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ABSTRACT

We document that U.S. Treasury convenience moved positively with inflation during the inflationary second half of the 20th century but not before WWII or after 2000. A macro-asset pricing model explains this shift through two channels. Inflationary supply shocks raise the opportunity cost of holding money and money-like assets, endogenously increasing convenience yields. In contrast, exogenous liquidity demand shocks elevate convenience but depress consumption and inflation. Model estimates show that spikes in liquidity demand after 2000 account for the weaker convenience–inflation link and the emergence of negative bond-stock betas, distinguishing liquidity from non-liquidity demand shocks.

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1 Introduction

The relationship between liquidity, interest rates, and inflation was central to the great macroeconomic debates of the 20th century (Keynes (1937), Friedman (1969)). Today, it is again relevant due to renewed concerns about inflation and the status of U.S. Treasuries. Recent research on Treasury markets indicates that investors value U.S. Treasury securities more highly than assets with the same cash flows, meaning that Treasury bonds have convenience value (Longstaff, 2004; Du et al., 2018b; Krishnamurthy and Vissing-Jorgensen, 2012).

Motivated by the old and new debates on the value of the Treasuries, we study how Treasury convenience interacts with the macroeconomy and asset prices. We formulate and estimate a macro-finance model with convenience to address two complementary questions: (i) How do macroeconomic shocks shape Treasury convenience? (ii) How do convenience or liquidity shocks propagate to macroeconomic and financial fluctuations?

As the main stylized fact behind our analysis, we document secular shifts in the relationship between Treasury convenience and inflation. Convenience rose with inflation during the Great Inflation of the second half of the 20th century, in stark contrast to the early part of the 20th century and the post-2000s.[1] Our baseline empirical result shows that a one percentage point increase in quarterly headline CPI inflation is associated with a convenience yield that is about 12 bps higher during 1952–1999 compared to the pre-WWII or post-2000 periods, a magnitude that is large relative to an average convenience yield of 42 bps. In contrast, the loadings of convenience on inflation in the pre-WWII and post-2000 periods are much smaller, close to zero, and often negative.

The three distinct periods that we examine—pre-WWII, 1952–1999, and 2000–2020—were characterized by different macroeconomic drivers and changing properties of U.S. Treasury bonds (Campbell et al. (2020), Pflueger (2025)). The 1952–1999 period featured stagflations and positive nominal bond-stock betas, indicative of supply shocks. A switch to negative bond-stock betas post 2000 followed along with low-inflation recessions, reminiscent of the pre-WWII experience. Using a finer sample split by bond-stock betas, we find that the relationship between convenience and inflation is strongly positive in subsamples where the nominal bond-stock beta is high; otherwise, it is close to zero and slightly negative within the bottom quartile of bond-stock beta observations.

[1] We follow Nagel (2016) in measuring the convenience value of three-month T-bills as the spread between bankers' acceptance and T-bills, extended with the spread between asset-backed commercial paper over T-bills. A higher spread corresponds to a higher value of Treasury convenience.

Hence, the relationship between inflation and convenience shifted along with the structural forces in the U.S. economy, being positive under supply-driven, countercyclical-inflation regimes (like the 1970s) and near-zero or negative under demand-driven, procyclical-inflation regimes (pre-WWII; 2000–2020). As a distinct feature, both the early 20th century and the post-2000 period were punctuated by financial crises, experiencing spikes in Treasury convenience that coincided with disinflations.

To provide a structural interpretation of these empirical facts, we embed the convenience yield within a macro asset pricing model. The model combines the New Keynesian block for macroeconomy and monetary policy with habits in the utility and liquidity preferences, delivering joint implications for asset prices and convenience yields. Specifically, the framework isolates convenience dynamics arising endogenously from shocks in the macroeconomy and monetary policy vs. exogenous liquidity shocks that themselves can become a source of fluctuations. To demonstrate the role of convenience, we explicitly distinguish between two sources of demand shocks. A shock to preference for liquid Treasuries leads to a "liquidity demand" shock in the consumption Euler equation. Instead, a time-preference or taste shock for consumption today vs. next period gives rise to a traditional "non-liquidity" demand shock.

The model thus features four classic shocks: a supply-side cost-push shock, a time-preference demand shock, a monetary policy shock, and a liquidity demand shock. In this setting, the interactions between Treasury convenience and the economy emerge via two types of mechanisms, which we refer to as the "money channel" and the "liquidity demand channel." Broadly, these channels distill the Treasury convenience dynamics into a component derived from macroeconomic and monetary shocks and a component stemming purely from exogenous liquidity demand. The "money channel" operates by changing the opportunity cost of holding money—the nominal short-term interest rate—thereby also affecting the cost of holding money substitutes. We demonstrate that the information embedded in the convenience-inflation comovement and bond-stock betas is essential for identifying the underlying mechanism and its transmission.

Linking directly to our evidence, the model implies that volatile supply shocks give rise to a positive convenience-inflation relationship via the money channel. Intuitively, higher inflation leads to an increase in the nominal policy rate. A higher nominal interest rate drives up the opportunity cost of holding liquid deposits, raising households' willingness to pay for close substitutes, such as liquid Treasuries. Compared to assets with no liquidity benefits, the nominal rate on liquid Treasuries hence rises less, resulting in a larger convenience yield. Because deposit rates in the model adjust sluggishly over time (as in the data), convenience depends not just on the current policy rate but also on the long-term policy rate target, which moves closely with inflation via the monetary policy rule. Supply shocks can hence replicate the trinity of a negative inflation-output gap correlation, a positive nominal Treasury bond-stock beta, and a positive convenience-inflation relationship, observed in the 1952–1999 period.

Our framework further clarifies the distinct economic and asset pricing consequences of shocks emanating from the demand for liquidity vis-à-vis the more traditional time-preference shocks driving demand in New Keynesian models. Among all four shocks, only the liquidity demand shock generates a negative convenience-inflation relationship, a hallmark of liquidity distress events such as banking crises. A positive shock to liquidity demand raises the convenience yield and thus the borrowing rate for households relative to the policy rate. This reduces consumption and output and, through the Phillips curve, lowers inflation. Therefore, inflation and convenience yield co-move negatively, consistent with the post-2000 evidence. The liquidity demand channel is thus similar in spirit to Keynes (1937)'s argument viewing exogenous variation in liquidity preference as a powerful force that can precipitate or deepen an economic depression.

We estimate shock volatilities using a simulated method of moments separately for 1952–1999 and 2000–2020. The estimates imply a high volatility of supply shocks, substantial monetary policy shocks, and almost no demand shocks for the 1952–1999 period. This contrasts with the post-2000 period where the model estimates a high volatility of liquidity and non-liquidity demand shocks, a substantial volatility of monetary policy shocks, and little supply shocks.[2] The model succeeds at matching the regression coefficient of the convenience spread on inflation, the inflation-output gap correlation, and the stock market beta of a 10-year nominal Treasury bond for each period. It also generates a high equity premium, equity Sharpe ratio, and stock return predictability from the lagged price-dividend ratio.

The economic impact of a liquidity demand shock in the model is quantitatively large. A 100 bps increase in the convenience spread, roughly the spike during the 2008–2009 financial crisis, leads to a 0.5 percentage point fall in inflation, a 0.9 percentage point fall in the output gap, and a 0.8 percentage point decline in the policy rate, with responses peaking between 4 and 10 quarters after the shock. While an adverse non-liquidity demand shock can also generate a positive correlation between inflation and the output gap, its effect on convenience is very small. Re-solving for

[2]We refrain from estimating the model for the pre-WWII period due to the very different monetary policy mechanism during this period (e.g., Friedman and Schwartz (1963)).

the asset pricing equilibrium shows that monetary policy shocks or non-liquidity demand shocks cannot explain the confluence of negative nominal bond-stock betas, low-inflation recessions, and a significantly lower convenience-inflation coefficient, observed during 2000–2020. Overall, the model implies that the comovement between inflation and convenience yields is a highly informative moment for identifying the nature of demand shocks in the economy attributed to liquidity versus other sources.

Complementary to the convenience-inflation relationship, the model-implied bond-stock betas can also reveal the dominance of supply, liquidity or non-liquidity demand shocks. Model-based counterfactuals replicate the empirical finding that the convenience-inflation relationship is most positive when bond-stock betas are high. Additionally, liquidity shocks in the model drive the hedging properties of nominal Treasuries, depressing bond-stock betas through the liquidity demand channel.

Finally, we use data on convenience yields and inflation expectations for the post-pandemic period to validate the model predictions out of sample. Consistent with a substantial role for liquidity demand shocks during the first half of the COVID pandemic in early 2020, we find that the initial rise of convenience spreads coincided with disinflation. However, the convenience-inflation relationship turned positive once the initial pandemic shock resolved and inflation took over in mid-2021. The reemergence of the money channel in the face of supply shocks can explain a positive shift in the convenience-inflation comovement post-2020, which robustly appears across different measures of convenience and inflation expectations.

Related Literature

Our model builds on the New Keynesian literature allowing for an explicit role for liquidity preferences. Agents derive utility from a liquidity aggregate, where Treasuries are substitutes with deposits or non-interest-bearing money, similarly to money in the utility function (Sidrauski, 1967; Krishnamurthy and Vissing-Jorgensen, 2012; Nagel, 2016). The macroeconomic block is a parsimonious three-equation New Keynesian model following Galí (2008), Rotemberg and Woodford (1997), or Clarida et al. (1999).[3] Consumption utility takes the habit form of Campbell et al. (2020), generating reasonable asset prices with high and predictable stock returns, volatile risk premia, and a high Sharpe ratio. Different from the prior literature on bond-stock comovements,

[3]For models of banking and money within a New Keynesian economy, see also Curdia and Woodford (2010), Gertler and Karadi (2011a), Drechsler et al. (2018), and Wang (2025).

however, convenience in our model enters into the Euler equation because the interest rate faced by households differs from the policy rate (Piazzesi et al. (2019)). The liquidity demand shock is interpreted as a shock to the usefulness of Treasuries in providing liquidity. This assumption can be motivated by a higher value of public liquidity relative to private liquidity in crises (Holmström and Tirole, 1998), as increased liquidity demand often coincides with banking crises (Brunnermeier, 2009; Krishnamurthy and Muir, 2025), or concerns about tail risks (Caballero and Krishnamurthy, 2008).

While the money channel goes back at least to Friedman (1969), this block in our model builds closely on Nagel (2016). Nagel (2016) establishes the connection between the nominal short rate, the opportunity cost of holding money and money-like assets, and convenience yields. Our contribution is to further study how the money channel arises endogenously from the economy and a monetary policy reaction function, while also allowing for an independent role of exogenous liquidity shocks. Bianchi et al. (2025) also feature an exogenous convenience yield shock as a driver of business cycles and asset prices but do not focus on understanding the determinants of convenience or the changing inflation-convenience relationship. Del Negro et al. (2017) and Li (2024) argue that liquidity shocks were important in the global financial crisis. Our focus is different as we take a broader approach to quantify the importance of liquidity shocks versus other macroeconomic shocks.

Convenience yields have been central to a variety of questions in economics and finance. Du et al. (2018b) and Jiang et al. (2021) document that violations of the covered interest parity in foreign exchange markets are correlated with international perceptions of the U.S. Treasury convenience. Binsbergen et al. (2022) construct stock-option implied risk-free rates and find that monetary policy affects the convenience yield. Hébert et al. (2023) provide complementary evidence that the gap between the stock market-implied risk-free rate and government rates acts as a shifter in the Euler equation, akin to a demand shock. Li (2024) presents the convenience yield as a channel of how quantitative easing policies affect the banking sector and financial crises. Hartley and Jermann (2024) explain the pricing of U.S. Treasury floating rate notes through a model where money-like assets, including Treasuries, differ in their degrees of moneyness.

Complementary to our work, Acharya and Laarits (2025) show that the hedging properties of Treasury bonds, measured by the bond-stock covariance, can help explain variation in the level of convenience yield. Cieslak and Pang (2021) use bond-stock comovement as an informative sign restriction to isolate the hedging premium component of Treasury yields in a structural asset-pricing

VAR setting. Fu et al. (2025) argue for a negative correlation between Treasury convenience and inflation expectations whereby fiscally-driven inflation expectations reduce convenience of Treasuries. Focusing on the post-1982 sample, Fu et al. (2025) base their evidence on expected inflation estimates from the Cleveland Fed model. These estimates, being derived from yields, can confound inflation expectations and convenience. By contrast, we document how the convenience-inflation relationship changed over time due to shifting supply and demand forces. We model this change, and its link to bond-stock betas, in a New Keynesian-asset pricing model with liquidity. Our interpretation highlights liquidity demand shocks, rather than expected inflation, as a causal factor behind the negative inflation-convenience relationship in recent decades.

The remainder of the paper is structured as follows. Section 2 describes our empirical results. Section 3 presents the model. Section 4 presents the model estimation, impulse responses and counterfactuals. Section 5 discusses the evidence from the post-COVID period. Finally, Section 6 concludes.

2 A Century of Inflation and Treasury Convenience

In this section, we present our main motivating fact: the relation between inflation and the Treasury convenience spread changed significantly over the past decades. In particular, the comovement between inflation and convenience was positive during the inflationary second half of the 20th century, different from the periods before or after.

2.1 Data and Measurement

Our primary measure of the convenience yield is the T-bill spread which proxies for the extra yield on a less liquid instrument with the same cash flows as a corresponding T-bill. We construct the T-bill spread following Nagel (2016), which is the spread between the three-month bankers' acceptance rate and the three-month T-bill rate before 1990, and the spread between the three-month term repo rate collateralized by Treasuries and the three-month T-bill rate after 1990. Since the repo data used by Nagel (2016) end in 2011, we rely on the three-month asset-backed commercial paper (ABCP) rate to supplement the most recent period.[4] We refer to the concatenated series as

[4] For the post-2011 sample, we cross-checked three-month commercial paper rates against three-month repo rates from JP Morgan markets (proprietary data), and found that they are similar. For replicability, we use the publicly available data on commercial paper rates.

the T-bill convenience yield or the T-bill spread. Bankers' acceptance, repo backed by U.S. Treasuries, and asset-backed commercial paper each represent a short-term borrowing instrument that is less liquid than the corresponding Treasury. Each of these instruments is guaranteed not only by the borrower but also by the guaranteeing bank or extremely safe collateral, making credit risk negligible.[5] While measuring convenience is non-trivial, most other measures of convenience are not available for our full sample or even the post-1952 sample. We also discuss robustness using the long-term Aaa-Treasury spread following Krishnamurthy and Vissing-Jorgensen (2012).

As a baseline inflation measure, we use the quarterly rate of change in the headline consumer price index, which is available for our full sample from St. Louis FRED. We report quarterly inflation in annualized percentage points.[6]

We control for known drivers of Treasury convenience, as documented by earlier research, including stock market volatility, the short-term interest rate, and total government debt supply. For market volatility and the short rate, we use data from Nagel (2016) available through 2011, which we extend through 2020. Market volatility is measured with the monthly average VIX after 1990; before 1990, it is a backward-fitted linear projection of the VIX on realized volatility of market returns, where projection coefficients are estimated on the post-1990 data. We denote the spliced market volatility measure as VIX. The short-term interest rate is the monthly average effective federal funds rate (FFR) available from FRED starting in July 1954 and the Federal Reserve Bank of New York's discount rate before July 1954. We denote the spliced series as FFR. For government debt supply, we use the total quantity of Treasury debt, at market value, excluding intra-governmental holdings and holdings by depository institutions and the Federal Reserve. The data construction follows Krishnamurthy and Li (2023).

We also control for bond-stock comovement following Acharya and Laarits (2025). We estimate rolling bond-stock betas as the regression coefficient of daily seven-year zero-coupon Treasury returns onto daily stock returns over a 120-day rolling window. We use the seven-year Treasury yield from Gürkaynak et al. (2007) because it is available for a longer sample than the ten-year yield.

We consider three periods for our empirical analysis that capture distinct macroeconomic

[5] Appendix A.2 provides a back-of-the-envelope calculation for an upper bound for the credit risk in our series.

[6] We use the seasonally adjusted Consumer Price Index for All Urban Consumers (CPI-U) published by the Bureau of Labor Statistics and available from 1947. For the earlier part of the sample, we use the seasonally unadjusted CPI-U series. The FRED tickers are CPIAUCSL and CPIAUCNS, respectively. The seasonally unadjusted CPI-U is also used by Shiller (2016) to cover a long period starting from the late 1800s.

regimes: the pre-WWII period (1923:01 through 1939:08), the second half of the 20th century (1952:01 through 1999:12), and the post-2000 period until the end of 2020 (2000:01 through 2020:12). The beginning of our sample excludes the extreme inflation in 1920–1921 in the aftermath of WWI followed by the 1922 deflation, as these outcomes were driven by distinct factors related to post-war recovery (Reed, 2014). We similarly exclude the period around WWII until 1951. During the WWII period, inflation was hard to measure due to product price controls and wartime rationing.[7] The sample ends in 2020, before the post-COVID inflation period, which we analyze separately in Section 5.

We set a break date for our analysis in January 2000, because the literature has argued that the nature of economic shocks changed from dominant supply to demand shocks around that time (e.g., Campbell et al. (2017), Stock and Watson (2007)). In particular, Campbell et al. (2020) estimate a break date in the relationship between inflation and the output gap around 2000, indicating that the economy changed from stagflationary recessions to low-inflation recessions. Table A1 in Appendix A.1 contains summary statistics for our key variables. Average inflation is 3.9% in the second half of the 20th century, which encompasses the high-inflation 1970s and 1980s, but is much lower in the other periods. The T-bill convenience is 42 bps on average, with a standard deviation of 46 bps. While our two main variables have some serial correlation, they are not overly persistent. The 12-month AR(1) for quarterly inflation is 0.27 and for the T-bill convenience yield is 0.54. These are substantially below the value of 0.95 that Stambaugh (1999) emphasizes as problematic.

2.2 The Changing Treasury Convenience-Inflation Relationship

Figure 1 illustrates the shifts in the correlation between inflation and the T-bill spread over time: from a negative correlation of –0.34 before WWII, to a strong positive correlation of 0.63 in the latter half of the 20th century, and down to zero after 2000.

To assess the statistical and economic significance of the changing convenience-inflation relationship, we estimate the following regression at a monthly frequency:

$$\textit{T-bill spread}_t = b_0 + b_1\pi_t + b_2\pi_t \times I_{\text{1952-1999},t} + b_3\pi_t \times I_{\geq 2000,t} + \Gamma' X_t + \varepsilon_t, \tag{1}$$

where we interact the inflation rate with period-specific dummy variables. The interaction coeffi-

[7] The Treasury-Federal Reserve Accord on March 4, 1951 ended the wartime interest rate controls and signaled a broader return to a non-wartime economy.

cients are interpreted relative to the pre-WWII period (the omitted category). π_t is the quarterly inflation rate from month $t-3$ to t in annualized units, $\pi_t = 400 \times (\frac{CPI_t}{CPI_{t-3}} - 1)$. The vector X_t represents time t controls.

Figure 1. T-bill convenience and inflation. We plot the T-bill convenience yield (constructed following Nagel (2016)), and quarterly inflation during three subperiods: 1923:01-1939:08, 1952:01-1999:12, and 2000:01-2020:12. The data frequency is monthly. In each subperiod, we normalize both series to have zero mean and unit standard deviation.

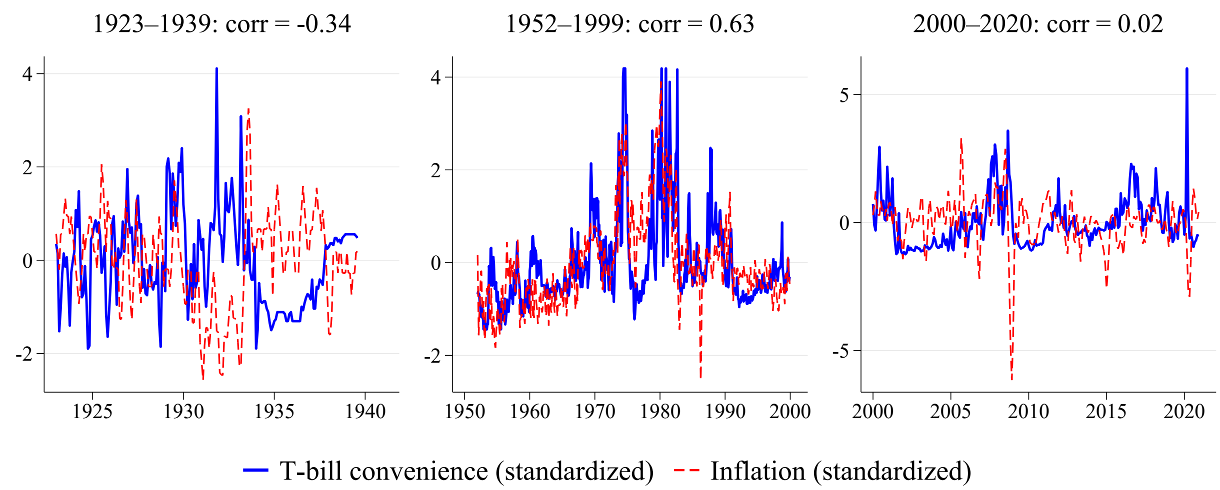


Table 1 shows that the interaction between inflation and the 1952–1999 dummy enters consistently positively. Thus, the relationship between inflation and the convenience spread is significantly stronger during the second half of the 20th century, which includes the Great Inflation of the 1970s and 1980s, than in the other periods. The estimates in column (1) imply that a one-percentage-point increase in inflation is associated with a 12 bps higher T-bill spread in the 1952–1999 period compared to pre-WWII. This magnitude is large relative to an average T-bill spread of 42 bps. Said differently, a one-standard-deviation increase in inflation corresponds to a 1.1 standard-deviation increase in the T-bill spread over 1952–1999, using full-sample standard deviations. By contrast, the negative baseline coefficient on inflation means that a one-percentage-point increase in inflation tends to be associated with a 1.4 bps decrease in the T-bill spread during the pre-WWII period. The relationship changes again during the 2000s, when the T-bill spread comoves negatively with inflation around the global financial crisis, and the average correlation is close to zero.

The shifts in the convenience-inflation relationship are robust to controlling for a comprehen-

Table 1. Shifts in T-bill convenience–inflation relationship. Monthly data runs from 1923:01 through 2020:12, excluding the 1939:09–1951:12 period. $I_{1952-1999}$ and $I_{\geq 2000}$ are dummy variables taking the value of one in the indicated subperiod. The pre-WWII period (1923:01–1939:08) is the omitted category. Newey-West t-statistics with 12 lags are shown in parentheses. The stars indicate significance at * p <0.10, ** p <0.05, *** p <0.01 levels.

	T-bill spread				
	(1)	(2)	(3)	(4)	(5)
Inflation	-0.015***	-0.015***	-0.013***	-0.0097***	-0.010***
	(-3.87)	(-4.62)	(-3.53)	(-3.93)	(-3.74)
Inflation x $I_{1952-1999}$	0.12***	0.059***	0.11***	0.053***	0.053***
	(7.93)	(3.71)	(7.73)	(3.74)	(3.54)
Inflation x $I_{\geq 2000}$	0.016**	-0.00064	0.0097	0.0089	0.0090
	(2.29)	(-0.11)	(1.41)	(1.27)	(1.28)
FFR		0.081***		0.082***	0.083***
		(8.11)		(8.15)	(8.07)
Debt/GDP			-0.35	0.20	0.20
			(-1.61)	(0.96)	(0.96)
VIX				0.010***	0.011***
				(4.81)	(3.54)
Baa spread					-0.012
					(-0.27)
$I_{1952-1999}$	-0.10	-0.11*	-0.052	-0.050	-0.060
	(-1.56)	(-1.87)	(-0.73)	(-0.73)	(-0.72)
$I_{\geq 2000}$	-0.022	0.10*	0.092	0.068	0.061
	(-0.46)	(1.94)	(1.01)	(0.84)	(0.71)
Constant	0.25***	0.026	0.32***	-0.27***	-0.26***
	(6.33)	(0.56)	(5.92)	(-3.16)	(-3.02)
$\bar{R}^2$	0.43	0.56	0.43	0.59	0.59
N	1028	1028	1028	1028	1028

sive set of variables suggested by earlier work on the determinants of Treasury convenience. We first include the FFR to directly proxy for the opportunity cost of money. Consistent with Nagel (2016), we find a strong positive comovement between T-bill spread and the FFR. The FFR absorbs about half the magnitude on the interaction between inflation and the 1952–1999 dummy, indicating that the opportunity cost of money plays a role for the positive relationship during this period. At the same time, the interaction between inflation and the 1952–1999 dummy remains positive and significant, indicating that the FFR might not fully control for the opportunity cost of holding liquid assets. In our model, we capture this fact through inertial deposit rates and the central bank's tendency to change interest rates in the direction of inflation. Controlling for the FFR in column (2), the post-2000 convenience-inflation relationship is indistinguishable from the pre-WWII period relationship, which is statistically and economically significantly negative.

Column (3) in Table 1 shows that our main finding is robust to controlling for the quantity of

Treasury debt. Prior research has found that an increase in the quantity of Treasury debt tends to reduce the convenience yield, especially on long-term Treasury bonds (Krishnamurthy and Vissing-Jorgensen, 2012; Krishnamurthy and Li, 2023). Another reason to control for the debt/GDP ratio is if fiscal policy is a driver of inflation (Cochrane, 2001; Corhay et al., 2023).[8] Column (3) shows that debt/GDP enters negatively with a coefficient of -0.35, roughly half the magnitude reported by Krishnamurthy and Vissing-Jorgensen (2012) for long-term bonds. The statistical significance of the debt/GDP variable is reduced because inflation and debt/GDP are negatively correlated in our sample. Our results also differ because we use T-bill convenience on the left-hand side, whereas Krishnamurthy and Vissing-Jorgensen (2012) use a measure of long-term convenience. Nagel (2016) shows that, after adjusting for the opportunity cost of holding money, debt quantity loses its explanatory power for the short-term convenience spread. We reproduce this finding in columns (4) and (5), again with little impact on the convenience-inflation relationship, which is at the center of our analysis.

The results also hold when we include other potential drivers of Treasury convenience in columns (4) and (5). T-bill convenience loads positively on equity volatility, which Nagel (2016) includes as a proxy for liquidity demand, but the coefficients on inflation and its interactions remain broadly unchanged compared to column (2). Controlling for the Baa-Aaa credit spread does not add independent explanatory power. This latter finding is consistent with the T-bill spread being largely immune to time-varying credit risk, as supported by our additional analysis in Appendix A.2. Appendix A.3 further reports regressions that control for the Baa-Aaa credit spread interacted with period dummies, again finding that our baseline results are robust. The switch in the inflation-spread relationship is thus specific to the convenience premium in Treasuries, as distinct from how inflation comoves with credit conditions (e.g., Kang and Pflueger (2015), Brunnermeier et al. (2025), Bhamra et al. (2023)) or with the market volatility.

We use the quarterly inflation from month $t-3$ to t for consistency with the model in Section 3 that runs at quarterly frequency. In Appendix A.4, we show that the changing inflation-convenience relationship is similar or even more pronounced using alternative definitions of inflation, including forward- versus backward-looking measures and annual versus quarterly horizons.

[8]Different from Fu et al. (2025), we follow Krishnamurthy and Vissing-Jorgensen (2012) in our construction of Debt/GDP and do not linearly de-trend.

2.3 Empirical Results Starting in 1952

We now report robustness results for a shorter sample starting in 1952. While the early 20th century provides a useful laboratory that informed the development of liquidity preference theories (Keynes, 1937), the conduct of monetary policy and data availability differ markedly from the rest of our sample (Romer and Romer, 2002; Asso et al., 2007). We therefore report regressions for the post-1952 period, which also forms the basis of our model estimation.

Table 2 treats the 1952–1999 period as the baseline. Column (1) confirms a consistently positive comovement between inflation and the T-bill spread during the second half of the 20th century, with magnitudes consistent with the full sample estimates in Table 1. The negative coefficient on inflation interacted with the post-2000 dummy shows that the relationship between convenience and inflation is statistically significantly lower after 2000 and close to zero. Column (2) shows that controlling for the FFR again absorbs about half the positive convenience-inflation relationship for the 1952–1999 period. Column (3) controls for debt/GDP, credit risk, and the VIX. In column (4), we show that the relationship is robust to including a 1979–1980 dummy to account for the highly volatile monetarist experiment. Moreover, this positive convenience-inflation relationship in the middle period is robust to alternative break dates, as shown in Appendix A.6. Figure 1 also illustrates that the positive inflation-convenience comovement is a broader feature of the second half of the 20th century, and not just the monetarist experiment.

Using a shorter sample starting in 1961, column (5) controls for rolling bond-stock betas. Recent work by Acharya and Laarits (2025) demonstrates that Treasury convenience reflects the hedging premium of Treasury bonds. Specifically, they show that convenience spreads are high when the covariance between aggregate equity and Treasury bond returns is low, as investors are willing to pay a premium for non-pecuniary liquidity attributes of the Treasuries in bad times. We construct a backward-looking 120-day stock-bond beta using daily CRSP stock market returns and seven-year zero-coupon bond returns from Gürkaynak et al. (2007).[9] The sample begins in 1961:11 when daily bond data become available. We find that the loading on the stock-bond beta is significantly negative, consistent with a hedging premium component of convenience as emphasized by Acharya and Laarits (2025). At the same time, the stock-bond beta does not affect the positive relationship between inflation and convenience before 2000 or the decline in this relationship post 2000, suggesting that the inflation-convenience comovement is distinct from the hedging premium

[9] We use the seven-year zero-coupon bond since it is the longest maturity consistently available from 1961:11 onward. The results are robust to using the ten-year zero-coupon bond after 1972:02.

Table 2. Regressions with different measures of convenience yield, post 1952. The table reports regressions of short- and long-term convenience spreads on inflation and controls. Columns (1)–(4) estimate specifications from Table 1 over a shorter 1952–2020 sample period. Column (4) controls for the high volatility period, including a dummy variable $I_{1979-80}$ equal to one in 1979 and 1980. Column (5) controls for the rolling stock-bond beta following Acharya and Laarits (2025). Columns (6) and (7) use proxies for long-term convenience as the dependent variable: the Moody's Aaa-Treasury spread and the Aaa-GSW spread based on the long-term GSW par yield. The Aaa-GSW (ca) spread is adjusted for the time-varying moneyness of call options in corporate bonds following Gilchrist and Zakrajšek (2012) and Duffee (1998). The sample period, depending on data availability, is indicated in column headers. Newey-West t-statistics with 12 lags are shown in parentheses. * p <0.10, ** p <0.05, *** p <0.01.

	1952:01-2020:12				1961:11-2020:12		
	(1) Tbill spr	(2) Tbill spr	(3) Tbill spr	(4) Tbill spr	(5) Tbill spr	(6) Aaa-Tsy	(7) Aaa-GSW (ca)
Inflation	0.10***	0.043***	0.043***	0.049***	0.040***	0.057***	0.017
	(7.20)	(2.73)	(3.04)	(2.84)	(2.73)	(3.48)	(1.13)
Inflation x $I_{\geq 2000}$	-0.10***	-0.059***	-0.042**	-0.048**	-0.060***	-0.067***	-0.026*
	(-6.62)	(-3.74)	(-2.55)	(-2.53)	(-4.05)	(-3.69)	(-1.71)
FFR		0.083***	0.082***	0.082***	0.10***	-0.014	0.011
		(7.86)	(7.14)	(7.03)	(8.60)	(-0.72)	(0.60)
Debt/GDP			0.24	0.25		-0.25	0.012
			(1.08)	(1.13)		(-0.66)	(0.04)
VIX			0.012***	0.012***		0.0033	0.0068
			(2.62)	(2.68)		(0.73)	(1.58)
Baa spread			-0.0045	-0.0027		0.44***	0.18*
			(-0.05)	(-0.03)		(4.51)	(1.95)
$I_{1979-80}$				-0.18		-0.44***	-0.41***
				(-0.88)		(-3.46)	(-3.10)
Stock-bond beta					-0.55***	-0.30	-0.16
					(-2.99)	(-1.63)	(-1.06)
$I_{\geq 2000}$	0.082	0.22***	0.10	0.12	0.19***	0.27*	0.18
	(1.37)	(3.39)	(1.22)	(1.33)	(2.98)	(1.81)	(1.39)
Constant	0.15***	-0.094	-0.35***	-0.37***	-0.16***	0.36*	0.43***
	(2.79)	(-1.55)	(-3.10)	(-3.01)	(-2.66)	(1.66)	(2.62)
$\bar{R}^2$	0.43	0.57	0.60	0.60	0.61	0.41	0.24
N	828	828	828	828	710	710	710

channel.

We next show that the relationship between long-term convenience spreads and inflation has also shifted over time. Existing studies of long-term convenience following Krishnamurthy and Vissing-Jorgensen (2012) rely primarily on the Aaa-Treasury spread provided by Moody's. However, in the first half of our sample, the Moody's Aaa-Treasury spread is confounded by multiple other factors unrelated to convenience: the flower bond clauses in Treasury bonds (Lehner et al., 2025), the widespread callability of corporate bonds (Duffee, 1998), and potential duration mismatch between the Treasuries and corporate bonds (van Binsbergen et al., 2025). To address these confounders, we construct the Aaa-GSW (call adjusted) spread, detailed in Appendix A.8. Specifically, we use the long-term GSW par yield, which excludes Treasuries with option features (callable and flower bonds) and closely matches the duration of Aaa corporate bonds in the Moody's index. To account for the time-varying moneyness of call options in corporate bonds, we follow Gilchrist and Zakrajšek (2012) and Duffee (1998) by orthogonalizing the Aaa–GSW spread with respect to the yield curve level, slope, and interest rate volatility.

Figure 2 plots the times series of T-bill and long-term convenience against inflation starting in 1952. The long-term convenience spreads, shown in Panels B and C, start in 1961 when GSW yield curves become available.[10] Looking across panels, we see that all measures of convenience are positively correlated with inflation in the 1952–1999 period. Long-term convenience is negatively correlated with inflation for the 2000–2020 period, while the correlation with short-term convenience is essentially zero, as in Figure 1. Figure 2 also reveals important differences between the long- and short-term convenience dynamics, especially during the monetarist experiment 1979–1980. This is perhaps not surprising given the coincidence of large monetary and cost-push shocks and the unprecedented interest rate volatility in this period.

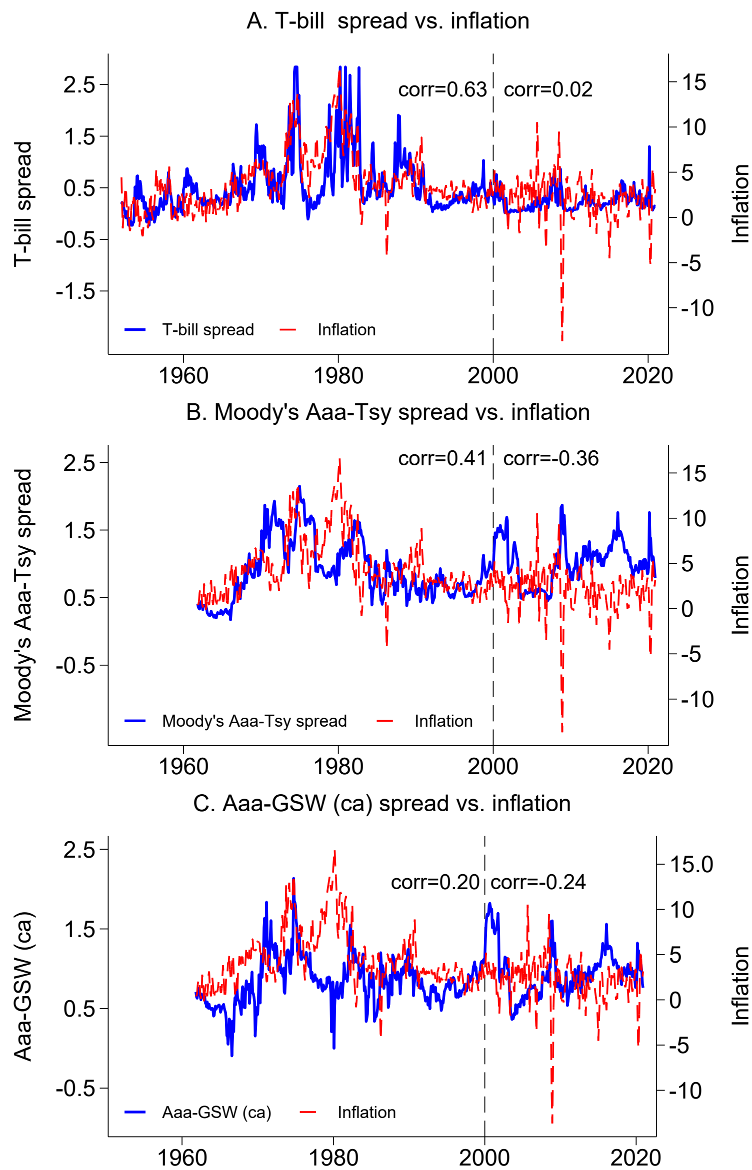
Table 2, columns (6) and (7) report the regressions for long-term convenience with our full set of controls. Column (6) shows that the Moody's Aaa-Treasury spread was positively correlated with inflation 1961–1999, but significantly negatively correlated 2000–2020. Column (7) shows that the Aaa-GSW (ca) spread has a still positive but statistically insignificant relationship for the 1961–1999 period, and a slightly negative relationship for the 2000–2020 period. Given the pervasive data concerns, such as maturity mismatches, callability, etc, present in long-term convenience proxies, we use the T-bill convenience spread as our benchmark in the main analysis. Nevertheless, measures of long-term Treasury convenience also indicate a shift in the relationship with inflation around 2000.

2.4 Macroeconomic Drivers According to Standard Moments

So far, we have seen that the convenience-inflation relationship changed in the second half of the 20th century and again around 2000. We now discuss moments for bond-stock betas and the inflation-output gap correlation, which we will also use to estimate our model, in Table 4. We present additional results on inflation-convenience comovement conditioning on bond-stock betas

[10]We start the Aaa-Tsy spread in 1961 for consistency with the Aaa-GSW (call adjusted) spread. Results are similar if we start in 1952.

Figure 2. Short- and long-term convenience spreads vs. inflation. The figure superimposes proxies for long- and short-term convenience against quarterly inflation, with the left axis showing various convenience yields and the right axis showing inflation, both in percentage points. The spreads are T-bill spread (Panel A), the Moody's Aaa-Treasury spread (Panel B), and the Aaa-GSW spread adjusted for the time-varying moneyness of call options in corporate bonds (Panel C). The sample starts in 1952:01 in Panel A and in 1961:11 in Panels B and C, and runs through 2020:12.



as an alternative sample split. All moments in this subsection are run on quarterly (quarter-end) data. While moving to a quarterly frequency reduces the number of observations available, it facilitates the match into our quarterly model.

During the second half of the 20th century, recessions were accompanied by high inflation. These "stagflations" are often taken as the hallmark of inflationary supply shocks, and indeed in the data the correlation between inflation and the output gap was significantly negative for 1952-1999 (see Table 4). At the same time, the stock market beta of 10-year nominal Treasury bonds was positive during this period at 0.215, as one should expect if bonds suffer from stagflationary dynamics along with stocks (Campbell et al. (2020)). Conversely, during the post-2000 period, the inflation-output gap correlation was significantly positive. As a case in point, inflation was low and even turned into a brief deflation during the financial crisis of 2008-2009. Concurring with this switch in the inflation-output gap correlation, the 10-year nominal Treasury bond-stock beta was economically and statistically significantly negative, as one would expect if the real cash flows of nominal bonds benefit from low inflation news just as the stock market falls.

Table 3 estimates the convenience-inflation relationship in subsamples directly determined by rolling nominal bond-stock betas as a financial market indicator of dominant supply vs. demand shocks. For each subsample, we regress the T-bill convenience spread onto quarterly inflation, as before. Moving from the leftmost to the rightmost column shows that the T-bill spread-inflation relationship changes from slightly negative to strongly and significantly positive in samples with a higher nominal Treasury bond-stock beta. For the sample where the rolling nominal Treasury bond-stock beta is above the median, a one percentage point increase in inflation is associated with a 10 bps increase in convenience, similar to the 1952-1999 coefficient in column (1) of Table 2. Conversely, when the rolling bond-stock beta is in its bottom quartile, the convenience-inflation relationship is significantly lower, close to zero, and even slightly negative, similar to the post-2000 period relationship in column (1) of Table 2. These empirical results are hence consistent with the interpretation that the shifting convenience-inflation relationship is linked to the dominance of macroeconomic supply shocks in the second half of the 20th century vs. demand shocks post-2000.

We regard the negative convenience-inflation relationship during the Great Depression also as broadly supportive of the role of liquidity-driven demand shocks for a negative convenience-inflation relationship. The Great Depression was a period of financial crises, bank runs, and deep deflation, just as the U.S. faced its most severe non-wartime recession. The Great Depression hence fits into the pattern of deflationary recessions and a negative convenience-inflation relationship,

Table 3. Sample split by bond-stock betas. This table estimates regressions of the T-bill spread onto annualized quarterly inflation. The sample is quarterly from 1961:Q4 to 2020:Q4. We use the daily seven-year nominal Treasury bond returns based on zero-coupon yields from Gürkaynak et al. (2007) to estimate a 120-day rolling bond-stock beta, available starting in 1961 (this availability restricts the sample starting date), as in Table 2 columns (5)–(7). Column (1) uses only observations when the bond-stock beta is within its bottom quartile. Columns (2), (3), and (4) use observations where bond-stock beta is below its median, above its median, and within its top quartile, respectively. Robust t-statistics are reported in parentheses. * $p< 0.10$, ** $p<0.05$, *** $p<0.01$.

	T-bill spread			
	(1)	(2)	(3)	(4)
	Beta Q1	Beta $<$ Med	Beta $>$ Med	Beta Q4
Inflation	-0.00528	0.0598**	0.104***	0.101***
	(-0.62)	(2.23)	(5.73)	(4.46)
Constant	0.283***	0.236***	0.166*	0.319***
	(7.61)	(3.89)	(1.96)	(3.09)
$\bar{R}^2$	-0.0131	0.159	0.311	0.307
N	59	119	120	61

similar to the post-2000 period. It is likely no coincidence that Keynes developed his "General Theory" during this period (Keynes (1937)). Unfortunately, many features of the economy were substantially different during the Great Depression from the modern day. These differences include measurement of output and especially potential output, the fact that Treasury bonds were initially still gold-linked, and the very different conduct of monetary policy. We hence only estimate our model for the two subperiods after 1952. We next turn to describing our model.

3 Model

This section describes the macro asset pricing model with liquidity preferences which explains the shifting inflation-Treasury convenience relationship via the competing "money" and "liquidity demand" channels.

3.1 Household Problem

3.1.1 Liquidity Aggregate

There are three different short-term interest rates, allowing us to model the spread between liquid and less liquid interest rates. We use I_t^l to denote the interest rate on illiquid loans (hence, superscript l), I_t^b to denote the interest rate on liquid Treasury bonds, and I_t^d to denote the interest rate on liquid deposits. The deposit rate represents the interest rate that households can earn by depositing money with a bank, i.e., the most liquid and money-like asset. Log interest rates are related to level interest rates via $i_t^l = \log\left(1 + I_t^l\right)$ and analogously for i_t^b and i_t^d. Unless noted otherwise, we use lowercase letters to denote logs throughout.

Different from the basic New Keynesian model, households have direct preferences over liquidity, similar to money in the utility function (Sidrauski (1967)) and the work by Krishnamurthy and Vissing-Jorgensen (2012). We assume that the liquidity aggregate Q_t is a composite of deposits and convenient government bonds

$$Q_t \quad = \quad D_t + \frac{\lambda_t}{1-\lambda_t} B_t. \tag{2}$$

Here, D_t denotes the real balance of zero-coupon bank deposits, and B_t is the sum of real balances of short- and long-term zero-coupon Treasury bonds. The parameter λ_t controls the contribution of government bonds to the liquidity aggregate. We assume that in steady state Treasury bonds offer positive liquidity but less than deposits, $0 < \bar{\lambda} < \frac{1}{2}$. A spike in λ_t can be interpreted as heightened uncertainty in the economy (Caballero and Krishnamurthy (2008)), tightened collateral constraints (Del Negro et al. (2017)), or a liquidity shock in the financial sector (Li (2024)), all of which would increase the preference for government debt relative to less liquid instruments.[11] We refer to λ_t as Treasury liquidity.

Following Nagel (2016), we consider the case of perfect substitutability between Treasury bonds and deposits for our calibration. However, this assumption is not crucial and the qualitative implications are similar if Treasuries and deposits are highly but not perfectly substitutable, as estimated by Krishnamurthy and Li (2023).[12] Piazzesi et al. (2019) provide a microfoundation

[11] For models that use similar shocks to the demand of liquid assets, see Anzoategui et al. (2019), Jiang et al. (2024), Itskhoki and Mukhin (2021), Kekre and Lenel (2024), Fukui et al. (2023), Bianchi et al. (2025), Engel and Wu (2023), Abadi et al. (2023).

[12] Appendix B.9 considers an extension to imperfect substitutability between deposits and Treasuries, and shows

for the complementarity between money and Treasuries for liquidity, where Treasuries serve as collateral for intermediaries.

3.1.2 Preferences

We follow the classic literature going back to Sidrauski (1967) and assume a simple separable utility function over consumption and liquidity

$$E_0 \sum_{t=0}^{\infty} \beta^t U\left(C_t, Q_t, H_t\right), \tag{3}$$

where

$$U\left(C_t, Q_t, H_t\right) = \frac{\left(C_t - H_t\right)^{1-\gamma}}{1-\gamma} + \alpha \log Q_t. \tag{4}$$

Here, C_t denotes market consumption, H_t denotes external habit over market consumption, and Q_t is the liquidity aggregate held by the representative household. Household subscripts are suppressed for conciseness. Modeling money in the utility function is the simplest way to capture the classic Friedman insight that the interest rate is the opportunity cost of holding money. Separable preferences over consumption and liquidity maintain the standard monetary transmission mechanism from the textbook New Keynesian model, and allow us to innovate on the consumption side of preferences.[13]

We extend the habit framework by combining preferences over the liquidity aggregate with the macro-asset pricing habit preferences of Campbell et al. (2020), which are known to account for a number of empirical features in macroeconomic dynamics, monetary policy, and asset prices. Relative risk aversion equals $-U_{CC}C/U_C = \gamma/S_t$, where surplus consumption is the share of market consumption available to generate utility $S_t = \frac{C_t - H_t}{C_t}$. The real consumption-based stochastic discount factor, pricing all real assets that have no special liquidity, such as illiquid loans or stocks,

that in that case, liquidity demand shocks need to be interpreted more broadly as incorporating shocks to the quantity of Treasury bonds. We include the debt/GDP ratio to control for this possibility in our empirical analysis.

[13]Galí (2008, Chapter 2.5) shows that non-separable preferences over consumption and liquidity can lead to the counterfactual implication that inflation rises following a contractionary monetary policy shock, while Woodford (2003b, Chapter 3.4) shows that for some quantifications the monetary policy transmission mechanism remains very similar to the separable case.

equals:

$$M_{t+1} = \beta \exp\left(-\gamma\left(\Delta c_{t+1} + \Delta s_{t+1}\right)\right). \quad (5)$$

Following Campbell et al. (2020), consumption habit H_t is assumed to be external, i.e., each household h takes habit as externally given. A model for habit implies a model for surplus consumption and vice versa. Log surplus consumption is assumed to follow:

$$s_{t+1} = (1-\theta_0)\bar{s} + \theta_0 s_t + \theta_1 x_t + \theta_2 x_{t-1} + \eta_t + \omega(s_t)\varepsilon_{c,t+1}, \quad (6)$$

where $\bar{s}$ is the steady-state value for log surplus consumption, x_t is the log output gap, i.e., log real output relative to log output without price-setting frictions, and $\varepsilon_{c,t+1} \equiv c_{t+1} - E_t c_{t+1}$ is the consumption surprise, which in equilibrium is an endogenous function of the fundamental shocks.

Most terms in the log surplus consumption dynamics (6) are standard. The term $\theta_0 s_t$ introduces persistence into log surplus consumption and implies that log habit approximately follows an exponentially-weighted moving average of past consumption. The term $\theta_1 x_t + \theta_2 x_{t-1}$ shifts log habits towards the most recent consumption lag, which is linked to the output gap x_t in equilibrium. Finally, surplus consumption is conditionally perfectly correlated with surprises to consumption $\varepsilon_{c,t+1}$, as it should if habit is pre-determined to first order.

The relationship between log surplus consumption and consumption surprises is heteroskedastic, with the sensitivity $\omega(s_t)$ downward-sloping and taking the functional form introduced by Campbell and Cochrane (1999) and shown in Appendix B.2.1. The downward-sloping sensitivity function captures the intuition that a negative consumption surprise is more painful and has a larger impact on marginal utility when consumption is already close to habit. The specific functional form of the sensitivity function ensures that the precautionary savings motive cancels exactly against the influence of s_t on the desire to smooth intertemporally, leading to linear policy rate dynamics that are independent of surplus consumption. By contrast, stocks and other long-term financial assets are highly non-linear in the state variable s_t.

The intertemporal substitution shock, η_t, in equation (6) is new and allows us to distinguish between liquidity and non-liquidity demand shocks. It introduces a simple way for intertemporal substitution incentives to vary over time, while preserving the linearity of the interest rate dynamics. From equation (5), it is clear that a predictable upward shift in Δs_{t+1} is economically similar to a predicted decline in the discount factor β, raising marginal utility today vs. marginal utility tomorrow. We model this impatience shock directly in surplus consumption dynamics to avoid introducing another state variable. Intuitively, a positive shock to η_t means that households expect better conditions tomorrow, leading them to wish to consume more today. Discount rate shocks are often used as a convenient way of introducing a demand shock in New Keynesian models. Correspondingly, η_t serves as the source of the non-liquidity demand shock in our New Keynesian block. Some prior asset pricing studies have examined the implications of "discount rate shocks" for stocks and bonds, though not usually within a model of monetary policy.[14]

3.1.3 Deposits and Monetary Policy

We assume that the deposit rate I_t^d is sticky and partially adjusts to the illiquid loan rate I_t^l,

$$I_t^d = \delta I_t^l + \rho^d I_{t-1}^d. \tag{7}$$

Parameter δ captures the imperfect pass-through of the illiquid rate to the deposit rate. The sticky adjustment, reflected in $0 < \rho^d < 1$, is a common feature of bank deposit rates in the data. We discipline the δ and ρ^d parameters using data in Section 4.2. We impose that $\delta/(1-\rho^d) < 1$, so that the steady-state deposit rate is below the steady-state loan rate, an important regularity that generates a positive convenience yield in steady state. The sluggish adjustment through ρ^d could arise from inertia in investor attention, while partial updating ($\delta/(1-\rho^d) < 1$) reflects bank market power.[15] In the special case where $\delta = \rho^d = 0$, deposits in the model can be interpreted as cash that carries a liquidity benefit but earns no interest.[16]

In our model, the central bank conducts monetary policy by setting the liquid bond rate, I_t^b, implicitly choosing deposit quantities to satisfy households' money demand function.[17] In practice, the Fed is not allowed to operate directly through private loan markets, since deciding which

[14] See e.g., Albuquerque et al. (2016), Gomez-Cram and Yaron (2021), and Gormsen and Lazarus (2025).

[15] A long-standing and growing literature has documented the presence of bank market power, see e.g. Barro and Santomero (1972); Startz (1979); Drechsler et al. (2017); Egan et al. (2022); Wang et al. (2022). Deposit rate sluggishness and depositor inattention are also well documented in the banking literature, e.g. Hannan and Berger (1991); Neumark and Sharpe (1992); Egan et al. (2025).

[16] As long as equation (7) is assumed to hold exactly, as shown in Appendix B.9.2, shocks to overall liquidity preference, α in equation (4), cannot be interpreted as a shock to the overall demand for liquidity. However, disturbances to the deposit rate pass-through (7) would act analogously to λ_t.

[17] If banks face a constant reserve requirement, Nagel (2016) shows that the implicit rule for deposits can be met by increasing or decreasing the amount of federal funds in the system, similarly to how the Fed operated for much of our sample period until the global financial crisis of 2008-2009.

borrower is creditworthy would be considered fiscal policy and outside the purview of the central bank. Specifying monetary policy in terms of a liquid policy rate is also consistent with how monetary policy was conducted throughout almost all of our sample.[18]

We assume that the policy rate follows a log-linear Taylor (1993)-type rule with inertia (Woodford (2003b))

$$i_t^b = (1-\rho^i)\underbrace{(\gamma^x x_t + \gamma^\pi \pi_t)}_{\text{Target Rate}} + \rho^i i_{t-1}^b + v_{i,t}. \tag{8}$$

Policy moves the policy rate towards a long-term target rate, which increases with inflation π_t and the log output gap x_t. A higher inertia parameter ρ^i implies that monetary policy raises interest rates slowly in response to an increase in inflation, as the short-term response to inflation, $(1-\rho^i)\gamma^\pi$, may be substantially smaller than the long-term response, γ^π. The monetary policy shock $v_{i,t}$ represents a deviation from this rule.

3.2 Firms

The supply side of the economy is standard, and we relegate the details to Appendix B. Partially monopolistic firms are assumed to set product prices but can adjust their product prices only in some periods according to Calvo (1983) with backward-indexation (Christiano et al., 2005). Supply shocks in the model are defined as cost-push shocks and formally arise as markup shocks to firms' market power over the variety they produce. We model households' labor-leisure trade-off from the consumption of home-produced goods in the manner of Greenwood et al. (1988), with an external home consumption habit exactly offsetting the equilibrium level of home goods. These assumptions ensure that we can abstract from the labor-leisure trade-off in solving for equilibrium asset prices. Since the model does not have real investment, the aggregate resource constraint implies that consumption equals output, $C_t = Y_t$. Assuming that productivity follows a learning-by-doing process as in Lucas (1988) then implies that in equilibrium the log real output gap equals

[18]The 1979-1982 monetarist experiment provides an exception, though interest rates featured prominently in the Federal Reserve's considerations even during this episode.

stochastically de-trended real consumption (up to a constant)

$$x_t \;=\; c_t - (1-\phi)\sum_{j=0}^{\infty}\phi^j c_{t-1-j}. \tag{9}$$

3.3 Asset Pricing Euler Equations

We start from the standard asset pricing Euler equation for a one-period illiquid loan

$$E_t\left[M^{\$}_{t+1}\left(1+I^l_t\right)\right] \;=\; 1. \tag{10}$$

Here, the nominal SDF $M^{\$}_{t+1}$ is the real SDF (5) divided by the gross rate of inflation, $M^{\$}_{t+1} \equiv M_{t+1}\exp(-\pi_{t+1})$.

In equilibrium, the representative household must be indifferent between marginally increasing Treasury bond holdings while decreasing consumption subject to the budget constraint, giving the Treasury bond Euler equation

$$E_t\left[M^{\$}_{t+1}\left(1+I^b_t\right)\right] \;=\; 1-\underbrace{\frac{\frac{\alpha}{Q_t}\frac{\lambda_t}{1-\lambda_t}}{U_c\left(C_t,Q_t,H_t\right)}}_{\zeta^b_t}. \tag{11}$$

Note that the Euler equation (11) for liquid Treasury bonds takes exactly the same form as in models with a reduced-form Treasury convenience benefit ζ^b_t, which has proven useful in understanding global currency fluctuations (Jiang et al. (2021)) and international business cycles (Jiang et al. (2024), Kekre and Lenel (2024)). Bianchi et al. (2025) introduce a similar wedge between the household and financial market Euler equations in their model of high-frequency market responses to monetary policy. We provide a new connection between this increasingly successful financial market shock and the real economy.

The analogous Euler equation for deposits is given by

$$E_t\left[M^{\$}_{t+1}\left(1+I^d_t\right)\right] \;=\; 1-\underbrace{\frac{\frac{\alpha}{Q_t}}{U_c\left(C_t,Q_t,H_t\right)}}_{\zeta^d_t}. \tag{12}$$

Equation (11) shows that Treasury bond convenience ζ^b_t increases with the liquidity weight of Treasury bonds, $\lambda_t/(1-\lambda_t)$, and also with the marginal consumption value of liquidity $\frac{\alpha/Q_t}{U_{c,t}}$. Equation (12) shows that the convenience of deposits ζ_t^d increases with the liquidity value of deposits measured in marginal consumption units. Combining the first-order conditions for loan rate (10), Treasury bonds (11) and deposits (12) with assumption (7) linking the deposit and loan rates, delivers the one-period Treasury convenience spread:

$$I_t^l - I_t^b = \frac{\lambda_t}{1-\lambda_t}\left((1-\delta)I_t^l - \rho^d I_{t-1}^d\right). \tag{13}$$

To interpret equation (13) note that in the special case where deposits are simply liquid cash ($\delta = \rho^d = 0$), the nominal loan rate I_t^l is the cost of holding non-interest-bearing cash, and $\lambda_t/(1-\lambda_t)$ is the liquidity value of Treasuries relative to cash. The convenience spread is hence higher when Treasuries provide greater liquidity services (higher λ_t) or when loan rates rise relative to sticky deposit rates.

3.4 Stocks and Long-Term Bonds

Because long- and short-term bonds enter as perfect substitutes into the liquidity aggregate, investors are indifferent between holding a one-period or long-term bond for one period, and the price of a nominal n-period zero-coupon bond, $P_{n,t}^{\$}$ satisfies the following recursion

$$E_t\left[M_{t+1}^{\$}\frac{P_{n-1,t+1}^{\$}}{P_{n,t}^{\$}}\right] = E_t\left[M_{t+1}^{\$}\exp(i_t^b)\right], \tag{14}$$

where the price of a one-period liquid nominal bond equals $P_{1,t}^{\$} = \exp(-i_t^b)$.

Different from bonds, consumption claims do not provide any liquidity services, following the standard recursion:

$$\frac{P_{n,t}^c}{C_t} = E_t\left[M_{t+1}\frac{C_{t+1}}{C_t}\frac{P_{n-1,t+1}^c}{C_{t+1}}\right]. \tag{15}$$

The price-consumption ratio for a claim to aggregate consumption is equal to the infinite sum of zero-coupon consumption claims:

$$\frac{P_t^c}{C_t} = \sum_{n=1}^{\infty}\frac{P_{n,t}^c}{C_t}. \tag{16}$$

We assume that stocks represent a levered claim on aggregate consumption following Abel (1990) and Campbell (1986), with an equity share δ^c. The assumption that stocks are a claim on consumption rather than on firm profits is merely to keep the model as simple as possible. Alternatively, one could assume sticky wages rather than sticky prices, which would leave inflation and output gap dynamics almost identical while implying that firm profits are proportional to consumption.

3.5 Log-Linear Convenience Spread

We now describe the log-linear solution for macroeconomic dynamics and convenience yields. Macroeconomic dynamics and the dynamics for the convenience spread are derived by log-linearizing the model around the flexible-price steady-state $\bar{c}$, $\bar{y}$, $\bar{\pi}$, $\bar{i}^l$, $\bar{i}^b$, $\bar{i}^d$, and $\bar{\lambda}$. For ease of notation, we use c_t, y_t, π_t, i_t^l, i_t^b, i_t^d, and λ_t to denote log deviations from steady-state values. We follow Campbell et al. (2020) in ignoring risk premia for one-period liquid and illiquid bonds, which are typically small, leading to a log-linear solution for macroeconomic dynamics. This facilitates solving for non-linear risk premia in long-term Treasury bonds and stocks.

We start by log-linearizing the equilibrium condition for our new state variable, namely the convenience spread in equation (13). This allows to express the convenience spread (up to a constant) as

$$\ell_t \equiv i_t^l - i_t^b \quad = (f^i - 1)i_t^b + f^\lambda \lambda_t - f^d i_{t-1}^d. \tag{17}$$

The first term on the right hand side captures the essence of the money channel. We show that f^i is strictly greater than one ($f^i > 1$), under the previously stated assumptions that $0 < \bar{\lambda} < \frac{1}{2}$ and $\delta/(1-\rho^d) < 1$ (see Appendix B.3.5 for a detailed derivation). The condition on $\bar{\lambda}$ means that Treasury bonds offer positive liquidity, but less than liquid deposits, in steady state. Intuitively, the opportunity cost of holding deposits rises as the loan rate increases relative to the deposit rate. Because the policy rate and the loan rate are linked, this leads to an increasing relationship between the log convenience spread and the log policy rate. Importantly, this derivation of the money channel depends only on liquidity and deposit rate setting, but not on monetary policy. The constant f^λ is a positive log-linearization constant linking the magnitude of λ_t to its impact on illiquid loan rates. Finally, the constant f^d reflects the sluggish deposit-rate adjustment and is proportional to ρ^d.

To show the dynamic evolution of the convenience spread, we substitute out the lagged deposit rate, which is not a state variable, by log-linearizing the deposit-rate setting equation (7), and we obtain

$$\ell_t = \rho^\ell \ell_{t-1} + (f^i - 1 - \frac{f^d}{\rho^i}) i_t^b + \frac{f^d}{\rho^i}(i_t^b - \rho^i i_{t-1}^b) + v_{\ell,t} \tag{18}$$

Here, the liquidity shock $v_{\ell,t}$ is a transformation of λ_t and λ_{t-1} given in equation (A50) in the Appendix. We assume that the liquidity shock $v_{\ell,t}$ is i.i.d and normally distributed. The first term in expression (18) is the serial correlation of the convenience spread, where the autoregressive coefficient ρ^ℓ is primarily driven by the sluggish deposit rate updating, and closely linked to ρ^d as in equation (17). The second term reflects the money channel arising from the current policy rate as in Nagel (2016). A higher policy rate increases the cost of holding liquid deposits, which is the foregone interest of holding liquid deposits rather than less liquid interest-bearing loans. Since Treasury bonds act as substitutes in the liquidity aggregate equation (2), the cost of holding Treasuries and hence ℓ_t also changes with the policy rate.

The third term in (18) is new and shows a positive dependence on the policy rate drift. This arises because deposit rates in the model are assumed to adjust sluggishly ($\rho^d > 0$). Comparing to the monetary policy rule (8) shows that the policy rate drift depends in particular on inflation, as monetary policy raises rates when inflation is higher, so this induces a positive relationship between the convenience spread and inflation even controlling for the current policy rate.

3.6 Log-Linear Macroeconomic Dynamics

The next equilibrium condition is the representative household's intertemporal first-order condition, which takes the exactly log-linear form

$$x_t \;=\; \rho^x x_{t-1} + (1-\rho^x) E_t x_{t+1} - \psi \left(i_t^l - E_t \pi_{t+1} \right) + v_{x,t}. \tag{19}$$

The derivation of (19) largely follows Campbell et al. (2020) and is provided in Appendix B.3.4. Intuitively, it follows from the Euler equation for the one-period illiquid loan (10), combined with the Fisher equation $r_t^l = i_t^l - E_t \pi_{t+1}$, and the consumption-output gap link (9). The parameters ρ^x and ψ are functions of the representative agent's preference parameters. The non-liquidity demand shock $v_{x,t}$ is new and proportional to η_t in log surplus consumption dynamics (6). This demand shock captures the typical New Keynesian demand taste shifter, whereby a decrease in patience raises consumption and output today at given interest rates and Treasury liquidity (e.g., Galí, 2008).

Using the definition of the convenience spread in (17), we re-write the Euler equation (19) in terms of our state variables

$$x_t = \rho^x x_{t-1} + (1-\rho^x)E_t x_{t+1} - \psi\left(i_t^b - E_t\pi_{t+1}\right) - \psi\ell_t + v_{x,t}. \tag{20}$$

An increase in government bond convenience acts just like a negative aggregate demand shock via the $-\psi\ell_t$ term. When Treasury convenience increases, households face a higher loan rate for a given Treasury bond rate, increasing their incentive to save and decreasing the incentive to consume this period. We hence obtain a wedge in the consumption Euler equation due to the difference ℓ_t between the rate at which the household borrows and saves i_t^l, given its Euler equation (10), and the policy rate i_t^b, similar to the model with microfounded liquidity demand by Piazzesi et al. (2019). While equation (20) shows the correspondence between the two demand-type shocks ℓ_t and $v_{x,t}$ in how they affect the output gap, below we highlight their distinct implications for convenience spreads and asset prices.

The final log-linearized equilibrium condition is given by the standard Phillips curve linking inflation and the output gap

$$\pi_t = \rho^\pi\pi_{t-1} + (1-\rho^\pi)E_t\pi_{t+1} + \kappa x_t + v_{\pi,t}. \tag{21}$$

Here, ρ^π and κ are log-linearization constants. We model the shock $v_{\pi,t}$ as a cost-push shock originating from time-varying firm markups, and refer to it as a supply shock for short (see Appendix B for details).[19] The slope parameter κ reflects the rise in marginal costs of production when output is running above potential, leading firms to optimally raise prices.

3.7 Model Solution

The model has four shocks: the cost-push supply shock $v_{\pi,t}$ in equation (21), the liquidity demand shock $v_{\ell,t}$ in equation (18), the monetary policy shock $v_{i,t}$ in equation (8), and the non-liquidity

[19] Alternative interpretations for a shock to the Phillips curve have been proposed, including fiscal dominance (Bianchi et al. (2023)), shifts in inflation expectations (Hazell et al. (2022)), deviations between marginal costs and output (Galí and Gertler (1999)), and policy makers learning about potential output (Primiceri (2006)). We do not describe the optimal monetary policy response, which differs depending on whether or not a shock to total factor productivity is observed (Nakamura et al. (2025)).

demand shock in equation (6),

$$v_{x,t} = \psi \eta_t. \tag{22}$$

All four shocks are assumed to be i.i.d, independent, normally distributed with mean zero and standard deviations σ_ℓ, σ_π, σ_i, and σ_x. We define a vector of these exogenous shocks $v_t = [v_{\ell,t}, v_{\pi,t}, v_{i,t}, v_{x,t}]$.

We solve for equilibrium model dynamics in two steps. In a first step, we use Blanchard and Kahn (1980) algorithm to solve the equilibrium dynamics of the output gap x_t in (20), inflation π_t in (21), policy rate i_t^b in (8), and convenience spread ℓ_t in (18), which still contain expectations. We solve for minimum state variable dynamics following

$$Z_t = P^Z Z_{t-1} + Q^Z v_t. \tag{23}$$

where the state vector is $Z_t = [x_t, \pi_t, i_t^b, \ell_t]$. Our quantification features a single non-explosive equilibrium of the form (23). This solution is fast and allows for efficient estimation via simulated method of moments (SMM).

In a second step, we solve for asset prices via value function iteration, using surplus consumption dynamics (6) and the link between consumption and the output gap (9). While our asset pricing solution builds on Pflueger (2025), solving our model is non-trivial and introduces another state variable (the convenience spread) and two new shocks (liquidity and non-liquidity demand shocks). The log surplus consumption ratio, s_t is a state variable for asset prices but not for macroeconomic dynamics, entering into the SDF and the non-linear sensitivity function $\omega(s_t)$. Details are provided in Appendix B.

4 Model Estimation and Results

We quantify the model in two steps. First, we set all parameters other than the volatilities of shocks to standard values. The non-volatility parameters are held constant across estimation periods to better delineate the effects of changing shock volatilities. Second, we estimate the volatilities of shocks $(\sigma_\ell, \sigma_\pi, \sigma_i, \sigma_x)$ to match data moments in an SMM step.

4.1 Calibrated Parameters

Parameters for the New Keynesian block of the model are set to values from the literature. Preference parameters are set as in Pflueger and Rinaldi (2022), matching the output gap and stock responses to identified monetary policy shocks in the data. The Phillips curve slope is set as in Rotemberg and Woodford (1997) and the backward-looking and forward-looking coefficients in the Phillips curve are derived from backward-looking price indexation as in Smets and Wouters (2007). The steady-state discount rate is chosen to match a log real risk-free rate of 0.94% annualized following Campbell and Cochrane (1999). Combined with a steady-state level inflation rate of $\bar{\Pi} = 2\%$ in annual units, this implies a steady-state illiquid log nominal loan rate of 2.95% annualized. The monetary policy rule has an inflation coefficient of $\gamma^{\pi} = 1.5$, and output gap coefficient of $\gamma^{x} = 0.5$ following Taylor (1993) and inertia of $\rho^{i} = 0.8$ in quarterly units following Clarida et al. (2000). A strongly anti-inflationary monetary policy rule, such as this one, is important to ensure that supply shocks lead to stagflations and positive bond-stock betas.

The deposit rate pass-through from the illiquid rate in equation (7) is set following the liquidity literature and empirical evidence from CALL report data. The long-term deposit-rate adjustment to the illiquid rate $\delta/(1-\rho^{d})$ is set to 1/3, within the range of 1/3 to 1/2 suggested by Nagel (2016). We use CALL report data (quarterly frequency from 1987 Q1 to 2020 Q1) to estimate $\rho^{d} = 0.92$ quarterly from a time-series regression of the form (7). The resulting value for the short-run pass-through is $\delta = 0.027$, which is also consistent with the time series regression in CALL report data reported in Appendix A.7. The steady-state weight on Treasury bonds in the liquidity aggregate, $\bar{\lambda}$, is set to match the steady-state convenience yield in the data. The firm's equity-to-asset ratio is set to $\delta^{c} = 0.5$ (50%) as in Campbell et al. (2020) to generate reasonable equity market volatility. The full set of period-invariant parameter values is listed in Appendix Table A6.

4.2 Estimated Shock Volatilities

Table 4 shows the targeted data and model moments. To identify shock volatilities $(\sigma_{\ell}, \sigma_{\pi}, \sigma_{i}, \sigma_{x})$, we match six moments in the data for each separate period 1952–1999 and 2000–2000. The moments include the convenience-inflation regression coefficient from our baseline empirical results, the nominal bond-stock beta, and the output gap-inflation correlation,[20] as well as the quarterly

[20]Since we use quarterly data for the model estimation, the convenience-inflation coefficients in Table 4 are very similar but not identical to Table 2.

Table 4. Targeted moments in the model and the data. This table reports the model-implied moments with the targeted moments in the data for the two subperiods 1952-1999 and 2000-2020. The T-bill spread, inflation, and output gap are all expressed in percentage points.

		1952–1999		**2000–2020**	
		Data	Model	Data	Model
Volatilities	Vol(Tbill spread)	0.579	0.476	0.219	0.291
		(0.030)		(0.017)	
	Vol(Inflation)	3.277	2.374	2.665	0.679
		(0.168)		(0.207)	
	Vol(Output gap)	2.287	2.665	1.955	1.606
		(0.117)		(0.152)	
Correlations	Corr(Inflation, Output Gap)	-0.119	-0.312	0.292	0.427
		(0.072)		(0.102)	
Bond-Stock Beta	Bond return $\sim \beta \cdot$ Stock return	0.215	0.157	-0.288	-0.049
		(0.071)		(0.054)	
Reg Coefs	Tbill spread $\sim b \cdot$ Inflation	0.109	0.091	-0.003	-0.010
		(0.010)		(0.009)	

volatilities of the T-bill spread ℓ_t, quarterly inflation, and the quarterly output gap. Our objective function sums the squared deviation between each of these model and data moments, divided by the squared standard deviation of the empirical target moments.[21] The table does not report the regression coefficient of the convenience yield on inflation after controlling for the policy rate. However, including this control in model-simulated data yields a positive coefficient for 1952–1999 and a negative coefficient for 2000–2020, matching the signs in the data.

The three moments on volatilities discipline the overall scale of shocks, while the comovements discipline the relative magnitudes of shocks. Importantly, the liquidity demand shock is the only shock in the model driving the negative convenience spread-inflation relationship. By contrast, non-liquidity demand shocks, monetary policy shocks, and supply shocks all drive a positive convenience and inflation relationship. Therefore, the correlation between convenience and inflation is highly informative about the presence of liquidity-demand shocks. The correlation between

[21] Because solving for asset prices is substantially slower than solving for macroeconomic moments, we employ a two-stage estimation procedure, which first targets all moments except the bond-stock beta using a gradient-based estimation method. In a second step, we run a local grid search around the best estimate from the first step, minimizing the full objective function over all target moments, including bond-stock betas.

inflation and the output gap disciplines the relative dominance of supply shocks vs. non-supply shocks.

Intuitively, supply shocks combined with a strongly anti-inflationary monetary policy rule tend to generate stagflations and a negative inflation-output gap correlation, consistent with the empirical correlation for the 1952–1999 period. By contrast, liquidity demand shocks, non-liquidity demand shocks, and monetary policy shocks tend to drive down inflation at the same time as the output gap, hence generating a positive inflation-output gap correlation. In order to fit the negative inflation-output gap correlation in the 1952–1999 period, the model hence requires a high volatility of supply shocks. Conversely, the positive inflation-output gap correlation for the 2000–2020 requires more volatile demand-side shocks and less volatile supply shocks. Finally, as we will see shortly, liquidity and non-liquidity demand shocks have distinct implications for the convenience-inflation relationship and bond-stock betas, thereby pinning down their relative magnitudes.

Table 5. Estimated model parameters. This table reports the estimated model parameters for each period. Standard errors are reported in parentheses. We apply the asymptotic variance formula for GMM estimation to calculate the standard errors.

parameters (yearly values)	1952–1999		2000–2020	
σ_ℓ (liquidity demand)	0.100	(0.002)	0.079	(0.004)
σ_π (cost-push)	0.793	(0.009)	0.138	(0.010)
σ_i (monetary policy)	1.294	(0.017)	0.906	(0.022)
σ_x (non-liquidity demand)	0.050	(0.038)	0.534	(0.009)

Table 5 reports estimated shock volatilities for both subperiods and standard errors. We find that supply shocks were substantially more volatile in the 1952–1999 period than in the 2000–2020 period. While the estimated volatility of liquidity demand shocks is roughly constant across periods, the decline in supply shock volatility means that in relative terms the importance of liquidity demand shocks increased in the 2000-2020 period. The volatility of monetary policy shocks is estimated to have been high in both periods, though somewhat lower in the post-2000 period. Because our model does not incorporate the zero lower bound, deviations from the monetary policy rule due to the zero lower bound could be interpreted as volatile monetary policy shocks during the post-2000 period. Although the estimated σ_ℓ may appear smaller than the other shocks, it is important to keep in mind that the liquidity demand shock $v_{\ell,t}$ enters into the macroeconomic Euler equation (20) through the convenience spread ℓ_t, which contains endogenous persistence. Even though $v_{\ell,t}$ itself is assumed to be i.i.d., its macroeconomic impact is therefore amplified roughly by a factor

Table 6. Model versus data on stocks, bonds, and macroeconomic volatilities. This table reports asset pricing moments for stocks and bonds in the model vs. data, as well as macroeconomic volatilities. Stocks and macroeconomic volatilities are averaged across the two model estimations/data periods. Bond-stock betas in the data are computed as the regression coefficient of quarterly excess returns for a zero-coupon ten-year nominal government bond onto quarterly excess returns for the S&P 500. Quarterly nominal bond excess returns are computed from ten-year zero-coupon nominal bond yields from Gürkaynak et al. (2010). We substitute the seven-year zero-coupon nominal yield for the ten-year zero-coupon nominal yield when the ten-year zero-coupon is not available 1961–1971. The model risk-neutral (RN) nominal bond-stock beta reports the regression coefficient of quarterly nominal bond excess returns onto full stock returns in simulated data. All simulations use 50000 periods with a 100-period burn-in period. Ten-year CPI inflation expectations are from the Survey of Professional Forecasters after 1990 and from Blue Chip before that, available from the Philadelphia Fed research website.

Stocks	Model		Data	
Equity Premium	9.01		7.21	
Equity Vol	16.72		17.25	
Equity SR	0.54		0.43	
AR(1) pd	0.95		0.94	
1 YR Excess Returns on pd	-0.35		-0.26	
1 YR Excess Returns on pd (R^2)	0.05		0.14	
Macroeconomic volatilities				
Std. Annual Cons. Growth	2.26		1.54	
Std. Annual Change FFR	2.28		1.79	
Quarterly Std. Ten-Year Inflation Forecast	0.16		0.29	
Bond-stock betas	1952-1999	2000-2020	1952-1999	2000-2020
Nominal bond-stock beta	0.16	-0.05	0.22	-0.29
RN Nominal bond-stock beta	0.07	-0.01		

$1/(1-\rho^{\ell}) \approx 10$. Finally, non-liquidity demand shocks are also estimated to have increased from the 1952–1999 period to the post-2000 period. Overall, the estimated volatility parameters are consistent with the view that the U.S. economy shifted from being exposed to supply shocks to a greater dominance of demand shocks around the turn of the millennium.[22]

Table 6 compares model and data moments for stocks, bonds, and macroeconomic volatilities. Although only the nominal bond-stock beta is targeted by the estimation, the model generates various stock and bond moments consistent with the data, including the equity premium, equity volatility, Sharpe ratio, and the persistence of the price-dividend ratio. Additional non-targeted macroeconomic volatilities are also consistent with the data, including the standard deviations of annual consumption growth, annual change of the FFR, and the ten-year inflation forecast. The model undershoots the quarterly inflation volatility, reported in Table 4. Since our macroeconomic block is quite stylized, there are plausibly additional inflation drivers especially at the quarterly frequency that we cannot capture. The model bond-stock beta for the post 2000 period is less negative than in the data. This is conservative in the sense that matching exactly the negative bond-stock beta in this last period would imply a greater role for liquidity demand shocks.

To better understand the role of the inflation–convenience yield relation and bond-stock beta in disciplining model parameters, in Appendix Tables A7 and A8 we present two alternative estimation strategies: (1) shutting off the bond-stock beta from the target moments; and (2) shutting off both asset-pricing moments from the target moments. Results remain largely unchanged when we remove only the bond-stock beta moment. However, when we also remove the inflation–convenience yield relation, the liquidity and non-liquidity demand shock volatilities become very imprecisely estimated, and the model is unable to clearly differentiate these two types of demand shocks. We conclude that the inflation–convenience yield relation is a highly informative moment for differentiating between liquidity and non-liquidity demand shocks.

4.3 Model Implications for Macroeconomic and Asset Price Dynamics

We next examine the economic implications of the liquidity demand shocks and supply shocks for the macroeconomy and asset prices through impulse responses. Figures 3 and 4 present responses

[22] A high supply shock volatility during the second half of the 20th century is consistent with the deposit rate channel of Drechsler et al. (2017), and their finding that inflationary supply shocks were dominant during this period, even though we target a different set of moments in the data. Beyond this, Drechsler et al. (2023) highlight a complementary mechanism, whereby deposit outflows can trigger bank credit crunches and amplify the real effects of supply shocks.

to a liquidity demand shock ($v_{\ell,t}$) and a supply shock ($v_{\pi,t}$), respectively. The top row in each figure displays macroeconomic responses: the policy rate, inflation, and output gap. The bottom row displays asset price responses: the convenience spread, the ten-year nominal yield, and cumulative stock market return (in excess of its steady state value). Due to the linear affine structure of the macroeconomic dynamics, the impulse responses in the top row and the risk-neutral impulse responses in the bottom row are invariant across the two estimation periods. However, the shock variances and hence the dominance of each shock differ across periods. Because bonds and stocks in the model price equilibrium variances of shocks, model-implied risk premia are different across the two periods.

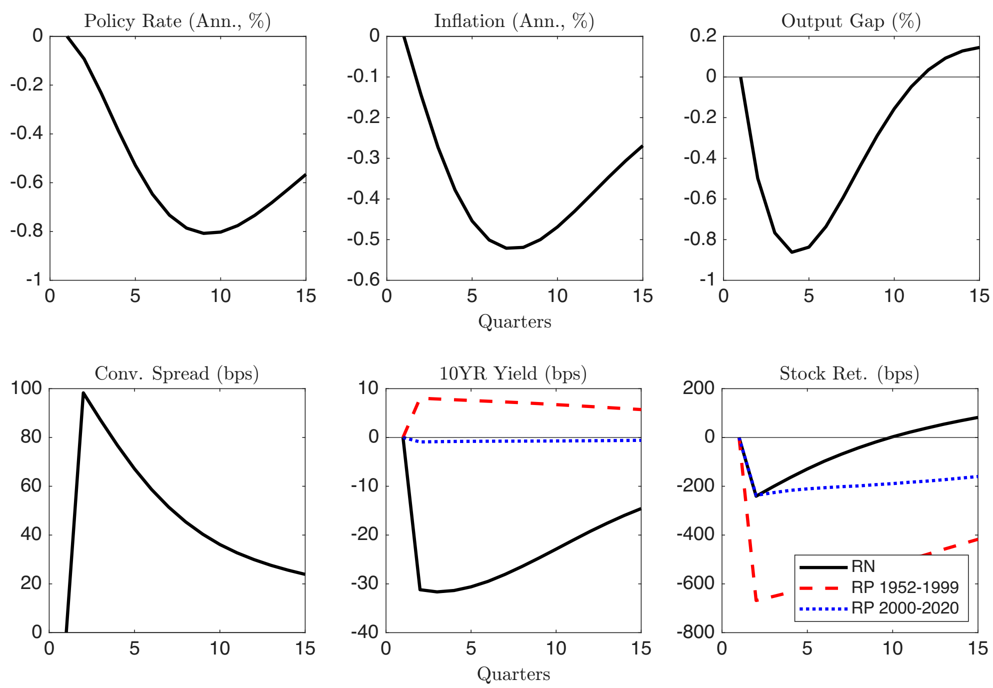
4.3.1 Impact of the Liquidity Demand Shock

Figure 3 considers a 100 bps increase in $v_{\ell,t}$. A shock of this size is consistent with magnitudes observed during major financial market disruptions, when liquidity demand shocks are likely to be especially relevant for macroeconomic fluctuations. In the top row, we see that the policy rate, i_t^b, declines by 0.8 percentage points or 80 bps, inflation, π_t declines by half a percentage point, and the output gap x_t declines by 0.8 percentage points.

Figure 3 shows that the liquidity demand channel in the model ties together several of our empirical findings. A liquidity demand shock induces inflation and the output gap to move in the same direction, as observed pre-WWII and 2000–2020. Further, the liquidity demand channel moves inflation and Treasury convenience in opposite directions, explaining our finding of a sharply lower empirical correlation between convenience and headline inflation in the post-2000 period. The intuition follows directly from equation (20): Households face a higher illiquid loan rate i_t^l at a given policy rate i_t^b, decreasing their demand to borrow and consume. Firms meet this weaker demand, reducing price pressure through the Phillips curve (21).

The macroeconomic responses to a liquidity demand shock are very similar to those of a comparably sized, negative, non-liquidity demand shock, as expected from ℓ_t and $v_{x,t}$ both appearing as shifters in the Euler equation (20). However, the bottom-left panel in Figure 3 shows that a liquidity demand shock raises the convenience spread while lowering inflation. This is different from a non-liquidity demand shock, which also implies a positive correlation between inflation and the output gap, but has almost no impact on the convenience spread. The macroeconomic impulse responses to a non-liquidity demand shock are standard and hence relegated to the Appendix Figure A6. The comovement between inflation and convenience is therefore a critical new moment,

Figure 3. Impulse responses to a liquidity demand shock. This figure shows impulse responses to a 1% (100 bps) increase in the liquidity demand shock $v_{\ell,t}$. Macro responses and risk-neutral (RN) asset-price responses are the same across the two periods and are denoted by black lines. Risk premium (RP) responses are shown in red dashed lines for 1952-1999 and blue dotted lines for 2000-2020. The total response is the sum of risk-neutral and risk premium responses (not plotted). The driving shock $v_{\ell,t}$ has mean zero and is drawn from the steady-state distribution in all other periods. All other shocks are drawn from their mean-zero steady-state distributions in all periods. Responses in the top row (policy rate, inflation, output gap) are in annualized percent. Responses in the bottom row (convenience spread, ten-year Treasury yield, cumulative stock return in excess of steady state) are in basis points (bps). Quarters after the impulse are shown on the x-axis. Impulse responses are averaged over 10^6 independent simulations with a 30-quarter burn-in period for asset prices in the bottom row. For the policy rate, inflation, and output gap, we reduce simulation noise by setting all non-driving shocks and the driving shock in the non-impulse period to zero.



allowing us to distinguish the nature of the demand shocks driving economic fluctuations.

The comovement of long-term nominal Treasury bonds with the stock market provides another marker of liquidity demand shocks. In the bottom row of Figure 3, we plot model ten-year Treasury bond yields, which are inversely related to bond prices in the bottom-middle, and cumulative stock returns since the shock period in the bottom-right. Risk-neutral responses are shown in solid black, with risk premia in blue dotted and red dashed, respectively, with risk-neutral and risk premium responses adding up to the total response (not plotted). We see that the ten-year Treasury bond yield mirrors the declines in inflation and the policy rate. Stocks fall due to the decline in the output gap. A peak decline in the output gap of roughly 0.8 percentage points is associated with a roughly 200 bps decline in risk-neutral stock returns. Investors become more risk-averse as consumption declines towards habit, so the required return on stocks increases, amplifying the fall in stock prices through time-varying risk premia.[23] The bond yield responses are dominated by the risk-neutral component, so the overall response for Treasury bond yields is a decline, implying a negative comovement between ten-year Treasury bond prices and stock returns. Hence, liquidity shocks in the model imply a negative nominal Treasury bond-stock return beta, consistent with the empirical hedging properties of nominal Treasury bonds during the 2000–2020 period.

Figure 3 also shows that the bond risk-premium response switches sign across periods, amplifying the positive bond-stock comovement for the pre-2000 period and the negative bond-stock comovement post-2000. When bonds' real cash flows are risky and fall in times of high marginal utility—as the model estimation implies for the 1952–1999 period—an increase in risk aversion leads investors to require higher risk discounts on stocks and bonds simultaneously. Conversely, when bonds are safe in terms of real cash flows—as in the model estimation for the 2000–2020 period—bond valuations benefit from an increase in risk aversion beyond risk-neutral valuation effects, increasing their valuations and driving down yields just as stocks fall. This is the endogenous "flight-to-safety" channel of Campbell et al. (2020, 2025).

4.3.2 Impact of the Supply Shock and the Money Channel

Figure 4 shows that the money channel explains the empirical evidence for the 1952–1999 period, as supply shocks shift convenience and inflation in the same direction. The intuition is that convenience moves with the opportunity cost of money, which rises with the policy rate and inflation due to incomplete and sluggish pass-through of policy to deposit rates.[24] Supply shocks also drive

[23] The macroeconomic and stock return responses to an identified monetary policy shock in the model are similar to Pflueger and Rinaldi (2022), who show that they can match the empirical evidence by Bernanke and Kuttner (2005). The risk premium amplification in the model is stronger in the pre-2000 period, because the overall risk in this period is greater.

[24] As shown by equation (17), we do not need restrictions on the monetary policy rule coefficients or active monetary policy for the presence of a money channel. Even though we do not formally capture this in our estimation, if inflation were subject to sunspot fluctuations and the inflation coefficient in the monetary policy rule were less than one during the 1970s, as argued by Clarida et al. (2000), this would further drive volatility in inflation and the policy rate. Such sunspot fluctuations would, therefore, act similarly to supply shocks for our purposes, generating a positive relationship between inflation and convenience.

additional characteristic features of the 1952–1999 data. They raise inflation just as the output falls, implying a negative inflation-output gap correlation. Finally, an adverse supply shock leads to an increase in ten-year Treasury nominal bond yields to reflect both the increase in inflation and the resulting expected policy rate increase, driving down bond prices just as stocks decline due to lower risk-neutral dividends and risk-bearing capacity. Hence, an additional characteristic of supply shocks in the model is that they generate a positive nominal bond-stock beta, consistent with the data for the 1952–1999 period (see Table 4). Overall, the money channel ties together our empirical evidence for a positive convenience-inflation relationship in the second half of the 20th century.

The model also delivers predictions for the effect of monetary policy shocks on convenience (see Appendix Figure A5). A monetary policy shock has opposing effects on the policy rate and inflation, leading to offsetting effects on the convenience spread via equation (18). As a result, monetary policy shocks cannot explain the key positive convenience-inflation comovement in the 1952-1999 period.

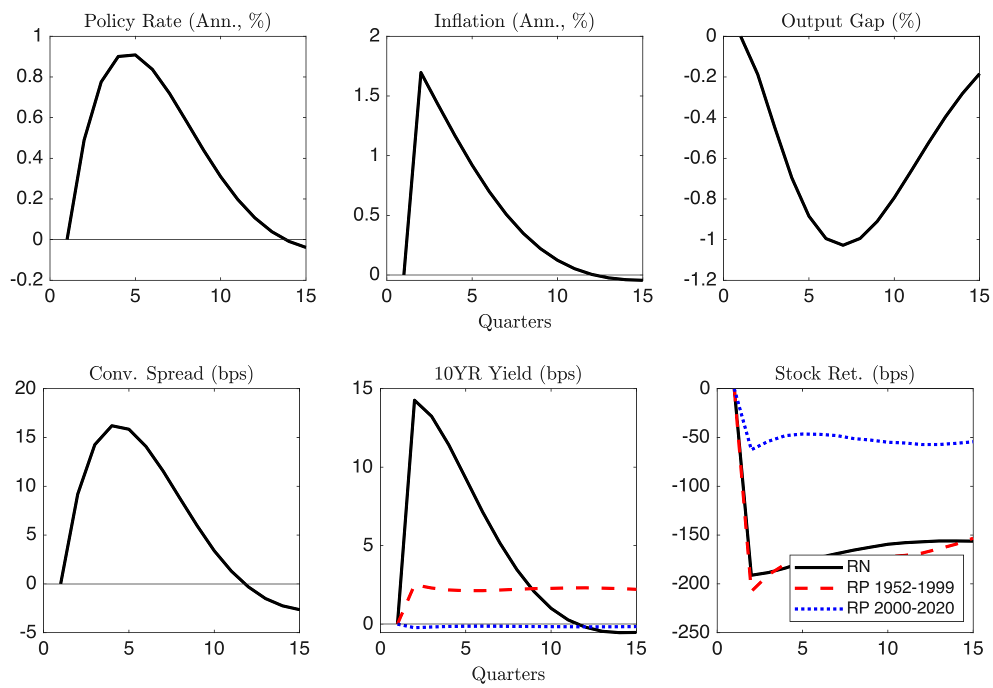
4.4 Bond-stock betas and the convenience-inflation relationship

We next use model counterfactuals to show that the money channel should be expected to be stronger when bond-stock betas are positive, as would be the case when supply shocks are present. This matches the empirical evidence in Table 3, where the convenience-inflation relationship is most positive in subperiods when bond-stock beta is also highest.

Figure 5 describes how the relationship between convenience spreads and inflation depends on the equilibrium distribution of shocks, and how this relationship links to the bond–stock beta. The figure reports regression coefficients from a regression of the convenience yield spread ℓ_t on inflation π_t, with both variables in annualized percentage units. On the data side (gray bars), we use the point estimates from the subperiod regressions in Table 3. On the model side (black line with triangles), we compute moments across four counterfactuals, moving from the 1952–1999 shock volatilities to the 2000–2020 shock volatilities in equal increments.

The figure shows that as we vary the equilibrium shock volatilities, the model matches the strengthening of the convenience spread–inflation relationship as the bond–stock beta increases. The leftmost triangle corresponds to the 2000–2020 estimation, where we already saw in Table 4 that the model generates an essentially zero but slightly negative convenience spread–inflation

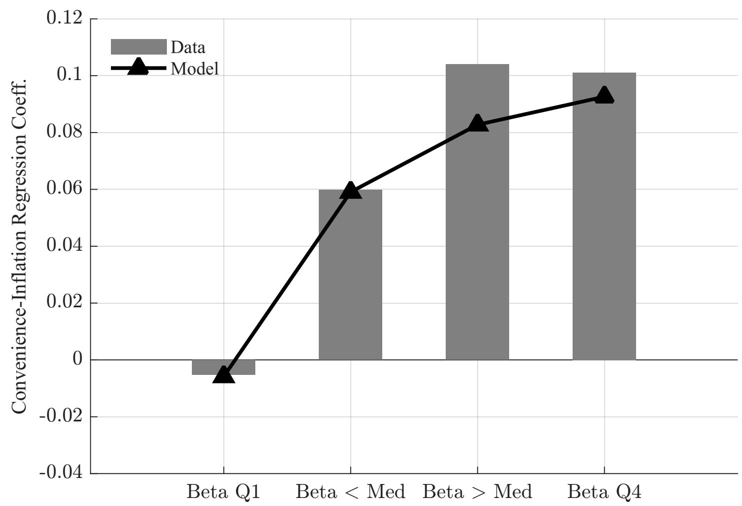
Figure 4. Impulse responses to a supply shock. This figure shows impulse responses to a one percentage point increase in the supply shock $v_{\pi,t}$. Macro responses and risk-neutral (RN) asset-price responses are the same across the two periods and are denoted by black lines. Risk premium (RP) responses are shown in red dashed lines for 1952-1999 and blue dotted lines for 2000-2020. The total response is the sum of risk-neutral and risk premium responses (not plotted). The driving shock $v_{\pi,t}$ has mean zero and is drawn from the steady-state distribution in all other periods. All other shocks are drawn from their mean-zero steady-state distributions in all periods. Responses in the top row (policy rate, inflation, output gap) are in annualized percent. Responses in the bottom row (convenience spread, ten-year Treasury yield, cumulative stock return in excess of steady state) are in basis points (bps). Quarters after the impulse are shown on the x-axis. Impulse responses are averaged over 10^6 independent simulations with a 30-quarter burn-in period for asset prices in the bottom row. For the policy rate, inflation, and output gap, we reduce simulation noise by setting all non-driving shocks and the driving shock in the non-impulse period to zero.



relationship, consistent with the data. At the same time, this equilibrium also generates the most negative bond–stock beta. The rightmost model triangle corresponds to the 1952–1999 period, where the model replicates both the positive convenience spread–inflation relationship and bond-stock betas in the data. As the shock volatilities move between these two extremes, the convenience spread–inflation coefficient increases along with the bond–stock beta in the model, similarly to the data.

Figure 6 shows that the joint observations of a negative bond-stock beta and a sharply lower

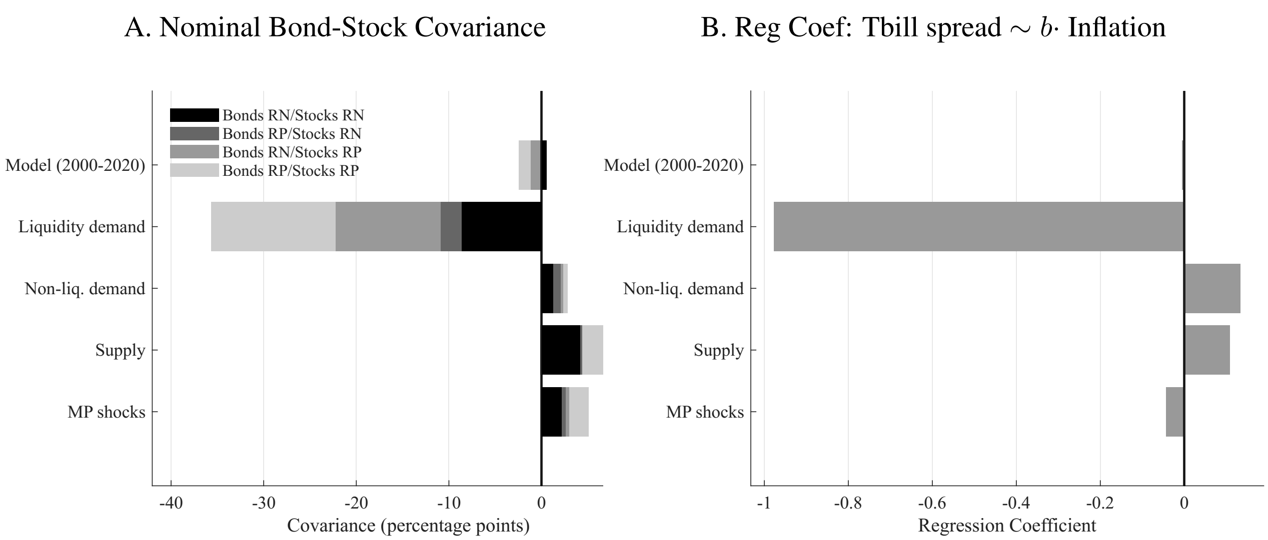
Figure 5. Convenience-inflation regression by bond-stock betas. This figure shows the regression coefficient β from a regression $\ell_t = a + b \cdot \pi_t + \varepsilon_t$ on the y-axis, with the spread and quarterly inflation both in annualized percent units. The empirical sample covers 1961-2000 and is split into subsamples according to a 120-day backward-looking rolling beta for the zero-coupon seven-year nominal Treasury bond. We use zero-coupon nominal Treasury yields from Gürkaynak et al. (2007) to compute daily bond returns. We consider the subsamples where the rolling bond-stock beta in its the bottom quartile (Beta Q1), below median (Beta $<$ Med), above median (Beta $>$ Med) and in its top quartile (Beta Q4). The model moments are from four different calibrations, where in each case we re-solve for equilibrium asset prices, moving from the 1952-1999 parameter values to the 2000-2020 parameter values in equal increments.



convenience-inflation relationship during the post-2000 period can only be explained by dominant liquidity demand shocks. Each row in Figure 6 increases the volatility of one shock at a time, using the 2000–2020 period as a baseline, and decomposes bond-stock comovements into risk-neutral and risk premium components.[25] Panel A shows that only liquidity demand shocks generate a negative bond–stock covariance in equilibrium, consistent with the empirically observed negative bond–stock betas during the 2000–2020 period. By contrast, increasing the volatility of supply shocks, monetary policy shocks, or non-liquidity demand shocks all imply positive bond–stock covariances, and hence cannot by themselves explain the post-2000 confluence of negative bond–stock betas and a sharply lower convenience spread–inflation coefficient. As the light gray bar in the second row shows, time-varying risk premia in bonds and stocks also comove

[25]This figure illustrates bond-stock covariances rather than betas, because bond-stock covariances satisfy an adding up constraint for the risk premium and risk-neutral components. The overall bond-stock covariance is the sum of the risk-neutral/risk-neutral, risk premium/risk-neutral, risk-neutral/risk premium, and risk premium/risk premium components.

Figure 6. Model bond-stock covariances for counterfactual demand-side shocks. Panel A shows the equilibrium covariance between ten-year nominal bond returns and stock returns in the model for different counterfactuals, starting from the model estimation for 2000-2020 in the top row. The subsequent rows increase each shock standard deviation by 100 bps while holding all other parameters constant at the 2000-2020 estimated values shown in Table 5. The covariance between the risk-neutral (RN) components of bond and stock returns is shown in black, the covariance between the risk premium (RP) component of bond returns with the risk-neutral component of stock returns in dark gray, the covariance between risk-neutral bond returns with the risk premium component of stock returns in medium gray, and the covariance between the risk premium components of bond and stock returns in light gray. The total covariance of bond and stock returns is the sum of these four components, corresponding to the combined length of the bars. Panel B shows the model regression coefficient of the T-bill-convenience spread onto quarterly inflation for the same combination of shock standard deviations.



strongly negatively in the counterfactual with more volatile liquidity demand shocks. Said differently, while the the impulse responses in Figure 3 show a small negative bond risk premium response to a liquidity demand shock for the 2000–2020 estimation, this response is amplified when liquidity demand shocks are priced as even more dominant, resembling an endogenous "flight-to-safety" effect to bonds when equilibrium concerns about liquidity shocks are prevalent.[26]

[26] While our finding that dominant liquidity demand shocks were important for explaining the negative bond-stock comovement post-2000 is qualitatively in line with the econometric analysis of Antolin-Diaz (2024), he does not consider the relationship with the macroeconomy or risk premia in a structural model, as we do. Cieslak and Pang (2021) also identify an increased importance of the hedging premium news in the post-1998 sample, which they argue drove the negative bond-stock comovement in this period.

4.5 Distinguishing Between Liquidity and Non-Liquidity Demand Shocks

Our analysis yields new insights regarding the distinction between liquidity and non-liquidity demand shocks, relative to the prior literature on bond-stock betas. As discussed in the context of the impulse responses in Figure 3, a liquidity demand shock tends to raise bond prices just as the economy and stock market fall, implying a negative bond-stock beta. By contrast, a negative non-liquidity demand shock, corresponding to an increase in patience, raises the risk-neutral valuations of all long-term assets, including both bonds and stocks, and thereby tends to push bond-stock betas upwards. Because risk premium responses also change with the equilibrium, the bond-stock comovement is pushed further upwards when volatile non-liquidity demand shocks are priced in equilibrium. Finally, it is intuitive that supply and monetary policy shocks tend to push towards positive bond-stock betas, since both lead to increases in bond yields and are associated with recessions.[27] Different from liquidity demand shocks, non-liquidity demand shocks hence cannot explain negative bond-stock betas during the post-2000 period.

Figure 6, Panel B shows that the liquidity demand shock is also the most potent driver of negative convenience spread–inflation comovements. By contrast, supply shocks and non-liquidity demand shocks generate positive convenience spread–inflation comovements through the money channel, different from the post-2000 evidence. Monetary policy shocks do generate a small negative convenience spread–inflation comovement, because they have offsetting effects on convenience via interest rates and inflation. The magnitude is small, especially in relation to the large and positive effect that monetary policy shocks have on the bond-stock comovement, as shown in Panel A. Monetary policy shocks alone, hence, cannot explain the empirically observed combination of a positive bond-stock beta and positive convenience-inflation relationship observed 1952-1999. Monetary policy shocks are also unable to generate a combination of a sharply lower convenience-inflation relationship together with negative bond-stock betas observed in the post-2000 period.

The non-liquidity demand shocks could in principle generate a positive convenience-inflation relationship. Nonetheless, we fail to find that non-liquidity demand shocks played a large role during 1952–1999, when a positive convenience-inflation relationship prevailed. The reason is that the negative inflation-output correlation of 1952–1999, which we target in estimation, cannot be rationalized by volatile demand shocks. We estimate a substantial non-liquidity demand shock volatil-

[27] For complete impulse responses to non-liquidity demand and monetary policy shocks, see Appendix Figures A6 and A5.

ity for 2000–2020 due to the positive inflation-output gap correlation during that period. However, non-liquidity demand shocks cannot explain the sharp decline in the convenience-inflation relationship after 2000, relative to the earlier decades.

Overall, in the model, the money channel drives a positive comovement between convenience spreads and inflation when supply shocks are large. By contrast, when liquidity demand shocks dominate, the inflation–convenience relationship weakens or becomes negative. These channels have additional implications for output, inflation, and asset prices that are consistent with the data. Our model suggests that supply shocks were dominant in the second half of the 20th century, driving stagflations, a positive nominal bond–stock beta, and a positive relationship between inflation and Treasury convenience. As supply shock volatility declined in the 2000s, demand shocks emanating from the demand for liquidity, such as during the global financial crisis of 2007–2009, drove a negative convenience-inflation relationship and negative bond-stock betas, at the same time contributing to a positive inflation-output gap correlation.

5 Discussion: Post-COVID Period Inflation and Convenience

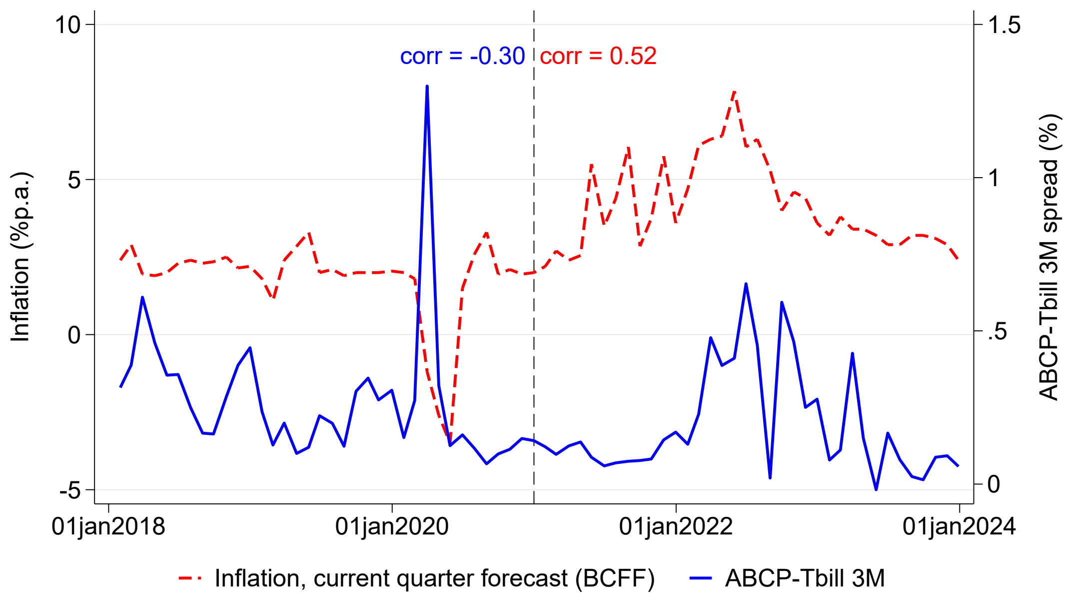
Our main analysis reveals structural shifts in the relation between inflation and the Treasury convenience yield over a long historical sample. These shifts are not isolated to historical episodes but remain relevant today.

Figure 7 superimposes the monthly T-bill convenience spread against the monthly inflation nowcast for the current quarter from the Blue Chip survey from 2018 to 2023.[28] We split the recent period into an early sample (2018–2020, low inflation) and a later sample (2021–2023, high inflation). While the correlation between Treasury convenience and inflation was negative during and immediately following the initial COVID-19 shock in March 2020, it became positive as inflationary pressures emerged after 2020. Figure 7 reports that the correlation increases from -0.30 during 2018–2020 to 0.52 during 2021–2023. This pattern remains robust across alternative convenience proxies. For example, using the three-month Refcorp–Treasury spread, the correlation shifts from -0.72 before 2020 to 0.72 after 2020. In Appendix A.9, we further confirm that the

[28]The BCFF survey is conducted in the last week of each month. For the purposes of the graph, using the Blue Chip nowcast allows us to pin down the timing of COVID disinflation, aligning it with the end-of-month T-bill spread. The results are very similar if we use realized quarterly inflation, yielding a negative inflation-spread correlation of -0.23 pre-2020 and a positive correlation of 0.34 post-2020. We do not use the Cleveland Fed inflation expectation because it is contaminated by term premia, as shown in Appendix A.5.

inflation-convenience yield correlation is negative in the period leading up COVID and becomes significantly positive afterwards when cost-push supply shocks become dominant, and this pattern is robust to controlling for FFR and Debt/GDP.

Figure 7. Convenience and inflation after 2018. We plot inflation against short-term convenience from 2018:01 to 2023:12. Inflation is the nowcast for the current quarter from Blue Chip Financial Forecasts (BCFF). The vertical dashed line marks the end of 2020.



During 2020, the economy witnessed a sharp drop in inflation due to the initial negative demand shock induced by the COVID-19 pandemic and then a gradual recovery from that shock. While the initial demand shock did not originate primarily as a liquidity or banking crisis, the pandemic shock was associated with heightened liquidity demand, as visible in Figure 7. One potential channel is that households and firms faced sudden income declines, necessitating immediate access to cash or liquid savings to cover essential expenses. Accordingly, different convenience yields spiked at the onset of the pandemic and then recovered during the rest of 2020.[29]

Our model offers a framework for interpreting the shift in the inflation–convenience comove-

[29] We note that the spike in the convenience yield does not conflict with the literature that documents a spike in the long-term Treasury yield at the onset of the pandemic (He et al., 2022). Despite an increase in Treasury yield during that time, yields of relatively less liquid safe bonds such as agency bonds increased by more, reflecting a scarcity of liquidity. Additionally, the spike in Treasury yields can be interpreted as investors resorting to the liquidity benefits of Treasuries in a "dash for cash" episode (Schrimpf et al., 2021; Vissing-Jorgensen, 2021; Duffie, 2023).

ment during this episode. It suggests that the roughly 100 bps increase in the convenience spread in March 2020 may have contributed roughly 0.5 percentage points to the sharp decline in inflation in 2020. Consistent with the impulse responses for the liquidity demand shock (Figure 3), we observe that the initial spike in the convenience spread precedes the decline in inflation. While non-liquidity demand shocks played an important role in this period, our analysis of the post-2000 data suggests that liquidity shocks should have also contributed meaningfully to disinflationary pressures. Beginning in 2021, however, a series of inflationary shocks pushed inflation above 2%, and the correlation between inflation and convenience switched from negative to positive. This reversal aligns with the model's prediction that large inflationary supply shocks can dominate the inflation–convenience relationship. According to the impulse responses in Figure 4, an inflationary supply shock raising inflation by five percentage points from two percent to seven percent would lead to an increase in convenience by 50 bps, similar to the peak in the ABCP spread in Figure 7.

Throughout our analysis, we emphasize the distinct impacts of liquidity demand and supply shocks on the relationship between Treasury convenience and inflation over the past century. In recent years, fiscal policy has been highlighted as a significant factor behind both elevated inflation and the decline in Treasury convenience. While our focus is not on the independent role of fiscal factors, it is important to recognize that these factors can influence inflation, inflation expectations and demand, with implications for the inflation-convenience relationship.

Several historical fiscal episodes are pertinent in this context. The fiscal dynamics underlying the Great Inflation from the 1960s to the early 1980s have long been debated. If fiscal policy contributed to the rise in trend inflation (e.g., Bianchi and Ilut, 2017), it may have also driven the positive comovement between inflation and convenience through the money channel. Similarly, fiscal stimulus during the post-Covid period, especially the American Rescue Plan Act in March 2021, may have contributed to inflation (Bianchi et al., 2023) and thereby increased convenience through the money channel, similarly to the 1970s and 1980s.

6 Conclusion

This paper argues that two competing mechanisms driving Treasury bond convenience – the "money channel" and the "liquidity demand channel" – dominated over distinct historical periods, leading to secular shifts in the comovement between Treasury convenience and inflation. Using a century of data, we show that during a large part of the inflationary 1970s and 1980s, higher inflation coincided with higher Treasury convenience. By contrast, high Treasury convenience was associated with low inflation during the first half of the 20th century and again during the post-2000 period. The output gap-inflation relationship and bond-stock betas also changed around the turn of the millennium, with both moments supporting a change in prevalence of supply shocks toward liquidity demand shocks. Splitting the sample by bond-stock betas as a financial market indicator for supply versus demand shocks paints a coherent picture whereby the positive relationship between inflation and convenience during the 1970s and 1980s was driven by periods with the highest bond-stock betas.

We provide a New Keynesian asset pricing model that embeds the money channel of Treasury convenience along with exogenous liquidity demand shocks. The model predicts that an inflationary supply shock raises inflation, the opportunity cost of holding money, and the price of holding convenient Treasuries. This money channel explains the positive convenience-inflation relationship that emerged during the 1970s. The model also implies that supply shocks lead to stagflationary recessions, and risky nominal Treasuries, as measured by the bond-stock beta, explaining the 1952–1999 evidence. Conversely, a liquidity demand shock increases the incentive to save and reduce consumption, lowering demand and, hence, inflation. Negative liquidity demand shock thus simultaneously induce negative inflation-convenience relationship and negative bond-stock betas, explaining the experience of the early 20th century and early 2000s.

Our results underscore that the relationship between Treasury convenience and inflation is informative about the sources of macroeconomic fluctuations – particularly, liquidity demand shock as a driver of inflation and the real economy. Distinct from non-liquidity demand shocks, liquidity demand shocks can explain the post-2000 confluence of negative convenience-inflation relationship seen in crisis periods, low-inflation recessions, and Treasury bonds as stock-market hedges.

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Online Appendix to
"Inflation and Treasury Convenience"

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A Data and Robustness of Empirical Results

In this section, we provide details on dataset construction and robustness checks of our main results.

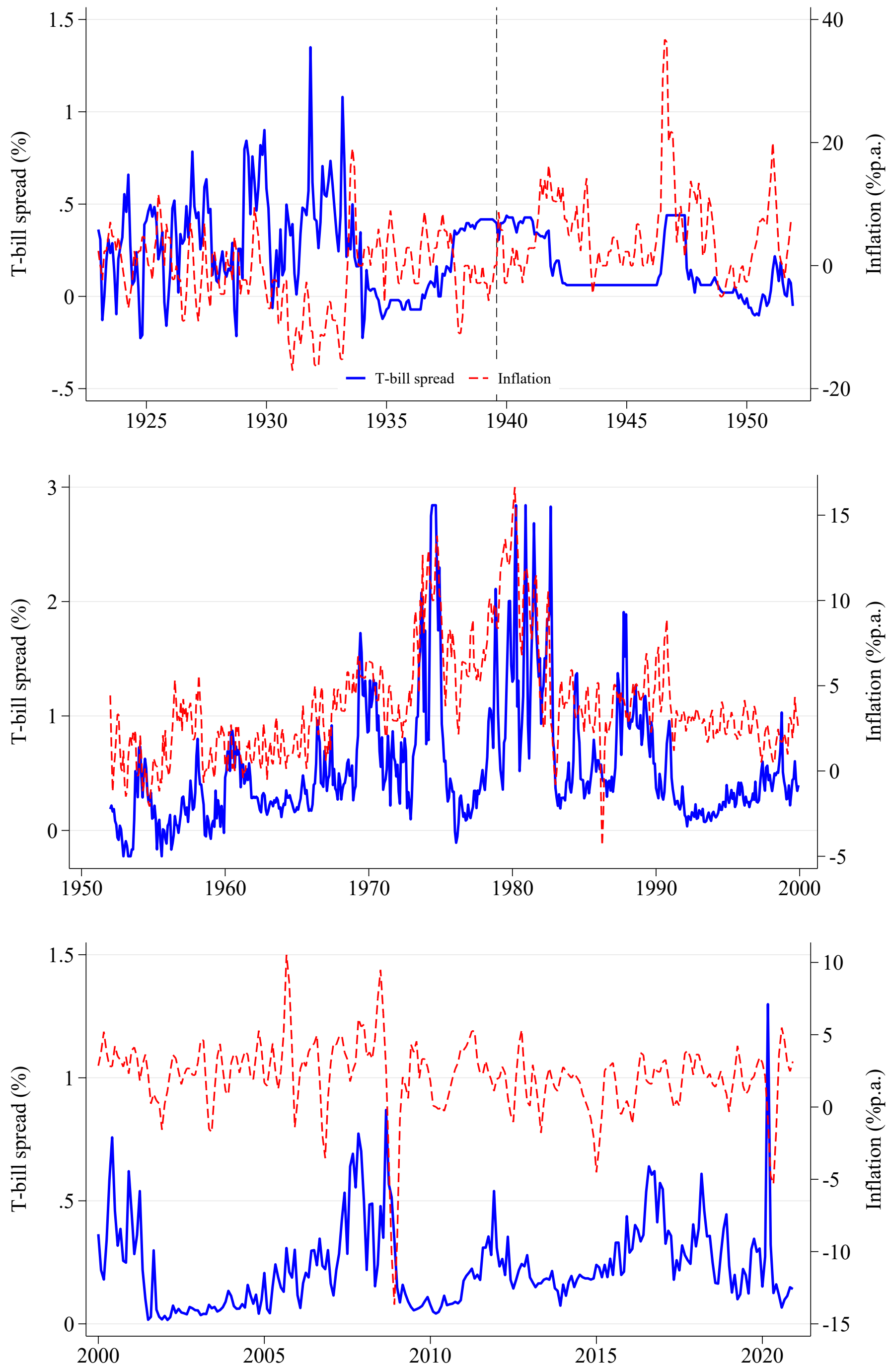
Table A1 reports the summary statistics for main variables used in empirical analysis.

Table A1. Summary statistics. This table presents summary statistics for our full sample 1923:01–2020:12, excluding the 1939:09–1951:12 period.

Variable	Obs	Mean	Std. Dev.	Mean		
	(1923-2020)			(1923-1939)	(1952-1999)	(2000-2020)
Tbill spread (%)	1028	0.42	0.46	0.27	0.56	0.23
Inflation (%)	1028	2.48	4.33	-1.13	3.90	2.10
VIX (%)	1028	19.7	8.45	25.9	17.4	19.9
Baa-Aaa spread (%)	1028	1.17	0.70	1.99	0.93	1.05
Debt/GDP	1028	0.29	0.15	0.17	0.26	0.47

Figure A1 visualizes the non-standardized series for the T-bill spread and inflation including the WWII period. We observe that during WWII, inflation is highly volatile but convenience yield is stable, due to the tight interest rate controls. Thus, we exclude WWII through 1951 from our analysis.

Figure A1. Time series of T-bill spread and inflation. This figure shows annualized quarterly inflation and the T-bill spread for the three subperiods used in our empirical analysis, 1923-1939:08, 1952–1999 and 2000–2020. The vertical line in the top panel marks the start of the WWII period until the end of 1951, which is excluded from the analysis.



A.1 Data sources

Our main measure of inflation is the annualized quarterly rate of change in the Consumer Price Index (CPI). We use the seasonally adjusted CPI for All Urban Consumers (CPI-U) published by the Bureau of Labor Statistics and available from 1947 via St. Louis FRED. For the earlier part of the sample, we use seasonally unadjusted CPI-U series. The FRED tickers are CPIAUCSL and CPIAUCNS, respectively. The seasonally unadjusted CPI-U is the same series as used by Shiller (2016) to cover a long period starting from the late 1800s.

We construct the T-bill convenience spread following Nagel (2016). We download the T-bill convenience series from Stefan Nagel's website, https://voices.uchicago.edu/stefannagel/code-and-data/, link *"Time-Series of Liquidity Premia ..."*. This series is constructed as the spread between 3-month banker acceptance rate and 3-month T-bill rate before 1990, and the spread between 3-month term repo rate collateralized by Treasuries and 3-month T-bill rate after 1990. This repo series ends in 2011. Therefore, we rely on the 3-month asset-backed commercial paper rates to supplement the recent period afterward, which is ticker "RIFSPPAAAD90NB" in FRED. For the post-2011 data, we cross-check the 3-month commercial paper rates with 3-month repo rates from JP-Morgan markets (proprietary data), and find they are similar. For replicability, we use the publicly-available data on commercial paper rates. Appendix A.2 discusses the credit risk in commercial paper rates arguing that it is very small.

The short-term interest rate (denoted FFR) and the proxy for market volatility (denoted VIX) come from Nagel (2016), with his long sample period covering 1920 though 2011. For robustness, we reproduce his constructions with available data. We extend his sample though 2020 using VIX and effective fed funds rate from FRED.

For government debt supply, we use the total quantity of Treasury debt held by the public, at market value, minus intra-governmental holdings and holdings by depository institutions and the Federal Reserve. The data construction follows Krishnamurthy and Li (2023). Total debt held by the public can be obtained from FRED, ticker "FYGFDPUN", from 1970 to 2016. Before 1970, we use the total debt measure in Nagel (2016) (the same data source as T-bill convenience), which originally come from Bohn (2008). Next, we calculate net debt supply as the book value of total debt held by the public minus financial sector holding and Federal Reserve holdings of Treasuries, which leads to a measure of non-bank private sector holding of Treasuries. Then we translate the book value into market values using the market-to-book ratio of all marketable Treasury securities. Data on market and book values are provided by the Federal Reserve Bank of Dallas, https://www.dallasfed.org/research/econdata/govdebt. For monetary policy, we use the end-of-month effective federal funds rate, downloaded from FRED with ticker "FEDFUNDS".

For the analysis in Section 5, we report results using the 3-month Refcorp-Treasury spread. Refcorp bonds are bonds issued by Resolution Funding Corporation (Refcorp), a government agency created in 1989 to resolve the savings and loan crisis of the 1980s. Refcorp is explicitly guaranteed by the U.S. government, and thus, the Refcorp-Treasury spread is free from default risks. The Refcorp data is from Bloomberg, tickers "C091[maturity]".

In Section 5, we also use the current-quarter nowcast for the quarter-over-quarter inflation from the Blue Chip Financial Forecasts (BCFF). The nowcast is expressed in annualized units. BCFF forecasts are usually collected during the last week of the month (except for December, which can be earlier) and are published on the first day of the subsequent month. We merge the month in which the survey is conducted with the convenience spread at the end of that month.

A.2 Credit risk in convenience yield measures

In this appendix, we address a natural concern about our convenience-yield measure: whether it embeds a material credit-risk component. We assemble evidence and back-of-the-envelope calculations showing that any such component is negligible.

First, for the sample before 2011, we directly use the T-bill convenience yield measure from Nagel (2016). This measure before 1991 is based on banker's acceptance, which is double-backed by both the borrowing firm and the bank that "accepts" the banker's acceptance. In order to default, both the borrowing firm and the guaranteeing bank would need to default. Because a banker's acceptance formally is an unconditional liability of the bank, it has at least the same credit quality as banks' P1 commercial paper. A simple back-of-the-envelope calculation gives a ballpark upper bound for the magnitude of credit risk in banker's acceptance. According to Moody's Investors Service (2018), the 90-day cumulative default probability for P-1 issuers is 0.0080%. To calculate the expected loss-given-default rate in the data, we use the recovery rates L_{loss} for AAA-rated bonds reported by the last column of Exhibit 27 in (Emery et al., 2009), which is 85.55%. Then the expected-loss rate over 90 days is:

$$\text{EL}_{90} = 0.0080\% \times (1 - 0.8555) = 0.1156 \text{ bps} \tag{A1}$$

Annualizing this gives us $0.1156 * 4 \approx 0.46$ bps. Even doubling this to allow for a sizable risk premium leaves the credit component below 1 basis point. From 1991 to 2011, Nagel (2016) uses 3-month term repos backed by U.S. Treasuries. The credit risk in this measure is arguably even smaller, because even if the counterparty were to default, the lender formally owns the Treasury backing the contract, and therefore is guaranteed to obtain the Treasury bond if the counterparty defaults.

For the sample after 2011, we use asset-backed commercial paper (ABCP) rates. While there were issues with ABCP during the financial crisis due to ABCP backed by mortgage-backed securities, we only use this series after regulation was substantially strengthened to minimize credit risk in ABCP. Post-2011, ABCP has stronger liquidity support, dual ratings, and is backed by bank liquidity/credit lines and collateralized by traditional, self-liquidating receivables (e.g., trade receivables, auto loans/leases, credit-card receivables, equipment leases), putting P-1 ABCPs at par with P-1 commercial papers. By the calculation above, the default-loss component is economically negligible. We have compared ABCP against 3-month repo rates backed by Treasuries, similar to the Nagel (2016) data, using data from JP Morgan markets. Unfortunately, our data from JP Morgan markets is proprietary. For replicability, we use the publicly available ABCP spread after 2011.

Finally, when we use the Aaa–Treasury spread as a long-horizon convenience-yield proxy, one might worry about embedded credit risk. Exhibit 31 of (Emery et al., 2009) shows zero realized default for Aaa bonds over the sample and low transition rates out of Aaa, implying a negligible default probability even after accounting for rating migration. Moreover, structural calibrations in Huang and Huang (2012) show that expected default losses explain only a small fraction of investment-grade credit spreads, which is the well-known "credit spread puzzle." Hence, the Aaa–Treasury spread suffers minimal contamination from default risk. Huang and Huang (2012) also pointed out that "among IG bonds, the shorter the maturity, the smaller the fraction of the observed yield spreads due to credit risk," so the credit risk component is even smaller for the T-bill convenience yield compared to the already tiny fraction in longer-maturity credit spread of investment-grade bonds.

A.3 Controlling period-specific credit risk

Even though credit risk in the T-bill spread measure is negligible, as argued above, another way to assess whether credit risk might be driving our results is to control for it. We use the Baa-Aaa spread from Moody's and include in both Tables 1 and 2. In Appendix Table A2, we control for the credit spread in isolation (column (2)) and for its interactions with the period dummies. One might hypothesize that credit risk mattered more during some of our subperiods than others, in which case it might be important to allow it to enter differently for different subperiods. Column (1) in Appendix Table A2 is identical to column (1) of Table 1. Comparing across the three columns shows that the T-bill-inflation coefficient is unchanged if we control for the credit spread, and its interactions with period dummies. This, hence, further indicates that credit risk does not drive our results.

Table A2. Controlling period-specific credit risk. This table replicates the baseline specification column (1) in Table 1, adding the Baa spread and its interactions with period dummies as additional controls. $I_{1952-1999}$ and $I_{\geq 2000}$ are dummy variables taking the value of one in the indicated subperiod. The pre-WWII period (1923:01–1939:08) is the omitted category. Newey-West t-statistics with 12 lags are shown in parentheses. The stars indicate significance at * p <0.10, ** p <0.05, *** p <0.01 levels.

	(1) T-bill spread	(2) T-bill spread	(3) T-bill spread
Inflation	-0.0145***	-0.00743	-0.0129***
	(-3.87)	(-1.58)	(-3.47)
Inflation x $I_{1952-1999}$	0.119***	0.106***	0.100***
	(7.93)	(6.51)	(5.79)
Inflation x $I_{\geq 2000}$	0.0157**	0.0171***	0.0142*
	(2.29)	(2.88)	(1.80)
Baa spread		0.118**	0.0279
		(2.15)	(0.84)
Baa spread x $I_{1952-1999}$			0.295**
			(2.34)
Baa spread x $I_{\geq 2000}$			-0.0248
			(-0.40)
$I_{1952-1999}$	-0.105	0.0356	-0.285**
	(-1.56)	(0.33)	(-2.39)
$I_{\geq 2000}$	-0.0224	0.0619	0.0274
	(-0.46)	(0.86)	(0.27)
Constant	0.255***	0.0292	0.201**
	(6.33)	(0.24)	(2.49)
$\bar{R}^2$	0.426	0.443	0.468
N	1028	1028	1028

A.4 Robustness with different inflation measures

In our baseline specification, we measure inflation as the quarter-over-quarter change in the CPI. Once asset prices enter, this timing is imperfect: the 3-month convenience yield is forward-looking, whereas realized quarterly inflation is backward-looking. To better align horizons, we define forward quarterly inflation as the change in CPI from the current to the next quarter, and denote it as $\pi_{t,t+3}$. The realized inflation measure is denoted as $\pi_{t-3,t}$ accordingly. Moreover, we will also check robustness with year-over-year (YoY) inflation, both using the realized value in the last year (denoted as $\pi_{t-12,t}$) and forward value in the next year (denoted as $\pi_{t,t+12}$).

Table A3 reproduces columns (1) and (5) of Table 1 with four inflation measures: $\pi_{t-3,t}$, $\pi_{t,t+3}$, $\pi_{t-12,t}$, and $\pi_{t,t+12}$. Across specifications, the coefficient on first-period inflation is negative and significant, consistent with a liquidity-demand shock that lowers inflation while raising the convenience yield. The interaction of inflation with the second-period dummy is positive and significant, with similar magnitudes within comparable control sets (columns 1–4 vs. 5–8). Summing the baseline inflation coefficient and its second-period interaction (rows 1 and 2) implies a positive net relation between inflation and the convenience yield in the second period.

By contrast, the interaction with the third-period dummy is near zero on average. However, when we control for confounding factors and use forward inflation measures that better match the convenience yield's expectation timing (columns (6) and (8)), the third-period coefficient becomes even more negative than the first-period coefficient, reinforcing the argument that significant liquidity demand shocks are present in the third period.

Table A3. Robustness to different inflation measures. This table replicates the main Table 1 column (1) and (5), using alternative inflation measures. One time period is one month, so $\pi_{t-3,t}$ is the realized inflation (percentage CPI change) in the last three months and $\pi_{t,t+3}$ is the forward inflation in the next three months. Monthly data runs from 1923:01 through 2020:12, excluding the 1939:09–1951:12 period. $I_{1952-1999}$ and $I_{\geq 2000}$ are dummy variables taking the value of one in the indicated subperiod. The pre-WWII period (1923:01–1939:08) is the omitted category. Newey-West t-statistics with 12 lags are shown in parentheses. The stars indicate significance at * p <0.10, ** p <0.05, *** p <0.01 levels.

	T-bill spread							
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Inflation measure:	$\pi_{t-3,t}$	$\pi_{t,t+3}$	$\pi_{t,t-12}$	$\pi_{t,t+12}$	$\pi_{t-3,t}$	$\pi_{t,t+3}$	$\pi_{t,t-12}$	$\pi_{t,t+12}$
Inflation	-0.015***	-0.010**	-0.028***	-0.019**	-0.010***	-0.0040	-0.020***	-0.0094*
	(-3.87)	(-2.13)	(-5.29)	(-2.26)	(-3.74)	(-1.16)	(-3.25)	(-1.79)
Inflation*$I_{1952\text{-}1999}$	0.12***	0.10***	0.15***	0.11***	0.053***	0.027*	0.078***	0.031**
	(7.93)	(5.89)	(9.14)	(5.50)	(3.54)	(1.68)	(3.56)	(2.24)
Inflation*$I_{\geq 2000}$	0.016**	0.0021	0.066***	0.0012	0.0090	-0.011*	0.0094	-0.031**
	(2.29)	(0.21)	(3.97)	(0.07)	(1.28)	(-1.84)	(0.35)	(-2.49)
FFR					0.083***	0.093***	0.076***	0.096***
					(8.07)	(8.67)	(6.24)	(8.85)
Debt/GDP					0.20	0.15	0.23	0.19
					(0.96)	(0.70)	(1.03)	(0.86)
VIX					0.011***	0.011***	0.011***	0.011***
					(3.54)	(3.29)	(3.66)	(3.26)
Baa spread					-0.012	0.0089	-0.055	0.0074
					(-0.27)	(0.21)	(-1.05)	(0.18)
$I_{1952-1999}$	-0.10	-0.052	-0.16**	-0.049	-0.060	0.0046	-0.14	0.0016
	(-1.56)	(-0.68)	(-2.45)	(-0.63)	(-0.72)	(0.05)	(-1.42)	(0.02)
$I_{\geq 2000}$	-0.022	-0.0079	-0.089	0.023	0.061	0.13	0.039	0.18*
	(-0.46)	(-0.15)	(-1.63)	(0.39)	(0.71)	(1.44)	(0.34)	(1.96)
Constant	0.25***	0.26***	0.24***	0.25***	-0.26***	-0.32***	-0.18*	-0.34***
	(6.33)	(6.10)	(6.43)	(5.83)	(-3.02)	(-3.68)	(-1.87)	(-4.04)
$\bar{R}^2$	0.43	0.34	0.48	0.31	0.59	0.57	0.60	0.57
N	1028	1028	1028	1028	1028	1028	1028	1028

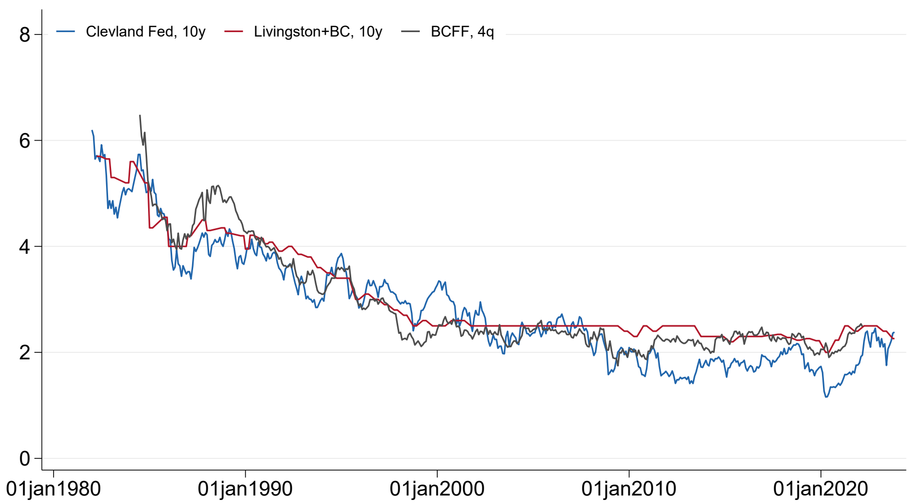
A.5 The Cleveland Fed index, inflation expectations, and term premia

We next investigate the suitability of the popular Cleveland Fed inflation expectations measures for analyzing convenience yields. This index is used, for example, in Fu et al. (2025). As stated by the Cleveland Fed "Our estimates are calculated with a model that uses Treasury yields, inflation data, inflation swaps, and survey-based measures of inflation expectations." While the Cleveland Fed uses a model that aims to separate term premia from expectations, such decompositions are somewhat reliant on the specific modeling choices and there is no guarantee that the resulting inflation expectations measure is indeed free of term premia and convenience.

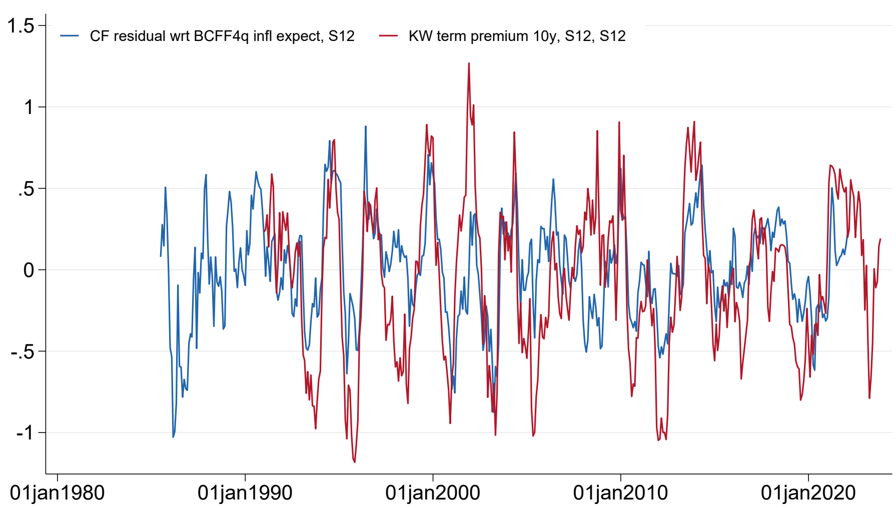
Panel A of Figure A2 shows 10-year inflation expectations from the Cleveland Fed against long-term and 4-quarter consensus inflation expectations from surveys. It is clear that the Cleveland Fed inflation expectations roughly move similarly at lower frequencies, but are substantially more volatile, raising the question whether being derived from bond yields they still contain time-varying term premia and, potentially, Treasury convenience.

Panel B of Figure A2 shows 12-quarter changes in the 10-year Cleveland Fed inflation forecast, residualized against survey expectations, together with a measure of contemporaneous changes in 10-year term premia from Kim and Wright (2011). The correlation is very high at 60%, further confirming the likely presence of term premia and convenience in the Cleveland Fed inflation expectations. While this may not be an issue if the objective is merely to obtain an unbiased forecast of long-term inflation, in a regression of a Treasury convenience spread on the left-hand side and the Cleveland Fed inflation expectations on the right-hand-side, it is likely to bias the results towards finding a negative regression coefficient. The intuition is simply that a shock that lowers the 10-year Treasury bond yield due to term premia or increased convenience, is likely to lower the Cleveland Fed measure even if inflation expectations truly did not move.

Figure A2. Cleveland Fed inflation expectations vs. inflation expectations and term premia. Panel A shows the 10-year inflation forecast from the Cleveland Fed model against 10-year inflation expectations from Blue Chip and Livingston surveys, and the 4-quarter consensus CPI inflation forecast from the Blue Chip Financial Forecasts. Cleveland Fed inflation expectations start in 1982:Q1, Livingston/Blue Chip forecasts start in 1982:Q1, and BCFF 4-quarter forecasts start in 1984:Q3. The sample ends in 2023:Q4. Panel B plots the residual from a regression of 12-month changes in Cleveland Fed inflation expectations onto BCFF 4-quarter inflation expectations against 12-month changes in the 10-year term premium from Kim and Wright (2011). Long-term CPI inflation forecasts from Blue Chip and Livingston surveys are available via the inflation-forecast website of the Philadelphia Fed.



(a) Cleveland Fed inflation expectations vs. surveys



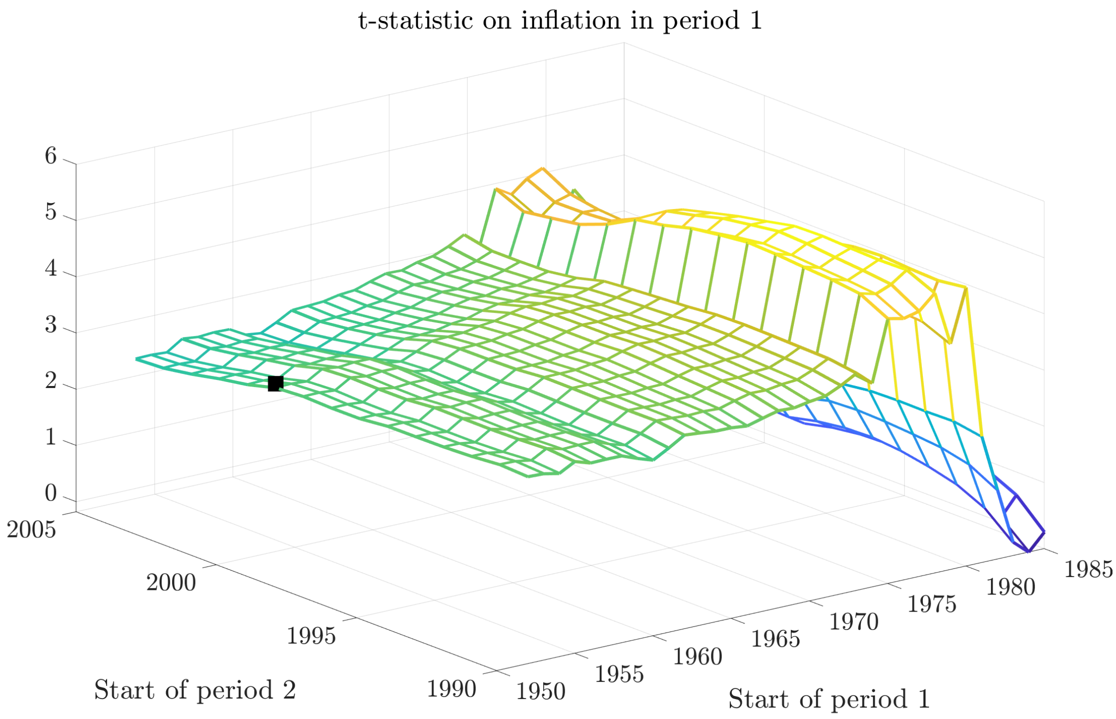
(b) Cleveland Fed inflation expectations vs. term premia

A.6 Robustness for different break dates

In this subsection, we present robustness to the timing of the break dates in our baseline regressions in Table 2. Figure A3 plots the t-statistics for the inflation loading of T-bill spread from the baseline regression (1), where the regression is estimated with different starting dates from 1952 through 1985, with the second period starting between 1990 and 2005:

$$\textit{T-bill spread}_t = b_0 + b_1\pi_t + b_2\pi_t \times I_{\text{Start of period 2},t} + \Gamma' X_t + \varepsilon_t \text{ if year} > \text{Start of period 1} \quad \text{(A2)}$$

Figure A3. T-statistics for T-bill loading on inflation for different period start dates. This figure reports results for the baseline regression in column (3) of Table 2 using different sample start dates ranging from 1952 to 1985 (start of period 1) and varying the cutoff dates for the second subperiod (start of period 2) between 1990 and 2004. The t-statistics is reported for the b_1 coefficient in regression equation (A2) and is based on Newey-West standard errors with 12 lags. Our baseline result from Table 2—sample starting in 1952 and break in 2000—is indicated with a black square.



The specification is identical to column (3) of Table 2, except that we vary subsample cutoffs.

The t-statistics on b_1 coefficient shows that the T-bill loading on inflation in the first period is positive and significant for a broad range of start dates in the 1950s and 1960s, consistent with the Inflation coefficient in Table 2. The t-statistic only drops below two if we start the sample as late as 1980, which is not surprising, because this misses the entire inflationary realizations of the 1970s. So, overall, the positive convenience-inflation relationship in the second half of the 20th century is robust to different break dates.

A.7 Deposit rate process

We directly estimate the deposit adjustment process in the theory using data from CALL reports. We define the deposit rate as the total checking deposit interest expense divided by the total dollar value of checking deposits. This measure reflects the average rate paid by the banking sector. The data are quarterly from 1987 Q1 to 2020 Q1, downloaded from WRDS. In Table A4, we regress the deposit rate on its one-quarter lag and the contemporaneous three-month T-bill rate. The regression yields an R^2 of 98.4%, suggesting that equation (7) closely fits the data. The deposit-rate stickiness is $\rho^d = 0.92$ and the implied passthrough coefficient is 0.023, similar to the main calibrated value of δ in our model. Results are similar if we use FFR instead of the T-bill rate.

Table A4. Deposit rate adjustment dynamics. This table presents estimates of how deposit rates respond to market rates and their own lagged values. The dependent variable is the deposit rate, defined as total interest payments on checking deposits divided by checking deposit balances. The sample covers 1987Q1–2020Q1 using quarterly Call Report data. Newey-West standard errors with 12 lags are in parentheses. Statistical significance: $^{*}p < 0.1$; $^{**}p < 0.05$; $^{***}p < 0.01$.

Variable	**Coefficient**
3-month T-bill rate	0.023**
	(0.009)
Lagged rate	0.918***
	(0.030)
Constant	−0.014
	(0.016)
Observations	132
R^2	0.984

A.8 Long-term convenience spread adjustments

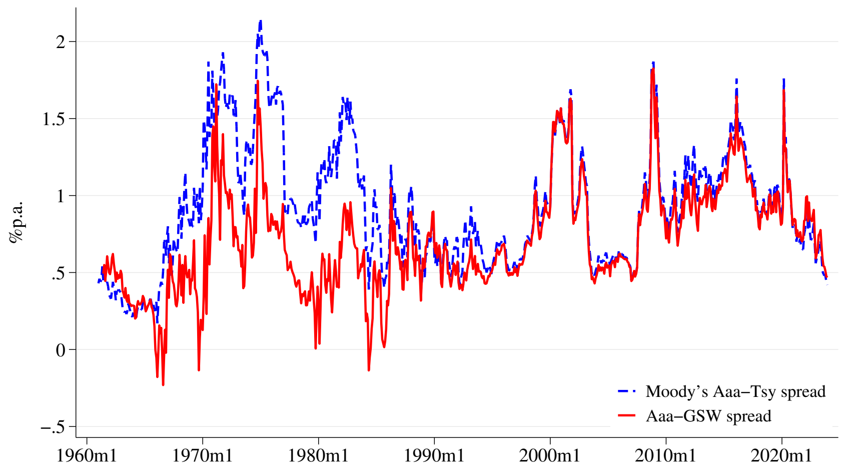
While in our main analysis we primarily focus on the T-bill spread, reflecting short-term convenience, it is worth discussing the long-term convenience spread as well. Much of the literature uses the Moody's Aaa-Treasury spread as a measure of long-term convenience, in the absence of alternatives available over the long historical sample (Krishnamurthy and Vissing-Jorgensen, 2012). However, the Moody's Aaa-Treasury spread can be confounded by other factors unrelated to convenience, as we acknowledge in Section 2.3. The issues pertain to the flower bond clauses in Treasury bonds, the callability of corporate bonds, and the duration mismatch between the Treasury and corporate bonds. In Table 2, we compare our baseline regressions using T-bill, the original Moody's Aaa-Treasury spread and the adjusted GSW-Treasury spread to account for some of these confounding factors. Below, we discuss the adjustments we undertake in more detail.

As recently highlighted by Lehner et al. (2025), the benchmark long-term Treasury yield used by Moody's includes the so-called flower bonds.[30] Flower bonds offer bondholders additional benefits similar to life insurance as they could be redeemed at par to cover the payment of estate taxes upon the holder's death. They were issued before 1966 with coupon rates below 4.5% and were effectively the only US long-term government bonds available in the early part of the sample. The analysis of Mayers and Smith (1987) indicates that the option became especially valuable when interest rose in the second half of the 1970s. This effect depresses the Treasury yield used by Moody's, and therefore overstates the value of Treasury convenience. Accordingly, using a carefully constructed Treasury benchmark that excludes flower bonds, Lehner et al. (2025) show that the long-term Aaa-Treasury convenience spread was significantly lower from the mid-1970s through the early 1980s than the Moody's Aaa spread would imply.

Flower bonds. To study the relationship between inflation and long-term convenience, we construct the long-term Treasury yield benchmark using the Gürkaynak et al. (2007, GSW) Treasury yield dataset. Importantly, GSW exclude securities with option-like features, including callable bonds and flower bonds. The GSW sample starts in June 1961. Between June 1961 and August 1971, the maximum available maturity is 7 years; it extends to 10 years in August 1971, to 15 years in November 1971, and to 20 years in July 1981. Since Moody's Aaa index is based on bonds with a remaining maturity of at least 20 years, we approximate the long-term government counterpart with a 20-year GSW par yield when it becomes available. For the prior years, we use

[30] Until June 2000, Moody's Aaa spread uses the yield on long-term US government securities from the Federal Reserve's G.13 statistical release, available via FRED with ticker LTGOVTBD.

Figure A4. Comparison of Moody's Aaa spread with Aaa–GSW spread. This figure compares the baseline Moody's Aaa-Treasury spread (dashed blue line) to an alternative constructed using Gürkaynak et al. (2007) yields, excluding flower bonds (solid red line). For the Aaa–GSW spread, we approximate the long-term government yield with a 20-year par yield from Gürkaynak et al. (2007) when it becomes available, and the longest available par yield prior to that.



the longest available maturity in GSW, assuming a flat par yield curve beyond the last available maturity knot.[31]

We denote the spread between the Moody's Aaa yield and the long-term GSW par yield as the Aaa–GSW spread. Figure A4 shows that the Aaa-GSW spread broadly replicates the findings reported by Lehner et al. (2025). It traces closely the Moody's Aaa-Treasury spread for the post-2000 period, but implies a lower spread during the 1970s and 1980s, although the two spreads remain highly correlated in the pre-2000 period as well, with a correlation of 0.7. The respective average spreads are 89 and 57 bps in the 1961:06–1999:12 sample.

Duration match. Further, we verify the quality of the duration match between long-term corporate bonds and the approximate long-term GSW par bond. Since we do not have historical market prices for the constituents of the Moody's Aaa index, we rely on the duration estimates from van Binsbergen et al. (2025) in the combined Lehman/Warga and Bank of America Merill Lynch datasets. This data becomes available from March 1974. The time-series correlation

[31] Alternatively, one could extrapolate the 20-year par yield using the Nelson-Siegel parameters reported by GSW to backfill the missing observations before 1981. However, GSW advise against this approach as the distant extrapolated yields can become unreliable. We verify that while extrapolation typically generates a nearly flat term structure, there are a few data points where extrapolated yields are economically meaningless.

between the average duration of Aaa corporate bonds with at least 20 years to maturity and the 20-year par bond duration estimated using the long-term GSW par yield[32] is 0.99 in levels and 0.79 in monthly changes. The average absolute duration difference is 0.35 years with a standard deviation 0.3 years.[33]

Callability of corporate bonds. Another factor affecting long-term convenience yield estimates are the call options embedded in corporate bonds. The call option allows the issuer to redeem the bond before maturity which becomes especially valuable when interest rates fall. Callability was a common feature of corporate bonds until the late 1990s, when after a flurry of redemptions, make-whole provisions became widespread, offering additional protection for bondholders (Brown and Powers, 2020).[34] The estimates by Duffee (1998) for the 1985–1995 sample imply that a 100 bps decrease in Treasury yields raises the long-term Aaa-Treasury spread by approximately 20 bps.[35] Thus, the call option could lead to an overstatement of the convenience spread in the post-Great Inflation period as interest rates fell through the late 1990s, but an understatement of the convenience spread as interest inflation and interest rates rose through the 1970s. By not adjusting for the call options in corporate bonds, the baseline Aaa-Treasury spread might hence over- or under-state the rise in convenience during the high-inflation period in the 1970s and 1980s.[36]

To adjust the convenience spread for the presence of the call option, we follow Gilchrist and Zakrajšek (2012) who, building on Duffee (1998), adjust for the moneyness of the call option with the yield curve level, slope, and interest rate volatility. We project the Aaa–GSW spread on these covariates[37] and use the fitted value as a proxy for the variation in the call, which we subtract from the Aaa–GSW spread. We denote this variable as "Aaa–GSW (ca) spread," as analyzed in Table 2.

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[32] We calculate the par bond duration as $D_n = 0.5\frac{1-(1+py_{n,t}/2)^{-2n}}{1-(1+py_{n,t}/2)^{-1}}$, where $py_{n,t}$ is the par yield on a bond with n-years to maturity paying semiannual coupons (see Campbell et al. (1997), equation (10.1.18)). We set $n = 20$ years. Before 1981, when the 20-year par yield becomes available in the GSW data, we approximate its level with the longest available par yield maturity, assuming flat par yield term structure beyond the maximum available maturity.

[33] We thank Yoshio Nozawa for providing the corporate bond duration estimates.

[34] A make-whole provision gives the issuer the option to call the bond early, but unlike traditional call options with fixed prices, the issuer must "make the investor whole" by paying the present value of all remaining coupon and principal payments the corresponding Treasury yield plus a make-whole spread.

[35] Make-whole provisions are common in investment-grade corporate bonds after late 1990s to early 2000s, following a flurry of bond redemptions .

[36] The call option could serve as inflation protection for the bondholder in the high inflation episode. Indeed, the option-adjusted estimates of bond excess premium in Gilchrist and Zakrajšek (2012) suggest that those spreads were higher in the late 1970s and early 1980s than the unadjusted spreads.

[37] We include the one-year yield as level, the spread between the 7-year yield and the one-year yield as slope, and the realized monthly volatility of the 7-year yield changes. All variables are constructed using GSW data.

The adjustments above highlight significant limitations in the long-term Treasury and corporate bond data in the early parts of our sample and motivate our focus on the short-term convenience which is not subject to those concerns.

A.9 Post-COVID period inflation and convenience

In Section 5, we show that the post-COVID period is consistent with supply shocks generating a positive inflation–convenience relationship, whereas during COVID the relationship is strongly negative, consistent with liquidity-demand shocks lowering inflation. In this appendix, we confirm the robustness of this result using alternative measures of convenience yield and inflation, and by controlling for key confounders such as the contemporaneous policy rate and the debt-to-GDP ratio.

Table A5 reports the formal regressions. Panel A regresses T-bill convenience on expected inflation, controlling for the policy rate and debt-to-GDP. We use two T-bill convenience measures: the 3-month ABCP–T-bill spread and the 3-month Refcorp STRIPS–T-bill spread. Both ABCP and Refcorp STRIPS are less liquid than T-bills and have negligible credit risk, so these spreads capture T-bill convenience. For both measures, the coefficient on expected inflation is negative during COVID and positive post-COVID, and its magnitude is similar with and without the debt-to-GDP control.

Panel B examines long-maturity Treasury convenience. We use the 10-year agency–Treasury STRIPS spread and the 10-year Refcorp–Treasury STRIPS spread, each based on maturity-matched zero-coupon securities. Agency and Refcorp bonds carry strong federal backing and risk profiles close to Treasuries, so these spreads provide clean measures of Treasury convenience. The results mirror Panel A: the coefficient on expected inflation is negative during COVID and positive post-COVID, and adding the debt-to-GDP control does not change these findings.

Table A5. Post-COVID inflation and convenience. This table presents estimates that regress various measures of short-term and long-term Treasury convenience yields onto inflation. The pre-sample is monthly from 2018:01 to 2020:12, which captures the initial COVID-19 shock, and the post-sample is monthly from 2021:01 to 2023:12, which captures the post-COVID inflation. E[inflation] is the nowcast of current-quarter inflation from Blue Chip Financial Forecasts. ABCP-Tbill 3M is the yield spread between three-month asset-backed commercial papers and three-month T-bills. Refcorp-Tbill 3M is the yield spread between three-month Refcorp STRIPs and three-month T-bills. Agency-Tsy STRIP 10Y is the yield spread between ten-year agency STRIPs (zero-coupon) and ten-year Treasury STRIPs (zero-coupon). Refcorp 10Y represents ten-year Refcorp STRIPs (zero-coupon). Debt/GDP is the ratio between total face value of government debt outstanding and nominal GDP. All regressions cover a 36-month period ($N = 36$). Newey-West t-statistics with 12 lags are shown in parentheses. Constant terms are included in regressions but not reported for conciseness. The stars indicate significance at * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$ levels.

Panel A: Short-Term Convenience Yields

	ABCP-Tbill 3M				Refcorp-Tbill 3M			
	Pre (2018-2020)		Post (2021-2023)		Pre (2018-2020)		Post (2021-2023)	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
E[inflation]	-0.054	-0.069***	0.069***	0.058***	-0.038**	-0.040**	0.12***	0.100***
	(-1.06)	(-2.94)	(4.38)	(3.12)	(-2.32)	(-2.57)	(4.42)	(3.61)
FFR	0.025	-0.31***	0.0099	0.0045	-0.11***	-0.13**	0.076***	0.068***
	(0.80)	(-3.69)	(1.16)	(0.57)	(-6.48)	(-2.64)	(5.76)	(4.98)
Debt/GDP		-3.51***		-2.06		-0.30		-3.16
		(-4.47)		(-1.26)		(-0.64)		(-1.58)
$\bar{R}^2$	0.045	0.66	0.24	0.25	0.61	0.61	0.43	0.44
N	36	36	36	36	36	36	36	36

Panel B: Long-Term Convenience Yields

	Agency-Treasury STRIP 10Y				Refcorp-Treasury STRIP 10Y			
	Pre (2018-2020)		Post (2021-2023)		Pre (2018-2020)		Post (2021-2023)	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
E[inflation]	-0.039***	-0.041***	0.022***	0.018**	-0.027**	-0.026*	0.055***	0.039***
	(-4.06)	(-4.57)	(4.14)	(2.31)	(-2.19)	(-1.99)	(7.06)	(4.10)
FFR	-0.031**	-0.062**	0.045***	0.043***	-0.055**	-0.023	0.045***	0.037***
	(-2.36)	(-2.44)	(11.19)	(9.45)	(-2.21)	(-0.88)	(6.10)	(5.64)
Debt/GDP		-0.34		-0.75		0.34		-3.15***
		(-1.36)		(-0.79)		(1.44)		(-2.90)
$\bar{R}^2$	0.66	0.68	0.72	0.72	0.52	0.53	0.49	0.56
N	36	36	36	36	36	36	36	36

B Model Appendix

In this appendix, we provide details of the model assumptions, log-linearization, numerical solution and estimation. Appendix B.8 provides additional model impulse responses for supply and monetary policy shocks. Appendix B.9 provides model extensions. Throughout the Appendix we use the notation $\tilde{\lambda}(s_t)$ for the sensitivity function for consistency with the code and the prior literature. In the main paper, the sensitivity is denoted by $\omega(s_t)$ to more clearly distinguish is from the liquidity shock, which is λ_t.

B.1 Model setup

B.1.1 Final good production

A final consumption good is produced by a representative perfectly competitive firm from a continuum of differentiated goods $Y_{i,t}$:

$$Y_t = \left(Y_{i,t}^{\frac{\epsilon_{p,t}-1}{\epsilon_{p,t}}} \right)^{\frac{\epsilon_{p,t}}{\epsilon_{p,t}-1}}. \tag{A3}$$

Here $\epsilon_{p,t} > 1$ is the elasticity of substitution across intermediate goods, which provides the source of supply shocks in the model. The resulting demand for the differentiated good i is downward-sloping in its product price $P_{i,t}$:

$$Y_{i,t} = Y_t \left(\frac{P_{i,t}}{P_t} \right)^{-\epsilon_{p,t}}. \tag{A4}$$

The aggregate price level is given by

$$P_t = \left(\int_0^1 P_{i,t}^{-(\epsilon_{p,t}-1)} di \right)^{-\frac{1}{\epsilon_{p,t}-1}}. \tag{A5}$$

B.1.2 Intermediate good producers

Intermediate goods firm i produces according to a Cobb-Douglas production function with capital share τ

$$Y_{i,t} = A_t N_{i,t}^{1-\tau}, \tag{A6}$$

where productivity equals A_t and N_t is the supply of the aggregate labor index. Each firm takes the downward-sloping demand schedule as given by (A5) and may choose a different amount of the aggregate labor index. With the final good equation (A3) aggregate output equals

$$Y_t = A_t N_t^{1-\tau}, \tag{A7}$$

where aggregate labor is defined:

$$N_t \equiv \left[\int_0^1 N_{i,t}^{\frac{(\epsilon_{p,t}-1)(1-\tau)}{\epsilon_{p,t}}} di \right]^{\frac{\epsilon_{p,t}}{(\epsilon_{p,t}-1)(1-\tau)}}. \tag{A8}$$

The labor provided by different households is assumed to be perfectly substitutable, so households take the real wage as given. The aggregate resource constraint in this economy is simple. Because there is no time-varying real investment, consumption equals output $C_t = Y_t$. Following Lucas (1988) and Campbell et al. (2020) we assume that productivity depends on past skills gained by all agents, and depends on past market labor, n_{t-1}:

$$a_t = \nu + a_{t-1} + (1-\phi)(1-\tau)n_{t-1}, \tag{A9}$$

where $0 \leq \phi < 1$ and $\nu > 0$ are constants. The assumption (A9) ensures that potential output increases with past output.

B.1.3 Price setting

Intermediate firms face standard price-setting frictions in the manner of Calvo (1983), where a fixed fraction of firms can change prices every period with equal probabilities across firms. When firms cannot update, their prices are indexed to lagged inflation (Smets and Wouters (2007), Christiano, Eichenbaum, and Evans (2005)). A firm that last reset its price at time t to $\tilde{P}_t$, charges a nominal time $t+j$ price $\tilde{P}_t\left(\frac{P_{t-1+j}}{P_{t-1}}\right)$. A firm that can update its product price maximizes the discounted sum of current and future expected profits while the price is expected to remain in place, discounted at the households' stochastic discount factor. For simplicity, price-setters are assumed to have rational inflation expectations.

B.2 Household preferences and SDF

In the main paper, we use $U(C_t, Q_t, H_t)$ to denote household utility, which is sufficient to describe the asset pricing Euler equation and intertemporal consumption choice. However, for a full micro-foundation, we need to consider the impact of habit on labor choice. For that purpose, we use a richer utility setup that leads to the same households consumption Euler equation but cancels out the impact of habit on labor choice. The assumptions on labor-leisure largely follow Pflueger and Rinaldi (2022). Specifically, we assume that household h maximizes

$$E_0 \sum_{t=0}^{\infty} \beta^t U\left(C_{h,t}, C_{h,t}^{home}, Q_{h,t}, H_t, H_t^{home}\right), \tag{A10}$$

where per-period utility equals

$$U\left(C_{h,t}, C_{h,t}^{home}, Q_{h,t}, H_t, H_t^{home}, \Theta_t\right) = \frac{\left((C_{h,t} - H_t) + (C_{h,t}^{home} - H_t^{home})\right)^{1-\gamma}}{1-\gamma} + \alpha \log Q_{h,t} \tag{A11}$$

Here, $C_{h,t}$ denotes market consumption, H_t denotes habit over market consumption. We introduce a labor-leisure choice in the manner of Greenwood et al. (1988). Introducing a home habit ensures that the surplus consumption ratio, s_t, enters into asset prices but does not enter as a state-variable into firms' profit maximization, and giving us a standard Phillips curve. $C_{h,t}^{home}$ denotes consumption of individual household and H_t denotes habit over individual household's consumption. Habit H_t is external and is taken as given by household h. Non-market or home consumption equals

$$C_{h,t}^{home} = A_t \frac{\int_{i=0}^{1} \left(1 - \frac{L_{h,i,t}^{\eta}}{1-\eta}\right) di}{1-\eta}, \tag{A12}$$

and external non-market consumption habit equals H_t^{home}, where $N_{h,i,t}$ is the labor provided by household h to intermediary producing firm i and A_t is aggregate productivity. It is assumed that in equilibrium $H_t^{home} = C_t^{home}$. As a result, this term determines the labor-leisure choice, but drops out of equilibrium preferences over market consumption, and hence the intertemporal trade-off determining asset prices. Since all households are identical, they choose the same market and home consumption at all times and we drop the subscripts h from now on to save on notation.

Canceling out the internal habit terms in equilibrium, we obtain the stochastic discount factor (SDF) M_{t+1} as given in equation (5) in the main paper

$$M_{t+1} = \beta \frac{\frac{\partial U_{t+1}}{\partial C}}{\frac{\partial U_t}{\partial C}} = \beta \exp\left(-\gamma(\Delta s_{t+1} + \Delta c_{t+1})\right). \tag{A13}$$

B.2.1 Habit dynamics

Habit dynamics are assumed to be given by (6) with the output gap-consumption link (9). The sensitivity function is denoted as $\omega(s_t)$ in equation (6), but here for consistency with the habit literature, we use the notation $\tilde{\lambda}(s_t)$ (the tilde differentiates itself from convenience yield demand λ_t), and we assume it to take the form first proposed by Campbell and Cochrane (1999):

$$\tilde{\lambda}(s_t) = \begin{cases} \frac{1}{\bar{S}}\sqrt{1-2(s_t-\bar{s})}-1 & s_t \leq s_{max} \\ 0 & s_t > s_{max} \end{cases}, \tag{A14}$$

$$\bar{S} = \sigma_c \sqrt{\frac{\gamma}{1-\theta_0}}, \tag{A15}$$

$$\bar{s} = \log(\bar{S}), \tag{A16}$$

$$s_{max} = \bar{s} + 0.5(1-\bar{S}^2). \tag{A17}$$

B.2.2 Household and government budget constraints

We assume that the fiscal authority issues bonds and uses lump-sum takes to repay these bonds in each period. Long-term bonds are in zero-net supply for concreteness and the amount of one-period bonds is exogenously given. The central banks sets the amount of deposits in the economy, by purchasing one-period government bonds and issuing money. We assume that the fiscal authority correspondingly issues more, leaving $B_{1,t}$ constant. For simplicity we assume deposits to be equal to the monetary aggregate, assuming a reserve constraint equal to one. The government's real budget constraint then becomes

$$\begin{aligned} B_t + D_t &= \frac{P_{t-1}}{P_t} B_{1,t-1}(1+R^b_{t-1}) + \frac{P_{t-1}}{P_t} D_{t-1} \\ & +\frac{P_{t-1}}{P_t} \sum_{i=2}^{\infty} \left(B_{i,t-1}(1+R^b_{i,t}) \right) - T_t, \end{aligned} \tag{A18}$$

where P_t is the aggregate price level in the economy at time t, T_t denotes real lump-sum taxes and the total face value of government debt equals

$$B_t = B_{1,t} + B_{2,t} + ... + B_{i,t}. \tag{A19}$$

$R^b_{i,t}$, denotes the nominal returns from buying an i-period bond, at time $t-1$ and selling it again at time t. In equilibrium $B_{i,t} = 0$ for $i \geq 2$, so the face value of government bonds equals the face value of short-term bonds.

The representative household's budget constraint can then be written as

$$\begin{aligned} D_t + B_t - L_t + C_t + T_t = & \frac{W_t}{P_t} N_t + \Pi_t + \frac{P_{t-1}}{P_t} D_{t-1}(1+R^d_{t-1}) \\ & + \frac{P_{t-1}}{P_t} B_{1,t-1}(1+R^b_{t-1}) - \frac{P_{t-1}}{P_t} L_{t-1}(1+R^l_{t-1}) \\ & + \frac{P_{t-1}}{P_t} \sum_{i=2}^{\infty} \left(B_{i,t-1}(1+R^b_{i,t}) \right), \end{aligned} \tag{A20}$$

where Π_t is the sum of firm and bank profits remitted to the household sector, W_t is nominal wages, and N_t is labor supply.

Amount of deposits, D_t is chosen by the central bank to satisfy the policy rate (8) subject to households' downward-sloping demand function, given by the Euler equation for liquid bonds (11).

B.3 Model Solution

B.3.1 Liquidity spread

Taking the difference between (11) vs. (10) and (12) vs. (10) gives the following expressions

$$\frac{I_t^l - I_t^b}{1 + I_t^l} = \frac{\frac{\alpha}{Q_t}\lambda_t}{U_c\left(C_t, H_t\right)} \tag{A21}$$

$$\frac{I_t^l - I_t^d}{1 + I_t^l} = \frac{\frac{\alpha}{Q_t}(1 - \lambda_t)}{U_c\left(C_t, H_t\right)}. \tag{A22}$$

Dividing (A21) with (A22) and substituting in (7) into (A22) give (13) in the main text.

B.3.2 Steady-state interest rates

We log-linearize the model around the flexible-price steady-state values $\bar{c}$, $\bar{\pi}$, $\bar{\imath}^l$, $\bar{\imath}^b$, $\bar{\theta}$, and $\bar{\lambda}$ with deviations c_t, π_t, i_t^l, i_t^b, θ_t, and λ_t. We define the log steady-state interest rates by $\bar{\imath}^l = \log(1 + \bar{I}^l)$, $\bar{\imath}^b = \log(1 + \bar{I}^b)$, $\bar{\imath}^d = \log(1 + \bar{I}^d)$. Also define the log deviations of interest rates from their steady states as

$$i_t^l = \log\frac{1 + I_t^l}{1 + \bar{I}^l}, \quad i_t^b = \log\frac{1 + I_t^b}{1 + \bar{I}^b}, \quad i_t^d = \log\frac{1 + I_t^d}{1 + \bar{I}^d}. \tag{A23}$$

The relationship between the illiquid steady-state real risk-free rate, the discount factor and other preference parameters is identical to Campbell and Cochrane (1999) and given by

$$\bar{r}^l = -\log\beta + \gamma g - \frac{1}{2}\gamma^2\sigma_c^2/\bar{S}^2. \tag{A24}$$

The steady-state illiquid nominal rate is given by

$$\bar{\imath}^l = \bar{r}^l + \bar{\pi}, \tag{A25}$$

where $\bar{\pi} = \log(1 + \bar{\Pi})$.

We now move to the steady-state values for the three different types of nominal interest rates in our model. The steady-state deposit process in (7) implies

$$\bar{I}^d = \frac{\delta}{1 - \rho^d}\bar{I}^l. \tag{A26}$$

And the steady-state convenience-yield according to (13) is

$$\bar{I}^l - \bar{I}^b = \frac{\bar{\lambda}}{1-\bar{\lambda}} \left((1-\delta)\bar{I}^l - \rho^d \bar{I}^d\right), \tag{A27}$$

Combining (A26) and (A27), we obtain the steady-state convenience yield as

$$\bar{I}^l - \bar{I}^b = \frac{\bar{\lambda}}{1-\bar{\lambda}} \left(1 - \frac{\delta}{1-\rho^d}\right) \bar{I}^l, \tag{A28}$$

and the steady-state policy rate

$$\bar{I}^b = \left(1 - \frac{\bar{\lambda}}{1-\bar{\lambda}} \left(1 - \delta \frac{1}{1-\rho^d}\right)\right) \bar{I}^l. \tag{A29}$$

The steady-state liquid log real rate then equals

$$\begin{aligned} \bar{r}^b &= \bar{i}^b - \bar{\pi}, && \text{(A30)} \\ &= \log\left(1+\bar{I}^b\right) - \bar{\pi}. && \text{(A31)} \end{aligned}$$

B.3.3 Output gap and Phillips curve

The log real output gap is defined as the difference between log real output and log real output in the absence of price-setting frictions. With the assumptions given above, it equals to stochastically detrended consumption, i.e. (9) in the main paper, as shown in Pflueger and Rinaldi (2022). Log-linearizing the intermediate firms' optimal price-setting condition gives the standard Phillips curve (21) in the main paper (e.g. Walsh (2017)) with

$$\rho^\pi = \frac{1}{1+\beta_g}, \; f^\pi = 1 - \rho^\pi, \; \beta_g = \beta \exp(-(\gamma-1)g). \tag{A32}$$

The Phillips curve shock $v_{\pi,t}$ arises from variation in the elasticity of substitution across intermediary goods $\epsilon_{p,t}$, similar to markup shocks. The Phillips curve slope, κ is an endogenous parameter that depends on households' labor-leisure choice, the frequency of price-setting, and steady-state markups. However, since we do not link κ back to these more fundamental parameters, we do not spell out these links here.

B.3.4 Consumption Euler equation

The Euler equation for the illiquid one-period real rate r_t^l is our starting point:

$$E_t\left[M_{t+1}\exp(r_t^l)\right] = 1 \tag{A33}$$

We make the assumption as in Campbell et al. (2020) that one-period log nominal yields can be approximated using the Fisher equation

$$\begin{aligned} r_t^l &= i_t^l - E_t\pi_{t+1}, & \text{(A34)} \\ r_t^b &= i_t^b - E_t\pi_{t+1}. & \text{(A35)} \end{aligned}$$

These approximations are appropriate if inflation risk premia on one-quarter bonds are small. We do not make these assumptions for two- and longer-term bonds.

Using (A13), we further expand (A33) as

$$0 = r_t^l - \gamma E_t\Delta c_{t+1} - \gamma E_t\Delta s_{t+1} + \frac{\gamma^2}{2}\left(1+\tilde{\lambda}(s_t)\right)^2\sigma_c^2 \tag{A36}$$

up to a constant. Substituting in for the sensitivity function $\tilde{\lambda}(s_t)$, using $E_t\Delta c_{t+1} = E_t x_{t+1} - \phi x_t$ and the definition of $v_{x,t}$ in (22), we get the exactly log-linear Euler equation with a non-liquidity demand shock

$$\begin{aligned} x_t &= f^x E_t x_{t+1} + \rho^x x_{t-1} - \psi r_t^l + v_{x,t}, & \text{(A37)} \\ f^x &= \frac{1}{\phi-\theta_1},\ \rho^x = \frac{\theta_2}{\phi-\theta_1},\ \psi = \frac{1}{\gamma(\phi-\theta_1)}, & \text{(A38)} \end{aligned}$$

As in Campbell et al. (2020) and Pflueger (2025), we impose the parameter restriction that $f^x = 1-\rho^x$. In (A37), the demand shock $v_{x,t}$ can be interpreted as a standard discount rate shock, or shock to intertemporal substitution. Substituting in from the Fisher equation (A34) gives equation (19) in the main paper.

B.3.5 Log-linearized convenience yield dynamics

We now derive the dynamics of the convenience yield in (18). We note that the log-linearized rates are

$$I_t^j = (1 + \bar{I}^j) i_t^j + \bar{I}^j, \tag{A39}$$

with $j \in \{l, b, d\}$. Linearizing (13) around the steady state gives:

$$\left(1+\bar{I}^l\right) i_t^l + \bar{I}^l - \left(1+\bar{I}^b\right) i_t^b - \bar{I}^b = \left(\frac{\bar{\lambda}}{1-\bar{\lambda}} + \frac{1}{(1-\bar{\lambda})^2}\hat{\lambda}_t\right)\left((1-\delta)\left((1+\bar{I}^l) i_t^l + \bar{I}^l\right) - \rho^d\left((1+\bar{I}^d) i_{t-1}^d + \bar{I}^d\right)\right),$$

$$\begin{aligned}\left(1+\bar{I}^l\right) i_t^l + \bar{I}^l - \left(1+\bar{I}^b\right) i_t^b - \bar{I}^b = & \frac{\bar{\lambda}}{1-\bar{\lambda}}\left((1-\delta)\left((1+\bar{I}^l) i_t^l + \bar{I}^l\right) - \rho^d\left((1+\bar{I}^d) i_{t-1}^d + \bar{I}^d\right)\right) \\ & + \frac{1}{(1-\bar{\lambda})^2}\hat{\lambda}_t\left((1-\delta)\bar{I}^l - \rho^d \bar{I}^d\right),\end{aligned}$$

$$\left(1+\bar{I}^l\right) i_t^l - \left(1+\bar{I}^b\right) i_t^b = \frac{\bar{\lambda}}{1-\bar{\lambda}}\left((1-\delta)(1+\bar{I}^l) i_t^l - \rho^d (1+\bar{I}^d) i_{t-1}^d\right) + \frac{1}{(1-\bar{\lambda})^2}\hat{\lambda}_t\left((1-\delta)\bar{I}^l - \rho^d \bar{I}^d\right).$$

This leads to

$$\begin{aligned} i_t^l = & \underbrace{\frac{1+\bar{I}^b}{1+\bar{I}^l}\frac{1}{1-\frac{\bar{\lambda}}{1-\bar{\lambda}}(1-\delta)}}_{\equiv f^i} i_t^b - \underbrace{\frac{1+\bar{I}^d}{1+\bar{I}^l}\frac{1}{1-\frac{\bar{\lambda}}{1-\bar{\lambda}}(1-\delta)}\frac{\bar{\lambda}}{1-\bar{\lambda}}\rho^d}_{\equiv f^d} i_{t-1}^d \\ & + \underbrace{\frac{1}{1-\frac{\bar{\lambda}}{1-\bar{\lambda}}(1-\delta)}\frac{(1-\delta)\bar{I}^l - \rho^d \bar{I}^d}{1+\bar{I}^l}\frac{1}{(1-\bar{\lambda})^2}}_{\equiv f^\lambda}\hat{\lambda}_t,\end{aligned} \tag{A40}$$

where we define f^i, f^d, and f^λ as loadings on i_t^b, i_{t-1}^d, and $\hat{\lambda}_t$, respectively. Substituting in for the steady-state liquid bond rate from equation (A29), the loading on the log policy rate i_t^b can be further expressed as

$$\begin{aligned} f^i &= \frac{1+\bar{I}^b}{1+\bar{I}^l}\frac{1}{1-\frac{\bar{\lambda}}{1-\bar{\lambda}}(1-\delta)} \\ &= \frac{1+\left(1-\frac{\bar{\lambda}}{1-\bar{\lambda}}\left(1-\frac{\delta}{1-\rho^d}\right)\right)\bar{I}^l}{1+\bar{I}^l}\frac{1}{1-\frac{\bar{\lambda}}{1-\bar{\lambda}}(1-\delta)}\end{aligned} \tag{A41}$$

Under the assumptions that $0 < \bar{\lambda} < \frac{1}{2}$ and $\frac{\delta}{1-\rho^d} < 1$, we have that

$$0 < 1 - \frac{\bar{\lambda}}{1-\bar{\lambda}}\left(1 - \frac{\delta}{1-\rho^d}\right) < 1. \tag{A42}$$

Next, we note that for any $0 < a < 1$ and $\bar{I}^l > 0$, we have $1 + a\bar{I}^l > a + a\bar{I}^l$, implying $\frac{1+a\bar{I}^l}{1+\bar{I}^l} > a$. It follows that

$$\begin{aligned} f^i &> \left(1 - \frac{\bar{\lambda}}{1-\bar{\lambda}}\left(1 - \frac{\delta}{1-\rho^d}\right)\right)\frac{1}{1 - \frac{\bar{\lambda}}{1-\bar{\lambda}}(1-\delta)} \\ &= 1 + \frac{\bar{\lambda}}{1-\bar{\lambda}}\frac{\rho^d\delta}{1-\rho^d}\frac{1}{1 - \frac{\bar{\lambda}}{1-\bar{\lambda}}(1-\delta)} \geq 1. \end{aligned} \tag{A43}$$

This shows that $f^i > 1$ provided that $\bar{\lambda} < \frac{1}{2}$ and $\frac{\delta}{1-\rho^d} < 1$.

Similarly, we can also rewrite f^d,

$$f^d = \frac{1 + \frac{\delta}{1-\rho^d}\bar{I}^l}{1+\bar{I}^l}\frac{1}{1 - \frac{\bar{\lambda}}{1-\bar{\lambda}}(1-\delta)}\frac{\bar{\lambda}}{1-\bar{\lambda}}\rho^d. \tag{A44}$$

For simplicity, we rewrite equation (A40) as

$$i_t^l - i_t^b = (f^i - 1)i_t^b + f^\lambda\hat{\lambda}_t - f^d i_{t-1}^d, \tag{A45}$$

which is exactly equation (17) in the main text.

Next, we will express convenience yield $\ell_t \equiv i_t^l - i_t^b$ as a function of its lagged value ℓ_{t-1}, monetary policy innovations $i_t^b - \rho^i i_{t-1}^b$, and current monetary policy rate i_t^b. To achieve this, we linearize the sluggish deposit adjustment equation in (7),

$$i_t^d = \frac{\delta(1+\bar{I}^l)}{1+\bar{I}^d}i_t^l + \rho_d i_{t-1}^d,$$

which leads to

$$i_{t-1}^d = \frac{1}{\rho_d}\left(i_t^d - \frac{\delta(1+\bar{I}^l)}{1+\bar{I}^d}i_t^l\right). \tag{A46}$$

Plugging (A46) into Equation (A40), we get

$$\begin{aligned}\left(1+\bar{I}^l\right)i_t^l-\left(1+\bar{I}^b\right)i_t^b&=\frac{\bar{\lambda}}{1-\bar{\lambda}}\left[(1-\delta)\left(1+\bar{I}^l\right)i_t^l-\left(1+\bar{I}^d\right)\left(i_t^d-\frac{\delta(1+\bar{I}^l)}{1+\bar{I}^d}i_t^l\right)\right]+\left(1+\bar{I}^l\right)f^\lambda\hat{\lambda}_t\\&=\frac{\bar{\lambda}}{1-\bar{\lambda}}\left[\left(1+\bar{I}^l\right)i_t^l-\left(1+\bar{I}^d\right)i_t^d\right]+\left(1+\bar{I}^l\right)f^\lambda\hat{\lambda}_t.\end{aligned} \tag{A47}$$

Equation (A47) allows us to express i_t^d as a function of i_t^l and i_t^b,

$$\begin{aligned}i_t^d&=\frac{(2\bar{\lambda}-1)(1+\bar{I}^l)i_t^l+(1-\bar{\lambda})(1+\bar{I}^b)i_t^b+(1-\bar{\lambda})(1+\bar{I}^l)f^\lambda\hat{\lambda}_t}{\bar{\lambda}(1+\bar{I}^d)}\\&=-\frac{(1-2\bar{\lambda})(1+\bar{I}^l)}{\bar{\lambda}(1+\bar{I}^d)}i_t^l+\frac{(1-\bar{\lambda})(1+\bar{I}^b)}{\bar{\lambda}(1+\bar{I}^d)}i_t^b+\frac{(1-\bar{\lambda})(1+\bar{I}^l)f^\lambda}{\bar{\lambda}(1+\bar{I}^d)}\hat{\lambda}_t.\end{aligned} \tag{A48}$$

Next, we set time subscript to $t-1$ in (A48) and replace the i_{t-1}^d term in (A45) to get

$$\begin{aligned}i_t^l-i_t^b&=(f^i-1)i_t^b+f^\lambda\hat{\lambda}_t-f^d\left(-\frac{(1-2\bar{\lambda})(1+\bar{I}^l)}{\bar{\lambda}(1+\bar{I}^d)}i_{t-1}^l+\frac{(1-\bar{\lambda})(1+\bar{I}^b)}{\bar{\lambda}(1+\bar{I}^d)}i_{t-1}^b+\frac{(1-\bar{\lambda})(1+\bar{I}^l)f^\lambda}{\bar{\lambda}(1+\bar{I}^d)}\hat{\lambda}_{t-1}\right)\\&=f^d\frac{(1-2\bar{\lambda})(1+\bar{I}^l)}{\bar{\lambda}(1+\bar{I}^d)}\left(i_{t-1}^l-i_{t-1}^b\right)-f^d\left(\frac{(1-\bar{\lambda})(1+\bar{I}^b)}{\bar{\lambda}(1+\bar{I}^d)}-\frac{(1-2\bar{\lambda})(1+\bar{I}^l)}{\bar{\lambda}(1+\bar{I}^d)}\right)i_{t-1}^b\\&\quad+f^\lambda\hat{\lambda}_t-f^d\frac{(1-\bar{\lambda})(1+\bar{I}^l)f^\lambda}{\bar{\lambda}(1+\bar{I}^d)}\hat{\lambda}_{t-1}+(f^i-1)i_t^b.\end{aligned} \tag{A49}$$

We will further simplify (A49) along three ways. First, we define

$$v_{\ell,t}=f^\lambda\hat{\lambda}_t-f^d\frac{(1-\bar{\lambda})(1+\bar{I}^l)f^\lambda}{\bar{\lambda}(1+\bar{I}^d)}\hat{\lambda}_{t-1}. \tag{A50}$$

Second, we will show that the coefficient in front of i_{t-1}^b is equal to f^d in equation (A49). For that purpose, we note that the steady-state policy rate (A29) implies

$$(1-\bar{\lambda})\bar{I}^b=\left(1-\bar{\lambda}-\bar{\lambda}\left(1-\delta\frac{1}{1-\rho^d}\right)\right)\bar{I}^l,$$

$$(1-\bar{\lambda})\bar{I}^b - (1-2\bar{\lambda})\bar{I}^l = \left(\bar{\lambda}\delta\frac{1}{1-\rho^d}\right)\bar{I}^l.$$

Replacing the right-hand side with $\bar{I}^d$ in (A26), we get

$$(1-\bar{\lambda})\bar{I}^b - (1-2\bar{\lambda})\bar{I}^l = \bar{\lambda}\bar{I}^d$$

which implies

$$\begin{aligned} \frac{(1-\bar{\lambda})(1+\bar{I}^b)}{\bar{\lambda}(1+\bar{I}^d)} - \frac{(1-2\bar{\lambda})(1+\bar{I}^l)}{\bar{\lambda}(1+\bar{I}^d)} &= \frac{\bar{\lambda}+(1-\bar{\lambda})\bar{I}^b-(1-2\bar{\lambda})\bar{I}^l}{\bar{\lambda}(1+\bar{I}^d)} \\ &= \frac{\bar{\lambda}+\bar{\lambda}\bar{I}^d}{\bar{\lambda}(1+\bar{I}^d)} = 1 \end{aligned} \tag{A51}$$

Third, the persistence of convenience yield spread in the model is

$$\begin{aligned} \rho^\ell &\equiv f^d \frac{(1-2\bar{\lambda})(1+\bar{I}^l)}{\bar{\lambda}(1+\bar{I}^d)} \\ &= \frac{1}{1-\frac{\bar{\lambda}}{1-\bar{\lambda}}(1-\delta)} \frac{\bar{\lambda}}{1-\bar{\lambda}} \rho^d \frac{1+\bar{I}^d}{1+\bar{I}^l} \cdot \frac{(1-2\bar{\lambda})(1+\bar{I}^l)}{\bar{\lambda}(1+\bar{I}^d)} \\ &= \frac{1-2\bar{\lambda}}{1-2\bar{\lambda}+\delta\bar{\lambda}}\rho^d \end{aligned} \tag{A52}$$

With simplifications in (A50), (A51), (A52), and using the convenience yield notation ℓ_t defined in (17), we rewrite equation (A49) as

$$\ell_t = \rho^\ell \ell_{t-1} + (f^i - 1 - \frac{f^d}{\rho^i})i_t^b + \frac{f^d}{\rho^i}(i_t^b - \rho^i i_{t-1}^b) + v_{\ell,t} \tag{A53}$$

Plugging in the monetary policy innovation implied by (8) into (A53), we get

$$\begin{aligned} \ell_t &= \rho^\ell \ell_{t-1} + \frac{f^d}{\rho^i}((1-\rho^i)(\gamma_x x_t + \gamma_\pi \pi_t) + v_{i,t}) + (f^i - 1 - \frac{f^d}{\rho^i})i_t^b + v_{\ell,t} \\ &= \rho^\ell \ell_{t-1} + \underbrace{f^d \frac{1-\rho^i}{\rho^i}\gamma_x}_{\equiv g^x} x_t + \underbrace{f^d \frac{1-\rho^i}{\rho^i}\gamma_\pi}_{\equiv g^\pi} \pi_t + \underbrace{(f^i - 1 - \frac{f^d}{\rho^i})}_{\equiv g^i} i_t^b + \underbrace{\frac{f^d}{\rho^i}}_{\equiv g^v} v_{i,t} + v_{\ell,t} \end{aligned} \tag{A54}$$

With definitions of g^x, g^π, g^i, and g^v, we finally get the following linearized convenience yield dynamics ready for model solution,

$$\ell_t = \rho^\ell \ell_{t-1} + g^x x_t + g^\pi \pi_t + g^i i_t^b + g^v v_{i,t} + v_{\ell,t} \tag{A55}$$

B.4 Macroeconomic equilibrium

We use a scaled state vector to solve for macroeconomic dynamics. We define $\xi_t = -\psi \ell_t$ and $v_{\xi,t} = -\psi v_{\ell,t}$, and denote the volatility of $v_{\xi,t}$ as σ_ξ. Then the updated macro block state vector is $\hat{Z}_t = \left[x_t, \pi_t, i_t^b, \xi_t\right]$ and the shock vector is $\hat{v}_t = [v_{\xi,t}, v_{\pi,t}, v_{i,t}, v_{x,t}]$. The dynamics are described by

$$\begin{aligned}
x_t &= (1-\rho^x) E_t x_{t+1} + \rho^x x_{t-1} - \psi i_t^b + \psi E_t \pi_{t+1} + \xi_t + v_{x,t}, && \text{(A56)}\\
\pi_t &= (1-\rho^\pi) E_t \pi_{t+1} + \rho^\pi \pi_{t-1} + \kappa x_t + v_{\pi,t}, && \text{(A57)}\\
i_t^b &= (1-\rho^i)\left(\gamma^x x_t + \gamma^\pi \pi_t\right) + \rho^i i_{t-1}^b + v_{i,t}, && \text{(A58)}\\
\xi_t &= \rho^\ell \xi_{t-1} - \psi g^x x_t - \psi g^\pi \pi_t - \psi g^i i_t^b - \psi g^v v_{i,t} + v_{\xi,t}. && \text{(A59)}
\end{aligned}$$

In matrix form, the model can be written as

$$0 = F E_t[\hat{Z}_{t+1}] + G \hat{Z}_t + H \hat{Z}_{t-1} + M v_t, \tag{A60}$$

where the matrices are given by

$$F = \begin{bmatrix} 1-\rho^x & \psi & 0 & 0 \\ 0 & 1-\rho^\pi & 0 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 \end{bmatrix}, \tag{A61}$$

$$G = \begin{bmatrix} -1 & 0 & -\psi & 1 \\ \kappa & -1 & 0 & 0 \\ (1-\rho^i)\gamma^x & (1-\rho^i)\gamma^\pi & -1 & 0 \\ -\psi g^x & -\psi g^\pi & -\psi g^i & -1 \end{bmatrix}, \tag{A62}$$

$$H = \begin{bmatrix} \rho^x & 0 & 0 & 0 \\ 0 & \rho^\pi & 0 & 0 \\ 0 & 0 & \rho^i & 0 \\ 0 & 0 & 0 & \rho^\ell \end{bmatrix}, \tag{A63}$$

$$M = \begin{bmatrix} 0 & 0 & 0 & 1 \\ 0 & 1 & 0 & 0 \\ 0 & 0 & 1 & 0 \\ 1 & 0 & -\psi g^v & 0 \end{bmatrix}. \tag{A64}$$

We use Uhlig (1999)'s formulation of Blanchard and Kahn (1980) to solve for an equilibrium of the form

$$\hat{Z}_{t+1} = B\hat{Z}_t + \Sigma\hat{v}_{t+1}.$$

We let Σ_v to denote the variance covariance matrix of $\hat{v}_{t+1}$.

B.5 Solving for asset prices

Next, the full state vector when we solve for long-term asset prices include not only $\hat{Z}_t$, but also the surplus consumption ratio relative to steady-state, $\hat{s}_t$ (see Appendix Section B.2.1), which affects risk premium. A key numerical challenge is high-dimensional integration with different shock sizes σ_j, $j \in \{\xi, \pi, i, \theta\}$, among which the θ dimension is highly nonlinear for asset prices due to the effect of the consumption-surplus ratio. For numerical computations, we will rotate the state-vector $\hat{Z}_t$ into $\tilde{Z}_t$, defined as $\tilde{Z}_t = A\hat{Z}_t$ for some invertible matrix A. Thus, the dynamics of $\tilde{Z}_t$ are given by:

$$\tilde{Z}_t = A\hat{Z}_t, \tag{A65}$$

$$\tilde{Z}_{t+1} = \underbrace{ABA^{-1}}_{\tilde{B}}\tilde{Z}_t + \underbrace{A\Sigma v_{t+1}}_{\epsilon_{t+1}}. \tag{A66}$$

We hence want a matrix, A, such that

$$\begin{aligned} Var\left(\epsilon_{t+1}\right) &= A\Sigma\Sigma_v\Sigma' A', \quad \text{(A67)} \\ &= \begin{bmatrix} 1 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 1 \end{bmatrix}. \quad \text{(A68)} \end{aligned}$$

and

$$A_1 \propto e_1. \quad \text{(A69)}$$

We can therefore find the three rows of A using the following steps:

1. Set $A_1 = \frac{e_1}{\sqrt{e_1 \Sigma\Sigma_v\Sigma' e_1'}}$.
2. We use the MATLAB function ***null*** to compute the null space of $A_1\Sigma\Sigma_v\Sigma'$. Let n_2 denote the first vector in $null\left(A_1\Sigma\Sigma_v\Sigma'\right)$. We then define the second row of A as the normalized version of n_2:

$$A_2 = \frac{n_2}{\sqrt{n_2\Sigma\Sigma_v\Sigma' n_2'}}. \quad \text{(A70)}$$

3. Let n_3 denote the first vector in $null\left(A_1\Sigma\Sigma_v\Sigma', A_2\Sigma\Sigma_v\Sigma'\right)$. We then define the third row of A as the normalized version of n_3:

$$A_3 = \frac{n_3}{\sqrt{n_3\Sigma\Sigma_v\Sigma' n_3'}}. \quad \text{(A71)}$$

4. Let n_4 denote the first vector in $null\left(A_1\Sigma\Sigma_v\Sigma', A_2\Sigma\Sigma_v\Sigma', A_3\Sigma\Sigma_v\Sigma'\right)$. We then define the fourth row of A as the normalized version of n_4:

$$A_4 = \frac{n_4}{\sqrt{n_4\Sigma\Sigma_v\Sigma' n_4'}}. \quad \text{(A72)}$$

It is then straightforward to verify that equation (A68) holds.

Expressing surplus consumption

Before deriving the recursions for the numerical asset pricing computations, we derive some convenient expressions. We use e_i to denote a row vector with 1 in position i and zeros elsewhere. The matrix

$$\Sigma_M = e_1\Sigma \tag{A73}$$

denotes the loading of consumption innovations onto the vector of shocks v_t, where e_1 is a basis vector with a one in the first position and zeros everywhere else. The volatility of consumption surprises equals:

$$\sigma_c^2 = \Sigma_M \Sigma_v \Sigma_M'. \tag{A74}$$

To simplify notation, we define $\hat{s}_t$ as the log deviation of surplus consumption from its steady state. The dynamics of $\hat{s}_t$ are given by (6) in the main paper, plus the specification of the sensitivity function from Campbell and Cochrane (1999):

$$\tilde{\lambda}(\hat{s}_t) = \lambda_0\sqrt{1-2\hat{s}_t} - 1, \hat{s}_t \leq s_{max} - \bar{s}, \tag{A75}$$

$$\tilde{\lambda}(\hat{s}_t) = 0, \hat{s}_t \geq s_{max} - \bar{s}. \tag{A76}$$

The steady-state surplus consumption sensitivity equals:

$$\lambda_0 = \frac{1}{\bar{S}}. \tag{A77}$$

Using the SDF equation (A13), definition of $m_{t+1} = \log(M_{t+1})$, and the sensitivity function (A14), we get:

$$\begin{aligned} \mathbb{E}_t\left[m_{t+1}\right] &= \log\beta - \gamma E_t \Delta\hat{s}_{t+1} - \gamma E_t \Delta c_{t+1} \\ &= -\bar{r}^l - \hat{r}_t^l - \frac{\gamma}{2}(1-\theta_0)(1-2\hat{s}_t). \end{aligned} \tag{A78}$$

This can be expressed in terms of the state variables:

$$\begin{aligned}\mathbb{E}_t\left[m_{t+1}\right] &= -\bar{r}^l - (e_3 - e_2 B)Y_t + \psi^{-1}\xi_t - \frac{\gamma}{2}(1-\theta_0)(1-2\hat{s}_t) \\ &= -\bar{r}^l - (e_3 - e_2 B - \psi^{-1}e_4)Y_t - \frac{\gamma}{2}(1-\theta_0)(1-2\hat{s}_t)\end{aligned} \tag{A79}$$

The updating rule for the log surplus consumption ratio can then be written in terms of the state variables.

Using (A79) we can instead write (up to a constant)

$$E_t\Delta\hat{s}_{t+1} = -E_t\Delta c_{t+1} - \frac{1}{\gamma}E_t m_{t+1}, \tag{A80}$$

$$= -e_1 B\hat{Z}_t + \phi e_1\hat{Z}_t + \frac{1}{\gamma}(e_3 - e_2 B - \psi^{-1}e_4)\hat{Z}_t - (1-\theta_0)\hat{s}_t \tag{A81}$$

Adding $\hat{s}_t$ to both sides allows us to re-express this in terms of state variables

$$\hat{s}_{t+1} = \theta_0\hat{s}_t + \underbrace{\left[-e_1\left(B-\phi I\right)\tilde{Z}_t + \frac{1}{\gamma}(e_3 - e_2 B - \psi^{-1}e_4)\right]A^{-1}}_{A_s}\tilde{Z}_t + \tilde{\lambda}(\hat{s}_t)\varepsilon_{c,t+1}. \tag{A82}$$

B.5.1 Recursion for zero-coupon liquid bond prices

We use $P^{b,\$}_{n,t}$ and $P^{b}_{n,t}$ to denote the prices of nominal and real n-period zero-coupon liquid bonds. The strategy is to develop analytic expressions for one- and two-period liquid bond prices. We then guess and verify recursively that the prices of nominal and real zero-coupon liquid bonds with maturity $n \geq 2$ can be written in the following form:

$$P^{b,\$}_{n,t} = B^{b,\$}_n(\tilde{Z}_t, \hat{s}_t), \tag{A83}$$

$$P^{b}_{n,t} = B^{b}_n(\tilde{Z}_t, \hat{s}_t). \tag{A84}$$

As discussed in the main paper, we assume that the short-term nominal interest rate contains no risk premium, so the one-period log nominal interest rate equals $i_t = r_t + E_t\pi_{t+1}$. Taking account of the constants, one-period liquid bond prices equal:

$$P_{1,t}^{b,\$} = \exp\left(-\hat{Z}_{3,t} - \bar{i}^b\right), \tag{A85}$$

$$P_{1,t}^{b} = \exp\left(-\hat{Z}_{3,t} + E_t\hat{Z}_{2,t+1} - \bar{r}^b\right), \tag{A86}$$

We next solve for longer-term liquid bond prices including risk premia. Substituting (A85) into the bond-pricing recursion in equation (14) gives:

$$\begin{aligned} P_{2,t}^{b,\$} &= \exp\left(\bar{i}^l - \bar{i}^b - \psi^{-1}\xi_t\right)\mathbb{E}_t\left[M_{t+1}^{\$}P_{1,t+1}^{b,\$}\right] & \text{(A87)}\\ &= \exp\left(-\psi^{-1}\xi_t + \bar{i}^l - \bar{i}^b - \bar{\pi}\right)\mathbb{E}_t\left[M_{t+1}P_{1,t+1}^{\$}\exp(-\hat{Z}_{2,t+1})\right], & \text{(A88)}\end{aligned}$$

We can now verify that the two-period nominal liquid bond price takes the form (A83):

$$\begin{aligned} B_2^{b,\$}(\tilde{Z}_t,\hat{s}_t) &= \exp\left(E_t\left(m_{t+1} - \psi^{-1}\xi_t - \hat{Z}_{3,t+1} - \hat{Z}_{2,t+1}\right) + \bar{i}^l - 2\bar{i}^b - \bar{\pi}\right)\\ &\quad\times\mathbb{E}_t\left[\exp\left(\left(-\gamma\left(\tilde{\lambda}(\hat{s}_t)+1\right)\Sigma_M - \underbrace{[(e_2+e_3)\Sigma]}_{v_{\$,b}}\right)v_{t+1}\right)\right].\end{aligned} \tag{A89}$$

Here, we define the vector $v_{\$,b}$ to simplify notation. Taking logs, substituting out for $E_t m_{t+1}$, and using the definition for the sensitivity function $\tilde{\lambda}(\hat{s}_t)$, we obtain

$$\begin{aligned} b_2^{b,\$}\left(\tilde{Z}_t,\hat{s}_t\right) &= -e_3\left[I+B\right]A^{-1}\tilde{Z}_t + \frac{1}{2}v_{\$,b}\Sigma_v v_{\$,b}'\\ &\quad+\gamma(\tilde{\lambda}(\hat{s}_t)+1)\Sigma_M\Sigma_v v_{\$,b}' - 2\bar{i}^b,\end{aligned} \tag{A90}$$

The closed-form solution for the two-period real liquid bond price becomes

$$\begin{aligned} P_{2,t}^{b} &= \exp\left(E_t\left(m_{t+1} - \psi^{-1}\xi_t - \hat{Z}_{3,t+1} + \hat{Z}_{2,t+2}\right) + \bar{i}^l - \bar{i}^b - \bar{r}^b\right)\\ &\quad\times\mathbb{E}_t\left[\exp\left((-\gamma(\tilde{\lambda}(\hat{s}_t)+1)\Sigma_M - \underbrace{(e_3 - e_2B)\Sigma}_{v_b})v_{t+1}\right)\right]\end{aligned} \tag{A91}$$

We define the vector v_b to simplify notation. Taking logs, substituting out for $E_t m_{t+1}$ using (A79), and using the definition for $\tilde{\lambda}(\hat{s}_t)$ gives:

$$\begin{aligned} b_2(\tilde{Z}_t, \hat{s}_t) &= -\left(e_3 - e_2 B\right)\left[I + B\right] A^{-1} \tilde{Z}_t \\ &\quad + \frac{1}{2} v_b \Sigma_v v_b' + \gamma\left(\lambda\left(\hat{s}_t\right) + 1\right) \Sigma_M \Sigma_v v_b' - 2\bar{r}^b. \end{aligned} \tag{A92}$$

For $n \geq 3$, we repeatedly substitute out for $E_t m_{t+1}$ to obtain the following recursion for nominal and real liquid bond prices, respectively:

$$\begin{aligned} B_n^{b,\$}(\tilde{Z}_t, \hat{s}_t) &= \exp\left(-\psi^{-1}\xi_t + (\bar{i}^l - \bar{i}^b - \bar{\pi})\right) \\ &\quad \times \mathbb{E}_t\left[\exp\left(m_{t+1} - \hat{Z}_{2,t+1} + b_{n-1}^{b,\$}\left(\tilde{Z}_{t+1}, \hat{s}_{t+1}, \hat{x}_t\right)\right)\right] \\ &= \mathbb{E}_t\left[\exp\left(-\bar{i}^b - e_3 A^{-1}\tilde{Z}_t - \frac{\gamma}{2}(1-\theta_0)(1-2\hat{s}_t)\right.\right. \\ &\quad \left.\left. -\gamma(1+\tilde{\lambda}(\hat{s}_t))\sigma_c \epsilon_{1,t+1} - e_2 A^{-1}\epsilon_{t+1} + b_{n-1}^{b,\$}\left(\tilde{Z}_{t+1}, \hat{s}_{t+1}\right)\right)\right]. \end{aligned} \tag{A93}$$

The value function iteration for real liquid bond prices then becomes

$$\begin{aligned} B_n^b(\tilde{Z}_t, \hat{s}_t) &= \exp\left(-\psi^{-1}\xi_t + \bar{i}^l - \bar{i}^b\right) \mathbb{E}_t\left[\exp\left(m_{t+1} + b_{n-1}^b\left(\tilde{Z}_{t+1}, \hat{s}_{t+1}\right)\right)\right] \\ &= \mathbb{E}_t\left[\exp\left(-\bar{r}^b - \left(e_3 - e_2 B\right) A^{-1}\tilde{Z}_t - \frac{\gamma}{2}\left(1-\theta_0\right)\left(1-2\hat{s}_t\right)\right.\right. \\ &\quad \left.\left. -\gamma\left(1+\lambda\left(\hat{s}_t\right)\right)\sigma_c \epsilon_{1,t+1} + b_{n-1}^b\left(\tilde{Z}_{t+1}, \hat{s}_{t+1}\right)\right)\right]. \end{aligned} \tag{A94}$$

Since (A93) and (A94) have the four-dimensional vector $\tilde{Z}_{t+1}$ on the right-hand-side, evaluating these expectations requires taking a four-dimensional expectation. Because $\hat{s}_{t+1}$ can be expressed as in equation (A82) the lagged output gap x_{t-1} is not required as a state variable, though we carry it around in the code for legacy reasons.

A simple debugging exercise uses that (A94) and (A93) also hold for $n = 2$ setting $b_1^b = -\hat{Z}_{3,t} + E_t\hat{Z}_{2,t+1}$ and $b_1^{b,\$} = -\hat{Z}_{3,t}$ everywhere. This equivalence allows us to check that the numerical integration works for bonds.

B.5.2 Consumption claim recursions

We now derive the recursion for zero-coupon consumption claims in terms of state variables $\tilde{Z}_t$, $\hat{s}_t$ and x_{t-1}. Let P^c_{nt}/C_t denote the price-dividend ratio of a zero-coupon claim on consumption at time $t+n$. The outline of our strategy here is that we first derive an analytic expression for the price-dividend ratio for P^c_{1t}/C_t. For $n \geq 1$ we guess and verify recursively that there exists a function $F_n(\tilde{Z}_t, \hat{s}_t, x_{t-1})$, such that

$$\frac{P^c_{nt}}{C_t} = F_n\left(\tilde{Z}_t, \hat{s}_t\right). \tag{A95}$$

The ex-dividend price-consumption ratio for a claim to all future consumption is then given by

$$\frac{P_t}{C_t} = F\left(\tilde{Z}_t, \hat{s}_t\right), \tag{A96}$$

where we define

$$F\left(\tilde{Z}_t, \hat{s}_t\right) = \sum_{n=1}^{\infty} F_n\left(\tilde{Z}_t, \hat{s}_t\right). \tag{A97}$$

We now derive the recursion of zero-coupon consumption claims in terms of state variables $\tilde{Z}_t$ and $\hat{s}_t$. The one-period zero coupon price-consumption ratio solves:

$$\frac{P^c_{1,t}}{C_t} = E_t\left[\frac{M_{t+1}C_{t+1}}{C_t}\right] \tag{A98}$$

We simplify

$$\begin{aligned}\frac{M_{t+1}C_{t+1}}{C_t} &= \exp\left(E_t m_{t+1} + E_t \Delta c_{t+1}\right.\\ &\quad \left.-\gamma(\hat{s}_{t+1} - E_t\hat{s}_{t+1}) - (\gamma-1)(c_{t+1} - E_t c_{t+1})\right)\end{aligned} \tag{A99}$$

Using the notation $f_n = \log(F_n)$, this gives the log one-period price-consumption ratio as:

$$\begin{aligned}f_1\left(\tilde{Z}_t, \hat{s}_t\right) &= \log\beta - (\gamma-1)g + \left[e_1\left[B - \phi I\right] - \left(e_3 - e_2 B - \psi^{-1}e_4\right)\right]A^{-1}\tilde{Z}_t\\ &\quad +\gamma(1-\theta_0)\hat{s}_t + \frac{1}{2}\left(\gamma\tilde{\lambda}(\hat{s}_t) + (\gamma-1)\right)^2\sigma_c^2.\end{aligned} \tag{A100}$$

Next, we solve for f_n, $n \geq 2$ iteratively. Note that:

$$\frac{P^c_{nt}}{C_t} = \mathbb{E}_t\left[\frac{M_{t+1}C_{t+1}}{C_t}\frac{P^c_{n-1,t+1}}{C_{t+1}}\right] = \mathbb{E}_t\left[\frac{M_{t+1}C_{t+1}}{C_t}F_{n-1}\left(\tilde{Z}_{t+1}, \hat{s}_{t+1}, x_t\right)\right]. \tag{A101}$$

This gives the following expression for f_n:

$$\begin{aligned} f_n(\tilde{Z}_t, \hat{s}_t) \quad = \quad & \log\Big[\mathbb{E}_t\Big[\exp\Big(\log\beta - (\gamma-1)g + \big[e_1\left[B - \phi I\right] - \big(e_3 - e_2 B - \psi^{-1}e_4\big)\big]A^{-1}\tilde{Z}_t \\ & +\gamma(1-\theta_0)\hat{s}_t - (\gamma(1+\tilde{\lambda}(\hat{s}_t)) - 1)\sigma_c\epsilon_{1,t+1} \\ & +f_{n-1}(\tilde{Z}_{t+1}, \hat{s}_{t+1})\Big)\Big]\Big]. \end{aligned} \tag{A102}$$

Here, $\epsilon_{1,t+1}$ denotes the first dimension of the shock ϵ_{t+1}. This expression clearly nests (A100) with $f_0 = 0$ everywhere. Combined with the analytical expression for f_1, this observation can be used to check that the numerical integration works as it should.

B.5.3 Risk-neutral zero-coupon liquid bond prices

We use the superscript ^{rn} for risk-neutral, superscript ^{cf} for cash flow, and ^{rp} for risk premium. Risk-neutral valuations are expected cash flows discounted with the risk-neutral discount factor, given by:

$$M^{rn}_{t+1} \quad = \quad \exp(-r^l_t). \tag{A103}$$

Note that since we are not interested in risk-neutral bond and stock prices, but only a decomposition of returns, multiplying M^{rn}_{t+1} by a constant discount rate does not matter. For any zero-coupon claim it would shift risk-neutral returns merely by a constant and therefore leave our decomposition into risk-neutral and risk-premium components unaffected. For a claim to all future consumption or stock returns, a constant discount rate could theoretically shift the weights between nearer-term consumption claims and longer-term consumption claims, and therefore change risk-neutral returns. However, since consumption growth is stationary we have found that this makes very little different to risk-neutral stock returns in any of our numerical applications.

We derive the two-period risk-neutral nominal liquid bond price analytically:

$$P_{2,t}^{b,\$,rn} = \exp\left(-\psi^{-1}\xi_t + \bar{i}^l - \bar{i}^b - \bar{\pi}\right) \mathbb{E}_t \left[M_{t+1}^{rn} P_{1,t+1}^{b,\$,rn} \exp(-\hat{Z}_{2,t+1})\right] \tag{A104}$$

$$= \exp\left(-\psi^{-1}\xi_t + \bar{i}^l - \bar{i}^b - \bar{\pi}\right) \mathbb{E}_t \left[M_{t+1}^{rn} \exp(-\hat{Z}_{3,t+1} - \hat{Z}_{2,t+1} - \bar{i}^b)\right]. \tag{A105}$$

We can hence verify that the two-period risk-neutral nominal liquid bond price takes the form (A83)

$$b_2^{b,\$,rn}(\tilde{Z}_t, \hat{s}_t) = -e_3 \left[I + B\right] A^{-1} \tilde{Z}_t + \frac{1}{2} v_{\$,b} \Sigma_v v_{\$,b}\prime - 2\bar{i}^b \tag{A106}$$

Here, the vector $v_{\$,b}$ is identical to the case with risk aversion. Comparing expressions (A106) and (A90) shows that they agree when $\gamma = 0$. We similarly solve for 2-period real liquid bond prices in closed form:

$$P_{2,t}^{b,rn} = \exp\left(-\hat{Z}_{3,t} + \mathbb{E}_t \hat{Z}_{2,t+1} - \bar{r}^b\right) \times \exp\left(\mathbb{E}_t\left(-\hat{Z}_{3,t+1} + \mathbb{E}_{t+1}\hat{Z}_{2,t+2} - \bar{r}^b\right)\right)$$
$$\times \mathbb{E}_t \left[\exp\left(-\underbrace{(e_3 - e_2 B)\Sigma}_{v_b} v_{t+1}\right)\right]. \tag{A107}$$

The vector v_b is again identical to the case with risk aversion. Taking logs gives:

$$b_2^{b,rn}(\tilde{Z}_t, \hat{s}_t) = -(e_3 - e_2 B) \left[I + B\right] A^{-1} \tilde{Z}_t + \frac{1}{2} v_b \Sigma_v v_b' - 2\bar{r}^b. \tag{A108}$$

Risk-neutral real liquid bond prices (A108) and liquid bond prices with risk aversion (A92) are identical when the utility curvature parameter γ equals zero.

For $n \geq 3$ the n-period risk neutral nominal and real liquid bond prices satisfy the following recursions, respectively:

$$B_n^{b,\$,rn}(\tilde{Z}_t, \hat{s}_t) = \mathbb{E}_t \left[\exp\left(-\bar{i}^b - e_3 A^{-1} \tilde{Z}_t - e_2 A^{-1} \epsilon_{t+1} + b_{n-1}^{b,\$,rn}(\tilde{Z}_{t+1}, \hat{s}_{t+1})\right)\right], \tag{A109}$$

$$B_n^{b,rn}(\tilde{Z}_t, \hat{s}_t) = \mathbb{E}_t\left[\exp\left(-\bar{r}^b - (e_3 - e_2 B)A^{-1}\tilde{Z}_t + b_{n-1}^{b,rn}(\tilde{Z}_{t+1}, \hat{s}_{t+1})\right)\right] \tag{A110}$$

B.5.4 Risk-neutral zero-coupon consumption claims

Next, we derive recursive solutions for the risk-neutral prices of zero-coupon consumption claims. Let $P_{nt}^{c,rn}/C_t$ denote the risk-neutral price-dividend ratio of a zero-coupon claim on consumption at time $t+n$. The risk-neutral price-consumption ratio of a claim to the entire stream of future consumption equals:

$$\frac{P_t^{c,rn}}{C_t} = \sum_{n=1}^{\infty} \frac{P_{nt}^{c,rn}}{C_t}. \tag{A111}$$

We start by deriving the analytic expression for F_1^{rn}. The one-period risk-neutral zero-coupon price-consumption ratio solves

$$\frac{P_{1,t}^{c,rn}}{C_t} = \mathbb{E}_t\left[M_{t+1}^{rn}\frac{C_{t+1}}{C_t}\right] \tag{A112}$$

$$= \exp(-\bar{r}^l - r_t^b + \psi^{-1}\xi_t)\mathbb{E}_t\left[\frac{C_{t+1}}{C_t}\right] \tag{A113}$$

Substituting out for expected consumption growth, this gives the following analytic expression for f_1^{rn}:

$$f_1^{rn}(z_t, \hat{s}_t, \hat{x}_{t-1}) = g - \bar{r}^l + \left[e_1\left[B - \phi I\right] - (e_3 - e_2 B - \psi^{-1}e_4)\right]A^{-1}\tilde{Z}_t + \frac{1}{2}\sigma_c^2. \tag{A114}$$

Next, we solve for f_n, $n \geq 2$ iteratively:

$$\frac{P_{nt}^{c,rn}}{C_t} = \exp\left(-(e_3 - e_2 B - \psi^{-1}e_4)A^{-1}\tilde{Z}_t - \bar{r}^l\right)\mathbb{E}_t\left[\frac{C_{t+1}}{C_t}F_{n-1}^{rn}\left(\tilde{Z}_{t+1}, \hat{s}_{t+1}\right)\right] \tag{A115}$$

This gives the following expression for f_n^{rn}:

$$f_n^{rn}(z_t, \hat{s}_t, \hat{x}_{t-1}) = \log \left[\mathbb{E}_t \left[\exp \left(g - \bar{r}^l + \left[e_1 \left[B - \phi I \right] - (e_3 - e_2 B - \psi^{-1} e_4) \right] A^{-1} \tilde{Z}_t + \sigma_c \epsilon_{1,t+1} + f_{n-1}^{rn} \left(\tilde{Z}_{t+1}, \hat{s}_{t+1} \right) \right) \right] \right]. \quad \text{(A116)}$$

B.6 Invariant model parameter values

Our model parameters are set in two steps. First, we use typical values from the literature for invariant parameters. Second, in each period, targeting data moments, we estimate shock volatility parameters. The first step is shown in Table A6.

As discussed in the main text, parameters for the New Keynesian block of the model are set to values from the literature. Preference parameters are set as in Pflueger and Rinaldi (2022), implying an Euler equation with forward- and backward-looking components and a plausible output gap response to identified monetary policy shocks. The Phillips curve slope is set as in Rotemberg and Woodford (1997) and the backward-looking and forward-looking coefficients in the Phillips curve are derived from backward-looking price indexation as in Smets and Wouters (2007). The steady-state discount rate is implied by a real risk-free rate of $\bar{r}^l = 0.93\%$ in annualized units via equation (A24), following Campbell and Cochrane (1999). Combined with a steady-state inflation of $\bar{\Pi} = 2\%$ in annual units this implies a steady-state illiquid nominal loan rate of 2.95% annualized in our model.

The deposit rate pass-through from policy rates is set following the liquidity literature and empirical evidence from CALL report data. The long-term deposit-rate adjustment to policy rate change $\delta/(1-\rho^d)$ is set to 1/3, within the range of 1/3 to 1/2 suggested by Nagel (2016). We use CALL report data (quarterly frequency from 1987 Q1 to 2020 Q1) to estimate $\rho^d = 0.92$ from a time-series regression of the form (7). The resulting value for the short-run pass through is $\delta = 0.027$, which is also consistent with the time series regression in CALL report data reported in Appendix A.7. The steady-state weight on Treasury bonds in the liquidity aggregate, $\bar{\lambda}$, is set to match the steady-state convenience yield in the data. The firm's equity-to-asset ratio is set to $\delta^c = 0.5$ (50%) as in Campbell et al. (2020) to generate reasonable equity market volatility.

Table A6. Model calibration. This table contains the calibration parameters for the New Keynesian model with convenience yields. Parameters are reported in units corresponding to inflation and interest rates in annualized percent, and output gap in percent, that is we report $\frac{\psi}{4}$, 4κ and $4\gamma^x$ compared to natural quarterly units. The discount rate and real risk-free rate are reported in annualized units.

Panel A: Preferences, Technology, and Monetary Policy			
Euler equation			Target
Interest rate slope	ψ	0.07	Pflueger and Rinaldi (2022)
Backward-looking component	ρ^x	0.45	Pflueger and Rinaldi (2022)
PC Parameters			
Slope	κ	0.02	Rotemberg and Woodford (1997)
Backward-looking PC	ρ^π	0.51	Fuhrer (1997a)
Monetary Policy			
MP inertia	ρ^i	0.80	Clarida et al. (2000)
Output gap weight	γ^x	0.50	Taylor (1993)
Inflation weight	γ^π	1.50	Taylor (1993)
Equities			
Equity share	δ^c	0.50	Campbell et al. (2020)
Panel B: Interest Rates and Liquidity			
Real risk-free rate	$\bar{r}^l$	0.94%	Campbell and Cochrane (1999)
Discount factor	β	0.90	From risk-free rate and eqn. (A24)
Steady-state level inflation	$\bar{\Pi}$	2%	Fed inflation target
Deposit rate pass-through	δ	0.027	Deposit rate sensitivity to the risk-free rate
Deposit rate sluggishness	ρ^d	0.92	Deposit rate sluggishness in the data
Bond liquidity weight	$\bar{\lambda}$	0.14	Level of T-bill convenience spread

B.7 Details on model estimation and standard errors

After setting the basic parameter values, we will use data moments to identify shock volatilities in each period. We estimate the four types of volatilities, including liquidity demand shock volatility σ_λ, cost-push supply shock volatility σ_π, monetary policy shock volatility σ_i, and non-liquidity demand shock volatility σ_x. The volatilities of these four different shocks drive the volatilities of various variables and also their correlations. Intuitively, convenience yield volatility is directly affected by σ_λ, while inflation volatility is directly affected by σ_π. The volatility of output gap is driven by both liquidity demand shock volatility σ_λ and non-liquidity demand shock volatility σ_x. Moreover, these volatilities also affect correlations of key variables, including output gap, convenience yield, and inflation. For example, liquidity demand shocks drive a positive inflation-output gap correlation but a negative inflation-convenience correlation. On the other hand, supply shocks drive a negative inflation-output gap correlation but a positive inflation-convenience correlation.

Estimation Algorithm. For each period, we estimate the four volatility parameters by targeting six data moments using quarterly data: volatilities of convenience yield (T-bill spread), quarterly inflation, and output gap, the correlation between quarterly inflation and output gap, and the bond-stock beta. Denote these data moments as m^{data}. For each parameter vector σ, we simulate the model at quarterly frequency for 50000 quarters (results are similar if we use a longer simulation horizon) and compute the model-implied moments $\hat{m}(\sigma)$. We estimate parameters to minimize the weighted squared deviations from the data moments,

$$\min_{\sigma}(m^{\text{data}} - \hat{m}(\sigma))' W^{-1} (m^{\text{data}} - \hat{m}(\sigma)) \tag{A117}$$

To improve the efficiency of our estimations, we will use theory-implied variance of data moments as diagonal elements of the weighting matrix W and for simplicity assume zero correlations among those moments.

In Table 4 of the main text, we illustrate the seven target moment values across three periods. We report the standard deviation of each moment value in bracket. The standard deviation of a volatility moment is $\hat{\sigma}/\sqrt{2 \times (N-1)}$, and the standard deviation of a correlation moment is $(1-\hat{\rho}^2)/\sqrt{N-3}$. The standard deviation of the regression coefficient is directly obtained from regression analysis.

Since the asset pricing solution involves nonlinearities, a direct estimation over four parameters on six data moments takes too long to finish. Therefore, we designed a two-stage estimation procedure to speed up the estimation. First, we only solve for the macro block, including the one-period convenience yield, and target all moments except the bond-stock beta in Table 4. This step involves a subset of m^{data}, $\hat{m}(\sigma)$, and weight matrix W. This gives us a reasonable guess for a good starting point of the full optimization. Denote the estimated parameter set as σ_{step1}. Next, once we obtain the solution σ_{step1}, we create a grid around it and evaluate those grid points to find the best solution that optimize the objective function in (A117) plus penalties on wrong signs, including bond-stock beta, large Sharpe ratio (above 1), regression coefficient of convenience yield on inflation with and without controlling the policy rate. To maximize the effectiveness of the grid without incurring unbearable computational burden, we use the Smolayak grid method that creates a sparse grid not subject to the curse of dimensionality. The resulting optimal parameter values from stage 2 are reported in Table 5 of the main text.

Calculating Standard Errors. After estimating the parameters, we evaluate the standard errors (or error bands) for parameter estimates, using the asymptotic variance-covariance matrix formula derived from Generalized Method of Moments (GMM) theory. We first calculate the covariance matrix of moment residuals $g(\hat{\sigma}) = m - \hat{m}(\hat{\sigma})$, by simulating the model for T^{data} periods (each simulation generates one residual vector) for 1000 times with distinct random seeds. We set T^{data} equal to the number of periods in the data to capture the finite-sample variability of the moments, ensuring that the uncertainty reflects the same sample size as the empirical data. We need to adjust the covariance matrix by sample length and denote the covariance matrix of moment residuals divided by T^{data} as S. Second, we numerically compute the Jacobian matrix that represents how sensitive the moments are to parameter changes,

$$D = \nabla g(\sigma)|_{\sigma=\hat{\sigma}}$$

For accuracy and stability, we use a large number of simulation runs (10000 periods in our case) to estimate the Jacobian. Note that the number of simulation runs in this step is only about numerical accuracy, not statistical inference. Finally, the asymptotic variance-covariance matrix of the estimated parameters is computed using

$$V = (D'WD)^{-1}D'WSWD(D'WD)^{-1},$$

Table A7. Parameter estimates under alternative moment sets. This table reports model parameter estimates obtained using alternative sets of target moments. Standard errors are shown in parentheses. The "baseline" columns correspond to the main specification in the main text. The "no beta" columns exclude the bond–stock beta moment from the objective function. The "no AP" columns exclude both the bond–stock beta moment and the convenience–inflation regression coefficient.

	1952–1999			2000–2020		
	baseline	no beta	no AP	baseline	no beta	no AP
σ_ℓ (liquidity demand)	0.100	0.100	0.172	0.079	0.079	0.000
	(0.002)	(0.006)	(0.003)	(0.004)	(0.007)	(0.271)
σ_π (cost-push)	0.793	0.793	0.699	0.138	0.170	0.228
	(0.009)	(0.008)	(0.006)	(0.010)	(0.012)	(0.005)
σ_i (monetary policy)	1.294	1.294	1.566	0.906	0.906	0.901
	(0.017)	(0.050)	(0.020)	(0.022)	(0.051)	(0.045)
σ_x (non-liquidity demand)	0.050	0.050	0.000	0.534	0.391	0.767
	(0.038)	(0.409)	(0.576)	(0.009)	(0.033)	(0.023)

where as explained before, S denotes the estimated covariance matrix of moment residuals, D is the Jacobian matrix of moment residual, and W is the weighting matrix. Then we report the square roots of diagonal matrix V as the standard deviation of estimated parameters, reported as numbers in parentheses in second columns of each period in Table 5.

The Role of Bond-Stock Beta and Inflation-Convenience Coefficient. Next, we evaluate how asset pricing moments affect our model estimations. We will do two exercises: First, we remove the bond-stock beta and we call this case "no beta"; Second, we remove both bond-stock beta and the regression coefficient of T-bill convenience yield on inflation, and we call this case "no AP". These exercises are informative about the role of these extra moments. We report the estimated parameter values under these alternative moment sets in Table A7 and the corresponding moment values in Table A8.

In Table A7, we find that removing the bond–stock beta moment does not affect parameter estimates for the 1952–1999 period, but leads to a smaller σ_x and a slightly larger σ_π for 2000–2020. Estimation errors increase overall because the set of targeted moments is smaller.

In Table A8, most model moments remain similar under the "no beta" specification in both periods. We therefore conclude that, conditional on including the convenience yield–inflation regression coefficient as a target moment, the bond–stock beta moment plays only a marginal

Table A8. Model moments under alternative moment sets. This table reports model-implied moments obtained under different sets of target moments. The "baseline" columns correspond to the main specification in the text. The "no beta" columns exclude the bond–stock beta moment from the objective function. The "no AP" columns exclude both the bond–stock beta moment and the convenience–inflation regression coefficient.

	1952–1999			2000–2020		
	baseline	no beta	no AP	baseline	no beta	no AP
Vol(Tbill spread)	0.476	0.476	0.584	0.291	0.289	0.254
Vol(Inflation)	2.374	2.374	2.180	0.679	0.726	0.857
Vol(Output Gap)	2.665	2.665	2.701	1.606	1.452	2.007
Corr(Inflation, Output Gap)	-0.312	-0.312	-0.149	0.427	0.341	0.253
Bond return $\sim \beta \cdot$ Stock return	0.157	0.157	0.068	-0.049	-0.005	0.079
Tbill spread $\sim b \cdot$ Inflation	0.091	0.091	-0.004	-0.010	-0.002	0.013

role in disciplining the parameters. This suggests that the convenience yield–inflation correlation is more informative than the bond–stock beta, despite their similarity.

We next consider the "no AP" case, shown in the third column of each period. In Table A7, for 1952–1999, the estimated liquidity-demand shock σ_ℓ becomes much larger while the non-liquidity demand shock σ_x becomes smaller, though the standard error on σ_x is also much larger. This change in the relative importance of liquidity versus non-liquidity demand shocks causes the model-implied convenience–inflation coefficient b to switch sign from positive to negative, as reported in Table A8, and this is opposite to the data.

For 2000–2020, when the convenience–inflation coefficient is excluded from the moment set, the implied liquidity-demand shock is estimated to be zero with a large error band, while the non-liquidity demand shock becomes more dominant. As a result, in the last column of Table A8, both the bond–stock beta and the convenience–inflation coefficient b flip to positive values, contrary to the data.

Comparing the no beta" and no AP" cases, we find that the convenience–inflation moment is crucial for the model to distinguish between liquidity and non-liquidity demand shocks, and thus for determining the sign of both the bond–stock beta and the convenience–inflation coefficient.

B.8 Impulse responses to monetary policy and non-liquidity demand shocks

Figure A5 shows impulse responses to a monetary policy shock in our baseline model. We see that inflation and the output gap both decline, whereas short-term convenience and the nominal policy rate increase. This occurs because the shock initially raises the policy rate and later induces overshooting as inflation falls following the contractionary shock. The risk-neutral component of the 10-year yield follows a path similar to the policy rate, but the risk-premium component rises sharply in the second period. Both the risk-neutral and risk-premium components of stock value decline. Overall, the monetary shock generates positive stock–bond co-movement and a sharp decline in inflation. These dynamics may help explain asset-pricing and macroeconomic patterns during the Volcker years.

Figure A6 reports impulse responses to a negative non-liquidity demand shock. We choose to show the response to a negative non-liquidity demand shock so that the macroeconomic responses are comparable to those to a positive liquidity demand shock, i.e. lead to a recession. Similarly to a liquidity demand shock, output gap and inflation fall in response to a negative non-liquidity demand shock, which drives down consumption and output, and hence inflation through the Phillips curve. Different from a non-liquidity demand shock, the convenience spread almost does not move at all, and if anything moves in the same direction as inflation. Bond and stock responses have a very small risk-neutral component, and are dominated by endogenous risk premia, which switch sign with the equilibrium.

B.9 Model extensions

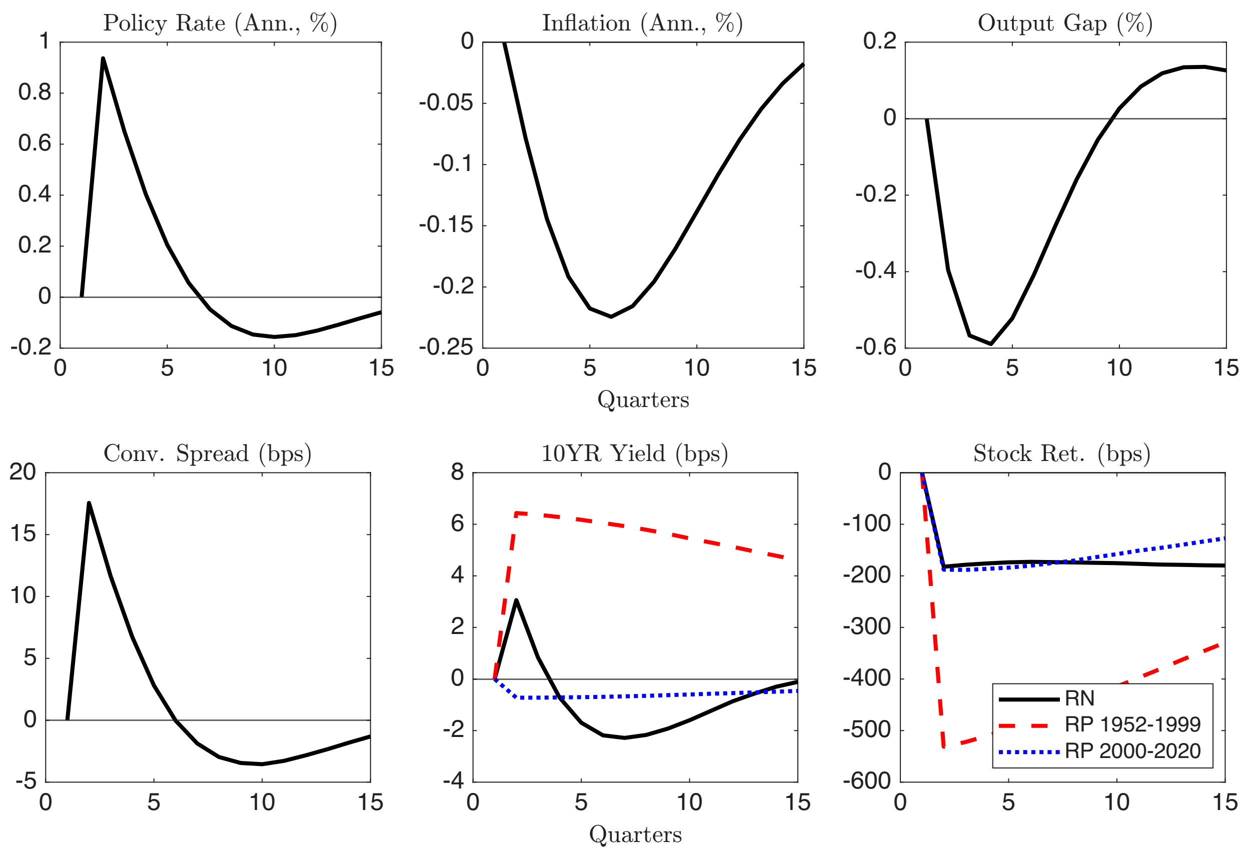
For simplicity, we simplify to the case with no deposit rate inertia, i.e. $\rho^d = 0$, throughout this subsection.

B.9.1 Imperfect substitutability

While our baseline model treats Treasuries and deposits as perfect substitutes, this assumption is made merely for simplicity. To see how the framework generalizes, assume that the liquidity aggregate is given as in Nagel (2016) and Krishnamurthy and Li (2023),

$$Q_t = \left(D_t^{\rho} + \frac{\lambda_t}{1-\lambda_t} B_t^{\rho} \right)^{1/\rho}, \tag{A118}$$

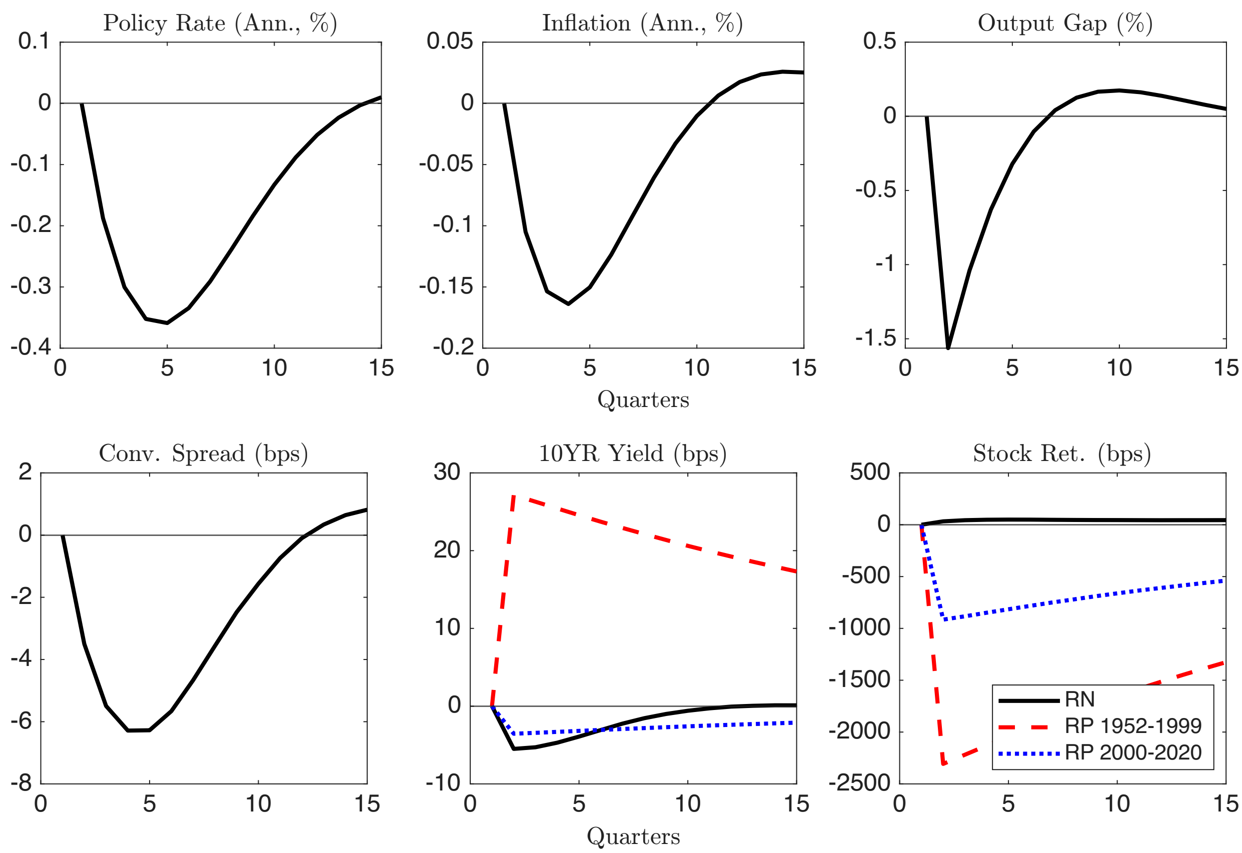
Figure A5. Baseline model responses to a monetary policy shock. This figure shows impulse responses to a 1 percentage point increase in the monetary policy shock $v_{i,t}$, where the impulse period is period 2. The driving shock has mean zero and is drawn from the steady-state distribution in all other periods. All other shocks are drawn from their mean-zero steady-state distributions in all periods. Responses for policy rate i^b, inflation π, and output gap x are in annualized percent units. The response for the convenience spread $i^l - i^b$, 10-year Treasury yield, and stock market are in annualized basis points units. Quarters are shown on the x-axis. Impulse responses are averaged over 10^6 independent simulations with a 30-quarter burn-in period for asset prices in the bottom row. For the policy rate, inflation, and output gap, we reduce simulation noise by setting all non-driving shocks and the driving shock in the non-impulse period to zero.



where the substitutability parameter ρ can be between zero and one. The case with $\rho = 1$ corresponds to perfect substitutability. For general substitutability, ρ, the liquidity premium becomes

$$I_t^l - I_t^b = \frac{\lambda_t}{1-\lambda_t}\left(\frac{B_t}{D_t}\right)^{\rho-1}(I_t^l - I_t^d), \tag{A119}$$

Figure A6. Baseline model responses to a non-liquidity demand shock. This figure shows impulse responses to a negative 1 percentage point decrease in the non-liquidity demand shock $v_{x,t}$, where the impulse period is period 2. The driving shock has mean zero and is drawn from the steady-state distribution in all other periods. All other shocks are drawn from their mean-zero steady-state distributions in all periods. Responses for policy rate i^b, inflation π, and output gap x are in annualized percent units. The response for the convenience spread $i^l - i^b$, 10-year Treasury yield, and stock market are in annualized basis points units. Quarters are shown on the x-axis. Impulse responses are averaged over 10^6 independent simulations with a 30-quarter burn-in period for asset prices in the bottom row. For the policy rate, inflation, and output gap, we reduce simulation noise by setting all non-driving shocks and the driving shock in the non-impulse period to zero.



showing that if $\rho < 1$ an increase in the quantity of bonds outstanding now acts similarly to a decrease in the preference for bonds, λ_t.

Log-linearizing the liquidity spread now gives an additional term depending on the log quantity of debt $\hat{b}_t \equiv \log B_t - \log \bar{B}$ relative to the log quantity of deposits $\hat{d}_t \equiv \log D_t - \log \bar{D}$,

$$\ell_t \equiv i_t^l - i_t^b = (f^i - 1)i_t^b + f^\lambda \hat{\lambda}_t - f^d i_{t-1}^d - f^b \left(\hat{b}_t - \hat{d}_t\right). \tag{A120}$$

Here, the log-linearization coefficient on $(\hat{b}_t - \hat{d}_t)$ is zero in the perfect substitutes case $\rho = 1$ but strictly negative otherwise.

Since $\hat{b}_t$ does not enter the Phillips curve or monetary policy rule, this shows that when deposits and Treasury bonds are imperfect substitutes, shocks to the log ratio of Treasury bonds to deposits $\hat{b}_t - \hat{d}_t$ act on the economy analogously to a liquidity demand shock for Treasuries. Intuitively, when $\rho < 1$, an increase in the amount of Treasuries outstanding relative to Treasuries lowers the marginal utility from holding another Treasury bond. This lowers the convenience yield on Treasuries, and compresses private borrowing rates relative to the monetary policy rate, acting to increase demand just like a negative liquidity demand shock, i.e. $\hat{b}_t \uparrow$ acts analogously to $\hat{\lambda}_t \downarrow$. The inflation-convenience relationship is therefore affected similarly by Treasury supply shocks and liquidity demand shocks, and we focus on the latter throughout the paper for simplicity.

B.9.2 Shocks to overall liquidity demand

A simple extension considers shocks to the overall liquidity weight in the utility function, α. Combining equations (10) and (11) gives

$$E_t\left[M_{t+1}^{\$}\right]\left(I_t^l - I_t^b\right) = \frac{\alpha_t / Q_t \lambda_t}{U_c\left(C_t, Q_t, H_t, N_t, \Theta_t\right)}. \tag{A121}$$

Combining equations (10) and (12) gives

$$E_t\left[M_{t+1}^{\$}\right]\left(I_t^l - I_t^d\right) = \frac{\alpha_t / Q_t (1 - \lambda_t)}{U_c\left(C_t, Q_t, H_t, N_t, \Theta_t\right)}. \tag{A122}$$

In these equations, it appears that an increase in α_t raises the convenience yield on both deposits and Treasury bonds. However, as long as we maintain assumption (7) this possibility is precluded, as α_t is not allowed to enter into the deposit spread by assumption in equilibrium. Substituting (A121) into (A122) then gives equation (13) and the Treasury convenience yield is not affected by α_t. Changes in α_t can therefore not be regarded as a shock to the overall demand for liquidity, as long as assumption (7) is assumed to hold.

Different assumptions are of course possible. For example, one could replace (7) by a relationship that depends on both I_t^l and on α_t to reflect the notion that α_t is a shock that affects the overall demand for liquidity, and lowers the deposit rates that households require. In that case, combining this alternative relationship with (A121) and (A122) makes it straightforward to see that α_t enters the Treasury convenience spread similarly to the deposit spread. By lowering the Treasury convenience yield, shocks to the overall liquidity preference α_t would then enter the log-linearized Euler equation analogously to λ_t, and affect the convenience-inflation relationship similarly to $\hat{\lambda}_t$ in our main model.