Inflation and Treasury Convenience

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Abstract

We document that U.S. Treasury convenience yields 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. The model estimates an increased relative importance of liquidity demand shocks after 2000. This channel weakens the convenience–inflation comovement and contributes to negative bond-stock betas, as distinct from non-liquidity demand shocks.

Keywords: Liquidity demand shocks; supply shocks; bond-stock betas; Keynesian recessions; money channel;

JEL classification: E44, E58, G01, G28

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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., 2018; Krishnamurthy and Vissing-Jorgensen, 2012).

Motivated by these old and new debates on the value of Treasury bonds, 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 repeated, 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-2000 period.[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. After 2000, the switch to negative bond-stock betas coincided 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: it was 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 a New Keynesian block for the 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. Moreover, 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 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 from exogenous liquidity demand. The "money channel" operates by changing the opportunity cost of holding money or liquid deposits, which is the foregone interest. This, in turn, also affects the cost of holding other liquid substitutes, such as Treasuries. 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-a-vis the more traditional time-preference shocks driving demand in New Keynesian models. Among all four shocks, the liquidity demand shock is the only one that unambiguously drives a negative inflation-convenience comovement, 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 comove negatively, consistent with the post-2000 evidence. The liquidity demand channel is thus similar in spirit to Keynes (1937)'s argument that exogenous variation in the preference for liquidity is 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. In our structural estimation, the share of Treasury convenience spread variance attributable to liquidity demand shocks rises from below 1% in 1952–1999 to 43% in 2000–2020, while the share attributable to supply shocks falls from 45% to below 6%. Monetary policy shocks, which are estimated to be volatile in both periods, are interpreted broadly, as capturing deviations from our parsimonious policy rule and other inflation-relevant disturbances that are not directly captured in our model. 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. An adverse non-liquidity demand shock can also generate a positive correlation between inflation and the output gap, but its effect on convenience is very small. Overall, the model implies that the comovement between inflation and convenience yields is a highly informative moment for distinguishing liquidity-driven demand from other demand shocks.

While monetary policy shocks generate substantial volatility in the model, their effect on the key convenience-inflation comovement is small. This arises because the convenience spread response to a positive monetary policy shock endogenously changes sign. The initial increase in the convenience spread reflects sluggish pass-through of the policy rate to the deposit rate, followed by a negative response after several quarters as the policy rate falls along with inflation. This feature further distinguishes monetary policy from cost-push shocks as drivers of the money channel, with cost-push shocks being the quantitatively important source of positive convenience-inflation comovements.

Complementary to the convenience-inflation relationship, the model-implied bond-stock betas can also reveal the dominance of supply, liquidity-demand 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, attributing both to dominant supply shocks. Instead, liquidity shocks in the model drive the hedging properties of nominal Treasuries, depressing bond-stock betas through the liquidity demand channel. Bond-stock betas thus serve as an overidentifying moment in our estimation, confirming the switch in the drivers of Treasury convenience.

Finally, we use post-pandemic data on convenience yields and inflation expectations as an out-of-sample check on the model's predictions. Consistent with a substantial role for liquidity demand at the beginning 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).[2] 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, 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. (2022)). 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 (2025) 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. (2018) 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. Bansal and Coleman (1996) propose a monetary asset-pricing model in which short-term government debt provides transaction services by backing checkable deposits,

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

thereby lowering its pecuniary return. Di Tella et al. (2025) provide complementary evidence on the gap between the stock market-implied zero-beta rate and government risk-free rates, which acts as a shifter in the Euler equation, akin to a demand shock. Li (2025) 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.

A related strand of work links stock-bond comovement to Treasury liquidity. Our finding that liquidity demand shocks were important for explaining the negative bond-stock comovement post-2000 is qualitatively in line with the econometric analyses of Baele et al. (2010) and Antolin-Diaz (2024). In contrast to those papers, we study the drivers of liquidity spreads and their relationship with the macroeconomy and risk premia within a structural model. 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.

In recent work, 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 repeatedly and significantly over the past decades.

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 1991, and the spread between the three-month term repo rate collateralized by Treasuries and the three-month T-bill rate after 1991. 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.[3] We refer to the concatenated series as 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.[4] 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.[5]

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,

[3]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.

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

[5]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.

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 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 from our main analysis, although we analyze this period separately in Appendix A.9. During the WWII period, monetary policy was conducted differently and inflation was hard to measure due to product price controls and wartime rationing.[6] 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

[6]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.

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.[7] These repeated shifts suggest that the inflation-convenience relationship changes with the fundamental drivers of the economy.

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

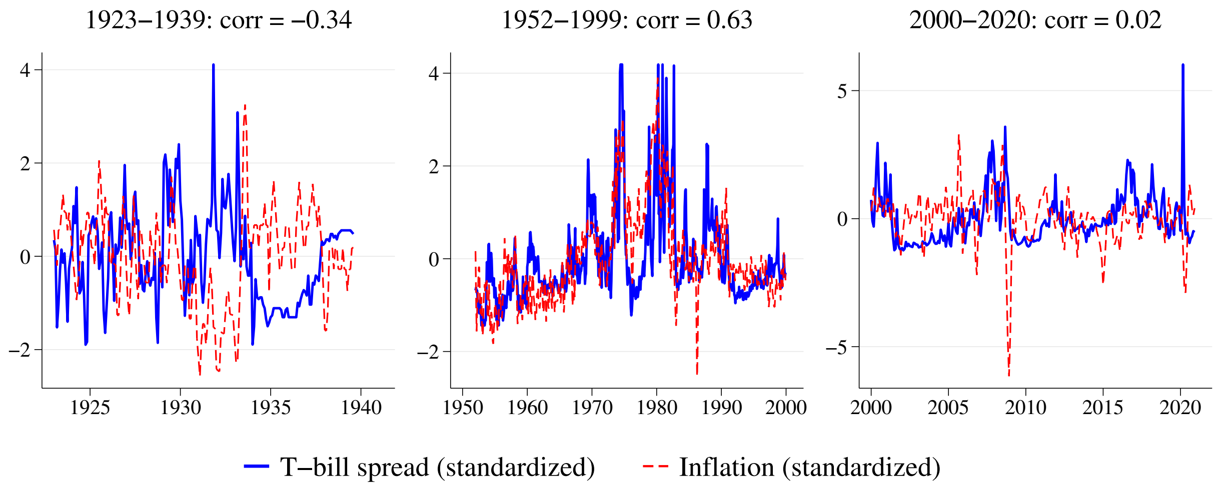
$$\textit{T-bill spread}_t = b_0 + b_1\pi_t + b_2\pi_t \times I_{\text{1952-1999},t} + b_3\pi_t \times I_{\geq 2000,t} + b_4 I_{\text{1952-1999},t} + b_5 I_{\geq 2000,t} + \Gamma' X_t + \varepsilon_t, \tag{1}$$

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

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, relative to the pre-WWII period, 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 1.5 bps decrease in

[7] While we exclude the WWII/immediate post-WWII period (1940–1951) from the regressions below, correlation was also positive at 0.31 (Appendix Figure A5).

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.



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 comprehensive 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 re-

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***	-0.0098***	-0.010***
	(-3.87)	(-4.62)	(-3.53)	(-4.01)	(-3.69)
Inflation x $I_{1952-1999}$	0.12***	0.059***	0.11***	0.053***	0.053***
	(7.93)	(3.71)	(7.73)	(3.74)	(3.53)
Inflation x $I_{\geq 2000}$	0.016**	-0.00064	0.0097	0.0060	0.0060
	(2.29)	(-0.11)	(1.41)	(0.90)	(0.91)
FFR		0.081***		0.082***	0.082***
		(8.11)		(8.15)	(8.07)
Debt/GDP			-0.35	0.19	0.19
			(-1.61)	(0.90)	(0.90)
VIX				0.010***	0.010***
				(4.75)	(3.48)
Baa spread					-0.0069
					(-0.16)
$I_{1952-1999}$	-0.10	-0.11*	-0.052	-0.050	-0.057
	(-1.56)	(-1.87)	(-0.73)	(-0.75)	(-0.68)
$I_{\geq 2000}$	-0.022	0.10*	0.092	0.076	0.072
	(-0.46)	(1.94)	(1.01)	(0.95)	(0.86)
Constant	0.25***	0.026	0.32***	-0.26***	-0.26***
	(6.33)	(0.56)	(5.92)	(-3.08)	(-2.97)
$\bar{R}^2$	0.43	0.56	0.43	0.59	0.59
N	1028	1028	1028	1028	1028

duce 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

[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.

(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.

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. 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. Moreover, this positive convenience-inflation relationship in the middle period is robust to alternative break dates, as shown in Appendix A.6.

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 $I_{1979-1980}$ equal to one in 1979 and 1980. Column (5) controls for the rolling bond-stock 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. * $p<0.10$, ** $p<0.05$, *** $p<0.01$.

	1952–2020				1961:11–2020		
	(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)	(3.03)	(2.83)	(2.73)	(3.48)	(1.12)
Inflation x $I_{\geq 2000}$	-0.10***	-0.059***	-0.045***	-0.051***	-0.060***	-0.067***	-0.028*
	(-6.62)	(-3.74)	(-2.69)	(-2.63)	(-4.05)	(-3.77)	(-1.87)
FFR		0.083***	0.081***	0.082***	0.10***	-0.014	0.011
		(7.86)	(7.14)	(7.02)	(8.60)	(-0.73)	(0.60)
Debt/GDP			0.22	0.23		-0.26	0.0043
			(1.02)	(1.08)		(-0.68)	(0.01)
VIX			0.011**	0.011***		0.0035	0.0063
			(2.55)	(2.62)		(0.82)	(1.56)
Baa spread			0.0044	0.0060		0.44***	0.18**
			(0.05)	(0.07)		(4.62)	(2.08)
$I_{1979-1980}$				-0.18		-0.44***	-0.41***
				(-0.89)		(-3.46)	(-3.09)
Bond-stock beta					-0.55***	-0.30	-0.17
					(-2.99)	(-1.62)	(-1.08)
$I_{\geq 2000}$	0.082	0.22***	0.11	0.13	0.19***	0.28*	0.19
	(1.37)	(3.39)	(1.30)	(1.41)	(2.98)	(1.82)	(1.40)
Constant	0.15***	-0.094	-0.34***	-0.36***	-0.16***	0.36*	0.43***
	(2.79)	(-1.55)	(-3.04)	(-2.97)	(-2.66)	(1.68)	(2.67)
$\bar{R}^2$	0.43	0.57	0.59	0.60	0.61	0.41	0.23
N	828	828	828	828	710	710	710

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 bond-stock 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 bond-stock beta is significantly negative, consistent with a hedging premium component of convenience as emphasized by Acharya and Laarits (2025). At the same time, the bond-stock beta does not affect the positive relationship between inflation and convenience before 2000 or the decline in this relationship after 2000. One explanation, consistent with our model, is that the convenience-inflation relationship reflects realized macroeconomic and liquidity shocks, while the hedging premium embedded in stock-bond beta additionally reflects changing risk aversion.

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 time 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 curve becomes 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

[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.

[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.

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 significantly positively correlated with inflation 1961–1999, and that this relationship declines significantly after 2000, turning slightly negative. 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. To position these facts in a broader macroeconomic and financial markets context, we now discuss the evolving properties of bond-stock betas and the inflation-output gap correlation. These moments inform the estimation of our model, and we report them in Table 4 alongside the model-implied moments. All moments in this subsection are estimated 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 0.214, 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

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.

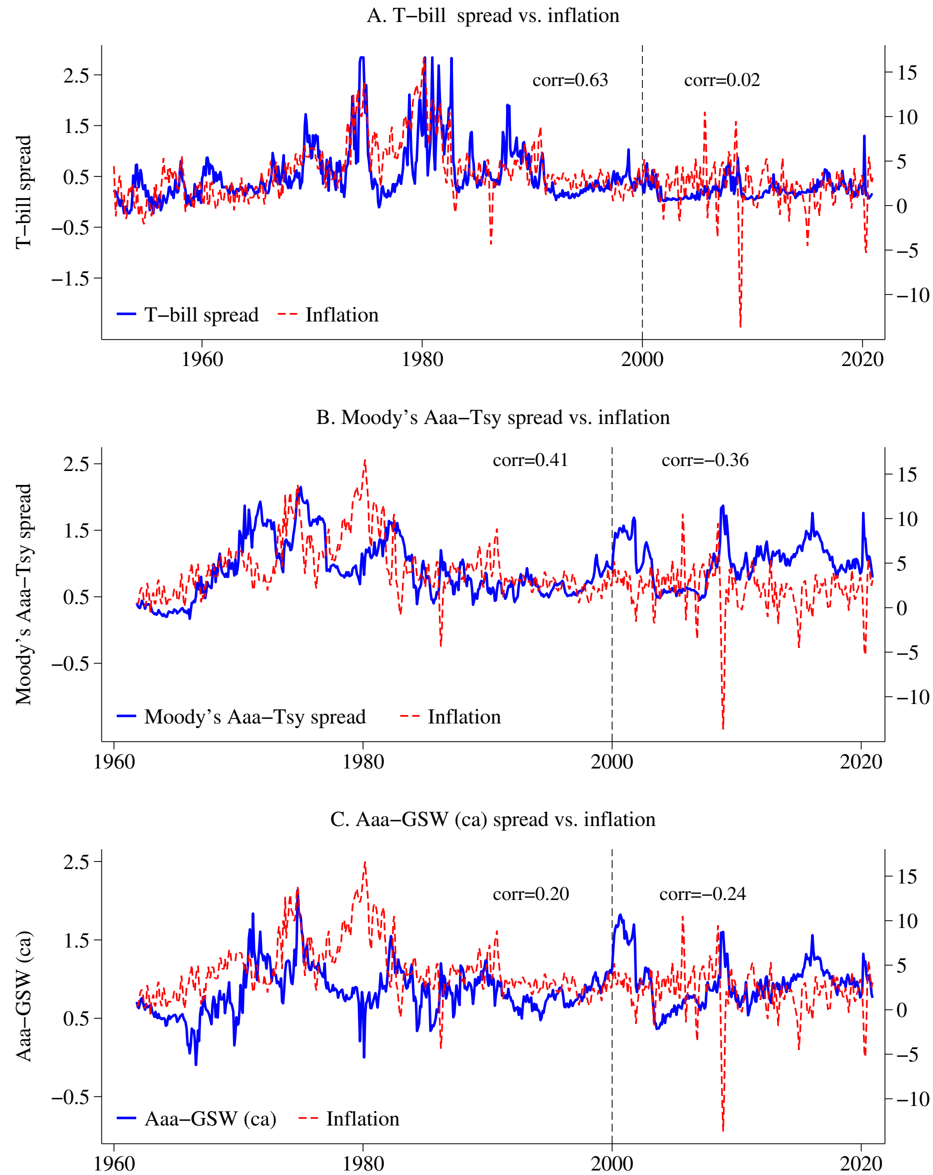


Table 3. Sample split by bond-stock betas. This table estimates regressions of the T-bill spread onto annualized quarterly inflation. The sample is quarterly from 1961:Q4 to 2020:Q4. We use the daily seven-year nominal Treasury bond returns based on zero-coupon yields from Gürkaynak et al. (2007) to 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.

	(1) Beta Q1	(2) Beta < Med	(3) Beta > Med	(4) Beta Q4
Inflation	-0.00528	0.0598**	0.106***	0.102***
	(-0.62)	(2.23)	(5.72)	(4.37)
Constant	0.283***	0.236***	0.153*	0.306***
	(7.61)	(3.89)	(1.75)	(2.77)
$\bar{R}^2$	-0.0131	0.159	0.313	0.302
N	59	119	118	60

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.

We present additional results on inflation-convenience comovement conditioning on bond-stock betas as an alternative sample split. 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. 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, similar to the post-2000 period. It is likely no coincidence that Keynes developed his "General Theory" during this period (Keynes (1937)). However, many features of the economy were substantially different during the Great Depression and WWII from the modern day, including Treasury bond gold clauses during the earliest years, yield curve and price controls during WWII, measurement issues for output and potential output, the role of deposits, and the conduct of monetary policy (Friedman and Schwartz, 1963). We therefore specify our model primarily for the two subperiods after 1952, although we provide robustness for the pre-1952 period in Appendix A.9. 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

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

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 (2025)), 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. (2022) provide a microfoundation 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

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

[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. (2025), Bianchi et al. (2025), Engel and Wu (2023), Abadi et al. (2023).

[12]Appendix B.10 considers an extension to imperfect substitutability between deposits and Treasuries, and shows 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.

where

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

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.

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, equals:

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

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:

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

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.[13]

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,

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

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

The 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.[14] 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.[15]

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.[16] In practice, the Fed is not allowed to operate directly through private loan markets, since deciding which 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.[17]

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

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

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

[14] 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).

[15] As long as equation (7) is assumed to hold exactly, as shown in Appendix B.10.2, shocks to overall liquidity preference, α in equation (4), cannot be interpreted as a liquidity demand shock. However, disturbances to the deposit rate pass-through (7) would act analogously to λ_t.

[16] 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.

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

$v_{i,t}$ 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) implies that in equilibrium the log real output gap equals 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

$$E_t\left[M_{t+1}^{\$}\left(1+I_t^l\right)\right] = 1. \tag{10}$$

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^d_t 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:

$$I^l_t - I^b_t = \frac{\lambda_t}{1-\lambda_t}\left((1-\delta)I^l_t - \rho^d I^d_{t-1}\right). \tag{13}$$

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^l_t 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

$$E_t\left[M^{\$}_{t+1}\frac{P^{\$}_{n-1,t+1}}{P^{\$}_{n,t}}\right] = E_t\left[M^{\$}_{t+1}\exp(i^b_t)\right], \tag{14}$$

where the price of a one-period liquid nominal bond equals $P^{\$}_{1,t} = \exp(-i^b_t)$.

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

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

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

$$\frac{P^c_t}{C_t} = \sum_{n=1}^{\infty}\frac{P^c_{n,t}}{C_t}. \tag{16}$$

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^l_t, i^b_t, i^d_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 us to express the convenience spread (up to a constant) as

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

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 $\rho^\ell < \rho^d$. 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 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 $v_{x,t} \equiv \gamma \psi \eta_t$ 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 an increase in impatience 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 ℓ_t, reflecting the difference between the rate at which the household borrows and saves and the policy rate, similar to the model with microfounded liquidity demand by Piazzesi et al. (2022). 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).[18] 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 demand shock $v_{x,t}$ in equation (6). 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 the 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), where these equilibrium relationships contain both the time-t variables and expectations. We solve for minimum state variable dynamics following

$$Z_t = P^Z Z_{t-1} + Q^Z v_t. \tag{22}$$

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 (22). 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

[18] 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)).

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 on the solution method, including robustness with respect to the numerical accuracy, 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 calibrated 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 $(\sigma_\ell, \sigma_\pi, \sigma_i, \sigma_x)$ for each subperiod 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 Christiano et al. (2005). 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 an 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

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 annualized percentage points. The bond-stock beta is the regression coefficient of quarterly excess returns on a ten-year nominal Treasury bond onto S&P 500 excess returns. Standard errors are reported in parentheses next to each data moment.

		1952–1999		**2000–2020**	
		Data	Model	Data	Model
Volatilities	Vol(T-bill spread)	0.579 (0.030)	0.476	0.219 (0.017)	0.291
	Vol(Inflation)	3.277 (0.168)	2.374	2.665 (0.207)	0.679
	Vol(Output gap)	2.287 (0.117)	2.665	1.955 (0.152)	1.606
Correlations	Corr(Inflation, Output Gap)	-0.119 (0.072)	-0.312	0.292 (0.102)	0.427
Bond-Stock Beta	Bond return $\sim \beta\cdot$ Stock return	0.214 (0.069)	0.157	-0.288 (0.052)	-0.049
Reg Coefs	T-bill spread $\sim b\cdot$ Inflation	0.109 (0.010)	0.091	-0.003 (0.009)	-0.010

$\rho^d = 0.92$ quarterly from a time-series regression of the form (7). Combined with this estimate, the calibration target $\delta/(1-\rho^d) = 1/3$ implies a short-run pass-through of $\delta = 0.027$, consistent with the direct regression estimate (0.023) 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 A8.

4.2 Estimation and Identification

4.2.1 Targeted Moments

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 2000–2020. The moments include the convenience-inflation regression coefficient from our baseline empirical results, the nominal bond-stock beta, and the output gap-inflation correlation,[19] as well as the annualized volatilities (in %) 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.[20]

4.2.2 Shock Identification

The three volatility moments discipline the overall scale of shocks, while the comovements discipline the relative magnitudes of shocks. Table 5 below summarizes the signed effects of the four structural shocks on endogenous variables, with all shocks normalized to positive values.

Table 5. Signed effects of structural shocks on key model variables. Effects are signed at the calibrated model parameters, with all shocks normalized to positive values. $\pm$ for the convenience spread response to the monetary policy shock reflects a positive immediate impact response that turns negative after several quarters.

Shock	Short rate i_t	Inflation π_t	Output gap x_t	Convenience spread ℓ_t
Cost-push ($v_\pi \uparrow$)	$+$	$+$	$-$	$+$
Non-liquidity demand ($v_x \uparrow$)	$+$	$+$	$+$	$+$
MP shock ($v_i \uparrow$)	$+$	$-$	$-$	$\pm$
Liquidity demand ($v_\ell \uparrow$)	$-$	$-$	$-$	$+$

The correlation between inflation and the output gap disciplines the relative importance of supply shocks vs. non-supply shocks, inheriting the standard signs in a New Keynesian model. In our calibration, 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. Liquidity demand shocks, negative 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

[19]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 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.

[20]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. See Appendix B.7 for more details of model estimation.

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, liquidity and non-liquidity demand shocks have distinct implications for the convenience-inflation relationship and bond-stock betas, thereby pinning down their relative magnitudes.

4.2.3 Information Content of Inflation-Convenience Comovement

Importantly, the liquidity demand shock is the only shock that unambiguously drives the negative convenience spread-inflation relationship, as seen in Table 5. Non-liquidity demand shocks and supply shocks generate a positive convenience-inflation relationship. The effect of the monetary policy shock on convenience changes sign, from initially positive on impact to negative after several quarters (indicated with $\pm$). The convenience-inflation beta is therefore a highly informative moment that allows us to identify the liquidity-demand shocks separately from other shocks, including the monetary policy shock.

Intuitively, the convenience-inflation beta synthesizes responses of inflation and convenience spread to shocks across various horizons. In the short run, both the liquidity demand shock and the monetary policy shock raise the convenience spread and lower inflation, generating negative inflation-convenience comovement. The distinction arises at longer horizons: for the liquidity demand shock, the convenience spread remains persistently elevated, so its negative comovement with inflation accumulates across all horizons into a large negative beta. For the monetary policy shock, instead, the disinflationary impact endogenously depresses the short rate and the convenience spread below steady state at longer horizons. At those horizons, inflation and convenience comove positively and both are below baseline, offsetting the initial negative policy effect and driving the inflation-convenience beta toward zero.

Beyond identification of exogenous liquidity shocks, the inflation-convenience comovement is also informative about the sources of the money channel, which operates by changing the opportunity cost of holding money and money substitutes. Here, too, the inflation-convenience coefficient disentangles the endogenous effect of inflationary cost-push shocks from that of monetary policy shocks, with the cost-push shocks unambiguously driving inflation and convenience in the same direction.

4.2.4 Estimated Shock Volatilities

Table 6 reports estimated shock volatilities and standard errors for both subperiods. 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.

Table 6. Estimated shock volatilities by period. This table reports the estimated volatilities of the four structural shocks for the two subperiods 1952–1999 and 2000–2020, respectively. All values are in annualized percent. Parameters are estimated by two-stage simulated method of moments (SMM). Standard errors, reported in parentheses, are computed from the GMM asymptotic variance formula. Details of model estimation are in Appendix B.7.

Parameters (annualized %)	1952–1999		2000–2020	
σ_ℓ (liquidity demand)	0.100	(0.033)	0.079	(0.037)
σ_π (cost-push)	0.793	(0.119)	0.138	(0.088)
σ_i (monetary policy)	1.294	(0.230)	0.906	(0.203)
σ_x (non-liquidity demand)	0.050	(0.531)	0.534	(0.086)

The estimated volatility of monetary policy shocks is high in both periods, though somewhat lower after 2000. Because the estimation does not target interest rate volatility and monetary policy shocks are not identified from the inflation-convenience comovement, their volatility is primarily disciplined by the volatility of macroeconomic aggregates. The parsimony of the model means that policy shocks are best interpreted more broadly, as absorbing deviations from a potentially misspecified monetary policy rule (e.g., during the ZLB) and other shocks driving inflation that are largely orthogonal to the inflation-convenience comovement.

4.2.5 Asset Pricing and Untargeted Macro Moments

Table 7 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 stock and bond moments consistent with the data, including the equity premium, equity volatility, equity 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. Appendix Table A13 provides additional non-targeted model autocorrelations, showing that the model-implied persistence of the T-bill convenience spread is within an empirically reasonable range, and lower than the deposit rate persistence parameter, ρ^d.

Table 7. Stocks, bonds, and macroeconomic volatilities in the model and the data. 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. (2007). 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 the risk-neutral component 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		7.46	
Equity Vol	16.72		17.23	
Equity SR	0.54		0.44	
AR(1) pd	0.95		0.94	
1 YR Excess Returns on pd	-0.35		-0.27	
1 YR Excess Returns on pd (R^2)	0.05		0.15	
Macroeconomic volatilities				
Std. Annual Cons. Growth	2.26		1.54	
Std. Annual Change FFR	2.28		1.83	
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		

4.3 Model Implications for Macroeconomic and Asset Price Dynamics

We next examine the economic implications of the model, in particular the liquidity demand shocks and supply shocks, for the convenience spread, asset prices, and the macroeconomy.

4.3.1 Variance Decompositions

To document the economic importance of shocks across periods, in Table 8 we calculate the model-based variance decompositions. In 1952–1999, supply shocks explain 45% of the unconditional variance in Treasury convenience and liquidity demand shocks explain very little. This pattern reverses in 2000–2020 where liquidity demand shocks explain 43% of the Treasury convenience variance and supply shocks explain less than 6%. Thus, the estimated shift from supply shocks to liquidity demand shocks represents a large reallocation in the sources of fluctuations in Treasury convenience. 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.[21]

As anticipated from volatility estimates in Table 6, the variance decomposition in Table 8 reveals large contributions of monetary policy shocks to macroeconomic volatilities. While this result may appear at odds with the literature, it is consistent with an interpretation of v_i shocks as broader residuals rather than tightly identified monetary policy shocks. Likewise, the large volatility shares from monetary policy shocks are not in conflict with monetary policy shocks having little impact on the inflation-convenience comovement, which we confirm by decomposing this comovement shock by shock in Figure 6, Panel A. The reason is that convenience responds immediately to a monetary policy shock and then reverses, while inflation shows little initial response and then declines with a lag (see Appendix Figure A6). Monetary policy shocks are hence not the first-order determinant of our mechanism.

The first column of Table 8 shows that liquidity demand shocks also matter for macroeconomic aggregates, such as inflation and the policy rate, but they do not dominate economic fluctuations on average, except during extreme liquidity episodes. On average, liquidity demand shocks pro-

[21] 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. (2024) highlight a complementary mechanism, whereby deposit outflows can trigger bank credit crunches and amplify the real effects of supply shocks.

Table 8. One-year unconditional variance decomposition. Each entry is the share of unconditional variance of the four-quarter sum attributable to each structural shock. Shocks: v_ℓ = liquidity demand, v_π = cost-push, v_i = monetary policy, v_x = non-liquidity demand.

	Share of variance (%)			
Variable	v_ℓ	v_π	v_i	v_x
Panel A: 1952–1999				
Conv. spread ℓ_t	0.9	45.3	53.8	0.0
Inflation π_t	0.0	83.6	16.4	0.0
Output gap x_t	0.0	56.8	43.2	0.0
Policy rate i_t^b	0.0	49.2	50.7	0.0
Panel B: 2000–2020				
Conv. spread ℓ_t	43.1	5.7	47.2	4.0
Inflation π_t	3.5	39.6	54.1	2.8
Output gap x_t	1.3	12.5	65.9	20.4
Policy rate i_t^b	4.1	10.8	77.6	7.5

duce variance shares for inflation and the policy rate somewhat below 5% after 2000. In order to interpret these shares, it is important to keep in mind that the post-2000 period featured long spans when liquidity demand was quiescent, punctuated by liquidity crises, when liquidity demand exerted pronounced downward pressure on inflation, output and policy rates. We next use impulse responses to show that a single large liquidity demand shock generates substantial economic effects in our model.

4.3.2 Impact of the Liquidity Demand Shock

Figure 3 presents responses to a liquidity demand shock ($v_{\ell,t}$). The top row 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 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, with distinct implications for risk premia, as displayed in the bottom middle and right panels of Figure 3.

The responses are generated considering a 100 bps increase in $v_{\ell,t}$. A liquidity 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 the macroeconomy. 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 0.9 percentage points.

Figure 3 shows that the liquidity demand 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 in 2000–2020. Further, it moves inflation and Treasury convenience in opposite directions, explaining our finding of a 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).

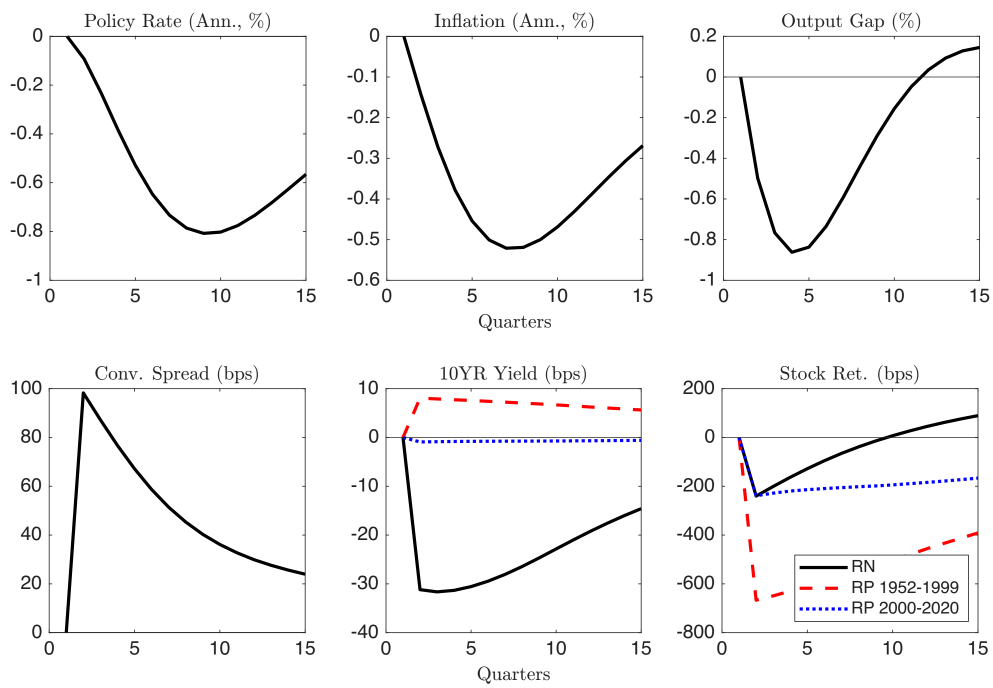
The macroeconomic responses to a positive liquidity demand shock resemble those of a negative non-liquidity demand shock. Both shocks lower the output gap, inflation, and the short rate, since ℓ_t and $v_{x,t}$ both enter as shifters in the Euler equation (20). However, the convenience spread provides a sharp distinction: a positive liquidity demand shock raises convenience (the bottom-left panel in Figure 3), while the non-liquidity demand shock has a negative and quantitatively very small effect. Section 4.5 develops this comparison further in terms of bond-stock betas.[22]

The comovement of long-term nominal Treasury bonds with the stock market provides another marker of liquidity demand shocks. The bottom row of Figure 3 plots responses of the ten-year Treasury bond yield (inversely related to bond prices) and cumulative stock returns since the shock period. Because bonds and stocks in the model price equilibrium variances of shocks, model-implied risk premia are different across the two estimation periods. Risk-neutral responses are shown in solid black, with risk premia in red dashed and blue dotted, for the pre- and post-2000 periods, respectively. Risk-neutral and risk premium responses add up to the total response (not plotted).

The bond yield responses are dominated by the risk-neutral component, so overall Treasury yields drop (and bond prices rise), mirroring the declines in inflation and the policy rate. Stocks fall due to the decline in the output gap. A peak reduction in the output gap of about 0.9 percentage points is associated with a roughly 200 bps decline in risk-neutral stock returns. Investors become

[22] The macroeconomic impulse responses to a non-liquidity demand shock are standard and hence relegated to Appendix Figure A7.

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.



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] 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 in 2000–2020.

[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.

Figure 3 also shows that the bond risk-premium response switches sign across periods, with liquidity shocks amplifying the positive bond-stock comovement for the pre-2000 period and the negative bond-stock comovement post-2000. We further clarify the importance of liquidity shocks for amplifying risk premia in a counterfactual analysis in Section 4.4.

4.3.3 Impact of the Supply Shock and the Money Channel

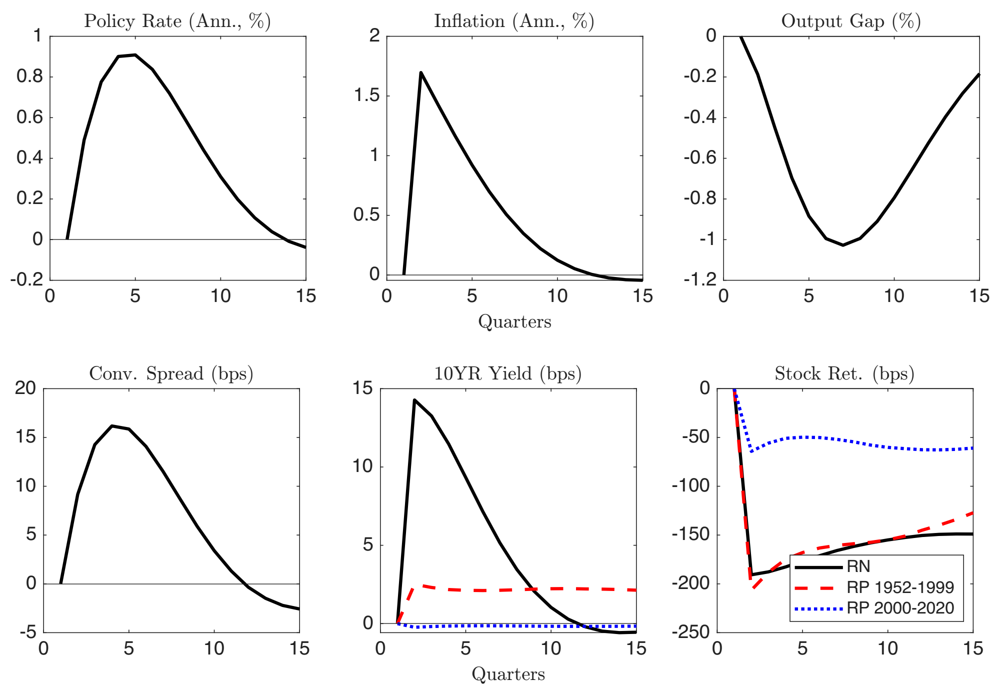
Our estimates indicate that supply shocks are an essential source of the convenience spread fluctuations in 1952–1999 via the money channel, absent significant liquidity demand shocks. The impulse responses in Figure 4 show that a supply shock persistently lifts convenience together with inflation. The 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 other typical characteristics of the 1952–1999 data. They raise inflation just as the output falls, implying a negative inflation-output gap correlation. An adverse supply shock increases ten-year Treasury nominal bond yields to reflect both higher inflation and the expected policy rate tightening. This depresses bond prices just as stocks decline due to lower risk-neutral dividends and risk-bearing capacity. Hence, an additional feature of supply shocks in the model is that they generate a positive nominal bond-stock beta, consistent with the 1952–1999 data (see Table 4).

The model also delivers predictions for how monetary policy shocks impact convenience (see Appendix Figure A6). Since a monetary policy shock moves the policy rate and inflation in opposite directions, it generates offsetting effects on the convenience spread across horizons (see Section 4.2.3). As a result, monetary policy shocks can neither explain the positive convenience-inflation relationship in 1952–1999 nor the negative relationship post-2000.

[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.

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.

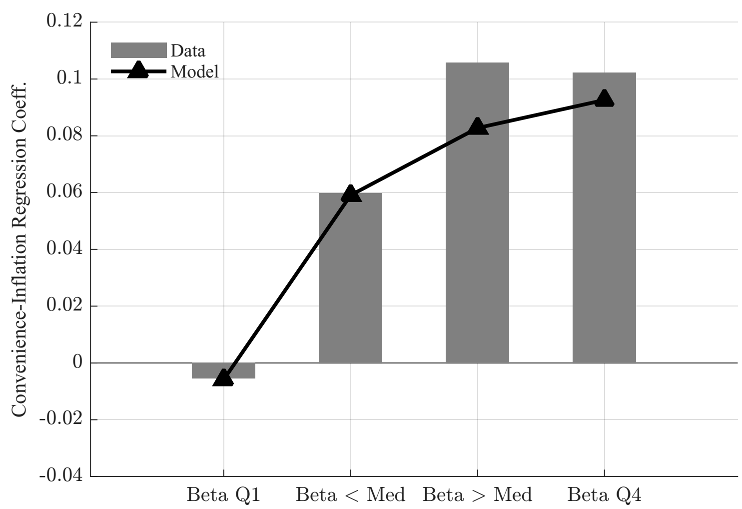


4.4 Bond-Stock Betas and the Convenience-Inflation Relationship

We now turn to model counterfactuals to analyze how changes in shock volatilities alter equilibrium outcomes. These counterfactuals illustrate the key role of liquidity shocks in explaining the inflation-convenience relationship and bond-stock betas post-2000.

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 coefficients from a regression of the convenience spread ℓ_t on inflation π_t, with both variables in annualized percentage units. For the model, we compute moments across four counterfactuals, moving from the 1952–1999 shock volatilities to the 2000–2020 shock volatilities in equal increments (black line with triangles). For the data, we use the point estimates from the subperiod regressions in Table 3 (gray bars).

Figure 5. Convenience-inflation regression by bond-stock betas. This figure shows the regression coefficient b 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 1961–2020 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 is in the bottom quartile (Beta Q1), below median (Beta $<$ Med), above median (Beta $>$ Med), and in the 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.



The model successfully replicates the empirical evidence in Table 3, where the convenience-inflation relationship is most positive in subperiods when bond-stock beta is also highest. 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 relationship, consistent with the data. At the same time, this equilibrium also generates the most negative bond–stock beta.

The money channel is expected to be stronger when large supply shocks are present, leading to positive bond-stock betas along with a positive inflation-convenience comovement. Accordingly, 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. Effects of model shocks on bond-stock covariance and convenience-inflation relationship. Each row increases one shock volatility by 100 bps relative to the 2000–2020 baseline. Panel A shows the model regression coefficient of the T-bill convenience spread onto inflation. Panel B shows the model covariance between 10-year nominal bond returns and stock returns, where both returns are quarterly log excess returns expressed in percentage points (not annualized). The covariance is decomposed into risk-neutral and risk premium components, which add up to the total covariance.

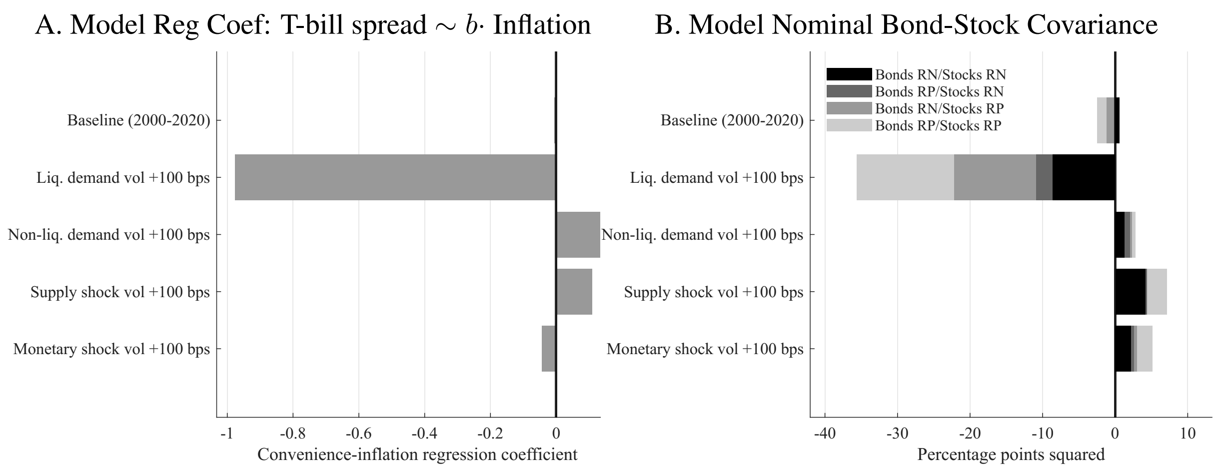


Figure 6 Panel A shows that the liquidity demand shock is the most potent driver of a negative convenience spread-inflation relationship.[25] By contrast, supply shocks and non-liquidity demand shocks generate positive convenience spread–inflation comovements through the money channel, different from the post-2000 evidence. Monetary policy shocks induce a negative but quantitatively very small convenience–inflation comovement, consistent with their sign-reversing effect on convenience spread discussed in Section 4.2.3. Therefore, policy shocks also cannot account for the negative inflation-convenience relationship in the more recent period.

Likewise, Figure 6 Panel B shows that only liquidity demand shocks generate a negative bond–stock covariance in equilibrium, consistent with the empirically observed negative bond–stock

[25]Each row in Figure 6 increases the volatility of one shock at a time, using the 2000–2020 period as a baseline.

betas during the 2000–2020 period.[26] 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.

Importantly, time-varying risk premia in bonds and stocks also comove strongly negatively in the counterfactual with more volatile liquidity demand shocks (light gray bar in the second row of Figure 6 Panel B). While Figure 3 shows a small negative bond risk premium response to a liquidity demand shock in 2000–2020, this response is amplified when liquidity demand shocks are priced as more dominant. When bonds are safe in terms of real cash flows, as in the 2000–2020 estimation, bonds benefit from an increase in risk aversion beyond risk-neutral valuation, driving down yields just as stocks fall. Dominant liquidity shocks thus contribute to the endogenous "flight-to-safety" channel via risk premia (Campbell et al., 2020, 2025), consistent with an increased importance of bond hedging premium from the late 1990s onward documented in Cieslak and Pang (2021).

Overall, the quantitative results in Figure 6 also clarify that the inflation-convenience relationship is the key identifying moment for our estimation, while bond-stock betas serve an overidentifying function once that moment is included.

4.5 Distinguishing Between Non-Liquidity and Liquidity Demand Shocks

Non-liquidity demand shocks in our model are distinguished from liquidity demand shocks via the convenience-inflation relationship, and from supply shocks via the inflation-output gap correlation.

Non-liquidity demand shocks could in principle generate a positive convenience-inflation relationship in 1952–1999. Nonetheless, we fail to find that non-liquidity demand shocks played a large role during this period. The reason is that volatile demand shocks cannot rationalize the negative inflation-output correlation during 1952–1999, which we target in the estimation. We estimate a substantial non-liquidity demand shock volatility 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; only liquidity demand can.

[26]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.

Our analysis also offers 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 positive liquidity demand shock tends to raise bond prices just as the economy and stock market fall, implying a negative bond-stock beta. In contrast, a negative non-liquidity demand shock, corresponding to an increase in patience, raises the risk-neutral valuations of all long-term assets, both bonds and stocks, and thereby tends to push bond-stock betas upwards. Because risk premia also change with the equilibrium, the bond-stock comovement is pushed further upwards when volatile non-liquidity demand shocks are priced in equilibrium. Non-liquidity demand shocks, hence, cannot explain negative bond-stock betas during the post-2000 period.

To summarize, large supply shocks are an essential element of the money channel inducing a positive comovement between convenience spreads and inflation. Instead, 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, negative bond-stock betas, and contributed to a positive inflation-output correlation.

4.6 Robustness

To better understand the role of the inflation–convenience relation and bond-stock beta in disciplining model parameters, we consider two alternative estimation strategies: (1) omitting the bond-stock beta moment ("no beta"); and (2) omitting both the bond-stock beta moment and the convenience–inflation regression coefficient ("no AP"). We summarize the main insights below and relegate detailed results to Appendix Tables A11 and A12.

When we remove only the bond-stock beta moment, results are largely unchanged for 1952–1999, and shift modestly for 2000–2020; most model moments remain similar. Bond-stock betas therefore primarily play an overidentifying role for the estimation. Instead, when we omit both asset-pricing moments, the relative volatilities of the liquidity and non-liquidity demand shocks become very imprecisely estimated. Post-2000 especially, the model is unable to clearly differentiate these two types of demand shocks, and it fails to match the sign of the convenience–inflation coefficient in the data. This is consistent with inflation–convenience relation being a highly informative moment for differentiating between liquidity and other shocks, in particular non-liquidity demand shocks.

We also extend the model estimation to cover the earlier 1923–1939 and 1940–1951 periods in Appendix Tables A5–A6. Applying the model to those earlier years is challenging, not only because of the data quality, but also because the nature of shocks and the conduct of monetary policy were different in this period (e.g., Friedman and Schwartz (1963)). While we treat the pre-1952 estimates mainly as qualitative, the results support a shifting dominance of liquidity demand shocks as a repeated pattern, rather than a single break. The Great Depression period in particular features a high volatility of liquidity demand shocks, which subsequently declines during the 1940s.

5 Discussion: Post-COVID Period Inflation and Convenience

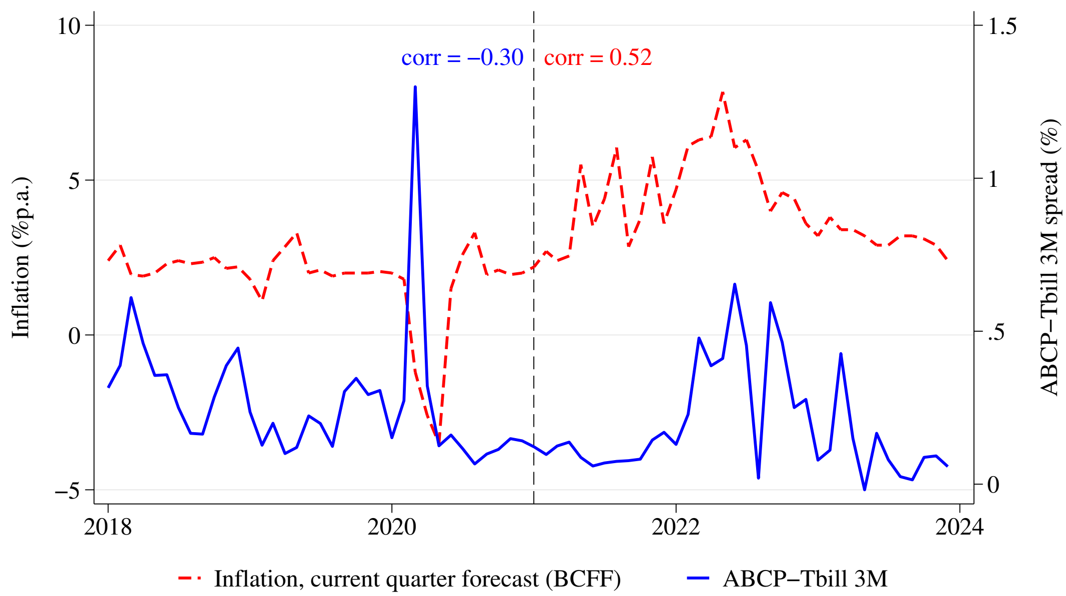
Our main analysis reveals structural shifts in the relation between inflation and the Treasury convenience yield over a long historical sample. These shifts are not isolated to historical episodes but remain relevant today.

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.[27] 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 during 2021. This pattern is robust across alternative convenience proxies, such as the Refcorp–Treasury spread, and to controlling for FFR and Debt/GDP, as we confirm in Appendix A.10.

During 2020, the economy witnessed a sharp drop in inflation due to the initial negative demand

[27] 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 qualitative pattern is similar if we use realized quarterly inflation: the inflation–spread correlation is negative in the early sample (2018–2020) and positive in the later sample (2021–2023). We do not use the Cleveland Fed inflation expectation because it is contaminated by term premia, as shown in Appendix A.5.

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.



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.[28]

Our model offers a framework for interpreting the shift in the inflation–convenience comovement 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 also observe that the initial spike in the convenience spread preceded the decline in inflation. Beginning

[28]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).

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.

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 (e.g., Bianchi and Ilut, 2017; Bianchi et al., 2023).

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 the 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 coincided with periods featuring 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. A liquidity demand shock thus simultaneously induces a negative inflation-convenience relationship and a negative bond-stock beta, 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. These relationships have changed repeatedly in the past century and, hence, may change again as the drivers of Treasury convenience continue to evolve.

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Online Appendix to
"Inflation and Treasury Convenience"

Anna Cieslak, Wenhao Li, and Carolin Pflueger

A Data and Robustness of Empirical Results

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

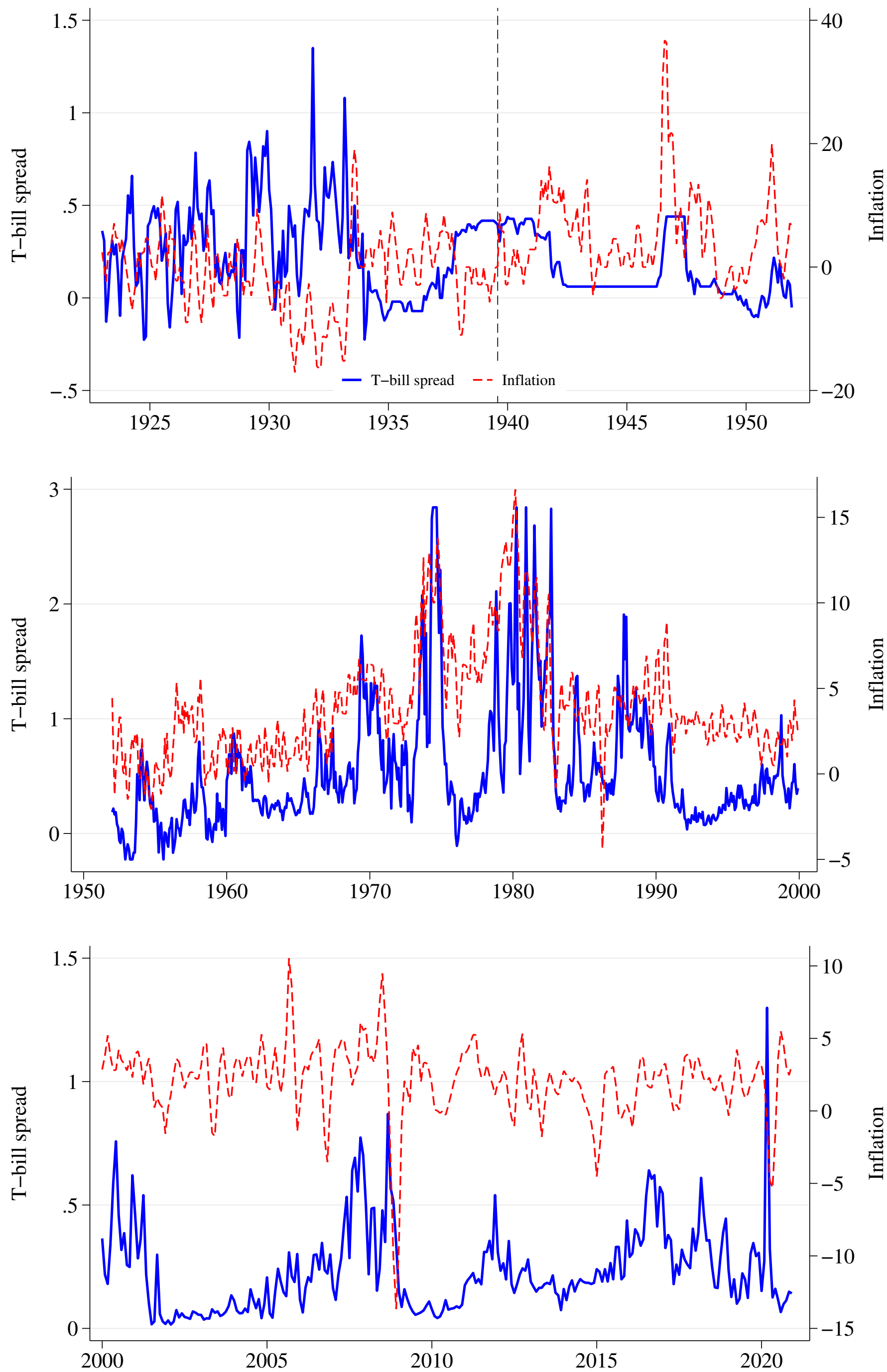
Table A1 reports the summary statistics for main variables used in empirical analysis.

Table A1. Summary statistics. This table presents summary statistics for our 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)
T-bill 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	8.42	25.9	17.4	20.0
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 over the full 1923–2020 sample, including the 1940–1951 period excluded from our analysis. The three subperiods used in the empirical analysis are 1923:01–1939:08, 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 1991, and the spread between 3-month term repo rate collateralized by Treasuries and 3-month T-bill rate after 1991. 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 positive T-bill-inflation relationship in 1952–1999 is somewhat attenuated but remains highly significant and economically similar 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	-0.015***	-0.0074	-0.013***
	(-3.87)	(-1.58)	(-3.47)
Inflation x $I_{1952-1999}$	0.12***	0.11***	0.10***
	(7.93)	(6.51)	(5.79)
Inflation x $I_{\geq 2000}$	0.016**	0.017***	0.014*
	(2.29)	(2.88)	(1.80)
Baa spread		0.12**	0.028
		(2.15)	(0.84)
Baa spread x $I_{1952-1999}$			0.29**
			(2.34)
Baa spread x $I_{\geq 2000}$			-0.025
			(-0.40)
$I_{1952-1999}$	-0.10	0.036	-0.28**
	(-1.56)	(0.33)	(-2.39)
$I_{\geq 2000}$	-0.022	0.062	0.027
	(-0.46)	(0.86)	(0.27)
Constant	0.25***	0.029	0.20**
	(6.33)	(0.24)	(2.49)
$\bar{R}^2$	0.43	0.44	0.47
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 columns (1) and (5), using alternative inflation measures. Columns (1)–(4) show robustness for the no-control specification of Table 1, column (1); columns (5)–(8) show robustness the full-control specification of Table 1, column (5), which adds the federal funds rate, Debt/GDP, the VIX, and the Baa spread. 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.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	$\pi_{t-3,t}$	$\pi_{t,t+3}$	$\pi_{t-12,t}$	$\pi_{t,t+12}$	$\pi_{t-3,t}$	$\pi_{t,t+3}$	$\pi_{t-12,t}$	$\pi_{t,t+12}$
Inflation	-0.015***	-0.010**	-0.028***	-0.019**	-0.010***	-0.0040	-0.020***	-0.0095*
	(-3.87)	(-2.13)	(-5.29)	(-2.26)	(-3.69)	(-1.17)	(-3.19)	(-1.80)
Inflation x $I_{1952-1999}$	0.12***	0.10***	0.15***	0.11***	0.053***	0.027*	0.078***	0.031**
	(7.93)	(5.89)	(9.14)	(5.50)	(3.53)	(1.68)	(3.54)	(2.25)
Inflation x $I_{\geq 2000}$	0.016**	0.0021	0.066***	0.0012	0.0060	-0.0089	0.0077	-0.029**
	(2.29)	(0.21)	(3.97)	(0.07)	(0.91)	(-1.55)	(0.30)	(-2.35)
FFR					0.082***	0.093***	0.076***	0.096***
					(8.07)	(8.66)	(6.21)	(8.85)
Debt/GDP					0.19	0.14	0.21	0.18
					(0.90)	(0.67)	(0.98)	(0.83)
VIX					0.010***	0.010***	0.011***	0.010***
					(3.48)	(3.19)	(3.58)	(3.19)
Baa spread					-0.0069	0.015	-0.049	0.014
					(-0.16)	(0.38)	(-0.95)	(0.34)
$I_{1952-1999}$	-0.10	-0.052	-0.16**	-0.049	-0.057	0.0087	-0.13	0.0055
	(-1.56)	(-0.68)	(-2.45)	(-0.63)	(-0.68)	(0.10)	(-1.37)	(0.06)
$I_{\geq 2000}$	-0.022	-0.0079	-0.089	0.023	0.072	0.13	0.047	0.18**
	(-0.46)	(-0.15)	(-1.63)	(0.39)	(0.86)	(1.45)	(0.43)	(1.98)
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)	(-2.97)	(-3.66)	(-1.85)	(-4.03)
$\bar{R}^2$	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

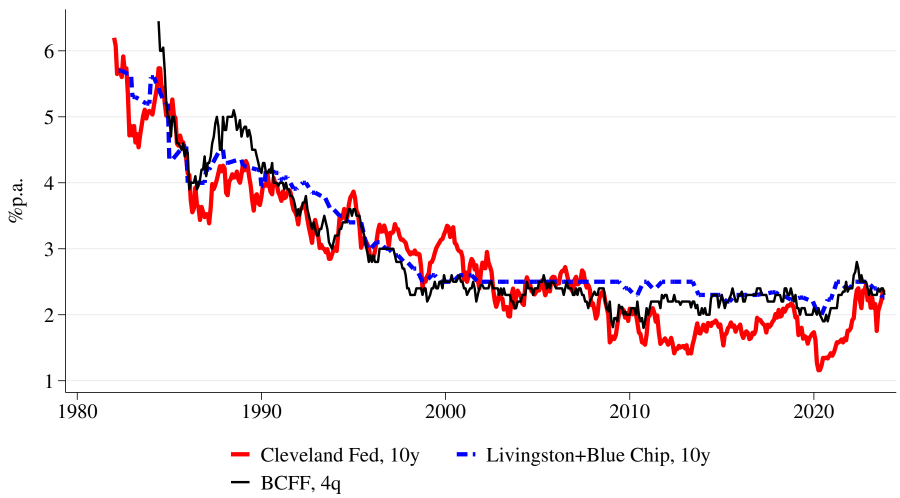
A.5 The Cleveland Fed Index, Inflation Expectations, and Term Premia

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.

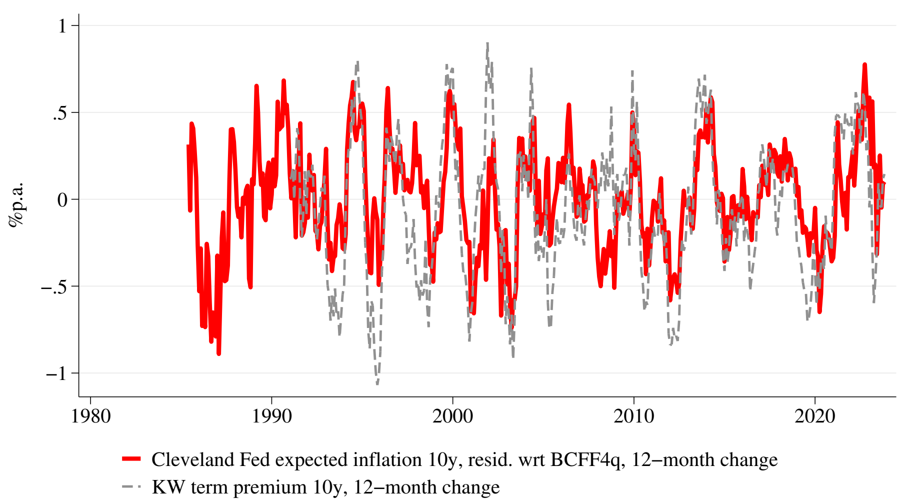
Panel A of Figure A2 shows 10-year inflation expectations from the Cleveland Fed against long-term and 4-quarter consensus inflation expectations from surveys. It is clear that the Cleveland Fed inflation expectations roughly move similarly at lower frequencies, but are substantially more volatile, raising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

Panel B of Figure A2 shows 12-month 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 0.63, further confirming the likely presence of term premia and convenience in the Cleveland Fed inflation expectations. Term premium contamination may not be an issue if the objective is merely to obtain an unbiased forecast of long-term inflation. However, 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 that a shock that lowers the 10-year Treasury bond yield due to term premium decline or increased convenience, is likely to lower the Cleveland Fed measure even if actual inflation expectations 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 (BCFF). 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



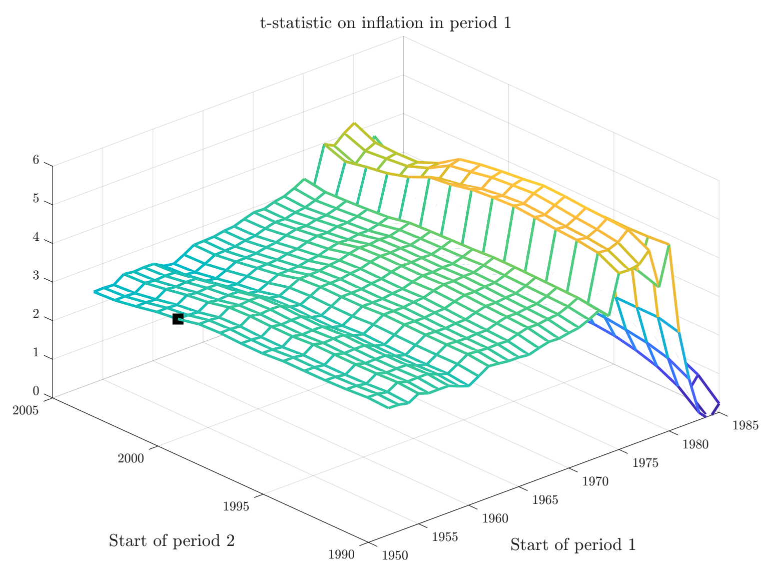
(b) Cleveland Fed inflation expectations vs. term premia

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:

$$\textit{T-bill spread}_t = b_0 + b_1\pi_t + b_2\pi_t \times I_{\text{Start of period 2},t} + b_3 I_{\text{Start of period 2},t} + \Gamma' X_t + \varepsilon_t \text{ if year} > \text{Start of period 1} \tag{A2}$$

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
	(0.015)
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.

Flower bonds. As recently highlighted by Lehner et al. (2025), the benchmark long-term Treasury yield used by Moody's includes the so-called flower bonds.[29] 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.

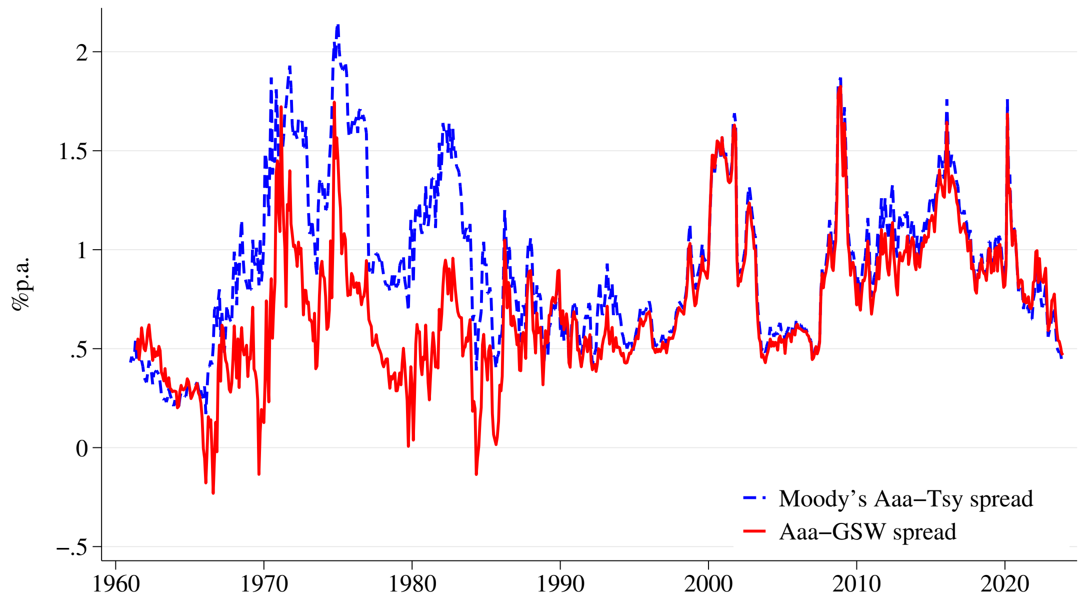
To study the relationship between inflation and long-term convenience, we construct the long-term Treasury yield benchmark using the Gürkaynak et al. (2007, GSW) Treasury yield dataset. Importantly, GSW exclude securities with option-like features, including callable bonds and flower bonds. The GSW sample 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 the longest available maturity in GSW, assuming a flat par yield curve beyond the last available maturity knot.[30]

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

[29] 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.

[30] 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.

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.



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 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[31] is 0.99 in levels and 0.79 in monthly changes. The average absolute duration difference is 0.35 years with a standard

[31]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.

deviation 0.3 years.[32]

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, pushing corporate bond yields up. 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).[33] 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. As rates rose during the 1970s, the declining call premium could make the raw spread understate the increase in the underlying spread; instead, the rising call premium could overstate it in the post-Great Inflation period as interest rates fell through the late 1990s. 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.[34]

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[35] 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.

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.

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

[33]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.

[34]The decline in the call option's value partially offset the bondholder's straight-bond losses as rates rose. Indeed, option-adjusted estimates of bond excess premium in Gilchrist and Zakrajšek (2012) that those time-series spreads rose higher in the late 1970s and early 1980s than the unadjusted spreads did.

[35]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.

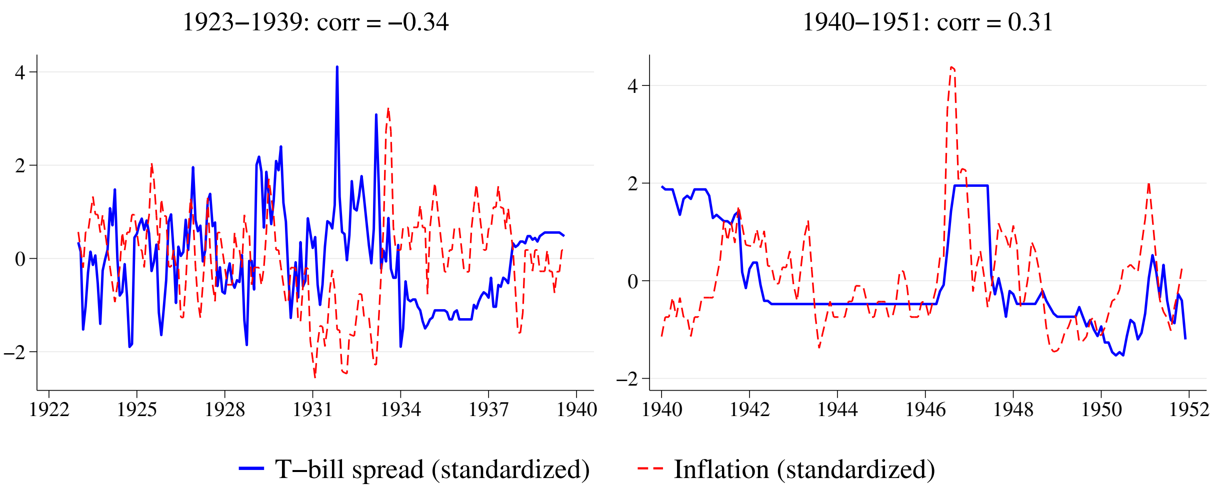
A.9 Pre-1952 Convenience and Inflation

This section extends our data and model estimation to the pre-1952 period. We caution that the historical period before 1952 was substantively different from the post-war period, so the quantitative application of our model to the pre-1952 period is challenging. Among other reasons, measurement of inflation and the output gap was different, and it is reasonable to think that monetary policy was not governed by an interest rate rule or even any rule at all prior to 1952 (Friedman and Schwartz, 1963). We nonetheless provide a simple estimation of our model for these earlier periods. We divide the pre-1952 period into the pre-WWII or Great Depression (1923–1939) and WWII and the immediate aftermath (1940–1951). In 1940, the Great Depression had largely concluded and fiscal pressures on inflation instead started to mount. Of course, the war economy led to numerous changes in the economy that are not captured by our model. Model results for these early periods are therefore to be treated with great caution.

We first document the data for the pre-1952 period, including WWII. Figure A5 shows our main measures, the T-bill convenience spread and inflation, for the pre-WWII (1923–1939) and WWII/immediate post-WWII (1940–1951) periods, paralleling Figure 1 of the main paper. The convenience-inflation correlation is -0.34 in 1923–1939 and $+0.31$ in 1940–1951. For comparison, the correlation in Figure 1 for the Great Inflation period is 0.63, and for the 2000–2020 period is 0.02. The Great Depression hence exhibits the most negative correlation between convenience and inflation, while WWII is more similar to the Great Inflation with its endogenous convenience dynamics, featuring a positive convenience-inflation correlation.

Table A5 reports the five moments that we target in our simple estimation for the Great Depression and the WWII periods, along with the corresponding model moments. Because the gold standard was effective for a portion of the Great Depression, and the government controlled the long-term yield curve during WWII, there is no meaningful data counterpart to the nominal bond-stock beta used in our main model estimation. We therefore conduct a "no-beta" estimation. The corresponding data and model moments for the post-1952 periods from the main paper are repeated for convenience in the rightmost columns. We see that during the Great Depression and WWII, the T-bill convenience spread was similarly volatile as during 2000–2020. The volatilities of inflation and output were extraordinarily high during the pre-1952 periods. The correlation between inflation and the output gap was generally positive pre-1952, particularly during the Great Depression. Most important for our estimation, the regression slope of T-bill convenience onto inflation was negative during the Great Depression, but positive during WWII. The negative coefficient for the

Figure A5. T-bill convenience spread and inflation in the pre-1952 period. Each panel shows the standardized T-bill convenience spread (blue solid) and standardized quarterly inflation (red dashed). The within-panel convenience-inflation correlation is shown in each panel title (-0.34 for 1923–1939; +0.31 for 1940–1951). In each subperiod, we normalize both series to have zero mean and unit standard deviation. This figure parallels Figure 1 of the main paper.



Great Depression is consistent with the empirical results in the main paper. The result for WWII is new, but consistent with the idea that WWII presented a conclusion to the Great Depression.

The model is able to match the positive comovement between inflation and output in both pre-1952 subperiods, and the switch from a negative convenience-inflation coefficient during the Great Depression to a positive convenience-inflation coefficient during WWII. It does so by estimating an important role for demand shocks during both pre-1952 subperiods, as shown in Table A6. The volatility of liquidity demand shocks is estimated to be particularly significant during the Great Depression, both compared to the subsequent WWII period and the entire post-war period. This makes sense in that the Great Depression was a period of banking panics and particularly severe financial market disruptions, and is in line with our empirical estimates in Table 1 and the repeated pattern emphasized in Figure 1.

While the basic insight that liquidity demand shocks were particularly volatile during the Great Depression is useful, our very simple model must, however, necessarily remain silent on many features of the pre-1952 economy. Matching the extremely high output and inflation volatilities during the pre-1952 periods is a challenge. This may be because monetary policy worked differently at the time, but also because shocks transmitted differently through the economy, and many shocks that were important at the time, such as changes in macroeconomic risk, are abstracted away from in our model. This gap between the model and the data on macroeconomic volatility has repercussions, and the monetary policy shock volatility, σ_i, is extremely imprecisely estimated for the WWII period. Christiano et al. (2003) provide a richer model of the Great Depression, also finding that liquidity shocks mattered as macroeconomic drivers. In contrast, we show that the inflation-convenience comovement is a key identifying moment for endogenous vs. exogenous liquidity shocks and has changed repeatedly over time.

Table A5. Targeted moments in the model and the data pre–1952. This table extends Table 4 to pre-1952. For 1952–1999 and 2000–2020, moments and estimates are identical to those in Table 4 (full two-stage SMM). For the pre-1952 periods, estimation uses Stage 1 SMM matching the five moments shown in this table. The T-bill spread, inflation, and output gap are expressed in percentage points.

	1923–1939		**1940–1951**		**1952–1999**		**2000–2020**	
	Data	Model	Data	Model	Data	Model	Data	Model
Volatilities								
Vol(T-bill spread)	0.265	0.323	0.157	0.188	0.579	0.476	0.219	0.291
	(0.023)		(0.016)		(0.030)		(0.017)	
Vol(Inflation)	6.304	0.637	7.215	0.371	3.277	2.374	2.665	0.679
	(0.549)		(0.744)		(0.168)		(0.207)	
Vol(Output gap)	9.015	1.587	3.872	0.960	2.287	2.665	1.955	1.606
	(0.785)		(0.399)		(0.117)		(0.152)	
Correlation								
Corr(Inflation, Output gap)	0.236	0.256	0.158	0.183	-0.119	-0.312	0.292	0.427
	(0.118)		(0.145)		(0.072)		(0.102)	
Regression coefficient								
$\hat{b}$: T-bill spread $\sim b\cdot$ inflation	-0.013	-0.016	0.008	0.008	0.109	0.091	-0.003	-0.010
	(0.005)		(0.003)		(0.010)		(0.009)	

A.10 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

Table A6. Estimated shock volatilities pre-1952. This table extends Table 6 to pre-1952. For 1952–1999 and 2000–2020, estimates are identical to those in Table 6 (full two-stage SMM). Standard errors are reported in parentheses. All values are annualized percentages.

Parameter	1923–1939	1940–1951	1952–1999	2000–2020
σ_ℓ (liquidity demand)	0.125 (0.047)	0.081 (0.019)	0.100 (0.033)	0.079 (0.037)
σ_π (cost-push)	0.165 (0.090)	0.104 (0.052)	0.793 (0.119)	0.138 (0.088)
σ_i (monetary policy)	0.562 (0.832)	0.001 (27.513)	1.294 (0.230)	0.906 (0.203)
σ_x (non-liquidity demand)	0.660 (0.171)	0.471 (0.121)	0.050 (0.531)	0.534 (0.086)

ratio. 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.

Table A7 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.

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. Adding the debt-to-GDP control preserves the sign and significance of these coefficients, though the post-COVID estimates attenuate somewhat.

Table A7. 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–Tsy STRIP 10Y is the yield spread between ten-year Refcorp STRIPs and ten-year Treasury STRIPs (both 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. The stars indicate significance at * p <0.10, ** p <0.05, *** p <0.01 levels.

Panel A: Short-Term Convenience Yields

	(2018–2020)		(2021–2023)		(2018–2020)		(2021–2023)	
	ABCP-Tbill 3M				Refcorp-Tbill 3M			
E[inflation]	-0.058***	-0.090***	0.067***	0.028	-0.042***	-0.045***	0.12***	0.055*
	(-2.77)	(-5.30)	(3.83)	(1.60)	(-4.98)	(-4.90)	(3.65)	(1.99)
FFR	0.047	-0.17***	0.0067	-0.0068	-0.10***	-0.12*	0.074***	0.053**
	(1.19)	(-3.36)	(0.70)	(-0.61)	(-6.56)	(-1.98)	(4.17)	(2.71)
Debt/GDP		-4.03***		-2.96***		-0.38		-4.61***
		(-4.63)		(-3.42)		(-0.40)		(-3.05)
Constant	0.31***	3.37***	-0.10*	2.07***	0.76***	1.04	-0.033	3.35***
	(4.39)	(4.97)	(-1.77)	(3.48)	(24.17)	(1.40)	(-0.30)	(3.28)
$\bar{R}^2$	0.068	0.48	0.24	0.37	0.58	0.57	0.40	0.53
N	36	36	36	36	36	36	36	36

Panel B: Long-Term Convenience Yields

	(2018–2020)		(2021–2023)		(2018–2020)		(2021–2023)	
	Agency-Treasury STRIP 10Y				Refcorp-Treasury STRIP 10Y			
E[inflation]	-0.040***	-0.043***	0.023***	0.014***	-0.029***	-0.024***	0.057***	0.029**
	(-7.26)	(-6.97)	(9.63)	(3.71)	(-4.19)	(-3.45)	(4.92)	(2.40)
FFR	-0.028***	-0.047***	0.046***	0.043***	-0.055***	-0.022	0.045***	0.036***
	(-2.77)	(-2.82)	(13.69)	(15.31)	(-2.95)	(-0.94)	(4.69)	(4.92)
Debt/GDP		-0.35*		-0.68***		0.63**		-2.11***
		(-1.85)		(-3.90)		(2.39)		(-5.57)
Constant	0.56***	0.82***	0.18***	0.68***	0.63***	0.15	0.11***	1.66***
	(29.89)	(5.31)	(16.08)	(5.34)	(23.61)	(0.72)	(3.15)	(6.53)
$\bar{R}^2$	0.64	0.65	0.72	0.73	0.51	0.55	0.48	0.58
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.9 provides additional model impulse responses for supply and monetary policy shocks. Appendix B.10 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 $Y_{i,t}$:

$$Y_t = \left(\int_0^1 Y_{i,t}^{\frac{\epsilon_{p,t}-1}{\epsilon_{p,t}}} di \right)^{\frac{\epsilon_{p,t}}{\epsilon_{p,t}-1}}. \tag{A3}$$

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 $N_{i,t}$ labor used by firm i. 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

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

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, giving us a standard Phillips curve. $C_{h,t}^{home}$ denotes home consumption of the individual household and H_t^{home} denotes the corresponding external habit. Both market and home habits, H_t and H_t^{home}, are external and are taken as given by household h. Non-market or home consumption equals

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

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 equilibrium external habit over home consumption, 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

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

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

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

$R^b_{i,t}$, denotes the nominal return 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

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

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

Subtracting (11) from (10) and (12) from (10), and using $E_t[M_{t+1}^{\$}](1 + I_t^l) = 1$ from (10), gives

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

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

Dividing (A21) by (A22) gives $(I_t^l - I_t^b)/(I_t^l - I_t^d) = \lambda_t/(1 - \lambda_t)$. Substituting (7) into $I_t^l - I_t^d$ then gives (13) in the main text.

B.3.2 Steady-State Interest Rates

We log-linearize the model around the flexible-price steady-state values $\bar{c}$, $\bar{\pi}$, $\bar{i}^l$, $\bar{i}^b$, $\bar{i}^d$, and $\bar{\lambda}$ with deviations c_t, π_t, i_t^l, i_t^b, i_t^d, and $\hat{\lambda}_t$. The deviation $\hat{\lambda}_t$ is the level deviation from the steady state. We define the log steady-state interest rates by $\bar{i}^l = \log(1 + \bar{I}^l)$, $\bar{i}^b = \log(1 + \bar{I}^b)$, $\bar{i}^d = \log(1 + \bar{I}^d)$. 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

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

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

$$\bar{r}^b = \bar{i}^b - \bar{\pi}, \tag{A30}$$

$$= \log\left(1+\bar{I}^b\right) - \bar{\pi}. \tag{A31}$$

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

$$r_t^l = i_t^l - E_t \pi_{t+1}, \tag{A34}$$

$$r_t^b = i_t^b - E_t \pi_{t+1}. \tag{A35}$$

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 $v_{x,t} \equiv \gamma \psi \eta_t$, we get the exactly log-linear Euler equation with a non-liquidity demand shock

$$x_t = f^x E_t x_{t+1} + \rho^x x_{t-1} - \psi r_t^l + v_{x,t}, \tag{A37}$$

$$f^x = \frac{1}{\phi - \theta_1}, \ \rho^x = \frac{\theta_2}{\phi - \theta_1}, \ \psi = \frac{1}{\gamma(\phi - \theta_1)}, \tag{A38}$$

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

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

For notational convenience, define

$$C_\lambda \equiv 1 - \frac{\bar{\lambda}}{1-\bar{\lambda}}(1-\delta).$$

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

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

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

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

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

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

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

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

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 the convenience 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

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

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

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

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 includes 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. 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:

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

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

We hence want a matrix, A, such that

$$Var\left(\epsilon_{t+1}\right) = A\Sigma\Sigma_v\Sigma'A', \tag{A67}$$

$$= \begin{bmatrix} 1 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 1 \end{bmatrix}. \tag{A68}$$

and

$$A_1 \propto e_1. \tag{A69}$$

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:

$$A_2 = \frac{n_2}{\sqrt{n_2\Sigma\Sigma_v\Sigma' n_2'}}. \tag{A70}$$

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:

$$A_3 = \frac{n_3}{\sqrt{n_3\Sigma\Sigma_v\Sigma' n_3'}}. \tag{A71}$$

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:

$$A_4 = \frac{n_4}{\sqrt{n_4 \Sigma \Sigma_v \Sigma' n_4'}}. \tag{A72}$$

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 $\hat{v}_{t+1}$, 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:

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

This can be expressed in terms of the state variables:

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

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:

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

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

$$P^b_{2,t} = \exp\left(E_t\left(m_{t+1} - \psi^{-1}\xi_t - \hat{Z}_{3,t+1} + \hat{Z}_{2,t+2}\right) + \bar{i}^l - \bar{i}^b - \bar{r}^b\right) \times \mathbb{E}_t\left[\exp\left((-\gamma(\tilde{\lambda}(\hat{s}_t)+1)\Sigma_M - \underbrace{(e_3 - e_2 B)\Sigma}_{v_b})v_{t+1}\right)\right] \tag{A91}$$

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:

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

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:

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

The value function iteration for real liquid bond prices then becomes

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

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

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

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

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

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

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

This gives the following expression for f_n:

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

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:

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

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 difference to risk-neutral stock returns in any of our numerical applications.

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

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

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

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

$$b^{b,\$,rn}_2(\tilde{Z}_t, \hat{s}_t) = -e_3\left[I + B\right]A^{-1}\tilde{Z}_t + \frac{1}{2}v_{\$,b}\Sigma_v v_{\$,b}\prime - 2\bar{i}^b \tag{A106}$$

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:

$$P^{b,rn}_{2,t} = \exp\left(-\hat{Z}_{3,t} + \mathbb{E}_t\hat{Z}_{2,t+1} - \bar{r}^b\right) \times \exp\left(\mathbb{E}_t\left(-\hat{Z}_{3,t+1} + \mathbb{E}_{t+1}\hat{Z}_{2,t+2} - \bar{r}^b\right)\right)$$
$$\times\mathbb{E}_t\left[\exp\left(-\underbrace{(e_3 - e_2B)\Sigma}_{v_b}v_{t+1}\right)\right]. \tag{A107}$$

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

$$b^{b,rn}_2(\tilde{Z}_t, \hat{s}_t) = -(e_3 - e_2B)\left[I + B\right]A^{-1}\tilde{Z}_t + \frac{1}{2}v_b\Sigma_v v_b' - 2\bar{r}^b. \tag{A108}$$

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:

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

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

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}:

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

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}:

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

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 A8.

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 Christiano et al. (2005). The steady-state discount rate is implied by a real risk-free rate of $\bar{r}^l = 0.94\%$ 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). Combined with $\rho^d = 0.92$, the calibration target $\delta/(1-\rho^d) = 1/3$ implies a short-run pass-through of $\delta = 0.027$, consistent with the direct regression estimate (0.023) 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 A8. 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	Christiano et al. (2005)
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 use data moments to identify shock volatilities in each period. We estimate four volatilities, namely the liquidity demand shock volatility σ_ℓ, cost-push supply shock volatility σ_π, monetary policy shock volatility σ_i, and non-liquidity demand shock volatility σ_x.

Estimation Algorithm. For each period, we estimate the four volatility parameters by targeting six data moments from quarterly data: the volatilities of the convenience yield (T-bill spread), quarterly inflation, and the output gap, the correlation between quarterly inflation and output gap, the bond-stock beta, and the convenience spread-inflation regression coefficient. Denote the vector of these data moments by m^{data}. For each parameter vector σ, we simulate the model 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,

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

In Table 4 of the main text, we report the six target moment values for both subperiods, where all volatilities are annualized for the convenience of interpretation. We report the standard errors of each moment in brackets. The standard deviation of a volatility moment is computed as $\hat{\sigma}/\sqrt{2 \times (T^{data} - 1)}$, and the standard deviation of a correlation moment is computed as $(1 - \hat{\rho}^2)/\sqrt{T^{data} - 3}$. We use the squares of these standard errors as the diagonal elements of the weighting matrix W and, for simplicity, set non-diagonal elements to zero. The bond-stock beta standard error is heteroskedasticity-robust; the T-bill spread–inflation regression coefficient is reported with the OLS standard error.

Since the asset pricing solution involves nonlinearities, a global minimization over four parameters to match six data moments would be impractically slow. 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 weighting matrix W. This gives us a reasonable guess for a good starting point for the full optimization. Denote the estimated parameter set from the first step as σ_{step1}. Next, we create a grid around σ_{step1} and evaluate those grid points to optimize the full objective function in (A117) plus penalties for wrong signs, including the bond-stock beta, a large Sharpe ratio (above 1), and the regression coefficients of the convenience yield on inflation with and without controlling for the policy rate. To maximize the effectiveness of the grid, we use the Smolyak 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 6 of the main text.

We next provide robustness for the numerical accuracy of the asset pricing solution and the estimation result when using a denser numerical grid for the value function iteration. Appendix Tables A9–A10 verify that the parameter estimates are numerically identical, and the asset pricing moments quantitatively indistinguishable, across three numerical grids of increasing density and left-tail coverage for the log surplus consumption ratio. The asset pricing solution is highly non-linear in the log surplus consumption ratio, and previous literature has found that some common grids are problematic, so it is important to verify that our results have converged with respect to the solution grid for the log surplus consumption ratio. We show results for the following three grids:

- **Baseline.** Let $S_{grid,1}$ denote a vector of 20 equally spaced points between 0 and S_{max} with S_{max} included but 0 excluded, and let $s_{grid,2}$ denote a vector of 30 equally spaced points between -50 and $\min(\log(S_{grid,1}))$. The grid for s_t then consists of the concatenation of $s_{grid,2}$ and $\log(S_{grid,1})$. This baseline grid, used throughout the paper, is significantly denser than the grid used by Campbell and Cochrane (1999), which featured only 13 equally spaced points for S_t.
- **Extended.** This grid uses the same $S_{grid,1}$ as the baseline grid. However, $s_{grid,2}$ consists of 100 equally spaced points between -100 and $\min(\log(S_{grid,1}))$. This is a finer grid with a wider coverage of the left tail of log surplus consumption.
- **Wachter.** This grid uses 100 equally spaced points for $S_{grid,1}$. Further, $s_{grid,2}$ consists of 900 equally spaced points between -300 and $\min(\log(S_{grid,1}))$. This is the highest-density grid, and matches exactly the grid used by Wachter (2005).

Table A9 reports the model estimation for these three grids. We see that all four shock volatilities are identical to three decimal places across grid variants in both subperiods. All model asset pricing moments, and in particular the model coefficient of 1YR excess stock returns onto the lagged price-dividend ratio, are quantitatively indistinguishable from our baseline. Macroeconomic dynamics, such as impulse responses, variance decompositions, covariance decompositions, by construction are invariant to the numerical grid for surplus consumption.

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) 1000 independent 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. Because

Table A9. Parameter estimates across grid densities. Estimated shock volatilities (annualized %) for the two subperiods, across three model variants differing only in the S_t solution grid density. "Baseline" uses $\approx$50 grid points covering $\log S_t \in [-50, \log S_{\max}]$; "Extended" uses $\approx$120 points covering $\log S_t \in [-100, \log S_{\max}]$; "Wachter" uses $\approx$1,000 points covering $\log S_t \in [-300, \log S_{\max}]$, matching the Wachter (2005) benchmark.

Parameter	Baseline	Extended	Wachter
Panel A: 1952–1999			
σ_ℓ (liquidity demand)	0.100	0.100	0.100
σ_π (cost-push)	0.793	0.793	0.793
σ_i (monetary policy)	1.294	1.294	1.294
σ_x (non-liquidity demand)	0.050	0.050	0.050
Panel B: 2000–2020			
σ_ℓ (liquidity demand)	0.079	0.079	0.079
σ_π (cost-push)	0.138	0.138	0.138
σ_i (monetary policy)	0.906	0.906	0.906
σ_x (non-liquidity demand)	0.534	0.534	0.534

Table A10. Asset pricing moments across grid densities. Model-implied asset pricing moments across three model variants differing only in the S_t solution grid density. "Baseline" uses $\approx$50 grid points; "Extended" uses $\approx$120 points; "Wachter" uses $\approx$1,000 points, matching the Wachter (2005) benchmark. Panel A reports stock moments averaged across the two periods. Panel B reports nominal and risk-neutral bond-stock betas by period.

Moment	Data	Baseline	Extended	Wachter
Panel A: Stocks and Macro Moments				
Equity Premium	7.46	9.01	8.96	9.19
Equity Vol	17.23	16.72	16.61	17.03
Equity SR	0.44	0.54	0.54	0.54
AR(1) pd	0.94	0.95	0.95	0.95
1 YR Excess Ret on pd	-0.27	-0.35	-0.35	-0.36
1 YR Excess Ret on pd (R2)	0.15	0.05	0.05	0.05
Std. Annual Cons. Growth	1.54	2.26	2.26	2.26
Std. Annual Change FFR	1.83	2.28	2.28	2.28
Quarterly Std. Ten-Year Infl. Forecast	0.29	0.16	0.16	0.16
Panel B: Bond-Stock Betas by Period				
Nominal bond-stock beta (1952–1999)	0.22	0.16	0.16	0.16
Nominal bond-stock beta (2000–2020)	-0.29	-0.05	-0.05	-0.05
RN nominal bond-stock beta (1952–1999)	–	0.07	0.07	0.07
RN nominal bond-stock beta (2000–2020)	–	-0.01	-0.01	-0.01

each residual vector is itself computed on a length-T^{data} simulated sample, the sample covariance of these residuals already estimates the finite-sample covariance matrix of the data moments, which we denote S; no further division by T^{data} is needed. 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) to estimate the Jacobian. Finally, the asymptotic variance-covariance matrix of the estimated parameters is computed

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

where S denotes the estimated covariance matrix of moment residuals, D is the Jacobian matrix of moment residuals, and W is the weighting matrix. We report the square roots of the diagonal matrix V as the standard deviation of estimated parameters in parentheses in the second column of each period in Table 6.

Estimation without Bond-Stock Beta and Convenience-Inflation 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 A11 and the corresponding moment values in Table A12.

In Table A11, 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 A12, 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 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 A11,

Table A11. 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
	(0.033)	(0.097)	(0.044)	(0.037)	(0.065)	(2.483)
σ_π (cost-push)	0.793	0.793	0.699	0.138	0.170	0.228
	(0.119)	(0.102)	(0.085)	(0.088)	(0.106)	(0.049)
σ_i (monetary policy)	1.294	1.294	1.566	0.906	0.906	0.901
	(0.230)	(0.759)	(0.282)	(0.203)	(0.494)	(0.409)
σ_x (non-liquidity demand)	0.050	0.050	0.000	0.534	0.391	0.767
	(0.531)	(6.310)	(7.981)	(0.086)	(0.338)	(0.209)

Table A12. 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(T-bill 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	0.134	-0.049	-0.005	-0.119
T-bill spread $\sim b\cdot$ Inflation	0.091	0.091	0.068	-0.010	-0.002	0.079

for 1952–1999 the estimated liquidity-demand shock σ_ℓ becomes larger while the non-liquidity demand shock σ_x collapses to zero, with a correspondingly large standard error, so that the estimation is pushed to a corner of the parameter space once the asset-pricing moments are dropped. Even so, the implied moments in Table A12 still match the data in sign for this period: both the bond–stock beta and the convenience–inflation coefficient b remain positive, as they are in the data.

For 2000–2020, by contrast, excluding the convenience–inflation coefficient is consequential. The implied liquidity-demand shock is now estimated to be essentially zero, again with a large error band, while the non-liquidity demand shock becomes dominant. As a result, in the last column of Table A12 the model-implied convenience–inflation coefficient b flips to a positive value, opposite to its negative value in the data. The bond–stock beta, however, remains negative as in the data. It is therefore the convenience–inflation relationship, the central object of our analysis, that the model fails to reproduce once this moment is excluded.

Comparing the "no beta" and "no AP" cases, we conclude that the convenience–inflation regression coefficient is the decisive identifying moment. It disciplines the relative volatility of liquidity versus non-liquidity demand shocks and, in particular, is what allows the model to reproduce the post-2000 reversal of the convenience–inflation relationship to a negative value. The bond–stock beta, by contrast, is robustly pinned down even when it is not targeted, and therefore plays an over-identifying rather than an identifying role.

B.8 Model Fit: Autocorrelations

Table A13 reports autocorrelations in the model and in the data, with a model simulation length of 50,000. The model T-bill convenience spread is persistent, but not extremely persistent with an autocorrelation coefficient of 0.809 for the 1952-1999 period and 0.765 for the 2000-2020 period. The economic magnitudes of these AR(1) coefficients correspond to half-lives of 3.3 and 2.6 quarters, respectively. In the data, the autocorrelation of the convenience spread tends to be somewhat lower, but is often measured with substantial noise. In 1952–1999, the data AR(1) coefficient for the T-bill convenience spread is 0.663 with a standard error of 0.056, corresponding to a half-life of just under 2 quarters.

For 2000–2020, the T-bill convenience spread AR(1) coefficient in the data is only 0.395. However, this empirical persistence is measured quite imprecisely, with a 95% confidence interval ranging from 0.097 (half-life 0.3 quarters) to 0.693 (half-life 1.9 quarters). The point estimate for the empirical AR(1) coefficient and the large standard error during this period reflect the limited length of the subsample together with the extreme liquidity episode around COVID, when the convenience spread spiked from 0.31% to 1.30% between 2019Q4 and 2020Q1 and then fell back to 0.16% after Fed interventions in 2020Q2–2020Q3. The model has no Fed balance sheet channel and therefore cannot replicate this specific reversal. With these special factors in mind, the model convenience spread persistence for the 2000–2020 period at 0.765 (half-life 2.6 quarters) is higher than in the data, but not economically meaningfully different from the empirical 95% confidence interval. Overall, simulations show that the persistence of the model convenience spread is meaningfully lower than the deposit rate persistence, and not economically far from the convenience spread persistence in data.

The second row presents inflation persistence in the model and in the data. We report the AR(1) of annual inflation (the four-quarter sum of quarterly inflation), which is the empirically relevant horizon for persistence comparisons and filters out high-frequency noise from quarter-to-quarter price index changes due to variation in highly transitory inflation components. The model matches empirical inflation persistence well for the 1952–1999 period. There is a gap between model and data inflation persistence for 2000–2020, but the gap is much reduced for core inflation dynamics during this period. Over the same period, core CPI inflation, which strips out the volatile food and energy components, has a much higher persistence than headline, with an AR(1) coefficient of 0.834 in 2000–2020. Finally, the output gap AR(1) coefficient is closely matched in both periods (0.916 vs. 0.954 in 1952–1999; 0.771 vs. 0.788 in 2000–2020).

Table A13. Autocorrelations in the model and data. For the data column, each entry is the quarterly AR(1) coefficient estimated over the indicated subsample. The `s.e.` entries (in parentheses) are data Newey–West standard errors with 12 lags. Model autocorrelations are computed analogously using model-simulated data of length 50,000 at the model parameter values for each subperiod. The autocorrelation moments listed in this table are not targeted in the model estimation.

		1952–1999			**2000–2020**		
		Data	s.e.	Model	Data	s.e.	Model
	T-bill spread	0.663	(0.056)	0.809	0.395	(0.152)	0.765
AR(1)	Inflation (annual)	0.973	(0.029)	0.948	0.672	(0.077)	0.966
	Output gap	0.916	(0.026)	0.954	0.771	(0.149)	0.788

B.9 Impulse Responses to Monetary Policy and Non-Liquidity Demand Shocks

Figure A6 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 A7 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 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.

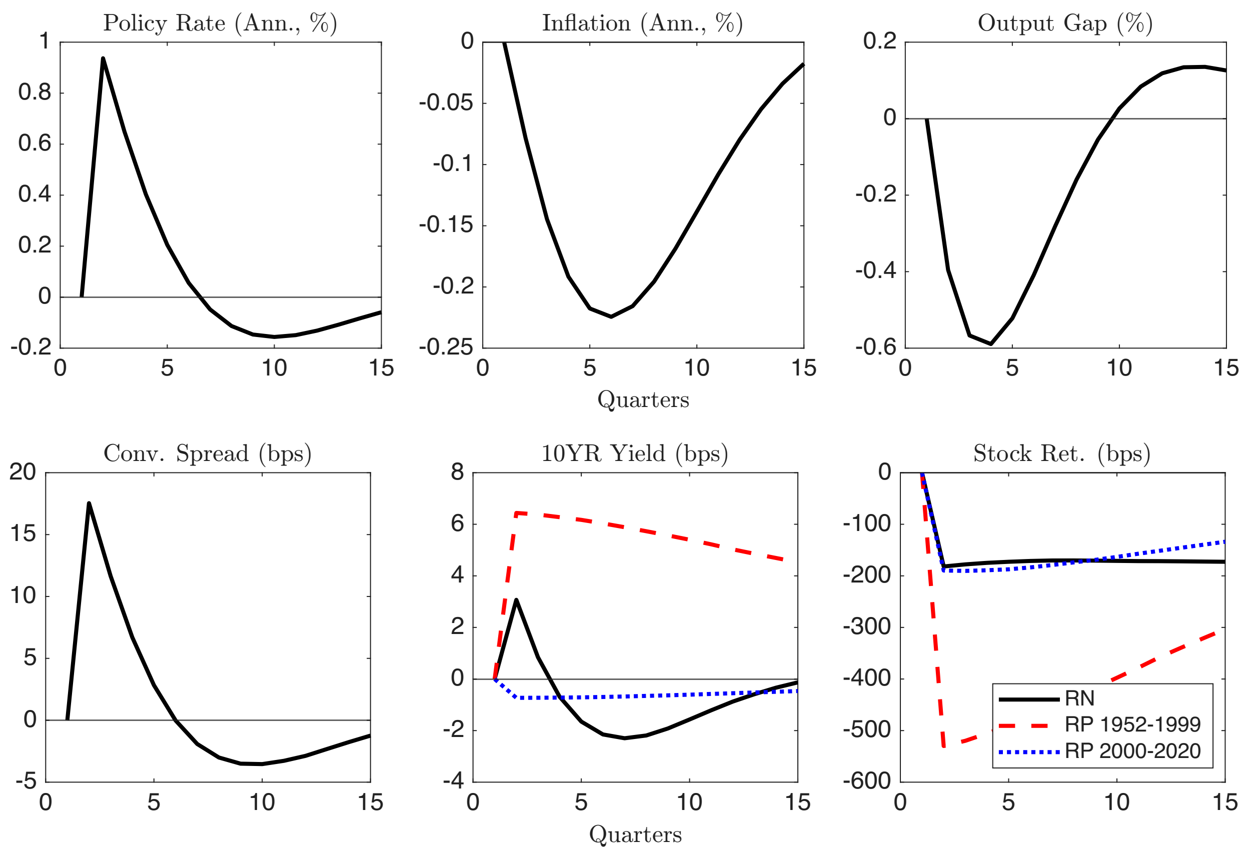
B.10 Model Extensions

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

B.10.1 Imperfect Substitutability

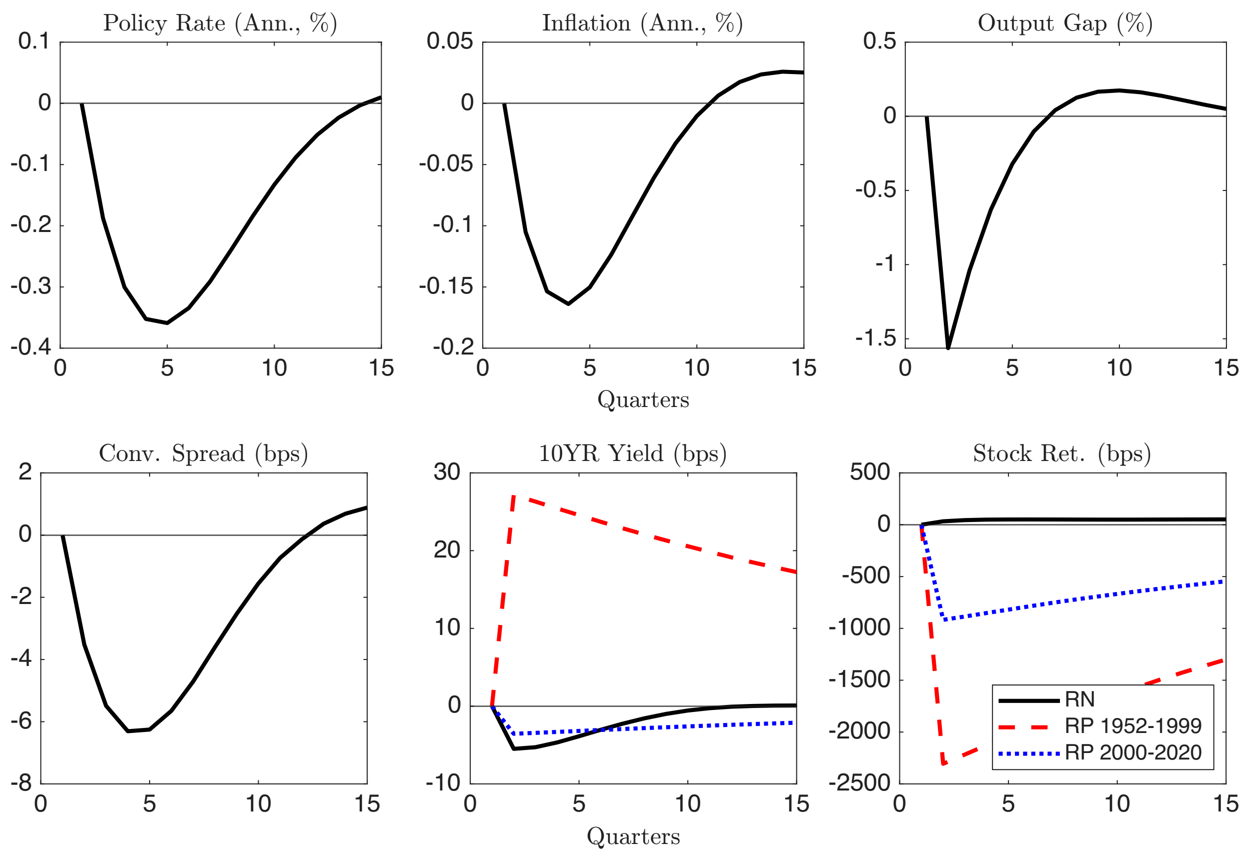
While our baseline model treats Treasuries and deposits as perfect substitutes, this assumption is made merely for simplicity. To see how the framework generalizes, assume that the liquidity aggregate is given as in Nagel (2016) and Krishnamurthy and Li (2023),

Figure A6. 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.



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

Figure A7. 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.



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

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

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.10.2 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

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

Combining equations (10) and (12) gives

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

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 by assumption α_t is not allowed to enter into the deposit spread in equilibrium. Substituting

(A121) into (A122) then gives equation (13) and the Treasury convenience yield is not affected by α_t. Thus, changes in α_t cannot be regarded as a shock to the convenience spread, as long as assumption (7) holds. However, deviations from (7) would act similarly to a shock to λ_t.