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# The cross section of the monetary policy announcement premium[☆]

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ABSTRACT

Using the expected option-implied variance reduction to measure the sensitivity of stock returns to monetary policy announcement surprises, this paper shows monetary policy announcements require significant risk compensation in the cross section of equity returns. We develop a parsimonious equilibrium model in which FOMC announcements reveal the Federal Reserve's private information about its interest-rate target, which affects the private sector's expectation about the long-run growth-rate of the economy. Our model accounts for the dynamics of implied variances and the cross section of the monetary policy announcement premium realized around FOMC announcement days.



## 1. Introduction

This paper provides empirical evidence that the exposure to monetary policy announcement surprises is priced in the cross section of equity returns. We show market expectations about firms' sensitivity to monetary policy announcements strongly predict their equity returns on Federal Open Market Committee (FOMC) announcement days. A long-short portfolio formed on our monetary policy sensitivity measure produces an average announcement-day return of 31.40 basis points (bps), which is both statistically and economically significant, even after controlling for standard risk factors. We provide evidence that risk exposure to monetary policy announcements is related to firms' sensitivity to growth-rate expectations, and we develop a parsimonious model in which firms have differing levels of sensitivity with respect to the Federal Reserve's interest-rate policy, which is revealed periodically during FOMC announcements. We show our model can quantitatively account for the cross section of FOMC announcement

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premiums and the dynamics of implied variance around announcements.

The motivation of our empirical exercise is twofold. First, pre-scheduled monetary policy announcements have a large impact on stock market valuations (Bernanke and Kuttner, 2005), the private sector's expectation about economic growth (Nakamura and Steinsson, 2018), and the cost of credit (Gertler and Karadi, 2015). In addition, FOMC announcement days are associated with realizations of significantly higher average equity-market returns compared with days without major macroeconomic announcements (see, e.g., Savor and Wilson, 2013; Lucca and Moench, 2015; and Ai and Bansal, 2018). Equity returns realized on announcement days can be informative about not only risk and risk compensation on financial markets, but also the transmission mechanism of monetary policies.

Second, the robustness of the CAPM on macroeconomic announcement days demonstrated by Savor and Wilson (2014) implies macroeconomic announcement days present a unique challenge and an opportunity to identify robust new risk factors other than the market factor. Savor and Wilson (2014) show the expected return-beta relationship predicted by CAPM holds very well not only for market beta-sorted portfolios, but also for many other investment strategies, such as the book-to-market, size, industry, idiosyncratic volatility, and down-side beta-sorted portfolios that are known to violate the CAPM in general. From a theoretical perspective, the CAPM may fail for two reasons: time-varying beta or the presence of additional risk factors. Beta is unlikely to change substantially over a short announcement window; therefore, if CAPM fails on announcement days, the reason is likely a second risk factor. The theory of Ai and Bansal (2018) implies the stochastic discount factor (SDF) is much more volatile on announcement days when information about the macroeconomy is revealed. Viewed through the lens of this theory, the significantly positive slope of the security market line is nothing but evidence of a highly volatile SDF on announcement days from the cross section of equity returns. Interpreted this way, the Savor and Wilson's (2014) evidence implies none of the known risk factors listed above are powerful enough to overturn the CAPM on announcement days. Our empirical evidence can be interpreted as the first known example of such a risk factor.

To evaluate whether firms with differing levels of sensitivity to monetary policy announcements also have differing expected returns on announcement days, we first develop a novel measure of market expectations of sensitivity to monetary policy surprises. A natural choice would be the elasticity (or beta) of firms' equity returns with respect to measures of monetary policy surprises. However, because FOMC announcements are infrequent (occurring eight times a year), estimates of beta are likely to be noisy and inaccurate, especially if the true level of sensitivity varies over time. To overcome this difficulty, we use information from the option-implied variance. Our intuition is that FOMC announcements resolve uncertainty about the macroeconomy and monetary policy and are associated with reductions in the option-implied variance. Firms that are more sensitive to monetary policy announcements should experience a greater implied variance reduction after announcements. Expectations for the implied variance reduction can therefore measure sensitivity to monetary policy announcements.

We find portfolios sorted on the expected implied variance reduction (EVR) yield a significant spread in average returns on FOMC announcement days but not on non-FOMC trading days. A long-short portfolio formed on our monetary policy sensitivity measure produces an average announcement-day return of 31.40 bps. In addition, the returns of EVR-sorted portfolios remain significant after controlling for market beta and other standard risk factors. To further demonstrate the spread on the EVR-sorted portfolios reflects risk compensation for monetary policy announcements, we use measures of monetary policy announcement surprises constructed by Nakamura and Steinsson (2018) to show the following (i) the average monetary policy announcement surprises are insignificantly different from zero, and therefore, rational expectations hold well in our sample period; and (ii) the returns of the EVR-sorted portfolios are monotonic in their sensitivity to monetary policy surprises.

Our findings suggest EVR is likely to proxy for an independent risk factor unknown to the literature given the previous work by Savor and Wilson (2014), who demonstrate that on announcement days, the CAPM holds well for market-beta-sorted portfolios, Fama-French size and book-to-market-sorted portfolios, and industry portfolios. Because market beta does not significantly change over an announcement day, time-varying beta is unlikely to explain the failure of the CAPM in accounting for the returns on EVR-sorted portfolios. The risk factor proxied by EVR must have an alpha that is significantly larger than that of other risk factors, such as size and book-to-market, to fail the CAPM on announcement days.

Motivated by the above findings, we provide evidence supportive of the "Fed information channel" documented by Nakamura and Steinsson (2018). That is, the FOMC announcement reveals the monetary authority's private information about its monetary policy target, which affects future economic growth through the conduct of monetary policy. We show EVR-sorted portfolios are significantly monotone in their exposure to measures of monetary policy surprises constructed by Nakamura and Steinsson (2018), which are shown to forecast economic growth. To corroborate the interpretation of the Fed information channel, we present two related pieces of evidence. First, we show the returns on the long-short portfolio sorted on EVR are positively correlated with revisions of forecasts of real gross domestic product (GDP) growth made by professional forecasters. In addition, we demonstrate EVR-sorted portfolios produce a similar return spread on other major macroeconomic announcement days, such as GDP and non-farm payroll announcements, which are likely to provide information about future economic growth.

To quantitatively account for the cross-sectional announcement returns, we develop a model in which FOMC announcement surprises require risk compensation because they reveal the Fed's private information about their interest-rate target, which affects the prospects for future economic growth. We also assume investors' preferences satisfy generalized risk sensitivity (Ai and Bansal, 2018). In our model, aggregate economic growth is driven both by productivity shocks and by the conduct of monetary policy, which is affected by shocks to the Fed's interest-rate target. The Fed has private information about its interest-rate target, which is revealed through periodic monetary policy announcements. We specify a cross section of stocks with dividend processes that differ in both their sensitivity to the Fed's interest-rate policy and their exposure to the publicly observable productivity shocks.

In our model, the size of the reduction in the implied variance on announcement days provides an accurate measure of stocks' risk exposure to surprises in monetary policy announcements. Sorting firms on the implied variance reduction is equivalent to sorting on sensitivity to news in FOMC announcements. The reason is that on non-announcement days, investors do not observe the true value of the interest-rate target and only update their beliefs about that value based on noisy signals contained in realized economic growth. Scheduled FOMC announcements reveal the Fed's interest-rate target, and as a result, investors' posterior beliefs jump on announcement days to incorporate the newly arrived information about the Fed's interest-rate target. Stocks that have a high exposure to policy surprises will experience a large drop in the implied variance on announcement days.

In this setup, our model matches several stylized features of the cross-sectional announcement returns. First, the average announcement-day excess return of the market is about 38.1 bps, and the spread on the announcement-day return of portfolios sorted on expected sensitivity is about 33.8 bps. Both are close to their empirical counterparts. Because of generalized risk sensitivity, announcement surprises carry news about the future prospects for the economy and are priced (Ai and Bansal, 2018). As a result, stocks that are more sensitive to monetary policy announcement surprises will receive a higher risk premium on announcement days.

Second, CAPM fails to account for the FOMC announcement returns of EVR-sorted portfolios in our model. In the model, the expected reduction in the implied variance accurately measures the sensitivity to monetary policy announcement surprises. The market beta, however, depends both on the sensitivity of the stock return to policy announcement surprises and, more importantly, on the sensitivity to the publicly observable productivity shocks, which account for a quantitatively larger fraction of variations in equity market valuations. EVR-sorted portfolios therefore exhibit a large dispersion in sensitivity to policy announcement surprises but a small dispersion in market beta, which is not enough to account for their announcement-day returns.

Third, even though the CAPM fails to account for the expected returns of EVR-sorted portfolios, it does explain the announcement-day returns of market-beta-sorted portfolios quite well. In the data, as documented by Savor and Wilson (2014), market-beta-sorted portfolios exhibit significant differences in their announcement premiums, which can be explained by the CAPM. In our model, market-beta-sorted portfolios exhibit a large dispersion in beta but a small dispersion in sensitivity to monetary policy surprises, because quantitatively, beta mostly reflects elasticity with respect to productivity shocks and is only weakly correlated with sensitivity to FOMC announcements. As a result, the announcement-day return of market-beta-sorted portfolios is mostly absorbed by differences in beta, making rejections of the CAPM difficult in a finite sample.

*1.0.0.0.1. Related literature.* Our paper is related to the literature that emphasizes the impact of monetary policy announcements on equity market returns. Bernanke and Kuttner (2005) demonstrate stock markets respond strongly to monetary policy announcements. Gürkaynak et al. (2005) document evidence that both monetary policy action and announcements have important impacts on asset markets.

Within this literature, most closely related to our paper are several recent papers emphasizing the impact of FOMC announcements on excess equity market returns. Lucca and Moench (2015) document an FOMC pre-announcement drift. Gorodnichenko and Weber (2016) show the volatility of stocks for firms with more price rigidity rises more than those for firms with more flexible prices after FOMC announcements. Neuhierl and Weber (2018) demonstrate the return drift around FOMC announcements depends on whether the monetary policy is expansionary or contradictory. Bollerslev et al. (2018) find that after the FOMC meetings, both volatility and volume increase, but the intra-day volume-volatility elasticity is systematically below unity. Cieslak et al. (2019) provide evidence for stock market returns over FOMC announcement cycles. Cieslak and Pang (2019) provide a decomposition of shocks that drive stock and bond market variations to explain stock and bond returns over the FOMC announcement cycle. Morse and Vissing-Jorgensen (2020) study the information transmission from the Fed to the stock market and its impact on the FOMC announcement premium. Whereas the above papers study the aggregate excess equity market returns around FOMC announcement days, our paper examines the heterogeneous impact of FOMC announcements on the cross section of the stock market.[1]

Our paper is also related to the broader literature on the macroeconomic announcement premium. Savor and Wilson (2013) document a significant equity market return on days with major macroeconomic announcements. Brusa et al. (2020) show the same holds for many other countries. Savor and Wilson (2014) demonstrate the CAPM holds well on macroeconomic announcement days but not on non-announcement days. Hu et al. (2019) provide evidence that the option-implied variance increases before announcements and drops afterward, and attribute the FOMC announcement premium to heightened stock market uncertainty. Amengual and Xiu (2018) argue the large declines in the option-implied variance after the FOMC announcements are associated with a resolution of policy uncertainty. Ernst et al. (2020) and Ghaderi and Seo (2020) study potential statistical bias in the estimation of the announcement premium. Engelberg et al. (2018) find

[1] Relatedly, Mueller et al. (2017) and Karnaukh (2018) study the impact of FOMC announcements on the foreign exchange market.

anomaly returns are higher on corporate news days and earnings announcement days. All of the above empirical evidence is broadly consistent with our equilibrium model in which announcements resolve macroeconomic uncertainty and are associated with reductions in the option-implied variance of equity market returns.

Our work is associated with papers that study monetary policy and the cross section of equity returns. Ozdagli and Velikov (2020) use observable firm characteristics to measure firms' exposure to monetary policy and find stocks with a more positive reaction to expansionary monetary policy surprises earn lower returns. Chava and Hsu (2020) find financially constrained firms earn lower returns than unconstrained firms after unanticipated increases in the federal funds rate target. The above papers study the monetary policy in general, but not necessarily monetary policy announcements. In fact, none of them focuses on returns realized on FOMC announcement days, nor do they find a significant premium realized following announcements.

Our theoretical model builds on recent developments in asset pricing models for the macroeconomic announcement premium. Ai and Bansal (2018) provide a revealed-preference theory for the macroeconomic announcement premium. Wachter and Zhu (2020) develop a rare-disaster-based model to explain the announcement premium and the robustness of CAPM on announcement days. Ai et al. (2018) study a production economy with macroeconomic announcements and its asset pricing implications. None of the above models addresses simultaneously the robustness of CAPM and the presence of additional risk factors on announcement days. The information channel we emphasize in our paper is consistent with recent work by Nakamura and Steinsson (2018), who provide empirical evidence and develop a theoretical model to show Fed's announcements affect beliefs not only about monetary policy, but also about economic fundamentals.

The rest of the paper is organized as follows. In Section 2, we develop a measure of expected sensitivity to monetary policy announcement surprises and present cross-sectional evidence for the relationship between expected sensitivity and expected returns. We provide evidence that the expected sensitivity to monetary policy announcements reflects firms' sensitivity to expectations about future economic growth in Section 3. In Section 4, we develop a continuous-time model with monetary policy announcements and explain cross-sectional equity returns. We present our quantitative analysis and demonstrate our model is able to account for the stylized features of the cross-sectional FOMC announcement premium in Section 5. Section 6 concludes.

## 2. Empirical evidence

In this section, we provide evidence that stocks that are more sensitive to monetary policy announcement surprises earn significantly higher premiums on FOMC announcement days. In addition, the CAPM cannot explain the cross section of the monetary policy announcement premium.

### 2.1. Measuring expected sensitivity

To study whether risk exposure to monetary policy announcements is priced in the cross section of equity returns, our strategy is to measure the sensitivity of firms' equity returns with respect to monetary policy announcements and to evaluate whether differences in the level of sensitivity are reflected in firms' expected returns realized on announcement days. This exercise requires constructing a firm-level measure of sensitivity with respect to FOMC announcement surprises, sorting stocks into portfolios based on such a measure, and estimating expected returns by computing the average returns of the sensitivity-sorted portfolios.

Because our purpose is to measure the expected returns of the sensitivity-sorted portfolios, the measure of sensitivity should be based on market expectations and cannot depend on information unavailable at the time of portfolio formation, to avoid any look-ahead bias. Because the FOMC makes announcements only eight times a year, any sensitivity estimates based on historical announcement data are likely to be noisy because of the lack of observations. In addition, if the true level of sensitivity is time varying, sensitivity estimates using historical announcements are likely to be inaccurate.

To overcome the above difficulty, our measure of sensitivity is based on the option-implied variance. In contrast to estimated sensitivity using historical announcement data, the option-implied variance is capable of capturing changes in market expectations in a timely manner. The construction of our measure is based on the intuition that FOMC announcements reduce uncertainty about the macroeconomy and monetary policy and are associated with reductions in the option-implied variance. In the cross section, firms that are more sensitive to monetary policy announcement surprises should experience higher implied variance reductions following announcements. To avoid look-ahead bias, we construct a measure of the EVR. In what follows, we first present evidence on the implied variance reductions on FOMC announcement days and then detail the construction of our measure of the EVR.

#### 2.1.1. Implied variance around monetary policy announcements

We first establish that significant reductions occur in the option-implied variance on FOMC announcement days both at the market level and at the firm level. We also show the firm-level implied variance reduction exhibits substantial heterogeneity. Collectively, the above evidence supports the two premises of our empirical exercises: (i) the implied variance reduction can be used to measure the firm-level sensitivity to monetary policy announcement surprises, and (ii) this sensitivity exhibits considerable heterogeneity across firms.

We use the squared option-implied volatility index, $VIX^2$, to measure the implied variance of the market return. We obtain data on VIX from the Chicago Board Options Exchange (CBOE). The CBOE's VIX is a model-free measure of implied variance computed from the S&P

**Table 1**
Implied variance around FOMC announcement days. This table reports the implied variance changes from one day before FOMC announcement to FOMC announcement days. In Panel A, we report changes in VIX$^2$ (monthly percentage squared units) around FOMC announcement days and their time-series statistics when testing whether the change is significantly different from zero. In Panel B, we report the cross-sectional average of changes in the firm-level option-implied variance (monthly percentage squared units) around FOMC announcement days and their time-series statistics. The firm-level implied variance uses the seven-day maturity variance. Our full sample period is from January 1996 to December 2017 with 176 FOMC announcement days. This period contains 6652 common stocks with traded options. Among these 6652 firms in our sample, 5446 firms have at least one observed option-implied variance on these 176 FOMC announcement days.

| | $VIX^2_{t-1}$ | $VIX^2_t$ | $VIX^2_{t-1} - VIX^2_t$ |
|---|---|---|---|
| Panel A: VIX$^2$ | | | |
| Mean | 41.940 | 39.526 | 2.414 |
| $t$-stats | | | (3.39) |
| Panel B: Average firm-level implied variance | | | |
| | $IV_{t-1}$ | $IV_t$ | $IV_{t-1} - IV_t$ |
| Mean | 281.165 | 276.210 | 4.954 |
| $t$-stats | | | (4.58) |

500 index option prices. For the firm-level implied variance, on each day and for each time to maturity, we follow Bakshi et al. (2003) to estimate the implied variance by averaging the weighted prices of out-of-money puts and out-of-money calls over a wide range of strike prices. We choose to use the seven-day implied variance because the short-maturity option prices are likely to be more sensitive to announcement surprises than the long-maturity option prices. Our firm-level option data are from OptionMetrics. The sample period is from January 1996 to December 2017. Our data period contains 176 pre-scheduled FOMC meetings. If an FOMC meeting lasts for two days, we treat the second day as the announcement day. We provide more detailed information about the firm-level implied variance and other data in Appendix A.

We document the patterns of the implied variance reduction on FOMC announcement days in Table 1. Panel A of the table compares the market-level implied variance ($VIX_t^2$) on FOMC announcement days with the same moment one day before announcements, $VIX^2_{t-1}$. Consistent with the previous literature (see, e.g., Savor and Wilson, 2013), VIX significantly decreases after FOMC announcements. On average, the daily reduction in VIX$^2$ is about 2.41 (monthly percentage squared units). We observe the same pattern at the firm level. As shown in Panel B of Table 1, the average reduction in the implied variance at the firm level is about 4.95 and significant.

Moreover, we find evidence of significant heterogeneity in announcement-day reductions in the implied variance across firms. We rank firms by their average announcement-day implied variance reduction and plot the histogram of these reductions in Fig. 1. The implied variance decreases after announcements for most firms, and the magnitude of the reduction differs substantially.

### 2.1.2. Expected variance reduction

Motivated by the above evidence, we measure the market-expected sensitivity to FOMC announcement surprises using the expected implied variance reduction. For an FOMC announcement day $t$, the expected implied variance reduction (EVR) is computed as[2]

$$EVR = IV_{t-2} - Median\ of\ Historical\ IV. \tag{1}$$

In the above equation, $IV_{t-2}$ is the seven-day implied variance ($IV$) for a firm computed from the closing price of its options two days before the FOMC announcement. We use closing prices two days before announcements to ensure our measure is not affected by the pre-FOMC announcement drift documented by Lucca and Moench (2015). The median of historical $IV$ in Eq. (1) is computed as the median value of seven-day $IV$ during day −15 and day −8, where the announcement day is normalized to day 0.[3] In our construction, the historical $IV$ is not affected by the upcoming FOMC announcement, whereas $IV_{t-2}$ is. A larger increase in $IV_{t-2}$ relative to its historical level (or a higher EVR) indicates the market expectation that the equity of the firm will be more sensitive to the upcoming FOMC announcement, and therefore, $IV$ will drop more after the announcement. In all empirical analyses, we use a demeaned EVR so that the intercept of a univariate regression with EVR can be interpreted as returns.

Our measure of the expected implied variance reduction has significant predictive powers for the actual implied variance reduction on FOMC announcement days. In Eq. (2), we report results from a panel regression of the actual implied variance reduction on our measure of the expected implied variance reduction. The regression coefficient on the expected variance reduction is significantly positive with a $t$-statistic of 9.57 (based on the day-

[2] Note our measure is different from the variance risk premium (VRP) in the literature; see, for example, Bollerslev et al. (2009), Drechsler and Yaron (2011), and Bollerslev et al. (2014). The VRP is the difference between the variance under the risk-neutral measure and that under the physical measure, whereas our measure is designed to capture the upcoming announcements, and we do not use any variance measure under the physical probability.

[3] Our results remain robust if we use the mean value of the historical $IV$ during the same period to adjust $IV_{t-2}$ or if we use a longer period to measure the historical $IV$, such as during day −22 and day −8.

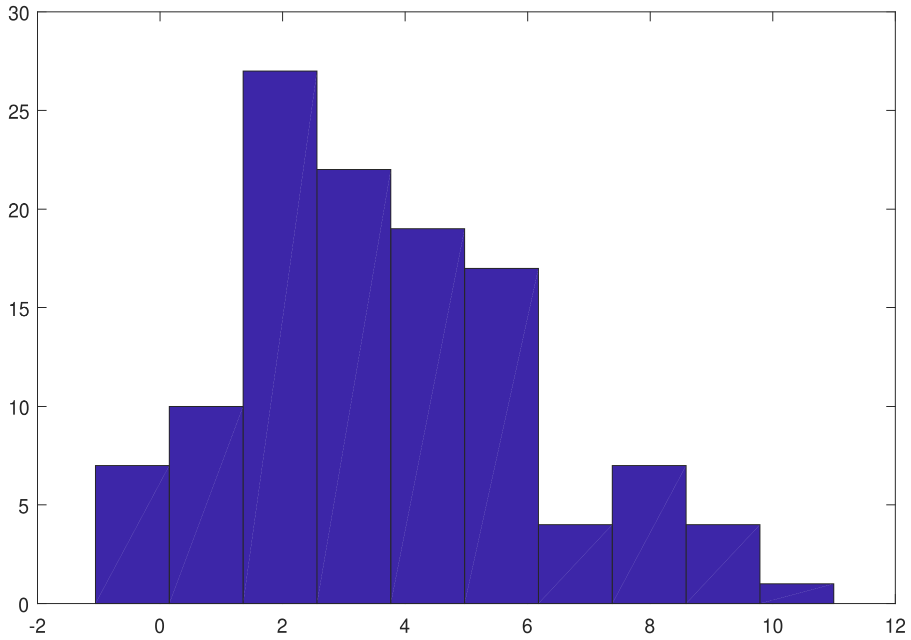


**Fig. 1.** Heterogeneous firm-level implied variance reduction around FOMC days. This figure plots the histogram of the time-series mean of the firm-level logarithm implied variance reduction (in %) around FOMC announcement days. For illustration purposes, we only report those firms with 160 or more observations out of the 176 FOMC announcement days in our data sample.

clustered standard error) and an $R^2$ of 6.3%[4]:

$$Actual\ IV\ Reduction = 0.0021 + \underset{(9.57)}{0.2164} \times EVR. \qquad (2)$$

Actual *IV* reduction is defined as $IV_{t-2} - IV_t$. The key advantage of using the implied variance to measure expected sensitivity is that the implied variance reflects market expectations in a timely manner. Alternative measures such as market beta on previous announcement days do not incorporate the same forward-looking information contained in the implied variance and may not efficiently measure market expectations on the upcoming announcement.

### *2.2. Portfolios sorted on expected sensitivity*

Using the EVR measure constructed above, we sort stocks into decile portfolios and examine their returns on FOMC announcement days. Consider an FOMC announcement day. We compute our EVR measure for each stock two days ahead of the announcement day and sort stocks into decile portfolios using the EVR measure. Stocks in the top (bottom) portfolio have the highest (lowest) EVR and are most (least) sensitive to the upcoming FOMC announcement. We keep track of the value-weighted portfolio returns and rebalance the portfolios until the next sorting date, which is two days before the next FOMC announcement day. We repeat the above procedure and compute the average FOMC announcement-day return and non-FOMC announcement-day return for each of the portfolios.[5]

Panel A of Table 2 reports the portfolio returns on FOMC announcement days and non-FOMC announcement days. To save space, we present the returns for portfolios 1, 2, 9, and 10, and the return for the consolidated portfolios 3 to 8. On FOMC announcement days, the top-decile portfolio with the highest expected sensitivity earns higher expected returns than the bottom decile portfolio with the lowest expected sensitivity. On average, the long-short portfolio earns a raw return of 31.40 bps on FOMC announcement days. This tradable strategy that invests only on the eight FOMC announcement days earns an average annual return of 2.51% (31.40 bps $\times$ 8). We also report alphas from the three-factor model of Fama and French (1992), denoted as FF3, and those from the five-factor model of Fama and French (2015), denoted as FF5. The risk-adjusted performance measured by these alphas monotonically increases from the bottom decile to the top decile. These alphas confirm that returns of the portfolios sorted on expected sensitivity on FOMC announcement days are not driven by CAPM beta and standard firm characteristics such as size, book-to-market ratio (B/M), operating profitability, and conservative investment. On average, after controlling for the standard risk factors, firms with higher expected sensitivity still earn a higher FOMC announcement premium. On announcement days, the long-short portfolio has a significant annual alpha of 2.28% or 2.19% (28.48 bps or 27.33 bps $\times$ 8), respectively. By contrast, most portfolios, as well as the long-short portfolio,

[4] In Online Appendix IA, we show EVR outperforms many other models in predicting the actual implied variance reduction on FOMC announcement days.

[5] Non-FOMC announcement days refer to any trading days without an FOMC announcement.

**Table 2**
Portfolios sorted on expected sensitivity. We conduct a tradable strategy two days before the FOMC announcement days. We measure firm-level sensitivity to monetary policy announcements by the expected implied variance reduction (EVR) around FOMC announcements, which is the difference between the implied variance two days before the announcement (normalized to day 0) and the median value of the implied variance during day −8 and day −15. We demean EVR, sort firms based on this measure into decile portfolios with the third portfolio containing 60% of all firms and each of the remaining four portfolios containing 10%, and document the value-weighted portfolio returns. These portfolios are rebalanced two days before the next FOMC announcement day. Panel A reports the average value-weighted portfolio EVR (demeaned, monthly percentage squared units), the time-series average and Newey-West *t*-statistics with 12 lags of portfolio returns, and alphas from Fama-French three-factor and five-factor models (bps) on FOMC announcement days and non-FOMC announcement days. Panel B reports the daily regression results of Eq. (3), which include the coefficients of market excess return (i.e., CAPM beta), the interaction of market excess return and the FOMC dummy, the non-FOMC dummy, and the FOMC dummy, as well as the Newey-West *t*-statistics.

| | 1 | 2 | 3 – 8 | 9 | 10 | (10 – 1) |
|---|---|---|---|---|---|---|
| Panel A: Average returns | | | | | | |
| EVR | -192.76 | -55.14 | -0.63 | 51.83 | 206.37 | |
| FOMC Return | 35.95 | 28.56 | 29.01 | 47.04 | 67.35 | 31.40 |
| | (1.92) | (2.35) | (3.33) | (3.45) | (3.06) | (2.67) |
| FF3 $\alpha$ | -12.49 | -7.85 | -0.44 | 9.73 | 15.98 | 28.48 |
| | (-1.57) | (-1.98) | (-0.26) | (2.22) | (2.04) | (2.83) |
| FF5 $\alpha$ | -13.88 | -8.46 | 0.19 | 9.24 | 13.46 | 27.33 |
| | (-2.17) | (-2.24) | (0.16) | (2.20) | (1.96) | (2.76) |
| Non-FOMC Return | 3.48 | 3.36 | 3.60 | 4.30 | 3.29 | -0.19 |
| | (1.44) | (1.73) | (2.65) | (2.36) | (1.34) | (-0.15) |
| | | | Panel B: CAPM | | | |
| | 1 | 2 | 3 – 8 | 9 | 10 | (10 – 1) |
| $R^M - r_f$ | 1.40 | 1.15 | 0.92 | 1.22 | 1.44 | 0.03 |
| $(R^M - r_f)\times$ FOMC Dummy | 0.08 | -0.03 | -0.02 | -0.07 | 0.14 | 0.05 |
| | (0.70) | (-0.51) | (-1.05) | (-1.39) | (1.02) | (0.25) |
| Non-FOMC Dummy | -0.70 | -0.22 | 0.56 | 0.55 | -0.97 | -0.27 |
| | (-0.60) | (-0.25) | (1.96) | (0.74) | (-0.81) | (-0.20) |
| FOMC Dummy | -12.43 | -8.08 | -0.64 | 9.38 | 16.23 | 28.66 |
| | (-1.73) | (-1.64) | (-0.40) | (2.06) | (2.03) | (3.14) |

do not earn significant returns on non-FOMC announcement days.

We next report results from CAPM regressions for our sorted portfolios. To distinguish between FOMC announcement days and non-FOMC announcement days, we consider the following regression:

$$R_t^i - r_{f,t} = \alpha_{Non}^i \cdot \mathbf{1}_{Non} + \alpha_{FOMC}^i \cdot \mathbf{1}_{FOMC} + \beta^i \left(R_t^M - r_{f,t}\right) + \beta_{FOMC}^i \left(R_t^M - r_{f,t}\right) \cdot \mathbf{1}_{FOMC} + \varepsilon_t^i, \quad (3)$$

where $R_t^i$ is the daily return of the sorted portfolio $i$, $R_t^M$ is the daily return of the market, and $r_{f,t}$ is the daily risk-free rate. The variables $\mathbf{1}_{Non}$ and $\mathbf{1}_{FOMC}$ are dummy variables that take values of 1 only on non-FOMC and FOMC announcement days, respectively. Panel B of Table 2 reports results for the above CAPM regression. Note the FOMC dummy is monotonically increasing across portfolios, and the FOMC dummy of the long-short portfolio is positive and statistically significant. The spread between the high-sensitivity portfolio and the low-sensitivity portfolio averages 28.66 bps on FOMC announcement days after controlling for market returns. The above portfolio return results remain robust if we exclude the height of the recent financial crisis period (July 2008 to June 2009) or if we exclude firms whose earnings announcement dates coincide with the FOMC announcement days, as shown in Online Appendix IA.

In Eq. (3), $\beta_{FOMC}^i$ measures the changes of the market betas of EVR-sorted portfolios on FOMC days relative to that on non-FOMC days. A positive $\beta_{FOMC}^i$ indicates a higher market beta of the EVR-sorted portfolio on FOMC days. In Panel B of Table 2, we show the market betas of the EVR-sorted portfolios do not change significantly on FOMC days. This result implies the spread on EVR-sorted portfolios on FOMC announcement days cannot be due to a higher beta on FOMC days and must come from a higher market price of risk captured by the EVR measure on FOMC announcement days.

### 2.3. Beta-sorted portfolios

The literature shows the CAPM holds well on macroeconomic announcement days (see, e.g., Savor and Wilson, 2014 and Wachter and Zhu, 2020). Two reasons account for the robustness of the CAPM on announcement days: (i) Portfolio betas are well estimated and are unlikely to change during a very short period, such as one announcement day; and (ii) variations in alphas, if any, are a lot smaller compared with the announcement-day premium.

Consider the projection of the excess return of portfolio $i$, $R^i - r_f$, onto the market excess return:

$$R_t^i - r_{f,t} = \alpha^i + \beta^i \left(R_t^M - r_{f,t}\right) + \varepsilon_t^i.$$

Taking expectations on both sides, $\mathbb{E}\left[R^i - r_f\right] = \alpha^i + \beta^i \mathbb{E}\left[R^M - r_f\right]$. Because the average market return $\mathbb{E}\left[R^M - r_f\right]$ on FOMC announcement days is known to be large (roughly 40 bps per day), and because $\beta^i$ is unlikely to change at a daily frequency, for any factor other than market beta to predict returns in the cross section, it must produce a large and significant $\alpha^i$ compared with the market announcement-day premium. This significant $\alpha^i$ on announcement days imposes a high standard for any other factors to predict announcement-day returns in the cross section. In fact, Savor and Wilson (2014) show

**Table 3**
Portfolios sorted on CAPM beta. This table reports the CAMP beta-sorted portfolio results. We first compute the beta exposure of each firm based on the past 12-month daily returns two days before the FOMC announcement. We sort firms into five portfolios with the third portfolio containing 60% of the firms and each of the remaining four portfolios containing 10%, and document the value-weighted portfolio returns. These portfolios are rebalanced two days before the next FOMC announcement day. Panel A reports the time-series average and Newey-West *t*-statistics (with 12 lags) of the portfolio returns (bps) on FOMC announcement days and non-FOMC announcement days. Panel B reports daily regression results including the coefficients of market excess return (i.e., CAPM beta), the interaction of market excess return and the FOMC dummy, the non-FOMC dummy, and the FOMC dummy, as well as the Newey-West *t*-statistics.

| | 1 | 2 | 3 – 8 | 9 | 10 | (10 – 1) |
|---|---|---|---|---|---|---|
| Panel A: Average returns | | | | | | |
| FOMC Return | 13.05 | 17.40 | 30.66 | 51.74 | 68.23 | 55.17 |
| | (2.27) | (2.98) | (2.89) | (2.55) | (2.70) | (2.41) |
| Non-FOMC Return | 3.09 | 3.63 | 3.57 | 2.61 | 4.05 | 0.95 |
| | (3.25) | (3.10) | (2.30) | (1.02) | (1.33) | (0.34) |
| Panel B: CAPM | | | | | | |
| | 1 | 2 | 3 – 8 | 9 | 10 | (10 – 1) |
| $R^M - r_f$ | 0.47 | 0.68 | 1.03 | 1.57 | 1.73 | 1.25 |
| $(R^M - r_f)\times$ FOMC Dummy | 0.04 | -0.01 | -0.03 | 0.02 | 0.25 | 0.21 |
| | (1.12) | (-0.32) | (-1.31) | (0.30) | (2.55) | (1.72) |
| Non-FOMC Dummy | 1.14 | 1.19 | 0.32 | -1.92 | -0.84 | -1.97 |
| | (1.73) | (1.89) | (0.88) | (-1.68) | (-0.49) | (-0.93) |
| FOMC Dummy | -3.97 | -4.57 | -1.85 | 0.35 | 4.85 | 8.82 |
| | (-0.92) | (-1.34) | (-1.18) | (0.06) | (0.56) | (0.75) |

known factors are not powerful enough to produce a significant $\alpha^i$ on announcement days.[6] The robustness of the CAPM relationship on announcement days also implies that empirically, if we find an instance of a failure of the CAPM, it is likely due to a truly powerful risk factor other than the market return, and it is unlikely due to time-varying betas. Any structural model that explains the failure of the CAPM on announcement days should have at least two independent risk factors.

In what follows, we reproduce the FOMC announcement premium for CAPM beta-sorted portfolios documented in the literature by using the same sample of firms as in the EVR-sorting procedure. This sorting result serves as an important discipline for the structural model we develop in Section 4. Our sorting procedure is similar to the one using the expected sensitivity measure, EVR, introduced in Section 2.2. For each FOMC meeting, we sort all stocks into decile portfolios based on their CAPM beta, which is calculated using the daily return during the past 12 months, two days before the FOMC announcement. We then document the daily portfolio returns until the next rebalancing date, which is two days before the next FOMC announcement.

Panel A of Table 3 reports the portfolio returns on FOMC announcement days and non-FOMC announcement days. We find that, consistent with Savor and Wilson (2014), on FOMC days, high-beta stocks do earn higher returns: on average, the long-short portfolio generates a raw return of 55.17 bps. On the other hand, portfolio returns on non-FOMC days are rather low with no obvious pattern.

In Panel B, we report the results from the CAPM regression, Eq. (3), for beta-sorted portfolios, where we include an FOMC dummy, a non-FOMC dummy, market excess returns, and the interaction between the FOMC dummy and market excess returns. By construction, beta monotonically increases across the portfolios. The coefficients on the FOMC dummy are insignificant for all portfolios, including the long-short portfolio. The result that the returns on the beta-sorted portfolios can be explained by the CAPM is consistent with the previous literature, for example, Savor and Wilson (2014). In addition, we note the coefficients of the interaction term between the FOMC dummy and market excess returns are not significant for most of the portfolios, indicating the CAPM beta does not change significantly on FOMC days except for the top-decile portfolio. The stability of betas across announcement days and non-announcement days for EVR- and beta-sorted portfolios provides the key motivation for our theoretical model in Section 4, where the cross section of expected returns are due to heightened volatility of the SDF on announcement days rather than differences in betas across announcement and non-announcement days.

### 2.4. Robustness of findings

In this subsection, we assess the robustness of the univariate portfolio sorts on expected sensitivity to FOMC announcements and CAPM beta through the risk-price estimations. We first follow the standard procedure in Fama and MacBeth (1973). To gain estimation efficiency, we run cross-sectional regressions using individual firms on each FOMC announcement day. Our choice of individual firms rather than portfolios sets a higher bar for significant results; see discussions in Ang et al. (2020) and Gagliardini et al. (2016). In the first stage, we compute the beta of each firm using the standard risk factors and the past 12-month daily excess returns two days before each FOMC announcement. In the second stage, we regress the cross-sectional stock excess returns on FOMC announcement days on the firm-level EVR and those betas of the standard risk factors. In all our regression specifications,

[6] See also Harvey et al. (2016) and Harvey and Liu (2020) for a critique of the fragility of risk factors in the cross section of equity returns in general.

**Table 4**
Daily excess return of individual stocks. The table reports the regression results of daily stock excess returns. For the Fama-MacBeth regression in Panel A, we first compute the beta of each firm based on the past 12-month daily excess returns two days before each FOMC announcement. Then, on each FOMC announcement day, we regress the cross-sectional stock excess returns on the estimated EVR and the beta from the first step. We report the average loadings on EVR and other factors' betas and their associated Newey-West *t*-statistics. In each cross-sectional regression, we remove the influential points and winsorize stock returns at the 1st and 99th percentiles. We demean EVR in cross-sectional regressions for interpretation convenience. For the pooled regression in Panel B, we report the coefficients in Eq. (4). *t*-statistics are calculated using the trading-day clustered standard errors. The unit of regression coefficients is in percentage. Our full sample period is from January 1996 to December 2017 with 176 FOMC days.

| | | (1) | (2) | (3) | (4) |
|---|---|---|---|---|---|
| Panel A: Fama-MacBeth regressions on FOMC announcement days | | | | | |
| | Constant | 0.40 | 0.01 | 0.02 | 0.05 |
| | | (3.11) | (0.13) | (0.40) | (1.19) |
| | EVR | 0.20 | 0.19 | 0.20 | 0.21 |
| | | (2.52) | (3.09) | (3.46) | (3.58) |
| | Market $\hat{\beta}$ | | 0.35 | 0.33 | 0.29 |
| | | | (2.68) | (2.69) | (2.65) |
| | SMB $\hat{\beta}$ | | | 0.03 | 0.04 |
| | | | | (1.01) | (1.16) |
| | HML $\hat{\beta}$ | | | 0.06 | 0.05 |
| | | | | (0.95) | (0.84) |
| | RMW $\hat{\beta}$ | | | | -0.05 |
| | | | | | (-1.22) |
| | CMA $\hat{\beta}$ | | | | -0.01 |
| | | | | | (-0.37) |
| $R^2$(%) | 0.49 | 5.07 | 8.94 | 10.30 | |

Panel B: Pooled regression

| $1_{Non}$ | $1_{FOMC}$ | $EVR$ | $EVR \cdot \mathbf{1}_{FOMC}$ | $\hat{\beta}$ | $\hat{\beta} \cdot \mathbf{1}_{FOMC}$ | $R^2$(%) |
|---|---|---|---|---|---|---|
| 0.042 | 0.296 | 0.055 | 0.146 | -0.017 | 0.120 | 0.10 |
| (1.80) | (2.65) | (2.51) | (1.43) | (-0.79) | (1.20) | |

we use the demeaned EVR along the cross-sectional dimension so that the intercept of a univariate regression on EVR can be interpreted as the average announcement-day risk premium.

Panel A of Table 4 reports the time-series average of the risk prices from Fama-MacBeth regressions on FOMC announcement days. The first regression has only EVR as the explanatory variable. The constant of the regression, which can be interpreted as the average risk premium on FOMC announcement days, is large and significant: about 40 bps. The estimated average risk price of EVR is 20 bps with a *t*-statistic equal to 2.52. In the second specification, we use both EVR and the market beta in a joint regression. The risk prices for EVR and market beta are 19 bps and 35 bps with *t*-statistics equal to 3.09 and 2.68, respectively. Echoing our results in the last two subsections, both EVR and the market factor carry a significant risk premium.

In columns (3) and (4), we present Fama-MacBeth regressions on FOMC announcement days, including the rest of the Fama and French (1992) three factors and Fama and French (2015) five factors. Our estimate of the risk price of the EVR factor remains significant and robust in magnitude in the presence of the market factor, whereas none of the other factors have a significant risk price. This finding confirms our previous results that EVR captures an independent risk factor other than the market in explaining the cross section of returns on announcement days, whereas none of the other factors (except for the market factor) carries a significant risk premium on announcement days.

Next, to test whether the risk price of the EVR factor changes between FOMC days and non-FOMC days, we run the following pooled regression in addition to the Fama-MacBeth regression:

$$R_t^j - r_{f,t} = \gamma_0 \cdot \mathbf{1}_{Non} + \gamma_{0,FOMC} \cdot \mathbf{1}_{FOMC} + \gamma_1 \cdot EVR_t^j + \gamma_2 \cdot EVR_t^j \cdot \mathbf{1}_{FOMC} + \delta_1 \cdot \widehat{\beta}_t^j + \delta_2 \cdot \widehat{\beta}_t^j \cdot \mathbf{1}_{FOMC} + \varepsilon_t^j, \tag{4}$$

where the dependent variable is firm *j's* excess return on date *t*. The variables $EVR_t^j$ and $\widehat{\beta}_t^j$ are estimated two days before date *t* for firm *j*. Panel B of Table 4 reports the regression results of Eq. (4). Standard errors are clustered at the daily level. After controlling for the market beta and its interaction term with the FOMC dummy, the non-FOMC slope coefficient for EVR equals 5.5 bps and is statistically significant. The predictability of EVR for stock returns on non-FOMC announcement days is likely due to the presence of other macroeconomic announcements and other information events that resolve uncertainty about the macroeconomy, which is consistent with our empirical evidence for other macroeconomic announcements in Section 3.3 and the model we present in Section 4. Our results also suggest that the risk price of EVR is higher on FOMC days. The point estimate for $\gamma_2$, the difference between the risk price of EVR on FOMC days and that on non-FOMC days, is 14.6 bps.[7] The fact that the estimated risk price of EVR is much higher

[7] In an unreported alternative pooled regression analysis that directly calculates the risk price of EVR on FOMC days and non-FOMC days, we

on FOMC days is likely because the risk premium realized on FOMC announcement days is larger in magnitude than that on other macroeconomic announcement days.

## 3. Inspecting the mechanism

In this section, we present several pieces of empirical evidence that are consistent with the hypothesis that EVR measures firms' sensitivity to economic growth-rate expectations and the EVR-sorted portfolios produce premiums through the "Fed information channel" (Nakamura and Steinsson, 2018). That is, FOMC announcements change the private sector's growth-rate expectations, and high-EVR stocks are riskier because they are more sensitive to Fed-policy-induced changes in expectations about future economic growth.

Given the difficulty in identifying the exact economic mechanism, due to the data limitation, our purpose is to provide some indirect evidence consistent with the Fed information channel and to motivate the assumptions of our structural model. To this end, we establish three empirical facts. First, we show the EVR-sorted portfolios are monotone in their betas to monetary policy surprises. In particular, the returns on the high-EVR portfolio tend to be positive upon a surprise hike in interest rates and negative upon a surprise cut in interest rates (after controlling for the market excess returns). This finding is consistent with the Fed information channel emphasized by Nakamura and Steinsson (2018): a surprise increase in interest rates signals the Fed's private information about higher economic growth and is associated with positive revisions of private sectors' forecasts of future economic growth.

Second, consistent with the Fed information channel, we show that the returns on the long-short portfolio are positively correlated with monetary policy surprises, as well as positively correlated with innovations in growth-rate expectations measured by forecast revisions of the real GDP growth from the Survey of Professional Forecasters.

Third, to corroborate our hypothesis of the Fed information channel, we show EVR is also priced similarly in the cross section of stocks on other macroeconomic announcements that are informative about future economic growth.

Finally, to distinguish our theory from models in which time-varying risk premium comes from shocks to the volatility of macroeconomic fundamentals, we show several volatility-related risk factors, such as news about future market variance (Campbell et al., 2018), co-skewness (Harvey and Siddique, 2000), and co-kurtosis (Dittmar, 2002), cannot subsume the risk price of EVR, especially on FOMC announcement days.

### 3.1. Sensitivity to FOMC announcement surprises

Nakamura and Steinsson (2018) construct measures of monetary policy surprises and show these measures capture the Fed's private information about future economic growth. In this section, we use two measures of monetary policy surprises constructed by Nakamura and Steinsson (2018) to examine the exposure of EVR-sorted portfolios to monetary policy surprises. The first measure is a composite measure constructed as the first principal component of the unanticipated change over the 30-minute FOMC announcement windows in a basket of five interest rates. The second is based on changes in the federal funds rate (FFR) on FOMC announcement days.

First, we show rational expectations hold well in the period of our portfolio-sorting exercise. In Table 5, we report the first and second moments of the two measures of monetary policy surprises: the composite measure of policy news constructed in Nakamura and Steinsson (2018) (labeled as Policy News) and the changes in the FFR on announcement days. Both measures of monetary policy surprises exhibit substantial variations; however, in both measures, the average surprises are not significantly different from zero. This result implies the absence of any systematic biases in the market's forecast about monetary policy announcements during this period, and the market excess return and that of the portfolios sorted on expected sensitivity EVR must be compensation for risk rather than a reflection of biases in expectations.

Second, we show the cross-sectional returns of portfolios sorted on EVR monotonically react to policy surprises. We compute the betas of portfolios sorted on EVR with respect to both measures of monetary policy surprises. We run the following regression on FOMC announcement days:

$$R_t^i - r_{f,t} = \alpha^i + \beta_{News}^i \Delta g_t + \beta_{Mkt}^i \left(R_t^M - r_{f,t}\right) + \varepsilon_t^i, \qquad (5)$$

where $R_t^i$ is the return of portfolio $i$, and $\Delta g_t$ stands for policy surprises on FOMC announcement-day $t$. We plot the slope coefficient of the above regression, $\beta_{News}^i$, and its 95% confidence band for each of the portfolios in Fig. 2.

The slope coefficient $\beta_{News}^i$ is significantly monotonically increasing for both measures of policy surprises. After an unexpected interest-rate hike, the return for the high-EVR portfolio increases and that for the low-EVR portfolio decreases, after controlling for market excess returns. The fact that the high-EVR portfolio is more sensitive to interest-rate hikes is consistent with Nakamura and Steinsson's (2018) interpretation that monetary policy announcements convey the Fed's private information about the economy, and surprising interest-rate hikes are associated with information that economic fundamentals are stronger than expected. Consistent with the above information effect, Nakamura and Steinsson (2018) document that in response to an unexpected increase in the real interest rate (a monetary tightening), survey estimates of expected output growth rise.

### 3.2. Long-short portfolio

To further corroborate the finding that the return spreads on EVR-sorted portfolios are compensation for risk exposure to Fed's private information, in this subsection, we examine the returns and characteristics of the long-short portfolio. In Fig. 3, we plot the time series of the compounded portfolio growth if we invest $1 in the high-

find the FOMC-day risk price of EVR is 20.1 bps, both positive and significant.

**Table 5**
Monetary policy shocks. This table reports the first and second moments and the time-series $t$-statistics when testing whether the change is significantly different from zero for monetary policy news shocks (Policy News) and federal funds rate shocks (FFR) from Nakamura and Steinsson (2018) and Acosta and Saia (2020). The data sample spans from 1996 to 2019. The policy news shock is a composite measure constructed as the first principal component of the unanticipated change in five interest rates over a 30-minute window around the FOMC announcement. The FFR shock is constructed using the price change of the federal funds futures over a 30-min window around the FOMC announcement.

| | Policy News | FFR |
|---|---|---|
| Mean | 0.0031 | -0.0043 |
| std | 0.0320 | 0.0400 |
| $t$-stat | (1.32) | (-1.49) |

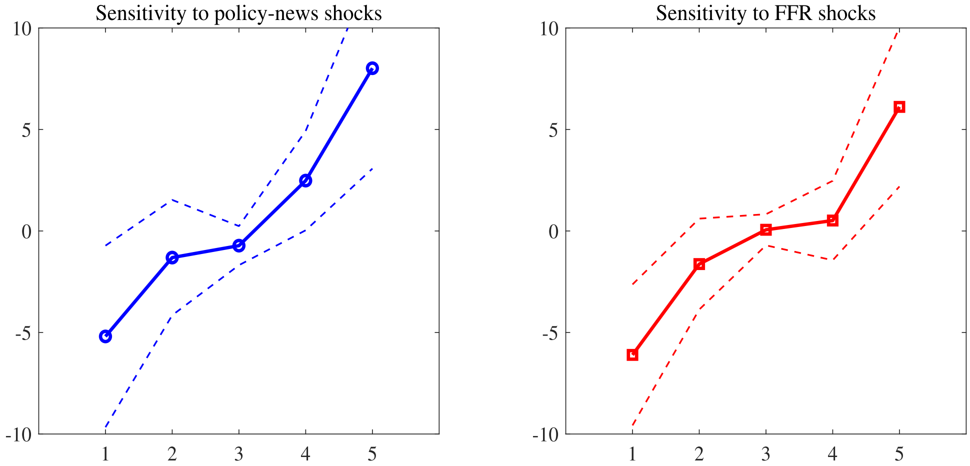


**Fig. 2.** Portfolio return sensitivity to monetary policy shocks. This figure plots the return sensitivity to monetary policy shocks for EVR-sorted portfolios. We regress the portfolios' excess returns on monetary-policy-news shocks or federal funds rate (FFR) shocks on FOMC announcement days, controlling for the market excess returns. The left (right) panel plots the coefficients of monetary policy news shocks (FFR shocks) and their corresponding 95% confidence interval. Monetary-policy-news shocks and FFR shocks are from Nakamura and Steinsson (2018) and Acosta and Saia (2020).

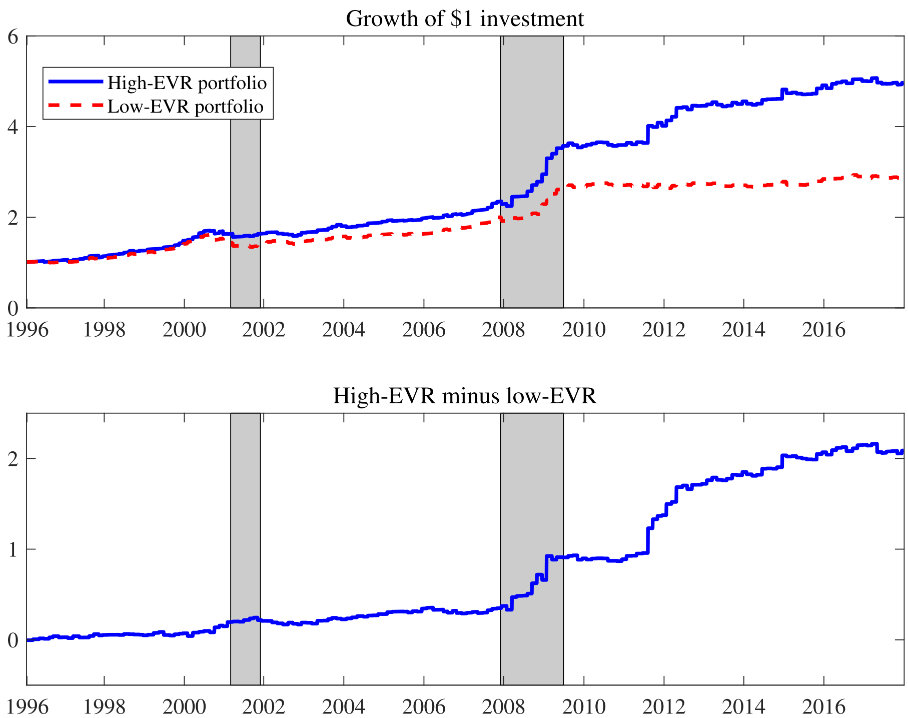


**Fig. 3.** High-EVR and low-EVR portfolio performance. The upper panel of this figure presents the time series of portfolio growth if we invest $1 in the high-EVR (solid line) and low-EVR (dashed line) portfolios on each FOMC announcement day and invest in the money market on non-FOMC announcement days. The lower panel plots the difference in the performance of the high-EVR and the low-EVR portfolios. Recessions defined by NBER are shaded in dark gray. The sample period is from 1996 to 2017.

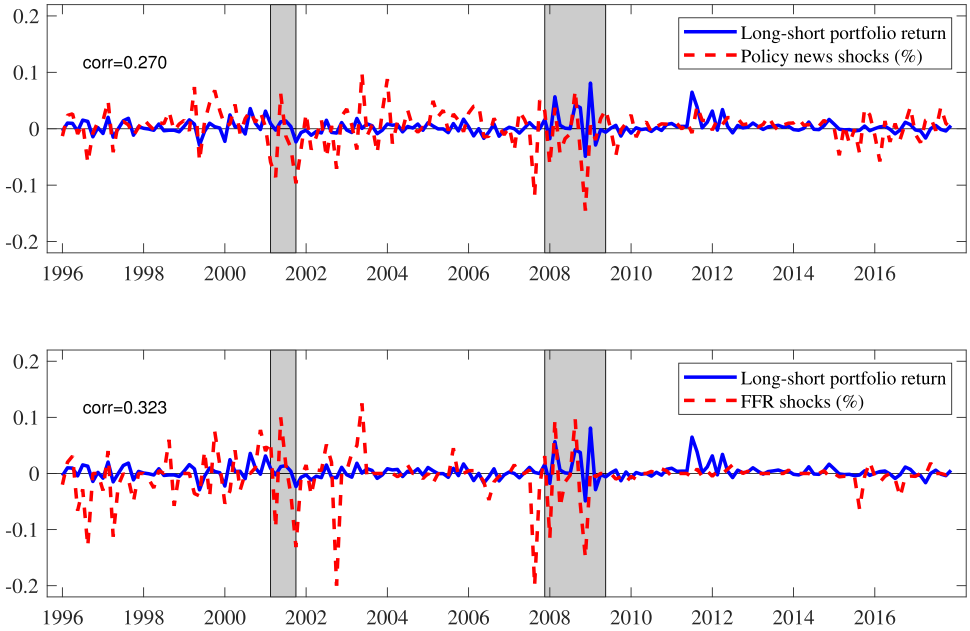


**Fig. 4.** Long-short portfolio return and monetary policy shocks. This figure plots the time series of the EVR-sorted long-short portfolio returns on FOMC announcement days (solid line) and monetary policy news shocks and federal funds rate (FFR) shocks (dashed lines) from 1996 to 2017. Shaded areas represent NBER recessions. Monetary policy news shocks and FFR shocks are from Nakamura and Steinsson (2018) and Acosta and Saia (2020).

and low-EVR portfolios on each FOMC announcement day and invest in the money market with the risk-free rate on non-FOMC announcement days. NBER-dated recessions are shaded. The lower panel plots the performance difference between the high- and low-EVR portfolios. The high-EVR stocks perform better than the low-EVR stocks during the recessions, especially during the 2008–2009 recession, and from 2012 going forward. In Fig. 4, we plot the time series of the long-short portfolio returns and monetary policy shocks on FOMC days, where we use two measures of monetary policy shocks as in the last section. Not surprisingly, the long-short portfolio return is positively correlated with policy news shocks and FFR shocks. The long-short portfolio return has a correlation of 27% ($p$-value $=$ 0.0003) with policy-news shocks, and a correlation of 32% ($p$-value $=$ 0.0000) with FFR shocks. In a regression of the long-short portfolio returns on monetary policy shocks, we find a 1 bp unexpected interest-rate hike is associated with an average return of 11 bps for the long-short portfolio on FOMC days with the $t$-statistic of 3.26.

We next show the returns on the long-short portfolio are correlated with innovations in investors' beliefs about the growth prospect of the economy. This finding is further evidence for the Fed information channel: if the excess return on the long-short portfolio is risk compensation for Fed's private information released upon announcements, it should comove with investors' belief revisions after FOMC announcements. We use the quarterly real GDP growth forecasts from the Survey of Professional Forecasters to measure investors' beliefs about future economic growth. The innovation in expected real GDP growth is calculated as the revision of the GDP forecast immediately following an FOMC announcement relative to the previous forecast of the same quarter GDP. In our sample, the long-short portfolio returns have a correlation of 23.9% ($p$-value $=$ 0.0251) with the forecast revisions of the GDP growth.

The above evidence shows high-EVR stocks are more sensitive to news contained in monetary policy announcements. We find firms in the high-EVR portfolio have a higher leverage ratio and are less financially constrained based on the Whited and Wu (2006) index than firms in the low-EVR portfolio. Intuitively, high-leverage firms are more sensitive to interest-rate policy changes. In addition, less financially constrained firms are likely to rely more on external financing and therefore are more sensitive to interest-rate news. These findings are consistent with Chava and Hsu (2020). Further, we find the high-EVR portfolio has more firms from the banking, pharmaceutical, trading, and insurance industries, whereas the low-EVR portfolio contains more firms from the computers, business service, utilities, aircraft, and computer software industries. Presumably, firms from the financial industry are more sensitive to interest-rate news, and firms in the utility and aircraft industries produce stable cash flows and are less sensitive to monetary policy news by comparison.

### *3.3. Other macroeconomic announcements*

To further corroborate the hypothesis of the Fed information channel, here, we show EVR carries a significant risk price on other macroeconomic announcement days that are likely to provide information about future economic growth.

As shown in Savor and Wilson (2013) and Ai and Bansal (2018), both GDP and non-farm payroll (NFP) announcements are associated with significant

**Table 6**
Fama MacBeth regressions on GDP and NFP announcement days. The table reports the results from Fama-MacBeth regressions. We first compute the beta of each firm based on the past 12-month daily excess returns one day before each GDP Advance (the first release) or Non-Farm Payrolls (NFP) announcement day (non-overlapping with FOMC announcement days). We then construct the EVR measure one day before each announcement day in the same manner as we did for FOMC announcement days. Next, on each GDP Advance or NFP announcement day (non-overlapping with FOMC days), we regress the cross-sectional stock excess returns on EVR and the beta from the first step. We report the average loadings on EVR and other factor betas and their associated Newey-West *t*-statistics. The unit of regression coefficients is in percentage. In each cross-sectional regression, we remove the influential points and winsorize stock returns at the 1st and 99th percentiles. We demean EVR in cross-sectional regressions for interpretation convenience. Our full sample period is from January 1996 to December 2017 with 76 GDP Advance days and 264 NFP announcement days (non-overlapping with FOMC days). In our data period, seven NFP announcements are released on non-trading days. In this case, we use stock returns of the next trading day after the NFP announcements.

| | Panel A: GDP Announcements | | | | Panel B: NFP Announcements | | | |
|---|---|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (1) | (2) | (3) | (4) |
| Constant | -0.06 | -0.06 | -0.04 | -0.02 | 0.16 | -0.01 | 0.01 | 0.02 |
| | (-0.35) | (-0.80) | (-0.56) | (-0.31) | (1.58) | (-0.24) | (0.16) | (0.47) |
| EVR | 0.17 | 0.17 | 0.17 | 0.16 | 0.10 | 0.13 | 0.13 | 0.13 |
| | (2.02) | (2.54) | (2.42) | (2.27) | (1.82) | (2.58) | (2.71) | (2.84) |
| Market $\hat{\beta}$ | | -0.00 | 0.02 | -0.00 | | 0.15 | 0.13 | 0.11 |
| | | (-0.02) | (0.17) | (-0.00) | | (2.07) | (2.03) | (1.84) |
| SMB $\hat{\beta}$ | | | -0.07 | -0.07 | | | -0.02 | -0.02 |
| | | | (-1.07) | (-1.27) | | | (-0.59) | (-0.66) |
| HML $\hat{\beta}$ | | | -0.05 | -0.05 | | | -0.01 | -0.01 |
| | | | (-1.13) | (-1.03) | | | (-0.26) | (-0.17) |
| RMW $\hat{\beta}$ | | | | 0.03 | | | | -0.03 |
| | | | | (0.79) | | | | (-1.04) |
| CMA $\hat{\beta}$ | | | | -0.06 | | | | -0.03 |
| | | | | (-2.21) | | | | (-1.02) |
| $R^2$(%) | 0.32 | 4.01 | 6.20 | 7.53 | 0.40 | 5.50 | 8.78 | 10.30 |

announcement-day premiums. Because these two announcements are important in measuring the overall health of the macroeconomy, they are likely to be more important than other announcements in providing information about future economic growth. In Table 6, we report the time-series average of the risk prices from Fama-MacBeth regressions on GDP and NFP announcement days, where we exclude the corresponding announcement day if it overlaps with an FOMC announcement.

In univariate regressions, EVR continues to carry a significant risk price. The estimated average risk price of EVR on a GDP (NFP) announcement day is 17 bps (10 bps) with a *t*-statistic equal to 2.02 (1.82). In the second specification, we use both EVR and the market beta in a joint regression. The risk price on a GDP (NFP) announcement day for EVR is 17 bps (13 bps) with a *t*-statistic equal to 2.54 (2.58). When we include Fama-French three and five factors in our third and fourth specifications, the point estimate of the risk price of EVR continues to be statistically significant and robust in magnitude. These results confirm our EVR measure can positively predict returns in the cross section on GDP and NFP announcement days beyond the standard factors. Comparing the regression results on GDP and NFP announcement days with that on FOMC announcement days in Table 4, we find the return predictability of EVR is larger in magnitude and more significant on FOMC announcement days in all four regression specifications.

### 3.4. Higher-moments-related risk factors

In this subsection, we provide additional evidence to distinguish our theory from alternative mechanisms. The key premise of our theory is that FOMC announcements reveal the Fed's private information about the economy and the resolution of macroeconomic uncertainty is associated with realizations of market risk premium. This key premise distinguishes our theory from the previous literature in which the predictability of returns is due to variations in the volatility of macroeconomic fundamentals, for example, Bansal and Yaron (2004), Bollerslev et al. (2009), and Savor and Wilson (2013). In the above models, macroeconomic volatility is a persistent state variable that affects expected returns in the future, whereas our mechanism does not rely on the heteroscedasticity of macroeconomic shocks.

We examine several measures of higher moments of market returns, and we show none of them explains the risk price of EVR. If the risk price of EVR is due to exposure to changes in the volatility of macroeconomic fundamentals or their higher moments, it should disappear once we control for risk factors constructed from higher moments. The robustness of the risk prices of EVR with respect to risk factors constructed from shocks to higher moments therefore provides further support for our model of "the Fed information channel."

We consider three higher-moment-related risk factors: the expected market variance and the second and the third moment of realized market returns. First, we examine whether the risk exposure to news about future market variance can explain the risk price of EVR. Our measure of expected variance is constructed as the fitted value of the quarterly realized variance. We use constrained least squares to ensure all the forecasts are non-negative. Our empirical exercise focuses on two measures of expected variance adopted in the previous literature. The first measure $\mathbb{E}_t[Var1]$ follows Campbell et al. (2018) and uses the last quarter's realized variance, the market excess return over the last quarter, the market price-earning ra-

**Table 7**

Explore other risk explanations. The table reports the regression results of daily excess stock returns controlling for the market beta, changes in market expected variance, co-skewness risk, and co-kurtosis risk. For the Fama-MacBeth regression in Panel A, we first compute the market beta, sensitivity to changes in the expected market variance, co-skewness, and co-kurtosis of each firm based on the past 12-month daily excess returns two days before each FOMC announcement. Then, on each FOMC announcement day, we regress the cross-sectional stock excess returns on EVR and other betas from the first step. We report the average loadings on EVR and other risk factor betas and their associated Newey-West $t$-statistics. We demean EVR in cross-sectional regressions for interpretation convenience. $\widehat{\beta}_{\Delta EV1}$ ($\widehat{\beta}_{\Delta EV2}$) is the individual stock's exposure to the news about future market variance $\mathbb{E}_t[Var1]$ ($\mathbb{E}_t[Var2]$), which is calculated using the approach in Campbell et al. (2018) (Savor and Wilson, 2014), as defined in Eq. (6). We report coefficients in the following pooled regression in Panel B:

$$R_t^j - r_{f,t} = \gamma_0 \cdot \mathbf{1}_{Non} + \gamma_{0,FOMC} \cdot \mathbf{1}_{FOMC} + \gamma_1 \cdot EVR_t^j + \gamma_2 \cdot EVR_t^j \cdot \mathbf{1}_{FOMC} + \kappa_1 \cdot \widehat{\beta}^j_{Other,t} + \kappa_2 \cdot \widehat{\beta}^j_{Other,t} \cdot \mathbf{1}_{FOMC} + \delta_1 \cdot \widehat{\beta}^j_{Mkt,t} + \delta_2 \cdot \widehat{\beta}^j_{Mkt,t} \cdot \mathbf{1}_{FOMC} + \varepsilon_t^j.$$

a For each stock, $\widehat{\beta}^j_{Other,t}$ is the sensitivity to changes in the expected market variance, co-skewness, or co-kurtosis, and is estimated using daily returns during the past 12 months. $t$-statistics are calculated using the trading-day clustered standard errors. The unit of regression coefficients is in percentage. Our full sample period is from January 1996 to December 2017 with 176 FOMC announcements.

| | (1) | (2) $\Delta EV1$ | (3) $\Delta EV2$ | (4) | (5) | (6) | (7) |
|---|---|---|---|---|---|---|---|
| Panel A: Fama-MacBeth regressions on FOMC announcement | | | | | | | |
| Constant | 0.40 | -0.01 | 0.00 | 0.40 | 0.01 | 0.39 | 0.01 |
| | (3.11) | (-0.16) | (0.02) | (3.10) | (0.14) | (3.11) | (0.27) |
| EVR | 0.20 | 0.18 | 0.18 | 0.19 | 0.18 | 0.20 | 0.19 |
| | (2.52) | (3.03) | (3.02) | (2.54) | (3.02) | (2.56) | (3.22) |
| $\widehat{\beta}_{Mkt}$ | | 0.37 | 0.36 | | 0.35 | | 0.35 |
| | | (2.76) | (2.69) | | (2.72) | | (2.70) |
| $\widehat{\beta}_{\Delta EV}$ | | -0.30 | -0.32 | | | | |
| | | (-1.94) | (-2.27) | | | | |
| $\widehat{\beta}_{CoSkew}$ | | | | -0.49 | -0.45 | | -0.18 |
| | | | | (-1.07) | (-1.24) | | (-0.86) |
| $\widehat{\beta}_{CoKurt}$ | | | | | | -3.51 | 2.03 |
| | | | | | | (-1.25) | (1.19) |
| $R^2(\%)$ | 0.49 | 5.52 | 5.54 | 1.21 | 5.56 | 1.17 | 5.99 |

Panel B: Pooled regressions

| | $1_{Non}$ | $1_{FOMC}$ | $EVR$ | $EVR \cdot \mathbf{1}_{FOMC}$ | $\widehat{\beta}_{Other}$ | $\widehat{\beta}_{Other} \cdot \mathbf{1}_{FOMC}$ | $\widehat{\beta}_{Mkt}$ | $\widehat{\beta}_{Mkt} \cdot \mathbf{1}_{FOMC}$ | $R^2(\%)$ |
|---|---|---|---|---|---|---|---|---|---|
| $\Delta EV1$ | 0.03 | 0.20 | 0.06 | 0.15 | 0.03 | -0.61 | -0.00 | -0.19 | 0.10 |
| | (1.56) | (1.66) | (2.50) | (1.47) | (0.49) | (-1.72) | (-0.14) | (-1.60) | |
| $\Delta EV2$ | 0.01 | 0.34 | 0.06 | 0.14 | -0.05 | -0.24 | 0.01 | -0.22 | 0.10 |
| | (0.22) | (2.61) | (2.49) | (1.41) | (-0.68) | (-0.57) | (0.48) | (-1.59) | |
| $CoSkew$ | 0.04 | 0.31 | 0.06 | 0.15 | -0.00 | 0.00 | -0.02 | 0.11 | 0.10 |
| | (1.87) | (2.80) | (2.51) | (1.43) | (-0.22) | (0.80) | (-0.85) | (1.13) | |
| $CoKurt$ | 0.04 | 0.28 | 0.06 | 0.15 | -0.00 | 0.00 | -0.02 | 0.13 | 0.10 |
| | (2.00) | (2.78) | (2.50) | (1.44) | (-1.15) | (0.97) | (-0.88) | (1.46) | |

tio, the U.S. treasury yield spread, the default spread, and the value spread as forecast variables. Our second measure of expected variance $\mathbb{E}_t[Var2]$ follows Savor and Wilson (2014) and uses a different set of predictors: the last quarter's realized variance, the market excess returns on FOMC days, and those on non-FOMC days, respectively, over the last quarter. We calculate individual stock $j$'s sensitivity to the first difference of market expected variance in the following multivariate regression:

$$R_t^j - r_{f,t} = \alpha^j + \beta^j_{Mkt}\left(R_t^M - r_{f,t}\right) + \beta^j_{\Delta EV}\Delta\mathbb{E}_t[Var] + \varepsilon_t^j. \tag{6}$$

We obtain the estimates of $\beta^j_{\Delta EV}$ at the firm level using daily returns during the past 12 months. Given these estimates, we run a firm-level Fama-MacBeth regression with both EVR and $\beta^j_{\Delta EV}$, as well as $\beta^j_{Mkt}$ as explanatory variables. We report our results in columns (2) and (3) of Panel A in Table 7. In the first and second rows of Panel B in Table 7, we present the pooled regression results of Eq. (4), with $\widehat{\beta}^j_{\Delta EV,t}$ and $\widehat{\beta}^j_{\Delta EV,t} \cdot \mathbf{1}_{FOMC}$ as additional explanatory variables. Consistent with the literature, we find the volatility news has a negative and significant risk price. However, after controlling for this volatility news, the risk price of EVR remains highly significant.

Second, we study whether beta to realized market variance can explain the risk price of EVR. We follow Harvey and Siddique (2000) and define co-skewness of a stock as the beta to squared market excess returns in the following multivariate regression[8]:

$$R_t^j - r_{f,t} = \alpha^j + \beta^j_{Mkt}\left(R_t^M - r_{f,t}\right) + \beta^j_{CoSkew}\left(R_t^M - r_{f,t}\right)^2 + \varepsilon_t^j. \tag{7}$$

We obtain the co-skewness $\beta^j_{CoSkew}$ and repeat our Fama-MacBeth regressions as above. Columns (4) and (5) of Panel A in Table 7 report the results. First, we find the co-skewness risk is not priced on FOMC days in our sample. In addition, comparing column (4) with column (1), we see the point estimate and the significance of the risk

[8] See Dittmar (2002) for a study of nonlinear pricing kernels that include both a quadratic and a cubic term of market returns.

price of EVR change only slightly after controlling for co-skewness. Finally, in the third row of Panel B in Table 7, we report the result of a pooled regression that augments the right-hand side of Eq. (4) with $\widehat{\beta}^j_{CoSkew,t}$ and $\widehat{\beta}^j_{CoSkew,t} \cdot \mathbf{1}_{FOMC}$. Consistent with the results from the Fama-MacBeth regressions, the risk price for EVR remains significant while that for the co-skewness is not. Moreover, the additional risk price of EVR on FOMC days is about 15 bps, which is similar in magnitude to the same coefficient reported in Table 4 without controlling for the co-skewness risk.

Third, we examine whether the beta of individual stocks to the third moment of realized market returns can explain the risk price of EVR. We follow Christoffersen et al. (2018) and define the beta to the cubic market excess returns as co-kurtosis and measure co-kurtosis in the following multivariate regression:

$$R^j_t - r_{f,t} = \alpha^j + \beta^j_{Mkt}\left(R^M_t - r_{f,t}\right) + \beta^j_{CoSkew}\left(R^M_t - r_{f,t}\right)^2 + \beta^j_{CoKurt}\left(R^M_t - r_{f,t}\right)^3 + \varepsilon^j_t. \tag{8}$$

Column (6) of Panel A in Table 7 reports the result of the Fama-MacBeth regression that controls for co-kurtosis risk, and column (7) reports the result of the regression that simultaneously controls for market beta, co-skewness, and co-kurtosis. Similar to the results of co-skewness, we also find the co-kurtosis risk is not priced on FOMC announcement days in our sample. The risk price of EVR and market beta are significant, and the magnitude and significance of the risk price of EVR remain robust after controlling for the co-kurtosis risk. The fourth row in Panel B of Table 7 presents the pooled regression results in which we control for the co-kurtosis risk. Again, the magnitude of the risk price of EVR remains largely unchanged compared with Table 4. Our findings are robust if we use other measures of the co-skewness and co-kurtosis risk in the literature such as Ang et al. (2006) and Guidolin and Timmermann (2008). These results are presented in Table IA.4 in Online Appendix IA.

Motivated by the above empirical evidence, in the next section, we develop a model of the Fed information channel, in which both productivity shocks and monetary policy shocks command a risk premium to account for the predictability of EVR and the robustness of the CAPM on announcement days.

## 4. A model of the Fed information channel

In this section, we set up a continuous-time equilibrium model in which the Fed's monetary policy announcements reveal its private information about the growth prospects for the economy and study the implications for the cross section of announcement returns. We first present the model setup and solution. Then, we discuss the model's implications on the pre-FOMC announcement drift and the slope of the security market line. We provide details of the model solutions in Appendix B.

### 4.1. Model setup

#### 4.1.1. Preference and technology

We consider a continuous-time representative-agent economy in which the representative agent has a recursive preference with constant relative risk aversion $\gamma$ and constant intertemporal elasticity of substitution (IES) $\psi$. We assume the growth rate of the aggregate output follows a dynamic process of the form

$$\frac{dY_t}{Y_t} = \theta_t dt + \sigma dB_{A,t}. \tag{9}$$

Here, the Brownian motion, $B_{A,t}$, denotes the productivity shock, and $\sigma$ is the volatility of the productivity shock. The expected growth rate, $\theta_t$, follows a continuous-time AR(1) process,

$$d\theta_t = a(\mu - \theta_t)dt + \sigma_\theta dB_{\theta,t}, \tag{10}$$

where $\mu$ denotes the long-run mean and $a > 0$ is the rate of mean reversion of $\theta_t$. The Brownian motion $B_{\theta,t}$ is independent of $B_{A,t}$. In Appendix B.1, we present a parsimonious three-equation New Keynesian model in which $\theta_t$ represents the central bank's target interest rate and $B_{\theta,t}$ denotes the monetary policy shock. Intuitively, in New Keynesian models, the monetary authority sets the nominal interest rate, which, in the presence of price rigidity, affects the real interest rate. The expected consumption growth, through households' intertemporal saving decisions, is a linear function of the real interest rate and hence a linear function of $\theta_t$. In addition, because our model does not have investment, we use the terms *consumption* and *output* interchangeably.

#### 4.1.2. Information and announcements

The Federal Reserve's interest rate target, $\theta_t$, which affects the expected economic growth rate, is not observable to investors but is periodically communicated by the monetary authority through pre-scheduled FOMC announcements. Investors can learn about $\theta_t$ through two sources of information. First, the aggregate output (9) contains noisy information about $\theta_t$. Second, at pre-scheduled FOMC announcement times, $T, 2T, 3T, \cdots$, Fed communication provides additional noisy signals about $\theta_t$. For $n = 1, 2, 3, \cdots$, we denote $s_n$ as the signal observed at time $nT$ and assume $s_n = \theta_{nT} + \varepsilon_n$, where $\varepsilon_n$ is i.i.d. and normally distributed with mean zero and variance $\sigma_s^2$.

We assume investors' prior belief about the latent variable $\theta_t$ is represented by a normal distribution. As a result, the posterior distribution of $\theta_t$ is also Gaussian, and it can be summarized by the first two moments. We define $\hat{\theta}_t \equiv \mathbb{E}_t[\theta_t]$ as the posterior mean and $q_t \equiv \mathbb{E}_t\left[\left(\theta_t - \hat{\theta}_t\right)^2\right]$ as the posterior variance of $\theta_t$ given information up to time $t$. In the interior of $((n-1)T, nT)$ without announcements, investors update their beliefs based on the observed consumption process using a Kalman-Bucy filter:

$$d\hat{\theta}_t = a\left(\mu - \hat{\theta}_t\right)dt + \frac{q_t}{\sigma} d\tilde{B}_{A,t}, \tag{11}$$

where the innovation process is defined by $d\tilde{B}_{A,t} = \frac{1}{\sigma}\left(\frac{dY_t}{Y_t} - \hat{\theta}_t dt\right)$. Intuitively, $d\tilde{B}_{A,t}$ represents the surprises in output growth, which is the difference between the realized and the expected growth. The posterior variance, $q_t$, satisfies the Riccati equation:

$$dq_t = \left(\sigma_\theta^2 - 2aq_t - \frac{1}{\sigma^2}q_t^2\right)dt. \tag{12}$$

Because the monetary authority makes announcements periodically at time $t = nT$, where $n$ is an integer, investors update their beliefs based on the signals $s_n$ about $\theta_t$ using Bayes' rule:

$$\hat{\theta}_{nT}^{+} = q_{nT}^{+}\left(\frac{1}{\sigma_s^2}s_n + \frac{1}{q_{nT}^{-}}\hat{\theta}_{nT}^{-}\right); \quad \frac{1}{q_{nT}^{+}} = \frac{1}{\sigma_s^2} + \frac{1}{q_{nT}^{-}}, \tag{13}$$

where $\hat{\theta}_{nT}^{+}$ and $q_{nT}^{+}$ are the posterior mean and variance after announcements, and $\hat{\theta}_{nT}^{-}$ and $q_{nT}^{-}$ are the posterior mean and variance before announcements, respectively. The uncertainty reduction upon the announcements in our model is characterized by $q_{nT}^{+} - q_{nT}^{-}$.

#### 4.1.3. The cross section of equity

We assume a cross section of equity claims, indexed by $i$. Equity $i$ is the claim to the following dividend process:

$$\frac{dD_t^i}{D_t^i} = \left[\mu + \xi_i\left(\hat{\theta}_t - \mu\right)\right]dt + \eta_i\sigma d\tilde{B}_{A,t} + \sigma_i dB_{i,t}, \tag{14}$$

where $dB_{i,t}$ is the idiosyncratic shock to each firm $i$, which is uncorrelated with the productivity shock $dB_{A,t}$ and monetary shock $dB_{\theta,t}$. The term $\sigma_i$ is the idiosyncratic volatility, and the parameters $(\xi_i, \eta_i)$ measure the sensitivity of the dividend with respect to monetary shocks and productivity shocks, respectively. We assume $\xi_i$ is uniformly distributed over the interval $\left[\underline{\xi}, \overline{\xi}\right]$ and $\eta_i$ is uniformly distributed over $\left[\underline{\eta}, \overline{\eta}\right]$, and the distributions of $\xi_i$ and $\eta_i$ are independent.

In our model, we assume $\xi$ and $\eta$ are constant over time. In the data, this assumption is unlikely to hold. In fact, the EVR measure changes significantly at the firm level over time. However, as we explained earlier, time-varying beta is unlikely to explain the cross section of EVR-sorted portfolios, so assuming constant $\xi$ and $\eta$ is without loss of generality. It is straightforward to extend our model to allow for time variation in $\xi$ and $\eta$ to match the patterns of portfolio transition in the data without affecting the portfolio sorting results.

Define the price-to-dividend ratio of firm $i$ as $p\left(\hat{\theta}_t, t \middle| \xi_i, \eta_i\right)$, which depends on both $\xi_i$ and $\eta_i$. The function $p\left(\hat{\theta}_t, t \middle| \xi_i, \eta_i\right)$ is defined as

$$p\left(\hat{\theta}_t, t \middle| \xi_i, \eta_i\right)D_t^i = \mathbb{E}_t\left[\int_0^\infty \frac{\pi_{t+s}}{\pi_t}D_{t+s}^i ds \middle| \hat{\theta}_t, q_t\right], \tag{15}$$

where the law of motion of $D_t^i$ is given in (14) and the solution for state price density $\pi_t$ is provided later in this section. For simplicity, we henceforth denote the firm-specific price-to-dividend ratio as $p^i\left(\hat{\theta}_t, t\right)$. The return on firm $i$'s equity during the period $(t, t+\Delta)$ is given by

$$R_{t,t+\Delta}^i = \frac{p^i(\hat{\theta}_{t+\Delta}, t+\Delta)\frac{D_{t+\Delta}^i}{D_t^i} + \int_t^{t+\Delta}\frac{D_s^i}{D_t^i}ds}{p^i\left(\hat{\theta}_t, t\right)}. \tag{16}$$

### 4.2. Model solution

In this section, we present the solution for the asset prices in our model. We also show the expected variance reduction in the model identifies the structural parameter $\xi_i$, which determines the sensitivity of stock returns to monetary policy announcements. The link between expected variance reduction and the sensitivity of returns to monetary policy surprises in our model provides a theoretical foundation for the empirical measure of EVR we constructed in Section 2.1.

#### 4.2.1. Asset prices

In the interior of $((n-1)T, nT)$, $n = 1, 2, \cdots$ without announcements, the law of motion of the real pricing kernel, $\pi_t$, satisfies the following stochastic differential equation (SDE):

$$d\pi_t = \pi_t\left[-r\left(\hat{\theta}_t, t\right)dt - \sigma_{\pi,t}d\tilde{B}_{A,t}\right], \tag{17}$$

where

$$r\left(\hat{\theta}_t, t\right) = \rho + \frac{1}{\psi}\hat{\theta}_t - \frac{1}{2}\gamma\left(1 + \frac{1}{\psi}\right)\sigma^2 + \frac{\frac{1}{\psi} - \gamma}{a+\rho}q_t + \frac{\left(\frac{1}{\psi} - \gamma\right)\left(1 - \frac{1}{\psi}\right)}{2(a+\rho)^2}\left(\frac{q_t}{\sigma}\right)^2 \tag{18}$$

is the risk-free interest rate and

$$\sigma_{\pi,t} = \gamma\sigma + \frac{\gamma - \frac{1}{\psi}}{a+\rho}\frac{q_t}{\sigma} \tag{19}$$

is the market price of the innovations in consumption growth, $\tilde{B}_{A,t}$. Eq. (19) contains both the compensation for the i.i.d. shock $B_{A,t}$ and that for the changes in the belief about $\theta_t$, captured by the term $\frac{\gamma - \frac{1}{\psi}}{a+\rho}\frac{q_t}{\sigma}$. The two sources of fundamental risk in the economy are $B_{A,t}$ and $B_{\theta,t}$. However, investors do not observe $\theta_t$ and cannot distinguish whether a change in output growth is due to productivity shock $B_{A,t}$ or monetary shock $B_{\theta,t}$. Innovations in consumption growth, from the investors' perspective, affect both the contemporaneous consumption growth rate and investors' beliefs about $\theta_t$. As in Ai (2010), $\sigma_{\pi,t}$ summarizes risk compensation from both channels.

As we show in Appendix B.3, each firm's price-to-dividend ratio $p^i\left(\hat{\theta}_t, t\right)$ must satisfy the partial differential equation of (B.23). In addition, at announcements $nT, n = 1, 2, 3, \ldots$, the boundary condition satisfies

$$p^i\left(\hat{\theta}_{nT}^{-}, nT^{-}\right) = \mathbb{E}\left[\frac{e^{\frac{\frac{1}{\psi}-\gamma}{a+\rho}\hat{\theta}_{nT}^{+}}p^i\left(\hat{\theta}_{nT}^{+}, nT^{+}\right)}{\left(\mathbb{E}\left[e^{\frac{1-\gamma}{a+\rho}\hat{\theta}_{nT}^{+}}\middle|\hat{\theta}_{nT}^{-}, q_{nT}^{-}\right]\right)^{\frac{\frac{1}{\psi}-\gamma}{1-\gamma}}}\middle|\hat{\theta}_{nT}^{-}, q_{nT}^{-}\right]. \tag{20}$$

That is, $SDF_{nT^-,nT^+} = e^{\frac{\frac{1}{\psi}-\gamma}{a+\rho}\hat{\theta}_{nT}^{+}} / \left(\mathbb{E}\left[e^{\frac{1-\gamma}{a+\rho}\hat{\theta}_{nT}^{+}}\middle|\hat{\theta}_{nT}^{-}, q_{nT}^{-}\right]\right)^{\frac{\frac{1}{\psi}-\gamma}{1-\gamma}}$ is the announcement SDF that prices post-announcement payoff into pre-announcement prices.

Under the assumption of generalized risk sensitivity (Ai and Bansal, 2018), $\gamma > \frac{1}{\psi}$, the term $e^{\frac{\frac{1}{\psi}-\gamma}{a+\rho}\hat{\theta}^{+}_{nT}}$ is negatively correlated with the posterior belief, $\hat{\theta}^{+}_{nT}$, which is updated immediately following announcements. As a result, an asset with a payoff that increases after the announcement about $\theta_{nT}$ requires a positive risk premium, and the magnitude of the announcement premium increases with the sensitivity of the asset's payoff to the announcement about $\theta_{nT}$.

*4.2.2. Expected variance reduction*

Below we show that in our model, both the sensitivity of stock returns to announcement surprises and the expected variance reduction are strictly increasing functions of $\xi_i$. As a result, the expected variance reduction in the model is a perfect measure of the sensitivity of stock returns to monetary policy announcements, which determines the cross section of the monetary policy announcement premium.

The definition of return in Eq. (16) implies the log announcement return of stock $i$ is

$$\ln R^i_{nT^-,nT^+} = \ln p^i\left(\hat{\theta}^{+}_{nT}, nT^+\right) - \ln p^i\left(\hat{\theta}^{-}_{nT}, nT^-\right). \tag{21}$$

Using a log-linear approximation, $\ln p^i\left(\hat{\theta}_t, t\right) \approx \frac{p^i_\theta\left(\hat{\theta}_t,t\right)}{p^i\left(\hat{\theta}_t,t\right)}\hat{\theta}_t + g(t)$, where $g(t)$ is a deterministic function of time, $p^i_\theta\left(\hat{\theta}_t, t\right) \equiv \frac{\partial p^i\left(\hat{\theta}_t,t\right)}{\partial \hat{\theta}_t}$, and $\frac{p^i_\theta\left(\hat{\theta}_t,t\right)}{p^i\left(\hat{\theta}_t,t\right)}$ is the sensitivity of the stock return with respect to $\hat{\theta}_t$. In Appendix B.4, we show the sensitivity $\frac{p^i_\theta\left(\hat{\theta}_t,t\right)}{p^i\left(\hat{\theta}_t,t\right)} \approx \frac{\xi_i-\frac{1}{\psi}}{a+\bar{p}^i}$, where $\bar{p}^i$ is the inverse of the steady-state level of the price-to-dividend ratio for firm $i$. Hence, the log announcement return in Eq. (21) can be written as:

$$\ln R^i_{nT^-,nT^+} - \mathbb{E}_{nT^-}\left[\ln R^i_{nT^-,nT^+}\right] = \frac{\xi_i - \frac{1}{\psi}}{a+\bar{p}^i}\left(\hat{\theta}^{+}_{nT} - \hat{\theta}^{-}_{nT}\right). \tag{22}$$

Clearly, the return of a stock with a higher $\xi_i$ is more sensitive to shocks to $\hat{\theta}_t$ and therefore requires a higher level of risk compensation on announcement days.

In addition, the expected variance reduction in our model is also a strictly increasing function of $\xi_i$. For illustration, note the implied variance reduction following the announcement is:

$$\begin{aligned}\Delta IV^i &= Var\left[\ln R^i_{nT^-,nT+\frac{30}{360}}\right] - Var\left[\ln R^i_{nT^+,nT+\frac{30}{360}}\right] \\ &= Var\left[\ln R^i_{nT^-,nT^+}\right].\end{aligned} \tag{23}$$

In the above expression, $R^i_{nT^-,nT+\frac{30}{360}}$ is the 30-day return including the announcement return at $nT$, and $R^i_{nT^+,nT+\frac{30}{360}}$ is the 30-day return without the announcement return. The first equality holds because our model contains no variance risk premium, and the option-implied variance equals the variance under the physical probability measure. The second equality is because returns are uncorrelated over time in our model. Using Eq. (22), the expected implied variance reduction following announcements is then given by

$$EVR^i = \mathbb{E}\left[\Delta IV^i\right] = \left(\frac{\xi_i - \frac{1}{\psi}}{a+\bar{p}^i}\right)^2\left(\frac{q^{-2}_{nT}}{q^{-}_{nT} + \sigma_s^2}\right). \tag{24}$$

Evidently, $EVR^i$ is a strictly increasing function of $\xi_i$, which determines the sensitivity of the dividend with respect to monetary policy shocks. Therefore, in our model, sorting on the expected implied variance reduction upon FOMC announcements is equivalent to sorting on $\xi_i$ and perfectly identifies stock returns' sensitivity to monetary policy announcements, $\frac{p^i_\theta\left(\hat{\theta}^{+}_{nT},nT^+\right)}{p^i\left(\hat{\theta}^{+}_{nT},nT^+\right)}$.

*4.3. Information release and risk compensation*

The key implication of our model is that the realization of the risk premium is associated with the release of information. This feature distinguishes our model from the standard long-run risk setup such as Bansal and Yaron (2004) and Savor and Wilson (2013), where the time-varying risk premium is due to time variations in the volatility of macroeconomic fundamentals. However, macroeconomic volatility is unlikely to change substantially just on an announcement day, whereas information arrives on financial markets continuously. In our model, macroeconomic shocks are homoscedastic, and the time-varying risk premium is driven by the time-varying information structure.

To highlight the difference between our model and the literature on time-varying volatility, in Eqs. (9), (10), and (14), we assume all macroeconomic volatilities are constant. Despite the homoscedasticity assumption, equity market returns are predictable. In particular, the magnitude of the realized equity premium during an event of information release is proportional to the amount of uncertainty reduction upon such an event. This implication can be seen by calculating the volatility of the log announcement SDF, $\ln\left(SDF_{nT^-,nT^+}\right)$, in Eq. (20):

$$Std\left[\frac{\frac{1}{\psi}-\gamma}{a+\rho}\hat{\theta}^{+}_{nT}\right] = \frac{\gamma - \frac{1}{\psi}}{a+\rho}\sqrt{q^{-}_{nT} - q^{+}_{nT}}. \tag{25}$$

That is, the equity premium is proportional to the square root of the amount of uncertainty reduction represented by the difference between the posterior variance before and after announcements. As a result, our model has two testable implications on the announcement-day premiums: pre-FOMC announcement drift and the slope of the security market line on announcement days.

*4.3.1. Pre-FOMC announcement drift*

First, in our model, risk premiums are realized upon the resolution of uncertainty. If information is revealed hours ahead of the FOMC announcement, the risk premium will realize during the same period. Empirically, as documented by Lucca and Moench (2015), most of the FOMC announcement premium is realized hours ahead of announcements. In our model, if the information is released to the market

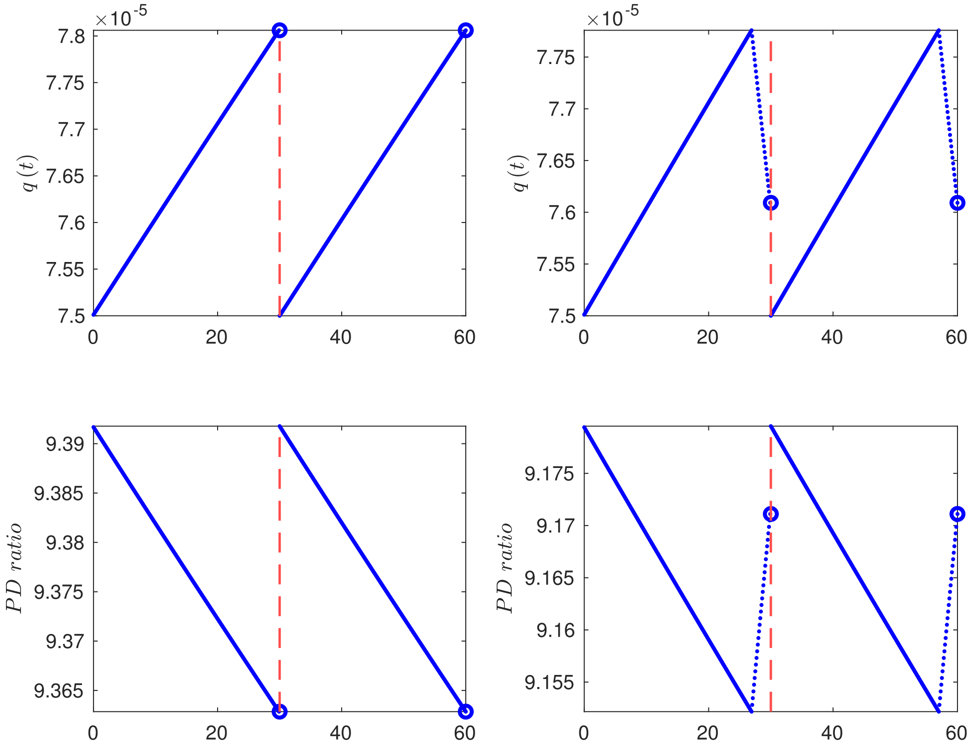


**Fig. 5.** Equilibrium without and with pre-released information. This figure plots $q_t$, the posterior variance of $\hat{\theta}_t$ in the top panel and the average price-to-dividend ratio in the bottom panel. The left column is the model in which all information is released at time $T$ = day 30 and day 60, respectively. The right column is a model in which the agent starts to observe additional signals at time $\tau$ = day 27 and day 57, which are three days before FOMC announcements. For simplicity, we choose $\xi = (\bar{\xi} + \underline{\xi})/2 = 4$ and $\eta = (\bar{\eta} + \underline{\eta})/2 = 4$ as the level of market sensitivity and $\hat{\theta}$ at the steady-state value $\mu = 1.5\%$ (see Table 8). For the model of the right column, we choose $\sigma_s = 7.23\%$ so that the two models have the same $q(0) = 0.0075\%$. All other parameter values are reported in Table 8.

hours ahead of announcements, the pre-FOMC announcement drift can be interpreted as risk compensation associated with the information release.

To demonstrate this implication, we compare a version of our model in which all information revelation happens at announcements with a version of the model in which information is partially revealed before announcements. In the latter version of the model, we assume that starting at time $\tau < T$, all investors observe a pre-announcement additional signal $\zeta_t$, which carries information about the upcoming announcement of $\theta_t$:

$$\zeta_t = \theta_t dt + \sigma_\zeta dB_{\zeta,t}, \tag{26}$$

where the constant $\sigma_\zeta$ is the inverse of the pre-announcement signal precision and $B_{\zeta,t}$ is a mutually independent Brownian motion noise. We provide details of the model solution in Online Appendix IB.1.

In Fig. 5, we plot the posterior variance (top panel) and the average price-to-dividend ratio (bottom panel) implied by our model for two announcement cycles, where announcements are made on day 30 and day 60, respectively. The left column is the model in which the only information release occurs on day 30 and day 60. In this case, the announcement resolves uncertainty and the posterior variance drops immediately at announcement time $T$. Associated with the jump in the posterior variance, a jump occurs in the price-to-dividend ratio at the same time. That is, in this example, all of the announcement premium is realized at the point of announcement time $T$.

The right column is the model with an additional signal released at time $\tau < T$. We choose $\tau$ so that the additional information is released three days before the FOMC announcement and set $\sigma_\zeta = 0.5\%$ so that the pre-announcement signal $\zeta_t$ is very informative about $\theta_t$. As a result, the posterior variance $q_t$ drops sharply but continuously from $\tau$ to $T$. At announcement time $T$, a smaller drop in $q_t$ occurs, because information is already partially revealed before $T$. The average price-to-dividend ratio in the bottom panel follows closely the pattern of the posterior variance. A pre-FOMC announcement drift starting from time $\tau$ until the announcement time $T$. An additional risk premium is released at announcement time $T$ as the posterior variance drops further. The fact that the risk premium is realized at the same time as the release of information is the key difference between our model and standard consumption-based asset pricing models in which the time-varying risk premium comes from the time-varying volatility of fundamentals.

#### 4.3.2. The slope of the security market line

The second implication of our model is that the slope of the security market line is much larger on announcement days than on non-announcement days. Note the slope of the security market line is the market risk premium. In

**Table 8**
Model parameters. This table reports the parameter values used in the model. All parameters are annualized. We assume announcements are made at a monthly frequency, that is, $T = \frac{1}{12}$.

| Parameter | Description | Values |
|---|---|---|
| $\rho$ | Time discount rate | 0.01 |
| $\gamma$ | Risk aversion | 10 |
| $\psi$ | IES | 2 |
| $\mu$ | Stationary mean of unobserved interest rate target | 1.5% |
| $\sigma$ | Volatility of productivity shock | 3% |
| $a$ | Mean reversion rate of unobserved interest rate target | 0.085 |
| $\sigma_\theta$ | Volatility of unobserved interest rate target | 0.75% |
| $\sigma_i$ | Idiosyncratic volatility | 56% |
| $\sigma_s$ | Informativeness of FOMC announcement | 4.37% |
| $\sigma_\zeta$ | Precision of pre-FOMC information release | 0.5% |
| $[\bar{\xi}, \underline{\xi}]$ | Sensitivity of dividend to monetary shocks | [2,6] |
| $[\bar{\eta}, \underline{\eta}]$ | Sensitivity of dividend to productivity shocks | [2,6] |

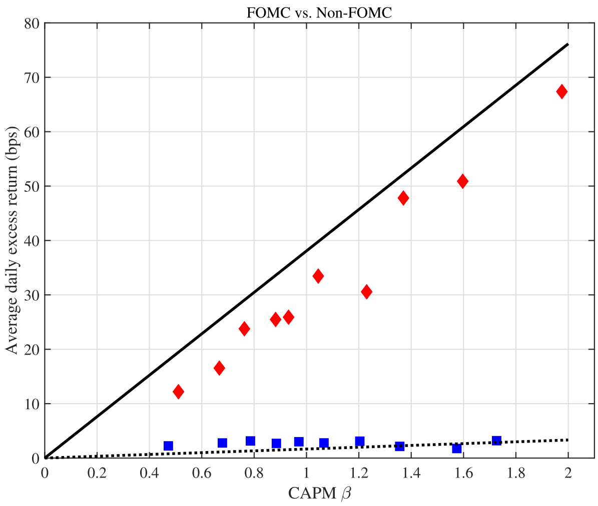


**Fig. 6.** Security market line on announcement days and non-announcement days. This figure plots the security market line, that is, the expected return-beta relationship implied by the model. The solid line is the security market line on FOMC announcement days, and the dotted line is the security market line on non-FOMC announcement days. The diamonds are the average excess returns of beta-sorted portfolios on FOMC announcement days in the data. The squares are the average excess returns of beta-sorted portfolios on non-FOMC-announcement days in the data.

our model, because the volatility of the SDF is particularly high on announcement days, the expected return-beta relationship is much more significant on announcement days than on non-announcement days. In Fig. 6, we replicate the Savor and Wilson (2014) result and plot the expected return-beta relationship for beta-sorted portfolios on FOMC announcement days (diamonds) and on non-FOMC announcement days (squares).

Our result is consistent with the evidence in Savor and Wilson (2014) in the sense that the betas of the portfolios are very stable over the announcement and non-announcement days, but the return spread between the portfolios is much larger on announcement days. The reason is that the slope of the security market line reflects the volatility of the SDF. In Fig. 6, the solid line is the model-implied security market line on announcement days. Note that in the model, an asset with a $\beta = 1$ must earn the announcement premium, which is 38.1 bps in our calibration. The dotted line is the model-implied security market line on non-announcement days. Because the model-implied non-announcement-day average return is close to zero, the security market line on non-announcement days is almost flat, as in the data. Clearly, higher-beta portfolios earn a higher premium on announcement days not because they have a time-varying beta, but because the market return is, on average, higher.

As we commented earlier, the significance of Savor and Wilson's (2014) result is that it points out the following challenge for the discovery of new risk factors. Because the market factor commands a significant premium on announcement days (many times larger than on non-announcement days), most known risk factors, such as value and momentum, are not powerful enough to overturn the CAPM on announcement days. In the following section, we show that despite the high market risk premium on announcement days, portfolios sorted on EVR in our model still earn a significant risk premium relative to the CAPM on announcement days, as in the data.

## 5. Quantitative implications

In this section, we calibrate our model and examine its quantitative implications on the FOMC announcement premium and the return on EVR-sorted portfolios.

### 5.1. Calibration

We calibrate the parameters of aggregate output in our model directly to match the dynamics of the consumption process and report this calibration in Table 8. Following Bansal and Yaron (2004) and Ai (2010), among others, we set the discount rate $\rho = 0.01$, the risk aversion $\gamma = 10$, and IES $\psi = 2$. We choose the average consumption growth $\mu = 1.5\%$ and the standard deviation of consumption growth $\sigma = 3.0\%$ to match the first and second moments of aggregate consumption in the sample period 1929–2018. We also choose the autocorrelation $a$ and the standard deviation of $\theta_t$ to be in line with standard long-run risk models, such as Bansal and Yaron (2004) and Ai (2010).

**Table 9**
Market excess returns on FOMC announcement and non-FOMC announcement days. The top panel reports the S&P 500 excess return on FOMC announcement days from Lucca and Moench (2015) (January 1980 to March 2011) and the corresponding excess returns on non-FOMC announcement days. The bottom panel is the model-implied average daily excess returns for 400 stocks with 22 valid years (122 years of simulation and 100 years of burn-in), a total of 264 FOMC announcement days. We simulate 500 independent sample paths with a daily frequency and report the value-weighted average excess returns on FOMC announcement and non-FOMC announcement days. All numbers are in basis points.

| | Data | |
|---|---|---|
| | Non-FOMC | FOMC |
| Excess Return | 2.0 | 36.6 |
| | Model | |
| | Non-FOMC | FOMC |
| Excess Return | 1.7 | 38.1 |

The parameters of investors' beliefs and the cross section of dividend processes are specific to our model. We choose them to match the dynamics of implied variance in our sample. We set $\sigma_i = 56\%$ to match the cross-sectional average implied variance of stocks on non-announcement days. Our model gives an implied variance of 274.28 in monthly percentage squared units, which is close to 276.21 in our data, reported in Table 1. The qualitative implication of the model on the announcement premium does not depend on the choice of $\sigma_s$, but quantitatively, the magnitude of the announcement premium and the volatility of the announcement return are increasing in the precision of the signal. We choose the parameter for the informativeness of FOMC announcements, $\sigma_s = 4.37\%$, so that our model matches an average announcement-day excess return of 38.1 bps. Finally, we choose $\underline{\xi} = \underline{\eta} = 2$ and $\overline{\xi} = \overline{\eta} = 6$ to match the slope in the cross section of expected returns. The calibrated parameter values are listed in Table 8.

We simulate our model for 122 years. We discard the first 100 years and keep the remaining 22 years so that the time span of our simulation is the same as in our data, and we can compare not only the point estimates but also the $t$-statistics in our model and their counterparts in the data.

### 5.2. Aggregate FOMC announcement premium

In Table 9, we report the average market excess returns on FOMC announcement days and non-FOMC announcement days for both the data and the model. In our calibration, the average market excess returns on FOMC announcement days and non-FOMC announcement days are 38.1 bps and 1.7 bps, respectively, which are very close to the same moments (36.6 bps and 2.0 bps) reported in Lucca and Moench (2015). The average firm-level implied variance reduction on announcement days in our model is 9.0 in monthly percentage squared units, comparable to the same number we reported in Table 1 in our data. Our model matches this pattern in the data quite well.

In our model, the two sources of risk are productivity risk, as captured by the Brownian motion $d\tilde{B}_{A,t}$, and news about the interest-rate target, as captured by changes in (the beliefs about) $\theta_t$ (i.e., $\hat{\theta}^{+}_{nT} - \hat{\theta}^{-}_{nT}$). Prior to the announcements, investors do not observe the true value of $\theta_t$, and they update their beliefs based on observed output growth. Because output growth is driven by Brownian motion shocks, the posterior belief, $\hat{\theta}_t$, updates continuously. At pre-scheduled announcement times, FOMC announcements are associated with discrete jumps in the posterior belief from $\hat{\theta}^{-}_{nT}$ to $\hat{\theta}^{+}_{nT}$.

Because investors' preferences satisfy generalized risk sensitivity, marginal utility is decreasing in the continuation value, which is a function of the posterior belief $\hat{\theta}_t$. Announcements carry news about $\hat{\theta}_t$, so they correlate with marginal utilities and are risky from the investor's perspective. As a result, stocks that are more sensitive to monetary policy news about $\theta_t$ require a larger amount of compensation following announcements. We now turn to the implications of our model in the cross section of equity returns.

### 5.3. Portfolios sorted on expected sensitivity

In our model, because announcement premiums represent compensation for shocks to beliefs about $\hat{\theta}_t$, stocks that are more sensitive to $\hat{\theta}_t$ require a higher level of compensation in terms of announcement returns. The sensitivity of stock returns to announcement surprises depends primarily on the sensitivity of dividend growth with respect to announcements, $\xi_i$. From Eq. (14), we see a stock with a higher sensitivity to monetary policy shocks, which is captured by the parameter $\xi_i$, is more sensitive to $\theta_t$. Therefore, its price-to-dividend ratio and return respond more to news about $\theta_t$, and it requires a high level of risk compensation.

In model simulations, because EVR is a perfect measure of $\xi_i$, high $\xi_i$ stocks are allocated to high-EVR portfolios and therefore have high announcement-day returns on average. In Table 10, we compare the average FOMC-day and non-FOMC-day return of portfolios sorted on expected sensitivity in our model and in the data. Our model replicates the monotone pattern of the FOMC announcement premiums for portfolios sorted on the expected sensitivity, and the spread on the announcement-day return of portfolios is about 33.8 bps, which is close to the empirical counterpart of 31.4 bps.

In addition, CAPM does not explain the FOMC announcement premium in our model. We run the same CAPM regression for the portfolios sorted on expected sensitivity and compare our model output and the data counterparts in Table 11.[9] As shown, the spread in announce-

[9] Because we have shown in Section 2 that the CAPM beta does not change significantly on FOMC days, we do not include the interaction of

**Table 10**

Announcement premium for portfolios sorted on expected sensitivity. This table documents the announcement and non-announcement returns for portfolios sorted on expected sensitivity in basis points. The top panel is based on data from January 1996 to December 2017 with 176 FOMC days. We sort stocks based on the expected implied variance reduction two days before FOMC days and record the long-short portfolio returns on FOMC days. We report FOMC and non-FOMC returns on the decile portfolios, the long-short portfolio, and the associated Newey-West $t$-statistics (in parentheses). The bottom panel is the model-implied average daily excess returns in basis points and the associated Newey-West $t$-statistics for decile portfolios. In our simulation, we use 400 stocks with 22 valid years (122 years of simulation and 100 years of burn-in), a total of 264 FOMC days. For each stock, we simulate 500 independent daily sample paths. We then sort these 400 stocks into 10 portfolios based on expected sensitivity $\xi$ and report the mean and $t$-statistics of each portfolio's excess returns on FOMC and non-FOMC days. We simulate 500 independent daily sample paths and report the average results over all the sample paths.

| | | | Data | | | |
|---|---|---|---|---|---|---|
| | 1 | 2 | 3–8 | 9 | 10 | (10-1) |
| FOMC Return | 35.95 | 28.56 | 29.01 | 47.04 | 67.35 | 31.40 |
| | (1.92) | (2.35) | (3.33) | (3.45) | (3.06) | (2.67) |
| Non-FOMC Return | 3.48 | 3.36 | 3.60 | 4.30 | 3.29 | -0.19 |
| | (1.44) | (1.73) | (2.65) | (2.36) | (1.34) | (-0.15) |
| | | | Model | | | |
| | 1 | 2 | 3–8 | 9 | 10 | (10-1) |
| FOMC Return | 19.84 | 24.82 | 38.64 | 50.59 | 53.65 | 33.81 |
| | (1.78) | (1.91) | (2.06) | (2.05) | (2.04) | (2.07) |
| Non-FOMC Return | 1.36 | 1.45 | 1.67 | 1.87 | 1.97 | 0.61 |
| | (0.25) | (0.26) | (0.31) | (0.29) | (0.30) | (0.15) |

**Table 11**

CAPM for portfolios sorted on expected sensitivity. This table documents the CAPM regression for portfolios sorted on expected sensitivity. The top panel is based on data from January 1996 to December 2017 with 176 FOMC announcement days. We sort stocks based on the expected implied variance reduction two days before FOMC announcement days. We run a CAPM regression on the market excess return, non-FOMC dummy, and FOMC dummy. We report the CAPM beta and coefficients of the non-FOMC dummy and FOMC dummy on the decile portfolios, the long-short portfolio, and the associated Newey-West $t$-statistics (in parentheses). The bottom panel reports the model-implied CAPM regression coefficients for decile portfolios. In the simulation, we use 400 stocks with 22 valid years (122 years of simulation and 100 years of burn-in), a total of 264 FOMC announcement days. We then sort these 400 stocks into 10 portfolios based on expected sensitivity $\xi$ and run CAPM regressions on the simulated market excess return, non-FOMC dummy, and FOMC dummy. We report the mean and $t$-statistics of these regression coefficients. We simulate 500 independent daily sample paths and report the average results over all the sample paths.

| | | | Data | | | |
|---|---|---|---|---|---|---|
| | 1 | 2 | 3–8 | 9 | 10 | (10-1) |
| CAPM Beta | 1.41 | 1.15 | 0.92 | 1.22 | 1.44 | 0.03 |
| Non-FOMC Dummy | -0.71 | -0.22 | 0.56 | 0.55 | -0.98 | -0.27 |
| | (-0.60) | (-0.25) | (1.96) | (0.75) | (-0.82) | (-0.21) |
| FOMC Dummy | -9.85 | -9.03 | -1.22 | 7.22 | 20.46 | 30.31 |
| | (-1.35) | (-1.96) | (-0.78) | (1.79) | (2.48) | (2.85) |
| | | | Model | | | |
| | 1 | 2 | 3–8 | 9 | 10 | (10-1) |
| CAPM Beta | 0.80 | 0.85 | 1.00 | 1.15 | 1.19 | 0.39 |
| Non-FOMC Dummy | 0.03 | 0.04 | 0.00 | -0.04 | -0.01 | -0.04 |
| | (0.06) | (0.08) | (-0.02) | (-0.08) | (-0.02) | (-0.05) |
| FOMC Dummy | -10.28 | -7.23 | 0.47 | 6.64 | 8.09 | 18.37 |
| | (-3.55) | (-2.56) | (0.63) | (2.35) | (2.80) | (4.09) |

ment returns is large and significant across $\xi$-sorted portfolios, whereas beta is only slightly increasing in the expected sensitivity. As a result, as in the data, the coefficients on the FOMC dummy are significant for most portfolios. Moreover, our model replicates the monotonic pattern of the FOMC dummy across portfolios sorted on expected sensitivity quite well.

To understand our result, in the left panel of Fig. 7, we plot the sensitivity of a stock's announcement return to $\hat{\theta}_t$ as a function of $\xi_i$, normalized by the same sensitivity measure of the market return, that is, $\frac{p^i_\theta\left(\hat{\theta}_t,t\right)}{p^i\left(\hat{\theta}_t,t\right)} / \frac{p_\theta\left(\hat{\theta}_t,t\right)}{p\left(\hat{\theta}_t,t\right)}$. Here, we report our results using a global numerical solution method, which is detailed in Online Appendix IB.2. Because the sensitivity of the announcement return to $\hat{\theta}_t$ depends on both $\xi$ and $\eta$, we plot for three different values of $\eta$: $\eta_i = 2, 4, 6$.

In our model, as we showed in Section 4.2, the sensitivity of the stock return to announcement surprises depends

the FOMC dummy and market excess returns in the model calibration for parsimony.

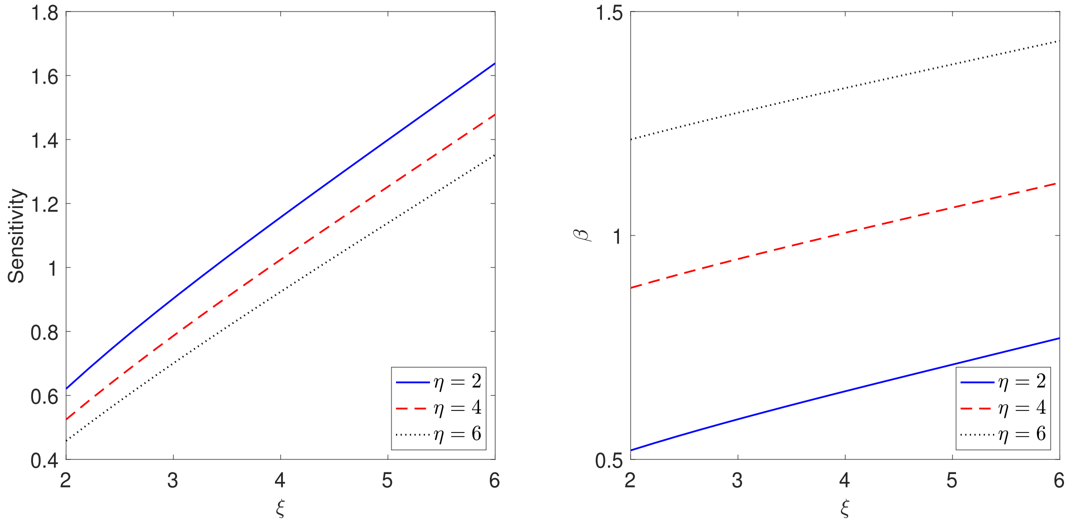


**Fig. 7.** Model-implied sensitivity to FOMC announcement surprises and market returns. The left panel plots the model-implied individual stock's sensitivity of FOMC announcement returns with respect to announcement surprises, normalized by the market sensitivity, as functions of $\xi$ for different values of $\eta$, that is, $\frac{p^i_\theta(\hat{\theta}_t,t)/p^i(\hat{\theta}_t,t)}{p_\theta(\hat{\theta}_t,t)/p(\hat{\theta}_t,t)}$. The right panel plots the model-implied sensitivity of individual stock's returns with respect to market returns at the announcement, that is, CAPM $\beta$ on FOMC announcement days (see Eq. (27)), as functions of $\xi$ for different values of $\eta$. We fix $\hat{\theta}$ at the steady-state value $\mu = 1.5\%$ (see Table 8) for simplicity and $t = nT^+ = 0$ at the announcement. In addition, we choose $\xi = \left(\bar{\xi} + \underline{\xi}\right)/2 = 4$ and $\eta = \left(\bar{\eta} + \underline{\eta}\right)/2 = 4$ as the level of market sensitivity.

only on $\frac{p^i_\theta\left(\hat{\theta}_t,t\right)}{p^i\left(\hat{\theta}_t,t\right)}$. From the figure, we see this sensitivity measure is increasing in $\xi$ but decreasing in $\eta$, where the impact of $\eta$ is quantitatively small. The parameter $\xi$ determines the sensitivity of dividend growth to the interest-rate target $\theta_t$, which is revealed through announcements. As a result, the return of stocks with higher values of $\xi$ is more sensitive to announcements. By comparison, the impact of $\eta$ is much smaller.[10] The sensitivity of the stock returns to announcement surprises depends mostly on $\xi$ and not on $\eta$. Therefore, as shown in Eq. (23), sorting on the expected implied variance reduction is equivalent to sorting on $\xi$ in our model.

In the right panel of Fig. 7, we plot the local CAPM beta as a function of $\xi$ for $\eta_i = 2, 4, 6$. Local betas are computed as follows:

$$\beta^i = \frac{Cov\left[dR^i_t, dR_t\right]}{Var\left[dR_t\right]} = \frac{\eta_i\sigma + \frac{p^i_\theta\left(\hat{\theta}_t,t\right)}{p^i\left(\hat{\theta}_t,t\right)}\frac{q_t}{\sigma}}{\eta\sigma + \frac{p_\theta\left(\hat{\theta}_t,t\right)}{p\left(\hat{\theta}_t,t\right)}\frac{q_t}{\sigma}}. \tag{27}$$

In our model, on non-announcement days, because investors do not observe $\theta_t$, the only shock that affects stock market returns is the surprise in output growth, $d\tilde{B}_{A,t}$. As we have explained, innovations in output growth affect both the contemporaneous growth rate of consumption and the posterior belief about the Fed's interest-rate target, $\hat{\theta}_t$. Therefore, qualitatively, the sensitivity of the stock return to innovations in consumption growth, captured by the parameter $\eta_i$, and the sensitivity of the stock return to the news about the target interest rate, determined by the parameter $\xi_i$, will both affect the estimated CAPM $\beta^i$. Therefore, the estimate of $\beta^i$ increases in both $\xi_i$ and $\eta_i$.

Quantitatively, the estimated $\beta^i$ increases strongly with respect to $\eta_i$ but only mildly with respect to $\xi_i$. As a result, EVR-sorted portfolios display a significant dispersion in $\xi_i$ but a small dispersion in estimated $\beta^i$. In the model, although investors do not observe the true value of $\theta_t$, periodic FOMC announcements provide information about $\theta_t$, and the posterior variance of $\theta_t$, $q_t$, is much smaller relative to $\sigma$ in Eq. (27). Because the estimated $\beta^i$ is more sensitive to $\eta_i$ and less sensitive to $\xi_i$, the EVR-sorted portfolios have a small dispersion in $\beta^i$, which cannot fully account for the difference in the announcement returns in CAPM regressions. As shown in Table 11, the FOMC announcement dummies are quite significant and monotonically increasing in our model, as in the data.

### 5.4. Beta-sorted portfolios

Savor and Wilson (2014) show the CAPM explains the announcement premiums of CAPM beta-sorted portfolios very well. As we showed in the last section, the SDF in our model is driven by two sources of risk, and a one-factor model such as the CAPM cannot fully account for the cross section of announcement premiums. However, as we demonstrate below, despite the failure of the CAPM in explaining the EVR-sorted portfolios, our model can account

[10] The log-linear approximation we presented in Section 4.2 does not capture the dependence of $\frac{p^i_\theta\left(\hat{\theta}_t,t\right)}{p^i\left(\hat{\theta}_t,t\right)}$ on $\eta_i$, but the global solution we used here does. High $\eta$ stocks are less sensitive to announcement surprises. The stocks with a higher $\eta$ require a higher risk premium and therefore have a lower cash-flow duration. As a result, its price-to-dividend ratio is less sensitive to news about the dividend growth rate.

**Table 12**
Announcement premium for beta-sorted portfolios. This table documents the announcement and non-announcement returns for beta-sorted portfolios in basis points. The top panel is based on data from January 1996 to December 2017 with 176 FOMC announcement days. We sort stocks based on the CAPM regression coefficient of a single stock excess return on the market excess return. We report FOMC announcement and non-FOMC announcement returns on decile portfolios, the long-short portfolio, and the associated Newey-West *t*-statistics (in parentheses). The bottom panel is the model-implied average excess returns in basis points and the associated Newey-West *t*-statistics for decile portfolios. In the simulation, we use 400 stocks with 22 valid years (130 years with 100 years of burn-in and 8 years of pre-sample to estimate single stocks' CAPM beta coefficients), a total of 264 FOMC announcement days. We then sort these 400 stocks into 10 portfolios based on the estimated beta coefficients and report the mean portfolio $\xi$, the mean portfolio $\eta$, and the mean and *t*-statistics of each portfolio's excess returns on FOMC announcement and non-FOMC announcement days. We simulate 500 independent daily sample paths and report the average results over all the sample paths.

| | | | Data | | | |
|---|---|---|---|---|---|---|
| | 1 | 2 | 3–8 | 9 | 10 | (10-1) |
| FOMC Return | 13.05 | 17.40 | 30.66 | 51.74 | 68.23 | 55.17 |
| | (2.27) | (2.98) | (2.89) | (2.55) | (2.70) | (2.41) |
| Non-FOMC Return | 3.09 | 3.63 | 3.57 | 2.61 | 4.05 | 0.95 |
| | (3.25) | (3.10) | (2.30) | (1.02) | (1.33) | (0.34) |
| | | | Model | | | |
| | 1 | 2 | 3–8 | 9 | 10 | (10-1) |
| FOMC Return | 27.37 | 32.18 | 38.55 | 43.75 | 46.17 | 18.81 |
| | (1.98) | (2.02) | (2.05) | (2.02) | (2.00) | (1.81) |
| Non-FOMC Return | 0.82 | 1.09 | 1.66 | 2.23 | 2.54 | 1.72 |
| | (0.97) | (1.17) | (1.66) | (1.71) | (1.82) | (1.74) |
| $\mathbb{E}[\xi_i]$ | 2.59 | 3.13 | 4.02 | 4.84 | 5.33 | 2.74 |
| $\mathbb{E}[\eta_i]$ | 2.60 | 3.05 | 3.99 | 4.98 | 5.45 | 2.85 |

**Table 13**
CAPM for beta-sorted portfolios. This table documents the CAPM regression for beta-sorted portfolios in basis points. The top panel is based on data from January 1996 to December 2017 with 176 FOMC announcement days. We sort stocks based on the CAPM regression coefficient of a single stock excess return on the market excess return. We then run the CAPM regression on the market excess return, the non-FOMC dummy, and the FOMC dummy. We report CAPM beta and coefficients of the non-FOMC dummy and FOMC dummy on the decile portfolios, the long-short portfolio and the associated Newey-West *t*-statistics (in parentheses). The bottom panel reports the model-implied CAPM regression coefficients for decile portfolios. In the simulation, we use 400 stocks with 22 valid years (130 years with 100 years of burn-in and 8 years of pre-sample to estimate single stocks' CAPM beta coefficients), a total of 264 FOMC announcement days. We then sort these 400 stocks into 10 portfolios based on the estimated beta coefficients and report the mean portfolio $\xi$, the mean portfolio $\eta$ and the mean of the estimated CAPM beta, and the mean and the *t*-statistics of the coefficients of the non-FOMC and FOMC dummies. We simulate 500 independent daily sample paths and report the average results over all the sample paths.

| | | | Data | | | |
|---|---|---|---|---|---|---|
| | 1 | 2 | 3–8 | 9 | 10 | (10-1) |
| CAPM Beta | 0.47 | 0.68 | 1.02 | 1.58 | 1.73 | 1.26 |
| Non-FOMC Dummy | 1.13 | 1.19 | 0.32 | -1.92 | -0.86 | -1.99 |
| | (1.72) | (1.89) | (0.89) | (-1.68) | (-0.50) | (-0.93) |
| FOMC Dummy | -2.78 | -4.93 | -2.63 | 1.04 | 12.50 | 15.28 |
| | (-0.67) | (-1.47) | (-1.67) | (0.17) | (1.46) | (1.34) |
| | | | Model | | | |
| | 1 | 2 | 3–8 | 9 | 10 | (10-1) |
| CAPM Beta | 0.70 | 0.81 | 1.00 | 1.19 | 1.29 | 0.59 |
| Non-FOMC Dummy | -0.34 | -0.25 | -0.01 | 0.24 | 0.40 | 0.73 |
| | (-0.66) | (-0.49) | (-0.06) | (0.47) | (0.77) | (0.95) |
| FOMC Dummy | 0.92 | 1.52 | 0.36 | -1.69 | -2.91 | -3.82 |
| | (0.33) | (0.55) | (0.49) | (-0.61) | (-1.05) | (-0.92) |
| $\mathbb{E}[\xi_i]$ | 2.59 | 3.13 | 4.02 | 4.84 | 5.33 | 2.74 |
| $\mathbb{E}[\eta_i]$ | 2.60 | 3.05 | 3.99 | 4.98 | 5.45 | 2.85 |

for the pattern of announcement returns of CAPM beta-sorted portfolios quite well.

Table 12 presents the announcement premium for beta-sorted portfolios in the data and from the model. As in the data, the announcement premiums are significant and monotonically increasing in beta in our model. The average returns on non-FOMC days are much smaller for all portfolios, as are the spreads between these portfolios. In Table 13, we present the results for the CAPM regressions for beta-sorted portfolios in the model. As in the data, the coefficients for non-FOMC and FOMC dummies are both insignificant. The FOMC announcement premium in the long-short portfolio disappears once we control for the market beta. In our model, even though the CAPM does not hold exactly, beta is still a monotone function of risk exposure. As a result, CAPM alpha becomes insignificant in finite samples once we control for the market return.

## 6. Conclusion

In this paper, we provide empirical evidence and an equilibrium model for the cross section of FOMC announcement-day returns. We show that stocks that are more sensitive to monetary policy announcement surprises require a higher level of risk compensation following FOMC announcements. Our evidence is supportive of the recent literature that emphasizes the importance of risk compensation in macroeconomic announcements, in particular, monetary policy announcements. To account for the cross section of FOMC announcement returns, we develop an equilibrium model in which FOMC announcements reveal the Fed's private information about prospects for future economic growth and stock returns differ in their sensitivity to economic growth rates.

## Appendix A. Implied variance and data

In this section, we provide more details of the firm-level implied variance and other data we used in the paper. To measure the firm-level sensitivity to monetary policy announcement surprises around FOMC days, we use equity-options data from OptionMetrics for the period of January 1, 1996 to December 31, 2017. We exclude options with missing or negative bid-ask spread, zero bid, or zero open interest. We restrict the sample to out-of-the-money (OTM) options to estimate the model-free implied variance (Bakshi et al., 2003). To ensure our results are not driven by misleading prices, we follow Conrad et al. (2013) and exclude options that do not satisfy the standard option price bounds. We further remove options with a maturity less than three days. For a firm on a given day and a given maturity, we do not compute the implied variance if the number of OTM options is less than four.

Define $IV_t(\tau)$ as the time-$t$ price of the $\tau$-maturity quadratic payoff on the underlying stock, $IV_t(\tau) \equiv e^{-r_f\tau}\mathbb{E}_t^{\mathbb{Q}}\left[r_{t,t+\tau}^2\right]$, where $r_f$ is the continuously compounded interest rate. Bakshi et al. (2003) show $IV_t(\tau)$ can be recovered from the prices of OMT call and put options as follows:

$$IV_t(\tau) = \int_{S_t}^{\infty} \frac{1-\ln(K/S_t)}{K^2/2} C_t(\tau;K)dK + \int_0^{S_t} \frac{1+\ln(S_t/K)}{K^2/2} P_t(\tau;K)dK, \tag{A.1}$$

where $S_t$ is the price of underlying stock, and $C_t(\tau;K)$ and $P_t(\tau;K)$ are call and put prices with maturity $\tau$ and strike $K$, respectively.

We compute $IV_t(\tau)$ for each firm on each day and each day-to-maturity. In theory, computing $IV_t(\tau)$ requires a continuum of strike prices, whereas in practice, we only observe a discrete and finite set of them. Following Jiang and Tian (2005) and others, we discretize the integrals in Eq. (A.1) by setting up a total of 1001 grid points in the moneyness $(K/S_t)$ ranging from 1/3 to 3. First, we use cubic splines to interpolate the implied volatility inside the available moneyness range. Second, we extrapolate the implied volatility using the boundary values to fill the rest of the grid points. Third, we calculate option prices from these 1,001 implied volatilities using the formula of Black and Scholes (1973).[11] Next, we compute $IV_t(\tau)$ if the number of OTM options is more than four (e.g., Conrad et al., 2013 and others). Lastly, to obtain the seven days to maturity $IV_t(7)$ for a firm on a given day, we interpolate or extrapolate $IV_t(\tau)$ with available $\tau$. This process yields a daily time series of the risk-neutral expected quadratic payoff for each eligible firm with a fixed maturity of seven days. Due to the extrapolation, we find some negative values. We treat them as missing observations.

We obtain stock return data from the Center for Research in Security Prices (CRSP) and merge it with the OptionMetrics data. Our data period contains 6652 individual firms with traded options. In our empirical analysis, we only consider stocks that have a CRSP share code of 10 or 11, and we exclude stocks with a price less than \$5 or daily return larger than 500% or less than -500%. We also exclude stocks with annualized implied variance larger than 25.

Fama-French risk factors are from Kenneth French's Data library. Monetary policy news shocks and FFR shocks are from Nakamura and Steinsson (2018) and Acosta and Saia (2020).

The dates of FOMC meetings are from the website of the Fed. Following Savor and Wilson (2014), we only include the pre-scheduled FOMC meetings during our data period (1996–2017). About eight regularly pre-scheduled FOMC meetings occur each year. When the meeting lasts for two days, we consider the second day the FOMC announcement day. In total, our data period contains 176 FOMC announcement days. Among the 6652 firms in our sample, 5446 firms have at least one observed option-implied variance on these 176 FOMC announcement days.

## Appendix B. Details of the continuous-time model

### B1. A three-equation new Keynesian model

We assume aggregate output $Y_t = A_t N_t$, where the productivity $A_t = e^{\int_0^t \sigma dB_{A,s}}$ follows a log-normal distribution, where $B_{A,t}$ is the shock to productivity modeled as a Brownian motion, and $\sigma$ is the volatility. We guess, in equilibrium, $N_t$, the labor supply, is of the form $N_t = e^{\int_0^t \mu_s ds}$ for some stochastic process $\{\mu_t\}$. In this case, the growth rate of the aggregate output can be written as

$$\frac{dY_t}{Y_t} = \mu_t dt + \sigma dB_{A,t}. \tag{B.1}$$

Here, the growth rate of aggregate output is affected by both the growth rate of productivity through $\sigma dB_{A,t}$ and the growth rate of labor supply through $\mu_t dt$.

[11] We apply these steps to the calculation of individual risk-neutral expected quadratic payoffs. The individual equity options are American. Therefore, directly using the mid-quotes of individual options prices is inappropriate because the early exercise premium may confound our results. To avoid this issue, we use the implied volatilities provided by OptionMetrics. These implied volatilities are computed using a proprietary algorithm based on the Cox et al. (1979) model, which takes into account the early exercise premium.

The equilibrium in standard New Keynesian models is characterized by three equations.[12] The first equation is the household intertemporal Euler equation that relates the real interest rate to the expected consumption growth:

$$\hat{r}_t - \Pi_t = \frac{1}{\psi}\mu_t + cons, \tag{B.2}$$

where $\hat{r}_t$ is the nominal interest rate, $\Pi_t$ is the inflation, and $\mu_t$ is the expected consumption growth from Eq. (B.1). The household intertemporal Euler equation implies the sensitivity of expected consumption growth with respect to the real interest rate is equal to IES ($\psi$). The unspecified constant, denoted as *cons*, includes the time discount rate $\rho$ and the term that is related to precautionary savings. The second equation is the Philips curve:

$$d\Pi_t = (\kappa_0 + \kappa_1\mu_t)dt - \kappa_2 dB_{A,t} - \kappa_3 dB_{\theta,t}, \tag{B.3}$$

where $\kappa_0$, $\kappa_1$ $\kappa_2$, and $\kappa_3$ are positive constants that describe firms' optimal price-setting rule and $B_{\theta,t}$ is the shock to the Fed's interest-rate target. This equation captures the idea that in the presence of price rigidity, optimal pricing implies inflation increases when labor supply is high or when aggregate productivity is low. The last equation is the Taylor rule:

$$\hat{r}_t = \phi_0 + \phi_\pi \Pi_t + \phi_A B_{A,t} + \theta_t, \tag{B.4}$$

where $\phi_\pi$ and $\phi_A$ are the coefficients that determine the Fed's nominal interest-rate response to inflation and output gap, respectively. The stochastic process $\theta_t$ describes the Fed's interest-rate target and is assumed to be a continuous-time AR(1) process as in (10). We assume these coefficients satisfy the Blanchard and Kahn (1980) condition so that the equilibrium is unique.

To derive the equilibrium conditions, we guess and verify that the equilibrium inflation takes the form of

$$\Pi_t = -\chi_0 - \chi_1\theta_t - \kappa_2 B_{A,t}, \tag{B.5}$$

where

$$\chi_1 = \frac{\psi\kappa_1}{a + \psi\kappa_1(\phi_\pi - 1)}, \text{ and } \chi_0 = \frac{(\phi_0 - cons)\psi\kappa_1 + \kappa_0}{\psi\kappa_1(\phi_\pi - 1)}.$$

To verify Eq. (B.5) is an equilibrium, note that under the conjectured functional form of inflation, (B.5), the Taylor rule (B.4) implies the real interest rate is

$$\begin{aligned}\hat{r}_t - \Pi_t &= [\phi_0 - (\phi_\pi - 1)\chi_0] + [1 - (\phi_\pi - 1)\chi_1]\theta_t \\ &\quad + [\phi_A - (\phi_\pi - 1)\kappa_2]B_{A,t} \\ &= [\phi_0 - (\phi_\pi - 1)\chi_0] + [1 - (\phi_\pi - 1)\chi_1]\theta_t,\end{aligned}$$

where we assume $\kappa_2 = \frac{\phi_A}{\phi_\pi - 1}$. Using the household intertemporal Euler Eq. (B.2), we have

$$[\phi_0 - (\phi_\pi - 1)\chi_0] + [1 - (\phi_\pi - 1)\chi_1]\theta_t = \frac{1}{\psi}\mu_t + cons,$$

which implies the equilibrium $\mu_t$ must satisfy

$$\mu_t = \psi\{[\phi_0 - cons - (\phi_\pi - 1)\chi_0] + [1 - (\phi_\pi - 1)\chi_1]\theta_t\}.$$

[12] See Gali (2008) for a discrete-time version of the three-equation model.

Mapping back into Eq. (9) in the main text, we have

$$\chi_0 = \frac{\phi_0 - cons}{\phi_\pi - 1}, \text{ and } \chi_1 = \frac{1 - 1/\psi}{\phi_\pi - 1}. \tag{B.6}$$

Finally, Eqs. (B.5) and (10) imply

$$\begin{aligned}d\Pi_t &= -\chi_1 d\theta_t - \kappa_2 dB_{A,t} = -\chi_1\left[a(\mu - \theta_t)dt + \sigma_\theta dB_{\theta,t}\right] \\ &\quad - \kappa_2 dB_{A,t}.\end{aligned}$$

By matching the coefficients, we verify that the conjectured functional form of inflation must satisfy the Philips curve (B.3) as long as

$$\kappa_0 = -\frac{a\mu(1 - 1/\psi)}{\phi_\pi - 1}, \kappa_1 = \frac{a(1 - 1/\psi)}{\psi(\phi_\pi - 1)}, \\ \text{and } \kappa_3 = \frac{\sigma_\theta(1 - 1/\psi)}{\phi_\pi - 1}. \tag{B.7}$$

*B2. The value function of the representative agent*

In this subsection, we start from the solution to the posterior belief of a representative agent. We then derive the solution to the value function and the associated boundary condition at the announcement.

Assume pre-determined announcements occur every period at time $nT$ $(n = 1, 2, \ldots)$. On non-announcement days, investors solve optimization problems in the interior $((n-1)T^+, nT^-))$. On the announcement days, investors solve the optimization problems at the boundary $nT$. For simplicity, we focus on an equilibrium in which all announcement cycles are identical. Under this assumption, we only need to characterize one representative announcement cycle, $[0, T]$. We denote 0 (or equivalently $nT^+$) as the time right *after* the announcements and $T$ (or equivalently $nT^-$) as the moment right *before* the announcements. *Posterior variance*

Because announcements provide information containing the true value of $\theta_t$ at $nT$, the posterior variance after each announcement drops from $q_{nT}^-$ to $q_{nT}^+ = q_0$. In the interior of $(0, T)$, the standard optimal filtering [see Theorem 10.3 from Liptser and Shiryaev (2001)] implies the posterior mean and variance of $\theta_t$ are given by Eqs. (11) and (12). The general closed-form solution for $q_t$ is given by

$$q(t) = \frac{\sigma_\theta^2\left(1 - he^{-2\hat{a}t}\right)}{\left(\hat{a} - a\right)he^{-2\hat{a}t} + a + \hat{a}}, \tag{B.8}$$

where $\hat{a} = \sqrt{a^2 + (\sigma_\theta/\sigma)^2}$ and the constant $h = \frac{\sigma_\theta^2 + (\hat{a} - a)q_0}{\sigma_\theta^2 - (\hat{a} + a)q_0}$ is chosen to satisfy the initial condition $q(0) = q_0$.

*Preference*

Using the results from Duffie and Epstein (1992), the representative agent's preference is specified by a pair of aggregators $(f, \mathcal{A})$ such that the utility of the representative agent, $V_t$, is the solution to the following SDE:

$$dV_t = [-f(C_t, V_t) - \frac{1}{2}\mathcal{A}(V_t)||\sigma_V(t)||^2]dt + \sigma_V(t)dB_t,$$

for some square-integrable process $\sigma_V(t)$. We adopt the convenient normalization $\mathcal{A}(V) = 0$ as Duffie and Epstein (1992), where the normalized aggregator $\bar{f}$ is,

$$\bar{f}(C_t, V_t) = \frac{\rho}{1 - 1/\psi}\frac{C_t^{1-1/\psi} - ((1-\gamma)V_t)^{\frac{1-1/\psi}{1-\gamma}}}{((1-\gamma)V_t)^{\frac{1-1/\psi}{1-\gamma} - 1}} \tag{B.9}$$

for $\psi \neq 1$. Because our model does not have investment, we use consumption $C_t$ to replace output $Y_t$ in the rest of the appendix. The Hamilton–Jacobi–Bellman (HJB) equation for the recursive utility satisfies

$$\bar{f}\left(C_t, V\left(\hat{\theta}_t, t, C_t\right)\right) + \mathcal{L}\left[V\left(\hat{\theta}_t, t, C_t\right)\right] = 0, \tag{B.10}$$

where $\mathcal{L}$ is the infinitesimal generator and $\mathcal{L}[V_t] = \lim_{\Delta\to 0}\frac{1}{\Delta}\mathbb{E}_t[V_{t+\Delta} - V_t]$. Due to the homogeneity, considering the value function of the form

$$V\left(\hat{\theta}_t, t, C_t\right) = \frac{1}{1-\gamma}H\left(\hat{\theta}_t, t\right)C_t^{1-\gamma}, \tag{B.11}$$

we show the following lemma must hold:

*Lemma 1.* $H\left(\hat{\theta}_t, t\right)$ *satisfies the following HJB equation*

$$\begin{aligned} 0 = {} & \frac{\rho}{1-\frac{1}{\psi}}\left(H\left(\hat{\theta}_t, t\right)^{-\frac{1-\frac{1}{\psi}}{1-\gamma}} - 1\right) + \left(\hat{\theta}_t - \frac{1}{2}\gamma\sigma^2\right) \\ & + \frac{1}{1-\gamma}\frac{H_t\left(\hat{\theta}_t, t\right)}{H\left(\hat{\theta}_t, t\right)} \\ & + \left[\frac{1}{1-\gamma}a\left(\mu - \hat{\theta}_t\right) + q_t\right]\frac{H_\theta\left(\hat{\theta}_t, t\right)}{H\left(\hat{\theta}_t, t\right)} \\ & + \frac{1}{2(1-\gamma)}\frac{H_{\theta\theta}\left(\hat{\theta}_t, t\right)}{H\left(\hat{\theta}_t, t\right)}\left(\frac{q_t}{\sigma}\right)^2. \end{aligned} \tag{B.12}$$

*Proof.* Given (B.9), we have $\bar{f}(C_t, V) = \frac{\rho}{1-\frac{1}{\psi}}C_t^{1-\gamma}\left[H\left(\hat{\theta}_t, t\right)^{1-\frac{1-\frac{1}{\psi}}{1-\gamma}} - H\left(\hat{\theta}_t, t\right)\right]$. Furthermore, using Ito's lemma we get

$$\begin{aligned} \frac{d\left[H\left(\hat{\theta}_t, t\right)C_t^{1-\gamma}\right]}{C_t^{1-\gamma}} = {} & (1-\gamma)H\left(\hat{\theta}_t, t\right)\left(\hat{\theta}_t dt + \sigma d\tilde{B}_{A,t}\right) \\ & - \frac{1}{2}\gamma(1-\gamma)H\left(\hat{\theta}_t, t\right)\sigma^2 dt \\ & + H_t\left(\hat{\theta}_t, t\right)dt + H_\theta\left(\hat{\theta}_t, t\right)\left[a\left(\mu - \hat{\theta}_t\right)dt + \frac{q_t}{\sigma}d\tilde{B}_{A,t}\right] \\ & + \frac{1}{2}H_{\theta\theta}\left(\hat{\theta}_t, t\right)\left(\frac{q_t}{\sigma}\right)^2 dt + (1-\gamma)H_\theta\left(\hat{\theta}_t, t\right)q_t dt, \end{aligned}$$

$$\begin{aligned} \frac{\mathcal{L}\left[V\left(\hat{\theta}_t, t, C_t\right)\right]}{C_t^{1-\gamma}} = {} & \frac{\mathcal{L}\left[H\left(\hat{\theta}_t, t\right)C_t^{1-\gamma}\right]}{(1-\gamma)C_t^{1-\gamma}} \\ = {} & \left(\hat{\theta}_t - \frac{1}{2}\gamma\sigma^2\right)H\left(\hat{\theta}_t, t\right) \\ & + \frac{1}{(1-\gamma)}\left[H_t\left(\hat{\theta}_t, t\right) + H_\theta\left(\hat{\theta}_t, t\right)a\left(\mu - \hat{\theta}_t\right) \right. \\ & \left. + \frac{1}{2}H_{\theta\theta}\left(\hat{\theta}_t, t\right)\left(\frac{q_t}{\sigma}\right)^2\right] + H_\theta\left(\hat{\theta}_t, t\right)q_t. \end{aligned}$$

Therefore, divide both sides of (B.10) by $C_t^{1-\gamma}H\left(\hat{\theta}_t, t\right)$, we get (B.12). □

We approximate $H\left(\hat{\theta}_t, t\right)$ by an exponential linear form:

$$H\left(\hat{\theta}_t, t\right) = e^{B\hat{\theta}_t + h(t)}, \tag{B.13}$$

where B satisfies Eq. (B.14) and $h(t)$ is characterized by the ordinary differential Eq. (B.15) with the boundary condition (B.17). Therefore, $H_t\left(\hat{\theta}_t, t\right) = H(\theta_t, t)h'(t)$, $H_\theta\left(\hat{\theta}_t, t\right) = BH(\theta_t, t)$, $H_{\theta\theta}\left(\hat{\theta}_t, t\right) = B^2H(\theta_t, t)$. Substituting them into (B.12), we would get

$$\begin{aligned} 0 = {} & \frac{\rho(1-\gamma)}{1-\frac{1}{\psi}}\left(e^{-\frac{1-\frac{1}{\psi}}{1-\gamma}\left[B\hat{\theta}_t + h(t)\right]} - 1\right) \\ & + (1-\gamma)\left(\hat{\theta}_t - \frac{1}{2}\gamma\sigma^2\right) + h'(t) \\ & + \left[a\left(\mu - \hat{\theta}_t\right) + (1-\gamma)q_t\right]B + \frac{1}{2}B^2\left(\frac{q_t}{\sigma}\right)^2. \end{aligned}$$

Using $e^z - 1 \approx z$ to approximate and simplify, the above equation becomes

$$\begin{aligned} 0 = {} & -\rho\left[B\hat{\theta}_t + h(t)\right] + (1-\gamma)\left(\hat{\theta}_t - \frac{1}{2}\gamma\sigma^2\right) + h'(t) \\ & + \left[a\left(\mu - \hat{\theta}_t\right) + (1-\gamma)q_t\right]B + \frac{1}{2}B^2\left(\frac{q_t}{\sigma}\right)^2. \end{aligned}$$

Matching the coefficients of $\hat{\theta}_t$ and $t$ yields

$$B = \frac{1-\gamma}{a+\rho}, \tag{B.14}$$

$$\begin{aligned} h'(t) = {} & \rho h(t) + \frac{1}{2}\gamma(1-\gamma)\sigma^2 - Ba\mu - (1-\gamma)Bq_t \\ & - \frac{1}{2}B^2\left(\frac{q_t}{\sigma}\right)^2. \end{aligned} \tag{B.15}$$

*Boundary conditions*

The boundary condition satisfies:

$$H\left(\hat{\theta}_{nT}^-, nT^-\right) = \mathbb{E}\left[H\left(\hat{\theta}_{nT}^+, nT^+\right)|\hat{\theta}_{nT}^-, q_{nT}^-\right], \quad n = 1, 2, \ldots, \tag{B.16}$$

where $\hat{\theta}_{nT}^+ \sim \mathcal{N}\left(\hat{\theta}_{nT}^-, q_{nT}^- - q_{nT}^+\right)$ comes from Eq. (13). The intuition is that the continuation value of the value function right before the announcement must equal to its expected value right after the announcement, conditional on the information at $nT^-$ before the announcement.

Using the conjectured functional form, the boundary condition can be rewritten as,

$$e^{B\hat{\theta}_T + h(T)} = \mathbb{E}\left[e^{B\hat{\theta}_0 + h(0)}|\hat{\theta}_T, q_T\right] = e^{B\hat{\theta}_T + \frac{1}{2}B^2(q_T - q_0) + h(0)},$$

which gives

$$h(T) = h(0) + \frac{1}{2}B^2(q_T - q_0). \tag{B.17}$$

Eqs. (B.15) and (B.17) could be jointly used to solve the $h(t)$ function in closed form. Note we do not need the functional form of $h(t)$ to solve for asset prices. However, we would need it to compute welfare gains.

*B3. Asset prices*

In this subsection, we first derive the pricing kernel and the risk-free rate. We then present the price-to-dividend ratio for each firm with the boundary condition at the announcement. Finally, we calculate the cumulative return and the risk premium.

*State price density and the risk-free rate*

*Lemma 2. In the interior of* $(0, T)$*, the law of motion of the state price density,* $\pi_t$*, satisfies the SDE of Eq. (17), where the risk-free interest rate is*

$$r\left(\hat{\theta}_t, t\right) = \rho + \frac{1}{\psi}\hat{\theta}_t - \frac{1}{2}\gamma\left(1 + \frac{1}{\psi}\right)\sigma^2 + \frac{\frac{1}{\psi} - \gamma}{1-\gamma}\frac{H_\theta\left(\hat{\theta}_t, t\right)}{H\left(\hat{\theta}_t, t\right)}q_t + \frac{\left(\frac{1}{\psi} - \gamma\right)\left(1 - \frac{1}{\psi}\right)}{2(1-\gamma)^2}\left(\frac{H_\theta\left(\hat{\theta}_t, t\right)}{H\left(\hat{\theta}_t, t\right)}\frac{q_t}{\sigma}\right)^2, \tag{B.18}$$

*and the market price of the productivity shock is*

$$\sigma_\pi(t) = \gamma\sigma - \frac{\frac{1}{\psi} - \gamma}{1-\gamma}\frac{H_\theta\left(\hat{\theta}_t, t\right)}{H\left(\hat{\theta}_t, t\right)}\frac{q_t}{\sigma}. \tag{B.19}$$

*Proof.* Pricing kernel is defined as

$$\frac{d\pi_t}{\pi_t} = \frac{d\bar{f}_C(C, V)}{f_C(C, V)} + \bar{f}_V(C, V)dt, \tag{B.20}$$

where $\bar{f}_C(C,V) = \rho H\left(\hat{\theta}_t, t\right)^{\frac{\frac{1}{\psi}-\gamma}{1-\gamma}} C_t^{-\gamma}$ and $\bar{f}_V(C, V) = \rho\frac{\frac{1}{\psi}-\gamma}{1-\frac{1}{\psi}}H\left(\hat{\theta}_t, t\right)^{-\frac{1-\frac{1}{\psi}}{1-\gamma}} - \rho\frac{1-\gamma}{1-\frac{1}{\psi}}$. Applying Ito's lemma,

$$\begin{aligned}
\frac{d\bar{f}_C(C, V)}{f_C(C, V)} &= \frac{d[H^{\frac{\frac{1}{\psi}-\gamma}{1-\gamma}}C_t^{-\gamma}]}{H^{\frac{\frac{1}{\psi}-\gamma}{1-\gamma}}C_t^{-\gamma}} \\
&= \Bigg\{-\gamma\hat{\theta}_t + \frac{1}{2}\gamma(\gamma+1)\sigma^2 + \frac{\frac{1}{\psi}-\gamma}{1-\gamma}\frac{H_t}{H} \\
&\quad + \frac{\frac{1}{\psi}-\gamma}{1-\gamma}\frac{H_\theta}{H}a\left(\mu - \hat{\theta}_t\right) \\
&\quad + \frac{1}{2}\left[\frac{\left(\frac{1}{\psi}-\gamma\right)\left(\frac{1}{\psi}-1\right)}{(1-\gamma)^2}\left(\frac{H_\theta}{H}\right)^2 + \frac{\frac{1}{\psi}-\gamma}{1-\gamma}\frac{H_{\theta\theta}}{H}\right]\left(\frac{q_t}{\sigma}\right)^2 \\
&\quad - \gamma\frac{\frac{1}{\psi}-\gamma}{1-\gamma}\frac{H_\theta}{H}q_t\Bigg\}dt + \left[-\gamma\sigma + \frac{\frac{1}{\psi}-\gamma}{1-\gamma}\frac{H_\theta}{H}\frac{q_t}{\sigma}\right]d\tilde{B}_{A,t}.
\end{aligned}$$

Matching the drifts and diffusions of (17) and (B.20), we can get (B.19) and

$$\begin{aligned}
r\left(\hat{\theta}_t, t\right) &= -\frac{\frac{1}{\psi}-\gamma}{1-\gamma}\Bigg\{\frac{H_t}{H} + \frac{H_\theta}{H}a\left(\mu - \hat{\theta}_t\right) \\
&\quad + \frac{1}{2}\left[\frac{\frac{1}{\psi}-1}{1-\gamma}\left(\frac{H_\theta}{H}\right)^2 + \frac{H_{\theta\theta}}{H}\right]\left(\frac{q_t}{\sigma}\right)^2 - \gamma\frac{H_\theta}{H}q_t\Bigg\} \\
&\quad + \gamma\hat{\theta}_t - \frac{1}{2}\gamma(\gamma+1)\sigma^2 - \rho\frac{\frac{1}{\psi}-\gamma}{1-\frac{1}{\psi}}H^{-\frac{1-\frac{1}{\psi}}{1-\gamma}} \\
&\quad + \rho\frac{1-\gamma}{1-\frac{1}{\psi}}.
\end{aligned} \tag{B.21}$$

Using the HJB equation to simplify $r\left(\hat{\theta}_t, t\right)$ by multiplying $\left(\frac{1}{\psi} - \gamma\right)$ on both sides of (B.12),

$$\begin{aligned}
0 &= \rho\frac{\frac{1}{\psi}-\gamma}{1-\frac{1}{\psi}}\left(H^{-\frac{1-\frac{1}{\psi}}{1-\gamma}} - 1\right) + \left(\frac{1}{\psi} - \gamma\right)\left(\hat{\theta}_t - \frac{1}{2}\gamma\sigma^2\right) \\
&\quad + \frac{\frac{1}{\psi}-\gamma}{1-\gamma}\frac{H_t}{H} \\
&\quad + \left(\frac{1}{\psi} - \gamma\right)\left[\frac{1}{1-\gamma}a\left(\mu - \hat{\theta}_t\right) + q_t\right]\frac{H_\theta}{H} \\
&\quad + \frac{\frac{1}{\psi}-\gamma}{2(1-\gamma)}\frac{H_{\theta\theta}}{H}\left(\frac{q_t}{\sigma}\right)^2,
\end{aligned}$$

and adding up with (B.21), we can get Eq. (B.18). Finally, substituting back the approximated $H\left(\hat{\theta}_t, t\right)$ defined in Eq. (B.13), we will get (18) and (19) in the main text. □

*Price-to-dividend ratio* The present value relationship (15) implies

$$\pi_t D_t^i + \lim_{\Delta\to 0}\frac{1}{\Delta}\left\{\mathbb{E}_t\left[\pi_{t+\Delta}p^i\left(\hat{\theta}_{t+\Delta}, t+\Delta\right)D_{t+\Delta}^i\right] - \pi_t p^i\left(\hat{\theta}_t, t\right)D_t^i\right\} = 0. \tag{B.22}$$

The above can be used to show the solution for $p^i\left(\hat{\theta}_t, t\right)$ satisfies the following lemma.

*Lemma 3. In the interior* $(0, T)$*, the price-to-dividend ratio of firm i,* $p^i\left(\hat{\theta}_t, t\right)$ *must satisfy the following PDE:*

$$1 - p^i\left(\hat{\theta}_t, t\right)\varpi^i\left(\hat{\theta}_t, t\right) + p_t^i\left(\hat{\theta}_t, t\right) - p_\theta^i\left(\hat{\theta}_t, t\right)\nu^i\left(\hat{\theta}_t, t\right) + \frac{1}{2}p_{\theta\theta}^i\left(\hat{\theta}_t, t\right)\left(\frac{q_t}{\sigma}\right)^2 = 0, \tag{B.23}$$

*where* $\varpi^i\left(\hat{\theta}_t, t\right)$ *and* $\nu^i\left(\hat{\theta}_t, t\right)$ *are defined by*

$$\begin{aligned}
\varpi^i\left(\hat{\theta}_t, t\right) &= \rho - \frac{1}{2}\gamma\left(1 + \frac{1}{\psi}\right)\sigma^2 + \gamma\sigma^2\eta_i \\
&\quad + \frac{1}{\psi}\hat{\theta}_t - \mu - \xi_i\left(\hat{\theta}_t - \mu\right) \\
&\quad + \frac{\frac{1}{\psi}-\gamma}{a+\rho}q_t(1-\eta_i) + \frac{1}{2}\frac{\left(\frac{1}{\psi}-\gamma\right)\left(1-\frac{1}{\psi}\right)}{(a+\rho)^2}\left(\frac{q_t}{\sigma}\right)^2, \\
\nu^i\left(\hat{\theta}_t, t\right) &= a\left(\hat{\theta}_t - \mu\right) + (\gamma - \eta_i)q_t - \frac{\frac{1}{\psi}-\gamma}{a+\rho}\left(\frac{q_t}{\sigma}\right)^2,
\end{aligned}$$

*with the boundary condition at the announcement satisfying*

$$p^i\left(\hat{\theta}_T, T\right) = \frac{\mathbb{E}\left[e^{\frac{\frac{1}{\psi}-\gamma}{a+\rho}\hat{\theta}_0}p^i\left(\hat{\theta}_0, 0\right) \mid \hat{\theta}_T, q_T\right]}{e^{\frac{\frac{1}{\psi}-\gamma}{a+\rho}\hat{\theta}_T + \frac{(1-\gamma)\left(\frac{1}{\psi}-\gamma\right)}{2(a+\rho)^2}(q_T - q_0)}}. \tag{B.24}$$

*Proof.* Eq. (B.22) implies $1 + p^i\left(\hat{\theta}_t, t\right)\frac{\mathcal{L}\left[\pi_t p^i\left(\hat{\theta}_t, t\right) D_t^i\right]}{\pi_t p^i\left(\hat{\theta}_t, t\right) D_t^i} = 0$. Using Eqs. (14) and (17), we have

$$\begin{aligned}
\frac{\mathcal{L}\left[\pi_t p^i\left(\hat{\theta}_t, t\right) D_t^i\right]}{\pi_t p^i\left(\hat{\theta}_t, t\right) D_t^i} &= -r\left(\hat{\theta}_t, t\right) \\
&+ \frac{1}{p^i}\left[p_t^i + p_\theta^i a\left(\mu - \hat{\theta}_t\right) + \frac{1}{2} p_{\theta\theta}^i \frac{q_t^2}{\sigma^2}\right] + \frac{p_\theta^i}{p^i} q_t \eta_i \\
&+ \left[\mu + \xi_i\left(\hat{\theta}_t - \mu\right)\right] - \left(\gamma\sigma - \frac{\frac{1}{\psi} - \gamma}{1-\gamma}\frac{H_\theta}{H}\frac{q_t}{\sigma}\right) \\
&\times \left[\eta_i \sigma + \frac{p_\theta^i}{p^i}\frac{q_t}{\sigma}\right].
\end{aligned}$$

Then we can derive the PDE for firm $i$'s price-to-dividend ratio as

$$\begin{aligned}
0 = &-\rho - \frac{1}{\psi}\hat{\theta}_t + \frac{1}{2}\gamma\left(1 + \frac{1}{\psi}\right)\sigma^2 + \frac{\gamma - \frac{1}{\psi}}{1-\gamma}\frac{H_\theta}{H} q_t \\
&- \frac{1}{2}\frac{\left(\frac{1}{\psi} - \gamma\right)\left(1 - \frac{1}{\psi}\right)}{(1-\gamma)^2}\left(\frac{H_\theta}{H}\frac{q_t}{\sigma}\right)^2 + \frac{p_\theta^i}{p^i} q_t \eta_i + \mu \\
&+ \xi_i\left(\hat{\theta}_t - \mu\right) + \frac{1}{p^i}\left[1 + p_t^i + p_\theta^i a\left(\mu - \hat{\theta}_t\right) + \frac{1}{2} p_{\theta\theta}^i \frac{q_t}{\sigma^2}\right] \\
&- \left(\gamma\sigma - \frac{\frac{1}{\psi} - \gamma}{1-\gamma}\frac{H_{\theta\theta}}{H}\frac{q_t}{\sigma}\right)\left[\eta_i \sigma + \frac{p_\theta^i}{p^i}\frac{q_t}{\sigma}\right].
\end{aligned}$$

Using the approximated form $\frac{H_\theta}{H} = \frac{1-\gamma}{a+\rho}$, we can simplify the above to get (B.23).

We next solve the boundary condition. In general, under the recursive utility, the SDF for a small interval $\Delta$ is,

$$SDF_{t,t+\Delta} = e^{-\rho\Delta}\left(\frac{C_{t+\Delta}}{C_t}\right)^{-\frac{1}{\psi}}\left[\frac{W_{t+\Delta}}{\left(\mathbb{E}\left[W_{t+\Delta}^{1-\gamma}\right]\right)^{\frac{1}{1-\gamma}}}\right]^{\frac{1}{\psi}-\gamma},$$

where $W_t = \left[(1-\gamma)V\left(\hat{\theta}, t, C_t\right)\right]^{\frac{1}{1-\gamma}}$. Thus,

$$SDF_{t,t+\Delta} = e^{-\rho\Delta}\left(\frac{C_{t+\Delta}}{C_t}\right)^{-\frac{1}{\psi}}\left[\frac{\left(H_{t+\Delta}C_{t+\Delta}^{1-\gamma}\right)^{\frac{1}{1-\gamma}}}{\left(\mathbb{E}\left[H_{t+\Delta}C_{t+\Delta}^{1-\gamma}\right]\right)^{\frac{1}{1-\gamma}}}\right]^{\frac{1}{\psi}-\gamma}.$$

At the announcement, $\Delta \to 0$, the consumption is continuous while the continuation value jumps (Ai and Bansal, 2018); therefore,

$$SDF_{t,t+\Delta} = \frac{H_{t+\Delta}^{\frac{\frac{1}{\psi}-\gamma}{1-\gamma}}}{\left(\mathbb{E}[H_{t+\Delta}]\right)^{\frac{\frac{1}{\psi}-\gamma}{1-\gamma}}} = \frac{e^{\frac{\frac{1}{\psi}-\gamma}{a+\rho}\hat{\theta}_t}}{\left(\mathbb{E}\left[e^{\frac{1-\gamma}{a+\rho}\hat{\theta}_t}\right]\right)^{\frac{\frac{1}{\psi}-\gamma}{1-\gamma}}}.$$

Under our simplified notation, $p^i\left(\hat{\theta}_0, 0\right) = p^i\left(\hat{\theta}_{nT}^+, nT^+\right)$ and $p^i\left(\hat{\theta}_T, T\right) = p^i\left(\hat{\theta}_{nT}^-, nT^-\right)$, where $\hat{\theta}_0 \sim \mathcal{N}\left(\hat{\theta}_T, q_T - q_0\right)$ from Eq. (13). Using the above SDF, the boundary condition for the price-to-dividend ratio at the announcement $nT$ can be derived as

$$\begin{aligned}
p^i\left(\hat{\theta}_T, T\right) &= \mathbb{E}\left[\frac{e^{\frac{\frac{1}{\psi}-\gamma}{a+\rho}\hat{\theta}_0} p^i\left(\hat{\theta}_0, 0\right)}{\left(\mathbb{E}\left[e^{\frac{1-\gamma}{a+\rho}\hat{\theta}_0} \mid \hat{\theta}_T, q_T\right]\right)^{\frac{\frac{1}{\psi}-\gamma}{1-\gamma}}}\middle|\hat{\theta}_T, q_T\right] \\
&= \frac{\mathbb{E}\left[e^{\frac{\frac{1}{\psi}-\gamma}{a+\rho}\hat{\theta}_0} p^i\left(\hat{\theta}_0, 0\right) \mid \hat{\theta}_T, q_T\right]}{\left(e^{\frac{1-\gamma}{a+\rho}\hat{\theta}_T + \frac{1}{2}\left(\frac{1-\gamma}{a+\rho}\right)^2 (q_T - q_0)}\right)^{\frac{\frac{1}{\psi}-\gamma}{1-\gamma}}} \\
&= \frac{\mathbb{E}\left[e^{\frac{\frac{1}{\psi}-\gamma}{a+\rho}\hat{\theta}_0} p^i\left(\hat{\theta}_0, 0\right) \mid \hat{\theta}_T, q_T\right]}{e^{\frac{\frac{1}{\psi}-\gamma}{a+\rho}\hat{\theta}_T + \frac{(1-\gamma)\left(\frac{1}{\psi}-\gamma\right)}{2(a+\rho)^2}(q_T - q_0)}},
\end{aligned}$$

which corresponds to Eqs. (20) and (B.24). □

Combining the PDE with the boundary condition in Lemma 3 would finally pin down the solutions for each firm's price-to-dividend ratio. In Section B.24, we show details about how to numerically solve the price-to-dividend ratio. *Risk premium*

We characterize the return process and risk premium by the following lemma.

*Lemma 4. The cumulative return follows the SDE of*

$$\frac{dR_t^i}{R_t^i} = \mu_{R,t}^i dt + \sigma_{R,t}^i d\tilde{B}_{A,t} + \sigma_i dB_{i,t}, \tag{B.25}$$

*where $\mu_{R,t}^i$ and $\sigma_{R,t}^i$ are the risky asset return and volatility for firm i, respectively:*

$$\begin{aligned}
\mu_{R,t}^i = &\frac{1}{p^i\left(\hat{\theta}_t, t\right)}\left[1 + p_t^i\left(\hat{\theta}_t, t\right) + p_\theta^i\left(\hat{\theta}_t, t\right) a\left(\mu - \hat{\theta}_t\right)\right. \\
&\left. + \frac{1}{2} p_{\theta\theta}^i\left(\hat{\theta}_t, t\right)\frac{q_t^2}{\sigma^2}\right] \\
&+ \mu + \xi_i\left(\hat{\theta}_t - \mu\right) + \frac{p_\theta^i\left(\hat{\theta}_t, t\right)}{p^i\left(\hat{\theta}_t, t\right)} q_t \eta_i,
\end{aligned} \tag{B.26}$$

$$\sigma_{R,t}^i = \eta_i \sigma + \frac{p_\theta^i\left(\hat{\theta}_t, t\right)}{p^i\left(\hat{\theta}_t, t\right)}\frac{q_t}{\sigma}. \tag{B.27}$$

*In the interior of $(0, T)$, the instantaneous risk premium is*

$$\mu_{R,t} - r\left(\hat{\theta}_t, t\right) = \left[\gamma\sigma - \frac{\frac{1}{\psi} - \gamma}{a+\rho}\frac{q_t}{\sigma}\right]\left[\eta_i \sigma + \frac{p_\theta^i\left(\hat{\theta}_t, t\right)}{p^i\left(\hat{\theta}_t, t\right)}\frac{q_t}{\sigma}\right]. \tag{B.28}$$

The above lemma could be applied to calculate the local CAPM beta of Eq. (27).

*Proof.* The cumulative return is defined as

$$\frac{dR_t^i}{R_t^i} = \frac{1}{p^i\left(\hat{\theta}_t, t\right) D_t^i}\left[D_t^i dt + d\left[p^i\left(\hat{\theta}_t, t\right) D_t^i\right]\right]. \tag{B.29}$$

Applying Ito's lemma, we have

$$\frac{d\left[p^i\left(\hat{\theta}_t, t\right)D_t^i\right]}{p^i\left(\hat{\theta}_t, t\right)D_t^i} = \left\{\frac{1}{p^i}\left[p_t^i + p_\theta^i a\left(\mu - \hat{\theta}_t\right) + \frac{1}{2}p_{\theta\theta}^i\frac{q_t^2}{\sigma^2}\right]\right.$$

$$\left. + \mu + \xi_i\left(\hat{\theta}_t - \mu\right) + \frac{p_\theta^i}{p^i}q_t\eta_i\right\}dt$$

$$+ \left(\eta_i\sigma + \frac{p_\theta^i}{p^i}\frac{q_t}{\sigma}\right)d\tilde{B}_{A,t} + dB_{i,t}.$$

Matching the drift and diffusion terms with Eq. (B.25), we can get (B.26) and (B.27). The instantaneous risk premium (B.28) can be obtained from

$$\mu_{R,t}^i - r\left(\hat{\theta}_t, t\right) = -Cov_t\left[\frac{d\left[p^i\left(\hat{\theta}_t, t\right)D_t^i\right]}{p^i\left(\hat{\theta}_t, t\right)D_t^i}, \frac{d\pi_t}{\pi_t}\right].$$

□

*B4. Expected sensitivity and implied variance reduction*

In this subsection, we derive the model-implied expected sensitivity and the implied variance reduction and show both are linear functions of $\xi_i$. Hence, EVR in our model perfectly identifies the structural parameter $\xi_i$, which determines the sensitivity of stock returns to monetary policy announcements.

We approximate the price-to-dividend ratio by the exponential linear form:

$$p\left(\hat{\theta}_t, t\right) = e^{A\hat{\theta}_t + g(t)}, \tag{B.30}$$

where $A$ satisfies Eq. (B.31). Substituting this into (B.23) yields,

$$e^{-A\hat{\theta}_t - g(t)} - \varpi\left(\hat{\theta}_t, t\right) + g'(t) - A\nu\left(\hat{\theta}_t, t\right) + \frac{1}{2}A^2\left(\frac{q_t}{\sigma}\right)^2 = 0.$$

Using a first-order approximation of term $e^{-A\hat{\theta}_t - g(t)}$ around the steady level of $\hat{\theta}_t$, $\mu$, $e^{-A\hat{\theta}_t - g(t)} \approx \bar{p} - A\bar{p}\left(\hat{\theta}_t - \mu\right)$, where $\bar{p}$ is the inverse of the steady-state price-to-dividend ratio. Matching the coefficient on $\hat{\theta}_t$ in Eq. (B.4) yields,

$$A = \frac{\xi - \frac{1}{\psi}}{a + \bar{p}}. \tag{B.31}$$

Using Eq. (16), because the dividend is continuous, the log announcement return is written as

$$\ln R_{nT^-, nT^+} = \ln p\left(\hat{\theta}_{nT}^+, nT^+\right) - \ln p\left(\hat{\theta}_{nT}^-, nT^-\right).$$

Eqs. (22) and (24) follow immediately from the expression of $A$ in (B.31),

$$\Delta IV_i = Var\left[\ln R_{nT^-, nT^+}^i \mid \hat{\theta}_{nT}^-, q_{nT}^-\right] = \left(\frac{\xi_i - \frac{1}{\psi}}{a + \bar{p}^i}\right)^2\left(q_{nT}^- - q_{nT}^+\right)$$

$$= \left(\frac{\xi_i - \frac{1}{\psi}}{a + \bar{p}^i}\right)^2\left(\frac{q_{nT}^{-2}}{q_{nT}^- + \sigma_s^2}\right), \tag{B.32}$$

where the last equivalence we used is Eq. (13).

## Supplementary material

Supplementary material associated with this article can be found, in the online version, at doi:10.1016/j.jfineco.2021.07.002.

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