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

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January 24, 2024

Abstract

We show that the stock market index price reaction to monetary policy announcements by 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.

*We are grateful for comments and suggestions to Benjamin Hébert, Ralph Koijen, Tyler Muir, Carolin Pflueger, and seminar participants at City University of Hong Kong. We thank Raymon Yue for excellent research assistance.

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1 Introduction

In their seminal study of the stock market reactions to monetary policy announcements by the Federal Open Market Committee (FOMC), Bernanke and Kuttner (2005) (BK) find that most of the stock market reaction is explained by changes of expected excess returns and very little by changes in expected real interest rates. They interpret this result as evidence of monetary policy 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 in FOMC announcement windows is explained by changes in the default-free term structure of yields, without equity 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, however. 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 the stock price falls in the case of an unexpected monetary tightening, the level of dividends is sticky in the short run, and hence the divided-price ratio rises. Based on the fact that movements dividend-price ratio are generally associated with movements in expected excess return, this then leads to the conclusion that expected excess returns must have moved in response to the monetary surprise. However, that the dividend-price ratio is generally associated with movements in expected excess return, as captured by the VAR estimated on a long sample of monthly data, does not necessarily mean that high-frequency movement in the dividend-price ratio around a selected set of specific events such as FOMC announcements is 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, say, with 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

$$P_t = \sum_{n=1}^{\infty} B_{n,t} G_{n,t}, \tag{1}$$

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 can be explained just to 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 on 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 might take.

What we find is very different from BK. Our estimates suggest that essentially all movement in the stock market index in FOMC announcement windows can be explained just with movements in the yield curve. There is very little room for additional effects from the effects of monetary surprises on the equity premium and the small component unaccounted for by yield curve changes is not statistically significant. This result holds for different definitions of the monetary 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 rates measured relative to pre-announcement Federal funds futures as in BK.

We can also ask how prices would have changed around FOMC announcements if the yield curve had not changed at all and only dividend futures prices have changed. In other words, we hold the yield curve fixed and allow risk-neutral expectations of future dividends to change. For this analysis we hold $B_{n,t}$ fixed at pre-announcement values and look at the change in implied price due to changes in $G_{n,t}$. We find that this implied price change is very close to zero. This further confirms the central result that yield curve movements, not movements in the equity premium explain stock prices changes in FOMC announcement windows.

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 zero-coupon yields of bonds at maturities beyond 7 years do not change in response to the monetary surprise. If this assumption holds, then only dividend futures with maturity up to 7 years are needed to compute the stock market index price change implied by the change in the yield curve. Empirically, the bulk of yield reactions to FOMC announcements happens at maturities lower than 7 years, so this assumption is reasonable. Nevertheless, to check robustness, we use the approach of Knox and Vissing-Jorgensen (2022) to estimate futures prices at maturities above 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 based on observed changes in the forward rate curve in the FOMC announcement window, this framework allows us to calculate the stock market index price change implied by changes in the yield curve. We obtain results very similar to our baseline analysis with dividend futures. Moreover, dividend futures prices are only available at daily frequency, but with the Campbell-Shiller approach, we can also use high-frequency changes in stock prices and forward rates within 30-minute windows around FOMC announcements. The results are similar.

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 the reaction. For the Federal funds rate surprise measure, we attribute almost the entire reaction to the term premium.

Finally, the dominant role of yield curve changes in stock price reaction FOMC announcements 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.

BK's findings have spawned a substantial theoretical literature that aims for models of asset pricing 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. Monetary policy affects the quantity of risk in Bianchi et al. (2022). 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. In contrast, Caballero and Simsek (2023), for example, focus on a traditional interest-rate and cash-flow channels of monetary policy. Our findings lend support to the latter traditional approach.

Our decomposition of announcement window stock market index returns is related to, but different from, the decomposition of Knox and Vissing-Jorgensen (2022). They also use dividend futures data, but they are interested in a decomposition into price changes due to risk-free rate news, news about expected excess returns, and news about cash flows. This decomposition requires a number of assumptions. As we show, if one is interested only in the change 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.

Bauer et al. (2023) find that there are significant changes in measures of investor risk appetite around FOMC announcements. However, these risk appetite measures are rather indirect proxies that do not directly capture changes in the equity premium. Our model-free evidence on the large role of yield curve changes suggests that there is little room left for equity premium changes in response to unexpected monetary tightening or easing. That said, we focus, like BK, on regression of stock market reaction on monetary policy surprise measures constructed from unexpected changes in interest rates. Since monetary policy surprise measures are not perfectly correlated with stock market index returns around FOMC announcements, this does not rule out that the equity premium may play a bigger role in explaining risky asset price reactions that are orthogonal to monetary policy surprise measures constructed from interest rates, as examined in Kroencke et al. (2021).

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

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

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

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

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

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

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

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

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

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

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

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

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

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

where $f_{n,m,t} \equiv \frac{1}{n-m}(n y_{n,t} - m y_{m,t})$ 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 assumptions. In addition, Table A.1 in Appendix A.1 shows that the means of $\frac{B_{n,t+}}{B_{n,t-}} - \frac{B_{m+1,t+}}{B_{m+1,t-}}$ are not statistically significantly different from 0 on FOMC announcement days for $m = 7$. We also show that this difference does not comove with monetary shocks.

Now with (6), we have

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

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.

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

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

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

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

This calculation implicitly assumes that expectations about dividends beyond maturity m do not change in response to the news conveyed by the FOMC announcement.

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-Jorgensen (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

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

with $\rho = 1/(1+\exp(\overline{d-p}))$ 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 2022 to estimate $\overline{d-p} = -3.44$. We use this value in our calculations.

We can now use the present-value identity to obtain the price change implied by changes in the yield curve. Fort this purpose, we decompose the return r_{t+n} into an excess return and forward rates. Let $f_{n,t}$ be the log forward rate for the period between $t+n-1$ to $t+n$:

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

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

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

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} and $x_{n,t+n}$:

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

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

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

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 maturity of less than one 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 maturing of one, two, and three 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 September 2022 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 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.

Dates of FOMC meetings between 1988 and 2019 are obtained from Bauer and Swanson (2023a). We obtain the remaining FOMC meeting dates until December 2022 from the Federal Reserve Board website.

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 maturities.

For the aggregate stock market prices, we use daily prices on the S&P 500 index from CRSP. 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

[2] We thank Miguel Acosta for providing the data on his website.
[3] We thank the authors for kindly sharing the data.
[4] We thank the authors for kindly sharing the data.

Table 1
Summary Statistics

This table reports the summary statistics. In Panel A, POLICY is the first principal component of 30-minute price changes in five interest-rate futures prices of maturities up to one year as in Nakamura and Steinsson (2018) and as updated by Acosta (2023). FFR is the price change of Federal funds futures expiring at the end of the month after the FOMC meeting in percentage points. Panel B shows the actual log and percentage price change of the S&P 500 index around scheduled FOMC announcements, Δp and ΔP, respectively, along with the corresponding counterfactual log and percentage price change based purely on yield changes based on the Campbell-Shiller present-value identity, Δp_B or the dividend futures approach, ΔP_B, as well as the counterfactual price change using only changes in dividend futures prices and holding fixed the yield curve, ΔP_G. Δp^{HF} and Δp_B^{HF} are the high-frequency actual and implied log price changes of the S&P 500 index using data from Bauer and Swanson (2023a). All price changes are shown in basis points.

	Mean	S.D.	Median	5th	95th
A. Monetary Surprise Measures					
POLICY	0	1	0.08	-1.71	1.46
FFR	-0.41	3.93	0	-6.35	5.64
B. Price Changes on Announcement Days					
Feb 1994–Dec 2022:					
Δp	27.38	116.09	12.95	-140.09	213.80
Δp_B	11.22	117.76	1.58	-148.08	185.69
Oct 2002–Dec 2022:					
ΔP	33.22	123.87	12.96	-134.96	261.56
ΔP_B	4.93	68.37	5.74	-94.38	91.53
ΔP_G	-0.84	8.88	0.15	-13.59	8.28
Feb 1994–Dec 2019:					
Δp^{HF}	-2.17	56.58	-4.71	-93.37	82.45
Δp_B^{HF}	2.29	74.76	0.88	-98.20	102.96

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. Based on the long sample from 1994 to 2022, the counterfactual log price change implied by yield curve changes using the Campbell-Shiller method, Δp_B, has almost the same volatility as the actual log price change Δp. In the shorter sample from 2002 to 2022 in which we have dividend futures data available, the counterfactual percentage price change just based on yield curve changes using the dividend futures method, ΔP_B, also has high volatility, but only about three quarters of the volatility of actual percentage price changes. In contrast, the counterfactual price change ΔP_G that keeps the yield curve fixed and uses only changes in dividend futures prices has much smaller standard deviation. 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.

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

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

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 responses 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.

4.1 Treasury yield response

Panel A in Table 2 shows that monetary policy surprises have effect 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 20Y maturity that is still more than a third of the yield change at a 1Y 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

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

[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.

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 one-day changes in zero-coupon Treasury yields. In Panel B, the dependent variables are one-day changes in 1-year forward yields. 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 shocks are standardized in the full sample period to have unit standard deviations. The sample only uses scheduled announcement days and is from February 1995 to September 2022. All regressions include a constant term whose estimates are not reported. We report t-statistics calculated using heteroskedasticity-robust standard errors in the brackets.

Panel A: Δy_n

	1Y		2Y		5Y		7Y		20Y		30Y	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
POLICY	3.74		4.23		3.85		3.20		1.41		0.97	
	[10.47]		[9.41]		[8.11]		[5.85]		[2.73]		[1.76]	
FFR		2.32		2.03		1.74		1.38		0.65		0.54
		[5.60]		[3.58]		[3.25]		[2.49]		[1.29]		[1.01]
N	221	221	221	221	221	221	221	221	221	221	221	221
Adj. R^2	0.47	0.18	0.40	0.09	0.25	0.05	0.17	0.03	0.05	0.01	0.02	0.00

Panel B: Δf_n

	1Y		2Y		5Y		7Y		20Y		30Y	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
POLICY	3.74		4.72		2.84		1.57		-0.02		0.40	
	[10.47]		[7.58]		[4.38]		[1.68]		[-0.02]		[0.54]	
FFR		2.32		1.73		1.36		0.50		0.81		0.25
		[5.60]		[2.27]		[2.30]		[0.55]		[0.96]		[0.33]
N	221	221	221	221	221	221	221	221	221	221	221	221
Adj. R^2	0.47	0.18	0.30	0.04	0.09	0.02	0.03	-0.00	-0.00	0.01	-0.00	-0.00

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 market index. Here we have a sample that covers later time periods and, to match with our dividend futures data, we use the S&P 500 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 6.4sd movement from the sample mean is associated with a $6.4 \times 19.89 \approx 128$ bps increase in the market price. Using the POLICY surprise measures instead, the regression produces a statistically significant coefficient of -25.25 (t-stat -2.37). Combined with column (1) in Panel A of Table 2, it implies that a POLICY shock that decreases the 1-year nominal Treasury yield by 25 bps is associated with a $25.25 \times (25/3.74) \approx 169$ bps increase in the market prices.

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_B, 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 -27.89, which is close in magnitude to the estimate of -41.82 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.

Table 3

Stock Market Response to Monetary Policy Surprises

The first two columns report the OLS estimates from regressing price changes 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 shocks are standardized to have unit standard deviations. Columns (3) and (4) use ΔP_B as the dependent variable in Panel A and Δp_B 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 September 2022 in Panel A and from February 1995 to September 2022 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	-41.82		-27.89		0.63		-13.93		-42.45	
	[-2.79]		[-4.07]		[0.79]		[-0.97]		[-2.76]	
FFR		-23.07		-13.67		-0.94		-9.40		-22.13
		[-1.26]		[-1.74]		[-0.72]		[-0.51]		[-1.15]
Constant	37.75	36.50	6.45	5.56	-0.82	-0.75				
	[3.87]	[3.61]	[1.22]	[1.00]	[-1.12]	[-1.04]				
N	149	149	149	149	149	149				
Adj. R^2	0.09	0.02	0.13	0.03	-0.00	0.00				

Panel B: Campbell-Shiller PV Identity Method

	Δp		Δp_B		$\Delta p - \Delta p_B$	
	(1)	(2)	(3)	(4)	(5)	(6)
POLICY	-25.25		-25.66		0.41	
	[-2.37]		[-2.46]		[0.03]	
FFR		-19.89		-13.07		-6.82
		[-1.84]		[-1.26]		[-0.51]
Constant	30.27	30.27	9.17	9.17		
	[3.99]	[3.95]	[1.21]	[1.18]		
N	221	221	221	221		
Adj. R^2	0.04	0.03	0.04	0.01		

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 coefficients is almost exactly zero.

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 movement around FOMC days is associated with movements in expected excess returns to exactly the same extent as any other dividend-price ratio movement on all other non-FOMC days. Our methods do not require such a strong assumption and 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 forward rate changes

Since the implied price changes ΔP_B and Δp_B 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 in the summations that yield ΔP_B and Δp_B on the monetary policy surprise measures. This provides a further check on the assumption that our dividend futures method is based on.

This means that for ΔP_B, we regress the $\frac{G_{n,t-}}{P_{t-}}(B_{n,t+} - B_{n,t-})$ 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-Jorgensen (2022) method as outlined in Appendix A.2. Figure 1 shows the result. In both panels, the negative coefficients flatten out at maturity of close to 7 years (in fact, it even reverts a bit for POLICY in panel A, but the coefficients at very long maturities come with big standard errors). Our dividend futures method assumes this flattening out, as we assume, to get to expression (8), 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. Here 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 7 years to be consistent with the assumption we make in the dividend futures method. As the figure shows, the estimates are broadly in line with this assumption. Overall, with both methods, changes in the forward rate curve out to about 7-year maturity drive almost all the variation in the implied price change measures.

4.4 High-frequency responses

So far we used daily yield changes on FOMC announcement days. To further eliminate noise in yield changes that are not related to monetary surprise, we use the intraday yield responses in the 30-minute window around FOMC announcements to calculate a higher-frequency measure. We use high-frequency S&P 500 futures price change, the high-frequency monetary policy surprise (MPS) measure, and the the residuals from regressing MPS on six macro and financial variables ($\text{MPS}^{\perp}$) from Bauer and Swanson (2023a).[8]

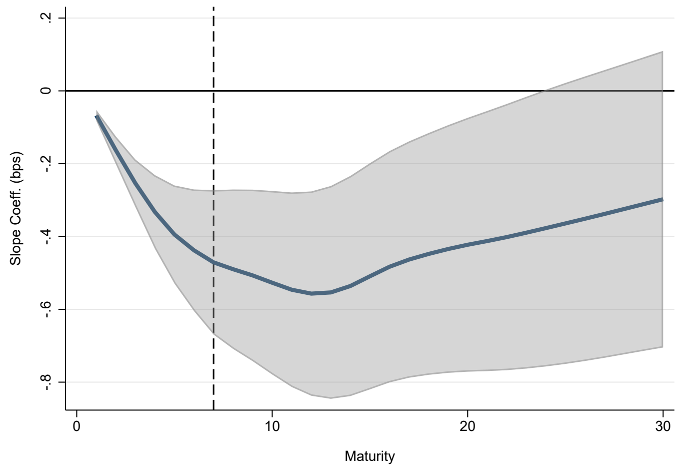
With high-frequency yield responses, we can construct the log prices implied by forward rate changes using Campbell-Shiller present-value identity approach as

$$\Delta p_{B,t}^{HF} = \sum_{n=1}^{30} \rho^{n-1}(f_{n,t-}-f_{n,t+}) \tag{17}$$

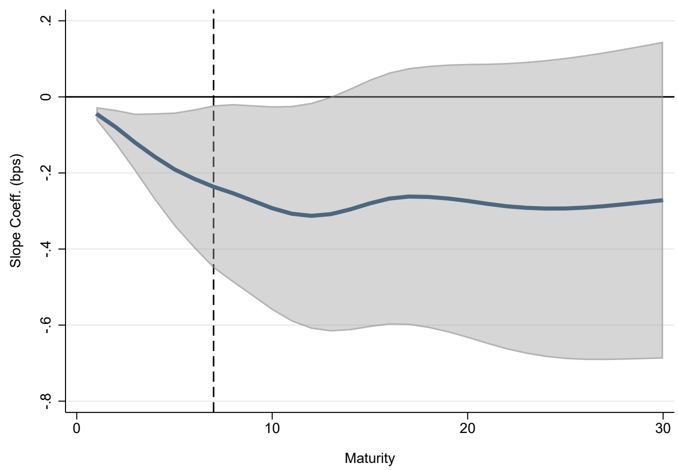
$$= \sum_{n=1}^{30} \rho^{n-1}\left[n(y_{n,t-}-y_{n,t+})-(n-1)(y_{n-1,t-}-y_{n-1,t+})\right], \tag{18}$$

using data on high-frequency changes in yields. We have high-frequency data for responses of 2-, 5-, 10-, and 30-year yields. 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

[8] We thank the authors for providing the data.



(A) POLICY

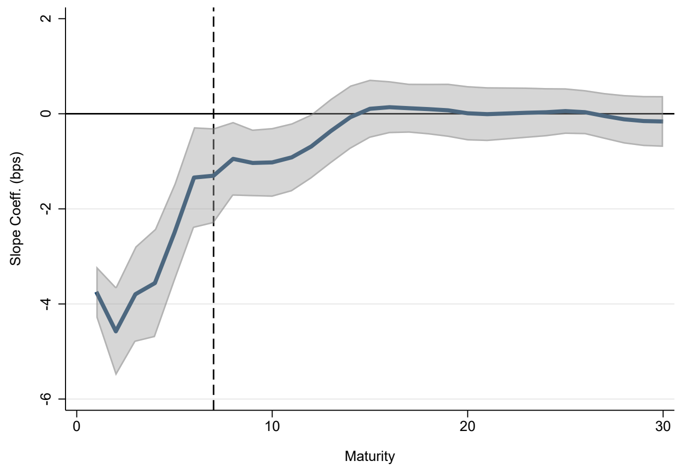


(B) FFR

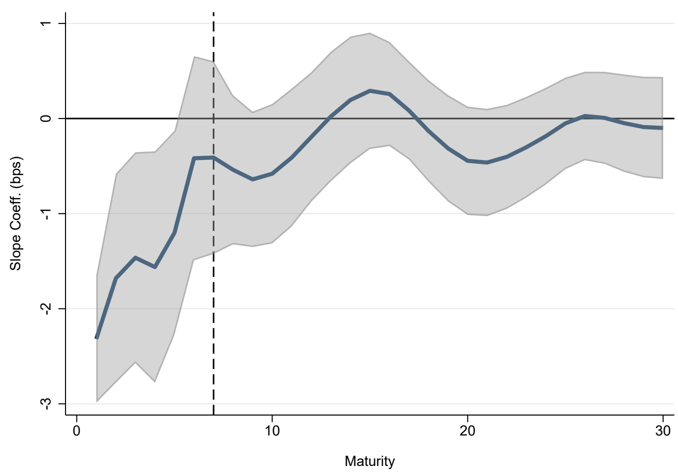
FIGURE 1

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-Jorgensen (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). Black dash 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). Black dash line indicates the 7-year maturity. The gray area represents the 95% confidence interval.

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 log returns of S&P 500 futures around scheduled FOMC announcements on monetary surprise measures. The monetary surprise measures include the unadjusted monetary policy surprises from Bauer and Swanson (2023a) (MPS), residuals from regressing MPS on six macro and financial variables (MPS$^{\perp}$), the policy shocks from Nakamura and Steinsson (2018) (POLICY), and the unexpected changes in the target Federal funds rates (FFR). All shocks are standardized to have unit standard deviations. The next four columns use the implied price changes based on the 30-minute-window yield responses as the dependent variable. Yield responses on 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. We report the t-statistics calculated using heteroskedasticity-robust standard errors in brackets. The sample period runs from February 1994 to September 2019 for MPS and MPS$^{\perp}$ and from February 1995 to September 2019 for POLICY and FFR.

	Δp^{HF}				Δp_B^{HF}				$\Delta p^{HF} - \Delta p_B^{HF}$			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
MPS	-26.47				-36.63				10.16			
	[-6.51]				[-5.84]				[1.31]			
MPS$^{\perp}$		-28.71				-31.86				3.15		
		[-6.62]				[-5.99]				[0.42]		
POLICY			-23.05				-36.31				13.25	
			[-5.17]				[-6.16]				[1.72]	
FFR				-13.68				-7.91				-5.76
				[-2.42]				[-2.02]				[-0.76]
Constant	-2.17	-2.17	-2.54	-1.64	2.29	2.29	0.15	1.58				
	[-0.62]	[-0.64]	[-0.71]	[-0.42]	[0.51]	[0.49]	[0.03]	[0.30]				
N	208	208	200	200	208	208	200	200				
Adj.R^2	0.22	0.25	0.17	0.06	0.24	0.18	0.24	0.01				

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 B shows that our results are robust to using linear interpolation methods to estimate responses of unobserved maturities.

Table 4 shows that regardless of which monetary surprise measure we use, implied price changes based on high-frequency yield responses are very similar in magnitude to the actual high-frequency price changes. For all measures, the difference between actual price changes and price changes implied by yield changes is not statistically significant.

4.5 Regressing actual price changes on implied price changes

This exercise does not require data on monetary surprise measures. We thus check two sample periods: the post-1994 period as previous and the full sample period starting from 1988 as identified by Bauer and Swanson (2023a). We also check both scheduled and unscheduled announcements.

At daily frequency, we have two measures of implied price changes, ΔP_B and Δp_B. We regress ΔP and Δp on them, respectively. At high-frequency, we have the measure based on the Campbell-Shiller present-value identity Δp_B^{HF} available. We regress Δp on it.

Table 5 reports the results. At daily frequency, both measures show that actual and implied price changes are virtually uncorrelated. This is expected as the monetary surprise measures only capture a small fraction of variation in both the actual and implied price changes. At high-frequency, for scheduled announcements, the implied price change displays a statistically significant positive relationship with actual price changes. It captures about 9% variation in actual price change. For unscheduled announcements, the relationship is negative and statistically insignificant, possibly due to a small number of observations.

4.6 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. There is 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 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

Table 5

Regressing Actual Price Changes on Implied Price Changes

This table reports the results of regressing actual price changes of S&P 500 market index on implied price changes. Columns (1) and (3) use ΔP as dependent variables and ΔP_B as independent variables. The remaining columns use Δp as dependent variables and Δp_B as independent variables. All variables are quoted in basis points. We report the t-statistics calculated using heteroskedasticity-robust standard errors in brackets. "Scheduled" and "Unscheduled" use scheduled and unscheduled FOMC announcements, respectively. "Post-1994" uses the sample period running from February 1994 to December 2022. "All" uses the sample period running from February 1988 to December 2022.

	Daily					High-Frequency		
	Scheduled				Unscheduled	Scheduled		Unscheduled
	Post-1994		All			Post-1994	All	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
ΔP_B	0.15		0.15					
	[0.66]		[0.66]					
Δp_B		-0.09		-0.02	-0.16	0.23	0.24	-0.76
		[-0.78]		[-0.17]	[-0.81]	[4.29]	[4.62]	[-1.18]
Constant	32.46	28.36	32.46	24.29	42.13	-2.69	-2.01	19.28
	[3.19]	[3.81]	[3.19]	[3.73]	[2.11]	[-0.73]	[-0.68]	[1.41]
N	151	231	151	284	55	208	262	61
Adj. R^2	-0.00	0.00	-0.00	-0.00	0.00	0.09	0.09	0.11

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 is measured with less precision than the daily or high-frequency price changes that we looked at earlier. Also, since we do not observe short-rate forecasts for horizons beyond 2 years, we need to make some assumptions to obtain expectations at longer horizons.

First, we assume that forecasters perceive short rates, i_t, to follow an AR(1) process:

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

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

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

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

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

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 $\tilde{\mathbb{E}}_t i_{t+1}$. Since these forecasts are only available at a monthly frequency, for a calendar date t in month $k(t)$, we use the beginning-of-month survey forecasts in month $k(t)$ for $\tilde{\mathbb{E}}_{t-} i_{t+1}$ and forecasts in month $k(t)+1$ for $\tilde{\mathbb{E}}_{t+} i_{t+1}$. 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 C, we provide more details on how we estimate γ using bi-annual surveys. 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^j_{t+1} 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:

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

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

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

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

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

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

The second part, reflecting changes in the term premium, is the residual:

$$\Delta P_{B,\lambda,t} \equiv \Delta P_{B,t} - \Delta P_{B,s,t}. \tag{25}$$

Now consider the Campbell-Shiller present-value identity approach. We can write (15) as

$$\Delta p_{B,t} = \sum_{n=1}^{\infty} \rho^{n-1} (f_{n,t-} - f_{n,t+}) \tag{26}$$

$$= \underbrace{\sum_{n=1}^{\infty} \rho^{n-1} \left(\mathbb{E}_{t-} i_{t+n-1} - \mathbb{E}_{t+} i_{t+n-1} \right)}_{\Delta p_{B,s,t}} + \sum_{n=1}^{\infty} \rho^{n-1} (\theta_{n,t-} - \theta_{n,t+}) \tag{27}$$

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

$$\Delta p_{B,s,t} \equiv \sum_{n=0}^{\infty} \rho^{n} \left(\mathbb{E}_{t-} i_{t+n} - \mathbb{E}_{t+} i_{t+n} \right), \tag{28}$$

which, using (21), we can rewrite as

$$\Delta p_{B,s,t} = i_{t-} - i_{t+} + \frac{\rho}{1-\rho\gamma} (\mathbb{E}_{t-} i_{t+1} - \mathbb{E}_{t+} i_{t+1}). \tag{29}$$

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

$$\Delta p_{B,\theta,t} \equiv \Delta p_{B,t} - \Delta p_{B,s,t}. \tag{30}$$

We then repeat our earlier exercise of regressing these implied price changes on monetary surprise measures. Table 6 and Table 7 report the results using short-rate forecasts from BCEI and BCFF, respectively. The short-term interest rate forecasted in the two surveys is the same—the 3-month T-bill rate—but the within-month timing of the BCEI and BCFF differs, which could potentially make a difference in this analysis because we cannot measure forecast changes in tight windows around the FOMC announcement.

That said, the results in Table 6 and Table 7 are very similar. Moreover, in both cases, the

Table 6

Decomposing Implied Price Responses: Short-Rate Expectations and Term Premium, BCEI

The first two columns report the results from regressing implied price changes induced by changes in short rate expectations from BCEI, $\Delta P_{B,s}$ in Panel A and $\Delta p_{B,s}$ in Panel B, on monetary surprise measures. The monetary surprise measures include the policy shocks from Nakamura and Steinsson (2018) (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 the implied price changes induced by changes in term premium, $\Delta P_{B,\lambda}$ in Panel A and $\Delta p_{B,\theta}$ in Panel B, as the dependent variable. 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 September 2022 in Panel A and from February 1995 to September 2022 in Panel B.

	Short Rate		Term Premium	
	(1)	(2)	(3)	(4)
A. Dividend Futures Method				
POLICY	-14.19		-13.70	
	[-2.85]		[-1.41]	
FFR		-1.38		-12.29
		[-0.23]		[-1.11]
Constant	5.64	5.00	0.81	0.56
	[2.02]	[1.64]	[0.13]	[0.09]
N	149	149	149	149
Adj. R^2	0.12	-0.01	0.02	0.02
B. Campbell-Shiller PV Identity Method				
POLICY	-10.88		-14.78	
	[-3.35]		[-1.23]	
FFR		1.70		-14.76
		[0.39]		[-1.22]
Constant	5.47	5.47	3.70	3.70
	[2.12]	[2.04]	[0.46]	[0.46]
N	221	221	221	221
Adj. R^2	0.07	-0.00	0.01	0.01

results based on the dividend futures method and the Campbell-Shiller present-value identity method are similar, too. For the POLICY shock 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 the term premium. In contrast, for the FFR surprise measure, the term premium drives 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.

4.7 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. Bauer and Swanson (2023b) show that public economic news predict monetary surprise measures. As they point out, such predictability can arise, for example, because investors are learning about the monetary policy rule parameters while econometricians have a hindsight knowledge advantage by using data in such predictability regressions that was not available to investors at the time they priced assets prior to the FOMC announcements.

Whether market's responses to monetary shocks are driven by response to economic news or Fed's private information does not affect our results above on the role of yield curve movements in explaining the stock market response to monetary policy surprises. But it would nevertheless be interesting to 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 the aforementioned two papers to orthogonalize the monetary surprise measure with respect to measures of information available before the FOMC announcement by regressing the monetary surprise measures $\psi \in \{\text{POLICY}, \text{FFR}\}$ on Blue Chip forecast revisions and public news:

$$\psi_t = \alpha + \beta \text{news}_{t-1} + \xi_t. \tag{31}$$

Table 7

Decomposing Implied Price Responses: Short-Rate Expectations and Term Premium, BCFF

The first two columns report the results from regressing implied price changes induced by changes in short-rate expectations from BCFF, $\Delta P_{B,s}$ in Panel A and $\Delta p_{B,s}$ in Panel B, on monetary surprise measures. The monetary surprise measures include the policy shocks from Nakamura and Steinsson (2018) (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 the implied price changes induced by changes in term premium, $\Delta P_{B,\lambda}$ in Panel A and $\Delta p_{B,\theta}$ in Panel B, as the dependent variable. 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 September 2022 in Panel A and from February 1995 to September 2022 in Panel B.

	Short Rate		Term Premium	
	(1)	(2)	(3)	(4)
A. Dividend Futures Method				
POLICY	-18.59		-9.30	
	[-3.67]		[-0.97]	
FFR		-4.94		-8.73
		[-0.73]		[-0.77]
Constant	11.58	10.84	-5.13	-5.28
	[4.00]	[3.36]	[-0.83]	[-0.86]
N	149	149	149	149
Adj. R^2	0.19	0.01	0.01	0.00
B. Campbell-Shiller PV Identity Method				
POLICY	-13.42		-12.24	
	[-3.80]		[-1.05]	
FFR		-2.35		-10.72
		[-0.57]		[-0.92]
Constant	9.54	9.54	-0.37	-0.37
	[4.20]	[3.91]	[-0.05]	[-0.05]
N	221	221	221	221
Adj. R^2	0.13	-0.00	0.01	0.00

Specifically, we include the 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). Consistent with these earlier papers, we find that economic news contains information of future POLICY shocks. But FFR shocks are much less predictable. The results are shown in Appendix D. Based on these results, we focus on the decomposition of POLICY shocks in the following analysis.

As the predictable and unpredictable components are constructed in the first-step regression in (31), using them in a second-step regression as independent variables can lead to bias in standard errors. We deal with this generated regressor issue using a bootstrap method. Specifically, we create bootstrap samples by randomly drawing with replacement clusters of price changes and 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 (31) to construct the predictable and unpredictable components of monetary surprises, and regress price changes on them. We obtain the p-values by comparing the sample regression t-statistic to the quantiles of the distribution of the t-statistic in the bootstrap regressions.

Table 8 shows that the stock market responds to both components with a similar economic magnitude. 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. Responses of price changes implied by yield curve changes are extremely close to the total price response to the residual component. For the predictable component the point estimates for actual price change and yield-change implied price change differ to some extent, but the difference is far from being statistically significant

Table 8

Stock Market Response to Predictable and Unpredictable Monetary Policy Surprises

The first two columns report the OLS estimates of regressing price changes on the S&P 500 market index on predictable (Predicted) and unpredictable (Residual) POLICY shocks on scheduled FOMC announcement days. Both components are standardized to have unit standard deviations. Columns (3) and (4) use ΔP_B as the dependent variable in Panel A and Δp_B 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 measured in basis points. We report in parentheses the p-values based on the distribution of t-statistic from block cluster bootstrapped samples of price changes and standardized monetary surprises, survey forecast revisions, and financial news. The sample period runs from November 2002 to September 2022 in Panel A and from February 1995 to September 2022 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	-31.01		-12.70		1.15		-18.30		-32.16	
	(0.13)		(0.12)		(0.47)		(0.26)		(0.12)	
Residual		-28.65		-23.38		0.06		-5.26		-28.71
		(0.01)		(0.00)		(0.93)		(0.67)		(0.02)
Constant	37.78	36.48	5.94	5.72	-0.86	-0.79				
	(0.00)	(0.01)	(0.18)	(0.19)	(0.37)	(0.39)				
N	149	149	149	149	149	149				
Adj. R^2	0.07	0.04	0.03	0.09	0.01	-0.01				

B. Campbell-Shiller PV Identity Method

	Δp		Δp_B		$\Delta p - \Delta p_B$	
	(1)	(2)	(3)	(4)	(5)	(6)
Predicted	-19.14		-13.95		-5.19	
	(0.27)		(0.26)		(0.73)	
Residual		-19.77		-22.16		2.39
		(0.01)		(0.03)		(0.86)
Constant	30.27	30.27	9.17	9.17		
	(0.00)	(0.00)	(0.20)	(0.16)		
N	221	221	221	221		
Adj. R^2	0.02	0.02	0.01	0.03		

at conventional levels.

As Table 9 shows, the distinction between predictable and unpredictable components of the POLICY shock matters much more for our decomposition of the implied price change into a short-rate expectations component and a term premium component. Based on the point estimates, the predictable component is more strongly related to short-term interest rate survey expectations changes than the unpredictable component, while the term premium component mostly moves only with the unpredictable component. Statistically, however, there is not much we can conclude with much confidence. Further decomposing (into reactions to predictable and unpredictable components) the decomposition (into short-rate expectations and term premia) seems to cut the data a bit too thinly to leave sufficient statistical power.

4.8 UK evidence

Recent years have seen an increasing effort of collecting data on high-frequency responses of asset prices to monetary policy announcement in economies other than the US. We utilize the newly published UK Monetary Policy Event-Study Database to study whether in UK stock market responses in monetary policy announcement windows are also captured by yield curve movements.

The UK Monetary Policy Event-Study Database provides data on responses of Libor rates, 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 the 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. In the sample period between 1997 and 2021, the first principal component explains 93.4% of variance in the four contracts responses. As Panel A in Table 10 shows, consistent with the US evidence, the Gilt yield curve responds

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

Table 9

Implied Price Response to Predictable and Unpredictable Monetary Policy Surprises

This table reports the results from regressing implied price changes induced by changes in short-rate expectations (SR) and implied price changes induced by changes in term premium (TP) on predictable (Predicted) and unpredictable (Residual) POLICY shocks on scheduled FOMC announcement days. Both components are standardized to have unit standard deviations. Panel A uses $\Delta P_{B,s}$ for SR and $\Delta P_{B,\lambda}$ for TP. Panel B uses $\Delta p_{B,s}$ for SR and $\Delta p_{B,\theta}$ for TP. The first four columns use BCEI forecasts to measure short-rate expectations and the last four columns use BCFF forecasts. All dependent variables are quoted in basis points. We report in parentheses the p-values based on the distribution of t-statistics from block cluster bootstrapped samples of price changes and standardized monetary surprises, survey forecast revisions, and financial news. The sample period runs from November 2002 to September 2022 in Panel A and from February 1995 to September 2022 in Panel B.

A. Dividend Futures Method

	BCEI				BCFF			
	SR		TP		SR		TP	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Predicted	-13.56		0.86		-17.52		4.81	
	(0.07)		(0.92)		(0.06)		(0.52)	
Residual		-8.09		-15.29		-10.73		-12.65
		(0.12)		(0.17)		(0.05)		(0.27)
Constant	5.86	5.17	0.08	0.55	11.84	10.96	-5.90	-5.24
	(0.13)	(0.36)	(0.99)	(0.93)	(0.01)	(0.09)	(0.38)	(0.46)
N	149	149	149	149	149	149	149	149
Adj. R^2	0.16	0.03	-0.01	0.03	0.24	0.06	-0.00	0.02

B. Campbell-Shiller PV Identity Method

	BCEI				BCFF			
	SR		TP		SR		TP	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Predicted	-14.65		0.70		-16.87		2.92	
	(0.01)		(0.96)		(0.03)		(0.78)	
Residual		-6.10		-16.06		-7.97		-14.19
		(0.07)		(0.16)		(0.01)		(0.21)
Constant	5.47	5.47	3.70	3.70	9.54	9.54	-0.37	-0.37
	(0.08)	(0.25)	(0.63)	(0.64)	(0.00)	(0.04)	(0.96)	(0.96)
N	221	221	221	221	221	221	221	221
Adj. R^2	0.13	0.02	-0.00	0.01	0.21	0.04	-0.00	0.01

to monetary policy surprises with a coefficient that declines with maturities. Based on this finding, we construct a measure of implied price changes induced by yield curve responses only by a step-function interpolation as in Section 4.4. We choose $\rho = 0.967$ based on the average log dividend-price ratio on the FTSE All Share Index between 1997 and 2021 which is -3.37.

Column (1) in Panel B of Table 10 shows that a 1sd negative monetary policy surprise is associated with a 14.5 bps increase in the FTSE All Share Index that is statistically significant at conventional levels. Such a magnitude of response is fully captured by our measure of implied price changes constructed under the Campbell-Shiller approach using high-frequency yield responses.

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 regularity 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,

[10] In our sample period, week 6 only contains 146 days, while all other weeks contain at least 500 days. Including week 6 does not quantitatively change our results.

Table 10

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

Panel A reports the OLS estimates of regressing the 30-minute-window changes in Gilt yields around the Monetary Policy Committee's announcements on monetary policy surprise (MPS) in UK. MPS is calculated as the first principal component of the 30-minute-window changes in the first four quarterly Short Sterling Futures prices and standardized to have 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 implied price changes based on the 30-minute-window Gilt yield responses of maturities between 1 to 10 years as the dependent variable. 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 March 2021.

Panel A: Gilt Yield

	1Y (1)	2Y (2)	5Y (3)	10Y (4)
MPS	3.83	3.41	2.59	1.67
	[12.54]	[9.81]	[9.34]	[5.74]
Constant	-0.24	-0.27	-0.19	-0.07
	[-1.84]	[-2.20]	[-1.30]	[-0.40]
N	266	266	266	266
Adj. R^2	0.77	0.74	0.54	0.28

Panel B: FTSE All Share Index

	Δp (1)	Δp_B (2)	$\Delta p - \Delta p_B$ (3)
MPS	-14.50	-14.91	0.41
	[-7.17]	[-5.98]	[0.12]
Constant	1.04	0.68	0.36
	[0.65]	[0.49]	[0.19]
N	265	265	265
Adj. R^2	0.23	0.30	-0.00

without considering dividends or risk-free rate, from t to $t+4$ assuming the changes are zero over the weekends.

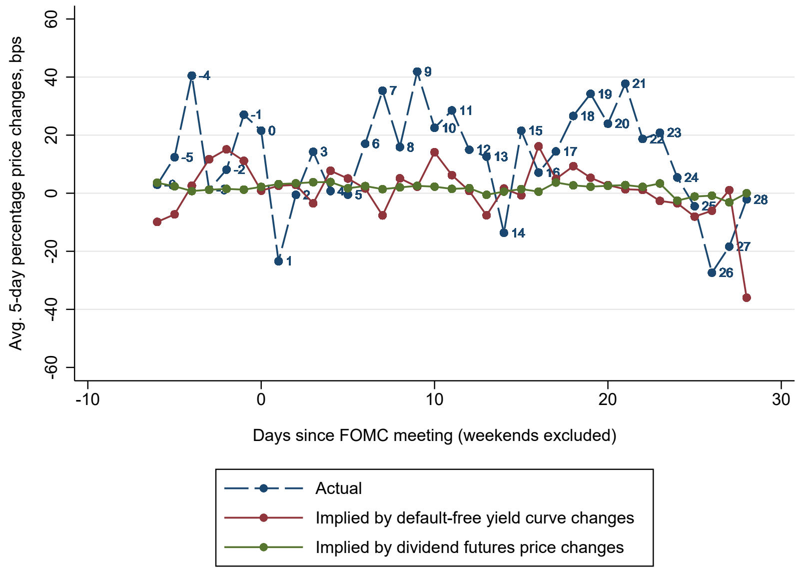
Panel A in Figure 3 shows that the cyclical pattern documented in Knox and Vissing-Jorgensen (2022) also holds in our shorter sample with dividend futures data availability from 2002 to 2022, but the pattern is not as pronounced as in their original study. Average price changes are higher in even weeks (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 clearly 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 the price change implied by dividend futures price changes is basically flat and not connected to the cycle pattern.

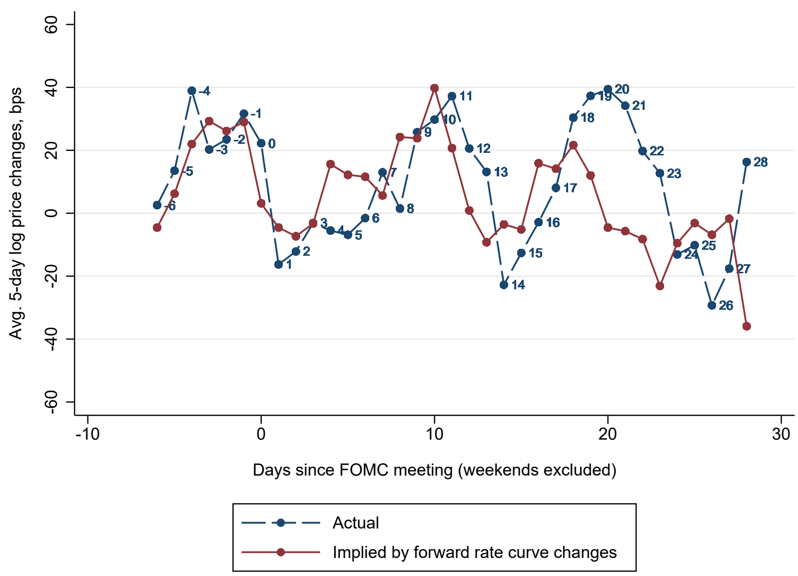
Panel B looks at the longer sample between 1995 and 2022, using log price changes and the Campbell-Shiller present-value identity method to calculate price changes implied by forward rate curve changes. In this longer sample, the pattern in actual price changes is more pronounced than in Panel A. The implied log price changes Δp_B 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 FOMC cycle.

In the shorter sample period since 2002, consistent with Figure 3, the point estimates in panel A, column (1) in Table 11 indicate that actual price changes are cyclical over the FOMC cycle. Average daily returns are 4.63 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 half of the even-week effect. In contrast, the estimated



(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 Nov 2002 and Dec 2022. 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 Feb 1995 and Dec 2022.

Table 11

Regressing Price Changes on FOMC Cycle Dummies

This table reports the coefficients from regressing actual price changes and implied price changes induced 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 2022 for the dividend futures method and from January 1994 to December 2022 for the Campbell-Shiller present-value identity method, respectively.

	Dividend Futures					Campbell-Shiller PV		
	ΔP	ΔP_B	ΔP_G	$\Delta P - \Delta P_B$	$\Delta P - \Delta P_G$	Δp	Δp_B	$\Delta p - \Delta p_B$
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
A. Even weeks								
Week 0, 2, 4	4.63	1.99	-0.13	2.64	4.76	7.56	5.23	2.33
	[1.34]	[1.36]	[-0.52]	[0.63]	[1.38]	[2.73]	[2.09]	[0.56]
Constant	1.37	-0.80	0.45	2.16	0.92	-0.85	-1.50	0.65
	[0.60]	[-0.83]	[2.59]	[0.79]	[0.41]	[-0.47]	[-0.90]	[0.24]
N	4639	4639	4639	4639	4639	7054	7054	7054
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	3.47	3.78	-0.27	-0.31	3.75	8.15	7.98	0.17
	[0.66]	[1.68]	[-0.77]	[-0.05]	[0.71]	[2.02]	[2.15]	[0.03]
Week 2	7.81	0.30	0.01	7.51	7.80	7.00	2.63	4.38
	[1.60]	[0.14]	[0.03]	[1.24]	[1.61]	[1.75]	[0.74]	[0.72]
Week 4	2.63	1.77	-0.12	0.86	2.75	7.49	4.88	2.61
	[0.53]	[0.85]	[-0.34]	[0.14]	[0.56]	[1.82]	[1.34]	[0.43]
Constant	1.37	-0.80	0.45	2.16	0.92	-0.85	-1.50	0.65
	[0.60]	[-0.83]	[2.59]	[0.79]	[0.41]	[-0.47]	[-0.90]	[0.24]
N	4639	4639	4639	4639	4639	7054	7054	7054
Adj. R^2	-0.00	0.00	-0.00	-0.00	-0.00	0.00	0.00	-0.00

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 to calculate log price changes implied by changes in the forward rate curve. Column (6) shows that log price changes are 7.56 bps higher in even weeks and the estimate is statistically significant at the 5% level. Using the implied price change Δp_B, we obtain a coefficient estimate of 5.23 which is also statistically significant. At 2.33 bps, 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 responses to unexpected monetary policy actions are mostly attributable to the effect of monetary policy surprises 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 and 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. However, our results in this paper highlight, that for studying the stock market reaction to surprises outside of FOMC announcement windows it would also be important to use a method that can account flexibly for shape and level changes in the yield curve without imposing strong functional form assumptions.

Appendix

A Constructing 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-Jorgensen (2022). Knox and Vissing-Jorgensen (2022) show that under the assumption of a Gordon growth model, we have

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

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

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

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 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 September 2022.

	$n=9$			$n=15$			$n=20$			$n=30$		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
Mean	0.26			-0.80			-1.61			5.13		
	[0.55]			[-0.28]			[-0.35]			[0.57]		
POLICY		-0.41			-0.04			2.82			14.57	
		[-0.78]			[-0.01]			[0.54]			[1.37]	
FFR			0.18			2.09			4.04			11.41
			[0.35]			[0.66]			[0.86]			[1.00]
N	220	220	220	220	220	220	220	220	220	220	220	220
Adj. R^2	-0.00	-0.00	-0.00	-0.00	-0.00	-0.00	-0.00	-0.00	-0.00	-0.00	0.01	0.00

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

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

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

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

Based on this assumption, we can calculate

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

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 Implied Price Changes 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-Jorgensen (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 September 2022.

	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	-41.82		1.73		8.30		-43.55		-50.13	
	[-2.79]		[0.37]		[1.07]		[-2.60]		[-2.79]	
FFR		-23.07		-3.25		-12.46		-19.82		-10.61
		[-1.26]		[-0.63]		[-0.88]		[-0.94]		[-0.36]
Constant	37.75	36.50	1.13	1.33	-2.03	-1.20				
	[3.87]	[3.61]	[0.29]	[0.34]	[-0.25]	[-0.15]				
N	149	149	149	149	149	149				
Adj.R^2	0.09	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 Additional results on high-frequency responses

Table B.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 B.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 log returns of S&P 500 futures around scheduled FOMC announcements on monetary surprise measures. The monetary surprise measures include the unadjusted monetary policy surprises from Bauer and Swanson (2023a) (MPS), residuals from regressing MPS on six macro and financial variables ($\text{MPS}^{\perp}$), the policy shocks from Nakamura and Steinsson (2018) (POLICY), and the unexpected changes in the target Federal funds rates (FFR). All shocks are standardized to have unit standard deviations. The next four columns use the implied price changes based on the 30-minute-window yield responses as the dependent variable. Yield responses on unobserved maturities are linearly interpolated. The last four 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 February 1994 to September 2019 for MPS and $\text{MPS}^{\perp}$ and from February 1995 to September 2019 for POLICY and FFR.

	Δp^{HF}				Δp_B^{HF}				$\Delta p^{HF} - \Delta p_B^{HF}$			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
MPS	-26.47				-40.37				13.90			
	[-6.51]				[-6.58]				[1.83]			
$\text{MPS}^{\perp}$		-28.71				-35.26				6.56		
		[-6.62]				[-6.66]				[0.88]		
POLICY			-23.05				-39.92				16.86	
			[-5.17]				[-6.92]				[2.24]	
FFR				-13.68				-9.90				-3.77
				[-2.42]				[-2.48]				[-0.51]
Constant	-2.17	-2.17	-2.54	-1.64	2.93	2.93	0.76	2.33				
	[-0.62]	[-0.64]	[-0.71]	[-0.42]	[0.66]	[0.63]	[0.17]	[0.44]				
N	208	208	200	200	208	208	200	200				
Adj. R^2	0.22	0.25	0.17	0.06	0.29	0.22	0.29	0.01				

C 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,

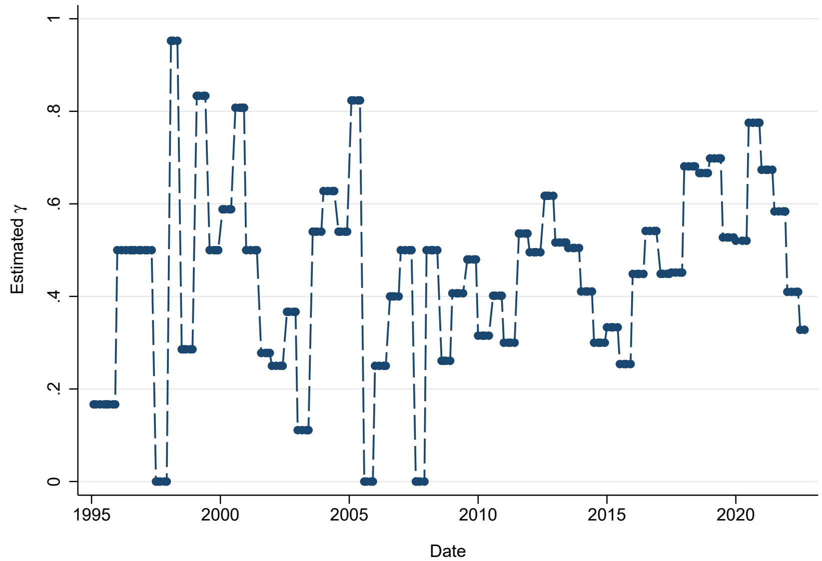
$$i_{t+n+1} - \mu = \gamma(i_{t+n} - \mu) + \tilde{\eta}_{t+n+1}, \tag{C.1}$$

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

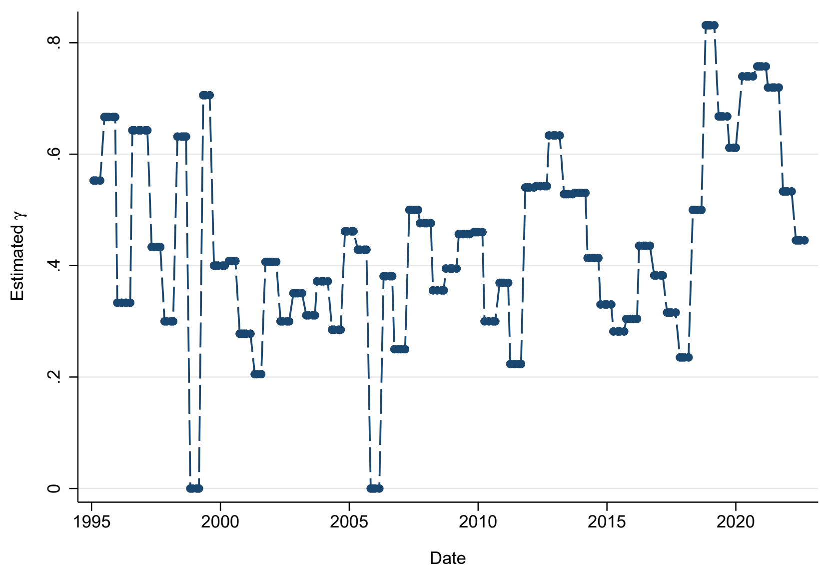
$$\tilde{\mathbb{E}}_t i_{t+n+1} - \mu = \gamma(\tilde{\mathbb{E}}_t i_{t+n} - \mu), \quad n \geq 1. \tag{C.2}$$

In every long-range survey, we use the annual forecasts to calculate the demeaned forecasts $\tilde{\mathbb{E}}_t i_{t+n} - \mu$ for $n = 1, 2, \ldots, 6$. 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 C.1 plots the estimated γ in our sample period dated by FOMC announcement months.



(A) BCEI



(B) BCFF

Figure C.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 and 2022.

D 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 D.1.

Table D.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 September 2022.

	POLICY			FFR		
	(1)	(2)	(3)	(4)	(5)	(6)
Rev_{rGDP}	0.11		0.02	0.08		0.07
	[1.17]		[0.17]	[1.36]		[0.97]
Rev_{CPI}	0.23		0.20	0.13		0.17
	[2.38]		[2.00]	[1.43]		[1.59]
$\Delta \log$ S&P 500		0.15	0.14		0.06	0.03
		[1.78]	[1.58]		[0.76]	[0.34]
Δyield curve slope		-0.12	-0.14		-0.08	-0.09
		[-1.60]	[-1.82]		[-0.98]	[-1.20]
$\Delta \log$ BCOM		0.20	0.08		0.07	-0.04
		[2.31]	[0.93]		[0.80]	[-0.42]
Constant	0.00	0.00	0.00	0.00	0.00	0.00
	[0.01]	[0.05]	[0.04]	[0.01]	[0.02]	[0.02]
N	221	221	221	221	221	221
Adj. R^2	0.07	0.09	0.10	0.02	0.00	0.02

References

Acosta, Miguel, 2023, The Perceived Causes of Monetary Policy Surprises, *Working Paper* .

Bauer, Michael D, Ben S Bernanke, and Eric Milstein, 2023, Risk appetite and the risk-taking channel of monetary policy, *Journal of Economic Perspectives* 37, 77–100.

Bauer, Michael D., and Eric T. Swanson, 2023a, A Reassessment of Monetary Policy Surprises and High-Frequency Identification, *NBER Macroeconomics Annual* 37, 87–155.

Bauer, Michael D., and Eric T. Swanson, 2023b, An Alternative Explanation for the "Fed Information Effect", *American Economic Review* 113, 664–700.

Bernanke, Ben S., and Kenneth N. Kuttner, 2005, What Explains the Stock Market's Reaction to Federal Reserve Policy?, *The Journal of Finance* 60, 1221–1257.

Bianchi, Francesco, Sydney C. Ludvigson, and Sai Ma, 2022, Monetary-based asset pricing: A mixed-frequency structural approach, *NBER Working Papers* .

Caballero, Ricardo J., and Alp Simsek, 2023, A monetary policy asset pricing model, *NBER Working Papers* .

Campbell, John Y., and John Ammer, 1993, What Moves the Stock and Bond Markets? A Variance Decomposition for Long-Term Asset Returns, *The Journal of Finance* 48, 3–37.

Campbell, John Y., and Robert J. Shiller, 1988, The Dividend-Price Ratio and Expectations of Future Dividends and Discount Factors, *The Review of Financial Studies* 1, 195–228.

Cieslak, Anna, Adair Morse, and Annette Vissing-Jorgensen, 2019, Stock Returns over the FOMC Cycle, *The Journal of Finance* 74, 2201–2248.

Cieslak, Anna, and Hao Pang, 2021, Common shocks in stocks and bonds, *Journal of Financial Economics* 142, 880–904.

Drechsler, Itamar, Alexi Savov, and Phillip Schnabl, 2017, A model of monetary policy and risk premia, *The Journal of Finance* 73, 317–373.

Filipović, Damir, Markus Pelger, and Ye Ye, 2022, Stripping the Discount Curve - a Robust Machine Learning Approach, *Working Paper* .

Hanson, Samuel G., and Jeremy C. Stein, 2015, Monetary Policy and Long-Term Real Rates, *Journal of Financial Economics* 115, 429–448.

Karnaukh, Nina, and Petra Vokata, 2022, Growth Forecasts and News about Monetary Policy, *Journal of Financial Economics* 146, 55–70.

Kekre, Rohan, and Moritz Lenel, 2022, Monetary policy, redistribution, and risk premia, *Econometrica* 90, 2249–2282.

Knox, Benjamin, and Annette Vissing-Jorgensen, 2022, A Stock Return Decomposition Using Observables, *Working Paper* .

Kroencke, Tim A., Maik Schmeling, and Andreas Schrimpf, 2021, The fomc risk shift, *Journal of Monetary Economics* 120, 21–39.

Martin, Ian, 2017, What is the Expected Return on the Market?, *The Quarterly Journal of Economics* 132, 367–433.

Nakamura, Emi, and Jón Steinsson, 2018, High-Frequency Identification of Monetary Non-Neutrality: The Information Effect, *The Quarterly Journal of Economics* 133, 1283–1330.

Pflueger, Carolin, and Gianluca Rinaldi, 2022, Why does the fed move markets so much? a model of monetary policy and time-varying risk aversion, *Journal of Financial Economics* 146, 277–281.

Politis, Dimitris N., and Halbert White, 2004, Automatic Block-Length Selection for the Dependent Bootstrap, *Econometric Reviews* 23, 53–70.

Van Binsbergen, Jules H, 2020, Duration-based stock valuation: Reassessing stock market performance and volatility, Technical report, National Bureau of Economic Research.

van Binsbergen, Jules H., and Ralph S.J. Koijen, 2017, The Term Structure of Returns: Facts and Theory, *Journal of Financial Economics* 124, 1 – 21.