Show Me the Money: The Monetary Policy Risk Premium*

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Abstract

Measuring the monetary policy exposure of an individual stock is challenging because the large idiosyncratic volatility of individual stock returns leads to imprecise estimates of the exposure. In this study, we generate a parsimonious monetary policy exposure (MPE) index based on observable firm characteristics that are likely to drive the exposure of firms to monetary policy. Using this index, we find that stocks whose prices react more positively to expansionary monetary policy shocks earn lower average returns, consistent with the intuition that monetary policy is expansionary in bad times when marginal value of wealth is high and thereby provides a hedge for the investors of high MPE stocks. A long-short trading strategy designed to exploit this effect achieves an annualized return of 10.56% (t-statistic of 6.19), an annualized Sharpe Ratio of 0.98, and a Fama-French-Carhart alpha of 7.56% (t-statistic 6.51) between 1975 and 2014. Our findings build a bridge between the literature that connects firm characteristics to expected returns and the literature that studies macroeconomic aggregates as predictors of asset returns.

JEL classification: E12, E31, E44, E52, G12, G14.

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1 Introduction

A large strand of macroeconomics literature studies monetary policy and a substantial subset of it examines the effects of monetary policy on asset prices. Since the most immediate effects of policy actions are felt in financial markets, understanding the differential responses in the cross section of equities is crucial for our understanding of the broader impact of monetary policy. However, while there seems to be a consensus on the fact that monetary policy affects aggregate risk premia, its effects on the cross-section of risk premia are not as well-understood.[1]

The main challenge in studying the impact of monetary policy on the cross-section of equity risk premia arises from the difficulty in measuring firms' exposure to monetary policy. A direct approach, where one regresses individual stock returns on monetary policy surprises, is not fruitful because the majority of stocks have a very high volatility and lack a long enough history, leading to imprecise coefficient estimates.[2] As an alternative approach, the literature has identified various firm characteristics that affect stocks' reaction to monetary policy shocks.[3] In this study, we combine the two approaches to build a parsimonious "monetary policy exposure" (MPE) index at the individual firm level. To this end, we regress individual stock returns around scheduled FOMC meetings on a set of firm-level characteristics likely to capture the exposure, interacted with the monetary policy surprises on those meetings. We then create our MPE index by adding up the product of the estimated coefficients on the interactions with the corresponding characteristics.

[1] Bernanke and Kuttner (2005) show that a surprise 25-basis-point cut in the Fed fund rate target is associated with a 1% increase in broad stock indexes, and they attribute this increase predominantly to changes in risk premia. In the cross-section, Thorbecke (1997) finds no differential impact of monetary policy changes on portfolios sorted by size. Maio (2014) uses the VAR framework of Bernanke and Kuttner (2005) to break down the impact of Fed fund rate changes on the returns of portfolios sorted on size and book-to-market, and argues that the cash-flow effects are stronger for the small cap and value stocks.

[2] Velikov (2015) documents that although stocks' historical covariances with monetary policy surprises predict returns under this direct approach, these covariances are transient in nature and do not capture future monetary policy exposure.

[3] For example, Bernanke and Kuttner (2005) find that stocks of firms in cyclical industries react more to monetary policy shocks. Ehrmann and Fratzcher (2004) document that S&P 500 stocks with small size, low cash-flow, poor ratings, low leverage, and valuation are more sensitive to monetary policy. Ozdagli (2015) shows that firms with higher financial frictions react less to policy shocks and Weber (2015) documents that sticky price firms are more exposed to monetary policy.

Equipped with the MPE index, we study the asset pricing implications of monetary policy. Although several classes of theoretical models imply that monetary policy is an important source of risk in the stock market, their predictions regarding the sign of the risk premia differ widely. For example, in the cash-in-advance models (e.g. Bansal and Coleman, 1996 and Balduzzi, 2007) the sign of risk premium depends on the elasticity of substitution between cash and credit goods. New Keynesian models (e.g. Li and Palomino, 2014 and Weber, 2015) and intermediary asset pricing models (e.g. Adrian et al., 2014, Drechsler et al., 2015), where monetary policy is the "driver" of business cycles, suggest that monetary policy affects the pricing kernel through its impact on real variables and can lead to a positive risk premium. Alternatively, we provide a simple model in which monetary policy serves as a "mitigator" of business cycles, consistent with the central bank's objective of economic stability. In particular, monetary policy is likely to be expansionary during bad times when the marginal value of consumption and wealth is high. Thus, assets that pay off after an expansionary monetary policy serve as a hedge against bad times and require a lower expected return.

Consistent with the theories that guide our empirical approach, we find that that our MPE index is a strong predictor of returns in the cross-section of equities. Specifically, stocks with relatively low MPE index (i.e. those that tend to perform poorly when there is an expansionary monetary policy shock) have significantly higher average returns than those firms with high MPE index (i.e. those that tend to perform well when there is an expansionary monetary policy shock). A long-short trading strategy designed to exploit this effect achieves an annualized return of 10.56% (t-statistic of 6.19) and an annualized Sharpe Ratio of 0.98 between 1975 and 2014. The Fama-French-Carhart alpha of 7.56% suggests that this pattern does not seem to be driven by known predictors of returns.

This predictability is a robust feature of the data because it also holds when we recalculate the MPE index using only historical data available to investors at the time of a given FOMC announcement. This strategy generates an annualized return of 10.20% with a a Fama-French-Carhart alpha of 7.08%. Spanning tests also confirm that the strategy performance continues to hold even after controlling for the underlying characteristics used to construct the mone-

tary policy exposure index. Finally, our results do not seem to be driven by the pre-FOMC announcement drift documented by Lucca and Moench (2015) either. To show that, we reestimate the value-weighted portfolio returns based on MPE index after excluding days -1 and 0 around all scheduled FOMC meetings, and show that the predictability still holds.

Our study contributes to the literature in several ways. Inspired by the extensive corporate finance literature that generates indices of distress (e.g. Altman, 1968; Ohlson, 1980; Dichev, 1998; Campbell et al., 2008) or financial constraints (e.g. Kaplan and Zingales, 1997; Whited and Wu, 2006), this is the first paper to build an index of monetary policy exposure using firm characteristics. This method has an important advantage. While monetary policy surprises can be reliably estimated only after 1994, utilizing the characteristics and estimated coefficients allows us to extend the sample back to early 1970's.[4] The monetary policy exposure index we build is well-grounded in theory, easy to construct, and a robust predictor of returns in the cross-section.

More importantly, we build a bridge between the literature that connects firm characteristics to their expected returns, as in Fama and French (1993), and the literature that studies macroeconomic aggregates as predictors of asset returns, as in Lettau and Ludvigson (2001). Even though there has been an extensive theoretical literature that studies the impact of monetary policy on risk premia, the empirical evidence has been scant. Building the index of monetary policy exposure allows us to test the predictions of several different classes of theoretical models and show that monetary policy exposure is a strong predictor of returns in the cross-section of equities. Moreover, while the recent literature has become skeptical of predictability in the cross-section of stock returns, our monetary policy exposure index is theoretically motivated and our results are robust to the stricter thresholds for statistical significance suggested by Harvey, Liu, and Zhu (2016) and Novy-Marx (2016).

[4]Before 1994, the timing of rate changes was ambiguous, because the FOMC's policy did not involve announcing target rate changes at pre-scheduled dates.

2 Theoretical Considerations

Standard asset pricing theory tells us that risk-premium of monetary policy is the product of price of monetary policy risk and exposure to monetary policy. In this section, we will summarize the implications of different theories for these two components and show that monetary policy risk premium can be positive or negative depending on the underlying assumptions. Therefore, we conclude that determination of this risk premium is an empirical question which we focus on in the following sections.

2.1 The Price of Monetary Policy Risk

Bansal and Coleman (1996) and Chan, Perosi, and Lang (1996) are two early examples that analyze the effect of monetary policy on equity premia. Both of these models feature cash-in-advance constraints from the seminar paper of Lucas and Stokey (1987). The information in these papers is later crystallized in Balduzzi (2007), which shows that the price of monetary policy risk depends on elasticity of substitution between cash and credit goods. To see this, suppose the preferences are given by

$$E_t\left(\sum_{s=0}^{\infty}\beta^s U\left(c_{1t+s}, c_{2t+s}\right)\right)$$

where c_1 is the cash good and c_2 is the credit good and total output is $y = c_1 + c_2$. Under the common assumption that cash-in-advance constraint is binding, the consumption of cash good is equal to real money balances, m. In this framework, the stochastic discount factor, Λ_{t+1}, is determined by the marginal utility of credit good, U_{c_2}, that is

$$\Lambda_{t+1} = \frac{\beta U_{c_2}\left(c_{1t+1}, c_{2t+1}\right)}{U_{c_2}\left(c_{1t}, c_{2t}\right)} = \frac{\beta U_{c_2}\left(m_{t+1}, y - m_{t+1}\right)}{U_{c_2}\left(m_t, y - m_t\right)}.$$

For a given level of output, y, an expansionary policy (reduction in interest rates or increase in money balances) makes the consumers substitute away from credit good and the net effect on stochastic discount factor will be driven by $U_{c_2c_1} - U_{c_2c_2}$. When $U_{c_2c_1} - U_{c_2c_2} < 0$ the

price of expansionary monetary policy shock is negative and when $U_{c_2c_1} - U_{c_2c_2} > 0$ the price of monetary policy risk is positive. Therefore, in cash-in-advance models, monetary policy commands a risk premium even in the absence of any real effects on output. Moreover, this risk premium can be positive or negative depending on the substitutability between cash and credit goods, regardless of the effect of monetary policy on output, dy/dm_{t+1}.

More recently, the New Keynesian monetary models that incorporate nominal and real rigidities have received increasing attention in the literature studying the risk of equities, e.g. Li and Palomino (2014), and Weber (2015).[5] In these two models, consumption and leisure are additively separable and hence an expansionary policy shock increases consumption and decreases the marginal utility thereof, leading to a positive risk price for monetary policy. Therefore, stocks that react more positively to an expansionary policy shock should command a higher risk premium. If, on the other hand, consumption and leisure are substitutes an increase in leisure following a contractionary shock can reduce the marginal utility of consumption, leading to a negative risk price of an monetary policy shocks.

Since the recent financial crisis, a growing number of papers incorporate macroeconomic models of financial frictions, such as those in Bernanke and Blinder (1988), to study asset pricing. The central theme of this literature is that the marginal investor is likely a financial intermediary, so the stochastic discount factor should depend on the health of the financial sector, or funding liquidity, which determines the marginal value of wealth. Thus, assets that pay off in times of high marginal value of wealth, i.e. times with low funding liquidity, should be less risky.[6] Empirically, Adrian, Etula, and Muir (2014) show that a single-factor model consisting of security broker-dealers' leverage explains a large fraction of returns of size, value, momentum, and treasury portfolios. Drechsler, Savov, and Schnabl (2015) provide a dynamic asset pricing model linking equity risk premia to monetary policy through this funding liquidity channel. In their model, lowering the nominal interest rate reduces the cost of leverage and

[5] See also Uhlig (2007) and Campbell, Pflueger, and Viceira (2015).

[6] These papers differ in the proxy they use for times of high marginal value of wealth. For example, He and Krishnamurthy (2013) and Brunnermeier and Sannikov (2014) argue that it is the equity of financial intermediaries while Brunnermeier and Pedersen (2009), Geanakopoulos (2009), and Adrian and Boyarchenko (2012) feature intermediary leverage as the relevant state variable.

effectively reduces the external finance premium, increasing risk taking and, in turn, decreasing risk premia.

Both the New Keynesian models and the funding liquidity model operate under the view that monetary policy is a "driver" of business cycles, that is, monetary policy affects the marginal value of wealth by affecting real variables which in turn transforms it into a source of priced risk. An alternative view arises from viewing the monetary policy as a "mitigator" of business cycles. In particular, monetary policy is more likely to be expansionary during bad times when the marginal value of wealth is high. The assets that are more likely to pay off after an expansionary monetary policy are precisely those that provide investors with additional funds in times of need and therefore have a lower risk premium. This mitigator view seems to be more in line with the role of Federal Reserve in stabilizing the economy.[7] The appendix B provides a simple mathematical model to clarify the "driver" and "mitigator/stabilizer" channels in more detail.

[7] Note that this is different from the "signalling channel" of monetary policy where an increase in interest rate reveals good news about the future of the economy, as in Romer and Romer (2000), Campbell, Evans, Fisher, and Justiniano (2012), and Nakamura and Steinsson (2015). While the signaling channel would also imply a negative risk price of expansionary shock, the intuition is actually closer to long-run risk models where positive interest rate shocks are generally good news, which makes long-duration assets, which are more interest sensitive, valuable hedges, reducing their risk premia.

2.2 Theories of Exposure to Monetary Policy

Monetary policy has a large effect on stock prices and this effect varies significantly across firms, consistent with the fact that there are many channels that work heterogeneously for firms with different characteristics. We give the definitions of these characteristics in appendix A and briefly discuss below the theory underlying the choice of these particular characteristics.

Financial Constraints (Credit Channel): The effect of firms' financial constraints on the monetary policy transmission has been at the heart of the policy and academic discussions.[8] While there has been an expansive literature focusing on the implications of this credit channel on real variables, the evidence on stock prices has been relatively scarce. Perez-Quiros and Timmermann (2000) find that stock prices of smaller (and therefore financially constrained) firms are more responsive to monetary policy, measured by money supply. Lamont, Polk, and Saa-Requejo (2001) recognize that modern monetary policy is characterized by the choice of policy rates but do not find any significant relationship when they use the change in policy rate. Ozdagli (2015) uses unexpected component of the policy rate change as in Bernanke and Kuttner (2005), because stocks should not react to expected changes in monetary policy, and finds that more constrained firms are less responsive to monetary policy if these firms rely less on external finance and hence less affected by the cost of external finance.[9] We use a firm's percentile rank according to the financial constraint index created by Whited and Wu (2006). Our choice of rank stems from the common practice in the literature that discretely separates the firms into financially constrained and unconstrained groups using financial constrained proxies, a practice which we take to the limit with percentile rank.

Cash and Short-Term Investments (Liquidity Effect): These are the most liquid assets of the firm and directly related to the monetary base, broadly defined. On the one hand, if firms keep their cash in a non-interest-bearing account, stocks of firms with higher amount of cash

[8] See Gertler and Gilchrist (1994) for the seminal empirical paper in this literature although the theory of this credit channel goes back earlier.

[9] Ozdagli (2015) also shows that the relationship between stock prices and monetary policy does not need to mirror the relationship between investment and monetary policy because market value of equity is the value function of a firm's optimization problem where investment is the choice variable and there is no theoretical basis that a value function and a choice variable should move in the same direction.

can react more negatively to an interest rate increase because the interest rate is the opportunity cost of holding cash. On the other hand, if they deposit their cash in a short-term savings or another interest bearing account, an increase in interest rate can actually help them obtain additional liquid funds.

Cash flow duration (Discount Rate Effect): While duration has been a construct widely employed by fixed-income analysts due to its clear relationship to the interest rate sensitivity of bond prices, the application to equity markets has only been recently studied by Dechow, Sloan, and Soliman (2004). Firms that expect to have cash flows farther in the future, and therefore have greater equity duration, would be more affected by an increase in interest rates since later cash flows are discounted at a higher rate.

Cash flow volatility: Cash flow volatility may capture the monetary policy sensitivity of a firm's stock price in multiple ways. For example, volatility can be related with cash flow duration and can capture aspects thereof not perfectly captured by standard cash flow duration measures. Mechanically, firms with lower volatility may have lower default likelihood and therefore longer lives and high duration of cash flows. A lower volatility may also imply a lower value of option to delay investment and increase cash flow duration by increasing investment today in exchange of cash flows in the future. Alternatively, higher cash flow volatility may imply that the firm needs to rely on external financing more often and therefore increases the importance of cost of external financing which is directly affected by monetary policy.

Operating Profitability (Sticky Prices): Nominal frictions in form of sticky prices and wages are an important ingredient in the New-Keynesian macroeconomic models. While the data on price stickiness at the firm level is difficult to find for majority of firms, operating profitability can still provide a window to the effects of sticky prices.[10] In particular, if the input prices, e.g. wages, are sticky, an expansionary monetary policy will have a large effect on the revenues of the firm without changing total cost of inputs as much, driving stock prices up. The resulting percentage increase in stock price will be stronger for firms of which revenues

[10] An important exception for firm-level price stickiness data is Weber (2015) who uses hand-collected data for S&P 500 firms and finds that firms with high price stickiness have higher returns but that these return differentials can be explained by CAPM β.

are closer to their input costs, i.e. those that have low profitability, because of the operating leverage effect created by relatively fixed input costs. On the other hand, if a firm's output prices are more sticky than input prices, an expansionary monetary policy will lead to a greater increase input costs than the firm's revenues, eating away the firm's profits, but also to smaller degree for firms with greater profitability.

Other: We also control for industry effects but do not include this in our MPE index to keep the index parsimonious. We also include firm ratings in our estimation sample, the 1994–2008 period for which we have the Bernanke-Kuttner surprises, but do not use them in our index because the ratings information does not go back to 1975, when our full sample analysis starts.

3 Monetary Policy Exposure Estimation

In this section, we discuss the estimation of monetary policy exposure. Following Kuttner (2001) and Bernanke and Kuttner (2005), we measure policy shocks as the unexpected component of the Federal funds target rate change due to target rate announcements on the day of FOMC meetings. We focus on scheduled FOMC metings between February, 1994, and June, 2008. Starting in February, 1994, the FOMC's policy of announcing target rate changes at prescheduled dates virtually eliminated the timing ambiguity associated with rate changes prior to this date. We end mid-2008, because this is when Ken Kuttner's data stops as the fed funds target rate reaches the zero lower bound.[11]

For an FOMC meeting happening on day d of month m, the surprise change in the Federal funds rate is given by:

$$\Delta i^u = \frac{D}{D-d}(f^0_{m,d} - f^0_{m,d-1}) \tag{1}$$

where D is the number of days in the month and $f^0_{m,d}$ is the federal funds rate implied by the federal funds future expiring in the current month. Following Bernanke and Kuttner (2005),

[11] See http://econ.williams.edu/profile/knk1/ for details. Given the Fed's recent move toward policy normalization, this will be the more relevant policy measure for the future as well.

our identification of monetary policy exposure comes from regressing the returns around the scheduled FOMC announcement on the surprise component of policy rate change on that announcement date.[12]

We prefer this approach from Bernanke and Kuttner (2005) because previous literature finds that federal funds futures outperform target-rate forecasts based on other financial market instruments or based on alternative methods, such as sophisticated time series specifications and monetary policy rules.[13] Another advantage of looking at one-day changes in near-dated federal funds futures is that federal funds futures do not exhibit predictable time-varying risk premia and forecast errors over daily frequencies.[14]

An important challenge to calculating the monetary policy exposure of individual stocks comes from the fact that many stocks have very high volatility and lack a long enough history, leading to imprecise coefficient estimates when one simply runs a regression of individual stock returns on policy surprises. Instead, we note that the literature has identified various firm characteristics that influence stocks' reaction to monetary policy shocks and use the interaction of these characteristics with policy surprises as explanatory variables. Our main specification is

$$r_{it} = \alpha + \Sigma_{k=1}^{n} \beta_k x_{it} + \Sigma_{k=1}^{n} \gamma_k MPS_t \times x_{it} + \text{Controls}_{it} + \varepsilon_{it}, \tag{2}$$

where t is the date of the scheduled FOMC meetings, r_{it} is the stock return surrounding these FOMC meetings, and x_{it} are the firm characteristics that capture the exposure of a firm to monetary policy, as discussed in section 2.2. MPS_t is the monetary policy surprise due to FOMC announcements on these meeting dates; $MPS_t = -\Delta i^u$ so that a positive surprise is expansionary. Controls_{it} includes meeting, industry, and rating fixed effects, as well as interactions of the industry and rating fixed effects with the monetary policy surprises. The

[12] We drop the FOMC announcement on March 18, 2008, as an outlier because on that day, the S&P 500 index increased by about 4 percent despite a 17 basis points contractionary policy surprise, reflecting the positive news about JP Morgan's purchase of Bear Stearns. Including this date changes the sensitivity of the CRSP value-weighted index to a one percentage point policy surprise from about 3.5 percent, in line with earlier estimates of Bernanke and Kuttner (2005), to 1.5 percent.

[13] Evans (1998) and Gürkaynak, Sack, and Swanson (2007).

[14] See, for example, Piazzesi and Swanson (2008).

regressions are estimated at the meeting-firm level. Table 1 shows the descriptive statistics for various characteristics, x_{it}, determine the monetary policy exposure.

Table 2 presents the results regarding how different characteristics capture the exposure of stock prices to monetary policy. Following Ippolito, Ozdagli, and Perez (2015), we focus on two-day returns to allow sufficient time for the stocks to react to monetary policy while still staying within the two-day blackout period that follows an FOMC meeting in order to avoid the contamination by other FOMC-related news. The first column gives the unconditional relationship and shows that, consistent with earlier results in Bernanke and Kuttner, a one percentage point expansionary surprise (reduction in fed funds target rate) leads, on average, to a 4 percent increase in stock prices. Column 2 shows that higher cash holdings lead to lower stock price sensitivity to monetary policy, consistent with the notion that short-term deposits in interest-bearing accounts provide a hedge against interest rate changes. Column 3 shows that, consistent with our intuition, firms with higher cash flow duration are more responsive to surprise changes in federal funds target rate. Column 4 presents the effect of financial constraints – consistent with Ozdagli (2015), stock prices of more constrained firms, i.e., those with higher Whited-Wu measure, less responsive to monetary policy surprises. Column 5 show that firms with greater cash flow volatility are more responsive to monetary policy surprises. Moreover, column 6 shows that firms with higher operating profitability are less responsive to monetary policy, consistent with the intuition that sticky input prices, such as wages, may lead to lower sensitivity of firm's profits to changes in input prices.

Finally, the second to last column of Table 2 puts all these variables together and show that the coefficients preserve most of their size and significance. Using the estimates in column 7, our monetary policy exposure (MPE) index becomes:

$$\begin{aligned} \text{MPE} = -3.14 \times \text{Cash} + 1.44 \times \text{CF Duration Rank} - 2.12 \times \text{Whited-Wu Rank} \\ 8.69 \times \text{CF Volatility} - 9.03 \times \text{Operating Profitability} \end{aligned} \tag{3}$$

The last column uses this index directly as a regressor interacted with monetary policy surprise we see that it is highly significant, with a t-statistic of 7.70. Note that while we have used Ratings and Industry Fixed Effects interacted with monetary policy surprise (MPS) as additional controls we do not include them into our MPE index. Ratings were seldom available for firms towards the beginning of our sample period, which starts in 1975; omitting industry effects allows us to focus on cross-sectional differences in stock returns that goes beyond differences in industry returns and keep our index parsimonious.

3.1 External Validity: Forward Guidance

In the previous section, we have followed the footsteps of Bernanke and Kuttner (2005) and used the monetary policy surprises implied by the fed funds futures to estimate our monetary policy exposure index. However, the nature of monetary policy has changed since 2008, once the short-term rates have hit the zero lower bound (ZLB). During this period, the Bernanke-Kuttner measure of monetary policy is uninformative because monetary policy actions did not have any significant effect on the short-term rates but instead influenced the expectations of longer-term rates through large scale asset purchases and forward guidance. Therefore, we are interested in how well our monetary policy exposure index does in the recent ZLB period, which would provide evidence of external validity of our measure.

We follow Rigobon and Sack (2004) and Gürkaynak, Sack, and Swanson (2005) and use the Eurodollar (ED) futures with maturities ranging from 1 to 8 quarters, in order to a measure the change in expectations about future interest rates around FOMC announcements.[15] If our MPE index has any merit, we should find that firms with high MPE respond more strongly to these policy-induced shocks to rate expectations (ZLB shocks). Moreover, the effect of these ZLB shocks should be stronger for longer-term rates that the monetary policy has focused on. Table 3 provides regressions similar to the one in the last column of Table 2 using these ZLB shocks. Consistent with our intuition, the size and significance of the effect of ZLB shocks become greater as we go from shocks to short-term rate expectations, captures by ED1, to

[15] ED futures capture the expectations of 3-month LIBOR at the maturity date of these futures.

longer-term expectations, captured by ED8. Moreover, the coefficients for ED5 to ED8 are very close to one as in the last column of Table 2. Overall, these results confirm the validity of our MPE index outside of the period used for its estimation, thereby providing support for its external validity.

4 Asset Pricing Implications

Equipped with our monetary policy exposure (MPE) index, in this section we explore the asset pricing implications of monetary policy. Table 4 reports average value-weighted excess returns, alphas, and loadings on the four Fama-French-Carhart factors for five portfolios constructed by sorting stocks based on their monetary policy exposure, as well as a portfolio that is long stocks in the lowest quintile of monetary policy exposure and short stocks in the highest quintile of monetary policy exposure. Monetary policy exposure is estimated using equation 3 from Section 3. In each month, firms are grouped into quintiles based on their monetary policy exposure using NYSE breakpoints. While data on the monetary policy surprises are only available for the shorter period used for the estimation of the monetary policy index in Table 2, data on the the characteristics used to construct the monetary policy exposure index are available going back to January 1975 and up to December 2014, so we use this longer time period for the asset pricing tests.

We can observe a monotonically decreasing pattern in the average returns to the five portfolios. The long/short portfolio achieves an annualized average return of over 10%, with an associated t-statistic of 6.19. The strategy generates a significant alpha of 63 basis points per month (t-statistic of 6.51) with respect to the four- factor model. The decrease in performance can be attributed to the significant loadings on the SMB and HML factors of 0.69 (t-statistic of 21.36) and 0.51 (t-statistic of 14.83), respectively. These loadings are not surprising, given that the monetary policy index employs the Whited-Wu index and CF Duration. Stocks with low MPE tend to be more financially constrained and to have lower cash flow duration. Dechow, Sloan, and Soliman (2004) and Weber (2016) show that cash flow duration is negatively

related to book-to-market, which explains the value tilt of the strategy, while log of total assets negatively enters the Whited-Wu index, which explains the size tilt of the strategy.

4.1 Out-of-sample Portfolio Performance

Despite the strong performance, the strategy from Table 4 is not implementable, because it uses information that would not have been available to investors at the beginning of the sample. To address this concern, in this section we perform an out-of-sample test using only historically available information. That is, monetary policy exposure is estimated using expanding windows, including only FOMC meetings prior to the portfolio formation month. The sample begins in 08/1996, using 20 FOMC meetings between 02/1994 and 07/1996. Following 06/2008, all 115 meetings between 02/1994 and 06/2008 are used for the estimation of the MPE.

We can observe that the average returns and alphas generated by this implementable strategy are very close to the full-sample ones. The expanding windows strategy generates 0.85% per month compared to 0.88% per month for the full sample ones. The four-factor alpha for the expanding windows strategy is 0.59% per month compared to 0.63% per month for the full sample one. The slightly lower statistical significance for the out-of-sample test can be attributed to the lower length of the sample period used.

Another interesting observation from Table 5 and, to a lesser extent from Table 4, is that the loadings on the market factor are all smaller than one, which indicates that the sample is tilted towards smaller beta stocks. This sample selection is caused by the construction of the Cash Flow Volatility variable used in the estimation of the monetary policy exposure. It requires the existence of eight consecutive quarterly observations, which cuts our sample by about a quarter.[16] This issue may understate our results because the missing observations are likely to have higher true cash flow volatility. As evidenced by Table 2, Cash Flow Volatility is positively associated with monetary policy exposure. Thus, the firms with missing Cash Flow Volatility should have lower average returns. Consistent with this hypothesis, the portfolio of firms with missing Cash Flow Volatility observations, but with available return and market

[16] See Appendix A for more details on the construction of the variables.

capitalization data, achieves an average monthly return of 20 basis points (t-statistic of 1.04), which is lower than the lowest monetary policy exposure portfolio in Table 5. Finally, the Fama-French- Carhart four factor loadings of the missing Cash Flow Volatility portfolio are given by: $R_{\text{Missing CF Vol Ptf}} = \underset{[-2.49]}{-0.329} + \underset{[38.22]}{1.181} \times R_{\text{MKT}} + \underset{[1.29}{0.051} \times R_{\text{SMB}} + \underset{[6.94]}{0.292} \times R_{\text{HML}} - \underset{[-0.97]}{0.024} \times R_{\text{UMD}}$. The loading on the market factor of 1.181 shows that it is indeed the construction of the Cash Flow Volatility variable that is causing the sample selection.

4.2 Robustness to FOMC meetings

Another potential concern with looking at the stock market reactions to scheduled FOMC meetings is the extent to which results are driven by the pre- FOMC drift documented by Lucca and Moench (2015). Thus, to alleviate this concern, Table 6 replicates the analysis in Table 4 using the same MPE index, but excludes the returns for all stocks on the day prior to and on the date of a scheduled FOMC meeting. That is, for months with FOMC meetings, all stocks are assumed to have zero returns on the day before the FOMC meeting and on the day of the FOMC meeting. We can observe that using the adjusted returns does not weaken the predictability of monetary policy exposure. If anything, the average returns and alphas to the long/short portfolio are somewhat higher. Excluding days -1 and 0 around FOMC days results in an average monthly return of 0.89% ($t = 6.50$) compared to 0.88% ($t = 6.19$), and a Fama-French-Carhart alpha of 0.67% ($t = 6.68$) compared to 0.63% ($t = 6.51$). Thus, the evidence indicates that the monetary policy risk premium we document is not accrued only around FOMC meetings, which is consistent with the notion that investors require a premium to hold stocks that perform poorly in times of expansionary monetary policy shocks.

4.3 Robustness to controlling for underlying characteristics

In this section we examine the predictability of monetary policy exposure controlling for the characteristics used to construct it.

Table 7, Panel A documents the average value-weighted returns in excess of the risk-free

rate to long/short portfolio strategies constructed by quintile sorts on the underlying characteristics using NYSE breakpoints. We can observe that, in the sample period examined, only Cash, Operating Profitability and to an extent CF Duration seem to be significant predictors of returns in the cross-section. Panel B documents results from spanning tests, in which the returns to the strategies from Panel A are regressed on a constant and the returns to the long/short MPE portfolio from Table 4. We can observe that the returns to the Size, CF Duration, and Operating Profitability strategies can largely be explained by the monetary policy exposure strategy. After controlling for monetary policy exposure, none of the underlying characteristics seems to be a significant predictor of returns in the cross section.

Similarly, Table 8 reports results from spanning tests, in which the returns to the long/short portfolio from Table 4 are regressed on a constant term and the returns to the underlying characteristics used to construct the monetary policy exposure measure. The results imply that the strategy based on monetary policy exposure is outside of the span of the strategies based on the underlying characteristics, even when all of them are considered simultaneously. Specification (6) shows that the monetary policy exposure strategy has a significant alpha of 28 basis points per month ($t = 2.96$) with respect to the underlying strategies.

5 Conclusion

In this paper, we find that stocks that react more positively to expansionary monetary policy shocks earn lower average returns. A long/short trading strategy designed to exploit this effect achieves an annualized return of 10.56% (t-statistic of 6.19), an annualized Sharpe Ratio of 0.98, and a Fama-French-Carhart alpha of 7.56% (t-statistic 6.51) between 1975 and 2014. Measuring the monetary policy exposure of an individual stock is a difficult identification problem because idiosyncratic volatility of individual stocks lead to large standard errors, especially when a stock does not have a long enough history. Instead, in the spirit of the literature on financial constraints, we generate a parsimonious monetary policy exposure index based on observable firm characteristics that are likely to drive the exposure of firms to monetary policy.

Therefore, our approach builds a bridge between the literature that connects firm characteristics to expected returns and the literature that study macroeconomic aggregates as a predictor of asset returns. While we derive our monetary policy exposure index to study its implications for asset pricing, our index can also be useful in future research about the relationship between firms' financing and investment decisions and monetary policy.

Table 1: Descriptive Statistics

This table reports descriptive statistics for variables used in estimating monetary policy exposure. The variable definitions are in Appendix A. Column one reports the number of firm-meetings. Column two reports time-series averages of cross-sectional averages for the seven variables. The rest of columns report time-series averages of cross-sectional percentiles. All statistics are multiplied by 100 for exposition.The sample excludes financial firms, firms with non-traded CRSP closing prices on the day before the meeting, and firms with CRSP closing prices lower than or equal to $5 on the meeting days. The sample covers 115 meetings between 01/1994 and 06/2008.

Variable			Percentiles				
	n	Mean	1^{st}	5^{th}	50^{th}	95^{th}	99^{th}
Cash	275,574	0.09	0.00	0.00	0.04	0.28	0.49
CF Duration Rank	275,574	0.54	0.02	0.07	0.52	0.89	0.97
Whited-Wu Rank	275,574	0.52	0.01	0.04	0.40	0.84	0.93
CF Volatility	275,574	0.11	0.01	0.02	0.05	0.29	0.65
Operating Profitability	275,574	0.06	-0.03	0.00	0.05	0.14	0.22

Table 2: Monetary Policy Exposure Estimation

This table reports regression estimates from panel ordinary least square regression estimations from the following specification: $r_{it}^{0,1} = \alpha + \Sigma_{k=1}^{n}\beta_k x_{it} + \Sigma_{k=1}^{n}\gamma_k MPS_t \times x_{it} + \text{Controls}_{it} + \varepsilon_{it}$, where $r_{it}^{0,1}$ is the 2-day cumulative return on stocks surrounding scheduled FOMC meetings (days 0 and 1), x_{it} are firm characteristics described in Appendix A, MPS_t is a monetary policy surprise measure, developed by Kuttner (2001), multiplied by -1 so that $MPS > 0$ is expansionary, and Controls_{it} includes meeting, industry, and rating fixed effects, the interactions of the industry and rating fixed effects with the monetary policy surprises. The regressions are estimated at the meeting-firm level. Independent variables are winsorized at the 1% level within each cross-section.The sample excludes financial firms, firms with non-traded CRSP closing prices on the day before the meeting, and firms with CRSP closing prices lower than or equal to $5 on the meeting days. The sample covers 115 meetings between 01/1994 and 06/2008.

Var	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Cash		0.22					0.26	
		[2.67]					[2.75]	
CF Duration Rank			0.18				0.20	
			[4.82]				[4.88]	
Whited-Wu Rank				-0.04			-0.10	
				[-0.80]			[-2.05]	
CF Volatility					0.04		0.02	
					[0.45]		[0.21]	
Operating Profitability						0.06	0.30	
						[0.30]	[1.24]	
MPE								0.05
								[6.24]
MPS	4.20							
	[25.21]							
MPS x Cash		-4.74					-3.14	
		[-3.19]					[-1.84]	
MPS x CF Duration Rank			2.63				1.44	
			[3.93]				[1.87]	
MPS x Whited-Wu Rank				-2.65			-2.12	
				[-3.10]			[-2.28]	
MPS x CF Volatility					7.92		8.69	
					[5.06]		[5.01]	
MPS x Op. Profitability						-15.18	-9.03	
						[-4.17]	[-2.17]	
MPS x MPE								1.08
								[7.70]
Meeting FE	No	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Industry FE	No	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Industry FE $\times$ MPS	No	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Rating FE	No	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Rating FE $\times$ MPS	No	Yes	Yes	Yes	Yes	Yes	Yes	Yes
$\bar{R}^2(\%)$	0.23	7.51	8.08	7.58	7.88	7.57	8.27	8.29

Table 3: Forward Guidance Table

This table reports regression estimates from panel ordinary least square regression estimations from the following specification: $r_{it}^{0,1} = \alpha + \beta x_{it} + \gamma \text{EDX}_t \times \text{MPE}_{it} + \text{Controls}_{it} + \varepsilon_{it}$, where $r_{it}^{0,1}$ is the 2-day cumulative return on stocks surrounding scheduled FOMC meetings (days 0 and 1), MPE is monetary policy exposure, estimated using equation (3) from the text, EDX_t is the change in the price of the X-quarter ahead eurodollar futures contract, following Rigobon and Sack (2004) and Gürkaynak, Sack, and Swanson (2005), multiplied by -1 so that $EDX > 0$ is expansionary, and Controls_{it} includes meeting, industry, and rating fixed effects, and the interactions of the industry and rating fixed effects with the monetary policy surprises. The regressions are estimated at the meeting-firm level. Independent variables are winsorized at the 1% level within each cross-section.The sample excludes financial firms, firms with non-traded CRSP closing prices on the day before the meeting, and firms with CRSP closing prices lower than or equal to $5 on the meeting days. The sample covers 48 meetings between 01/2009 and 12/2014.

Var	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
MPE	-0.01	-0.02	-0.02	-0.02	-0.02	-0.02	-0.02	-0.02
	[-0.94]	[-1.19]	[-1.34]	[-1.37]	[-1.49]	[-1.53]	[-1.40]	[-1.37]
ED1 x MPE	0.60							
	[0.83]							
ED2 x MPE		0.63						
		[1.34]						
ED3 x MPE			0.70					
			[1.95]					
ED4 x MPE				0.78				
				[2.65]				
ED5 x MPE					0.99			
					[3.91]			
ED6 x MPE						1.11		
						[4.90]		
ED7 x MPE							1.19	
							[5.91]	
ED8 x MPE								1.17
								[6.46]
Meeting FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Industry FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Industry FE $\times$ EDX	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Rating FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Rating FE $\times$ EDX	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
n	88,597	88,597	88,597	88,597	88,597	88,597	88,597	88,597
$\bar{R}^2(\%)$	26.52	26.54	26.50	26.50	26.48	26.46	26.45	26.45

Table 4: Full-sample Portfolio Performance

This table reports average excess returns, alphas, and loadings on the four Fama-French-Carhart factors. In each month, firms are sorted by their monetary policy exposure into quintiles based on NYSE breakpoints. Monetary policy exposure is estimated using equation (3) from the text. For each portfolio, average value-weighted returns in excess of the risk-free rate, alphas and loadings to the four factors of the Fama-French-Carhart model are reported. T-stats are in brackets. Sample period is 01/1975 to 12/2014.

Var	r^e	α_{ff4}	β_{mkt}	β_{smb}	β_{hml}	β_{umd}
(L)	1.32	0.57	0.99	0.61	0.25	-0.15
	[5.39]	[7.04]	[52.41]	[22.30]	[8.73]	[-8.40]
(2)	0.95	0.23	1.02	0.27	0.26	-0.09
	[4.16]	[2.95]	[56.56]	[10.22]	[9.36]	[-5.42]
(3)	0.83	0.18	0.93	-0.02	0.24	-0.02
	[4.21]	[2.39]	[53.57]	[-0.77]	[9.10]	[-0.98]
(4)	0.62	0.05	0.94	-0.12	0.01	0.02
	[3.10]	[0.67]	[57.34]	[-4.90]	[0.57]	[1.50]
(H)	0.44	-0.06	0.97	-0.09	-0.26	0.00
	[2.06]	[-1.18]	[83.49]	[-5.17]	[-14.61]	[0.23]
(L-H)	0.88	0.63	0.02	0.69	0.51	-0.16
	[6.19]	[6.51]	[1.00]	[21.36]	[14.83]	[-7.16]

Table 5: Out-of-sample Portfolio Performance
This table reports average excess returns, alphas, and loadings on the four Fama-French-Carhart factors. In each month, firms are sorted by their monetary policy exposure into quintiles based on NYSE breakpoints. Monetary policy exposure is estimated using equation (3) from the text, using only historically available information for all meetings starting in 02/1994 and until 06/2008. The first estimation uses 20 FOMC meetings between 02/1994 and 07/1996. For each portfolio, average value-weighted returns in excess of the risk-free rate, alphas and loadings to the four factors of the Fama-French-Carhart model are reported. T-stats are in brackets. Sample period is 08/1996 to 12/2014.

Var	r^e	α_{ff4}	β_{mkt}	β_{smb}	β_{hml}	β_{umd}
(L)	1.36	0.59	0.89	0.67	0.50	-0.12
	[3.70]	[4.74]	[30.19]	[17.83]	[12.45]	[-4.87]
(2)	0.98	0.26	0.94	0.26	0.49	-0.06
	[2.96]	[2.17]	[33.33]	[7.15]	[12.76]	[-2.79]
(3)	1.01	0.36	0.90	0.10	0.34	-0.03
	[3.31]	[3.14]	[33.10]	[2.84]	[9.04]	[-1.35]
(4)	0.78	0.20	0.91	-0.05	0.10	0.03
	[2.68]	[2.02]	[39.67]	[-1.82]	[3.15]	[1.37]
(H)	0.51	0.01	0.96	-0.08	-0.30	0.00
	[1.57]	[0.13]	[59.31]	[-3.77]	[-13.62]	[0.20]
(L-H)	0.85	0.59	-0.08	0.75	0.80	-0.12
	[3.35]	[4.18]	[-2.38]	[17.82]	[17.89]	[-4.46]

Table 6: Excluding FOMC Meetings

This table reports average excess returns, alphas, and loadings on the four Fama-French-Carhart factors. In each month, firms are sorted by their monetary policy exposure into quintiles based on NYSE breakpoints. Monetary policy exposure is estimated using equation (3) from the textt. For each portfolio, average returns in excess of the risk-free rate, alphas and loadings to the four factors of the Fama-French-Carhart model are reported. The returns to the portfolios are measured by excluding the two-day (days -1 and 0) daily returns around 339 days with scheduled or unscheduled FOMC meetings between 01/1975 and 12/2014 from the monthly returns. T-stats are in brackets. Sample period is 01/1975 to 12/2014.

Var	r^e	α_{ff4}	β_{mkt}	β_{smb}	β_{hml}	β_{umd}
(L)	1.14	0.40	0.95	0.55	0.27	-0.12
	[4.78]	[4.11]	[41.47]	[16.68]	[7.69]	[-5.60]
(2)	0.76	0.06	0.97	0.24	0.28	-0.08
	[3.42]	[0.62]	[45.27]	[7.76]	[8.52]	[-3.74]
(3)	0.62	-0.02	0.89	-0.03	0.26	0.01
	[3.22]	[-0.22]	[43.94]	[-0.99]	[8.36]	[0.27]
(4)	0.41	-0.15	0.90	-0.12	0.05	0.04
	[2.10]	[-1.86]	[46.53]	[-4.34]	[1.54]	[1.89]
(H)	0.24	-0.26	0.93	-0.09	-0.20	0.02
	[1.16]	[-3.61]	[54.77]	[-3.61]	[-7.59]	[1.06]
(L-H)	0.89	0.67	0.01	0.64	0.47	-0.14
	[6.50]	[6.68]	[0.58]	[19.00]	[13.10]	[-6.26]

Table 7: Spanning Tests

The table documents results from time-series regressions of the returns to the strategy from table 4 on returns to strategies constructed from the variables used to derive the monetary policy exposure. All strategies are consructed on the same sample from a quintile sort using NYSE breakpoints. T-stats are in brackets. Sample period is 01/1975 to 12/2014.

Panel A: Average Returns					
	$y = R_{\text{Cash}}$	$y = R_{\text{CF Duration}}$	$y = R_{\text{Whited-Wu}}$	$y = R_{\text{CF Volatility}}$	$y = R_{\text{Op. Prof.}}$
r^e	0.53	-0.32	0.30	-0.13	0.62
	[3.01]	[-1.85]	[1.32]	[-0.78]	[4.37]
Panel B: Regressions of the form $R_{t,X} = \alpha + \beta_{\text{MPE}} R_{MPE,t} + \varepsilon_t$					
α	0.31	0.06	-0.25	-0.13	0.19
	[1.75]	[0.35]	[-1.13]	[-0.75]	[1.46]
β_{MPE}	24.29	-43.51	62.37	0.00	48.75
	[4.38]	[-8.26]	[9.16]	[0.00]	[12.34]

Table 8: Spanning Tests

The table documents results from time-series regressions of the returns to the strategy from table 4 on returns to strategies constructed from the variables used to derive the monetary policy exposure. All strategies are consructed on the same sample from a quintile sort using NYSE breakpoints. T-stats are in brackets. Sample period is 01/1975 to 12/2014.

	Regressions of the form $R_{t,MPE} = \alpha + \beta_X R_{X,t} + \varepsilon_t$					
Coefficient	(1)	(2)	(3)	(4)	(5)	(6)
Const	0.80	0.79	0.81	0.89	0.57	0.28
	[5.64]	[5.89]	[6.14]	[6.24]	[4.54]	[2.96]
β_{Cash}	16.05					20.32
	[4.38]					[6.31]
$\beta_{\text{CF Duration}}$		-28.99				-23.07
		[-8.26]				[-7.87]
$\beta_{\text{Whited-Wu}}$			24.14			33.49
			[9.16]			[12.58]
$\beta_{\text{CF Volatility}}$				0.00		-26.09
				[0.00]		[-6.29]
$\beta_{\text{Operating Profitability}}$					49.98	46.98
					[12.34]	[15.01]

Appendices

A Variable Definitions

- Cash - Cash and short-term investments (CHEQ) scaled by by the sum of total liabilities (LTQ) and market capitalization
- Cash Flow Duration Rank - Cash flow duration rank, estimated using the cash flow duration measure of Dechow, Sloan, and Soliman (2004):

$$\text{CF Duration}_t = \frac{\Sigma_{s=1}^{T} s \times \text{CF}_{t+s}/(1+r)^t}{P_t} + T + \frac{1+r}{r} \times \frac{P_t - \Sigma_{s=1}^{T} \text{CF}_{t+s}/(1+r)^s}{P_t}$$

Cash flows are measured assuming clean surplus accounting:

$$\begin{aligned} \text{CF}_t &= \text{E}_t - (\text{BV}_t - \text{BV}_{t-1}) \\ &= \text{BV}_{t-1} \times \left[\underbrace{\frac{\text{E}_t}{\text{BV}_{t-1}}}_{\text{ROE}} - \underbrace{\frac{(\text{BV}_t - \text{BV}_{t-1})}{\text{BV}_{t-1}}}_{\text{Growth in BE}} \right] \end{aligned}$$

and forecasted following Nissim and Penman (2001). ROE follows a first-order autoregressive process with an autocorrelation coefficient equal to the long-run average rate of mean reversion in ROE and a long-run mean equal to the cost of equity. Nissim and Penman (2001) show that past sales growth is a better predictor of future equity growth, so the growth in book equity similarly follows a first-order autoregressive process with an autorocorrelation coefficient equal to the long-run average rate of mean reversion in sales growth and a mean equal to the long-run GDP growth rate. Following Dechow, Sloan, and Soliman (2004) and Nissim and Penman (2001) the autorrelation coefficients for ROE and Sales growth used are 0.57 and 0.24, respectively, long-run cost of equity is assumed to be 12%, long-run growth in GDP is assumed to be 6%, and the terminal period T is assumed to be 10 years. Instead of the cash-flow duration estimate, its

percentile rank witihin each cross-section is used.

- Whited-Wu Rank - Financial constraints index rank, estimated following Whited and Wu (2006). The index is calculated as:

$$\begin{aligned}\text{WW}_{i,t} = -0.091 \times \text{CF}_{i,t} - 0.062 \times \text{DIVPOS}_{i,t} + 0.021 \times \text{TLTD}_{i,t} - 0.044 \times \text{LNTA}_{i,t} \\ + 0.102 \times \text{ISG}_{i,t} - 0.035 \times \text{SG}_{i,t}\end{aligned}$$

  where CF is the ratio of cash flow to total assets, DIVPOS is the cash dividend indicator variable, TLTD is the ratio of the long-term debt to total assets, LNTA is log of total assets, ISG is the firms' three-digit industry growth, and SG is sales growth. The variables use the most recent COMPUSTAT quarterly observations. Instead of the Whited-Wu index estimate, its percentile rank witihin each monthly cross-section is used.

- Cash Flow Volatility - Standard deviation over the last 20 quarters of cash flows, measured by operating cash flow scaled by total assets. Operating cash flow is measured by sales (SALEQ) minus cost of goods sold (COGSQ) minus selling, administrative and general expenses (XSGAQ) minus working capital change (WCAPQ minus lagged WCAPQ). A minimum of 8 consecutive quarters is required.

- Operating Profitability - Sales (SALEQ) minus cost of goods sold (COGSQ), scaled by market value of assets. Market value of assets equals total assets (ATQ) minus shareholder equity (SEQQ) plus market capitalization.

B Driver vs Stabilizer Effect of Monetary Policy

For simplicity, we assume that we live in an endowment economy so that consumption is equal to the output, y_t, and follows the following diffusion process in continous time

$$\frac{dy_t}{y_t} = \mu dt + \theta_y dw_y$$

where μ is the drift and dw_y is the Brownian increment and σ_y is the standard deviation of these shocks to output growth.[17] Output is affected both by real shocks dw_A and by changes in monetary policy dm so that

$$\theta_y dw_y = \theta_A dw_A + \theta_M dm$$

from which we have

$$\frac{dy_t}{y_t} = \mu dt + \theta_A dw_A + \theta_M dm.$$

where both θ_A and θ_M are positive so that an increase in money supply (or reduction in interest rate) increases output.

Moreover, changes is monetary policy is driven by real shocks, dw_A, and pure policy shocks uncorrelated with real shocks, dw_M, that is,

$$dm = -\eta_A dw_A + \eta_M dw_M$$

where $-\eta_A < 0$ captures the stabilizer effect, the fact that monetary policy tightens in response to expansionary real shocks, and $\eta_M > 0$ captures the effect of expansionary pure policy surprises. Without loss of generality we are assuming $\eta_A^2 + \eta_M^2 = 1$ so that dm has a unit variance, which simplifies our calculations below.

Assuming constant relative risk aversion preferences with relative risk aversion coefficient of γ and discount rate ρ, the stochastic discount factor between time zero and t becomes $\Lambda_{0,t} = \exp(-\rho t)\, c_t^{-\gamma}/c_0^{-\gamma}$, which leads to the diffusion process

$$\begin{aligned} d\ln\Lambda &= -\rho dt - \gamma d\ln y_t \\ &= -\rho dt - \gamma\left[\left(\mu - \frac{1}{2}\theta_y^2\right) dt + \theta_y dw_y\right] \end{aligned}$$

[17] We omit time-subscripts throughout for the sake of parsimony. Similar results can be obtained with discrete time using the log-normal approximation in Cochrane (2001). We use continuous time as it yields closed-form solutions without any approximation.

where the second line follows from Ito's lemma. Another round of Ito's lemma yields

$$
\begin{aligned}
\frac{d\Lambda}{\Lambda} &= -\left(\rho+\gamma\left(\mu-\frac{1}{2}\theta_y^2\right)-\frac{1}{2}\gamma^2\theta_y^2\right)dt-\gamma\theta_y dw_y \\
&= -\left(\rho+\gamma\left(\mu-\frac{1}{2}\theta_y^2\right)-\frac{1}{2}\gamma^2\theta_y^2\right)dt-\gamma\left[\theta_A dw_A+\theta_M dm\right] \\
&= -\left(\rho+\gamma\left(\mu-\frac{1}{2}\theta_y^2\right)-\frac{1}{2}\gamma^2\theta_y^2\right)dt-\gamma\left[\left(\theta_A-\eta_A\theta_M\right)dw_A+\eta_M\theta_M dw_M\right]
\end{aligned}
$$

Stocks react both to real shocks and changes in money supply, that is, for a stock i, we have

$$
\begin{aligned}
dR_i &= \left(r+\phi_i\right)dt+\sigma_{A,i}dw_A+\sigma_{M,i}dm+\sigma_i dw_i \\
&= \left(r+\phi_i\right)dt+\sigma_{A,i}dw_A+\sigma_{M,i}\left(-\eta_A dw_A+\eta_M dw_M\right)+\sigma_i dw_i \\
&= \left(r+\phi_i\right)dt+\left(\sigma_{A,i}-\eta_A\sigma_{M,i}\right)dw_A+\sigma_{M,i}\eta_M dw_M+\sigma_i dw_i
\end{aligned}
$$

where $\sigma_{A,i}$ and $\sigma_{M,i}$ capture the exposure of stock i to real shocks and policy changes respectively and dw_i is the individual (non-systematic) risk of the stock. (Our empirical approach allows us to find $\sigma_{M,i}$ by identifying pure policy shocks dw_M.)

The term ϕ_i captures the risk premium given by the product of price and quantity of risk,

$$
\begin{aligned}
\phi_i &= -\frac{1}{dt}E_t\left[\frac{d\Lambda}{\Lambda}dR_i\right] \\
&= \sigma_{A,i}\gamma\theta_A+\sigma_{M,i}\gamma\theta_M \\
&= \gamma\left(\sigma_{A,i}-\eta_A\sigma_{M,i}\right)\left(\theta_A-\eta_A\theta_M\right)+\gamma\sigma_{M,i}\eta_M^2\theta_M .
\end{aligned}
$$

Hence the effect of the monetary policy exposure on risk premium is given by

$$
\begin{aligned}
\frac{\partial\phi_i}{\partial\sigma_{M,i}} &= \gamma\left[-\eta_A\left(\theta_A-\eta_A\theta_M\right)+\eta_M^2\theta_M\right] \\
&= \gamma\left[\theta_M-\eta_A\theta_A\right].
\end{aligned}
$$

where the second line follows from $\eta_A^2+\eta_M^2=1$. Here, the first term in square brackets cap-

tures the "driver" effect that works through the effect of monetary policy on real variables as in New Keynesian models whereas the second term captures the "stabilizer/mitigator" effect of monetary policy that works through the contractionary response of monetary policy to expansionary real shocks. The net effect will depend on whether the driver effect is dominated by the stabilizer effect.

Moreover, we also see that even the conditional version of the standard CAPM would not capture the risk premia correctly since we have two aggregate shocks, dw_A and dw_M.

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