward rate changes $\rho^{n-1}(f_{n,t-} - f_{n,t+})$ on the monetary policy surprises for maturities up to 30 years. The monetary policy surprise measures include the policy news shock from [Nakamura and Steinsson \(2018\)](#) as updated by [Acosta \(2023\)](#) (POLICY) and the unexpected changes in the target Federal funds rates (FFR). The gray area represents the 95% confidence interval.

for 2-, 5-, 10-, and 30-year maturities. We interpolate the responses of other maturities by assuming they follow a step function, i.e., responses of yields with maturities between 2- and 5-year are the same as the 5-year responses. Our choice of using a step function is motivated by the observation from Panel A in Table 2 that longer maturity yields respond much less than shorter maturity yields. Thus, using such a step function for interpolation provides a conservative measure of how much price variation can be captured by looking at yield responses only. Appendix C shows that our results are robust to using linear interpolation methods to estimate responses of unobserved maturities.

We can also construct the log prices implied by forward rate changes using the Campbell-Shiller present-value identity method as

$$\Delta p_{F,t}^H = \sum_{n=1}^{30} \rho^{n-1} [n(y_{n,t-}^H - y_{n,t+}^H) - (n-1)(y_{n-1,t-}^H - y_{n-1,t+}^H)]. \quad (18)$$

Panel C, Table 1 reports the summary statistics for the high-frequency actual and counterfactual S&P 500 price changes around FOMC announcements. In both methods, the counterfactual price changes exhibit slightly higher volatility compared to the actual price changes.

Panel A, Table 4 shows that, regardless of which monetary surprise measure we use, both the actual price changes and counterfactual changes based on high-frequency yield responses using the dividend futures method respond to monetary surprises with similar magnitudes, and their difference is not statistically significant. Panel B, Table 4 shows that a similar conclusion holds for the longer sample period when calculating counterfactual price changes using the Campbell-Shiller present-value identity method.

4.5 Short-rate expectations and term premia

The evidence so far suggests that movements in the default-free yield curve largely explain the stock market response to monetary policy surprises in FOMC announcement windows,

TABLE 4
High-Frequency Stock Market Response to Monetary Policy Surprises

The first four columns report the OLS estimates of regressing the 30-minute-window actual price changes of S&P 500 futures around scheduled FOMC announcements on monetary surprise measures. Actual price changes are percentage changes in Panel A and log changes in Panel B. The monetary surprise measures include the policy shocks from [Nakamura and Steinsson \(2018\)](#) as updated by [Acosta \(2023\)](#) (POLICY), the unexpected changes in the target Federal funds rates (FFR), the unadjusted monetary policy surprises from [Bauer and Swanson \(2023a\)](#) (BS), and residuals from regressing BS on six macro and financial variables ($BS^\perp$). All surprises are standardized to have unit standard deviations. The next four columns use the counterfactual price changes based on the 30-minute-window yield responses as the dependent variable. Counterfactual price changes are calculated using the dividend futures method in Panel A and the Campbell-Shiller present-value identity method in Panel B. Yield responses of unobserved maturities are interpolated with a step function. The last four columns test the difference between slope coefficients. All dependent variables are quoted in basis points. All regressions include a constant term that is not reported. We report the t -statistics calculated using heteroskedasticity-robust standard errors in brackets. The sample period runs from November 2002 to December 2019 in Panel A and February 1994 to December 2019 in Panel B.

Panel A: Dividend Futures Method

	ΔP^H				ΔP_B^H				$\Delta P^H - \Delta P_B^H$			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
POLICY	-27.06				-30.96				3.90			
	[-4.39]				[-3.91]				[0.41]			
FFR		-12.15				-8.43				-3.72		
		[-1.23]				[-1.94]				[-0.38]		
BS			-30.36				-31.23				0.87	
			[-6.20]				[-3.28]				[0.09]	
$BS^\perp$				-32.64				-22.77				-9.87
				[-5.60]				[-3.66]				[-1.20]
N	128	128	128	128	128	128	128	128				
Adj. R^2	0.17	0.03	0.22	0.27	0.10	0.00	0.11	0.06				

Panel B: Campbell-Shiller PV Identity Method

	Δp^H				Δp_F^H				$\Delta p^H - \Delta p_F^H$			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
POLICY	-23.36				-37.24				13.89			
	[-5.10]				[-6.31]				[1.79]			
FFR		-13.46				-7.75				-5.70		
		[-2.42]				[-2.04]				[-0.77]		
BS			-26.47				-36.28				9.81	
			[-6.51]				[-5.85]				[1.28]	
$BS^\perp$				-28.71				-31.56				2.85
				[-6.62]				[-6.00]				[0.38]
N	200	200	208	208	200	200	208	208				
Adj. R^2	0.16	0.06	0.22	0.25	0.25	0.01	0.24	0.18				

leaving little room for changes in the equity premium. This does not mean, however, that risk premia play no role. The default-free yield curve also embodies risk premia in the form of the term premium in long-term yields. Part of the yield change in response to monetary policy surprises may be a change in the term premium. We now look at how much of the stock market price response can be traced to a change in the term premium and how much can be attributed to changes in expectations of future short-term interest rates.

For this purpose, we use Blue Chip survey forecasts of 3-month Treasury bill rates. Since survey data is only available at a monthly frequency, the decomposition into short-rate expectations and term premia is much noisier than the decomposition of daily or high-frequency price changes we looked at earlier. Also, since we do not observe short-rate forecasts for horizons beyond 2 years, we need to make additional assumptions to impute expectations at longer horizons. Specifically, we assume that forecasters perceive short rates, i_t , to follow an AR(1) process:

$$i_{t+1} - \mu = \gamma(i_t - \mu) + \eta_{t+1}, \quad (19)$$

where μ denotes the perceived long-run mean. This implies that the forecasters report

$$\tilde{\mathbb{E}}_t i_{t+n} = \gamma^{n-1}(\tilde{\mathbb{E}}_t i_{t+1} - \mu) + \mu, \quad n \geq 1, \quad (20)$$

as their expectation of n -period-ahead short rates. Thus, the revisions in these expectations

$$\tilde{\mathbb{E}}_{t-} i_{t+n} - \tilde{\mathbb{E}}_{t+} i_{t+n} = \gamma^{n-1}(\tilde{\mathbb{E}}_{t-} i_{t+1} - \tilde{\mathbb{E}}_{t+} i_{t+1}), \quad n \geq 1, \quad (21)$$

can be backed out based on the observed short-horizon expectations revision on the right-hand side of this equation.

In our empirical implementation, we directly use the forecasts of 1-year-ahead 3-month Treasury bill rates to measure $\tilde{\mathbb{E}}_t i_{t+1}$. Given their monthly frequency, we use the latest survey forecast before the FOMC meeting date as $\tilde{\mathbb{E}}_{t-} i_{t+1}$ and the first available survey forecast after

the meeting as $\tilde{\mathbb{E}}_{t+1} i_{t+1}$. To estimate γ , we rely on the bi-annual long-range forecasts from Blue Chip as described in Section 3.3. We allow the AR(1) coefficient γ and the perceived long-run mean μ to have low-frequency variation at a bi-annual frequency. In Appendix D, we provide more details on how we estimate γ and μ using bi-annual long-range survey forecasts. We obtain γ estimates that are between 0.3 and 0.7 most of the time.

Equipped with short-rate forecasts, we can now decompose the price changes implied by yield curve changes. Let rx_{t+1}^j be the one-period excess return on a j -maturity zero-coupon bond realized at $t + 1$. We can decompose the forward rate $f_{n,t}$ into an expected short rate and a risk premium component:

$$f_{n,t} = \tilde{\mathbb{E}}_t i_{t+n-1} + \underbrace{\mathbb{E}_t \sum_{j=1}^n rx_{t+n-j+1}^j}_{n\lambda_{n,t}} - \underbrace{\mathbb{E}_t \sum_{j=1}^{n-1} rx_{t+n-j}^j}_{(n-1)\lambda_{n-1,t}}. \quad (22)$$

Here $n\lambda_{n,t}$ is the term premium earned by an n -maturity zero-coupon bond and $\theta_{n,t} = n\lambda_{n,t} - (n-1)\lambda_{n-1,t}$ is the forward term premium.

Consider now the dividend futures method. We have

$$\begin{aligned} \Delta P_{B,t} &= \sum_{n=1}^{\infty} \frac{G_{n,t-}}{P_{t-}} (B_{n,t+} - B_{n,t-}) \\ &= \sum_{n=1}^{\infty} \frac{P_{n,t-}}{P_{t-}} \left(e^{n(y_{n,t-} - y_{n,t+})} - 1 \right) \\ &\approx \underbrace{\sum_{n=1}^{\infty} \frac{P_{n,t-}}{P_{t-}} \left(\tilde{\mathbb{E}}_{t-} \sum_{k=0}^{n-1} i_{t+k} - \tilde{\mathbb{E}}_{t+} \sum_{k=0}^{n-1} i_{t+k} \right)}_{\Delta P_{B,s,t}} + \sum_{n=1}^{\infty} \frac{n P_{n,t-}}{P_{t-}} (\lambda_{n,t-} - \lambda_{n,t+}). \end{aligned} \quad (23)$$

This means we have now decomposed the implied price change into two parts. The first part, reflecting changing expectations of short-term interest rates, can be estimated with survey

data. Based on (21), we can rewrite this first part as

$$\Delta P_{B,s,t} = i_{t-} - i_{t+} + \frac{(\tilde{\mathbb{E}}_{t-} i_{t+1} - \tilde{\mathbb{E}}_{t+} i_{t+1})}{1 - \gamma} \left(1 - \sum_{n=1}^{\infty} \frac{\gamma^{n-1} P_{n,t-}}{P_{t-}} \right). \quad (24)$$

We use all available dividend futures, i.e., up to maturity 7 years, to calculate P_n in the summation in the above expression. With γ between 0.3 and 0.7 in most periods, γ^{n-1} is effectively zero for the terms with maturities beyond 7 years, so this data availability restriction is not material. The second part, reflecting changes in term premia, is the residual:

$$\Delta P_{B,\lambda,t} \equiv \Delta P_{B,t} - \Delta P_{B,s,t}. \quad (25)$$

For the Campbell-Shiller present-value identity method, we can write (15) as

$$\Delta p_{F,t} = \underbrace{\sum_{n=1}^{\infty} \rho^{n-1} (\mathbb{E}_{t-} i_{t+n-1} - \mathbb{E}_{t+} i_{t+n-1})}_{\Delta p_{F,s,t}} + \sum_{n=1}^{\infty} \rho^{n-1} (\theta_{n,t-} - \theta_{n,t+}) \quad (26)$$

The first sum represents the implied price changes induced only by changes in short-rate expectations:

$$\Delta p_{F,s,t} \equiv \sum_{n=0}^{\infty} \rho^n (\mathbb{E}_{t-} i_{t+n} - \mathbb{E}_{t+} i_{t+n}), \quad (27)$$

which, using (21), we can rewrite as

$$\Delta p_{F,s,t} = i_{t-} - i_{t+} + \frac{\rho}{1 - \rho\gamma} (\mathbb{E}_{t-} i_{t+1} - \mathbb{E}_{t+} i_{t+1}). \quad (28)$$

We then obtain the implied price changes driven by forward risk premium changes as a residual:

$$\Delta p_{F,\lambda,t} \equiv \Delta p_{F,t} - \Delta p_{F,s,t}. \quad (29)$$

We then repeat our earlier exercise of regressing these implied price changes on monetary policy surprise measures. Table 5 reports the results using short-rate forecasts from both

BCEI and BCFF. While both surveys forecast the 3-month T-bill rate, they are conducted at different times within the month. This timing difference could potentially lead to different results in this analysis, as we cannot measure forecast changes in the tight windows around FOMC announcements.

That said, the results in Table 5 are very similar between BCEI and BCFF. Moreover, in both cases, the results based on the dividend futures method and the Campbell-Shiller present-value identity method are similar, too. For the POLICY surprise measure, we attribute about half of the stock market response implied by yield changes to changes in short-term interest rate expectations and about half to changes in term premia. In contrast, for the FFR surprise measure, changes in term premia drive virtually the entire response. The bottom line is that to understand the reaction of the stock market to monetary policy surprises, changes in risk premia are still important, but changes in the term premium are key, not the equity premium.

4.6 Predictable components of measured monetary surprises

Karnaukh and Vokata (2022) find that monetary policy surprises constructed using bond yields are predictable with the pre-FOMC Blue Chip professionals' revisions in GDP growth forecasts. Cieslak (2018) shows that monetary policy surprises are predictable ex post by Federal funds rate level and employment growth. Bauer and Swanson (2023b) show that public economic news predicts monetary surprise measures. As they point out, such predictability can arise, for example, because investors are learning about the monetary policy rule parameters from observable data, while econometricians running predictability regressions ex post have a hindsight knowledge advantage by using data that was not available to investors when they priced assets prior to FOMC meetings.

Whether the measured monetary policy surprises are predictable in ex post regressions or not does not affect our results above on the role of yield curve movements in explaining the stock market response to these surprises. But it would nevertheless be interesting to

TABLE 5

Decomposing Implied Price Changes Response: Short-Rate Expectations and Term Premia

The first two columns report the results from regressing counterfactual price changes implied by changes in short rate expectations, $\Delta P_{B,s}$ in Panel A and $\Delta p_{F,s}$ in Panel B, on monetary surprise measures. Columns (3) and (4) use the counterfactual price changes implied by changes in the term premium, $\Delta P_{B,\lambda}$ in Panel A and $\Delta p_{F,\lambda}$ in Panel B, as the dependent variable. The first four columns use BCEI forecasts to measure short-rate expectations and the last four columns use BCFF forecasts. The monetary surprise measures include the policy shocks from Nakamura and Steinsson (2018) as updated by Acosta (2023) (POLICY) and the unexpected changes in the target Federal funds rates (FFR). Both surprises are standardized to have unit standard deviations. All dependent variables are quoted in basis points. We report the t -statistics calculated using heteroskedasticity-robust standard errors in brackets. The sample period runs from November 2002 to July 2023 in Panel A and from February 1995 to July 2023 in Panel B.

	BCEI				BCFF			
	Short-Rate		Term Premia		Short-Rate		Term Premia	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
A. Dividend Futures Method								
POLICY	-12.98		-10.65		-17.60		-6.03	
	[-2.82]		[-1.26]		[-3.76]		[-0.72]	
FFR		-1.00		-9.61		-4.45		-6.16
		[-0.17]		[-0.96]		[-0.68]		[-0.59]
Constant	5.58	4.93	1.46	1.23	11.21	10.44	-4.17	-4.28
	[2.05]	[1.68]	[0.27]	[0.23]	[4.01]	[3.38]	[-0.76]	[-0.78]
N	155	155	155	155	155	155	155	155
Adj. R^2	0.10	-0.01	0.01	0.01	0.18	0.00	0.00	0.00
B. Campbell-Shiller PV Identity Method								
POLICY	-10.49		-13.62		-13.13		-10.99	
	[-3.37]		[-1.17]		[-3.90]		[-0.97]	
FFR		1.73		-13.73		-2.22		-9.77
		[0.41]		[-1.17]		[-0.55]		[-0.86]
Constant	5.29	5.24	3.87	3.87	9.18	9.13	-0.02	-0.02
	[2.10]	[2.01]	[0.49]	[0.49]	[4.15]	[3.85]	[-0.00]	[-0.00]
N	226	226	226	226	226	226	226	226
Adj. R^2	0.07	-0.00	0.01	0.01	0.13	-0.00	0.00	0.00

see whether the close connection between stock price changes and yield curve changes are concentrated in predictable or the unpredictable component of monetary policy surprises.

Thus, we follow [Karnaikh and Vokata \(2022\)](#) and [Bauer and Swanson \(2023b\)](#) to orthogonalize the monetary surprises with respect to information available before FOMC announcements by regressing the monetary surprise measures $\psi \in \{\text{POLICY}, \text{FFR}\}$ on Blue Chip forecast revisions and public news:

$$\psi_t = \alpha + \beta \text{news}_{t-1} + \xi_t. \quad (30)$$

Specifically, we include the forecast revisions for real GDP and CPI, defined as the average revisions of the forecasts for the current and future 3 quarters; 3-month changes in log prices of the S&P 500 index (from 65 trading days before the FOMC announcement to the day before the FOMC announcement); 3-month changes in the yield curve slope defined as the spread between 10-year and 3-month Treasury yields; 3-month changes in the Bloomberg Commodity Index (BCOM). Consistent with these earlier papers, we find that economic news contains information of future POLICY surprises. But FFR surprises are much less predictable. The results are shown in [Appendix E](#). Based on these results, we focus on the decomposition of POLICY surprises in the following analysis.

As the predictable and unpredictable components are constructed in the first-step regression in [\(30\)](#), we conduct inference based on a bootstrap approach that accounts for this generated regressor problem. Specifically, we create bootstrap samples by randomly drawing with replacement clusters of price changes, standardized monetary surprises, survey forecast revisions, and financial news on the same FOMC announcement day from the original data. To preserve the autocorrelation structure of these variables, we also use a block bootstrap with block length determined as in [Politis and White \(2004\)](#). In each bootstrap sample, we then re-run the regression [\(30\)](#) to construct the predictable and unpredictable components of monetary surprises, and regress price changes on them. We obtain p -values by comparing the sample regression t -statistic to the quantiles of the bootstrapped distribution of t -statistics.

TABLE 6
Stock Market Response to Predictable and Unpredictable Monetary Policy Surprises

The first two columns report results from regressing price changes of the S&P 500 market index on predictable (Predicted) and unpredictable (Residual) POLICY surprises on scheduled FOMC announcement days. Both explanatory variables are standardized to have unit standard deviations. Columns (3) and (4) use ΔP_B as the dependent variable in Panel A and Δp_F in Panel B. Columns (5) and (6) in Panel A use ΔP_G as the dependent variable. The rest of columns test the difference between slope coefficients. All dependent variables are quoted in basis points. We report in parentheses the p -values based on the distribution of t -statistics from block bootstrapped samples of price changes, standardized monetary surprises, survey forecast revisions, and financial news. The sample period runs from November 2002 to September 2023 in Panel A and from February 1995 to September 2023 in Panel B.

A. Dividend Futures Method										
	ΔP		ΔP_B		ΔP_G		$\Delta P - \Delta P_B$		$\Delta P - \Delta P_G$	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Predicted	-36.13 (0.07)		-10.55 (0.12)		1.23 (0.46)		-25.58 (0.10)		-37.36 (0.06)	
Residual		-21.17 (0.06)		-19.91 (0.00)		0.07 (0.92)		-1.26 (0.90)		-21.24 (0.06)
Constant	33.24 (0.01)	30.82 (0.02)	6.44 (0.12)	6.23 (0.11)	-0.90 (0.36)	-0.79 (0.38)				
N	156	156	156	156	156	156				
Adj. R^2	0.09	0.02	0.03	0.09	0.02	-0.01				

B. Campbell-Shiller PV Identity Method										
	Δp		Δp_F		$\Delta p - \Delta p_F$					
	(1)	(2)	(3)	(4)	(5)	(6)				
Predicted	-22.63 (0.18)		-13.45 (0.27)			-9.19 (0.52)				
Residual		-15.72 (0.04)		-20.91 (0.04)			5.20 (0.66)			
Constant	27.21 (0.00)	26.55 (0.00)	9.84 (0.16)	9.44 (0.14)						
N	228	228	228	228						
Adj. R^2	0.03	0.01	0.01	0.03						

Table 6 shows that the stock market responds to both components with similar economic magnitudes. However, the coefficient on the predictable component is estimated with much more statistical uncertainty, with the consequence that we cannot reject at conventional levels that the coefficient is equal to zero. The results also show that the magnitudes of the responses to the unpredictable component of monetary surprises are very similar between the actual price changes and the counterfactual price changes implied by yield changes. For the predictable component, the point estimates for actual price changes and counterfactual price changes differ to some extent, but the difference is far from being statistically significant at conventional levels.

4.7 Large monetary policy surprises

Similar to BK and many other studies of asset price reactions to FOMC announcements, we focus on asset price changes projected on monetary policy surprise measures. The fact that the R^2 in regressions of announcement-window price changes on monetary policy surprise measures is not very high indicates that other news orthogonal to these monetary surprises reaches the market in these windows. For example, even in the high-frequency analysis in Table 4, the R^2 is only up to 27% in 30-minute windows depending on the monetary surprise measure used. Since the news about Federal Reserve policy is arguably the dominating news reaching the market in these short time windows, this suggests that FOMC announcements must convey other news that is relevant for stock prices, but at the same time not captured by standard monetary policy surprise measures. For example, even if an announcement does not convey any news about the path of short-term interest rates (which the monetary policy surprise measures capture), it might still convey news about policy uncertainty (which the monetary policy surprise measures do not capture) that affects risk premia.

Such other news may be more important in some announcement windows than in others. There are some announcement days on which there is just not much news about the path of interest rates because market participants have already anticipated the FOMC's decisions

based on FOMC members’ speeches and other communications ahead of the meeting. In such FOMC announcement windows, stock price movements may mostly reflect other types of news. While a thorough investigation of these other news components orthogonal to conventional monetary policy surprise measures is beyond the scope of this paper, it would be useful to check whether our results hold up if we focus on FOMC announcements in which standard measures of monetary policy surprises have large realizations, indicating that the news captured by these surprises is likely dominant in the announcement window.

Table 7 extends the analysis from Table 4 by using only announcement windows in which the absolute value of the monetary surprise realization exceeds one full-sample standard deviation. The results show that the R^2 from regressing actual or counterfactual price changes on monetary policy surprise measures is about twice as big as in Table 4 in most specifications. Importantly, for this subset of announcement windows with large monetary policy surprises, the projections of counterfactual price changes in columns (5) to (8) remain very similar to those of actual price changes in columns (1) to (4). However, here the projections account for a significantly bigger share of total actual and counterfactual price changes compared to Table 4. These results are consistent with the idea that in some announcement windows, conventional monetary policy surprise measures do not capture much relevant news that moves asset prices, while in others, when these measures are substantial, they capture a much bigger share.

4.8 UK evidence

In recent years, there has been a growing effort to collect data on the high-frequency responses of asset prices to monetary policy announcements in economies other than the US. We utilize the newly published UK Monetary Policy Event-Study Database (Braun et al. 2023) to study whether the UK stock market response in monetary policy announcement windows is also explained by yield curve movements.

The UK Monetary Policy Event-Study Database provides data on responses of Libor rates,

TABLE 7
High-Frequency Stock Market Response to Large Monetary Policy Surprises

The first four columns report the OLS estimates of regressing the 30-minute-window actual price changes of S&P 500 futures around scheduled FOMC announcements on large monetary surprise measures, defined as those with absolute values higher than full-sample standard deviations. Actual price changes are percentage changes in Panel A and log changes in Panel B. The monetary surprise measures include the policy shocks from Nakamura and Steinsson (2018) as updated by Acosta (2023) (POLICY), the unexpected changes in the target Federal funds rates (FFR), the unadjusted monetary policy surprises from Bauer and Swanson (2023a) (BS), and residuals from regressing BS on six macro and financial variables ($BS^\perp$). All surprises are standardized to have unit standard deviations. The next four columns use the counterfactual price changes based on the 30-minute-window yield responses as the dependent variable. Counterfactual price changes are calculated using the dividend futures method in Panel A and the Campbell-Shiller present-value identity method in Panel B. Yield responses of unobserved maturities are interpolated with a step function. The last four columns test the difference between slope coefficients. All dependent variables are quoted in basis points. All regressions include a constant term that is not reported. We report the t -statistics calculated using heteroskedasticity-robust standard errors in brackets. The sample period runs from November 2002 to December 2019 in Panel A and February 1994 to December 2019 in Panel B.

Panel A: Dividend Futures Method												
	ΔP^H				ΔP_B^H				$\Delta P^H - \Delta P_B^H$			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
POLICY	-26.63 [-4.13]				-22.14 [-5.24]				-4.49 [-0.64]			
FFR		-10.02 [-0.89]				-6.89 [-1.55]				-3.13 [-0.28]		
BS			-33.46 [-6.90]				-29.15 [-2.80]				-4.31 [-0.40]	
BS [⊥]				-29.26 [-4.26]				-17.33 [-3.87]				-11.93 [-1.45]
N	21	13	27	32	21	13	27	32				
Adj. <i>R</i> ²	0.27	0.01	0.49	0.38	0.69	0.13	0.06	0.40				

Panel B: Campbell-Shiller PV Identity Method												
	Δp^H				Δp_F^H				$\Delta p^H - \Delta p_F^H$			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
POLICY	-23.96 [-5.26]				-34.75 [-5.25]				10.80 [1.28]			
FFR		-12.55 [-2.07]				-7.02 [-1.81]				-5.54 [-0.71]		
BS			-27.76 [-6.43]				-35.11 [-4.90]				7.35 [0.84]	
BS [⊥]				-27.38 [-5.64]				-31.36 [-5.49]				3.97 [0.47]
N	44	34	48	57	44	34	48	57				
Adj. <i>R</i> ²	0.28	0.15	0.41	0.39	0.51	0.06	0.27	0.39				

Treasury (Gilt) yields of maturities up to 10 years, and the FTSE All Share Index in the UK Monetary Policy Committee’s announcement windows.⁹ Similar to [Nakamura and Steinsson \(2018\)](#) and [Bauer and Swanson \(2023a\)](#), we measure monetary policy surprises as the first principal component of responses in the first four quarterly Short Sterling Futures contracts that are based on 3-month Libor rates (SONIA rates since May 2021). In the sample period from 1997 to 2023, the first principal component explains 91% of the variance in the four contracts’ response. As Panel A in Table 8 shows, consistent with the US evidence, the Gilt yield curve responds to monetary policy surprises with coefficients that decline with maturity. Based on this finding, we construct a measure of yield-curve-implied price changes using a step-function interpolation as in Section 4.4. We choose $\rho = 0.966$ based on the average log dividend-price ratio of the FTSE All Share Index between 1997 and 2023, which is -3.36. We use the Campbell-Shiller approach using high-frequency yield responses to calculate the counterfactual price changes.

Column (1) in Panel B of Table 8 shows that a 1 s.d. negative monetary policy surprise is associated with a 14 bp increase in the FTSE All Share Index that is statistically significant at conventional levels. As column (2) shows, this response is fully captured by our measure of counterfactual price changes implied by yield curve movements.

5 Price changes over the FOMC cycle

Financial market participants experience monetary policy surprises not only on FOMC announcement days. The stock market may also respond to news about monetary policy that comes out between FOMC meeting days. In this regard, [Cieslak et al. \(2019\)](#) document a striking pattern. Using data from 1994 to 2016, they find that average stock index returns are much higher in even weeks than odd weeks in FOMC cycle time. To better understand this puzzling regularity in stock price changes, it is useful to check whether this regular-

⁹ There is also a quarterly Monetary Policy Report publication followed by a press conference. Prior to August 2015, the report publication and corresponding press conference typically occurred a week after the policy announcement. For this reason, we focus on the asset price responses in announcement windows only.

TABLE 8
High-Frequency Asset Price Response to Monetary Policy Surprises in the UK

Panel A reports the results from regressing the 30-minute-window changes in UK Gilt yields around the Monetary Policy Committee's announcements on a measure of monetary policy surprise (MPS). MPS is calculated as the first principal component of 30-minute-window changes in the first four quarterly Short Sterling Futures prices, standardized to unit standard deviations. In Panel B, the first column uses the 30-minute-window log returns on the FTSE All Share Index as the dependent variable. The second column uses the counterfactual price changes based on the 30-minute-window Gilt yield changes of maturities between 1 to 10 years as the dependent variable, calculated based on the Campbell-Shiller present-value identity method. Yield responses on unobserved maturities are interpolated with a step function. The last column tests the difference between slope coefficients. All dependent variables are quoted in basis points. We report the t -statistics calculated using heteroskedasticity-robust standard errors in brackets. The sample period runs from June 1997 to June 2023.

Panel A: Gilt Yield				
	1Y (1)	2Y (2)	5Y (3)	10Y (4)
MPS	3.90 [13.95]	3.51 [10.84]	2.68 [9.99]	1.72 [6.06]
Constant	-0.20 [-1.51]	-0.27 [-2.14]	-0.18 [-1.19]	-0.05 [-0.32]
N	284	284	284	284
Adj. R^2	0.76	0.73	0.52	0.27
Panel B: FTSE All Share Index				
	Δp (1)	Δp_F (2)	$\Delta p - \Delta p_F$ (3)	
MPS	-13.98 [-7.02]	-15.32 [-6.33]	1.34 [0.38]	
Constant	0.84 [0.55]	0.56 [0.40]	0.28 [0.15]	
N	283	283		
Adj. R^2	0.23	0.29		

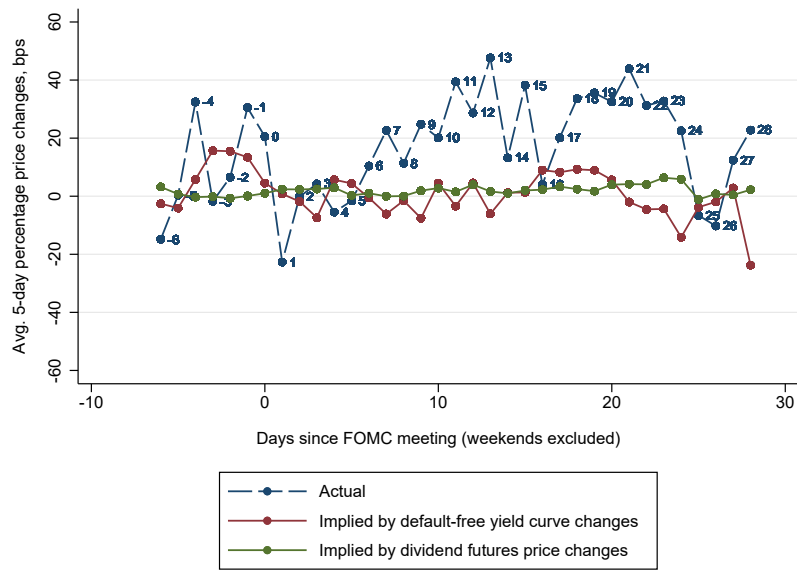
ity can be attributed to the effect of yield changes in the bond market or whether it is a stock-market specific phenomenon that reflects FOMC cycle time seasonality in the equity premium. Define day 0 as the day of a scheduled FOMC announcement and day t as the number of weekdays since the FOMC announcement. Since weekends are excluded, week 0 in FOMC cycle time is defined as days -1 to 3, week 1 is defined as days 4 to 8, and so on so forth. We study weeks from -1 (days -6 to -2) to 5 (days 24 to 28) as the total number of days within week 6 is much smaller.¹⁰ We then calculate the cumulative 5-day price changes, without considering dividends or risk-free rate, from t to $t + 4$ assuming the changes are zero over the weekends.

Panel A in Figure 3 shows that the cyclical pattern documented in Cieslak et al. (2019) also holds in our shorter sample with dividend futures data available from 2002 to 2023, but the pattern is not as pronounced as in their original study. Average price changes are higher in even weeks (days -1 to 3, 9-13, 19-23) than in odd weeks. As the figure also shows, the price changes ΔP_B implied by changes in the zero-coupon yield curve, calculated using our dividend futures method, are positively correlated with the actual price changes. So at least a portion of the FOMC cycle pattern in stock prices can be attributed to changes in the default-free yield curve. Panel A also shows price changes implied by dividend futures price changes are basically flat and not connected to the cycle pattern.

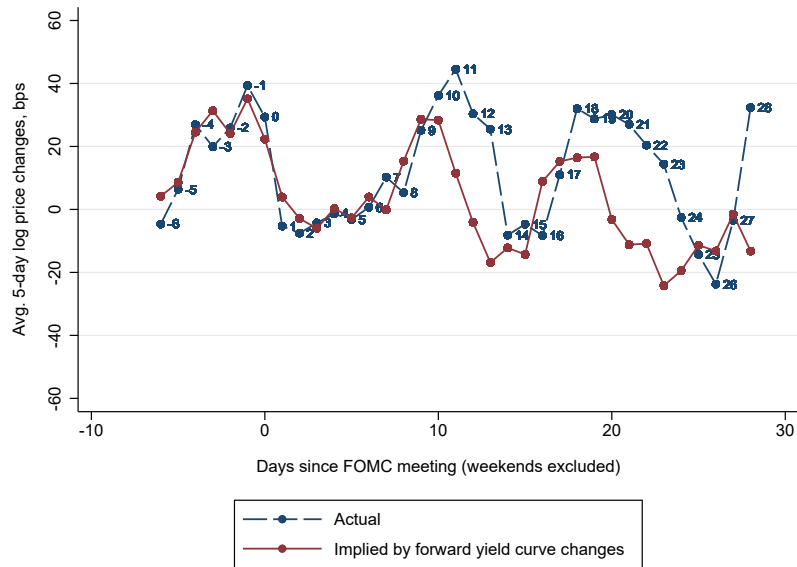
Panel B looks at the longer sample between 1994 and 2023, using log price changes and the Campbell-Shiller present-value identity method to calculate price changes implied by forward rate curve changes. In this longer sample, the pattern in actual price changes is more pronounced than in Panel A. The implied log price changes Δp_F in this panel also line up very well with the actual log price changes, leaving little room for other explanatory factors.

We perform a formal test to assess the statistical significance of the cyclical behavior by regressing daily price changes, either actual or implied, on dummies of even weeks in the

¹⁰ In our sample period, week 6 in FOMC cycle time only includes 140 days, while all other weeks contain approximately 500 to 1200 days. Including week 6 does not quantitatively change our results.



(A) Dividend futures method



(B) Campbell-Shiller present-value identity method

FIGURE 3
Price changes over the FOMC cycle

The numbers along the line indicate number of days since the FOMC meeting. The blue lines in each panel plot the average actual 5-day price changes from t to $t + 4$. In Panel A, the red line shows the average 5-day percentage price changes implied by zero-coupon yield curve changes based on the dividend futures method. The green line shows the price changes implied by changes in dividend futures prices only. The sample period runs from November 2002 and December 2023. In Panel B, the red line shows the average 5-day log price changes implied by forward rate changes based on the Campbell-Shiller present-value identity method. The sample period in Panel B runs from February 1995 and December 2023.

TABLE 9
Regressing Price Changes on FOMC Cycle Dummies

This table reports the coefficients from regressing actual price changes and counterfactual price changes implied by yield curve changes on FOMC cycle dummies. In Panel A, the independent variables are a dummy that equals to 1 in week 0 (days -1 to 3), 2 (days 9 to 13), and 4 (days 19 to 23) in FOMC cycle time and a constant. In Panel B, we include separate dummies for week 0, 2, and 4 in FOMC cycle time and a constant as the independent variables. All dependent variables are quoted in basis points. We report the t -statistics calculated using heteroskedasticity-robust standard errors in brackets. The sample period runs from October 2002 to December 2023 for the dividend futures method and from January 1994 to December 2023 for the Campbell-Shiller present-value identity method, respectively. Days belonging to week 6 in FOMC cycle time are excluded from the regressions.

	Dividend Futures					Campbell-Shiller PV		
	ΔP (1)	ΔP_B (2)	ΔP_G (3)	$\Delta P - \Delta P_B$ (4)	$\Delta P - \Delta P_G$ (5)	Δp (6)	Δp_F (7)	$\Delta p - \Delta p_F$ (8)
A. Even weeks								
Week 0, 2, 4	5.26 [1.48]	1.39 [1.06]	-0.34 [-1.21]	3.87 [0.93]	5.60 [1.60]	8.16 [2.94]	5.79 [2.34]	2.37 [0.57]
Constant	1.44 [0.60]	-0.33 [-0.35]	0.52 [2.79]	1.73 [0.62]	0.91 [0.39]	-0.87 [-0.47]	-1.85 [-1.12]	0.98 [0.36]
N	4700	4700	4700	4700	4700	7348	7348	7348
Adj. R^2	0.00	0.00	0.00	-0.00	0.00	0.00	0.00	-0.00
B. Week by week								
Week 0	4.06 [0.75]	3.02 [1.50]	-0.43 [-1.22]	1.04 [0.17]	4.49 [0.83]	9.30 [2.25]	8.97 [2.46]	0.33 [0.05]
Week 2	6.68 [1.28]	-0.41 [-0.21]	-0.34 [-0.66]	7.09 [1.13]	7.02 [1.37]	6.58 [1.63]	2.46 [0.68]	4.11 [0.66]
Week 4	5.22 [1.13]	1.33 [0.73]	-0.23 [-0.66]	3.89 [0.72]	5.45 [1.20]	8.52 [2.23]	5.61 [1.60]	2.91 [0.50]
Constant	1.44 [0.60]	-0.30 [-0.35]	0.52 [2.79]	1.73 [0.62]	0.91 [0.39]	-0.87 [-0.47]	-1.85 [-1.12]	0.98 [0.36]
N	4700	4700	4700	4700	4700	7348	7348	7348
Adj. R^2	-0.00	0.00	-0.00	-0.00	-0.00	0.00	0.00	-0.00

FOMC cycle.

In the shorter sample period since 2002, consistent with Figure 3, the point estimates in panel A, column (1) in Table 9 indicate that actual price changes are cyclical over the FOMC cycle. Average daily returns are 5.26 basis points higher in even weeks than in odd weeks. However, due to the small number of observations, the point estimate is not statistically significant at conventional levels. Column (2) shows that the price change implied by yield curve changes, ΔP_B , explains about a third of the even-week effect. In contrast, the estimated difference in ΔP_G between even and odd weeks is very close to zero.

Columns (6) to (8) look at the longer sample, using the Campbell-Shiller present-value identity method to calculate log price changes implied by changes in the forward rate curve. Column (6) shows that log price changes are 8.16 bp higher in even weeks and the estimate is statistically significant at the 5% level. Using the implied price change Δp_F , we obtain a coefficient estimate of 5.79 which is also statistically significant. At 2.37 bp, the difference between actual and implied price changes in column (8) is quite small and not statistically significant. Thus, most of the cyclical pattern in the FOMC cycle can be traced to cyclical movements in the stock market due to changes in the default-free yield curve.

This conclusion contrasts with the conclusion in Cieslak et al. (2019) that changes in the equity premium account for most of the FOMC cycle effect in stock returns. Their conclusion is based on the equity premium bound of Martin (2017), which is a rather indirect way of measuring the equity premium that does not account for possibly complex changes in the yield curve when monetary policy news reaches financial markets. Our results show that yield curve movements play a much bigger role than it originally seemed.

Our interpretation also differs from Cieslak and Pang (2021) who use a structural VAR with sign restrictions to decompose the FOMC cycle effect and attribute much of the FOMC cycle effect in stock returns to risk premium changes. However, there is partial overlap in that their definitions of risk premium shocks includes shocks to the term premium which, in our model-free framework, is part of the yield curve changes that enter into our calculation

of implied price changes.

6 Conclusion

Our findings in this paper overturn the conventional wisdom that stock market response to monetary policy surprises is mostly attributable to the effect of monetary policy on the equity premium. Using a model-free method based on dividend futures data, we find instead that the response is almost entirely explained by valuation effects due to the changes in the default-free yield curve, not the equity premium.

The key advantage of our method is that it does not require the assumption that stock market index returns in FOMC announcement windows obey the same VAR dynamics for returns, the dividend-price ratio, and interest rates as on any other day of the year. Using dividend futures prices as weights, the method delivers the change in the stock market index value implied by yield curve changes without virtually any functional form assumptions on the yield curve or the dynamics of stock returns. An alternative method that uses the Campbell-Shiller present-value identity weights instead of dividend futures prices produces very similar results.

The bottom line is that the effects of monetary policy on the stock market are of a more conventional nature than it seemed in the earlier literature. At least for the monetary policy surprises in FOMC announcement windows that we studied in this paper, a conventional yield curve channel is sufficient to explain the stock market reaction. This result does not necessarily extend to monetary policy surprises outside of FOMC announcement windows such as, for example, the market reaction to speeches or other forms of inter-meeting communication, but the methods developed in this paper would also be suitable for separating equity premium and yield curve effects of other monetary policy news events.

In our analysis, as in BK and in much of the subsequent literature on asset price reactions to monetary policy, we examine the projection of stock price changes on monetary policy surprise measures constructed from unexpected changes in near-term interest rates. Our result

that yield curve changes are sufficient to explain the magnitude of the stock price response focuses on this component of stock price changes. Yet there is also clearly substantial movement in stock prices within announcement windows that is orthogonal to these conventional monetary policy surprise measures. There is room left for monetary policy effects on the equity premium in this orthogonal component, but explaining these equity premium changes would require different theories than those that generate equity premium shocks correlated with surprises about near-term interest rates. This is an interesting area for future research.

Appendix

A Additional results on calculating measures of implied price changes

A.1 Testing assumptions on zero-coupon bond price changes

In this section we test the empirical validity of (6). Given that the maximum maturity we can observe for dividend futures prices is 7 years, we choose $m = 7$. We then calculate the difference in nominal zero-coupon bond price changes $\frac{B_{n,t+}}{B_{n,t-}} - \frac{B_{m+1,t+}}{B_{m+1,t-}}$ for $n = 9, 15, 20$, and 30. We test: (i) whether the difference is zero; (ii) whether the difference co-moves with monetary surprises.

The first block in Table A.1 shows that the mean of $\frac{B_{n,t+}}{B_{n,t-}} - \frac{B_{m+1,t+}}{B_{m+1,t-}}$ is, statistically, not significantly different from zero for all maturities considered. The second and the third blocks in Table A.1 show that the relationship between these differences and monetary shocks are not significantly different from zero. Overall, these results support our assumption in (6).

A.2 Alternative measures of ΔP_G

The first alternative measure of ΔP_G is based on estimating $G_{n,t}$ for longer maturities following the methodology in Knox and Vissing-Jørgensen (2022). Knox and Vissing-Jørgensen (2022) show that under the assumption of a Gordon growth model, we have

$$\frac{P_{n,t}}{P_t} \approx \frac{P_{m,t}}{P_t} \left(\frac{P_t - \sum_{k=1}^m P_{k,t}}{P_t - \sum_{k=1}^{m-1} P_{k,t}} \right)^{n-m}, \quad n > m, \quad (\text{A.1})$$

where m is the maximum maturity of dividend futures observed. Using the additional relationship that $P_{n,t} = G_{n,t}e^{-ny_{n,t}}$, we can estimate

$$G_{n,t} \approx e^{ny_{n,t} - my_{m,t}} G_{m,t} \left(\frac{P_t - \sum_{k=1}^m G_{k,t}e^{-ky_{k,t}}}{P_t - \sum_{k=1}^{m-1} G_{k,t}e^{-ky_{k,t}}} \right)^{n-m}, \quad n > m. \quad (\text{A.2})$$

TABLE A.1
Long-Maturity Zero-Coupon Bond Price Changes on FOMC Announcement Days

This table reports the means of difference in zero-coupon bond price changes and their slope coefficients from regressing on monetary surprises:

$$\frac{B_{n,t+}}{B_{n,t-}} - \frac{B_{8,t+}}{B_{8,t-}} = \alpha + \beta\psi_t + \eta_t,$$

for $n=9, 15, 20$, and 30 . The monetary policy surprise measures include the policy news shock from [Nakamura and Steinsson \(2018\)](#) as updated by [Acosta \(2023\)](#) (POLICY) and the unexpected changes in the target Federal funds rates (FFR). Both shocks are standardized to have unit standard deviations. We report the t -statistics calculated using heteroskedasticity-robust standard errors in brackets. All dependent variables are quoted in basis points. The sample period runs from February 1995 to December 2023.

	$n = 9$			$n = 15$			$n = 20$			$n = 30$		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
Mean	0.80 [1.76]			2.81 [0.92]			4.55 [0.94]			10.34 [1.13]		
POLICY		-1.31 [-2.07]			-5.42 [-1.49]			-4.94 [-0.78]			-6.42 [-0.51]	
FFR			-0.81 [-1.31]			-1.89 [-0.52]			-3.00 [-0.49]			-6.92 [-0.57]
N	230	230	230	230	230	230	230	230	230	230	230	230
Adj. R^2	0.01	0.03	0.01	-0.00	0.01	-0.00	-0.00	0.00	-0.00	0.00	-0.00	-0.00

We can then calculate (9) using the expanded set of maturities as

$$\Delta P_{G,t}^1 = \sum_{n=1}^{30} \frac{B_{n,t-}}{P_{t-}} (G_{n,t+} - G_{n,t-}). \quad (\text{A.3})$$

The second alternative measure of ΔP_G is obtained by simply assuming that

$$\frac{G_{n,t+}}{G_{n,t-}} = \frac{G_{m,t+}}{G_{m,t-}}, \quad \forall n > m, \quad \forall t. \quad (\text{A.4})$$

Based on this assumption, we can calculate

$$\Delta P_{G,t}^2 = \sum_{n=1}^m \frac{B_{n,t-}}{P_{t-}} (G_{n,t+} - G_{n,t-}) + \left(\frac{G_{m,t+}}{G_{m,t-}} - 1 \right) \left(1 - \sum_{n=1}^m \frac{G_{n,t-} B_{n,t-}}{P_{t-}} \right). \quad (\text{A.5})$$

Table A.2 shows that our conclusions from Panel B, Table 3 continue to hold. Both alternative measures of ΔP_G have economically and statistically insignificant responses to

TABLE A.2
Regressing Alternative Measures of Price Changes Implied by Changes in Dividend Futures
Prices on Monetary Policy Surprises

The first two columns report the OLS estimates of regressing price changes on the S&P 500 market index on monetary policy surprises on scheduled FOMC announcement days. The monetary policy surprise measures include the policy news shock from [Nakamura and Steinsson \(2018\)](#) as updated by [Acosta \(2023\)](#) (POLICY) and the unexpected changes in the target Federal funds rates (FFR). Both shocks are standardized to have unit standard deviations. Columns (3) and (4) use ΔP_G^1 as the dependent variable, which is based on estimated dividend futures prices following [Knox and Vissing-Jørgensen \(2022\)](#). Columns (5) and (6) use ΔP_G^2 as the dependent variable, which is based on the assumption that dividend futures price changes are the same for maturities longer than 7 years. The remaining columns report tests for the difference between slope coefficients. All dependent variables are quoted in basis points. We report the t -statistics calculated using heteroskedasticity-robust standard errors in brackets. The sample period runs from November 2002 to December 2023.

	ΔP		ΔP_G^1		ΔP_G^2		$\Delta P - \Delta P_G^1$		$\Delta P - \Delta P_G^2$	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
POLICY	-36.84 [-2.53]		1.49 [0.36]		4.92 [0.78]		-38.33 [-2.39]		-41.75 [-2.48]	
FFR		-24.33 [-1.37]		-3.31 [-0.66]		-11.12 [-0.92]		-21.02 [-1.02]		-13.21 [-0.49]
Constant	32.91 [3.39]	32.03 [3.25]	0.54 [0.14]	0.72 [0.19]	-1.07 [-0.15]	-0.47 [-0.07]				
N	158	158	158	158	158	158				
Adj. R^2	0.07	0.02	-0.01	-0.00	-0.00	0.01				

monetary policy surprises. We can also reject the hypothesis that implied price changes induced by dividend price changes and actual price changes have the same responses to monetary policy surprises.

B Replicating [Kekre et al. \(2024\)](#)

Panel B in [Table B.1](#) replicates the [Kekre et al. \(2024\)](#) analysis, using the same sample period (2004 to 2016), the same monetary policy surprises based on [Jarociński and Karadi \(2020\)](#) (JK), and the same GSW forward rate data. As in their analysis, we find a negative response of nominal forward rates to the JK surprises at the very long end of the curve. However, as [Panel A](#) shows, this effect is completely absent when we use forward rates from the FPY

TABLE B.1
Forward Yields' Response to Monetary Policy Surprises

This table reports the results of regressing daily changes in 1-year forward yields on daily changes in 1-year zero-coupon yields instrumented by monetary policy surprises on scheduled FOMC announcement days. The forward yields are based on zero-coupon yields constructed by [Filipović et al. \(2022\)](#) (FPY) and [Gürkaynak et al. \(2007\)](#) (GSW), respectively. The monetary policy surprise measures include the changes in the 3-month Federal funds futures prices from [Jarociński and Karadi \(2020\)](#) (JK) and the policy news shock from [Nakamura and Steinsson \(2018\)](#) as updated by [Acosta \(2023\)](#) (POLICY). All regressions include a constant term that is not reported. We report the t -statistics calculated using heteroskedasticity-robust standard errors in the brackets. The sample runs from January 2004 to December 2016.

Panel A: FPY												
	1Y		2Y		5Y		10Y		20Y		30Y	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
JK	0.99		1.41		1.04		0.59		0.12		-0.04	
	[5.21]		[4.80]		[3.42]		[1.48]		[0.21]		[-0.17]	
POLICY		1.03		1.50		1.10		0.59		0.10		-0.04
		[5.92]		[5.61]		[3.85]		[1.48]		[0.17]		[-0.16]
N	104	104	104	104	104	104	104	104	104	104	104	104
Adj. R^2	0.39	0.38	0.30	0.31	0.10	0.10	0.05	0.04	-0.01	-0.01	-0.01	-0.01
Panel B: GSW												
	1Y		2Y		5Y		10Y		20Y		30Y	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
JK	1.07		1.56		0.88		0.09		-0.52		-0.88	
	[5.40]		[4.48]		[2.16]		[0.25]		[-1.97]		[-1.93]	
POLICY		1.01		1.52		0.89		0.13		-0.43		-0.71
		[6.55]		[5.37]		[2.61]		[0.44]		[-1.78]		[-1.73]
N	104	104	104	104	104	104	104	104	104	104	104	104
Adj. R^2	0.34	0.33	0.26	0.28	0.04	0.05	-0.01	-0.01	0.03	0.02	0.04	0.03

dataset over the same sample period.

C Additional results on high-frequency responses

Table [C.1](#) reports the results we obtain when we linearly interpolate high-frequency yield responses on unobserved maturities between adjacent knots instead of assuming a step function.

TABLE C.1
High-Frequency Stock Market Response to Monetary Policy Surprises, Linearly
Interpolated Yield Responses

The first four columns report the OLS estimates of regressing the 30-minute-window actual price changes of S&P 500 futures around scheduled FOMC announcements on monetary surprise measures. Actual price changes are percentage changes in Panel A and log changes in Panel B. The monetary surprise measures include the policy shocks from Nakamura and Steinsson (2018) as updated by Acosta (2023) (POLICY), the unexpected changes in the target Federal funds rates (FFR), the unadjusted monetary policy surprises from Bauer and Swanson (2023a) (BS), and residuals from regressing BS on six macro and financial variables ($BS^\perp$). All surprises are standardized to have unit standard deviations. The next four columns use the counterfactual price changes based on the 30-minute-window yield responses as the dependent variable. Counterfactual price changes are calculated using the dividend futures method in Panel A and the Campbell-Shiller present-value identity method in Panel B. Yield responses of unobserved maturities are linearly interpolated. The last four columns test the difference between slope coefficients. All dependent variables are quoted in basis points. All regressions include a constant term that is not reported. We report the t -statistics calculated using heteroskedasticity-robust standard errors in brackets. The sample period runs from November 2002 to December 2019 in Panel A and February 1994 to December 2019 in Panel B.

Panel A: Dividend Futures Method												
	ΔP^H				ΔP_B^H				$\Delta P^H - \Delta P_B^H$			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
POLICY	-27.06 [-4.39]				-33.24 [-5.78]				6.18 [0.77]			
FFR		-12.15 [-1.23]				-10.47 [-2.17]				-1.68 [-0.17]		
BS			-30.36 [-6.20]				-32.44 [-4.96]				2.08 [0.26]	
BS [⊥]				-32.64 [-5.60]				-25.21 [-4.78]				-7.44 [-0.94]
N	128	128	128	128	128	128	128	128				
Adj. <i>R</i> ²	0.17	0.03	0.22	0.27	0.24	0.02	0.24	0.14				
Panel B: Campbell-Shiller PV Identity Method												
	Δp^H				Δp_F^H				$\Delta p^H - \Delta p_F^H$			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
POLICY	-23.36 [-5.10]				-41.01 [-7.12]				17.65 [2.33]			
FFR		-13.46 [-2.42]				-9.73 [-2.51]				-3.73 [-0.51]		
BS			-26.47 [-6.51]				-40.08 [-6.62]				13.61 [1.81]	
BS [⊥]				-28.71 [-6.62]				-35.02 [-6.69]				6.31 [0.86]
N	200	200	208	208	200	200	208	208				
Adj. <i>R</i> ²	0.16	0.06	0.22	0.25	0.30	0.01	0.29	0.22				

D Estimation of γ from Blue Chip long-term forecasts

We find that survey forecasts of the average short-term interest rate between 7 to 11 years ahead are a good proxy of the perceived long-run means of short rates. Regressing the forecasts of this average between horizons 7 to 11 years on the 6-year forecasts yields a coefficient of 0.94 for BCFF and 0.95 for BCEI, which means that movements in 6-year forecasts largely reflect movements in the perceived long-run mean that are shared by forecasts at longer horizons, rather than slope or curvature changes that affect the short end of the forecast curve but not the perceived long-run mean. Thus, we treat μ in the following perceived AR(1) process,

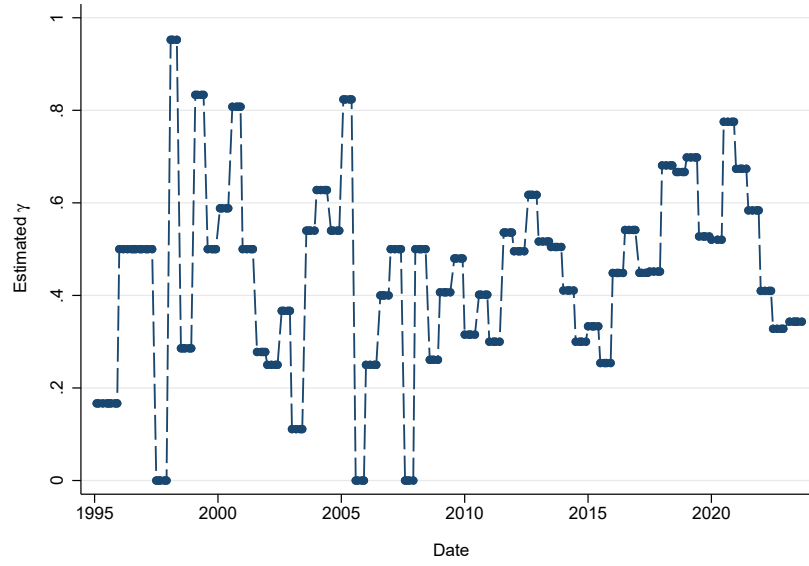
$$i_{t+n+1} - \mu = \gamma(i_{t+n} - \mu) + \tilde{\eta}_{t+n+1}, \quad (\text{D.1})$$

as observable and use the 7- to 11-year average to measure it. The perceived AR(1) dynamics imply

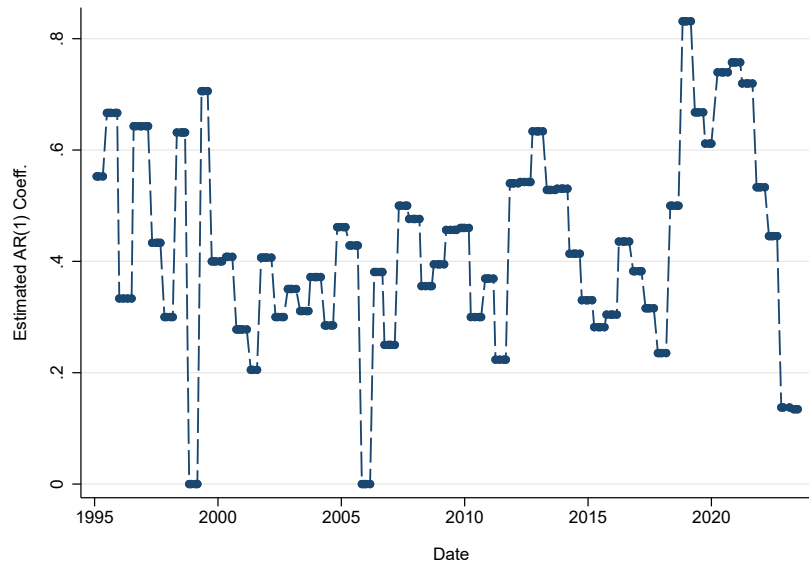
$$\tilde{\mathbb{E}}_t i_{t+n+1} - \mu = \gamma(\tilde{\mathbb{E}}_t i_{t+n} - \mu), \quad n \geq 1. \quad (\text{D.2})$$

In every long-range survey, we use the annual-horizon forecasts to calculate the demeaned forecasts $\tilde{\mathbb{E}}_t i_{t+n} - \mu$ for $n = 1, 2, \dots, 6$. We then fit an OLS regression in the cross-section of forecast horizons n (without a constant) of $\tilde{\mathbb{E}}_t i_{t+n+1} - \mu$ on $\tilde{\mathbb{E}}_t i_{t+n} - \mu$ to obtain an estimate of γ . Given the bi-annual frequency of estimated γ , we match monthly short-term forecasts to γ from the nearest month. For example, we match the short-term forecast reported in April 2022 to γ estimated in June 2022. For months with equal distance to long-range survey months, e.g., September 2022, we match them to the earlier long-range survey month.

Figure [D.1](#) plots the estimated γ in our sample period dated by FOMC announcement months.



(A) BCEI



(B) BCFF

FIGURE D.1

Estimated γ from Blue Chip long-range surveys

In each panel the blue dots plot the estimated γ from fitting an OLS regression to demeaned long-range annual forecasts from Blue Chip surveys. Panel A is for BCEI and Panel B is for BCFF. The sample period runs from 1995 to 2023.

TABLE E.1
Predicting Monetary Policy Surprise Measures with Economic News

This table reports regressions of monetary surprise measures on prior economic news. The economic news includes forecast revisions on real GDP and CPI, defined as the average revisions in the current and the future three quarters; 3-month changes in log prices of the S&P 500 index; 3-month changes in the yield curve defined as the spread between 10-year and 3-month Treasury yields; 3-month changes in the Bloomberg Commodity Index (BCOM). All explanatory variables are standardized to have unit standard deviations. We report the t -statistics calculated using heteroskedasticity-robust standard errors in brackets. The sample period runs from February 1995 to December 2023.

	POLICY			FFR		
	(1)	(2)	(3)	(4)	(5)	(6)
Rev_{rGDP}	0.13 [1.48]		0.04 [0.41]	0.09 [1.43]		0.08 [1.04]
Rev_{CPI}	0.17 [1.84]		0.13 [1.44]	0.13 [1.48]		0.16 [1.69]
$\Delta \log \text{S\&P 500}$		0.16 [1.95]	0.15 [1.68]		0.07 [0.81]	0.03 [0.36]
$\Delta \text{yield curve slope}$		-0.12 [-1.70]	-0.13 [-1.82]		-0.08 [-1.04]	-0.09 [-1.25]
$\Delta \log \text{BCOM}$		0.21 [2.43]	0.10 [1.15]		0.06 [0.73]	-0.05 [-0.51]
Constant	-0.00 [-0.05]	-0.01 [-0.17]	-0.01 [-0.18]	-0.01 [-0.16]	-0.00 [-0.07]	-0.02 [-0.24]
N	228	230	228	228	230	228
Adj. R^2	0.06	0.09	0.10	0.03	0.00	0.02

E Predicting monetary policy surprise measures with economic news

We report the results of using survey forecast revisions and financial news to predict monetary surprises measures in Table E.1.

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