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.

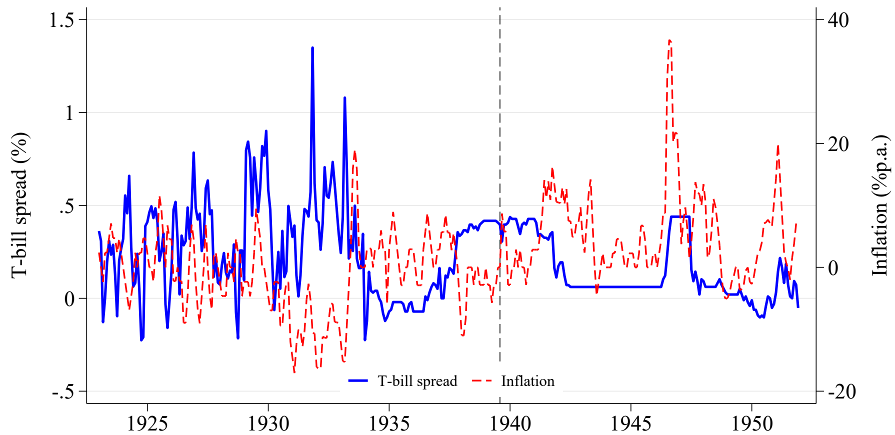
Table A1 reports the summary statistics for main variables used in empirical analysis.

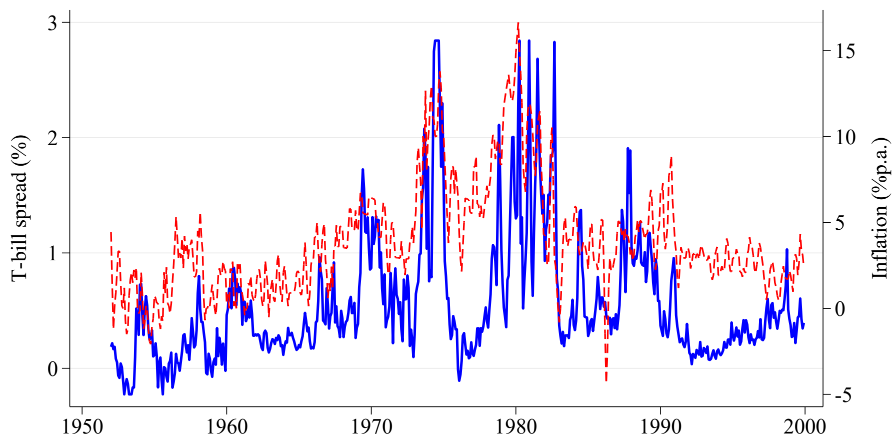
Table A1. Summary statistics. This table presents summary statistics for our full sample 1923:01–2020:12, excluding the 1939:09–1951:12 period.

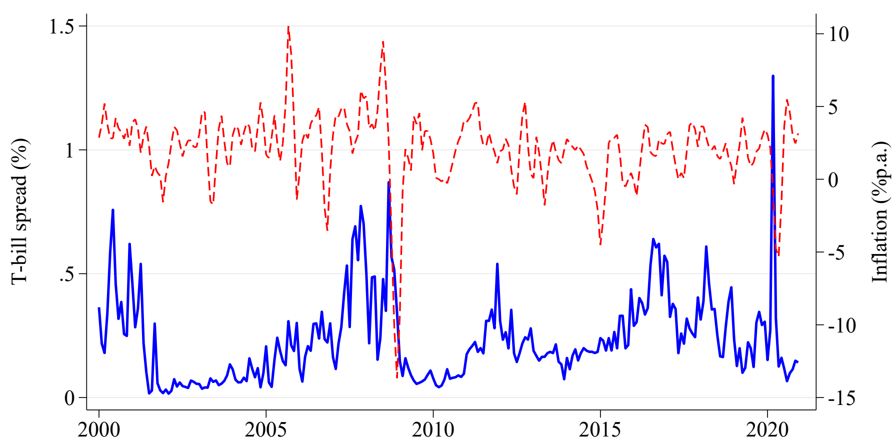
Variable	Obs	Mean	Std. Dev.	Mean		
	(1923-2020)			(1923-1939)	(1952-1999)	(2000-2020)
Tbill spread (%)	1028	0.42	0.46	0.27	0.56	0.23
Inflation (%)	1028	2.48	4.33	-1.13	3.90	2.10
VIX (%)	1028	19.7	8.45	25.9	17.4	19.9
Baa-Aaa spread (%)	1028	1.17	0.70	1.99	0.93	1.05
Debt/GDP	1028	0.29	0.15	0.17	0.26	0.47

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

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







A.1 Data sources

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

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

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

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

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

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

A.2 Credit risk in convenience yield measures

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

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

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

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

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

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

A.3 Controlling period-specific credit risk

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

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

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

A.4 Robustness with different inflation measures

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

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

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

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

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

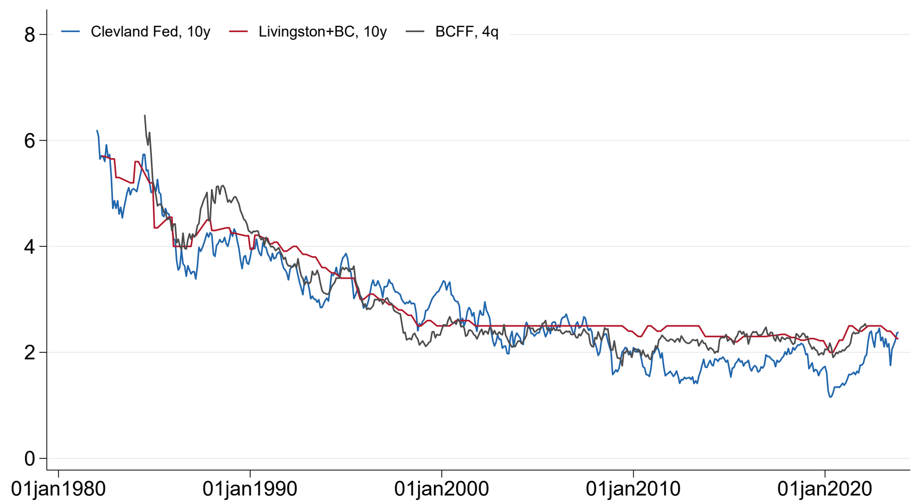
A.5 The Cleveland Fed index, inflation expectations, and term premia

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

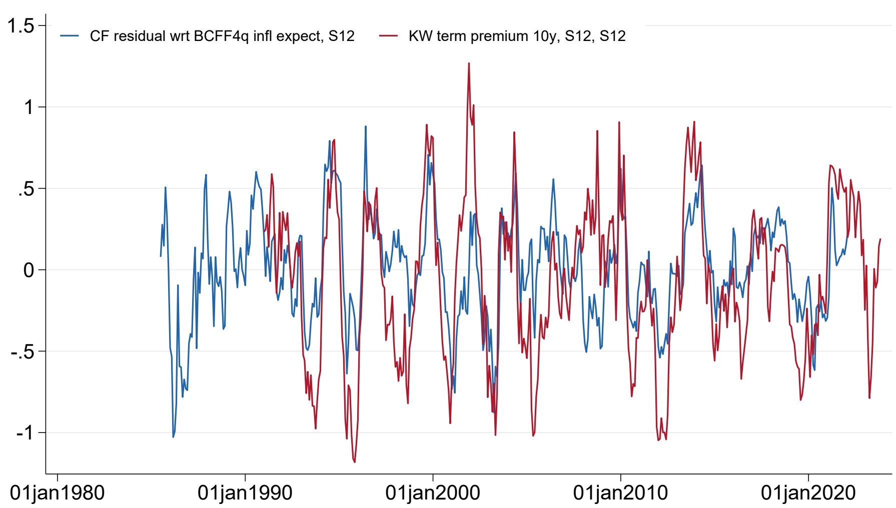
Panel A of Figure A2 shows 10-year inflation expectations from the Cleveland Fed against long-term and 4-quarter consensus inflation expectations from surveys. It is clear that the Cleveland Fed inflation expectations roughly move similarly at lower frequencies, but are substantially more volatile, raising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

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

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



(a) Cleveland Fed inflation expectations vs. surveys



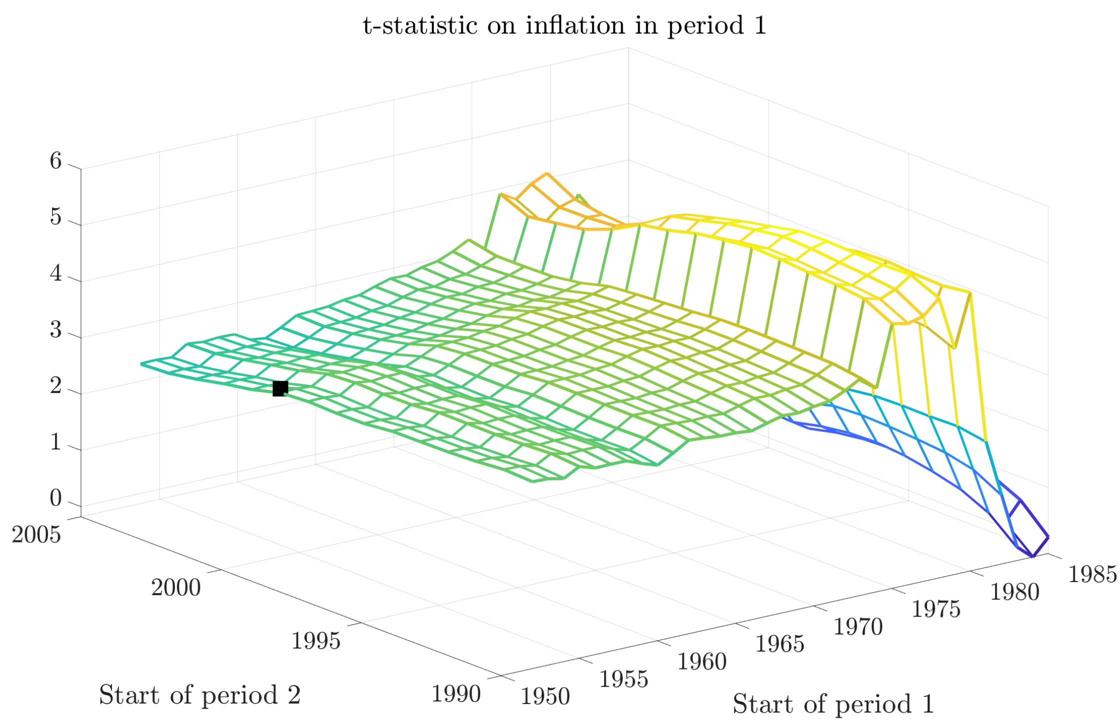
(b) Cleveland Fed inflation expectations vs. term premia

A.6 Robustness for different break dates

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

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

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



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

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

A.7 Deposit rate process

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

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

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

A.8 Long-term convenience spread adjustments

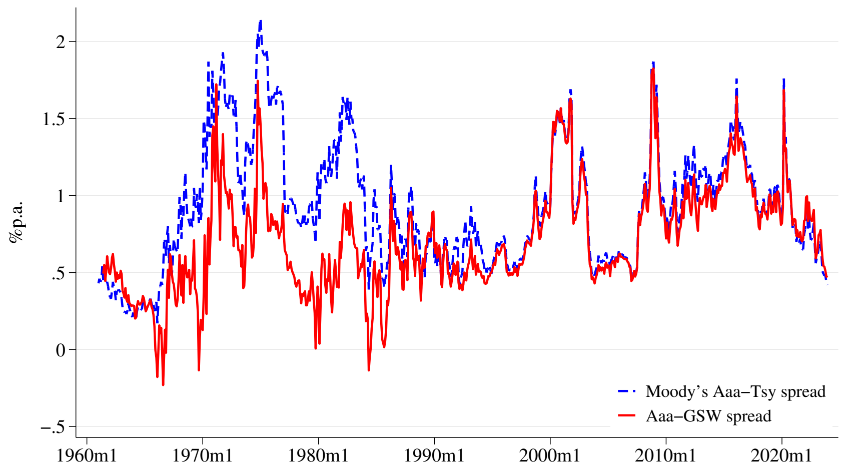
While in our main analysis we primarily focus on the T-bill spread, reflecting short-term convenience, it is worth discussing the long-term convenience spread as well. Much of the literature uses the Moody's Aaa-Treasury spread as a measure of long-term convenience, in the absence of alternatives available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

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

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

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

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



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

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

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

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

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

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

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

---

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

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

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

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

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

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

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

A.9 Post-COVID period inflation and convenience

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

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

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

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

Panel A: Short-Term Convenience Yields

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

Panel B: Long-Term Convenience Yields

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

B Model Appendix

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

B.1 Model setup

B.1.1 Final good production

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

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

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

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

The aggregate price level is given by

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

B.1.2 Intermediate good producers

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

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

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

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

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}:

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

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

B.1.3 Price setting

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

B.2 Household preferences and SDF

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

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

where per-period utility equals

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

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

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

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

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

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

B.2.1 Habit dynamics

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

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

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

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

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

B.2.2 Household and government budget constraints

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

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

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

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

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

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

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

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

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

B.3 Model Solution

B.3.1 Liquidity spread

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

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

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

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

B.3.2 Steady-state interest rates

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

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

B.3.3 Output gap and Phillips curve

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

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

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

B.3.4 Consumption Euler equation

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

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

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

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

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

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 of $v_{x,t}$ in (22), 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:

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

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

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

This leads to

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

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

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

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

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

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

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

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

Similarly, we can also rewrite f^d,

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

For simplicity, we rewrite equation (A40) as

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

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

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

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

which leads to

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

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

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

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

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

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

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

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

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

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

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

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

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

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

which implies

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

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

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

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

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

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

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

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

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

B.4 Macroeconomic equilibrium

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

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

In matrix form, the model can be written as

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

where the matrices are given by

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

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

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

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

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

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

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

B.5 Solving for asset prices

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

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

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

We hence want a matrix, A, such that

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

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

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

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

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

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

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

The steady-state surplus consumption sensitivity equals:

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

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

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

This can be expressed in terms of the state variables:

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

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

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

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

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

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

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

B.5.1 Recursion for zero-coupon liquid bond prices

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

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

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

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

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

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

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

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

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

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

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

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

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

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

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

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

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

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

The value function iteration for real liquid bond prices then becomes

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

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

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

B.5.2 Consumption claim recursions

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

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

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

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

where we define

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

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

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

We simplify

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

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

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

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

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

This gives the following expression for f_n:

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

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

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

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

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

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

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

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

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

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

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

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

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

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

$$b_2^{b,rn}(\tilde{Z}_t, \hat{s}_t) = -(e_3 - e_2B)\left[I + B\right]A^{-1}\tilde{Z}_t + \frac{1}{2}v_b\Sigma_v v_b' - 2\bar{r}^b. \tag{A108}$$

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:

$$B_n^{b,\$,rn}(\tilde{Z}_t, \hat{s}_t) = \mathbb{E}_t\left[\exp\left(-\bar{i}^b - e_3A^{-1}\tilde{Z}_t - e_2A^{-1}\epsilon_{t+1} + b_{n-1}^{b,\$,rn}(\tilde{Z}_{t+1}, \hat{s}_{t+1})\right)\right], \tag{A109}$$

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

B.5.4 Risk-neutral zero-coupon consumption claims

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

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

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

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

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

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

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

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

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

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

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

B.6 Invariant model parameter values

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

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

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

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

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

B.7 Details on model estimation and standard errors

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

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

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

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

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

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

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

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

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

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

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

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

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

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

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

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

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

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

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

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

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

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

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

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

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

B.9 Model extensions

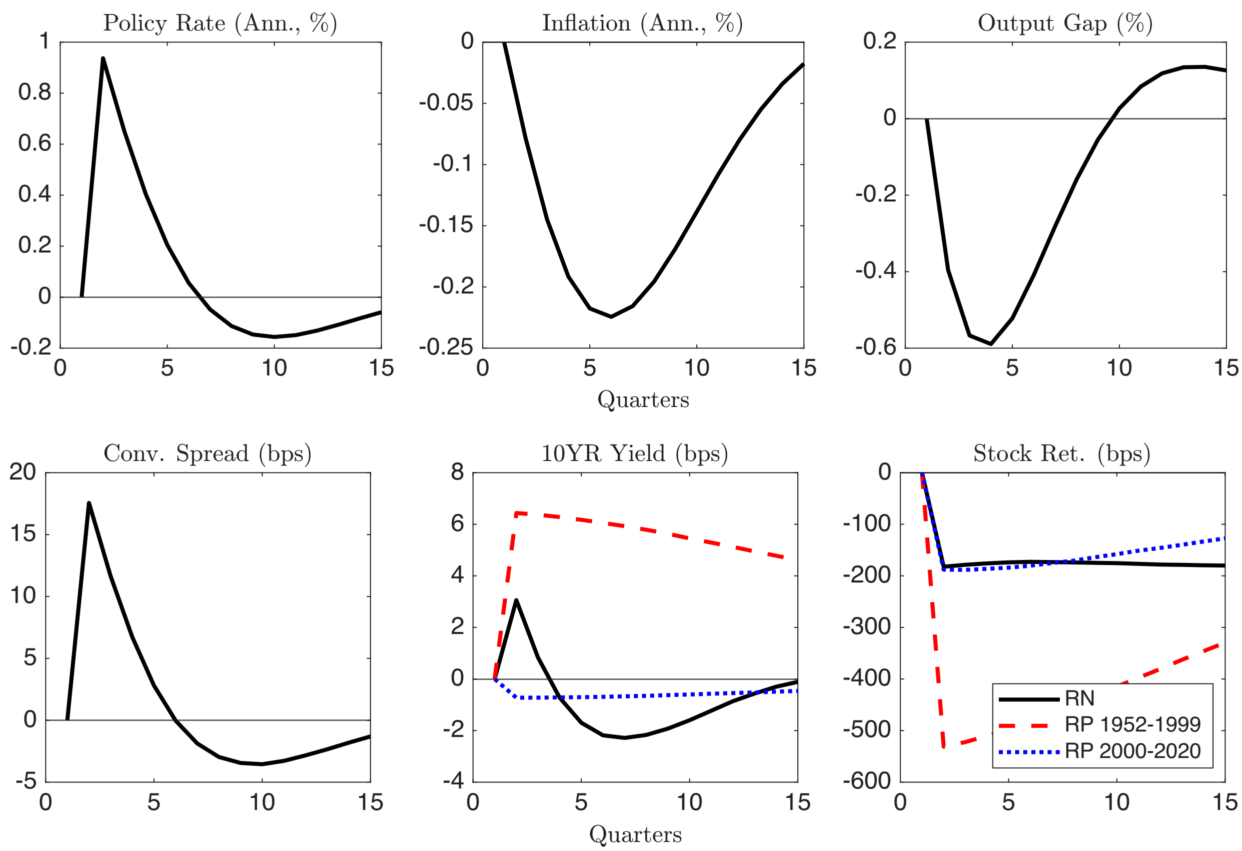
For simplicity, we simplify to the case with no deposit rate inertia, i.e. $\rho^d = 0$, throughout this subsection.

B.9.1 Imperfect substitutability

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

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

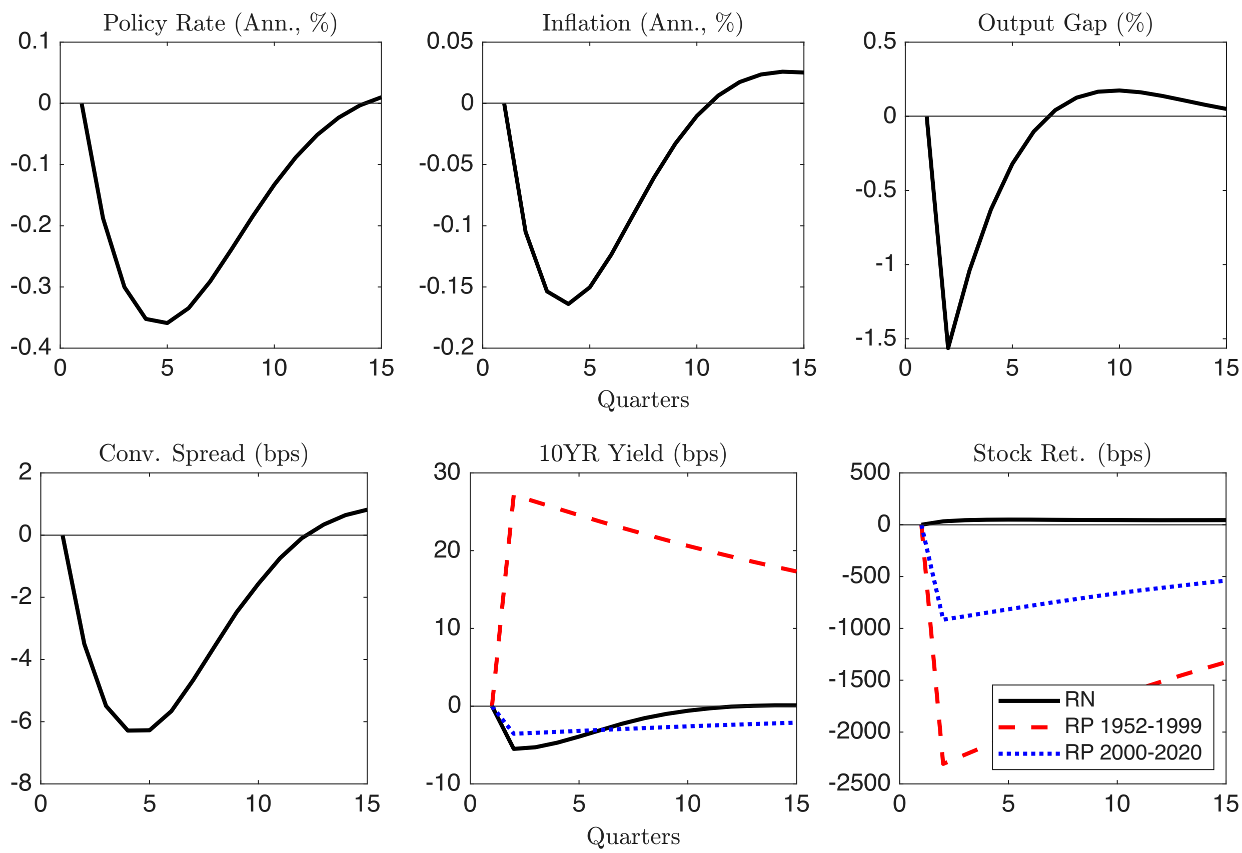
Figure A5. Baseline model responses to a monetary policy shock. This figure shows impulse responses to a 1 percentage point increase in the monetary policy shock $v_{i,t}$, where the impulse period is period 2. The driving shock has mean zero and is drawn from the steady-state distribution in all other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



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

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

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



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

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

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

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

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

B.9.2 Shocks to overall liquidity demand

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

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

Combining equations (10) and (12) gives

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

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

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