Can Inflation and Monetary Policy Predict Asset Prices?

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Abstract: We develop a continuous-time endowment economy model of the US with inflation and the central bank's interest rate adjustments as observable risk factors. We show that they have predictive power for consumption growth and can explain many features of the aggregate stock and bond market. We derive the price-dividend ratio, the equity premium, the risk-free rate, and the term structure of interest rates. We show in a calibrated model that inflation and the federal funds rate adequately predict those key asset pricing moments. The model offers a novel mechanism to explain the variation in the aggregate price-dividend ratio and the risk-free rate as it relies on observable rather than latent risk factors.

Keywords: Asset pricing, equity premium, federal funds rate, inflation, monetary policy, term structure of interest rates

JEL subject codes: E43, E44, G12

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1 Introduction

Asset prices are influenced by a multitude of factors with macroeconomic variables playing a pivotal role in shaping market behavior. A particularly important macroeconomic factor affecting asset prices is inflation, which is known to have a strong impact on economic growth (e.g., Fischer, 1993; Barro, 2013). To steer inflation rates and prevent them from reaching too high levels, central banks adjust interest rates, thereby affecting asset prices as well. Indeed, it is well documented that central banks' announcements on interest rate decisions often cause strong reactions in stock and bond markets, leading to increased return volatility and excess returns around the announcement days (e.g., Savor and Wilson, 2013; Lucca and Moench, 2015).

We are interested in the interplay between asset prices, inflation, and monetary policy and aim to answer the central research question raised in the title of this paper: Can inflation and monetary policy predict asset prices? To answer this question, we develop a continuous-time reduced-form endowment economy model of the US with inflation and the FED's interest rate adjustments as observable macroeconomic risk factors. Our model builds on a standard Barro (2006, 2009)-type disaster risk framework and imposes exogeneously given inflation and interest rate dynamics. We model inflation dynamics by a Vasicek-type process, whose mean-reversion level depends on the federal funds rate.[1] We assume that the nominal federal funds rate is set in every FOMC meeting,[2] which takes place eight times a year at fixed dates. From the agent's point of view those interest rate adjustments happen exogeneously and not in a completely predictable manner.

Our model introduces a novel mechanism to explain key characteristics of the aggregate stock and bond market by relying exclusively on observable macroeconomic factors. Traditional asset pricing models often rely on latent risk factors, such as long-run risk and stochastic volatility (e.g., Bansal and Yaron, 2004), time-varying disaster intensities (e.g., Wachter, 2013), or latent regime switches (e.g., Song, 2017) to explain the complexities of asset markets. While these factors can be inferred from macroeconomic data, they lack direct macroeconomic meaning, such as inflation or unemployment rates. Although traditional models provide valuable insights, their reliance on unobservable factors can limit their applicability and practical utility. In contrast, our model focuses solely on observable macroeconomic factors, offering a more straightforward and empirically grounded framework for understanding asset prices, although the dynamics of those factors still need to be estimated from the data.

[1] The federal funds rate is the interest rate at which depository institutions lend federal funds uncollateralized to each other overnight.

[2] FOMC is short for the Federal Open Market Committee, which is the body within the Federal Reserve that sets monetary policy. This committee holds eight regularly scheduled meetings per year and other meetings as needed (Federal Reserve, 2021).

Our analysis proceeds in several steps. First, we fit a model of consumption growth to US data. We find that inflation and interest rates both have explanatory power for future economic growth. Then, we embed the estimated fundamental dynamics in a general equilibrium asset pricing model with recursive preferences (Epstein and Zin, 1989; Weil, 1989) and provide analytical results for key asset pricing moments. Finally, our quantitative analysis reveals that the model successfully matches key asset pricing moments, both unconditional and conditional on the prevailing macroeconomic conditions. In particular, we find that inflation increases the equity premium and higher interest rates lower the equity premium. Besides, we can replicate the observed patterns and time variation in the price-dividend ratio and the risk-free rate. Our model also produces an unconditional equity premium, price-dividend ratio, and a risk-free rate that align with empirical observations. Besides, the discontinuous nature of our interest rate process allows the model to explain significant changes in asset prices following the FED's interest rate announcements, offering a novel perspective on the impact of monetary policy on financial markets. Additionally, the calibrated model generates reasonable nominal term structures for government bond yields and can explain under which macroeconomic conditions the term structure is upward sloping, flat, or downward sloping. In our simulations as well as in the data, an upward sloping nominal term structure is the normal case although it can also produce an inverse term structure if inflation rates are high compared to the federal funds rate.

Our paper contributes to the asset pricing literature about inflation as a priced risk factor. Within this field of literature, our paper is related to a number of recent works that study theoretical asset pricing models linking inflation to consumption growth. Hasseltoft (2009) shows that a long-run risk framework with time-varying first and second moments of consumption growth, inflation, and dividend growth can jointly explain key features of the stock and bond market. Hasseltoft and Burkhardt (2012) extend the model of Hasseltoft (2009) by incorporating a two-state Markov chain to govern expected consumption growth and expected inflation allowing for procyclical and countercyclical inflation regimes. They find that the time-varying correlation of consumption growth and inflation explains asset prices and correlations. David and Veronesi (2013) estimate a general equilibrium model with Markov switching dynamics for fundamentals and learning and contend that the time-varying signaling role of inflation drives the joint dynamics of stock and bond markets. Dergunov et al. (2022) build on the work of David and Veronesi (2013) and study a parsimonious regime switching long-run risk model with learning. They find that inflation might be relevant for the pricing of real assets because low consumption growth tends to be associated with either very high or very low inflation. In a similar vein, Piazzesi and Schneider (2007) analyze the role of inflation as a signal about future consumption growth. They document that news about expected future consumption growth and news about expected future inflation are negatively correlated in the data, allowing them to explain the upward sloping behavior of the nominal yield curves. Similarly, Bansal and Shaliastovich (2013) build on the negative correlation between inflation and consumption and develop a long-run risk model with time-varying volatilities of expected growth and inflation that can generate plausible risk premium variation in stocks, bonds, and currencies.

While the above mentioned papers abstract from the role of monetary policy, our paper also contributes to the strand of literature studying the theoretical implications of monetary policy decisions on asset prices. Song (2017) studies the behavior of the Treasury yield curve and the sign-switching stock-bond return correlation in a long-run risk model that allows for regime switches in the aggressiveness of monetary policy and in the conditional covariance of macroeconomic shocks. Campbell et al. (2020) analyze the stock-bond correlation in a New Keynesian model with habit formation preferences and monetary policy regimes. They find that a switch to accommodating monetary policy combined with increased volatility in the inflation target can explain negative stock-bond correlations. Similarly, using a bivariate regime-switching model for output and inflation, Baele and van Holle (2017) find that negative stock-bond correlations are associated with periods of accomodative monetary policy, but only in times of low inflation. Pflueger and Rinaldi (2022) build on Campbell et al. (2020) and develop a New Keynesian model of inflation and monetary policy with endogenously time-varying risk premia via habit formation preferences. Their model is able to explain the large fall in the stock market and the increase in long-term bond yields in response to a surprising policy rate increase. Piazzesi (2005) constructs a continuous-time model of the joint distribution of bond yields and the interest rate target set by the FOMC. She finds that introducing monetary policy improves yield curve models and allows to model important seasonalities around FOMC meetings. Similarly, Ang et al. (2011) evaluate the impact of changing monetary policy on the entire term structure using a reduced-form no-arbitrage model with drifting coefficients in the interest rate rule. Other papers focusing on the effects of monetary policy on the term structure include Ireland (2007), Bikbov and Chernov (2013), and Shaliastovich and Yamarthy (2015), among others.

Our model is distinctive from the above mentioned literature in the sense that it combines both strands of literature in a unified parsimonious theoretical asset pricing model. Although Song (2017), Campbell et al. (2020), and Pflueger and Rinaldi (2022) combine inflation and monetary policy in an asset pricing model and study features of stock and bond markets, our model is different from theirs. While Campbell et al. (2020) focus on explaining stock-bond correlations and unconditional asset pricing moments in a model with constant expected growth rates, we focus on matching conditional asset pricing moments. Similarly, Pflueger and Rinaldi (2022) study an asset pricing model with inflation and monetary policy but only focus on the effect of surprises in monetary policy decisions and do not study conditional asset pricing moments. Our paper is also distinctive from Song (2017) and more general the whole literature using latent factors in that sense that it relies on observable risk factors only, resolving one of the main limitations of the long-run risk model and other consumption-based models using latent factors. By identifying the federal funds rate and the inflation rate as great predictors of asset prices, our model is able to not only explain unconditional moments of the aggregate stock and bond markets but also conditional moments. This distinguishing feature of our model allows us to replicate important findings in the empirical literature on the interplay between asset returns, inflation, and monetary policy that the other models cannot. First, our model produces an equity premium that increases in the inflation rate.[3] Second, our model can explain significant changes in asset prices following the FED's interest rate announcements.[4] Third, our model is able to replicate the historical pattern and time-variation in the price-dividend ratio. Fourth, our model produces reasonable shapes of the term structure of interest rates.

The remainder of this paper is organized as follows. Section 2 describes our data and provides some empirical evidence regarding the relation between consumption, inflation, and the federal funds rate. Section 3 introduces our reduced-form asset pricing model, while Section 4 presents a number of theorems characterizing the key equilibrium asset pricing variables, such as the risk-free rate, the equity premium, the price-dividend ratio, and the term structure of interest rates. Section 5 presents the calibration and Section 6 investigates the quantitative asset pricing implications of the model. Section 7 provides sensitivity analyses. Finally, Section 8 concludes. An appendix provides additional material, such as proofs, an outline of the numerical solution approach, further estimation results and calibration details, and additional numerical analyses.

2 Empirical Evidence

In this section, we describe our data and provide estimation results for the empirical relation between consumption growth, inflation rates, and the federal funds rate. Moreover, we provide summary statistics for financial quantities, such as stock returns and bond yields.

[3] There is an extensive empirical literature on stock returns and inflation. Numerous studies have established a negative relation between inflation and stock returns. Early contributions to this field include Fama (1981),Geske and Roll (1983), Chen et al. (1986), Stulz (1986), Ferson and Harvey (1991), and Marshall (1992). More recently, Cohen et al. (2005), Bekaert and Wang (2010), and Schmeling and Schrimpf (2011) find predominantly negative inflation betas. Using a consumption-based asset pricing model, Boons et al. (2020) also demonstrate that inflation risk is priced into stock returns. Other papers in this area include, among others, Brandt and Wang (2003), Katz et al. (2017), and Fang et al. (2022).

[4] There is also an extensive literature empirically analyzing the effects of central banks' interest rate announcements on stock and bond prices. Those studies find that central banks' interest rate announcements have a huge impact on asset prices and risk premia (e.g., Bomfim, 2003; Bernanke and Kuttner, 2005; Savor and Wilson, 2013; Lucca and Moench, 2015; Liu et al., 2022).

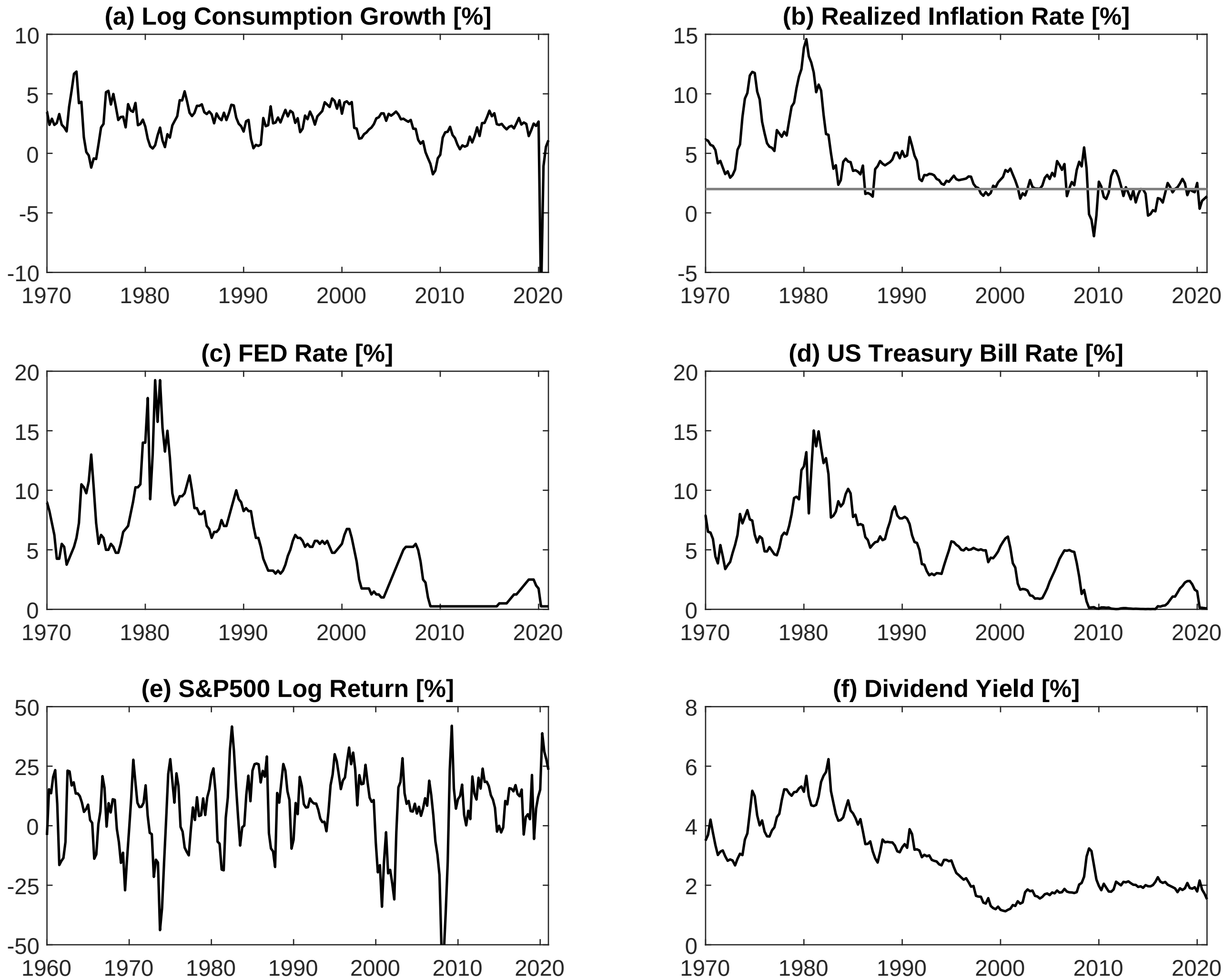

Figure 1: Quarterly US Data (Annualized). The figure depicts (a) the annualized real consumption growth rate, (b) the inflation rate, (c) the federal funds rate, (d) the yields on 3-month US Treasury bills, (e) the S&P500 return series, and (f) the dividend yield on the S&P500. The grey line in Graph (b) depicts the FED's 2% inflation target that has been valid since 2012.

2.1 Data and Summary Statistics

We measure aggregate consumption with quarterly NIPA data of real consumption expenditures on nondurables and services available from the Bureau of Economic Analysis. We use the growth rate in the consumer price index (CPI), obtained from the Bureau of Labor Statistics, as inflation measure.

Macroeconomic Data					
(a) Consumption Growth		(b) Inflation Rate		(c) Federal Funds Rate	
$\mathbb{E}[\Delta c]$	2.66	$\mathbb{E}[\pi]$	3.67	$\mathbb{E}[i]$	5.03
$\sigma[\Delta c]$	1.79	$\sigma[\pi]$	2.83	$\sigma[i]$	3.74
$\min[\Delta c]$	−12.34	$\min[\pi]$	−1.96	$\min[i]$	0.25
$q_{5\%}[\Delta c]$	0.04	$q_{5\%}[\pi]$	0.87	$q_{5\%}[i]$	0.25
$q_{95\%}[\Delta c]$	5.30	$q_{95\%}[\pi]$	10.32	$q_{95\%}[i]$	11.63
$\max[\Delta c]$	6.87	$\max[\pi]$	14.59	$\max[i]$	19.25
$\mathrm{AC}[\Delta c]$	66.00	$\mathrm{AC}[\pi]$	95.50	$\mathrm{AC}[i]$	94.93

Nominal Financial Data					
(d) Treasury Bill Rate		(e) Stock Market Return		(f) Dividend Yield	
$\mathbb{E}[r^f]$	4.45	$\mathbb{E}[r^s]$	9.97	$\mathbb{E}[y_d]$	2.91
$\sigma[r^f]$	3.17	$\sigma[r^s]$	15.58	$\sigma[y_d]$	1.13
$\min[r^f]$	0.02	$\min[r^s]$	−52.62	$\min[y_d]$	1.13
$q_{5\%}[r^f]$	0.05	$q_{5\%}[r^s]$	−19.86	$q_{5\%}[y_d]$	1.38
$q_{95\%}[r^f]$	9.84	$q_{95\%}[r^s]$	27.86	$q_{95\%}[y_d]$	5.14
$\max[r^f]$	15.02	$\max[r^s]$	46.65	$\max[y_d]$	6.24
$\mathrm{AC}[r^f]$	96.62	$\mathrm{AC}[r^s]$	76.03	$\mathrm{AC}[y_d]$	97.58

Table 1: Summary Statistics. The table reports summary statistics of the relevant variables. The sample period is from January 1960 until December 2020. All reported numbers are in percentage terms. Moments are calculated using overlapping annual observations constructed from quarterly US data.

Data on the federal funds rate is retrieved from the FRED.[5] Aggregate stock market data consists of monthly observations of returns, dividends, and prices of the S&P500 obtained from the homepage of Robert Shiller at Yale University.[6] Data on the 3-month US Treasury bill rate and on US Treasury bond yields with maturities from one to 30 years is obtained from the FRED. Our sample period is from January 1960 to December 2020.

Figure 1 shows time series plots of the data and Table 1 provides summary statistics. All data is annualized.[7] Panel (a) depicts the log growth rate of real consumption. Here some cyclical pattern is noticeable, representing normal fluctuations in the business cycle. Large drops in consumption are observable in the aftermath of the financial crisis in 2009. The large drop in consumption growth in March 2020 is striking. Here consumption fell by more than 12% due to the COVID-19 shock, see also Table 1. This challenges the assumption of normally distributed consumption data, as such a large drop can hardly be explained by the relatively low volatility of consumption growth. Panel (b) depicts

[5] Note that we use the federal funds rate rather than the effective federal funds rate. Since 2008 the federal funds rate is given as a range of 25bps (e.g., 1 – 1.25%). In those cases we use the upper bound of the range.

[6] See `http://www.econ.yale.edu/~shiller/`.

[7] As in Wachter (2013), data moments for consumption growth rates, inflation rates and stock returns in Panels (a), (b), (e), and (f) are calculated using overlapping annual observations constructed from quarterly US data.

the realized inflation rate. We observe that during the 1970's and the early 1980's the US economy underwent a stagflationary period with inflation rates far above the 2% inflation benchmark. With the appointment of Paul Volcker as Chairman of the Federal Reserve, the disinflation period began and inflation rates gradually reverted to the 2% level. In the aftermath of the financial crisis, inflation rates became negative and subsequently stayed at a relatively low level. Panel (c) depicts the federal funds rate. Here we can identify a strong positive comovement with inflation rates. In particular, we see very high interest rates during the 1970's and the early 1980's, which is exactly the time when inflation rates were unusually high. With inflation rates normalizing, the federal funds rate also decreases, reaching the zero lower bound from 2009 onward. The autocorrelation of both the inflation rate and the federal funds rate is about 95%, indicating that both processes are highly persistent, see Table 1.

Turning to financial data, Panel (d) depicts the 3-month US Treasury bill rate, which follows a similar pattern as the federal funds rate with a comparable autocorrelation of 97%. However, Treasury bill rates are typically slightly lower than the federal funds rate. Panel (e) shows the log returns of the S&P500. We find large drops in the returns during the 1970's as well as after the bursting of the dotcom bubble in the early 2000's and in the aftermath of the financial crisis in 2008. In particular, we can observe here a positive comovement with log consumption growth. Finally, Panel (f) shows the dividend yield series on the S&P500. Here we can observe that the dividend yield follows a similar pattern as the federal funds rate, i.e., it is high in times of high inflation and decreases as inflation rates fall. As the federal funds rate, the dividend yield is a very persistent process with an autocorrelation of about 97% as shown in Table 1.

2.2 Empirical Evidence

We start our empirical analysis by estimating a three-dimensional p-th order vector autoregression model for real consumption growth, inflation, and the federal funds rate. We define the vector $y_t = (g_t, \pi_t, i_t)^\top$, where g_t denotes annualized log consumption growth from time $t-1$ to t. Similarly, π_t is the inflation rate, and i_t is the federal funds rate. We run the following three-dimensional VAR(p)-model

$$y_{t+1} = \beta_0 + \sum_{k=1}^{p} \beta_k y_{t-k} + \varepsilon_t,$$

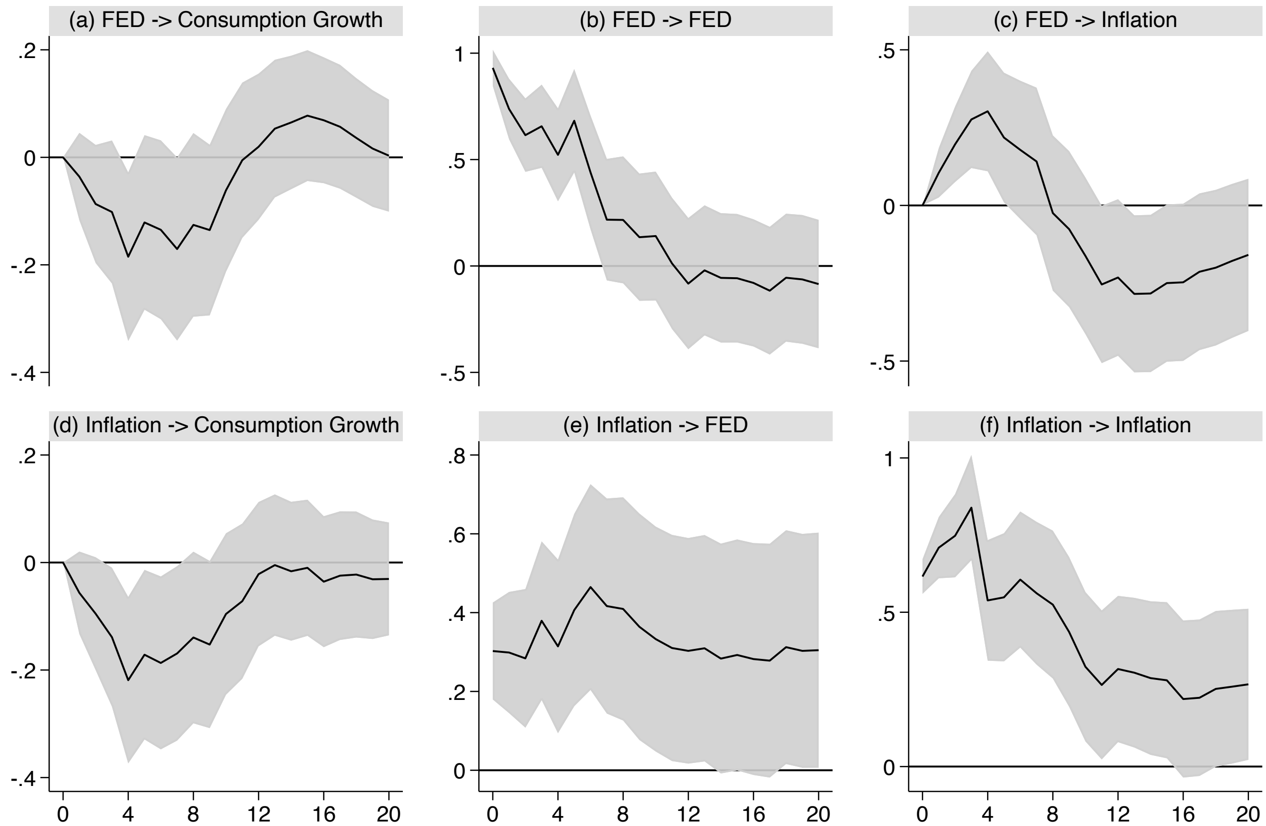

Figure 2: VAR(p)-Model: Impulse Response Functions. The figure depicts the Cholesky orthogonalized impulse responses of real consumption growth, the federal funds rate and the inflation rate to a one-standard deviation shock in the federal funds rate and in the inflation rate, respectively. The impulse response functions are calculated based on a VAR(9)-model for real consumption growth, inflation, and the federal funds rate (FED). Shaded areas indicate a 95% confidence interval. The horizontal axis represents time steps measured in quarters. Our sample period is from January 1960 until December 2020.

where β_0 is a three-dimensional vector of constants, β_k, $k = 1, \ldots, p$, are time-invariant (3×3) matrices, and ε_t is a three-dimensional vector of error terms (Sims, 1980). We first determine the optimal lag order with the Akaike information criterion resulting in nine lags.[8]

Figure 2 shows the impulse response functions to a one-standard deviation shock in inflation and the federal funds rate, respectively. The impulse response is based on the Cholesky orthogonalization of the

[8]Likelihood-ratio tests and the final prediction error criterion also pinpoint to an optimal lag order of nine. On the other hand, the Hannan-Quinn information criterion (Hannan and Quinn, 1979) suggests six lags. The qualitative results and the inferred stylized facts on the interplay between consumption growth, inflation, and the federal funds rate do not change if we run a VAR(6)-model instead of a VAR(9)-model and the model fit is about the same. The Bayesian information criterion (Schwarz, 1978) suggests an optimal lag order of one. We discuss VAR(1)-models in Appendix C.1.

VAR-model in which inflation shocks are ordered first. From the figure, we can infer several key stylized facts that align well with the existing macroeconomic literature: (a) Positive interest rate shocks tend to have a negative short-term impact on real economic growth. However, the long-term effect of such shocks is positive, as higher interest rates initially dampen investment and consumption, but over time they contribute to stabilizing inflation and fostering sustainable growth. (b) Positive shocks to the federal funds rate lead to a short-term rise in the federal funds rate itself. This effect gradually diminishes over time, suggesting a mean-reverting behavior of the federal funds rate. This reversion reflects the Federal Reserve's tendency to adjust rates back towards a neutral level after addressing temporary shocks. (c) Positive interest rate shocks also have an immediate, positive effect on inflation rates. In the short-run, higher interest rates may signal stronger inflationary pressures or more persistent inflation expectations. However, over the long-run, increased interest rates help curb inflation, as they suppress demand and slow down price growth, eventually bringing inflation under control. Similar effects have recently been documented by Ferreira et al. (2024) for Eurozone data.

Focusing on inflation shocks, we find that (d) positive inflation shocks tend to reduce real economic growth in both the short and long-run. While the negative impact is most pronounced initially, it gradually weakens over time as the economy adjusts to the higher price levels and inflationary pressures subside. (e) Following positive inflation shocks, the Federal Reserve typically responds by raising interest rates. This policy action aims to anchor inflation expectations and prevent further upward pressure on prices. (f) Positive inflation shocks temporarily drive inflation rates higher. However, this increase is usually transitory, as inflationary pressures dissipate over time, either due to policy interventions (such as interest rate hikes) or natural adjustments in the economy.

3 Model Setup

We consider a continuous-time reduced-form model of an endowment economy with a representative agent and study the asset pricing effects of monetary policy uncertainty. The model builds on a standard Lucas (1978) tree model and assumes that consumption growth follows a jump-diffusion process. The model captures the economic intuition gained in the empirical section, but to keep the setting simple, we impose exogenously given inflation and interest rate dynamics. Based on the prevailing interest and inflation rate, the representative agent forms expectations about the FED's next interest rate decision. Figure 3 summarizes the structure of the model.

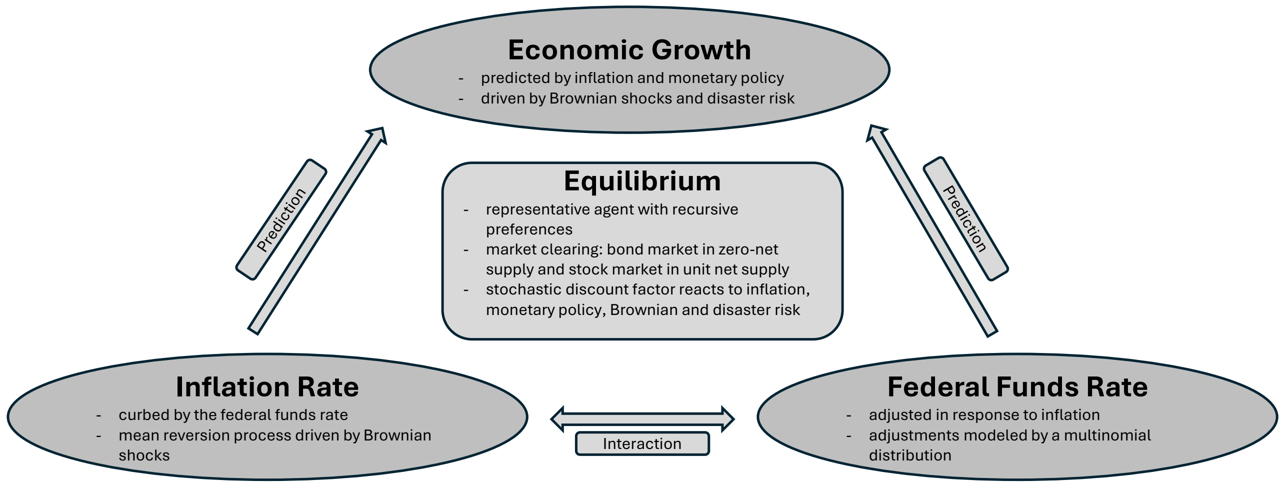

Figure 3: Model Structure. This figure illustrates the structure of the model and the interplay of its building blocks as outlined in Section 3.

3.1 The Economy

Aggregate Consumption Denoting the prevailing inflation rate by $\pi = (\pi_t)_{t\geq 0}$ and the federal funds rate set by the FED by $i = (i_t)_{t\geq 0}$, real aggregate consumption $c = (c_t)_{t\geq 0}$ follows the jump-diffusion process

$$\frac{\mathrm{d}c_t}{c_{t-}} = \mu_c(\pi_t, i_t)\mathrm{d}t + \sigma_c \mathrm{d}W_t^c - \ell_t \mathrm{d}N_t, \tag{3.1}$$

where $W^c = (W_t^c)_{t\geq 0}$ is a standard Brownian motion, and $N = (N_t)_{t\geq 0}$ is a Poisson process capturing macroeconomic disasters with constant jump intensity λ_c, i.e., the probability for a jump to occur during the infinitesimal interval $[t, t+\mathrm{d}t]$ is $\lambda_c\,\mathrm{d}t$.[9] The parameter $\ell_t \in (0,1)$ denotes the corresponding jump size, i.e., the loss of aggregate consumption when a macroeconomic disaster hits the economy, which is a random variable whose time-invariant distribution is independent of the Brownian and Poissonian shocks in the model. Conditional on no disasters, $\mu_c(\pi, i)$ is the expected consumption growth, which depends on the inflation rate and the federal funds rate to capture the effects of shocks to those vari-

[9] The model could be easily extended to time-varying disaster risk in the spirit of Wachter (2013) or to a disaster intensity which depends on the inflation rate and the federal funds rate. We study such an extension in Appendix E.4. Besides, one could also account for the duration of crises, as in Branger et al. (2016). However, to focus on the novel implications, we abstract from those extensions. Moreover, in some models, time-varying disaster intensities reduce the model's ability to match the data, e.g., Gabaix (2012).

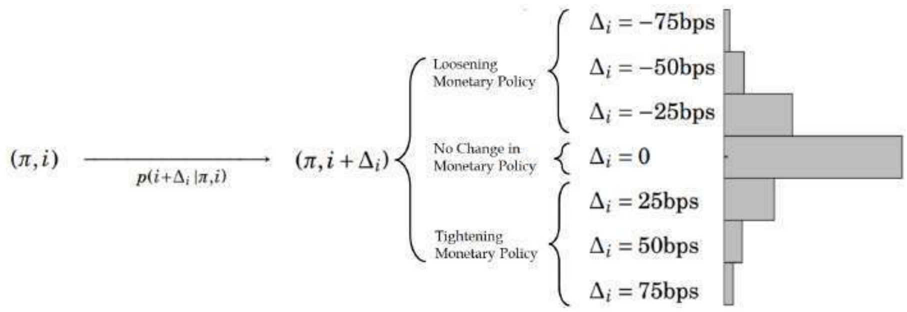

Figure 4: Mechanism of Interest Rate Adjustments. The figure illustrates the mechanism of interest rate adjustments in our model that take place eight times a year at fixed dates $\frac{n}{8}$, $n \in \mathbb{N}$. At those dates, the FED adjusts interest rates from i_{t-} to $i_t = i_{t-} + \Delta_i$ according to a probability distribution $p(\Delta_i \mid \pi, i)$ taking the current state (π, i) into account.

ables on economic growth as outlined in the previous section. Finally, σ_c is the constant volatility of consumption shocks.[10]

Inflation Rate The inflation rate π follows a modified mean-reverting Ornstein-Uhlenbeck process with state-dependent mean-reversion level

$$d\pi_t = \mu_\pi(\pi_t, i_t)dt + \sigma_\pi\left(\rho_{c\pi}dW_t^c + \sqrt{1-\rho_{c\pi}^2}dW_t^\pi\right), \tag{3.2}$$

where $W^\pi = (W_t^\pi)_{t\geq 0}$ is another standard Brownian motion independent of W^c. The parameter $\rho_{c\pi}$ denotes the instantaneous diffusive correlation coefficient between the two Brownian shocks to the consumption and the inflation process. The drift rate of inflation μ_π may explicitly depend on both inflation and the federal funds rate, and σ_π denotes the constant volatility of inflation shocks.

Federal Funds Rate Although the nominal federal funds rate i is set in every FOMC meeting, from the agent's point of view interest rate adjustments happen exogenously and not in a completely predictable manner.[11] Interest rate adjustments take place eight times a year at fixed dates $\frac{n}{8}$, $n \in \mathbb{N}$, and the nominal interest rate takes on values in $\mathscr{I} \in 0.0025\mathbb{N}$.

[10]One can show that in equilibrium a simple AK-production economy as in Pindyck and Wang (2013) can imply consumption dynamics similar to (3.1) if the capital dynamics are adequately specified.

[11]Notice that, according to its dual mandate, the FED seeks to promote the two coequal, albeit sometimes conflicting, objectives of maximum employment and price stability (Mishkin, 2007). Therefore, its interest rate adjustments are influenced by many factors, which are not fully captured by a Taylor rule, making it hard to predict future interest rate decisions. To focus on the novel asset pricing implications, we thus abstract from explicitly modeling the rational behind the FED's interest rate adjustments and instead propose a reduced-form model, where we estimate the distribution of interest rate adjustments from historical data.

We assume that the agent's subjective probability that the FED adjusts interest rates from $i \in \mathscr{I}$ to $i + \Delta_i \in \mathscr{I}$ in a FOMC meeting depends on (π, i) and is denoted by $\mathbb{P}(\Delta_i \mid \pi, i)$. Figure 4 illustrates the mechanism of interest rate adjustments in our reduced-form model. If $\Delta_i < 0$, we observe loosening monetary policy, if $\Delta_i > 0$, we observe tightening monetary policy. Between FOMC meetings, the interest rate is assumed to be constant.

Dividend Dynamics We think of the stock (market) as a claim to aggregate dividends $D = (D_t)_{t \geq 0}$. Empirically, dividends are more variable than consumption and more sensitive to macroeconomic disasters (Longstaff and Piazzesi, 2004). We assume that dividends are driven by an additional shock that is orthogonal to the consumption and inflation shocks (e.g., Bansal and Yaron, 2004; or Zhou and Zhu, 2015 for a continuous-time model), i.e.,

$$\frac{\mathrm{d}D_t}{D_{t-}} = \mu_d(\pi_t, i_t)\mathrm{d}t + \sigma_d\big(\rho_{cd}\mathrm{d}W_t^c + \widehat{\rho}_{d\pi}\mathrm{d}W_t^\pi + \widehat{\rho}_d\mathrm{d}W_t^d\big) + \big((1-\ell_t)^\phi - 1\big)\mathrm{d}N_t. \tag{3.3}$$

Here, μ_d denotes the expected dividend growth rate in normal times, and $\phi > 1$ is a leverage parameter for the dividend loss when a disaster hits the economy. $W^d = (W_t^d)_{t \geq 0}$ is a third Brownian motion, independent of W^c, W^π, ℓ, and N. The parameter ρ_{cd} is the instantaneous correlation between consumption and dividends, and

$$\widehat{\rho}_{d\pi} = \frac{\rho_{d\pi} - \rho_{cd}\rho_{c\pi}}{\sqrt{1-\rho_{c\pi}^2}}, \qquad \widehat{\rho}_d = \sqrt{1-\rho_{cd}^2 - \widehat{\rho}_{\pi d}^2},$$

where $\rho_{\pi d}$ denotes the instantaneous correlation between inflation and dividends.[12] For succinctness of notation, we will let $W = [W^c,\ W^\pi,\ W^d]^\top$ denote the three-dimensional Brownian motion and the corresponding volatility vectors

$$\Sigma_c = \sigma_c[1,\ 0,\ 0]^\top, \qquad \Sigma_\pi = \sigma_\pi\big[\rho_{c\pi},\ \sqrt{1-\rho_{c\pi}^2},\ 0\big]^\top, \qquad \Sigma_d = \sigma_d\big[\rho_{cd},\ \widehat{\rho}_{d\pi},\ \widehat{\rho}_d\big]^\top$$

in the following.

[12] To preserve tractability, Abel (1999) and Wachter (2013), among others, model dividends as leveraged consumption, i.e., $D = c^\phi$ with leverage parameter $\phi > 1$, implying a perfect co-movement of dividends and consumption. Our dynamics are more general and contain this special case if $\mu_d = \phi\mu_c + \frac{1}{2}\phi(\phi-1)\sigma_c^2$, $\rho_{cd} = 1$, $\sigma_d = \phi\sigma_c$, $\rho_{d\pi} = \rho_{c\pi}$.

3.2 Preferences

The representative agent has stochastic differential utility (SDU) introduced by Duffie and Epstein (1992a,b). These preferences are the continuous-time version of discrete-time recursive utility proposed in Kreps and Porteus (1978), Epstein and Zin (1989) and Weil (1989). The value function (syn. indirect utility function) of the representative agent J is defined by the recursion

$$J(t,c,\pi,i) = \mathbb{E}_t\Big[\int_t^\infty f\big(c_s, J(s,c_s,\pi_s,i_s)\big)\mathrm{d}s\Big]. \tag{3.4}$$

The aggregator function $f(c,J)$ is given by

$$f(c,J) = \begin{cases} \delta\theta J\left\{\left(\frac{c^{1-\frac{1}{\psi}}}{[(1-\gamma)J]^{\frac{1}{\theta}}}\right) - 1\right\}, & \text{for } \psi \neq 1, \\ \delta(1-\gamma)J\Big(\ln(c) - \frac{1}{1-\gamma}\ln\big((1-\gamma)J\big)\Big), & \text{for } \psi = 1, \end{cases}$$

where $\delta > 0$ is the subjective time preference rate, $\gamma > 1$ denotes the agent's risk aversion, ψ is the elasticity of intertemporal substitution (EIS), and $\theta = \frac{1-\gamma}{1-1/\psi}$. Note that in the special case $\gamma = \frac{1}{\psi}$, the preference structure collapses to time-additive CRRA utility. We assume that $\gamma > \frac{1}{\psi}$, i.e., we assume that the agent has a preference for early resolution of uncertainty.

3.3 Value Function

The indirect utility function J depends on the three state variables consumption c, inflation rate π, and federal funds rate i. Although the representative agent has an infinite time horizon, the indirect utility function also depends on time t since the FOMC meetings take place eight times a year after fixed time intervals, destroying the problem's time-homogeneity. Consequently, the problem is repeated after each FOMC meeting as in Wachter and Zhu (2022). Thus, the value function is a periodic function in the time dimension with period $T = \frac{1}{8}$, i.e.,

$$J(t,c,\pi,i) = J(t+nT,c,\pi,i), \qquad T = \frac{1}{8},\ n \in \mathbb{N}.$$

As a consequence of its periodicity, it is sufficient to determine the value function over the finite time interval $[0,T]$ only. When the FOMC has decided on the next interest rate adjustment Δ_i at time T, the indirect utility function jumps from $J(T_{-},c,\pi,i)$ to $J(T,c,\pi,i+\Delta_i)$. As the agent only knows a probability distribution of interest rate adjustments, this leads to the following boundary condition

$$J(T,c,\pi,i) = \sum_{\Delta_i \in \mathscr{I}} J(T,c,\pi,i+\Delta_i)\mathbb{P}(\Delta_i \mid \pi,i). \tag{3.5}$$

That is, the value function on the instant before the FOMC meeting equals the expectation of its value just after the announcement, i.e., $J(T,c,\pi,i) = \mathbb{E}_{T-}\left[J(T,c,\pi,i+\Delta_i)\right]$ (see Appendix A in Wachter and Zhu, 2022).

Applying Itô's Lemma, one can verify that the agent's value function (3.4) satisfies the following PDE between two FOMC meetings[13]

$$\begin{aligned} 0 = f(c,J) + J_t + J_c\mu_c(\pi,i)c + J_\pi\mu_\pi(\pi,i) + \frac{1}{2}J_{cc}c^2\sigma_c^2 + \frac{1}{2}J_{\pi\pi}\sigma_\pi^2 + J_{c\pi}c\sigma_c\sigma_\pi\rho_{c\pi} \\ + \lambda_c\,\mathbb{E}\left[J(t,c(1-\ell),\pi,i) - J\right] \end{aligned} \tag{3.6}$$

subject to the boundary condition (3.5). The value function is homogeneous of degree $1-\gamma$ in consumption, allowing us to solve a reduced-form value function with only two state variables (π,i) instead of three (c,π,i). The next theorem characterizes the reduced-form value function in our model.[14]

Theorem 3.1. *The value function is*

$$J(t,c,\pi,i) = \frac{1}{1-\gamma}c^{1-\gamma}G(t,\pi,i). \tag{3.7}$$

Here the function $G = G(t,\pi,i)$ satisfies the following PDE between two FOMC meetings

$$\begin{aligned} 0 = \delta\theta G^{1-1/\theta} + G_t + G\Big[(1-\gamma)\mu_c(\pi,i) - \gamma(1-\gamma)\frac{1}{2}\sigma_c^2 - \delta\theta + \lambda_c\mathbb{E}\left[(1-\ell)^{1-\gamma}-1\right]\Big] \\ + G_\pi\left[\mu_\pi(\pi,i) + (1-\gamma)\sigma_c\sigma_\pi\rho_{c\pi}\right] + G_{\pi\pi}\frac{1}{2}\sigma_\pi^2 \end{aligned} \tag{3.8}$$

subject to the boundary condition

$$G(T,\pi,i) = \sum_{\Delta_i \in \mathscr{I}} G(T,\pi,i+\Delta_i)\mathbb{P}(\Delta_i \mid \pi,i). \tag{3.9}$$

[13] Here, subscripts of J denote partial derivatives, e.g., $J_t = \frac{\partial J}{\partial t}$. For notational convenience we drop the dependencies of J and its derivatives.

[14] An alternative model specification where the federal funds rate follows a Markov chain whose transition intensity is time-dependent and increases dramatically around the FOMC meetings leads to an additional term in PDE (3.6) but avoids the technical issues stemming from the boundary condition (3.5) (cf. Piazzesi, 2005). We elaborate more on this alternative approach in Appendix B.2. It leads to virtually indistinguishable numerical results if the transition intensity is zero between any two FOMC meetings but very large at time nT, $n \in \mathbb{N}$.

Proof. See Appendix A.1. □

The numerical solution algorithm that we use to solve for the function $G(t,\pi,i)$ is discussed in Appendix B.

4 Analytical Asset Pricing Results

In this section, we determine several asset pricing implications of the model. We first identify the stochastic discount factor (SDF) from which the equilibrium real risk-free interest rate and the market prices of risk can be derived. Subsequently, we obtain the equilibrium characterizations of other economic quantities, such as the stock market's price-dividend ratio, the equity premium, and the term structure of interest rates.

4.1 The Stochastic Discount Factor and the Risk-Free Rate

Let $H = (H_t)_{t\geq 0}$ denote the real stochastic discount factor. Duffie and Epstein (1992a,b) show that the stochastic discount factor for SDU is given by

$$H_t = e^{\int_0^t f_J(c_s, J_s)\mathrm{d}s} f_c(c_t, J_t),$$

which, by Itô's Lemma, yields the dynamics

$$\frac{\mathrm{d}H_t}{H_{t-}} = \frac{\mathrm{d}f_c(c_{t-}, J_{t-})}{f_c(c_{t-}, J_{t-})} + f_J(c_{t-}, J_{t-})\mathrm{d}t.$$

The following theorem provides a characterization of the dynamics of the stochastic discount factor in our model.

Theorem 4.1. *The dynamics of the stochastic discount factor between two FOMC meetings are governed by*

$$\frac{\mathrm{d}H_t}{H_{t-}} = -r^f \mathrm{d}t - \Theta_W^\top \mathrm{d}W_t + \mathbb{E}\left[(1-\ell)^{-\gamma} - 1\right]\mathrm{d}N_t - \Theta_N \mathrm{d}t.$$

The equilibrium real risk-free interest rate r^f between two FOMC meetings is[15]

$$
\begin{aligned}
r^f = \underbrace{\delta}_{Discounting} + \underbrace{\frac{1}{\psi}\mu_c(\pi,i)}_{Smoothing} - \underbrace{\frac{1}{2}\gamma\Big(\frac{1}{\psi}+1\Big)\sigma_c^2}_{Standard\ Diffusion\ Risk} \underbrace{+\lambda_c\Big(\frac{\theta-1}{\theta}\mathbb{E}\big[(1-\ell)^{1-\gamma}\big]-\mathbb{E}\big[(1-\ell)^{-\gamma}\big]+\frac{1}{\theta}\Big)}_{Macroeconomic\ Disaster\ Risk} \\
+ \underbrace{\frac{\theta-1}{2\theta^2}\frac{G_\pi^2}{G^2}\sigma_\pi^2}_{Inflation\ Risk} \underbrace{+\frac{\theta-1}{\theta}\frac{G_\pi}{G}\sigma_\pi\sigma_c\rho_{c\pi}}_{Interaction\ Risk}.
\end{aligned}
\tag{4.1}
$$

The market prices of risk Θ_W and Θ_N are

$$
\Theta_W = \gamma\Sigma_c - \frac{\theta-1}{\theta}\frac{G_\pi}{G}\Sigma_\pi, \qquad \Theta_N = \lambda_c\mathbb{E}\big[(1-\ell)^{-\gamma}-1\big]. \tag{4.2}
$$

Proof. See Appendix A.2. □

Equation (4.1) offers a decomposition of the risk-free interest rate into its various components. The first two terms represent the subjective time preference rate and the desire of intertemporal consumption smoothing. The third term captures the formation of precautionary savings as insurance against diffusive shocks and reduces the risk-free rate. Those three terms also arise in a standard Lucas (1978) tree model. The fourth term represents precautionary savings in response to macroeconomic recurring disaster risk and as for diffusive risk reduces the risk-free rate. This effect is increasing non-linearly in the coefficient of risk aversion γ. In the second row, the first term reflects precautionary savings for inflation risk, while the second term captures the interaction between consumption and inflation risk. Note that in case of time-additive CRRA utility ($\theta = 1$) both of those terms vanish. We emphasize that compared to a standard Lucas (1978) tree model, the risk-free interest rate in (4.1) is not a continuous process but depends on the current federal funds rate i, which is adjusted in a discontinuous manner. This effect is captured in $\mu_c(\pi,i)$ but also hidden in the function G via the boundary condition (3.9). Consequently, when the FED adjusts the federal funds rate, the risk-free rate jumps accordingly.

Equation (4.2) provides a characterization of the market prices of risk in the economy. The market price of diffusive risk Θ_W comprises two terms. The (standard) first term reflects the compensation for aggregate diffusive consumption risk, whereas the second term represents the compensation for shocks to the inflation rate π and the federal funds rate i. Since the representative agent has a preference for early resolution of uncertainty, she not only cares about immediate consumption risk but also about the fact that her indirect utility depends on the state variables and is thus stochastic. In case of time-

[15]Following a standard no-arbitrage argument (Benninga and Protopapadakis, 1983), the corresponding equilibrium nominal risk-free interest rate can be calculated from the Fisher identity (Fisher, 1930).

additive CRRA utility the corresponding market prices of state variable risk would be zero. The market price of consumption jump risk Θ_N represents the compensation for the immediate impact of economic disasters on the consumption level.

4.2 Price-Dividend Ratio and Equity Premium

Let $P = (P_t)_{t\geq 0}$ denote the price of the claim to aggregate dividends following the dynamics in (3.3). Absence of arbitrage implies that P is the expected integral of future dividends discounted using the stochastic discount factor,

$$P(t,c,d,\pi,i) = \mathbb{E}_t\Big[\int_t^\infty \frac{H_s}{H_t} D_s \mathrm{d}s\Big].$$

Dividing by current dividends yields the price-dividend ratio,

$$\Omega(t,\pi,i) = \mathbb{E}_t\Big[\int_t^\infty \frac{H_s}{H_t}\frac{D_s}{D_t} \mathrm{d}s\Big],$$

which is independent of consumption and dividends. In Appendix A.4 we show that the price-dividend ratio solves the PDE specified in the following theorem.

Theorem 4.2. *The price-dividend ratio $\Omega(t,\pi,i)$ satisfies the following PDE between two FOMC meetings*

$$0 = 1 + \Omega_t + \Omega\big(\mu_{\widehat{D}} + \lambda_c \mathbb{E}\big[(1-\ell)^{\phi-\gamma} - 1\big]\big) + \Omega_\pi\Big(\mu_\pi(\pi,i) + \Sigma_\pi^\top \Sigma_{\widehat{D}}\Big) + \Omega_{\pi\pi}\frac{1}{2}\sigma_\pi^2 \tag{4.3}$$

subject to the boundary condition

$$\Omega(T,\pi,i) = \sum_{\Delta_i \in \mathscr{I}} \Omega(T,\pi,i+\Delta_i)\mathbb{P}(\Delta_i \mid \pi, i),$$

where the processes $\mu_{\widehat{D}}$ and $\Sigma_{\widehat{D}}$ are specified in Appendix A.3.

Proof. See Appendix A.4. □

To better understand the relation between the dividend yield $y_d = \frac{1}{\Omega}$ and the model parameters, we consider the following special case of our model: assuming a constant expected consumption and dividend growth rate in normal times, μ_c and μ_d, respectively, the state variables become irrelevant, allowing PDE (4.3) to be solved analytically. We obtain the following expression for the dividend yield

$$y_d = \delta + \frac{1}{\psi}\mu_c - \mu_d - \frac{1}{2}\gamma\Big(\frac{1}{\psi}+1\Big)\sigma_c^2 + \gamma\sigma_c\sigma_d\rho_{cd} + \lambda_c\Big(\frac{\theta-1}{\theta}\mathbb{E}\big[(1-\ell)^{1-\gamma}\big] - \mathbb{E}\big[(1-\ell)^{\phi-\gamma}\big] + \frac{1}{\theta}\Big). \tag{4.4}$$

Although determined in a very special case of our model with constant expected growth rates, we conjecture that the economic mechanism behind (4.4) carries over to our full model with state-dependent expected consumption and dividend growth rates $\mu_c(\pi, i)$ and $\mu_d(\pi, i)$, respectively.[16] In turn, fluctuations in expected growth rates, coming from time-varying inflation and the FED's interest rate adjustments, make the dividend yield state-dependent and may predict a systematic pattern. We confirm this mechanism numerically in our quantitative analysis in Section 6.

We now turn to the composition of the equity premium in this model. The equity premium arises from the exposure of the stock price to the various risk factors in the economy, i.e., diffusion risk, macroeconomic disaster risk, and monetary policy risk. Itô's Lemma implies that the stock price P_t satisfies

$$\frac{\mathrm{d}P_t}{P_{t-}} = \mu_p \mathrm{d}t + \Sigma_P^\top \mathrm{d}W_t + \left[(1-\ell)^\phi - 1\right]\mathrm{d}N_t - \lambda_c \mathbb{E}\left[(1-\ell)^\phi - 1\right]\mathrm{d}t$$

for processes μ_p and Σ_P specified in Appendix A.5. Here the term Σ_P represents normal times variation in dividends, whereas the term $[(1-\ell)^\phi - 1]$ models shocks to dividends in case of a disaster.

Let $r^s = \mu_p + y_d$ denote the instantaneous return on the stock. The instantaneous equity premium is therefore $r^s - r^f$.

Theorem 4.3. *The instantaneous equity premium relative to the risk-free interest rate is*

$$\begin{aligned} r^s - r^f = \underbrace{\gamma\sigma_c\sigma_d\rho_{cd}}_{\text{Diffusion Risk}} + \underbrace{\lambda_c \mathbb{E}\left[\left(1-(1-\ell)^{-\gamma}\right)\left((1-\ell)^\phi - 1\right)\right]}_{\text{Macroeconomic Disaster Risk}} - \underbrace{\frac{\theta-1}{\theta}\frac{G_\pi}{G}\frac{\Omega_\pi}{\Omega}\sigma_\pi^2}_{\text{Inflation Risk}} \\ - \underbrace{\frac{\theta-1}{\theta}\frac{G_\pi}{G}\Sigma_\pi^\top\Sigma_d + \gamma\frac{\Omega_\pi}{\Omega}\sigma_c\sigma_\pi\rho_{c\pi}}_{\text{Interaction Risk}}. \end{aligned} \tag{4.5}$$

Proof. See Appendix A.5. □

The equity premium consists of various components. The first two terms are analogous to expressions in Barro (2006): the first term denotes compensation for shocks to consumption and dividends due to the Brownian motions, while the second term follows from the risk of a macroeconomic disaster. The third term arises from the representative agent's objective to hedge against diffusive shocks to the inflation rate. The last term reflects the compensation for the interaction risk between inflation and consumption

[16] The case $\mu_d = \phi\mu_c$ is particularly interesting and will be the benchmark case in our model. Then, (4.4) simplifies to $y_d = \delta + \left(\frac{1}{\psi} - \phi\right)\mu_c - \frac{1}{2}\gamma\left(\frac{1}{\psi}+1\right)\sigma_c^2 + \gamma\sigma_c\sigma_d\rho_{cd} + \lambda_c\left(\frac{\theta-1}{\theta}\mathbb{E}\left[(1-\ell)^{1-\gamma}\right] - \mathbb{E}\left[(1-\ell)^{\phi-\gamma}\right] + \frac{1}{\theta}\right)$. Here, μ_c and μ_d correlate perfectly with each other, making the influence of the state variables more clearly visible. This structure of dividend growth rates is comparable with the specification in Bansal and Yaron (2004) and replicates the modeling of dividends as leveraged consumption as in Wachter (2013) in the special case when $\sigma_d = \phi\sigma_c$.

and dividends, respectively. With time-additive CRRA utility those two last terms would vanish. As for the risk-free rate, the equity premium is not a continuous process but depends on the federal funds rate i.

4.3 Term Structure of Interest Rates

Let $Z^\tau = Z^\tau(t,\pi,i)$ denote the price of a real default-free zero-coupon bond with notional one and maturity τ, hereafter called τ-bond. By absence of arbitrage, its time-t value is given by

$$Z^\tau(t,\pi,i) = \mathbb{E}_t\Big[\frac{H_\tau}{H_t}\Big],$$

and the real bond yield is

$$r_t^\tau = -\frac{1}{\tau - t}\ln\big(Z^\tau(t,\pi,i)\big).$$

Using standard arguments (see Appendix A.6), we can derive a partial differential equation for the price of the default-free zero-coupon bond.[17]

Theorem 4.4. *The price of any default-free zero-coupon bond with maturity τ satisfies the following PDE between two FOMC meetings*

$$r^f Z^\tau = Z_t^\tau + Z_\pi^\tau\big(\mu_\pi(\pi,i) + \Sigma_\pi^\top \Sigma_H\big) + Z_{\pi\pi}^\tau \frac{1}{2}\sigma_\pi^2 \tag{4.6}$$

subject to $Z^\tau(\tau,\pi,i) = 1$ and Σ_H being specified in Appendix A.3. A bond with maturity $\tau = k\cdot T$, $k \in \mathbb{N}$ satisfies the following boundary condition

$$Z^{kT}(T,\pi,i) = \sum_{\Delta_i \in \mathscr{I}} Z^{kT}(T,\pi,i+\Delta_i)\mathbb{P}(\Delta_i \mid \pi,i).$$

Proof. See Appendix A.6. □

An outline of our numerical solution algorithm to calculate the term structure of interest rates is given in Appendix B.3. The calculation of nominal bond prices and yields follows the approach in Piazzesi and Schneider (2007).

[17] Using similar arguments, we can easily extend this analysis to defaultable Treasury bonds as in Barro (2006), Wachter (2013), and Fleischer et al. (2024) or other fixed-income products that are exposed to credit risk, such as corporate bonds. In this case, the left-hand-side of (4.6) would be $\big(r^f + \lambda_c q_d \mathbb{E}[(1-\ell)^{-\gamma}\ell_d]\big)Z^\tau$, assuming the bond defaults with probability q_d if a disaster hits the economy and the loss given default equals ℓ_d. Note that the above mentioned papers assume that $\ell_d = \ell$. However, since we focus on US data rather than OECD data and the default probability of the US government is close to zero, we restrict our analysis to default-free bonds.

Parameter	Description	Value
	Consumption Dynamics	
$\widehat{\beta}_{0,c}$	Expected growth parameter	0.018
$\widehat{\beta}_{i,c}$	Expected growth parameter	0.366
$\widehat{\beta}_{i\times\pi,c}$	Expected growth parameter	−3.681
σ_c	Consumption volatility	0.0128
λ	Disaster intensity	0.088
α	Disaster size parameter	8
	Inflation Dynamics	
σ_π	Inflation volatility	0.0164
κ	Mean-reversion speed	0.171
$\overline{\pi}$	Mean-reversion level	0.0385
$\rho_{c\pi}$	Consumption/inflation correlation	0.0361
	Dividend Dynamics	
ϕ	Leverage factor	2.0
σ_d	Dividend volatility	0.0855
ρ_{cd}	Consumption/dividend correlation	0.2654
$\rho_{\pi d}$	Dividend/inflation correlation	−0.3307
	Preferences	
γ	Relative risk aversion	4.33
ψ	Elasticity of intertemporal substitution	0.32
δ	Subjective discount rate	0.0075

Table 2: Baseline Parameter Values. See the main text for the motivation of the assumed parameter values.

5 Calibration

In this section, we calibrate the asset pricing model from Section 3 such that it is in line with the stylized facts from Section 2. For this purpose, we estimate the modified VAR(1)-model (C.2), which discretizes our economy. In a second step, we set all insignificant parameters to zero and re-estimate (C.2). The technical details and the estimation results are given in Appendix C.

Consumption Growth To calibrate the consumption process (3.1), we choose the following drift rate as motivated in Appendix C.2

$$\mu_c(\pi, i) = \widehat{\beta}_{0,c} + \widehat{\beta}_{i,c}\, i + \widehat{\beta}_{i\times\pi,c}\, \pi \cdot i \tag{5.1}$$

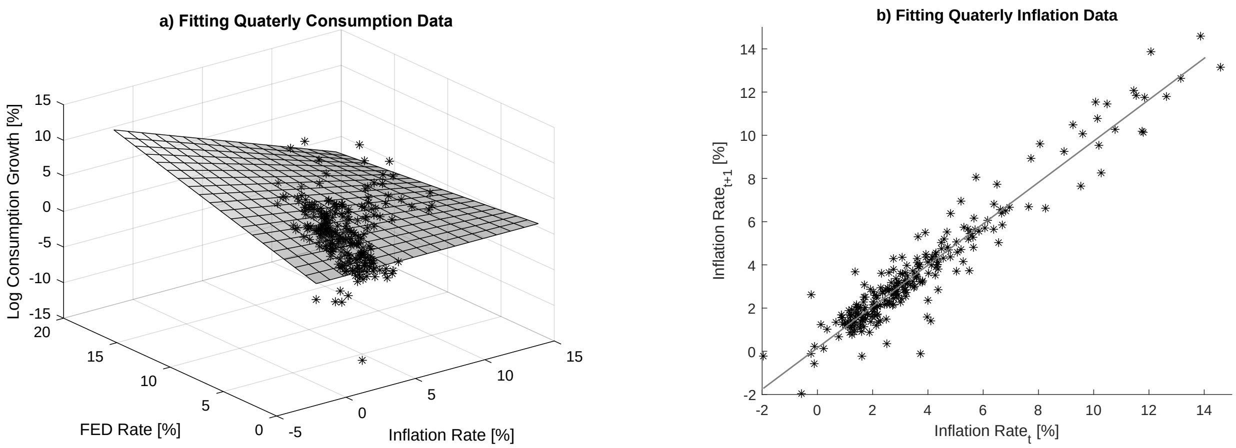

Figure 5: Fitting Consumption Growth and Inflation. The figure complements the regression results summarized in Panel (b) of Table C.2. Panel (a) depicts the results for real consumption growth rates and Panel (b) depicts the results for inflation. Our sample period is from January 1960 until December 2020.

with parameters $\widehat{\beta}_{0,c}, \widehat{\beta}_{i,c}, \widehat{\beta}_{i\times\pi,c}$ estimated from (C.2).[18] We emphasize that the model can generate a positive effect of interest rate shocks on economic growth if inflation is low, but also a negative effect if both inflation and the prevailing federal funds rate are already high. The latter effect is captured by the interaction term $\beta_{i\times\pi,c}\,\pi_t \cdot i_t$.

Panel (a) of Figure 5 illustrates the model fit to historical consumption growth data. It turns out that overall the fit is quite good and inflation and the federal funds rate seem to predict economic growth reasonably well. However, the model fails to capture macroeconomic disasters, such as the COVID-19 shock in March 2020, where consumption dropped by more than 12%. Such a huge drop in consumption cannot be explained by the relatively small volatility of $\sigma_c = 1.28\%$ as determined from the estimated cross-equation error variance–covariance matrix. To account for the risk of rare disasters, we follow Hambel et al. (2024) and choose a jump intensity of $\lambda = 8.8\%$ and a stochastic jump size with a power-distributed recovery rate $Z = 1 - \ell$ with parameter $\alpha = 8$.[19]

Such a model is in line with Barro and Jin (2021), who argue in favor of a combined long-run risk model with a disaster shock component as in our consumption dynamics. To rigorously test whether this calibration is in line with the historical consumption data, we simulate the model conditional on the observed values of π_t and i_t and compare the model outcome with the historical return series. A

[18]Using those results, we have in our sample period an expected consumption growth rate of $\mathbb{E}[\mu_c] = 2.65\%$. Consequently, the model replicates expected consumption growth in the data set pretty well, which is 2.66%, see Table 1.

[19]The power distribution for the recovery rate was originally introduced into the literature by Pindyck and Wang (2013).

Kolmogorov-Smirnov test cannot reject the hypothesis that the simulated data and the historical data have the same distribution with a p-value of 19%.

Inflation Dynamics To calibrate the inflation dynamics (3.2), we choose the following drift rate

$$\mu_\pi(\pi, i) = \kappa\big(\overline{\pi}(i) - \pi\big), \tag{5.2}$$

which captures the mean-reverting property of inflation, where the mean-reversion level $\overline{\pi}(i)$ may depend on the federal funds rate. Notice that in Panel (a) of Table C.2, the coefficient of the federal funds rate is statistically insignificant at the 5% level, suggesting that $\overline{\pi}$ is just a constant.[20] Discretizing (3.2) and using (5.2), we obtain $\kappa = \frac{1-\widehat{\beta}_{\pi,\pi}}{\Delta_t}$ and $\overline{\pi} = \frac{\widehat{\beta}_{0,\pi}}{1-\widehat{\beta}_{\pi,\pi}}$ with parameters $\widehat{\beta}_{0,\pi}$ and $\widehat{\beta}_{\pi,\pi}$ estimated from (C.2) and illustrated in Panel (b) of Figure 5. Since we have used quarterly data, we choose $\Delta_t = 0.25$. This yields a mean-reversion level of $\overline{\pi} = 3.85\%$ and a mean-reversion speed of $\kappa = 0.171$. From the estimated cross-equation error variance–covariance matrix, we determine the inflation volatility and the instantaneous correlation with consumption growth and obtain $\sigma_\pi = 1.64\%$ and $\rho_{c\pi} = 3.61\%$. We emphasize that, although the point estimate $\widehat{\beta}_{\pi,\pi} = 0.957$ seems to be close to one, one is outside the 95% confidence interval $[0.921,\ 0.993]$ of $\beta_{\pi,\pi}$, thus differs significantly from one.

Interest Rate Adjustments To model interest rate adjustments, Song (2017) and Campbell et al. (2020), among others, have built upon some variants of a Taylor-type rule, as introduced by Taylor (1993, 1999). However, in the past this rule has been widely criticized and retrospective comparisons with historical interest rate changes are not entirely convincing (Rudebusch, 2002; Svensson, 2003; Cochrane, 2011). Moreover, even if the central bank were fully committed to the Taylor rule, estimating components such as the output gap is notoriously difficult and the rule is also not designed to handle unconventional monetary policy.

Rather than following the normative approach by Taylor (1993, 1999), we specify several scenarios and model the FED's past actions in each scenario. In particular, we extract the historical probability distribution of interest rate adjustments for each specified scenario directly from the data and assume that

[20] As we have seen in Section 2.2, higher interest rate shocks have an immediate, positive short-term effect on inflation rates as they may signal stronger inflationary pressures or more persistent inflation expectations. However, on the long-run, higher interest rates help bring down inflation as they suppress demand and slow down price growth (e.g., Federal Reserve, 2021; Ferreira et al., 2024). Focusing on those effects, we analyze the influence of a state-dependent mean-reversion level in Section 7.

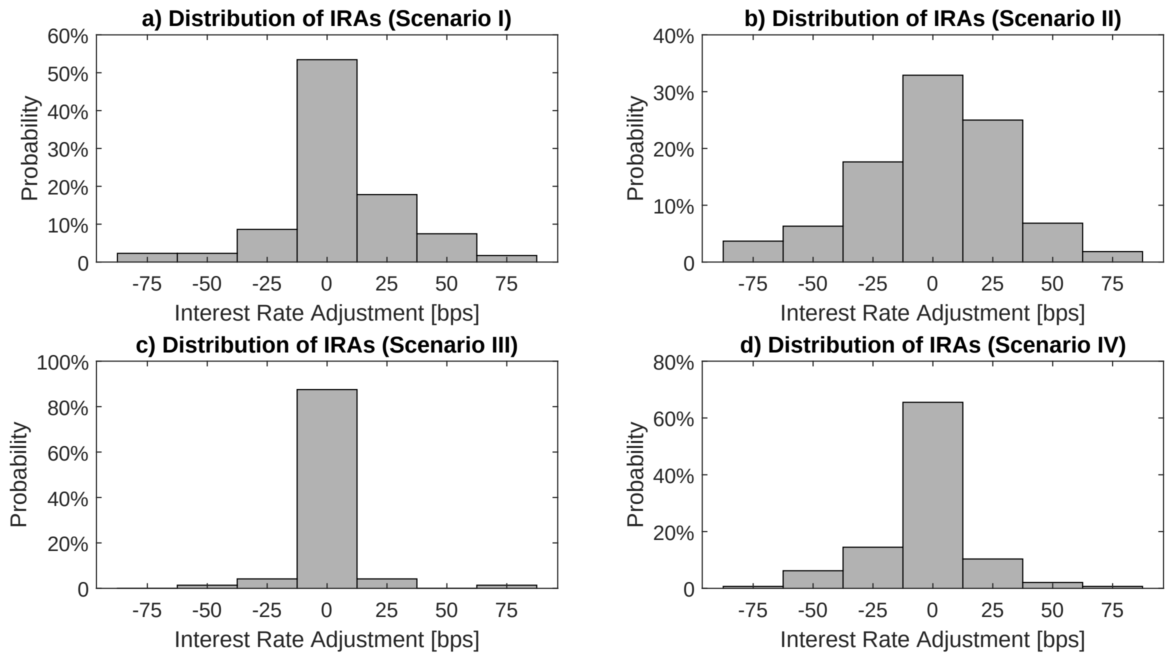

Figure 6: Historical Interest Rate Adjustments. Histograms show the probability distribution of interest rate adjustments (in basis points) for four different scenarios. In Scenario I, the inflation rate is above the 2% target and exceeds the federal funds rate. In Scenario II, the inflation rate is above the 2% target but is below the federal funds rate. Scenario III corresponds to the situation in which the inflation rate is below the 2% target and exceeds the federal funds rate, while in Scenario IV, the inflation rate is below the 2% target and below the federal funds rate. Table C.4 in Appendix E provides the concrete numbers of the probabilities for interest rate adjustments in the four scenarios underlying the shown histograms.

interest rate adjustments follow a multinomial distribution. More precisely, we consider the following four scenarios:[21]

I. The inflation rate is above the 2% target and exceeds the nominal interest rate, i.e., $\pi_t > 2\%$ and $\pi_t > i_t$.

II. The inflation rate is above the 2% target and below the nominal interest rate, i.e., $\pi_t > 2\%$ and $\pi_t \leq i_t$.

III. The inflation rate is below or equal to the 2% target and exceeds the nominal interest rate, i.e., $\pi_t \leq 2\%$ and $\pi_t > i_t$.

IV. The inflation rate is below or equal to the 2% target and below the nominal interest rate, i.e., $\pi_t \leq 2\%$ and $\pi_t \leq i_t$.

[21] In our numerical implementation, we exclude the possibility of negative federal funds rates, while other interest rates, such as the risk-free rate or long-term Treasury yields, can become negative both in nominal and real terms.

Figure 6 depicts four histograms showing the probability distribution of interest rate adjustments for each specific scenario. In Scenarios I and II, interest rate increases are more likely than a cut in interest rates. This is plausible as in those scenarios the inflation rate is above the 2% target and the FED aims to bring down inflation by raising interest rates. Further, we observe a rather high volatility in interest rate adjustments in Scenario II. This is because in that scenario the interest rate already exceeds the inflation rate, implying that the FED has already entered a cycle of raising interest rates. Thus, future interest rate adjustments may go in both directions, depending on whether the FED hangs on to its expansive monetary policy or returns to a more restrictive monetary policy. Scenario III is particularly interesting. Here the inflation rate is below the 2% target and interest rates are even lower. In this scenario, interest rate adjustments are rather unlikely. This corresponds to the period of the FED's zero-interest-rate policy and quantitative easing, beginning in the aftermath of the financial crisis in 2009. It is important to highlight that our model of interest rate adjustments can capture this unconventional monetary policy, whereas a Taylor-type rule fails to account for such a scenario. Finally, in Scenario IV, interest rate cuts are more likely as here the FED tries to push inflation to its 2% target by lowering interest rates.

Dividend Dynamics Our model builds on the standard model of leveraged consumption, $D = c^{\phi}$ (e.g., Wachter, 2013), but we adjust the dividend volatility term to match the observed volatility of aggregate dividends. First, we assume a leverage parameter of $\phi = 2.0$, which is well in line with the literature.[22] Second, instead of assuming $\sigma_d = \phi\sigma_c$, we estimate the volatility from historical data. We obtain $\sigma_d = 8.55\%$. Third, we estimate the instantaneous correlations with consumption and inflation risk to be $\rho_{cd} = 26.54\%$ and $\rho_{d\pi} = -33.07\%$.

Preference Parameters We calibrate the preference parameters to simultaneously explain the average risk-free interest rate, dividend yield, and equity premium in our model. First, we choose a degree of relative risk aversion of $\gamma = 4.33$ to match an average equity premium of 5.5% in our sample.[23] Moreover, we calibrate the elasticity of intertemporal substitution and the time preference rate such that the average model-implied risk-free rate and dividend yield are close to their historical averages of 0.78%

[22] For instance, Abel (1999) uses a leverage parameter of $\phi = 2.74$, Wachter (2013) assumes $\phi = 2.6$, Branger et al. (2016) use $\phi = 2$.

[23] This value is well in line with other long-run risk and disaster risk models. For instance Barro and Jin (2021) use a risk aversion of six in a combined long-run risk and disaster risk model. Other disaster risk models use γ in the range of three to six, too.

and 2.91%, respectively. This can be achieved by choosing $\psi = 0.32$ and $\delta = 0.75\%$.[24] Note that this calibration ensures preferences for early resolution of uncertainty since $\psi > \frac{1}{\gamma}$. We refer to Appendix C.3 for more calibration details.

6 Quantitative Analysis

We now turn to the quantitative analysis of our model, where we use the calibration given in Table 2. First, we investigate how our model fits aggregate consumption growth data. We then examine the quantitative asset pricing implications by first analyzing how the model-implied asset pricing moments depend on the state variables, followed by assessing the price impacts of monetary policy and the model's empirical performance.

6.1 Consumption Growth

Since we focus on consumption-based asset pricing, it is instructive to start our analysis by evaluating how our model can fit aggregate consumption growth data before turning to the asset pricing implications.

Figure 7 depicts the model-implied simulation results of *expected* economic growth (——) together with the historical growth data (——). Panel (a) shows the historical realized and simulated expected growth rate in normal times, $\mu_c(\pi, i)$ as calculated in (5.1). Although the prevailing inflation rate and federal funds rate can explain a substantial amount of the variation in the consumption growth data, those macroeconomic factors alone are not able to explain the striking spikes in 2009 and 2020, which is the reason why we also take disaster risk into account.[25]

Focusing on the relation between consumption growth and inflation, we see in Panel (b) that the comovement between those two quantities has varied substantially over time. In particular, we observe a strong negative comovement between 1965 and 2000, which, however, turns positive from the year 2000 onward. Specifically, coming from a correlation of −84% in the 1980's, correlations reached 55% in the 2010's. Our model is also able to capture the sign-switching behavior of the comovement between those two quantities. In particular, we can reproduce the negative correlation until the beginning of

[24]Concerning the EIS, there is mixed empirical evidence. Hansen and Singleton (1982), Attanasio and Weber (1989), and Vissing-Jorgensen and Attanasio (2003) estimate the EIS to be well above one. Hall (1988), Campbell (1999), and Vissing-Joergenen (2002) on the other hand, estimate its value to be well below one. Others use unit EIS for the sake of tractability Wachter, 2013. See also the survey article by Thimme (2017) and the references therein. Other studies often use an elasticity of $\psi = 1.5$, e.g., Bansal and Yaron (2004). Therefore, we perform a sensitivity analysis in Section 7.

[25]In Figure C.3 in Appendix E we show four different sample paths of simulated consumption growth rates.

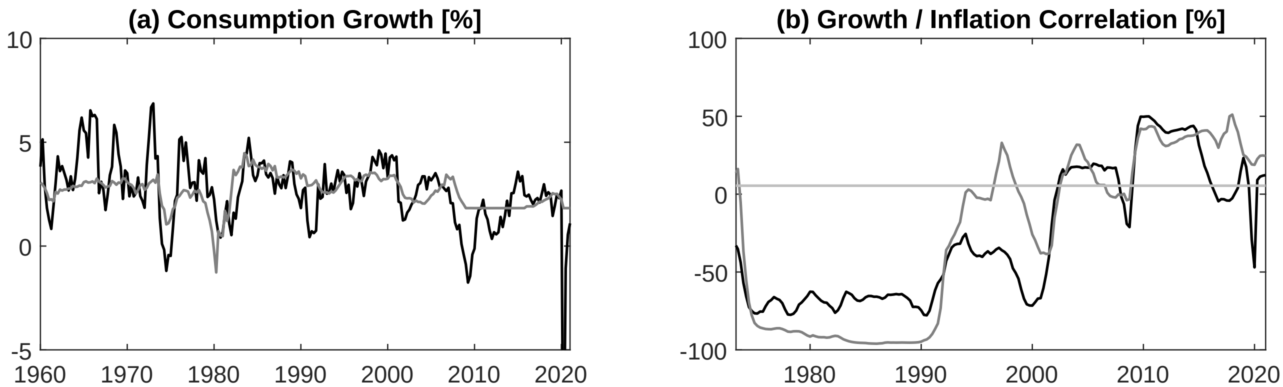

Figure 7: Data versus Model Results: Consumption Growth. The figure depicts the historical and simulated (a) annualized real consumption growth rate and (b) the time-varying correlation between consumption growth and inflation. Black lines (——) show historical time series, whereas gray lines (——) present model simulation results. We use quarterly data. All results are annualized.

the 2000's and the positive comovement thereafter. However, we have to admit that our model cannot match the negative trend in comovement starting in 2020 potentially triggered by the COVID-19 shock.

6.2 Asset Pricing Moments: State Dependency

Panel (a) of Figure 8 shows the dividend yield as a function of the inflation rate and the federal funds rate. The gray surface illustrates the functional form of the dividend yield using (4.3), and the black stars visualize historical data. It is shown that the dividend yield increases in the federal funds rate and to a lesser extend in the inflation rate. This pattern can also be seen from the historical data.

Similarly, Panel (b) visualizes the equity premium in our model as a function of those two macroeconomic factors using the functional form in (4.5). We observe that the equity premium increases in the prevailing inflation rate, while the effect of the federal funds rate is ambiguous: for low inflation rates, the effect of the federal funds rate on the equity premium is negative, while for higher inflation rates (e.g., 10%), the federal funds rate has almost no effect on the equity premium

Panel (c) depicts the nominal short-term risk-free rate together with the historical 3-month US Treasury bill rates. It turns out that the model explains the historical response of the short-term Treasury bill rate to inflation and the federal funds rate quite well. Both, historical rates and model-implied rates, increase in inflation and the federal funds rate, and our model captures the slope of this pattern.

Panel (d) shows the nominal 10y Treasury bond yield as computed from solving (4.6) alongside with the historical data.We see that the bond yield increases in both the inflation rate and the federal funds rate.

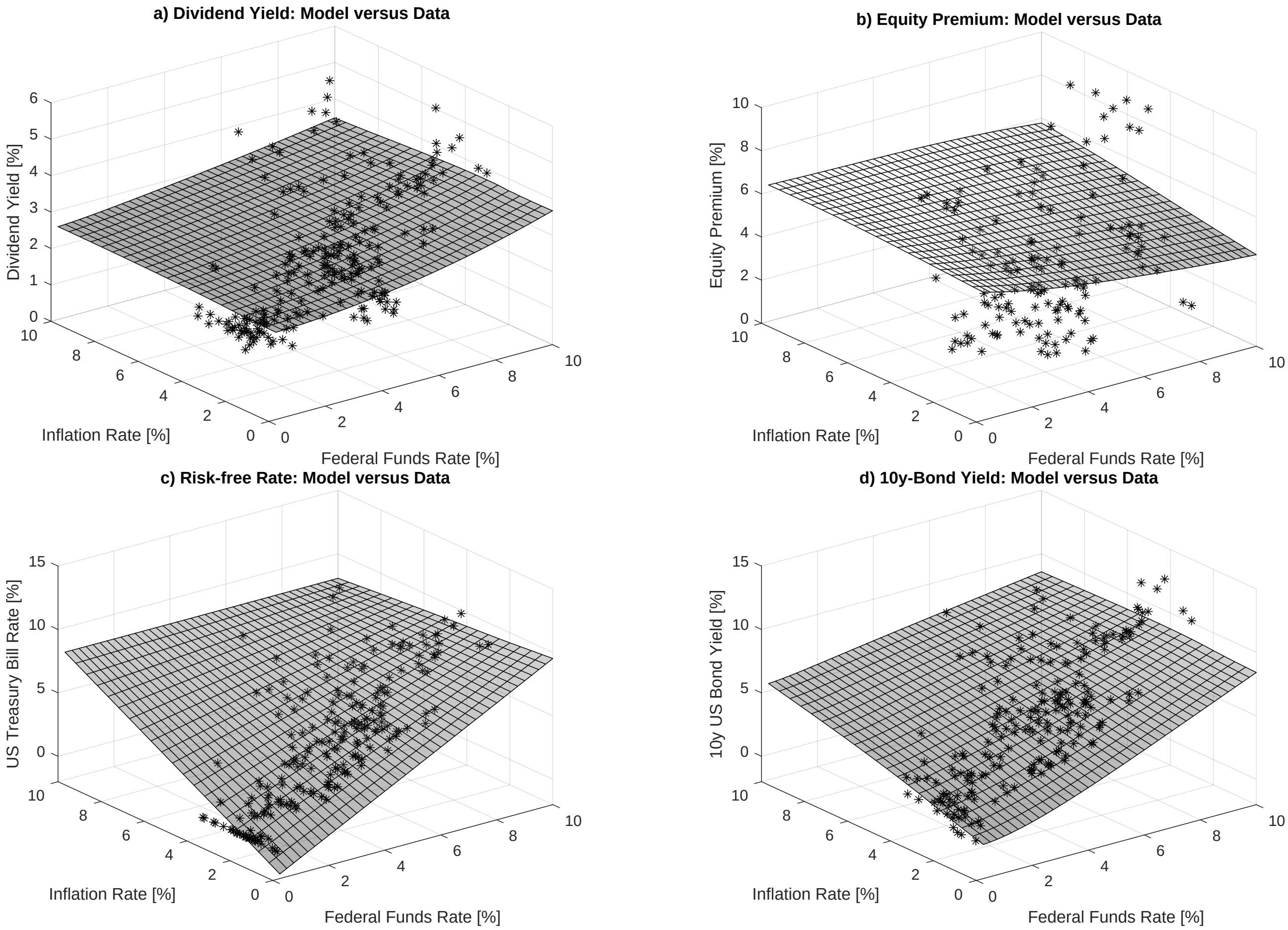

Figure 8: Key Asset Pricing Moments. The figure depicts the historical and model-implied (a) dividend yield, (b) equity premium, (c) risk-free rate, and (d) 10y Treasury bond yields conditional on the inflation rate and federal funds rate. The gray surfaces represent the model implied results, while black stars visualize historical data. All quantities are given in nominal terms. Our sample period is from January 1960 until December 2020 and we use quarterly data. All results are annualized.

Comparing our model results for the 10y bond yield with the short-term risk-free rate in Panel (c), we find that our model typically produces a normal term structure of interest rates. However, for very high inflation rates (e.g., 10%) in combination with a low federal funds rate, it can also produce an inverse term structure. This is well explained by the fact that investors expect inflation rates to stay high over the short-term and thus require a higher compensation for short-term than long-term bonds as inflation might return to normal levels in the long-run. However, for high inflation rates and a high federal funds rate, we obtain a normal term structure as in this case investors expect inflation rates to fall in the near future.

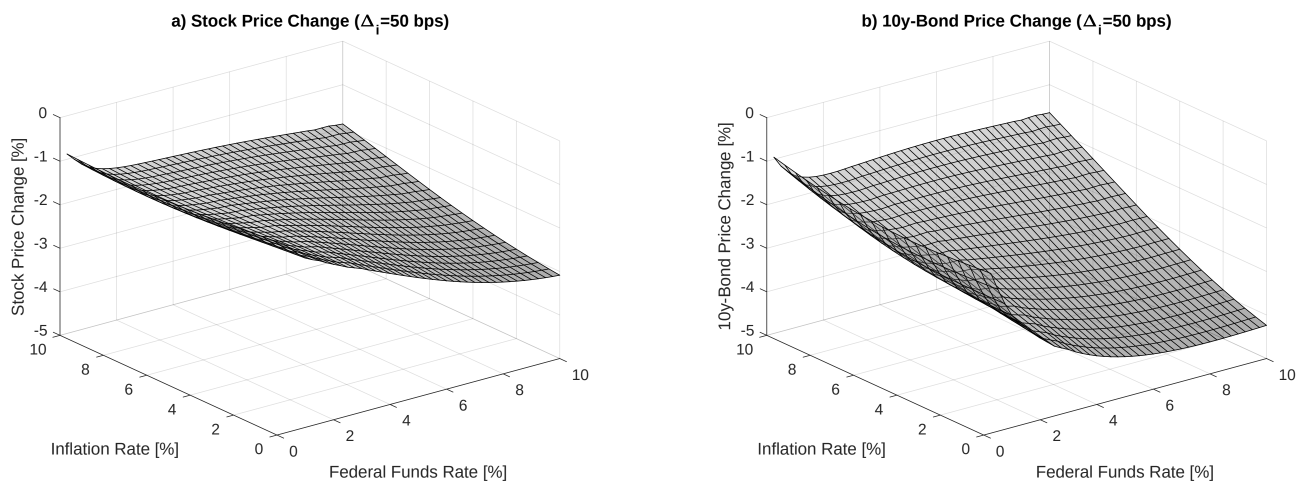

Figure 9: Price Impact of Monetary Policy. The figure depicts the model-implied (a) stock price change and (b) 10y bond price change in response to a change in the federal funds rate of Δ_i = 50bps conditional on the inflation and federal funds rate.

6.3 Price Impacts of Monetary Policy

We now turn to the analysis of the impact of monetary policy on stock and bond prices. As we assume that the federal funds rate process is discontinuous, which implies discontinuous asset pricing moments as well, our model can explain drastic changes in asset prices following a FED's interest rate announcement.[26] Figure 9 depicts the model-implied change in (a) the stock price and (b) the 10y bond price following an interest rate adjustment of Δ_i = 50bps conditional on the prevailing inflation and federal funds rate.[27] We find that the FED's interest rate adjustments can have a sizeable price impact on stock and bond prices. This is also in line with the empirical observations (e.g., Bomfim, 2003; Bernanke and Kuttner, 2005; Gürkaynak et al., 2005). In particular, a tightening monetary policy has a negative effect on stock and bond markets leading to strong losses. The opposite is true for a loosening monetary policy. This result is reasonable as an increase in interest rates makes it more expensive for companies to raise capital, which hurts future growth prospects as well as near-term earnings and thus explains a drop in the stock price. For bonds, an increase in the federal funds rate leads to an increase in its yield, resulting in a lower bond price. Besides, we find that bonds in general are more sensitive to interest rate adjustments than stocks. Here the effects are strongest for low inflation rates and high levels of the federal funds rate.

[26]Piazzesi (2005) also models the federal funds rate as discontinuous process. Precisely, she assumes that the federal funds rate follows a Markov chain, whose jump intensities increase drastically around FOMC meetings.

[27]Figure E.5 in Appendix E provides a similar analysis considering an interest rate adjustment of Δ_i = 25bps and Δ_i = 75bps, respectively.

(a) Stock Market

Moment	Model	US Data
$\mathbb{E}[y_d]$	2.91	2.91
$\sigma(y_d)$	0.45	1.13
$\mathrm{AC}(y_d)$	89.53	97.58
$\mathbb{E}[r_s - r_f]$	5.49	5.52
$\sigma(r_s)$	11.43	16.03
SR	48.00	34.42

(b) Bond Market

Moment	Model	US Data
$\mathbb{E}[r_f]$	4.46	4.45
$\sigma(r_f)$	2.97	3.17
$\mathrm{AC}(r_f)$	95.76	96.62
$\mathbb{E}[r^{10}]$	4.90	6.01
$\sigma(r^{10})$	2.55	2.99
$\mathrm{AC}(r^{10})$	97.11	97.05

(c) Data/Model Correlations

Correlation	Value
$\mathrm{corr}(g^d, \mu_c^m)$	47.01
$\mathrm{corr}(y_d^d, y_d^m)$	72.77
$\mathrm{corr}(r_f^d, r_f^m)$	87.52
$\mathrm{corr}(r^{5,d}, r^{5,m})$	91.50
$\mathrm{corr}(r^{10,d}, r^{10,m})$	90.62
$\mathrm{corr}(r^{30,d}, r^{30,m})$	90.51

Table 3: Data versus Model Asset Pricing Moments. Panel (a) and (b) report simulated and historical unconditional moments of key asset pricing quantities. AC represents the lag-1 autocorrelation, which measures the correlation between data points separated by one quarter. Panel (c) shows the correlation between the simulated time series of the respective asset pricing quantities and the historical time series. The model is simulated at a quarterly frequency and simulated data are aggregated to an annual frequency. Data moments are calculated using overlapping annual observations constructed from quarterly US data from January 1960 until December 2020. All data is given in nominal terms. y_d denotes the dividend yield, r_s is the return on stocks, r_f is the risk-free rate and r^5, r^{10}, and r^{30} denote the yield on 5y, 10y, and 30y Treasury bonds. All numbers are reported in percentage terms.

6.4 Asset Pricing Moments: Empirical Performance

Now we focus on the ex-post performance of our model. Figure 10 depicts the model-implied simulation results (——) together with the historical data (——). As those results are simulated based on the prevailing interest and inflation rate, we refer to them as conditional moments. A numerical decomposition of the conditional risk-free rate and the equity premium into its several components, as shown in Theorems 4.1 and 4.3, can be found in Appendix D. In Table 3 we further present the unconditional moments of the key asset pricing quantities.

Panel (a) of Figure 10 depicts the conditional model-implied and historical dividend yield series. We observe a sharp increase in the dividend yield if inflation is high as seen from 1975 to 1985. Our model can replicate this pattern quite well and we can explain this relation through (4.4). Even though this equation was derived in a model with no state variables, it explains that and how the expected consumption growth rate μ_c increases the dividend yield as its coefficient $\frac{1}{\psi} - \phi$ is positive due to the low level of the EIS. A high nominal federal funds rate contributes positively to consumption growth and in turn to the dividend yield. This explains the spikes in the previously mentioned period. However, from the year 2000 onward, our simulated dividend yield is rather flat and can only partly explain the

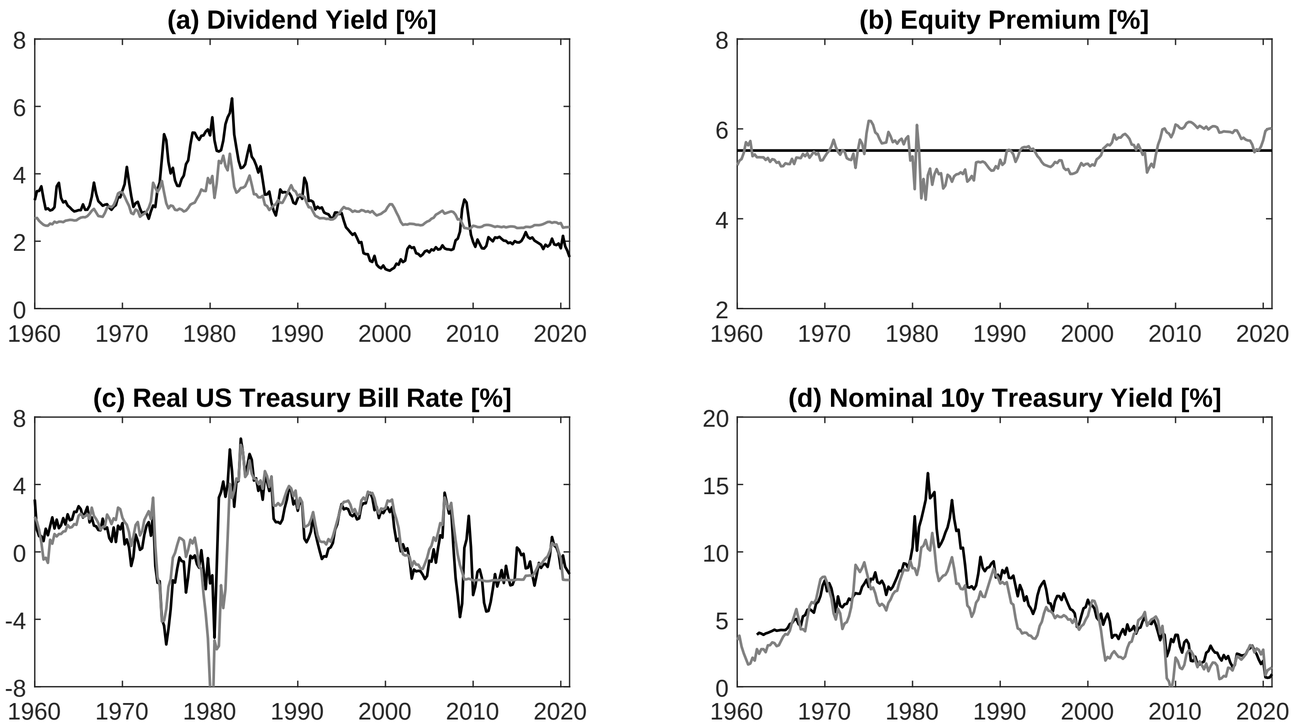

Figure 10: Data versus Model Results. The figure depicts the historical and simulated (a) S&P500 dividend yield series, (b) equity premium, (c) real US Treasury bill rate, and (d) nominal 10y Treasury yield. Black lines (——) show historical time series, whereas gray lines (——) present model simulation results. We use quarterly data. All results are annualized. Note that data on 10y Treasury bond yields is available from 1962 onward only.

trends in the data.[28] The correlation between the two series is about 72%, indicating a high predictive power of inflation and the federal funds rate for the dividend yield, see Table 3. Overall, the average dividend yield in the sample is 2.91% with a standard deviation of 1.13%, while the model implies an average dividend yield of 2.91% with a standard deviation of 0.45% in the sample period. Thus, our model underestimates the volatility of the dividend yield. The autocorrelation of the dividend yield implied in our model is 89.53% and thus slightly lower than the historical value of 97.58%.

[28]One possible explanation for this finding is the occurrence of a structural break in the relationship between the dividend yield and the inflation rate. To formally test this, we apply a supremum Wald test, which allows for the detection of unknown breakpoints in time series data. The test results indicate that we can reject the null hypothesis of no structural break at the 1% significance level, with the estimated break date occurring in the second quarter of 1995. More specifically, we observe a striking shift in the correlation between the dividend yield and the inflation rate around this breakpoint. Prior to 1995, the correlation between the dividend yield and the inflation rate is notably strong and positive at 71%. However, following the structural break, the relationship reverses, with the correlation turning significantly negative at –38%. It is well-known that latent regime shifts, as in Song (2017) or Dergunov et al. (2022), can be used to capture such time-varying correlations. We illustrate this mechanism in Appendix E.3.

Panel (b) depicts the time series of the conditional average equity premium implied by our model and compares it to the average equity premium in the data. The spikes around the 1980's are striking and mimic the spikes in inflation and the federal funds rate observed in the data. Thus, our state-depended equity premium reacts positively to higher inflation. This is captured by the last two terms in (4.5). While the equity premium was on average 5.52% in the sample period, our model predicts a slightly lower average equity premium of 5.49%. Considering the unconditional volatility of stock returns, our model produces a volatility that is too low compared to the volatility we observe in the data, i.e. 11.43% in the model vs. 16.03% in the data.[29] Thus, the Sharpe Ratio generated by our model (48.00%) exceeds the historical value (34.42%).

Panel (c) depicts the historical and model-implied *real* US Treasury bill rate. Here we can observe that our model is able to match the historical series very well. In particular, the correlation between the two series is about 87% and they share all major upward and downward trends. The strong comovement of the model-implied risk-free rate and the historical risk-free rate can be seen from (4.1), since a large portion of the fluctuations in the risk-free rate comes from inflation and the federal funds rate via the expected consumption growth rate μ_c. The average risk-free rate is 0.78% in the sample, and our model generates an average risk-free rate of 0.79%, matching this figure quite well. Focusing on the unconditional moments of the *nominal* rates given in Table 3, we find that our model produces an average nominal risk-free rate of 4.46% with a standard deviation of 2.97% compared to the historical risk-free rate of 4.45% with a standard deviation of 3.17%. Our model is also able to match the autocorrelation of the risk-free rate series, i.e., 95.76% in the model vs. 96.62% in the data. Considering the correlation between our model-implied and historical time series of nominal risk-free rates, we find that it is about 88% and thus even higher than for real risk-free rates.

Panel (d) depicts the historical and model-implied nominal yields of 10y Treasury bonds. We find that our model can replicate all major upward and downward trends of the historical data. The correlation between the two series is about 91%. Focusing on the unconditional moments, we find that our model slightly underestimates the average yield, i.e., our model generates an average yield of 4.90% compared to 6.01% in the data.[30] However, it can generate a similar volatility (2.55% vs. 2.99%) and autocorrelation (97.11 vs. 97.05) as in the data.

[29] We could generate a higher stock return volatility by incorporating time-varying disaster risk (Wachter, 2013), stochastic volatility (Bansal and Yaron, 2004) or an additional idiosyncratic disaster shock. However to focus on the novel implications of our model, we abstract from those extensions.

[30] Note that we abstract from credit risk and liquidity risk in our model. Incorporating modest credit risk in the spirit of Barro (2009); Wachter (2013); Fleischer et al. (2024) could generate a higher average bond yield.

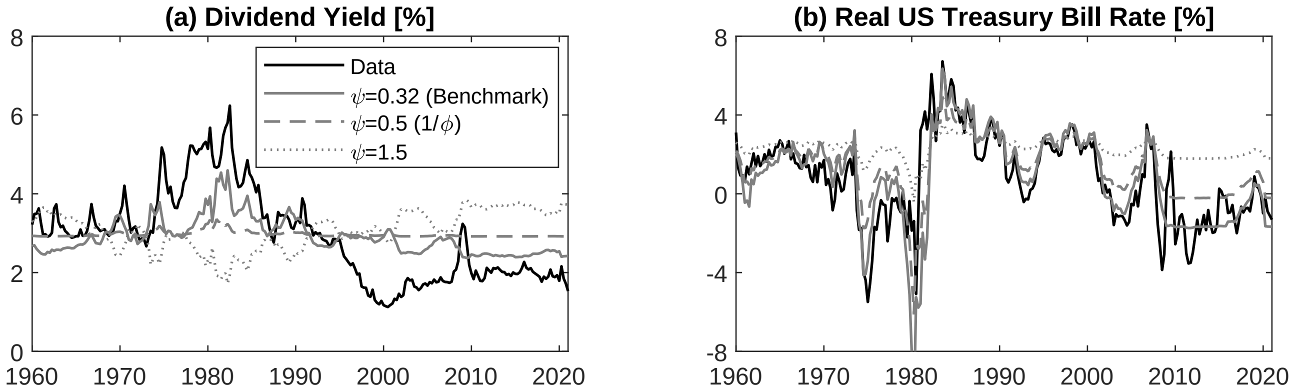

Figure 11: Sensitivity Analysis: Elasticity of Intertemporal Substitution. The figure depicts the historical and simulated (a) S&P500 dividend yield series and (b) real US Treasury bill rate for different values of the elasticity of intertemporal substitution. Black lines (——) show historical time series, while gray lines (——) present model simulation results using our benchmark calibration given in Table 2. Gray dashed lines (- - -) represent simulation results using an elasticity of 0.5, which corresponds to the case $\psi = 1/\phi$, whereas gray dotted lines (⋯⋯) show simulation results for an elasticity of 1.5. For the latter two cases we adjust δ accordingly to match the unconditional mean of the S&P500 dividend yield series. Other parameters are chosen as given in Table 2. We use quarterly data. All results are annualized.

7 Sensitivity Analysis

This section investigates how sensitive the empirical performance of our model is to key parameters, namely the elasticity of intertemporal substitution ψ together with the time preference rate δ, the relative risk aversion γ, and the mean-reversion level of the inflation process $\overline{\pi}(i)$. Unless otherwise stated, we change one of these parameters at a time and keep all other parameter fixed at the baseline levels given in Table 2.

7.1 Preference Parameters

Elasticity of Intertemporal Substitution Figure 11 depicts the historical and simulated dividend yield series and risk-free rate for different levels of the elasticity of intertemporal substitution. We change the time preference parameter δ accordingly such that we match the unconditional mean of the dividend yield for the different levels of the elasticity. We find that, although we can match the unconditional mean of the dividend yield, we are not able to reproduce the cyclical pattern of the dividend yield using an elasticity other than our benchmark case. Considering the case $\psi = 1.5$ (e.g., Bansal and Yaron, 2004), we find that here fluctuations of the dividend yield go in the opposite direction (⋯⋯). This can be well explained by (4.4) and the discussion in Footnote 16 as fluctuations in the dividend yield

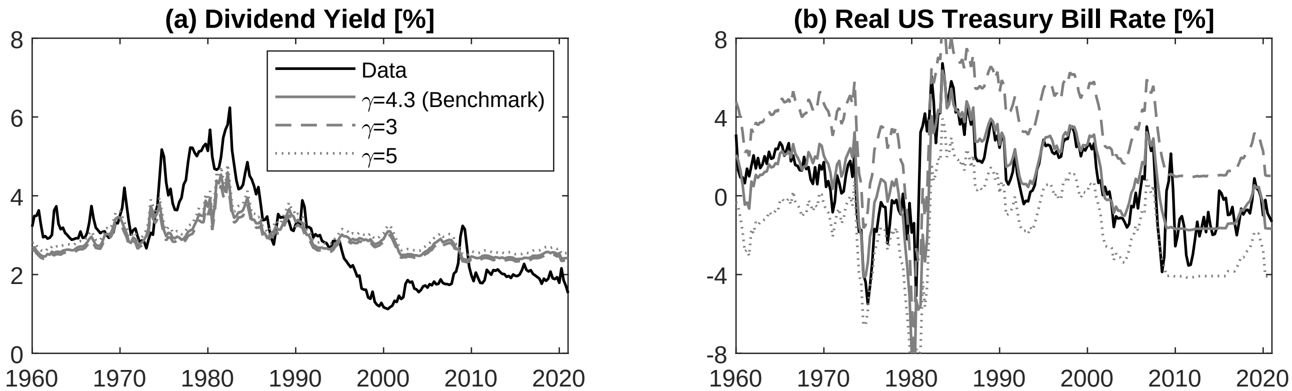

Figure 12: Sensitivity Analysis: Relative Risk Aversion. The figure depicts the historical and simulated (a) S&P500 dividend yield series and (b) real US Treasury bill rate for different values of the coefficient of relative risk aversion. Black lines (——) show historical time series, while gray lines (——) present model simulation results using our benchmark calibration given in Table 2. Gray dashed lines (- - -) represent simulation results using a relative risk aversion of 3, whereas gray dotted lines (······) show simulation results for a relative risk aversion of 5. Other parameters are chosen as given in Table 2. We use quarterly data. All results are annualized.

roughly mirror fluctuations in inflation and the federal funds rate. This effect on the dividend yield is captured by $(\frac{1}{\psi}-\phi)\mu_c(\pi,i)$. Notice that if $\psi=1.5$, the coefficient $\frac{1}{\psi}-\phi$ becomes negative turning around the pattern. The case $\psi=\frac{1}{\phi}=0.5$ produces a flat dividend yield series as in this case the $\mu_c(\pi,i)$ term vanishes and thus there is no state-dependency from the consumption smoothing motive anymore (- - -). Focusing on the effects on the risk-free rate, we find that our model still matches the general pattern of the risk-free rate in the data, but cannot reproduce the overall level of the risk-free rate. This is because we needed to increase δ to match the unconditional mean of the dividend yield.

In Figure E.6 in Appendix E, we provide a similar analysis, where we adjust δ to match the unconditional mean of the risk-free rate. However, in this case we cannot match the average dividend yield anymore. This analysis confirms the finding that varying the elasticity of intertemporal substitution too much destroys the model's ability to mach the historical pattern of the dividend yield.

Relative Risk Aversion Figure 12 illustrates the impact of changing the relative risk aversion away from its baseline value of $\gamma=4.33$. We find that risk aversion has a minor effect on the dividend yield series. However, although not shown in the figure, the relative risk aversion has a sizable impact on the magnitude of the equity premium. As can be seen from (4.5), the equity premium depends on risk aversion in a non-linear manner, implying that even small changes of risk aversion can have a significant impact on the equity premium. In order to bring the equity premium to its historical average, one would have to change the leverage parameter ϕ accordingly, which then would have a direct impact

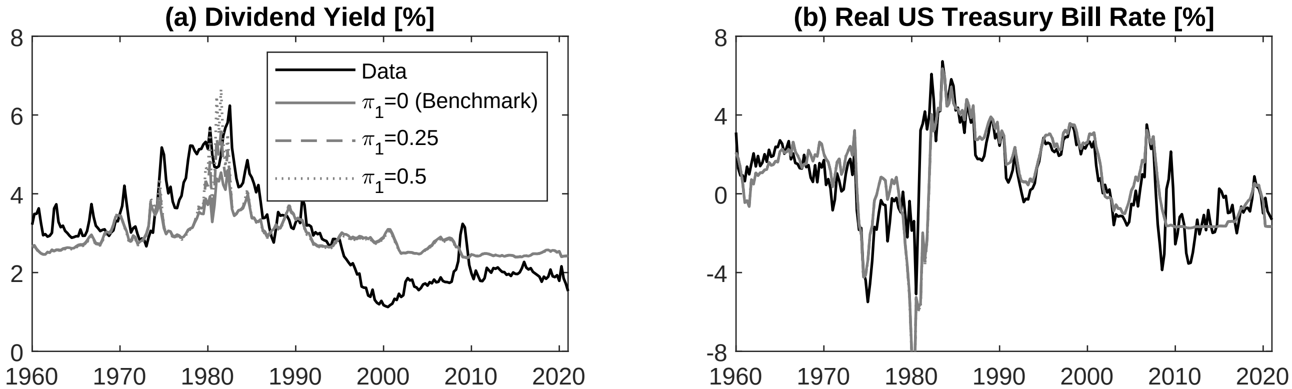

Figure 13: Sensitivity Analysis: Mean-Reversion Parameters. The figure depicts the historical and simulated (a) S&P500 dividend yield series and (b) real US Treasury bill rate for different values of the mean-reversion parameter π_1. Black lines (——) show historical time series, while gray lines (——) present model simulation results using our benchmark calibration given in Table 2. Gray dashed lines (- - -) represent simulation results using $\pi_1 = 0.25$, whereas gray dotted lines (······) show simulation results for $\pi_1 = 0.5$. Other parameters are chosen as given in Table 2. We use quarterly data. All results are annualized.

on the dividend yield series. Focusing on Panel (b), we find that the coefficient of relative risk aversion strongly affects the risk-free rate. As seen from (4.1), a higher coefficient of relative risk aversion leads to higher precautionary savings, pulling down the risk-free rate.

7.2 Inflation Dynamics

In our benchmark calibration, we assume that the mean-reversion level of inflation $\overline{\pi}$ is just a constant, motivated by the estimation results of the various VAR-models reported in Appendix C. However, as pointed out by Federal Reserve (2021) and in line with our empirical analysis from Section 2, there is a negative long-term effect of the federal funds rate on inflation rates. To capture this effect in our model, we now assume that the mean-reversion level negatively depends on the federal funds rate, i.e., $\overline{\pi}(i) = \pi_0 - \pi_1 i$ for non-negative constants π_0 and π_1. The latter parameter determines the sensitivity of the mean-reversion level of the inflation rate on the federal funds rate.

Figure 13 depicts the effect of changing the mean-reversion parameters π_0 and π_1. To still match the overall mean-reversion level of the inflation rate, we adjust the parameter π_0 accordingly to match the average level of inflation in our sample. As can be seen from the figure, the concrete choice of this parameter pair has a minor effect on the dividend yield series and the risk-free rate. Considering the dividend yield, we see that a positive value of π_1 produces a slightly higher dividend yield volatility with higher spikes around the 1980s, where inflation and federal funds rates were very high. Such a calibration aligns better with the data and captures the inflation-dampening effects of monetary policy. It also produces a slightly higher equity premium. For instance if $\pi_1 = 0.5$ ($\cdots\cdots$), the average model-implied equity premium is 5.64% and the volatility of the dividend yield increases from 0.45% to 0.65%. The effect of π_1 on the risk-free rate is negligible as the simulated risk-free rate series are almost indistinguishable.

8 Conclusion

This paper studies a parsimonious consumption-based asset pricing model with inflation and the central bank's interest rate adjustments as observable risk factors. It offers a novel mechanism to explain many features of the aggregate stock and bond market. In particular, it relies on observable macroeconomic factors only, which distinguishes it from the models using latent risk factors, such as long-run risk and stochastic volatility (e.g., Bansal and Yaron, 2004), time-varying disaster intensities (e.g., Wachter, 2013), or latent regime switches (e.g., Song, 2017).

A quantitative analysis shows that inflation and the federal funds rate both have explanatory power for future consumption growth and asset prices. We find that our model matches the key asset pricing moments conditional on the prevailing inflation and federal funds rate. In particular, we show that inflation increases the equity premium and higher interest rates lower the equity premium as they are intended to counter inflation. Besides, we can replicate the observed patterns and time variation in the price-dividend ratio and the risk-free rate. Our model is also able to explain unconditional moments of key asset pricing quantities, e.g., it produces an unconditional equity premium of 5.49% and a risk-free rate of 0.79% under reasonable values for the preference parameters. We have to acknowledge that our model underestimates the volatility of stock returns. However, this could be resolved by incorporating time-varying disaster risk or stochastic volatility. Besides, the model can explain drastic changes in asset prices following a FED's interest rate announcement, as we assume that the interest rate process is discontinuous, leading to discontinuous asset pricing moments, too. Furthermore, the calibrated model generates empirically reasonable term structures of government bond yields but underestimates the average yield on government bonds. Here a better fit could be achieved by incorporating some kind of credit risk.

A number of interesting extensions are possible for future research. First, our model could be studied with a different data set, e.g., Eurozone data or data on emerging markets. Especially in the latter case, we observe very high inflation rates in some countries and it would be interesting to see how our model performs in those scenarios. Besides, we have substantially higher default risk in those markets as compared to the US, making it interesting to study the model implications on defaultable bonds, see Footnote 17. Second, one could make the FED's interest rate adjustments endogenous and explicitly model the rational behind its interest rate decisions. Here it would be interesting to study the asset pricing implications and welfare effects if the investor would be able to perfectly foresee future interest rate adjustments. This would allow to draw novel conclusion about the welfare effects of central bank communication in the spirit of Blinder et al. (2008) and Candian (2021), among others.

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A Analytical Results

This appendix provides proofs for the analytical results in Sections 3 and 4.

A.1 Value Function

We conjecture that the value function in (3.6) has the form

$$J(t,c,\pi,i) = \frac{1}{1-\gamma}c^{1-\gamma}G(t,\pi,i).$$

Noting that

$$f(c,J) = \delta\theta\frac{1}{1-\gamma}c^{1-\gamma}G\big(G^{-1/\theta}-1\big)$$

and substituting the relevant partial derivatives into (3.6) yields

$$\begin{aligned}0 &= \delta\theta\frac{1}{1-\gamma}c^{1-\gamma}G\big(G^{-1/\theta}-1\big) + \frac{1}{1-\gamma}c^{1-\gamma}G_t + c^{1-\gamma}G\mu_c(\pi,i) + \frac{1}{1-\gamma}c^{1-\gamma}G_\pi\mu_\pi(\pi,i)\\ &- \frac{1}{2}c^{1-\gamma}G\gamma\sigma_c^2 + \frac{1}{2}\frac{1}{1-\gamma}c^{1-\gamma}G_{\pi\pi}\sigma_\pi^2 + c^{1-\gamma}G_\pi\sigma_c\sigma_\pi\rho_{c\pi}\\ &+ \lambda_c\mathbb{E}\Big[\frac{1}{1-\gamma}\big(c(1-\ell)\big)^{1-\gamma}G - \frac{1}{1-\gamma}c^{1-\gamma}G\Big].\end{aligned}$$

Dividing by $\frac{1}{1-\gamma}c^{1-\gamma}$ and collecting terms, we end up with

$$\begin{aligned}0 = \delta\theta G^{1-1/\theta} + G_t + G\Big[(1-\gamma)\mu_c(\pi,i) - \gamma(1-\gamma)\frac{1}{2}\sigma_c^2 - \delta\theta + \lambda_c\mathbb{E}\big[(1-\ell)^{1-\gamma}-1\big]\Big]\\ + G_\pi\big[\mu_\pi(\pi,i) + (1-\gamma)\sigma_c\sigma_\pi\rho_{c\pi}\big] + G_{\pi\pi}\frac{1}{2}\sigma_\pi^2.\end{aligned}$$

A.2 Dynamics of the Stochastic Discount Factor

Following Duffie and Epstein (1992a), the dynamics of the stochastic discount factor H are given by

$$\frac{\mathrm{d}H}{H_-} = \frac{\mathrm{d}f_c(c,J)}{f_c(c,J)} + f_J(c,J)\mathrm{d}t. \tag{A.1}$$

The relevant partial derivatives of the aggregator are

$$f_c(c,J) = \delta\big((1-\gamma)J\big)^{1-1/\theta}c^{-\phi}, \qquad f_J(c,J) = \delta(\theta-1)\big((1-\gamma)J\big)^{-1/\theta}c^{1-\phi} - \delta\theta.$$

Given our conjecture (3.7) of the optimal utility we get

$$f_c(c,J) = \delta c^{-\gamma} G^{1-1/\theta}, \qquad f_J(c,J) = \delta\big((\theta-1)G^{-1/\theta} - \theta\big).$$

To calculate the dynamics of the stochastic discount factor, we first compute

$$\frac{\mathrm{d}c^{-\gamma}}{c_-^{-\gamma}} = -\gamma\big(\mu_c(\pi,i)\mathrm{d}t + \Sigma_c^\top \mathrm{d}W\big) + \frac{1}{2}\gamma(\gamma+1)\sigma_c^2 \mathrm{d}t + \mathbb{E}\big[(1-\ell)^{-\gamma} - 1\big]\mathrm{d}N.$$

Next, we determine the dynamics of $G^{1-1/\theta}$. From Itô's Lemma it follows that $G = G(t,\pi,i)$ satisfies the following SDE

$$\frac{\mathrm{d}G}{G} = \mu_G \mathrm{d}t + \Sigma_G^\top \mathrm{d}W,$$

where its drift rate and volatility vector are

$$\mu_G = \frac{1}{G}\Big(G_t + G_\pi \mu_\pi(\pi,i) + \frac{1}{2}G_{\pi\pi}\sigma_\pi^2\Big), \qquad \Sigma_G = \frac{G_\pi}{G}\Sigma_\pi.$$

Another application of Itô's Lemma yields

$$\frac{\mathrm{d}G^{1-1/\theta}}{G_-^{1-1/\theta}} = \Big(\frac{\theta-1}{\theta}\mu_G - \frac{\theta-1}{2\theta^2}\|\Sigma_G^2\|\Big)\mathrm{d}t + \frac{\theta-1}{\theta}\Sigma_G^\top \mathrm{d}W.$$

Using Itô's product rule, we get for the dynamics of $\mathrm{d}f_c$

$$\begin{aligned}\frac{\mathrm{d}f_c}{f_c} &= \Big(-\gamma\mu_c(\pi,i) + \frac{1}{2}\gamma(\gamma+1)\|\Sigma_c^2\| + \frac{\theta-1}{\theta}\mu_G - \frac{\theta-1}{2\theta^2}\|\Sigma_G^2\| - \gamma\frac{\theta-1}{\theta}\Sigma_c^\top\Sigma_G\Big)\mathrm{d}t \\ &\quad + \Big(-\gamma\Sigma_c + \frac{\theta-1}{\theta}\Sigma_G\Big)^\top \mathrm{d}W + \mathbb{E}\big[(1-\ell)^{-\gamma} - 1\big]\mathrm{d}N. \end{aligned} \tag{A.2}$$

Given that G satisfies (3.8), it can be shown that

$$\begin{aligned}-f_J &= \delta + \frac{\theta-1}{\theta}\mu_G + \Big(\frac{1}{\psi} - \gamma\Big)\mu_c(\pi,i) - \frac{1}{2}\frac{\theta-1}{\theta}\gamma(1-\gamma)\|\Sigma_c^2\| + \frac{\theta-1}{\theta}(1-\gamma)\Sigma_c^\top\Sigma_G \\ &\quad + \frac{\theta-1}{\theta}\lambda_c\mathbb{E}\big[(1-\ell)^{1-\gamma} - 1\big]. \end{aligned} \tag{A.3}$$

Combining (A.1), (A.2) and (A.3), we get for the dynamics of the stochastic discount factor H

$$\frac{\mathrm{d}H}{H_-} = -r^f \mathrm{d}t + \Big(-\gamma\Sigma_c + \frac{\theta-1}{\theta}\Sigma_G\Big)^\top \mathrm{d}W + \mathbb{E}\big[(1-\ell)^{-\gamma} - 1\big]\mathrm{d}N - \lambda_c\mathbb{E}\big[(1-\ell)^{-\gamma} - 1\big]\mathrm{d}t, \tag{A.4}$$

where the risk-free rate is given by

$$r^f = \delta + \frac{1}{\psi}\mu_c(\pi,i) - \frac{1}{2}\gamma\Big(\frac{1}{\psi}+1\Big)\|\Sigma_c^2\| + \frac{\theta-1}{2\theta^2}\|\Sigma_G^2\| + \frac{\theta-1}{\theta}\Sigma_G^\top\Sigma_c + \lambda_c\Big(\frac{\theta-1}{\theta}\mathbb{E}\big[(1-\ell)^{1-\gamma}\big] - \mathbb{E}\big[(1-\ell)^{-\gamma}\big] + \frac{1}{\theta}\Big).$$

A.3 Dynamics of Discounted Dividends

In a next step, we calculate the dynamics of discounted dividends. From A.4 we have that

$$\frac{\mathrm{d}H}{H_-} = \mu_H \mathrm{d}t + \Sigma_H^\top \mathrm{d}W + \mathbb{E}\big[(1-\ell)^{-\gamma} - 1\big]\mathrm{d}N$$

with drift rate and volatility vector

$$\mu_H = -r^f - \lambda_c \mathbb{E}\big[(1-\ell)^{-\gamma} - 1\big], \qquad \Sigma_H = -\gamma\Sigma_c + \frac{\theta-1}{\theta}\Sigma_G.$$

Using Itô's product rule, the dynamics of discounted dividends $\widehat{D} = HD$, where D is given by (3.3), are

$$\frac{\mathrm{d}\widehat{D}}{\widehat{D}_-} = \mu_{\widehat{D}}\mathrm{d}t + \Sigma_{\widehat{D}}^\top \mathrm{d}W + \mathbb{E}\big[(1-\ell)^{\phi-\gamma} - 1\big]\mathrm{d}N$$

with drift rate and volatility vector given by

$$\mu_{\widehat{D}} = \mu_H + \mu_d + \Sigma_H^\top\Sigma_d, \qquad \Sigma_{\widehat{D}} = \Sigma_H + \Sigma_d.$$

A.4 Price-Dividend Ratio

Let $\omega(t,\pi,i,x) \equiv \log\Big(\frac{P}{D}\Big)$ denote the log price-dividend ratio. An application of Itô's Lemma yields[31]

$$\mathrm{d}\omega = \mu_\omega \mathrm{d}t + \Sigma_\omega^\top \mathrm{d}W$$

with drift rate and volatility vector

$$\mu_\omega = \omega_t + \omega_\pi \mu_\pi(\pi,i) + \frac{1}{2}\omega_{\pi\pi}\sigma_\pi^2, \qquad \Sigma_\omega = \omega_\pi \Sigma_\pi.$$

Similarly, the price-dividend ratio $\Omega = e^\omega$ satisfies

$$\frac{\mathrm{d}\Omega}{\Omega} = \big(\mu_\omega + \frac{1}{2}\|\Sigma_\omega\|^2\big)\mathrm{d}t + \Sigma_\omega^\top \mathrm{d}W.$$

[31]For notational convenience we drop the dependencies of ω when appropriate.

In a next step, we rewrite discounted asset prices HP as $\widehat{P}(\widehat{D},\Omega)=\widehat{D}e^{\omega}$. An application of Itô's product rule leads to the following dynamics

$$\frac{\mathrm{d}\widehat{P}}{\widehat{P}_-}=\left(\mu_{\widehat{D}}+\mu_\omega+\frac{1}{2}\|\Sigma_\omega\|^2+\Sigma_\omega^\top\Sigma_{\widehat{D}}\right)\mathrm{d}t+(\Sigma_\omega+\Sigma_{\widehat{D}})^\top\mathrm{d}W+\mathbb{E}\left[(1-\ell)^{\phi-\gamma}-1\right]\mathrm{d}N.$$

Applying the Feynman-Kač Theorem, we have that

$$\mathscr{L}\widehat{P}+\mathrm{e}^{-\omega}\widehat{P}=0, \tag{A.5}$$

where $\mathscr{L}\widehat{P}$ denotes the infinitesimal generator. The no-arbitrage condition implies

$$\frac{\mathscr{L}\widehat{P}}{\widehat{P}_-}=\mu_{\widehat{D}}+\mu_\omega+\frac{1}{2}\|\Sigma_\omega\|^2+\Sigma_\omega^\top\Sigma_{\widehat{D}}+\lambda_c\mathbb{E}\left[(1-\ell)^{\phi-\gamma}-1\right]. \tag{A.6}$$

Substituting (A.6) into (A.5) yields

$$0=\mu_{\widehat{D}}+\mu_\omega+\frac{1}{2}\|\Sigma_\omega\|^2+\Sigma_\omega^\top\Sigma_{\widehat{D}}+\lambda_c\mathbb{E}\left[(1-\ell)^{\phi-\gamma}-1\right]+e^{-\omega}.$$

Finally, we obtain the following PDE for the log price-dividend ratio ω

$$0=\mu_{\widehat{D}}+\omega_t+\omega_\pi\mu_\pi(\pi,i)+\frac{1}{2}(\omega_{\pi\pi}+\omega_\pi^2)\sigma_\pi^2+\omega_\pi\Sigma_\pi^\top\Sigma_{\widehat{D}}+\lambda_c\mathbb{E}\left[(1-\ell)^{\phi-\gamma}-1\right]+e^{-\omega}.$$

Note that this PDE involves the squared partial derivatives of ω. We therefore make the substitution $\Omega=\mathrm{e}^{\omega}$ to transform this PDE into a linear one that can be solved using standard finite difference methods as described in Appendix B. We end up with the following PDE for the price-dividend ratio Ω

$$0=1+\Omega\Big(\mu_{\widehat{D}}+\lambda_c\mathbb{E}\left[(1-\ell)^{\phi-\gamma}-1\right]\Big)+\Omega_t+\Omega_\pi\left(\mu_\pi(\pi,i)+\Sigma_\pi^\top\Sigma_{\widehat{D}}\right)+\Omega_{\pi\pi}\frac{1}{2}\sigma_\pi^2.$$

A.5 Risk Premium

We can write the stock price as $P=\mathrm{e}^{\omega}D$. An application of Itô's product rule yields the following stock price dynamics

$$\frac{\mathrm{d}P}{P_-}=\mu_P\mathrm{d}t+\Sigma_P^\top\mathrm{d}W+\left[(1-\ell)^{\phi}-1\right]\mathrm{d}N-\lambda_c\mathbb{E}\left[(1-\ell)^{\phi}-1\right]\mathrm{d}t$$

with drift rate and volatility vector

$$\mu_p = \mu_\omega + \mu_d + \Sigma_d^\top \Sigma_\omega + \frac{1}{2}\|\Sigma_\omega\|^2 + \lambda_c \mathbb{E}\big[(1-\ell)^\phi - 1\big], \qquad \Sigma_P = \Sigma_\omega + \Sigma_d.$$

In the following, we set $\Gamma_W = \Sigma_P$ and $\Gamma_N = \mathbb{E}\big[(1-\ell)^\phi - 1\big]$, which can be interpreted as the sensitivities of the stock price with respect to the Brownian shocks W and the Poissonian shock N. The expected return on the stock can be computed as the sum of its expected price change μ_p and its dividend yield $y_d = \mathrm{e}^{-\omega}$

$$r^s = \mu_p + y_d.$$

Replacing y_d by the PDE in Equation (4.3), rearranging terms and finally using the expression for the risk-free interest rate or alternatively multiplying risk exposures with the appropriate market prices of risk, we get for the expected equity premium

$$\begin{aligned} r^s - r^f &= \Theta_W^\top \Gamma_W - \Theta_N \Gamma_N \\ &= \Big(\gamma \Sigma_c - \frac{\theta - 1}{\theta} \Sigma_G\Big)^\top \big(\Sigma_\omega + \Sigma_d\big) + \lambda_c \mathbb{E}\Big[\big(1 - (1-\ell)^{-\gamma}\big)\big((1-\ell)^\phi - 1\big)\Big]. \end{aligned}$$

Simplifying yields

$$r^s - r^f = \gamma \Sigma_c^\top \Sigma_d + \gamma \omega_\pi \Sigma_c^\top \Sigma_\pi - \frac{\theta-1}{\theta}\frac{G_\pi}{G}\omega_\pi \sigma_\pi^2 - \frac{\theta-1}{\theta}\frac{G_\pi}{G}\Sigma_\pi^\top \Sigma_d + \lambda_c \mathbb{E}\Big[\big(1-(1-\ell)^{-\gamma}\big)\big((1-\ell)^\phi - 1\big)\Big].$$

Substituting $\omega_\pi = \frac{\Omega_\pi}{\Omega}$ into this expression finishes the proof.

A.6 Term Structure of Interest Rates

Let Z_t^τ denote the price of a default-free zero-coupon bond with maturity τ. We conjecture that

$$Z_t^\tau = \mathrm{e}^{z^\tau(t,\pi,i)}.$$

The dynamics of z^τ follow from an application of Itô's Lemma

$$\frac{\mathrm{d}z^\tau}{z_-^\tau} = = \mu_z \mathrm{d}t + \Sigma_z^\top \mathrm{d}W$$

with drift rate and volatility vector

$$\mu_z = \frac{1}{z^\tau}\left[z_t^\tau + z_\pi^\tau \mu_\pi(\pi,i) + \frac{1}{2} z_{\pi\pi}^\tau \sigma_\pi^2\right], \qquad \Sigma_z = \frac{1}{z^\tau} z_\pi^\tau \Sigma_\pi.$$

Then, the dynamics for $Z^\tau = \mathrm{e}^{z^\tau}$ are given by

$$\frac{\mathrm{d}Z^\tau}{Z_-^\tau} = \left(z^\tau \mu_z + \frac{1}{2}(z^\tau)^2 \|\Sigma_z\|^2\right)\mathrm{d}t + z^\tau \Sigma_z^\top \mathrm{d}W.$$

Next, we determine the dynamics of the discounted price of the zero-coupon bond $\widehat{Z}^\tau = HZ^\tau$. An application of Itô's product rule leads to the following dynamics

$$\frac{\mathrm{d}\widehat{Z}^\tau}{\widehat{Z}_-^\tau} = \left(\mu_H + z^\tau \mu_z + \frac{1}{2}(z^\tau)^2 \|\Sigma_z\|^2 + z^\tau \Sigma_z^\top \Sigma_H\right)\mathrm{d}t + (z^\tau \Sigma_z + \Sigma_H)^\top \mathrm{d}W + \mathbb{E}\left[(1-\ell)^{-\gamma} - 1\right]\mathrm{d}N.$$

Applying the Feynman-Kač Theorem, we have that

$$\mathscr{L}\widehat{Z}^\tau = 0. \tag{A.7}$$

No-arbitrage implies

$$\frac{\mathscr{L}\widehat{Z}^\tau}{\widehat{Z}_-^\tau} = \mu_H + z^\tau \mu_z + \frac{1}{2}(z^\tau)^2 \|\Sigma_z\|^2 + z^\tau \Sigma_z^\top \Sigma_H + \lambda_c \mathbb{E}\left[(1-\ell)^{-\gamma} - 1\right]. \tag{A.8}$$

Substituting (A.8) into (A.7), we obtain the following PDE for the log price z^τ

$$0 = \mu_H + z_t^\tau + z_\pi^\tau \mu_\pi(\pi,i) + \frac{1}{2}\left(z_{\pi\pi}^\tau + (z_\pi^\tau)^2\right)\sigma_\pi^2 + z_\pi^\tau \Sigma_\pi^\top \Sigma_H + \lambda_c \mathbb{E}\left[(1-\ell)^{-\gamma} - 1\right].$$

Finally, we get the following PDE for the price of the zero-coupon bond $Z^\tau = \mathrm{e}^{z^\tau}$

$$0 = Z^\tau\Big(\mu_H + \lambda_c \mathbb{E}\left[(1-\ell)^{-\gamma} - 1\right]\Big) + Z_t^\tau + Z_\pi^\tau\left(\mu_\pi(\pi,i) + \Sigma_\pi^\top \Sigma_H\right) + Z_{\pi\pi}^\tau \frac{1}{2}\sigma_\pi^2.$$

B Solution Approach

The PDEs in (3.8), (4.3) and (4.6) have to be solved numerically. We describe our solution algorithm for illustration purposes for the PDE in (3.8). The other two PDEs can be solved in a similar fashion.

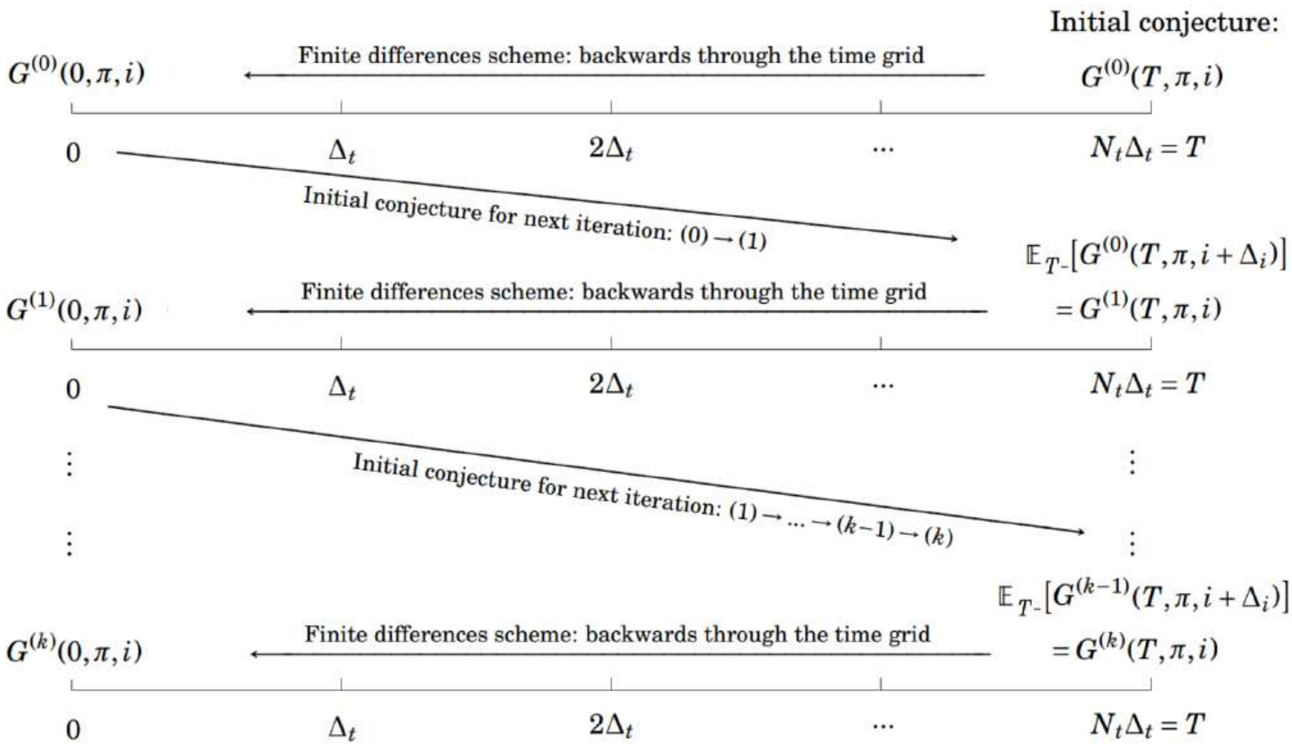

Figure B.1: Solving the PDE for the Value Function. This figure depicts a graphical illustration of our numerical solution approach to solve for the value function. The algorithm terminates when the deviation between two iterations has become negligibly small, i.e., when $\|G^{(k)}(0,\pi,i) - G^{(k-1)}(0,\pi,i)\| \leq \varepsilon$ for a small number $\varepsilon > 0$. In our implementation, we use $\varepsilon = 10^{-5}$. Then, we set $G = G^{(k)}$ and have found the value function. The same iterative method is employed for the computation of the price-dividend ratio Ω.

B.1 Benchmark Solution Approach

Basic Idea Although we face a problem with an infinite time horizon, the indirect utility function G depends on time t since the FOMC meetings take place after fixed time intervals. Therefore, the PDE for the indirect utility function has to be solved on $[0,T]$, denoting the time span between two FOMC meetings. In our implementation, we employ an iterative method to solve for the indirect utility function. We start with an initial conjecture and set $G^{(0)}(T,\pi,i) = 1$, where the superscript denotes the number of performed iterations. Note that the precise form of this boundary condition is not decisive but affects the number of iterations to be executed until the indirect utility function converges. Starting with this boundary condition, we work backwards through the time grid until we have the value for $G^{(0)}(0,\pi,i)$. We then update our conjecture and use $G^{(1)}(T,\pi,i) = \sum_{\Delta_i \in \mathscr{I}} G^{(0)}(0,\pi,i+\Delta_i)\mathbb{P}(\Delta_i \mid \pi,i)$ as initial conjecture for the next iteration. Using this initial conjecture, we again work backwards through the time grid until we have determined $G^{(1)}(0,\pi,i)$ and repeat this iteration method as long as the deviation between two iterations has become negligibly small, i.e., when $\|G^{(k)}(0,\pi,i) - G^{(k-1)}(0,\pi,i)\| \le \varepsilon$ for a small number $\varepsilon > 0$. In our implementation, we use $\varepsilon = 10^{-5}$. Then, we set $G = G^{(k)}$ and have found the value function. In Figure B.1 we present a graphical illustration of our solution algorithm.

Definition of the Grid We use a grid-based solution approach to solve the non-linear PDE in Equation (3.8). We discretize the (t,π)-space using an equally-spaced lattice. Its grid points are defined by

$$\{(t_n, \pi_j) \mid n = 0, \cdots, N_t, j = 0, \cdots, N_\pi\},$$

where $t_n = n\Delta_t$ and $\pi_j = j\Delta_\pi$ for some fixed grid size parameters Δ_t and Δ_π that denote the distances between two grid points. The numerical results are based on a choice of $N_\pi = 100$ and four time steps per month. However, the results hardly change if we use a finer grid or only one time step per month. In the following, let $G_{n,j,i}$ denote the approximated value function at the grid point (t_n, π_j, i). We apply an implicit finite-difference approach.

Finite Differences Approach In this paragraph, we describe the numerical solution approach in more detail. We have to solve the following semi-linear PDE

$$0 = G_t + \delta\theta G^{1-1/\theta} + M_1 G + M_2 G_\pi + M_3 G_{\pi\pi}$$

with state-dependent coefficients $M_s = M_s(t,\pi,i)$

$$\begin{aligned}
M_1 &= (1-\gamma)\big(\mu_c(\pi,i) - \frac{1}{2}\gamma\sigma_c^2\big) + \lambda_c \mathbb{E}\big[(1-\ell)^{1-\gamma} - 1\big] - \delta\theta, \\
M_2 &= \mu_\pi(\pi,i) + (1-\gamma)\sigma_c\sigma_\pi\rho_{c\pi}, \\
M_3 &= \frac{1}{2}\sigma_\pi^2.
\end{aligned}$$

In the finite difference approximation we use so-called "up-wind" approximations of the derivatives, which tends to stabilize the approach. Due to the implicit approach, we approximate the time derivative by forward finite differences. For each point (n,j,i) in the interior of the grid the relevant finite difference approximations of the derivatives are then given by

$$\mathrm{D}_\pi^+ G_{n,j,i} = \frac{G_{n,j+1,i} - G_{n,j,i}}{\Delta_\pi}, \qquad \mathrm{D}_\pi^- G_{n,j,i} = \frac{G_{n,j,i} - G_{n,j-1,i}}{\Delta_\pi},$$

$$\mathrm{D}^2_{\pi\pi} G_{n,j,i} = \frac{G_{n,j+1,i} - 2G_{n,j,i} + G_{n,j-1}}{\Delta_\pi^2}, \qquad \mathrm{D}^+_t G_{n,j,i} = \frac{G_{n+1,j,i} - G_{n,j,i}}{\Delta_t}.$$

Substituting these expressions into the PDE above, yields the following semi-linear equation for the grid point (t_n, π_j, i)

$$\begin{aligned} G_{n+1,j,i} \frac{1}{\Delta_t} &= G_{n,j,i} \Big[-M_1 + \frac{1}{\Delta_t} + \mathrm{abs}\Big(\frac{M_2}{\Delta_\pi}\Big) + 2\frac{M_3}{\Delta_\pi^2} \Big] \\ &+ G_{n,j-1,i} \Big[\frac{M_2^-}{\Delta_\pi} - \frac{M_3}{\Delta_\pi^2} \Big] + G_{n,j+1,i} \Big[-\frac{M_2^+}{\Delta_\pi} - \frac{M_3}{\Delta_\pi^2} \Big] \\ &+ \delta\theta G_{n,j,i}^{1-1/\theta}. \end{aligned}$$

Therefore, for a fixed point in time each grid point is determined by a non-linear equation. In the interior of the grid, we use centered finite differences. At the boundaries, we apply forward or backward differences.

B.2 Alternative Solution Approach

Instead of iteratively solving the model between two FOMC meetings taking the boundary condition (3.5) into account, one could also formulate a model where the federal funds rate i follows a Markov chain whose jump intensity $\lambda_i(t,\pi,i)$ is time-dependent and increases drastically around the FOMC meetings (e.g., Piazzesi 2005). If a jump occurs, the new value of i is sampled from a probability distribution as described in the main text. This leads to an additional jump component in the PDE

$$\begin{aligned} 0 = f(c,J) + J_t + J_c \mu_c(\pi,i) c + J_\pi \mu_\pi(\pi,i) + \frac{1}{2} J_{cc} \sigma_c^2 c^2 + \frac{1}{2} J_{\pi\pi} \sigma_\pi^2 + J_{c\pi} \sigma_c \sigma_\pi \rho_{c\pi} c \\ + \lambda_c \mathbb{E}\big[J(t, c(1-\ell), \pi, i) - J\big] + \lambda_i(t,\pi,i) \mathbb{E}\big[J(t,c,\pi,i+\Delta_i) - J(t,c,\pi,i)\big]. \end{aligned}$$

This PDE can be solved with a similar standard finite difference technique as outlined above. Assuming a jump intensity of zero between any two FOMC meetings and a very large intensity in a short interval around a meeting leads to almost identical results.

B.3 Term Structure of Interest Rates

Theorem 4.4 provides a simple algorithm to calculate the term structure of interest rates. The current price $Z^T(0,\pi,i)$ of the bond that matures at T can be computed by solving the PDE (4.6) over $[0,T]$ with the terminal condition $Z^T(T,\pi,i) = 1$.

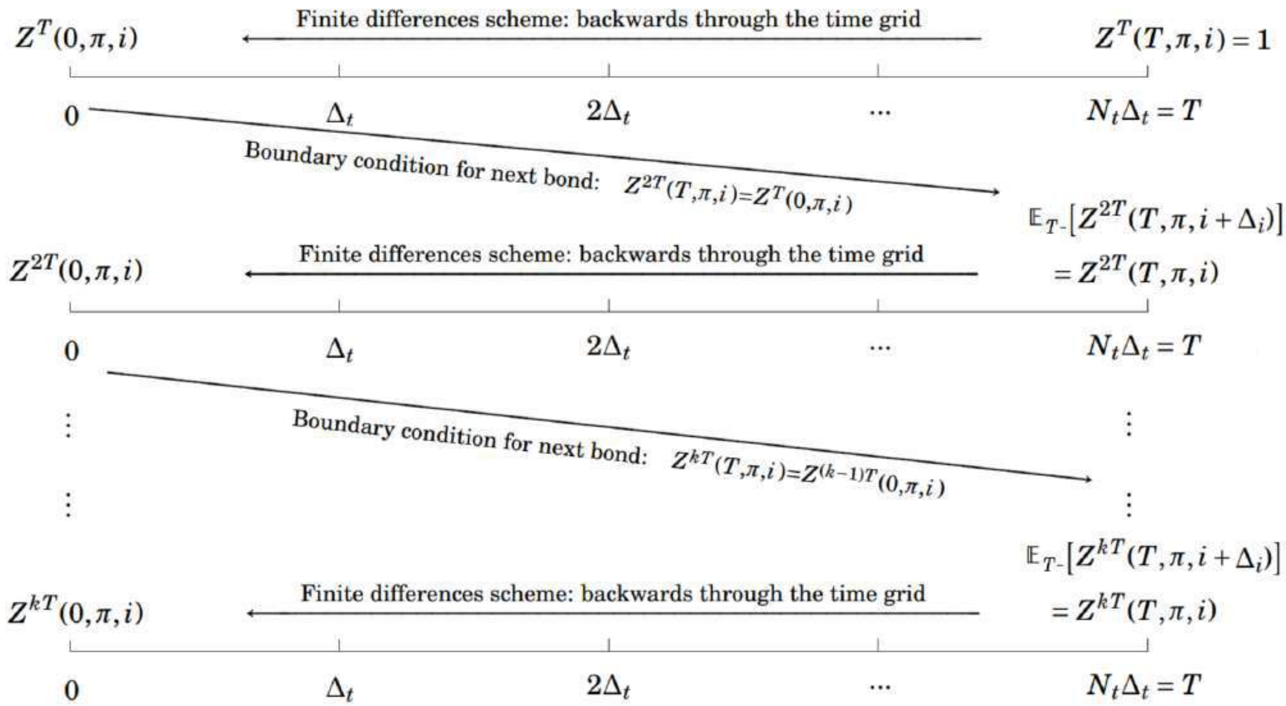

Figure B.2: Solving the PDEs for the Zero-coupon Bonds. This figure depicts a graphical illustration of our numerical solution approach to solve for the term structure of interest rates.

To calculate the price of the bond maturing at $\tau = 2T$, we note that the price of the $2T$-bond at time T equals the price of the T-bond at time 0 under the then prevailing federal funds rate i, i.e., $Z^{2T}(T,\pi,i) = Z^T(0,\pi,i)$. Taking this into account, it is sufficient to solve the PDE for the $2T$-bond on the interval $[0,T]$ subject to the boundary condition $Z^{2T}(T,\pi,i) = \sum_{\Delta_i \in \mathscr{I}} Z^{2T}(T,\pi,i+\Delta_i)\mathbb{P}(\Delta_i \mid \pi, i)$.

Having computed the prices $Z^T,\ldots,Z^{kT}$ of the first k zero-coupon bonds ($k \in \mathbb{N}$), the price of the zero-coupon bond that matures at $\tau = (k+1)T$ can be computed by solving the PDE (4.6) with boundary condition $Z^{(k+1)T}(T,\pi,i) = \sum_{\Delta_i \in \mathscr{I}} Z^{k+1)T}(T,\pi,i+\Delta_i)\mathbb{P}(\Delta_i \mid \pi, i)$. Note that the right-hand side is known since we have already computed the price of the kT-bond and $Z^{(k+1)T}(T,\pi,i) = Z^{kT}(0,\pi,i)$. A graphical illustration of the solution algorithm is presented in Figure B.2. A similar recursion in a discrete time model has been employed in Piazzesi and Schneider (2007).

(a) VAR(1)-Model with all Coefficients

	Consumption	*Inflation Rate*	*Federal Funds Rate*
Lagged Consumption Growth	0.822***	0.106***	0.218***
	(0.038)	(0.037)	(0.051)
Lagged Inflation	–0.040	0.951***	0.149***
	(0.031)	(0.031)	(0.042)
Lagged Federal Funds Rate	0.005	0.017	0.860***
	(0.023)	(0.023)	(0.032)
Constant	0.006***	–0.191	–0.443**
	(0.001)	(0.138)	(0.191)
R^2	0.708	0.922	0.914

(b) VAR(1)-Model with Consumption Growth Coefficients Set to Zero

	Consumption	*Inflation Rate*	*Federal Funds Rate*
Lagged Inflation	–0.301***	0.918***	0.080*
	(0.049)	(0.029)	(0.040)
Lagged Federal Funds Rate	0.181***	0.039*	0.906***
	(0.038)	(0.022)	(0.031)
Constant	0.029***	0.110	0.175
	(0.002)	(0.091)	(0.128)
R^2	0.134	0.920	0.907

Table C.1: Regression Results: VAR(1)-Models. Panel (a) reports the estimation results of a standard VAR(1)-model of consumption growth, inflation, and the federal funds rate. Panel (b) reports the estimation results of the restricted VAR(1)-model (C.1) of consumption growth, inflation, and the federal funds rate with several coefficients set to zero. Our sample period is from January 1960 until December 2020. Standard errors are reported in brackets. Asterisks correspond to the following p-values: $*p < 0.1, **p < 0.05, ***p < 0.01$.

C Empirical Analyses and Calibration Details

This section provides further empirical evidence of the relation between consumption growth, inflation rates, and the federal funds rate. In the whole section, we use quarterly data as described in Section 2. Those estimation results are the basis for our calibration strategy as described in Section 5.

C.1 VAR(1)-Models

Complementing the empirical results of Section 2, we now estimate a standard VAR(1)-model of consumption growth, inflation, and the federal funds rate. This model is the preferred model according to the Bayesian information criterion and can replicate most of the stylized facts from the VAR(9)-model in the main text. However, it fails to replicate the negative short-term effect of positive interest rates shocks on consumption growth as well as the negative long-term effect of positive interest rates shocks on inflation. The estimation results are reported in Panel (a) of Table C.1.

To connect the asset pricing model from Section 3 to the data described in Section 2 and to provide interpretable estimation results, we now follow Kraft et al. (2019) and estimate a restricted VAR(1)-model with several coefficients set to zero such that it adequately discretizes our consumption and inflation dynamics. More precisely, we set the coefficients of lagged consumption growth to zero and estimate the coefficients of

$$\begin{pmatrix} g_{t+1} \\ \pi_{t+1} \\ i_{t+1} \end{pmatrix} = \begin{pmatrix} \beta_{0,c} \\ \beta_{0,\pi} \\ \beta_{0,i} \end{pmatrix} + \begin{pmatrix} \beta_{\pi,c} & \beta_{i,c} \\ \beta_{\pi,\pi} & \beta_{i,\pi} \\ \beta_{\pi,i} & \beta_{i,i} \end{pmatrix} \begin{pmatrix} \pi_t \\ i_t \end{pmatrix} + \varepsilon_{t+1}, \tag{C.1}$$

where ε_{t+1} is a vector of error terms with zero mean and time-invariant variance-covariance matrix Σ. The regression results are reported in Panel (b) of Table C.1. Inflation has a significant negative effect on consumption growth, indicating that a positive shock to inflation today predicts lower real growth in the future. Notice that the point estimate of $\beta_{i,\pi}$ is positive but small and statistically insignificant at the usual 5% level with a p-value of 7.2%. This result implies that higher interest rates have at most a weak positive effect on inflation in line with the third stylized fact from Section 2.2. Note that for the unrestricted VAR(1)-model as reported in Panel (a), this coefficient is even smaller and not even significant at the 10% level corroborating this result.[32]

C.2 VAR(1)-Models with Interaction Term

To improve the model fit, we go one step further and add an interaction term between inflation and the federal funds rate. The idea behind such an interaction is that an increase in the federal funds rate, when inflation is already high, may signal more persistent inflation expectations, which could potentially affect both economic growth and inflation. To capture this effect, we adjust Regression (C.1) and incorporate an interaction term $\pi_t \cdot i_t$, leading to the model

$$\begin{pmatrix} g_{t+1} \\ \pi_{t+1} \\ i_{t+1} \end{pmatrix} = \begin{pmatrix} \beta_{0,c} \\ \beta_{0,\pi} \\ \beta_{0,i} \end{pmatrix} + \begin{pmatrix} \beta_{\pi,c} & \beta_{i,c} & \beta_{i\times\pi,c} \\ \beta_{\pi,\pi} & \beta_{i,\pi} & \beta_{i\times\pi,\pi} \\ \beta_{\pi,i} & \beta_{i,i} & \beta_{i\times\pi,i} \end{pmatrix} \begin{pmatrix} \pi_t \\ i_t \\ \pi_t \cdot i_t \end{pmatrix} + \varepsilon_{t+1}, \tag{C.2}$$

where ε_{t+1} is a vector of error terms with zero mean and time-invariant variance-covariance matrix Σ. We estimate this model with three-stage least squares (3SLS) as originally introduced by Zellner and

[32] The positive coefficient might seem to be counter-intuitive at the first glance, but it can be explained as follows: Higher interest rates might trigger inflation expectations, leading to a positive effect on inflation in the short-run, but eventually help curb inflation in the long-run as can also be seen from Figure 2 in the main text.

(a) VAR(1)-Model with Interaction Term

	Consumption	*Inflation Rate*	*Federal Funds Rate*
Lagged Inflation	0.009	0.895***	0.229***
	(0.085)	(0.043)	(0.058)
Lagged Federal Funds Rate	0.376***	0.031	0.986***
	(0.054)	(0.027)	(0.037)
Lagged Federal Funds Rate × Lagged Inflation	–3.786***	0.002	–0.018***
	(0.741)	(0.004)	(0.005)
Constant	0.017***	0.166	–0.312*
	(0.003)	(0.133)	(0.181)
R^2	0.190	0.919	0.913

(b) VAR(1)-Model with Interaction Term and Insignificant Coefficients Set to Zero

	Consumption	*Inflation Rate*	*Federal Funds Rate*
Lagged Inflation		0.957***	0.252***
		(0.043)	(0.055)
Lagged Federal Funds Rate	0.366***		0.972***
	(0.053)		(0.035)
Lagged Federal Funds Rate × Lagged Inflation	–3.681***		–0.019***
	(0.494)		(0.005)
Constant	0.018***	0.158*	–0.308*
	(0.002)	(0.133)	(0.174)
R^2	0.190	0.917	0.913

Table C.2: Regression Results: VAR(1)-Models with Interaction Term. Panel (a) reports the estimation results of the VAR(1)-model (C.2) of consumption growth, inflation, the federal funds rate, and an interaction term between inflation and the federal funds rate. Panel (b) reports the estimation results of the restricted VAR(1)-model (C.2) of consumption growth, inflation, the federal funds rate, and an interaction term between inflation and the federal funds rate with insignificant coefficients set to zero. Our sample period is from January 1960 until December 2020. Standard errors are reported in brackets. Asterisks correspond to the following p-values: $*p < 0.1, **p < 0.05, ***p < 0.01$.

Theil (1962). The corresponding regression results are reported in Panel (a) of Table C.2. It appears that the interaction term sizably increases the explanatory power of inflation and the federal funds rate for real consumption growth compared to the restricted VAR(1)-model from Section C.1. The coefficient of determination increases from $R^2 = 13.4\%$ to $R^2 = 19\%$, while the explanatory power of the model for inflation and the federal funds rate is about the same as in Section C.1. Since the estimates $\widehat{\beta}_{\pi,c}$, $\widehat{\beta}_{i,\pi}$, and $\widehat{\beta}_{i\times\pi,\pi}$ are not significantly different from zero, we set those parameters to zero as in Kraft et al. (2019). Subsequently, we run a version of Regression (C.2) with those parameters set to zero, see Panel (b) of Table C.2. This specification will serve as the benchmark specification for our model calibration. Overall, the model fit is very good and excluding non-significant parameters does not harm the goodness of fit. We illustrate the model fit in Figure 5. We use the estimated parameters along with the estimated

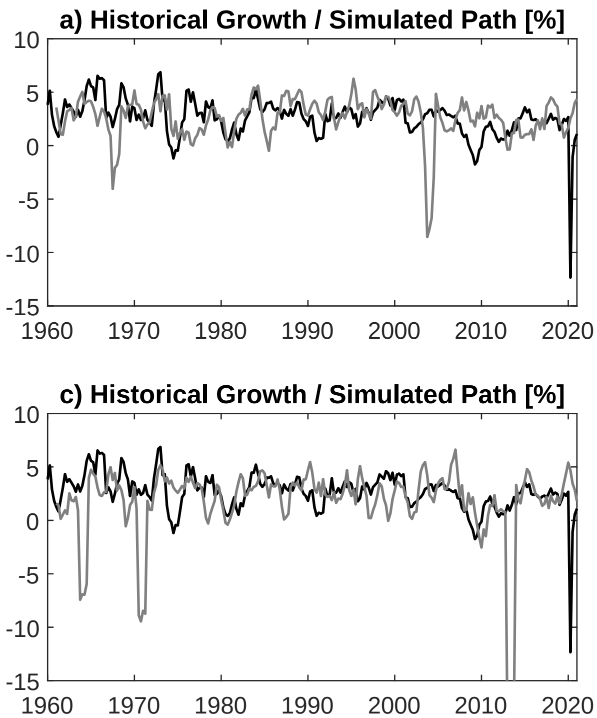

Figure C.3: Historical versus Simulated Growth Rates. The figure depicts the historical and simulated annualized real consumption growth rate. For simulated data, we show four different sample paths (a)-(d). Black lines (——) show historical time series, whereas gray lines (——) present model simulation results. We use quarterly data. All results are annualized.

cross-equation error variance–covariance matrix as a basis for our benchmark calibration as described in Section 5.

C.3 Additional Calibration Details

Consumption Growth We test empirically whether our calibrated consumption growth model replicates the historical distribution of consumption growth. Figure C.3 depicts simulated real consumption growth rates (——) with the calibration as described in Section 5 of the main text along with the historical growth rates (——). We perform a Kolmogorov-Smirnov test and test the null hypothesis that the historical growth rate distribution F and the simulated growth rate distribution $\widehat{F}$ are identical, i.e., $H_0 : F = \widehat{F}$. The test cannot reject this hypothesis with a p-value of 19.0%.

Elasticity of Intertemporal Substitution As discussed in Footnote 24, there is mixed evidence about the magnitude of the EIS. We choose the EIS such that we match the historical real US Treasury

Constant	0.0027***
	(0.005)
Inflation	–1.052***
	(0.018)
FED Rate	0.868***
	(0.013)
R^2	0.957

Table C.3: Regressing the Real US Treasury Bill Rate on Inflation and Federal Funds Rate. The table reports the results of an OLS regression of the real US Treasury bill rate on the inflation rate and the federal funds rate. Standard errors are reported in brackets. Asterisks indicate *** $p < 0.01$.

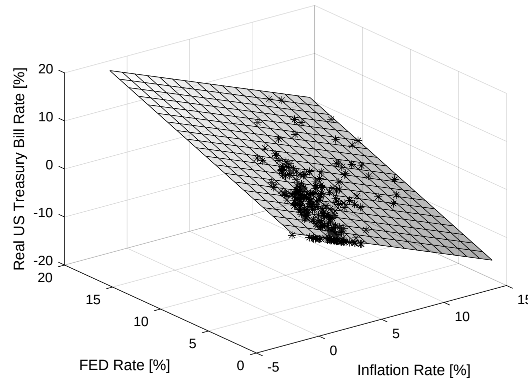

Figure C.4: Regressing the Real US Treasury Bill Rate on Inflation and Federal Funds Rate. The figure illustrates the results of an OLS regression of the real US Treasury bill rate on the inflation rate and the federal funds rate. The plane is parameterized with the parameters reported in Table C.3.

bill rate. Running the following regression of the real US Treasury bill rate on inflation and the federal funds rate

$$r_{f,t} = \beta_{0,r} + \beta_{i,r} i_t + \beta_{\pi,r}\pi_t + \epsilon_t,$$

we observe an almost perfect fit. The corresponding regression results are reported in Table C.3 and visualized in Figure C.4.

We conjecture and confirm in Section 6 that the detected state dependency of the Treasury bill rate primarily stems from the term capturing consumption smoothing, i.e., $\frac{1}{\psi}\mu_c(\pi_t, i_t)$, in (4.1). Therefore, we choose the EIS in such a way that the surface of consumption growth $\mu_c(\pi_t, i_t)$ depicted in Figure 5(a) is tilted to match the risk-free rate surface obtained from the regression outlined above.[33] This can be achieved by setting $\psi = 0.32$. Note that this calibration ensures preferences for early resolution of uncertainty since $\psi > \frac{1}{\gamma}$.

Interest Rate Adjustments Table C.4 complements Figure 6 and reports the probability distribution of interest rate adjustments (in basis points) for the four four different scenarios described in the main text.

[33] Similarly, Hansen and Singleton (1983) and Hall (1988), among others, regress the risk-free interest rate on consumption growth, i.e., $r_{f,t} = \beta_0 + \beta_1 g_t + \epsilon_t$. Then, the EIS is estimated by $\psi = 1/\widehat{\beta}_1$. These studies typically find values well below one in line with our estimate.

	–75bps	–50bps	–25bps	0bps	25bps	50bps	75bbs
Scenario I	2.46	2.46	9.20	57.06	19.02	7.97	1.84
Scenario II	3.91	6.71	18.72	34.92	26.54	7.26	1.95
Scenario III	0	1.41	4.23	88.72	4.23	0	1.41
Scenario IV	0.69	6.21	14.48	65.52	10.34	2.07	0.69

Table C.4: Distribution of Interest Rate Adjustments. The table reports the probability distribution of interest rate adjustments (in basis points) for four different scenarios. In Scenario I, the inflation rate is above the 2% target and exceeds the federal funds rate. In Scenario II, the inflation rate is above the 2% target but is below the federal funds rate. Scenario III corresponds to the situation in which the inflation rate is below the 2% target and exceeds the federal funds rate, while in Scenario IV, the inflation rate is below the 2% target and below the federal funds rate. The probabilities are given in percentage terms.

D Decomposition of the Risk-free Rate and the Equity Premium

We now want to explicate the mechanism through which our asset pricing model generates a state-dependent risk-free rate and equity premium. Therefore, we decompose both key asset pricing moments into their different components.

Risk-free Rate From Theorem 4.1, we know that the real risk-free rate can be decomposed as follows

$$r^f = \underbrace{\delta}_{\text{Discounting}} + \underbrace{\frac{1}{\psi}\mu_c(\pi,i)}_{\text{Smoothing}} - \underbrace{\frac{1}{2}\gamma\Big(\frac{1}{\psi}+1\Big)\sigma_c^2}_{\text{Standard Diffusion Risk}} + \underbrace{\lambda_c\Big(\frac{\theta-1}{\theta}\mathbb{E}\big[(1-\ell)^{1-\gamma}\big]-\mathbb{E}\big[(1-\ell)^{-\gamma}\big]+\frac{1}{\theta}\Big)}_{\text{Macroeconomic Disaster Risk}}$$
$$+ \underbrace{\frac{\theta-1}{2\theta^2}\frac{G_\pi^2}{G^2}\sigma_\pi^2}_{\text{Inflation Risk}} + \underbrace{\frac{\theta-1}{\theta}\frac{G_\pi}{G}\sigma_\pi\sigma_c\rho_{c\pi}}_{\text{Interaction Risk}}.$$

To identify the importance of the state-dependent terms, we report this decomposition for the prevailing inflation rates and federal funds rates at the beginning of several selected years in Table D.5. The constant terms in this decomposition are the discounting component $\delta = 0.75\%$, precautionary savings with respect to standard diffusion risk $-\frac{1}{2}\gamma\big(\frac{1}{\psi}+1\big) = 0.15\%$, and precautionary savings against macroeconomic disaster risk $\lambda_c\big(\frac{\theta-1}{\theta}\mathbb{E}\big[(1-\ell)^{1-\gamma}\big]-\mathbb{E}\big[(1-\ell)^{-\gamma}\big]+\frac{1}{\theta}\big) = 8.11\%$. Thus, the constant terms contribute as much as –7.54% to the risk-free rate due to the large amount of precautionary savings against disaster risk. The largest part of the variation of the risk-free rate over time stems from the desire for consumption smoothing represented by the term $\frac{1}{\psi}\mu(\pi_t,i_t)$ with values ranging between –1% and 12% with a sample mean of 8.27%. We also find that only a small portion of the variation of the risk-free

Year	*Inflation Rate*	*Federal Funds Rate*	*Constant Terms*	*Consumption Smoothing*	*Inflation Risk*	*Interaction Risk*	***Total*** r_f
1960	1.24	4.00	–7.51	9.58	0.02	0.00	2.09
1965	1.10	4.00	–7.51	9.64	0.01	0.00	2.15
1970	6.16	9.00	–7.51	9.49	0.04	0.00	2.03
1975	11.75	7.25	–7.51	4.07	0.05	0.00	–3.39
1980	13.87	14.00	–7.51	–0.75	0.22	0.00	–8.03
1985	3.53	8.50	–7.51	11.85	0.01	0.00	4.35
1990	5.20	8.25	–7.51	10.08	0.02	0.00	2.59
1995	2.87	5.75	–7.51	10.25	0.01	0.00	2.75
2000	2.79	5.50	–7.51	10.10	0.01	0.00	2.60
2005	2.84	2.50	–7.51	7.62	0.05	0.00	0.15
2010	2.62	0.25	–7.51	5.79	0.03	0.00	–1.70
2015	–0.23	0.25	–7.51	5.87	0.01	0.00	–1.63
2020	2.51	1.75	–7.51	7.07	0.05	0.00	–0.39
Average	3.67	5.00	–7.51	8.27	0.04	0.00	**0.79**

Table D.5: Decomposition of the Risk-free Rate. The table reports the prevailing inflation and federal funds rate at the beginning of several years along with a decomposition of the risk-free rate into its several components. All results are expressed in percentage terms and have been generated with the parameters reported in Table 2.

rate can be attributed to the inflation risk component $\frac{\theta-1}{2\theta^2}\frac{G_\pi^2}{G^2}\sigma_\pi^2$, while the interaction risk component $\frac{\theta-1}{\theta}\frac{G_\pi}{G}\sigma_\pi\sigma_c\rho_{c\pi}$ is negligible as can be seen from Table D.5.

Equity Premium We now want to explicate the mechanism through which our asset pricing model generates a large equity premium and how the several components contribute to this premium. Theorem 4.3 provides the following decomposition of the equity premium

$$r^s - r^f = \underbrace{\gamma\sigma_c\sigma_d\rho_{cd}}_{\text{Diffusion Risk}} + \underbrace{\lambda_c\mathbb{E}\big[\big(1-(1-\ell)^{-\gamma}\big)\big((1-\ell)^{\phi}-1\big)\big]}_{\text{Macroeconomic Disaster Risk}} - \underbrace{\frac{\theta-1}{\theta}\frac{G_\pi}{G}\frac{\Omega_\pi}{\Omega}\sigma_\pi^2}_{\text{Inflation Risk}} \\ \underbrace{-\frac{\theta-1}{\theta}\frac{G_\pi}{G}\Sigma_\pi^\top\Sigma_d + \gamma\frac{\Omega_\pi}{\Omega}\sigma_c\sigma_\pi\rho_{c\pi}}_{\text{Interaction Risk}}.$$

The decomposition of the equity premium into its several components is given in Table D.5. The diffusion risk premium depends on the exposure of the price of the dividend claim to diffusive risk and on the market price of risk. It is given by $\gamma\sigma_c\sigma_d\rho_{cd}$ and equal to 0.25%. The macroeconomic disaster risk premium is given by the disaster intensity multiplied with the negative covariance between the relative change in the SDF and the relative change in the price of the dividend claim when a dis-

Year	*Inflation Rate*	*Federal Funds Rate*	*Constant Terms*	*Inflation Risk*	*Interaction Risk*	***Total*** $r_s - r_f$
1960	1.24	4.00	5.26	–0.02	–0.05	5.19
1965	1.10	4.00	5.26	–0.02	–0.07	5.17
1970	6.16	9.00	5.26	–0.03	0.15	5.39
1975	11.75	7.25	5.26	–0.03	0.95	6.18
1980	13.87	14.00	5.26	–0.17	0.30	5.39
1985	3.53	8.50	5.26	–0.00	–0.27	4.99
1990	5.20	8.25	5.26	–0.01	0.07	5.31
1995	2.87	5.75	5.26	–0.02	–0.04	5.21
2000	2.79	5.50	5.26	–0.02	–0.01	5.22
2005	2.84	2.50	5.26	–0.06	0.46	5.65
2010	2.62	0.25	5.26	–0.04	0.87	6.10
2015	–0.23	0.25	5.26	–0.01	0.67	5.92
2020	2.51	1.75	5.26	–0.06	0.55	5.75
Average	3.67	5.00	5.26	–0.04	0.26	**5.49**

Table D.6: Decomposition of the Equity Premium. The table reports the prevailing inflation and federal funds rate at the beginning of several years along with a decomposition of the equity premium into its several components. All results are expressed in percentage terms and have been generated with the parameters reported in Table 2.

aster occurs. This component is by far the largest contributor to the equity premium and given by $\lambda_c \mathbb{E}\left[\left(1-(1-\ell)^{-\gamma}\right)\left((1-\ell)^{\phi}-1\right)\right] = 5.01\%$. While the desire for consumption smoothing greatly generates a state-dependent risk-free rate, the state-dependency of the equity premium is less pronounced. As for the risk-free rate, the inflation risk component $\frac{\theta-1}{\theta}\frac{G_\pi}{G}\frac{\Omega_\pi}{\Omega}\sigma_\pi^2$ and the interaction risk component $\frac{\theta-1}{\theta}\frac{G_\pi}{G}\Sigma_\pi^\top \Sigma_d + \gamma \frac{\Omega_\pi}{\Omega}\sigma_c \sigma_\pi \rho_{c\pi}$ are relatively insignificant compared to the other components. Still, the latter component plays a larger role than its counterpart for the risk-free rate and contributes on average 0.26% to the equity premium, i.e., about as much as the diffusion risk premium.

E Additional Analyses

E.1 Price Impact of Monetary Policy

Figure E.5 complements our analyses from Section 6.3 and illustrates the price impact of restrictive monetary policy on bond and stock prices if the central bank increases the federal funds rate by 25bps or 75bps, respectively. As one would expect, the higher the interest rate hike, the greater the impact on prices.

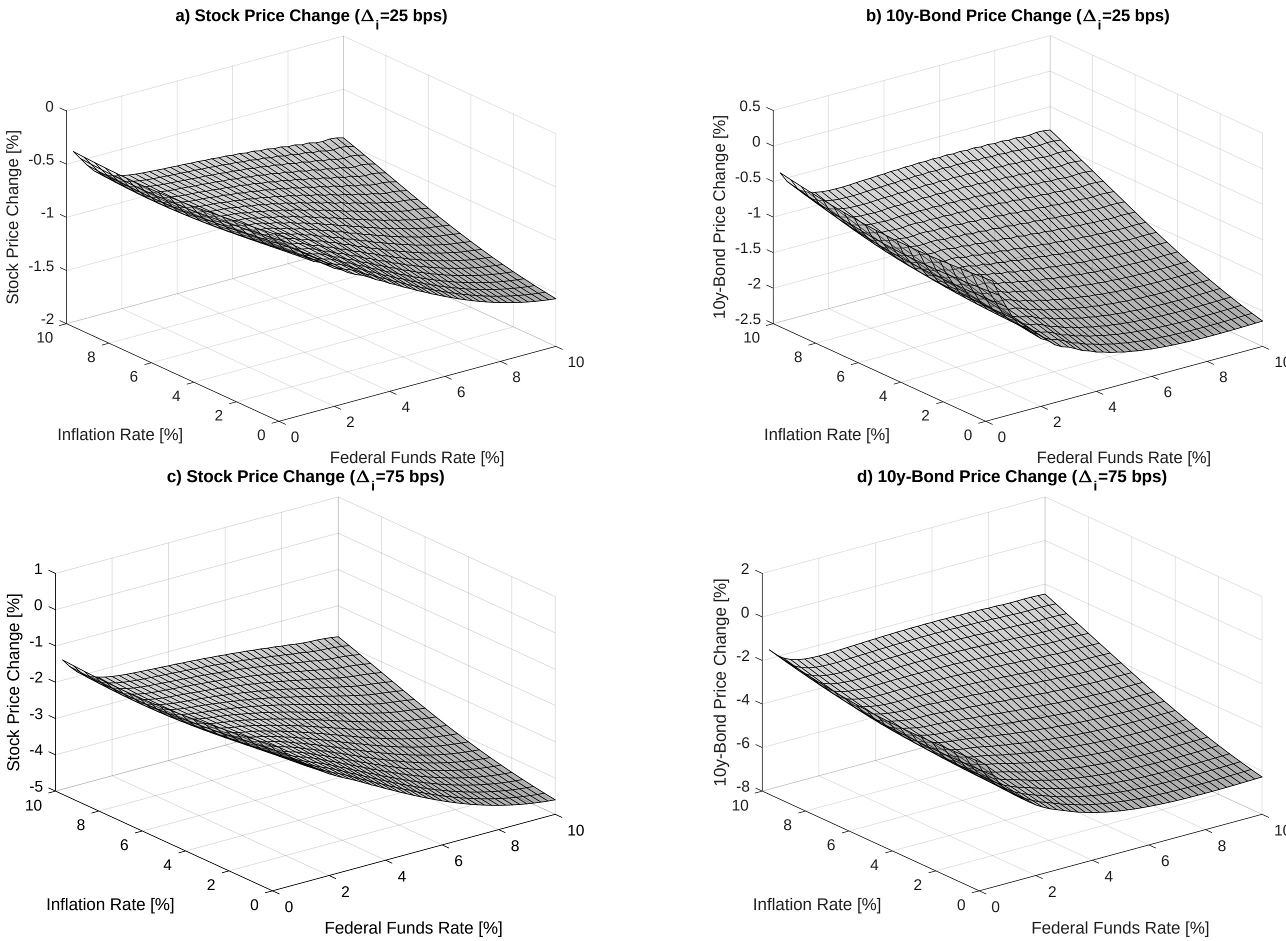

Figure E.5: Price Impact of Monetary Policy. The figure depicts the model-implied (a) stock price change and (b) 10y bond price change in response to a change in the federal funds rate of $\Delta_i = 25$bps conditional on the inflation and federal funds rate. Panels (c) and (d) report the respective price changes for a change in the federal funds rate of of $\Delta_i = 75$bps.

E.2 Sensitivity Analysis: Elasticity of Intertemporal Substitution

Figure E.6 illustrates a similar analysis as in Figure 11. We vary the elasticity of intertemporal substitution but adjust δ to match the unconditional mean of the risk-free rate in our data set. In this case we cannot match the average dividend yield anymore. This analysis confirms the finding that varying the elasticity of intertemporal substitution too much destroys the model's ability to match the historical pattern of the dividend yield. Moreover, even if we recalibrate the time preference rate such that the model matches the unconditional mean of the risk-free rate in the data, it fails to match the time series of the risk-free rate if the elasticity of intertemporal substitution deviates too much from its benchmark value.

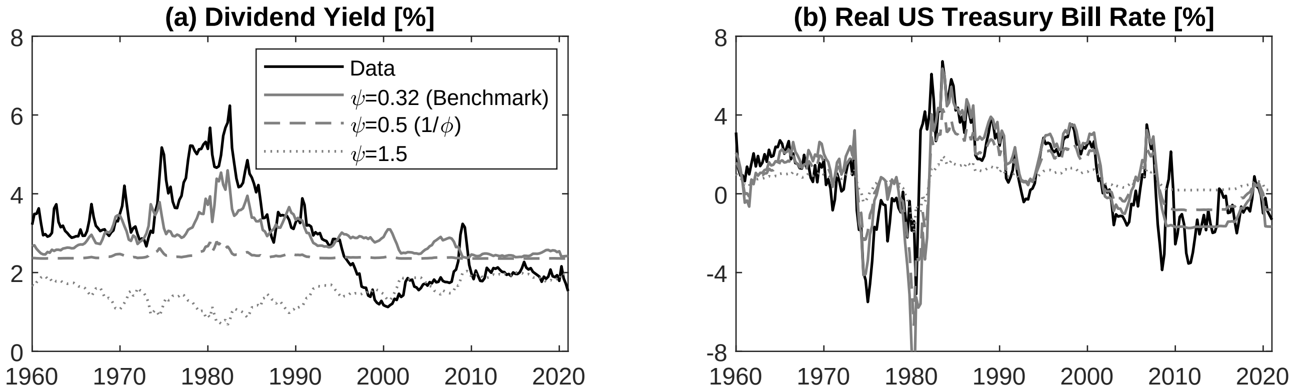

Figure E.6: Sensitivity Analysis: Elasticity. The figure depicts the historical and simulated (a) S&P500 dividend yield series and (b) real US Treasury bill rate for different values of the elasticity of intertemporal substitution. Black lines (——) show historical time series, while gray lines (——) present model simulation results using our benchmark calibration given in Table 2. Gray dashed lines (- - -) represent simulation results using an elasticity of 0.53, which corresponds to the case $\psi = 1/\phi$, whereas gray dotted lines (⋯⋯) show simulation results for an elasticity of 1.5. For the latter two cases we adjust δ accordingly to match the unconditional mean of the US Treasury bill rate. Other parameters are chosen as given in Table 2. We use quarterly data. All results are annualized.

E.3 Model Extension: Structural Break in Dividend Growth

As outlined in Footnote 28, the benchmark model cannot perfectly match the dividend yield series because of a structural break in this time series found by a supremum Wald test. More specifically, we observe a striking shift in the correlation between the dividend yield and the inflation rate around this break point. Prior to 1995, the correlation between the dividend yield and the inflation rate is notably strong and positive at 71%. However, following the structural break, the relationship reverses, with the correlation turning significantly negative at –25%. Similarly, we identify a structural break in the growth rate of the real dividend stream around this year, and the post-break dividend growth rate is slightly higher than the pre-break growth rate. To capture this break in our model, we follow Cecchetti et al. (1993) and assume that the expected dividend growth rate depends on a latent regime shift, i.e., the expected dividend growth rate is $\mu_d(\pi, i, S) = \varphi(S)\mu_c(\pi, i)$, where S is a two-stated Markov chain.

To illustrate the mechanism, we assume a pre-break leverage parameter of 2 as in our benchmark calibration and a post-break leverage parameter of 2.25. Moreover, we assume that the mean-reversion level negatively depends on the federal funds rate and choose $\pi_1 = 0.25$ as in Section 7.2. The remaining parameters are chosen as described in Section 5. The results are shown in Figure E.7. It turns out that capturing this structural break through a latent regime shift improves the fit to the data. Similarly, our model can also be extended in the spirit of Dergunov et al. (2022) to capture potential structural

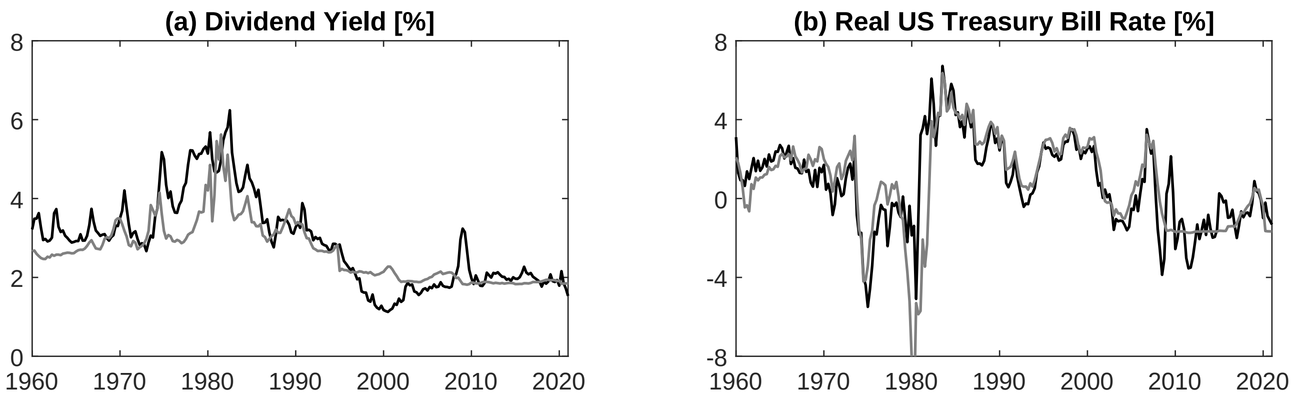

Figure E.7: Structural Break: Data versus Model Results. The figure depicts the historical and simulated (a) S&P500 dividend yield series and (b) real US Treasury bill rate. Black lines (——) show historical time series, while gray lines (——) present model simulation results for the model extension with the structural break in the dividend dynamics.

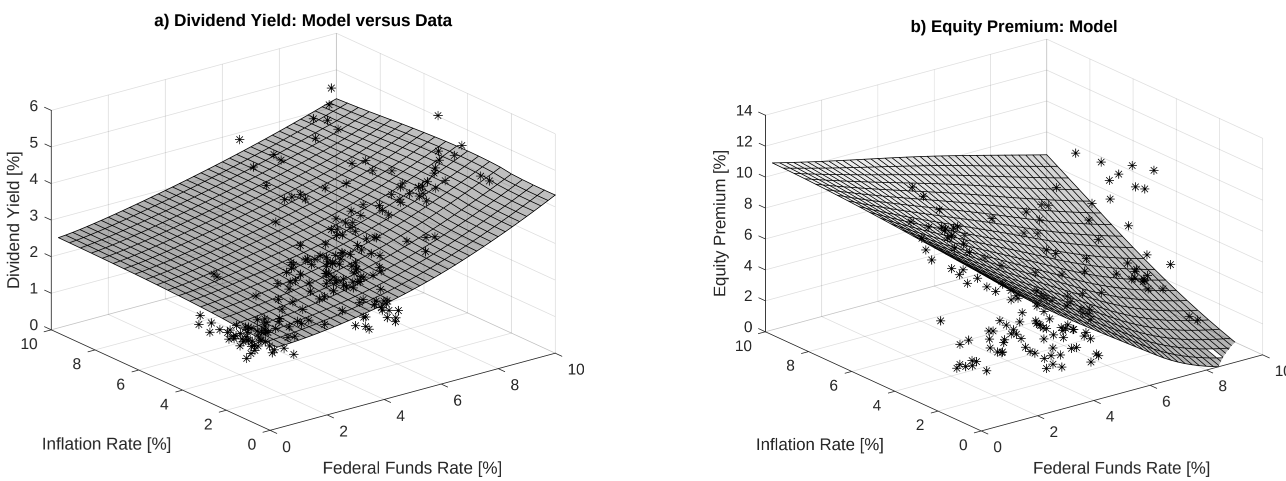

Figure E.8: Stochastic Disaster Intensity: Key Asset Pricing Moments. The figure depicts the historical and model-implied (a) dividend yield and (b) equity premium conditional on the inflation rate and federal funds rate. The gray surfaces represent the model implied results, while black stars visualize historical data. All quantities are given in nominal terms. Our sample period is from January 1960 until December 2020 and we use quarterly data. All results are annualized.

breaks in consumption growth and inflation dynamics, or in the spirit of Song (2017) to model different monetary policy regimes.

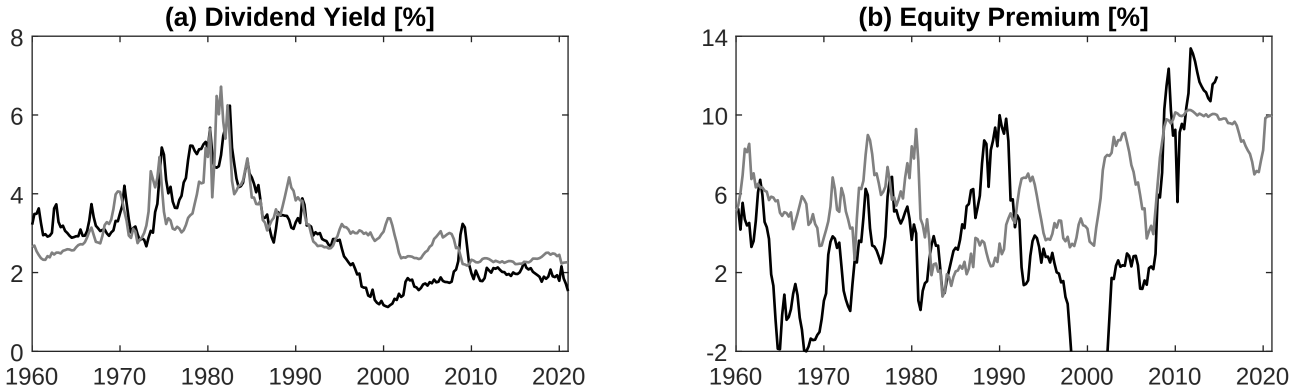

Figure E.9: Stochastic Disaster Intensity: Data versus Model Results. The figure depicts the historical and simulated (a) S&P500 dividend yield series and (b) real equity premium computed from a 10 year moving average of excess returns. Black lines (——) show historical time series, while gray lines (——) present model simulation results with a stochastic disaster intensity.

E.4 Model Extension: Stochastic Disaster Intensity

Our benchmark model assumes a constant disaster intensity λ_c. Now, we consider a model extension where the disaster intensity explicitly depends on the macroeconomic factors, i.e., $\lambda_c = \lambda_c(\pi, i)$. Such a model extension leads to stochastic disaster risk in the spirit of Wachter (2013). The analysis from Section 4 carries over to this more general case without affecting the analytical results. However, the effect of inflation and the federal funds rate becomes more pronounced compared to the benchmark case. In turn, the equity premium becomes more volatile, which is more in line with historical data.[34] To calibrate this model extension, we first determine how strongly the equity premium reacts to changes in inflation and the federal funds rate. For this purpose, we regress the estimated equity premium on inflation, the federal funds rate, and an interaction term, i.e.,

$$ep_t = \beta_0 + \beta_\pi \pi_t + \beta_i i_t + \beta_{i\times\pi} i_t \cdot \pi_t + \varepsilon_t,$$

leading to $\widehat{\beta}_0 = 0.0599^{***}$, $\widehat{\beta}_i = -0.8589^{***}$, $\widehat{\beta}_\pi = 0.1472$, $\widehat{\beta}_{i\times\pi} = 5.2244^{***}$, and $R^2 = 14.9\%$. Although the coefficient β_π is not significant, we want to stress that those regression results indicate a significant positive effect of inflation on the equity premium in line with the findings in Section 6. Setting the insignificant coefficient β_π to zero and re-estimating the regression yields $\widehat{\beta}_0 = 0.0634^{***}$, $\widehat{\beta}_i = -0.8688^{***}$, $\widehat{\beta}_{i\times\pi} = 6.2164^{***}$, and $R^2 = 14.7\%$. Next, we use these estimates to parametrize the disaster inten-

[34]The same mechanism holds true for stochastic volatility if we assume that the consumption volatility depends on the macroeconomic factors, i.e., $\sigma_c = \sigma_c(\pi, i)$.

sity function $\lambda_c(\pi, i)$.[35] If we stick to the benchmark parameter values from Section 5, we find that $\lambda_c(\pi, i) = 0.1521 - 2.09\,i + 14.92\,i \cdot \pi$ matches the time-variation in the historical price-dividend ratio and the equity premium reasonably well as shown in Figures E.8 and E.9. Overall, the model fit is good and the correlation between the historical equity premium and model-implied equity premium is 44%.

[35] Assuming that disaster risk is driven by an additional idiosyncratic Brownian motion as in Wachter (2013) could potentially explain the time-varying pattern of the equity premium and the stock market volatility even better but requires an additional state variable. We therefore refrain from this extension.