Inflation and Treasury Convenience

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Abstract

Using a century of data, we show that Treasury convenience yield and inflation comove positively during the inflationary second half of the 20th century, but negatively pre-WWII and post-2000. An inflation decomposition reveals that higher supply inflation predicts higher convenience, while lower demand inflation follows higher Treasury convenience. In our model, inflationary cost-push shocks raise the opportunity cost of holding money and money-like assets, inducing higher convenience. Conversely, liquidity demand shocks drive up convenience but lower consumption demand and inflation in the model. An estimated version of our model matches the changing inflation-convenience relationship in the data and shows that liquidity demand shocks dominated pre-WWII and post-2000, but supply shocks prevailed in the second half of the 20th century.

Keywords: Treasury convenience; inflation; demand shocks; money view;

JEL classification: E44, E58, G01, G28

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1 Introduction

How is Treasury convenience linked to inflation? The relationship between liquidity, interest rates, and inflation was central to the macroeconomic debates of the 20th century (Keynes (1937), Friedman (1969)). Today, it is again relevant due to renewed concerns about inflation and the status of the U.S. Treasuries. Recent research on Treasury markets indicates that investors value U.S. Treasury securities more highly than assets with the same cash flows, meaning that Treasury bonds have convenience value (Longstaff, 2004; Du et al., 2018b; Krishnamurthy and Vissing-Jorgensen, 2012). Besides serving as a significant source of fiscal capacity, Treasury convenience affects monetary policy transmission (Drechsler et al., 2018; Jiang et al., 2023), drives business cycle dynamics during global financial crises (Del Negro et al., 2017; Li, 2024), and is a critical component of dollar valuation and exchange-rate dynamics (Jiang et al., 2021; Du et al., 2018a).

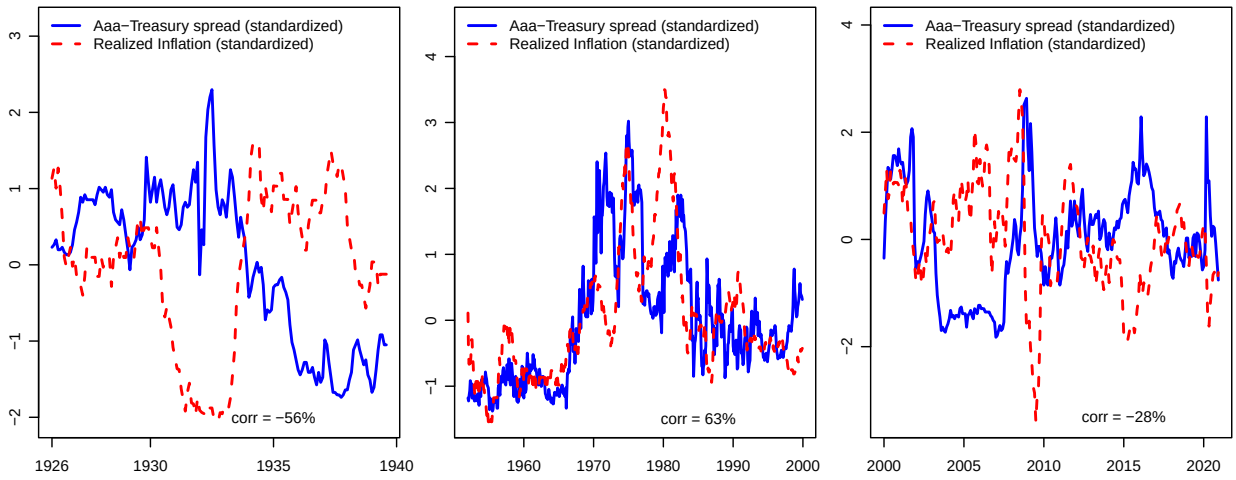
Using a century of U.S. data, we document secular shifts in the relationship between Treasury convenience and inflation, and link them to the macroeconomic drivers of inflation. Figure 1 illustrates a striking fact: Convenience comoves positively with inflation during the Great Inflation episode, but not before or after.[1] In periods typically associated with supply-side shocks, such as the 1970s, higher inflation tends to go along with higher – not lower – Treasury convenience. Conversely, in periods with preeminent demand-side shocks, such as financial crises, including the pre-WWII period and the pre-pandemic 2000s, lower inflation tends to coincide with higher Treasury convenience.

We explain these empirical findings with the changing dominance of two channels: the "money channel" and the "liquidity demand channel." Combining a standard New Keynesian model with liquidity demand for Treasuries, the money channel generates a positive inflation-convenience relationship in response to inflationary cost-push supply shocks, as in the 1970s. According to this channel, higher inflation expectations increase the cost of holding money and close money substitutes such as Treasuries, and raise the Treasury convenience premium that investors pay in equilibrium.[2] On the other hand, a positive shock to the demand for liquid Treasuries lowers aggre-

[1] We follow Krishnamurthy and Vissing-Jorgensen (2012) in measuring the convenience value of Treasury bonds with the spread between Aaa-rated corporate bond yields and long-term Treasury bond yields, consistent with the fact that Aaa-rated corporate bonds have never defaulted during our historical period. A higher Aaa-Treasury spread corresponds to a higher value of Treasury convenience.

[2] This channel encapsulates the monetarist view that the nominal rate of interest is the cost of holding money (Cagan, 1958; Mundell, 1963; Tobin, 1969; Friedman, 1969), augmented with the view that Treasury bonds of all maturities have some money-like qualities (Friedman and Schwartz, 1982; Holmström and Tirole, 1998; Longstaff,

Figure 1. Convenience yield and inflation. We plot the Treasury convenience, measured as Aaa-Treasury spread (Krishnamurthy and Vissing-Jorgensen, 2012), and the 12-month inflation during three subperiods: 1926:01-1939:08, 1952:01-1999:12, and 2000:01-2020:12. We exclude the 1939:09-1951:12 period covering WWII until lifting of interest rate controls. The data frequency is monthly. In each subperiod, we normalize both series to have zero mean and unit standard deviation.



gate macroeconomic demand, leading to demand-driven disinflation and a negative convenience-inflation relationship, as during the pre-WWII and post-2000 periods.[3] Intuitively, liquidity demand shocks tend to coincide with banking crises (Diamond and Van Tassel, 2023) and thus may reflect a higher value of public liquidity relative to private liquidity (Holmström and Tirole, 1998) as well as concerns about tail risks (Caballero and Krishnamurthy, 2008).

Our baseline empirical result shows that a one percentage point increase in 12-month headline CPI inflation is associated with a long-term convenience yield that is 13 bps higher in the second half of the 20th century (1952-1999) compared to the pre-WWII period, a magnitude that is large relative to an average convenience spread of 87 bps. We split the sample at the start of WWII and also in 2000, when inflation dynamics are broadly understood to have changed from supply- to demand-driven (e.g., Justiniano and Primiceri, 2008; Stock and Watson, 2007; Bekaert et al., 2021; Pflueger, 2025). In contrast, we show that the coefficients on inflation in the pre-WWII and post-2000 periods are generally negative and not statistically significantly distinguishable from each other. These relationships between inflation and long-term convenience hold while controlling for

2004; Krishnamurthy and Vissing-Jorgensen, 2012; Nagel, 2016; Krishnamurthy and Li, 2023).

[3]Keynes (1937) considered exogenous variation in liquidity preference as the key determinant of interest rates, business cycles, and inflation.

the federal funds rate, the quantity of Treasury debt as measured by the debt/GDP ratio, equity volatility, and the credit spread between Baa and Aaa corporate bond yields. Controlling for the federal funds rate, in particular, ensures that our results do not simply reflect monetary policy surprises or the monetary policy response to non-liquidity demand shocks. Results for short-term T-bill convenience yield, measured by the Repo-T-bill spread, display similar shifts, though the level of the federal funds rate now captures a bigger share of the short-term convenience variation in line with Nagel (2016).

Evidence from inflation components further supports the existence of a money channel vs. liquidity demand channel, whose relevance has changed over time. We first show that the long-term Treasury convenience yield exhibits a stable positive relationship with core inflation, which is often used as a simple and robust measure of broad-based inflation. Further, we document stable positive relationships during the second half of the 20th century between long-term convenience and survey inflation expectations as well as trend inflation estimates following Stock and Watson (2007). The correlation between energy inflation and convenience is weaker. Overall, the money channel appears to manifest in our data when inflationary shocks are longer-lived and pass onto broad measures of core inflation.

We next show that supply- and demand-driven inflation have opposite relationships with Treasury convenience, both contemporaneously and at leads and lags. We use Shapiro (2024)'s decomposition of core PCE inflation into demand and supply components, which is based on disaggregated product-level price and quantity data and available from 1969 onwards. We find that a higher long-term Treasury convenience yield is associated with higher supply-driven inflation, but lower demand-driven inflation. Lead-lag relationships further reveal that higher supply-driven inflation predicts a higher Treasury convenience yield, while a higher Treasury convenience yield predicts lower demand-driven inflation. The decomposition into supply- and demand-driven inflation thus supports the interpretation that supply shocks tend to raise inflation and the Treasury convenience yield, whereas shocks to Treasury liquidity demand tend to raise convenience but lower inflation.

To further explore how the relationship between inflation and convenience changes with the nature of macroeconomic shocks, we break our century of data into subsamples using a financial markets indicator of supply-driven inflation—the bond-stock betas (Campbell et al., 2020; Pflueger, 2025). We document that the inflation-convenience relationship monotonically increases with the betas. Intuitively, negative supply (cost-push) shocks tend to drive up inflation and lower Treasury bond valuations just as stocks fall due to the impending recession, inducing a positive

correlation between bonds and stocks. We find that a 1 percentage point increase in 12-month CPI inflation is associated with a 6 bps decline in long-term Treasury convenience when bond-stock betas are roughly within their bottom quintile, whereas they are associated with a 5 bps *increase* when bond-stock betas are within their top quintile. At the same time, our baseline results are robust to controlling for bond-stock betas directly, which might matter if bond-stock betas have an effect on the hedging value of the Treasuries (Acharya and Laarits, 2023). Hence, bond-stock betas further support the idea that supply-driven inflation contributes to a positive inflation-Treasury convenience relationship.

We explain the empirical findings by combining a parsimonious three-equation New Keynesian model as in Galí (2008), Rotemberg and Woodford (1997), or Clarida et al. (1999), with a standard model of convenience yield following Nagel (2016).[4] We model agents as deriving utility from a liquidity aggregate where Treasuries are substitutes with deposits or non-interest-bearing money, similarly to money in the utility function (Sidrauski, 1967). In the model, supply-type cost-push shocks lead to persistently higher inflation, raising the opportunity cost of holding money and, thus, also the opportunity cost of holding near-money assets, including Treasuries. Consequently, higher inflation leads to higher convenience yield via the money channel. We show in a simple quantification that the money channel is particularly prominent if the economy experiences supply cost-push shocks and inflation is persistent, as in the 1970s and 1980s.

The liquidity demand channel is modeled via a shock to the liquidity value of Treasuries. Convenience yield introduces a wedge in the household Euler equation between the risk-free discount rate and the Treasury yield. A positive shock to convenience yield increases households' borrowing rate relative to the policy rate and suppresses household aggregate demand. This, in turn, decreases inflation and induces a negative inflation-convenience comovement.[5] The liquidity demand channel can therefore explain the negative inflation-convenience relationship during the pre-WWII and post-2000 period, when liquidity shocks and financial crises were prevalent. This model predic-

[4]For models of banking and money within a New Keynesian economy see also Curdia and Woodford (2010), Gertler and Karadi (2011a), Drechsler et al. (2018), Piazzesi et al. (2019) and Wang (2022). Caballero and Simsek (2020) and Caballero and Simsek (2022b) develop models of optimal monetary policy when broad asset prices matter for aggregate fluctuations, and both financial and non-financial demand shocks are present. Our focus is different, in that we seek to build the most basic model of the economy and Treasury convenience that can replicate the changing inflation-convenience relationship that we document in the data.

[5]Shocks to the convenience of Treasury bonds have been increasingly used to explain a wider range of empirical facts (Anzoategui et al. (2019), Jiang et al. (2023), Itskhoki and Mukhin (2021), Kekre and Lenel (2021), Fukui et al. (2023), Bianchi et al. (2022b), Engel and Wu (2023), Abadi et al. (2023).

tion also speaks to our empirical results based on decomposing inflation into supply and demand components and the lead-lag relationships between those components and convenience. One implication is that a central bank may find it optimal to respond to liquidity demand shocks in financial markets, just like it would to other demand shocks driving inflation and the output gap.

The model implies that the comovement between inflation and convenience yields is a highly informative moment for identifying the nature of demand shocks driving economic fluctuations. In particular, it helps disentangle the role of liquidity demand from other sources of macroeconomic demand shocks. We estimate the volatilities of structural shocks in the model to match the observed volatilities of convenience spreads, inflation, and the output gap as well as their correlations across three distinct historical episodes. The results match the changing inflation-convenience relationship in the data. The estimated model highlights the central role of liquidity demand shocks during the Great Depression and the post-2000 period, and their relative attenuation during the high-inflation era of the 1970s and 1980s, when cost-push shocks were the dominant force.

The model also shows that a direct negative effect of inflation on Treasury convenience cannot explain our empirical findings. A direct negative effect of inflation on convenience could arise via several mechanisms: (a) Higher inflation means that the monetary authority has less capacity to substitute debt with money (Nagel, 2016; Krishnamurthy and Li, 2023). (b) When inflation is more volatile, Treasuries have less stable valuation (Krishnamurthy and Vissing-Jorgensen, 2012; Di Tella, 2020; Brunnermeier et al., 2022b; Liu et al., 2021). (c) Higher inflation increases the cost of financial intermediaries trading Treasury securities.[6] Contrary to the data, this assumption implies a more negative convenience-inflation relationship in high-inflation periods, such as the second half of the 20th century, than during low-inflation periods, such as the pre-WWII and post-2000 periods. Intuitively, with a direct inflation-liquidity link, Treasuries command little convenience during high inflation times, weakening the money channel and resulting in a less positive or even more negative Treasury convenience-inflation relationship during high inflation. Our empirical findings, therefore, challenge the notion that inflation directly depresses Treasury convenience, at least at historically experienced inflation rates in the U.S.

Finally, we use data on convenience yields and inflation expectations for the post-pandemic period to validate the model predictions out-of-sample. Consistent with the demand shock inter-

[6]See Duffie et al. (2007) for a general theory of how intermediation frictions affect asset prices. Jermann (2020) models how limits to arbitrage affect the Treasury-swap spread. Du et al. (2023) provide both theory and empirics on how intermediation costs affect Treasury pricing. d'Avernas et al. (2024) incorporates both bank regulation and central bank in a unified model of the Treasury market.

pretation of the COVID pandemic in the first half of 2020, we show that the initial rise of long-term convenience yields coincides with disinflation. However, the negative convenience-inflation relationship disappears and turns positive once the initial pandemic shock resolves. The reemergence of the money channel in the face of supply shocks can explain a positive shift in the convenience-inflation comovement post-2020, which robustly appears across different measures of convenience, such as agency-Treasury spreads and different measures of inflation expectations.

Measuring Treasury convenience—particularly at long maturities—faces significant data challenges in the earlier part of our sample, due to the scarcity of long-term Treasury bonds and the presence of option-like features in both long-term Treasuries and corporate bonds. We propose several adjustments to the original Moody's Aaa spread to account for these confounding factors. We find that the positive comovement between inflation and long-term convenience remains a robust feature across much of the high-inflation period, when cost-push shocks dominated economic fluctuations, with a potential exception for the monetarist experiment period. Although the unprecedented monetary policy actions and extreme interest rate volatility of 1979 and 1980 weaken the relationship between our proxies for the long-term convenience spread, the T-bill spread indicates that Treasury bonds continued to provide meaningful convenience benefits during that period.

Our work contributes to the growing literature that studies the determinants and effects of Treasury convenience, by examining the link to the macroeconomy and inflation. Nagel (2016) shows that the U.S. monetary policy drives the Treasury convenience yield by changing the opportunity cost of holding money and money-like assets. Diamond and Van Tassel (2023) provide international evidence and link convenience demand shocks to financial crises. Du et al. (2018b) and Jiang et al. (2021) document that violations of the covered interest parity in foreign exchange markets are correlated with international perceptions of the U.S. Treasury convenience. Binsbergen et al. (2022) construct stock-option implied risk-free rates and find that monetary policy affects the convenience yield. Hébert et al. (2023) provide complementary evidence that the gap between the stock market-implied risk-free rate and government rates acts as a shifter in the Euler equation akin to a demand shock. Li (2024) presents the convenience yield as a channel of how quantitative easing policies affect the banking sector and financial crises. Hartley and Jermann (2024) explain the pricing of U.S. Treasury floating rate notes through a model where money-like assets, including Treasuries, differ in their degrees of moneyness.

Complementary to our work, Acharya and Laarits (2023) show that the hedging properties of Treasury bonds, measured by the stock-bond beta, contribute to explaining variation in the

convenience yield. Fu et al. (2023) argue for a negative correlation between Treasury convenience and inflation expectations. Focusing on the post-1982 sample, they base their evidence on expected inflation estimates from the Cleveland Fed model. These estimates, being derived from yields, can confound inflation expectations and convenience. We instead rely on direct measures of inflation and inflation expectations from surveys, not derived from yields, and find a positive convenience-inflation relationship during high-inflation episodes.

The remainder of the paper is structured as follows. Section 2 discusses the data and measurement. Section 3 describes our empirical results. Section 4 presents the model and results from a simple quantification. Section 5 describes the results for the recent post-COVID period. Finally, Section 6 concludes.

2 Data and Measurement

The literature has proposed two main measures of the Treasury convenience yield that are available for our full sample back to the 1920s. Our primary proxy for the long-term convenience yield is the Aaa corporate bond yield minus a maturity-matched Treasury bond yield, as in Krishnamurthy and Vissing-Jorgensen (2012).[7] Second, we construct the short-term convenience yield following Nagel (2016) 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. Since the repo data used by Nagel (2016) ends in 2011, we rely on the 3-month commercial paper rates to supplement the most recent period.[8] We refer to the concatenated series as the T-bill convenience yield.

Our baseline inflation measure is the 12-month backward-looking 12-month change in the headline consumer price index, which is also available for our full sample.[9] As shown by Atkeson and Ohanian (2001) and Stock and Watson (2007), the one-year moving average of past inflation

[7]Following Krishnamurthy and Vissing-Jorgensen (2012), we subtract a long-term government bond yield until 2000 and the yield on a 20-year Treasury bond afterward to match the duration of the Moody's index, which contains corporate bonds with 20- to 30-years to maturity. See detailed construction in Appendix A.1.

[8]For the post-2011 sample, we cross-checked 3-month commercial paper rates against 3-month repo rates from JP-Morgan markets (proprietary data), and found that they are similar. For replicability, we use the publicly-available data on commercial paper rates.

[9]We use the Consumer Price Index for All Urban Consumers (CPI-U) published by the Bureau of Labor Statistics. We download the series from Shiller (2016), who reports the data starting from the late 1800s. The CPI series is also available from St. Louis FRED starting from 1947 (ticker: CPIAUCSL).

is one of the most robust predictors of future inflation.

We control for known drivers of Treasury convenience, as documented by earlier research, in particular, market volatility, the total government debt supply, and the stance of monetary policy. For market volatility, we use the VIX index. VIX data is only available since 1990. For the period before 1990, we use a linear projection of the VIX on realized volatility of S&P 500 index returns, where the projection coefficients are estimated on the post-1990 data. For government debt supply, we use the total quantity of Treasury debt, at market value, excluding intra-governmental holdings and holdings by depository institutions and the Federal Reserve. The data construction follows Krishnamurthy and Li (2023). For monetary policy, we use the end-of-month effective federal funds rate, available from the flow of funds data.

The Aaa-Treasury spread is commonly interpreted as reflecting the convenience yield of Treasuries, rather than default risk of Aaa bonds, because short- and long-term default rates for Aaa bonds are nearly indistinguishable from zero. As reported by Moody's (Emery et al., 2009), Aaa-rated bonds have never defaulted in history. To understand the credit risk of Aaa-rated bonds over the lifetime of the bond, we employ a simple calculation based on the rating migration rates, default rates, and loss rates reported by Emery et al. (2009) (see Appendix A.2 for details). This calculation implies that the annualized credit loss rate over a 20-year horizon for Aaa bonds is only 0.007%, but that of Baa is 0.31%, a 40-fold difference. See also the substantial literature on the credit spread puzzle, which has found that a liquidity component is needed to explain the spread on highly-rated and especially Aaa-rated corporate bonds, e.g. Huang et al. (2025). To further alleviate concerns about residual credit risk affecting the time-variation in the Aaa-Treasury spread, we also control for the Baa-Aaa spread in our main regressions.

We consider three periods for our empirical analysis that capture distinct macroeconomic regimes: the pre-WWII period (1926:01 through 1939:08), the second half of the 20th century (1952:01 through 1999:12), and the post-2000 period until the end of 2020 (2000:01 through 2020:12). The start of the sample is determined by the availability of daily stock returns used to construct stock return volatility. The sample ends in 2020, before the post-COVID inflation period, which we analyze separately in Section 5. We exclude the period of WWII until 1951 due to interest rate controls, which were lifted by the Treasury-Fed accord.[10] We set a break date

[10]From April 1942 until the Treasury-Fed Accord on March 4, 1951 interest rates were formally pegged at 3/8% on short-term Treasury bills and implicitly at 2.5% on long-term Treasury bonds, initially with the intention to allow cheaper WWII financing. See `https://www.federalreservehistory.org/essays/treasury-fed-accord`. We therefore restart the sample in 1952:01 after the peg was removed.

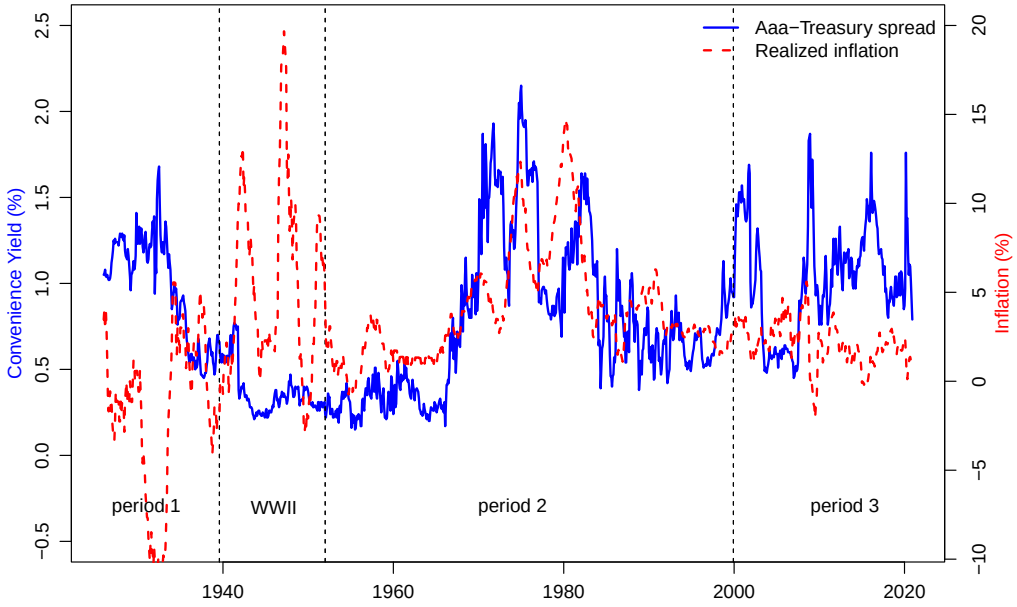
in January 2000, because the literature has found significant changes in the inflation dynamics and their relationship with the real economy around this time (e.g., Campbell et al. (2017), Stock and Watson (2007)). We verify the robustness of our findings to the specific period cutoffs in the subsequent analysis.

Table 1. Summary statistics. This table presents summary statistics for our full sample 1926:01–2020:12, excluding the 1939:09–1951:12 period.

Variable	Obs	Mean	Std. Dev.	Mean		
	(1926-2020)			(1926-1939)	(1952-1999)	(2000-2020)
Aaa-Tsy spread (%)	992	0.87	0.42	0.98	0.78	1.03
T-bill convenience yield (%)	992	0.43	0.46	0.28	0.56	0.23
Inflation (%)	992	2.60	3.64	-1.58	3.99	2.13
VIX (%)	992	19.64	8.16	27.39	17.30	19.96
Baa-Aaa spread (%)	992	1.14	0.70	2.03	0.93	1.05
Debt/GDP	992	0.30	0.15	0.17	0.26	0.47

Table 1 presents summary statistics for our key variables. Figure 2 visualizes the non-standardized series for the Aaa-Treasury spread and inflation, with vertical lines indicating the subperiod break dates. Average inflation is 4% in the middle period, which encompasses the high-inflation 1970s and 1980s, and much lower in the other periods. The Aaa-Treasury spread averages around 87 bps, with similar means across subperiods, while the T-bill convenience is about half that at 43 bps on average. A simple calculation illustrates the economic magnitude of the convenience yield. In 2024, the U.S. government debt had a duration of around 5 years and a total outstanding value of 27 trillion USD (market value of debt held by the public). A convenience yield of 87 bps would then translate roughly into extra U.S. fiscal capacity of $5 \times \frac{0.87}{100} \times 27 = 1.17$ trillion dollars. While the magnitude of the Aaa-Treasury spread is comparable to that of the Baa-Aaa spread, recall that the credit loss rates of Aaa bonds are negligible compared to the credit loss rates of Baa bonds. Thus, we interpret the Aaa-Treasury spread as primarily capturing non-credit related Treasury convenience, while the Baa-Aaa spread is a useful control for credit risk.

Figure 2. Time series of Aaa-Treasury spread and inflation. This figure shows the measures of inflation and the Aaa-Treasury spread. Vertical lines indicate the subperiods used in subsequent analysis, excluding the WWII period until the end of yield curve controls (marked as "WWII" on the graph). The three subperiods shown are 1926:01–1939:08, 1952:01–1999:12, and 2000:01–2020:12.



3 A Century of Evidence on Inflation and Treasury Convenience

In this section, we present our main empirical results. We first show that the relation between inflation and the Treasury convenience spread has changed in a quantitatively and statistically significant manner over the three periods we consider. In particular, the comovement between inflation and convenience was positive during the inflationary second half of the 20th century, and significantly different from the periods before or after. We then provide more finely grained evidence based on inflation components, lead-lags, and bond-stock betas. Taken together, this evidence supports the existence of competing "money" and "liquidity demand" channels associated with supply- vs. demand-driven inflation, and driving opposite signs for the inflation-Treasury convenience relationship.

3.1 The Changing Treasury Convenience-Inflation Relationship

As shown in Figure 1, the correlation between inflation and the Aaa-Treasury spread changes from negative -0.56 pre-WWII to positive 0.63 in the second half of the 20th century, and back to negative -0.28 post-2000. To assess the statistical and economic significance of those changes, we estimate the following regression at a monthly frequency:

$$spread_t^{Aaa} = b_0 + b_1\pi_t + b_2\pi_t \times I_{\text{1952-1999},t} + b_3\pi_t \times I_{\geq 2000,t} + \Gamma' X_t + \varepsilon_t, \tag{1}$$

where we interact the annual inflation rate with period-specific dummy variables. The interaction coefficients are interpreted relative to the pre-WWII period (the omitted category). π_t is the inflation rate over the 12 months from $t-12$ to t. The vector X_t captures time t controls.

Table 2. Shifts in long-term Treasury convenience–inflation relationship. Monthly data runs from 1926:01 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 (1926:01–1939:08) 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.

	Long-term convenience: Aaa-Treasury spread				
	(1)	(2)	(3)	(4)	(5)
Inflation	-0.038***	-0.038***	-0.022**	-0.013	-0.034***
	(-5.15)	(-5.31)	(-2.40)	(-1.00)	(-4.61)
Inflation x $I_{1952-1999}$	0.13***	0.12***	0.094***	0.078***	0.12***
	(6.45)	(3.87)	(3.29)	(2.64)	(5.28)
Inflation x $I_{\geq 2000}$	-0.034	-0.049	-0.083**	-0.083*	-0.072*
	(-0.93)	(-1.33)	(-2.02)	(-1.96)	(-1.81)
FFR		0.017	0.0096	0.0085	
		(0.86)	(0.50)	(0.46)	
Debt/GDP			-0.57	-0.58	-0.60
			(-1.43)	(-1.48)	(-1.61)
VIX			0.0093***	0.0058	
			(2.88)	(1.64)	
Baa spread				0.10	
				(1.59)	
$I_{1952-1999}$	-0.52***	-0.51***	-0.34***	-0.25	-0.42***
	(-5.12)	(-5.30)	(-2.72)	(-1.57)	(-3.41)
$I_{\geq 2000}$	0.27***	0.31***	0.56***	0.60***	0.51***
	(2.71)	(2.89)	(2.92)	(2.96)	(2.71)
Constant	0.92***	0.88***	0.76***	0.67***	1.03***
	(13.20)	(10.56)	(4.90)	(3.85)	(11.48)
$\bar{R}^2$	0.39	0.39	0.43	0.44	0.41
N	992	992	992	992	992

Table 2 shows that the key coefficient of interest – the interaction between inflation and the 1952–1999 dummy – is consistently estimated to be positive. Thus, the relationship between inflation and the Aaa-Treasury spread is significantly more positive over the second half of the 20th century – which includes the Great Inflation of the 1970s and 1980s – than over other periods. The coefficients in column (1) imply that a one percentage point increase in inflation is associated with a 13 bps higher Aaa-Treasury spread in the 1952–1999 period compared to pre-WWII. This magnitude is large relative to an average Aaa-spread of 87 bps reported in Table 1. Said differently, a one-standard-deviation increase in inflation is associated with roughly a one-standard-deviation greater Aaa-Treasury spread in the 1952–1999 period. By contrast, the negative baseline coefficient on inflation means that a one percentage point increase in inflation tends to be associated with a four bps decrease in the Aaa-Treasury spread during the pre-WWII period. The relationship is similar or even more negative during the 2000s.

The coefficients of long-term Treasury convenience on inflation and period interactions remain quantitatively unchanged and statistically significant when we control for the fed funds rate as a proxy for monetary policy. These results also hold controlling for potential other drivers of Treasury convenience, such as the government debt/GDP ratio, equity volatility, and even the Baa credit spread, underscoring that the switch in the inflation-spread relationship is specific to the convenience premium in Treasuries, as distinct from how inflation affects credit risk (e.g., Kang and Pflueger (2015), Brunnermeier et al. (2023), Bhamra et al. (2023)).[11]

Controlling for debt/GDP is important, as prior research has found that an increase in the quantity of Treasury debt tends to reduce the convenience of Treasury bonds (Krishnamurthy and Vissing-Jorgensen, 2012; Krishnamurthy and Li, 2023). Columns (3)–(5) of Table 2 show that debt/GDP enters negatively with a quantitatively meaningful coefficient of around -0.60, similar to the magnitude reported by Krishnamurthy and Vissing-Jorgensen (2012), but the time-varying relationship between convenience and inflation is little affected by controlling for debt/GDP. The statistical significance of the debt/GDP variable is reduced because inflation and debt/GDP are negatively correlated in our sample.[12] Debt/GDP is also a simple proxy for fiscal channels driving

[11] We also verify that adding the slope of the Treasury yield curve as a control does not affect these results. The coefficient loading on the slope is insignificant. To the extent that the slope proxies for the variation in the Treasury term premia, this suggests that the results are unlikely to be driven by the duration mismatches between the Aaa corporate bonds and the Treasuries. To keep regressions relatively parsimonious, we do not include the slope in subsequent specifications.

[12] Following long-standing literature (Krishnamurthy and Vissing-Jorgensen, 2012) we use raw debt/GDP, which shows that debt/GDP spiked during 2008-2009 just as inflation fell. By contrast, Fu et al. (2023) use linearly de-

inflation, though the relationship between fiscal channels and inflation may be complicated by endogenous debt issuance and term premia (Cochrane, 2001; Corhay et al., 2023).

Table 3. Shifts in T-bill convenience–inflation relationship. Monthly data runs from 1926:01 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 (1926–1939:08) 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.

	Short-term convenience: T-bill spread				
	(1)	(2)	(3)	(4)	(5)
Inflation	-0.029***	-0.027***	-0.016***	-0.021***	-0.029***
	(-4.69)	(-5.95)	(-2.83)	(-2.99)	(-4.52)
Inflation x $I_{1952-1999}$	0.15***	0.082***	0.071***	0.078***	0.15***
	(9.04)	(3.95)	(3.69)	(3.37)	(9.22)
Inflation x $I_{\geq 2000}$	0.070***	0.0040	0.013	0.013	0.063***
	(4.30)	(0.16)	(0.49)	(0.51)	(3.18)
FFR		0.075***	0.076***	0.076***	
		(6.00)	(6.10)	(6.18)	
Debt/GDP			0.22	0.22	-0.11
			(0.99)	(1.01)	(-0.50)
VIX			0.0092***	0.011***	
			(3.77)	(3.23)	
Baa spread				-0.048	
				(-0.87)	
$I_{1952-1999}$	-0.15**	-0.15**	-0.093	-0.14	-0.14*
	(-2.12)	(-2.15)	(-1.23)	(-1.38)	(-1.90)
$I_{\geq 2000}$	-0.091	0.11	0.051	0.030	-0.046
	(-1.48)	(1.48)	(0.46)	(0.27)	(-0.44)
Constant	0.24***	0.046	-0.23**	-0.19*	0.25***
	(4.87)	(0.90)	(-2.41)	(-1.87)	(4.07)
$\bar{R}^2$	0.49	0.58	0.60	0.60	0.49
N	992	992	992	992	992

Table 3 shows that secular shifts also occurred in the relationship between short-term Treasury convenience and inflation, with a positive coefficient on the interaction between inflation and the 1952–1999 dummy in column (1). However, different from the long-term spread, columns (2) through (4) of Table 3 show that the fed funds rate enters significantly, consistent with Nagel (2016)'s finding that short-term convenience is strongly positively correlated with the fed funds rate. While controlling for the fed funds rate in column (2) somewhat reduces the strength of the relationship between short-term convenience and inflation in the 1952-1999 period, the interaction between inflation and the 1952-1999 dummy remains positive, economically meaningful, and

trended debt/GDP growth.

statistically significant. In column (2) the post-2000 interaction is not significantly different from the baseline pre-WWII period, showing that controlling for the fed funds rate also helps uncover a negative relationship between inflation and T-bill convenience in the post-2000 period, similarly to Table 2. These findings are consistent with our model in Section 4, where the money channel links short-term convenience to the contemporaneous nominal policy rate, but long-term convenience to the persistent component in inflation. A relationship between inflation and short-term convenience may arise after controlling for the policy rate if our proxy for short-term convenience is somewhat more long-term than the overnight policy rate, as is plausible in practice. Overall, secular shifts in the inflation-convenience relationship are similar for long-term and short-term convenience, and while the contemporaneous policy rate matters little for long-term convenience, it is an important control to uncover these relationships for short-term convenience.

3.2 Evidence from Inflation Components

Having shown the changing relationship between headline inflation and Treasury convenience, we now turn to inflation components. We provide evidence for different measures of persistent and expected inflation, before turning to a decomposition of inflation into supply- and demand-inflation from disaggregated product data (Shapiro, 2024). We thereby provide direct evidence that the positive inflation-convenience relationship in the second half of the 20th century can be attributed to aggregate supply disturbances driving persistent inflation, whereas the negative inflation-convenience relationship is associated with changes in demand.

3.2.1 Persistent Inflation Components

Table 4 shows that the positive headline inflation-convenience relationship that we found for the second half of the 20th century holds robustly for different measures of expected and persistent inflation, including core inflation, survey-based inflation expectations, and trend inflation from the econometric model of Stock and Watson (2007).[13] If measures such as survey inflation expectations and core inflation reveal relatively more persistent supply shocks, they should be associated

[13]Cleveland Fed produces alternative estimates of inflation expectations used in several recent studies. The Cleveland Fed estimates are based on a model that relies on Treasury yields as inputs, in addition to other survey and market-based proxies for inflation expectations. Given those inputs, those estimates can reflect term premia and convenience yields. Since this can lead to spurious correlations between expected inflation and convenience (as we show in Appendix A.3), we do not rely on any estimates of expected inflation that involve yield curve information.

Table 4. Convenience yield and persistent inflation. The table reports regressions of the Aaa-Treasury spread on different measures of inflation: headline CPI inflation, the 4-quarter ahead GDP deflator forecast from the Survey of Professional Forecasters (SPF), trend inflation, and core and energy CPI inflation. All inflation variables are expressed in percent per annum. Trend inflation is based on the Stock and Watson (2007) unobserved components model with stochastic volatility, estimated following Chan (2018). Trend inflation and SPF forecasts are available quarterly, remaining variables are monthly. For headline, trend, core, and energy inflation, the sample is 1959–2020; the sample for the SPF forecasts is 1970:Q2–2020:Q4. Controls include covariates from column (4) in Table 2.

	Long-term convenience: Aaa-Treasury spread							
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Infl (head)	0.089***	0.081***						
	(4.24)	(3.21)						
Infl (head) x $I_{\geq 2000}$	-0.16***	-0.099**						
	(-3.87)	(-2.26)						
$\mathbb{E}_{4q}^{\text{SPF}}$(GDP defl)			0.077***	0.073**				
			(2.69)	(2.15)				
$\mathbb{E}_{4q}^{\text{SPF}}$(GDP defl) x $I_{\geq 2000}$			-0.28	-0.0046				
			(-1.11)	(-0.02)				
Trend infl					0.079***	0.054**		
					(3.78)	(2.17)		
Trend infl x $I_{\geq 2000}$					-0.33***	-0.40***		
					(-3.29)	(-4.39)		
Infl (core)							0.12***	0.12***
							(5.81)	(4.09)
Infl (core) x $I_{\geq 2000}$							0.021	0.092
							(0.16)	(0.79)
Infl (eng)							-0.0076	-0.0044
							(-1.18)	(-0.70)
Infl (eng) x $I_{\geq 2000}$							-0.0027	-0.00038
							(-0.39)	(-0.05)
$I_{\geq 2000}$	0.71***	0.43***	0.80	0.12	1.09***	1.24***	0.43*	0.12
	(6.29)	(2.88)	(1.56)	(0.23)	(5.06)	(5.69)	(1.74)	(0.46)
Constant	0.48***	0.34*	0.63***	0.74**	0.51***	0.40*	0.35***	0.29
	(5.33)	(1.72)	(4.66)	(2.45)	(5.36)	(1.89)	(3.81)	(1.47)
Controls	No	Yes	No	Yes	No	Yes	No	Yes
$\bar{R}^2$	0.31	0.45	0.084	0.28	0.27	0.46	0.38	0.48
N	744	744	202	202	248	248	744	744

with higher long-term convenience via the "money channel," and we would thus expect a consistently positive relationship between these measures of inflation and Treasury convenience across periods.

Columns (1) and (2) of Table 4 show that the results from Table 2 carry over to a shorter sample starting in 1959, when core and energy inflation become available. Columns (3) and (4) use the 4-quarter forecast for real GDP inflation from the Survey of Professional Forecasters (SPF), which is however only available for an even shorter sample starting in the second quarter of 1970 and at a quarterly frequency. As shown by Chernov and Mueller (2012), survey data on inflation contain additional useful information for understanding longer-term inflation dynamics. Here, we see that the positive relationship between inflation expectations and Treasury convenience in the second half of the 20th century continues to hold, but the interaction with the post-2000 dummy becomes insignificant. This result for the post-2000 period can be rationalized given that survey inflation expectations were very stable during this period, and did not incorporate shorter-lived inflation disturbances representative of liquidity demand shocks.

Columns (5) and (6) use a measure of trend inflation from the inflation forecasting literature (Stock and Watson, 2007), estimated following Chan (2018) at quarterly frequency. Trend inflation, based on the unobserved components trend-cycle model of inflation with stochastic volatility, can be thought of as filtering out a time-varying amount of short-term noise in realized inflation. As in Table 2, we find that trend inflation is positively associated with convenience during the second half of the 20th century, but negatively associated with convenience during the post-2000 period. The interaction with the post-2000 dummy may differ from the result for survey inflation expectations in columns (3) and (4), because the trend variable smooths transitory noise in inflation but likely still preserves some variation due to aggregate demand fluctuations during the post-2000 period.[14]

The last two columns of Table 4 decompose headline inflation into its core and energy components. Although simple, this decomposition has the advantage of being available in real-time without the forward-looking bias of most econometric models. It also reflects the way the Federal Reserve thinks about inflation, where core inflation is viewed to capture the more persistent inflation movements, whereas fluctuations around core, driven by food and energy prices, are viewed as more transitory.[15] Columns (7) and (8) of Table 4 show that the Aaa-Treasury spread displays

[14]Over the post-2000 period, the standard deviation of trend inflation was 0.56, but of 4-quarter GDP inflation forecasts was only 0.22 (all in annualized percent).

[15]For example, Fed Chair Jerome Powell stated in his 2023 Jackson Hole speech "Food and energy prices are influ-

a stable positive relationship with core inflation, for both the pre- and post-2000 samples, as indicated by the insignificant interaction with the post-2000 dummy. The relationship of long-term convenience with energy inflation is statistically indistinguishable from zero. The decomposition into core and energy inflation hence again confirms the stable positive relationship between the more persistent components of inflation with Treasury convenience.

Overall, different measures of persistent or broad-based inflation confirm the robustly positive inflation-convenience relationship during the second half of the 20th century. The inflation-convenience relationship is stable across subperiods for survey inflation expectations and core inflation, which is as expected if these measures filter out the more transitory demand components of inflation. Taken together, the results in Table 4 are in line with an interpretation whereby the money channel induces a consistently positive inflation-convenience relationship, though its effect on the headline inflation-convenience relationship may at times be concealed by competing liquidity demand shocks.

3.2.2 Demand- and Supply-Driven Inflation

We next show results for an explicit decomposition of inflation into its demand- versus supply-driven components from Shapiro (2024). Shapiro (2024) uses monthly price, quantity, and expenditure data for more than 100 goods and services categories in the personal consumption expenditures (PCE)[16] to separate demand from supply shocks at an individual product category level. The shocks are then aggregated to obtain demand- and supply-driven PCE inflation. The decomposition is available starting from 1969:12, thus covering the key part of the high inflation period and the 2000s.

Table 5 reports our baseline contemporaneous regressions of Aaa-Treasury spreads on supply and demand components of PCE inflation, with the decomposition for core PCE (which the Fed mostly focuses on) in the first three columns, and headline PCE in the subsequent three columns. The full sample regressions of the Aaa-Treasury spread onto supply and demand components of

enced by global factors that remain volatile, and can provide a misleading signal of where inflation is headed. In my remaining comments, I will focus on core inflation, which omits the food and energy components." (August 25, 2023, speech by Fed Chair Jerome Powell). While it is well known that energy inflation can reflect both supply and demand shocks, demand shocks have been important for energy prices since at least the early 1990s. For decompositions of energy prices see Kilian (2009) and Baumeister et al. (2022).

[16]The PCE and the CPI inflation (which we used so far) trace each other relatively closely. However, the construction of the two indices differs in the scope of the consumption basket, the weights of individual items, and how the basket changes are accounted for. Since 2000, the Federal Reserve has used PCE inflation as the preferred inflation measure.

inflation in columns (1) and (4) show that Treasury convenience has a positive relationship with supply-driven inflation, but a negative relationship with demand-driven inflation. The next columns add interactions between the inflation components and a post-2000 dummy and add controls. This analysis shows that the positive association with supply inflation is predominantly driven by the pre-2000 period, as one might expect if there were few supply shocks post-2000. The post-2000 relationship between supply inflation and the Aaa-Treasury spread is statistically indistinguishable from zero. While the relationship between long-term convenience and demand inflation loses power for the shorter subsamples, the point estimates are consistently negative for each subperiod, indicating that demand-driven inflation has a negative relationship with convenience. These relationships between convenience and the components of inflation are similar without and with our full set of controls.

Figure 3 shows that the different-signed relationships between long-term convenience with supply- vs. demand-inflation are even clearer when we analyze lead-lag patterns. We predict future changes in the annual inflation rate from month t to $t+h$ with a one-month change in the Aaa-Treasury spread from month $t-1$ to t, and vice versa. We interpret the results in the spirit of Granger causality. Specifically, we estimate the following predictive regressions separately considering the demand and supply drivers of core PCE inflation:

$$\pi_{t+h}^{core,j} - \pi_t^{core,j} = a_h + b_h \Delta spread_t^{Aaa} + \gamma_h X_t + \varepsilon_{t+h} \tag{2}$$

$$spread_{t+h}^{Aaa} - spread_t^{Aaa} = c_h + d_h \Delta \pi_t^{core,j} + \delta_h X_t + \epsilon_{t+h}, \tag{3}$$

where inflation is split into demand- and supply-driven components, $j = \{dem, sup\}$. We augment these regressions with the full set of controls, X_t, using the first differences of the variables included in Table 5. During the 1969–1999 period, an increase in supply-driven inflation predicts a significant increase in the Aaa-Treasury spread with a peak effect at about 20 months ahead (Panel A). In contrast, in the post-2000s period, a higher spread predicts a lower demand-driven inflation, with a maximum effect at around 24 months ahead (Panel B).[17] We find no evidence of higher spreads forecasting a higher supply inflation, or higher demand inflation forecasting higher spreads

[17] In terms of magnitudes, the estimates for the 1969–1999 period imply that a one-standard-deviation positive monthly change in the supply inflation (or 13 bps) predicts a 0.3 standard deviations (or 13 bps) increase in the Aaa-Treasury spread over the next 24 months (top left panel in Figure 3). Conversely, in the post-2000 period, a one-standard-deviation monthly increase in the Aaa-Treasury spread (or 11 bps) predicts a 0.23 standard deviations (or 15 bps) decline in demand inflation over the next 24 months (bottom right panel).

Table 5. Convenience yield onto demand and supply inflation components. The table reports regressions of long-term convenience spread on demand and supply components inflation for core inflation in the first three columns, and for headline inflation in the last three columns. $I_{\geq 2000}$ is a dummy variable equal to one from 2000 onward. The sample period is 1969:12–2020:12. Newey-West t-statistics with 12 lags are shown in parentheses. * p <0.10, ** p <0.05, *** p <0.01.

	Long-term convenience: Aaa-Treasury spread					
	(1)	(2)	(3)	(4)	(5)	(6)
	Core decomposition			Headline decomposition		
Infl (sup)	0.25***	0.24***	0.22***	0.15***	0.18***	0.18***
	(5.62)	(4.25)	(4.44)	(3.44)	(3.92)	(3.94)
Infl (dem)	-0.29***	-0.17	-0.11	-0.17**	-0.11	-0.076
	(-4.13)	(-1.12)	(-0.85)	(-2.18)	(-0.89)	(-0.77)
Infl (sup) x $I_{\geq 2000}$		-0.52***	-0.40*		-0.25**	-0.23***
		(-2.63)	(-1.88)		(-2.56)	(-2.76)
Infl (dem) x $I_{\geq 2000}$		-0.32	-0.085		-0.13	0.013
		(-1.51)	(-0.41)		(-0.88)	(0.10)
FFR			-0.041**			-0.049***
			(-2.56)			(-2.72)
VIX			0.0086**			0.0078*
			(2.13)			(1.96)
Debt/GDP			-0.48			-0.52
			(-1.13)			(-1.33)
Baa spread			0.16**			0.22***
			(2.05)			(2.93)
$I_{\geq 2000}$		0.81***	0.48*		0.51***	0.27
		(3.30)	(1.66)		(3.63)	(1.64)
Constant	0.90***	0.71***	0.77***	0.92***	0.71***	0.79***
	(13.26)	(4.59)	(2.91)	(12.86)	(6.15)	(3.32)
$\bar{R}^2$	0.26	0.35	0.45	0.16	0.28	0.42
N	613	613	613	613	613	613

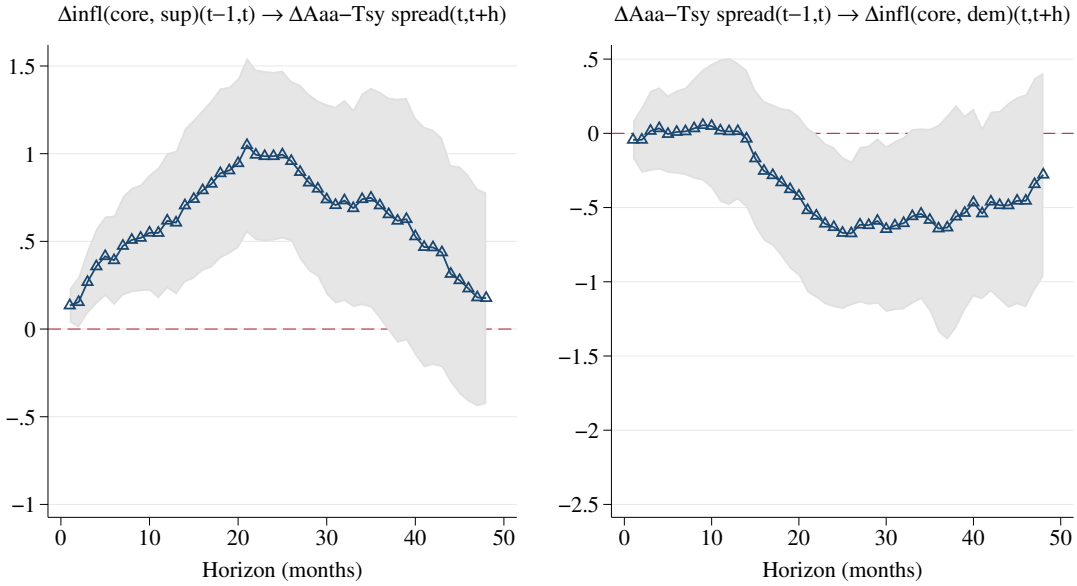
(shown in Appendix Figure A4). While these results are not a direct test of causality, they indicate that supply- vs. demand-driven factors alter the causal interpretation of the convenience-inflation relationship.

3.3 Results by Bond-Stock Betas

So far, we have divided different historical episodes based on narratives about underlying shocks. If the changing Treasury convenience-inflation relationship over broad historical periods is indeed due to the changing dominance of fundamental economic shocks, bond-stock betas should provide an additional criterion for the split. This approach builds on the literature characterizing comovements of stocks and bonds in supply- and demand-driven environments (e.g., Campbell et al.,

Figure 3. Predictive regressions with demand- vs. supply-driven inflation components This figure presents coefficients from predictive regressions (2) and (3). The controls include the monthly changes from $t-1$ to t of FFR, VIX, debt/GDP, and Baa-Aaa spread, i.e., the first differences of controls in Table 5. The right panels display the b_h coefficients in regression (2) predicting changes in the demand-driven component of core inflation with the Aaa-Treasury spread changes. The left panels display the d_h coefficients in regression (3) predicting the Aaa-Treasury spread changes with changes in the supply component of core inflation. The gray shading indicates 95% confidence intervals based on Newey-West standard errors with h lags.

Panel A: 1969:12–1999:12



Panel B: 2000:01-2020:12



2017, 2020; Ermolov, 2022; Pflueger, 2025), whereby inflation driven by aggregate supply shocks (or "bad inflation" as in Cieslak and Pflueger (2023)) induces a positive stock-bond beta and inflation due to aggregate demand (or "good inflation") induces a negative stock-bond beta. Indeed, the model in Section 4 predicts a positive Treasury convenience-inflation relationship when supply shocks are dominant, and bond-stock betas are positive, and vice versa.

Table 6 shows that the Treasury convenience-inflation relationship increases monotonically with the bond-stock beta, where a positive bond-stock beta indicates more volatile supply shocks and a negative bond-stock beta indicates more volatile aggregate demand shocks.[18] Notably, when we condition on a negative bond-stock beta, the inflation-convenience relationship becomes negative. This fact aligns with the interpretation that a negative inflation-convenience relationship arises when demand shocks are especially prominent. The magnitude of the convenience-inflation coefficient between the lowest beta and highest beta observations is substantially different. When the bond-stock beta is above 0.2 (roughly the top quintile of bond-stock beta observations), a one percentage point increase in inflation leads to an 11 bps higher convenience spread than when the bond-stock beta is below -0.1 (roughly the bottom quintile of bond-stock beta observations), roughly in line with our baseline magnitudes.

So far, we have seen that a sample split by bond-stock betas is informative about the economic drivers of negative vs. positive convenience-inflation relationships. Table 7 shows that our baseline results are robust to controlling for bond-stock betas directly. Bond-stock betas could directly affect the convenience yield by changing the hedging properties of Treasuries, as argued by Acharya and Laarits (2023). Or negative bond-stock betas might reflect more volatile demand shocks due to demand for liquid Treasuries, which in turn might be correlated with the level of convenience yields. Column (4) shows that while the bond-stock beta enters with a negative sign for the full sample (confirming the results in Acharya and Laarits (2023)), its inclusion does not substantively change the significantly positive coefficient on inflation. Moreover, subperiod regressions show that our main results are unaffected. In particular, columns (5) and (6) show that the coefficient on inflation is positive and significant for the 1961-1999 period, but negative and significant in the 2000-2020 period. Comparing to columns (2) and (3) shows that neither the magnitude nor

[18] These results are for the post-1961 subsample, dictated by the availability of Gürkaynak et al. (2007) zero-coupon yield curves. The bond-stock beta is the 120-day rolling beta of daily returns for a nominal zero-coupon Treasury bond with seven years to maturity (the longest maturity available over the full period) onto daily stock returns. Daily nominal Treasury bond returns are computed from zero-coupon bond yields. Daily stock market index returns are from Ken French's website.

Table 6. Aaa-Treasury spread inflation for subsamples split by Treasury bond-stock beta. The bond-stock beta is the 120-day rolling beta of daily returns for a nominal zero-coupon Treasury bond with seven years to maturity onto daily stock returns. Daily nominal Treasury bond returns are computed from zero-coupon bond yields (Gürkaynak et al., 2007). Daily stock market index returns are from Ken French's website. The sample period is from 1961:11–2020:12. Robust Newey-West t-statistics (column (1)) and robust t-statistics (all other columns) in parentheses. * p <0.10, ** p <0.05, *** p <0.01.

	(1) 1961-2020	(2) beta≤ -0.1	(3) beta≤ 0	(4) beta> 0.1	(5) beta> 0.2	(6) beta> 0.3
Inflation	0.043***	-0.064***	-0.0032	0.045***	0.052***	0.052***
	(2.48)	(-3.91)	(-0.20)	(9.61)	(7.12)	(5.93)
Constant	0.78***	1.19***	0.96***	0.64***	0.61***	0.62***
	(30.51)	(30.67)	(22.88)	(21.51)	(15.67)	(12.88)
$\bar{R}^2$	0.092	0.090	0.000	0.16	0.18	0.21
N	710	152	284	261	165	105

Table 7. Aaa-Treasury spread onto inflation controlling for bond-stock beta. This table presents regressions of Aaa-Treasury spread onto inflation and bond-stock beta. 120-day rolling Treasury bond-stock betas are constructed as described in Table 6. The sample period is from 1961:11–2020:12. Newey-West t-statistics with 12 lags are shown in parentheses. * p <0.10, ** p <0.05, *** p <0.01.

	(1) 1961-2020	(2) 1961-1999	(3) 2000-2020	(4) 1961-2020	(5) 1961-1999	(6) 2000-2020
Inflation	0.043**	0.078***	-0.072**	0.059***	0.079***	-0.070**
	(2.48)	(3.73)	(-2.01)	(3.47)	(3.83)	(-2.15)
Bond-stock beta				-0.63***	-0.19	-0.76
				(-3.68)	(-0.71)	(-1.48)
Constant	0.78***	0.53***	1.19***	0.76***	0.56***	1.09***
	(9.48)	(5.35)	(17.07)	(9.95)	(4.94)	(14.34)
$\bar{R}^2$	0.092	0.29	0.080	0.17	0.30	0.12
N	710	458	252	710	458	252

significance of the inflation coefficient are affected by the inclusion of the bond-stock beta, which itself becomes insignificant in columns (5) and (6).

Taken together, using bond-stock betas as a financial markets indicator shows that the Treasury convenience-inflation relationship is most positive when inflation is perceived to be driven by aggregate supply shocks, while the relationship is most negative when inflation is perceived

to be driven by aggregate demand shocks. At the same time, controlling for bond-stock betas in the regression leaves our baseline findings unchanged. Overall, these results further support an interpretation of two competing channels, whereby the positive Treasury convenience-inflation relationship in the 1970s and 1980s was due to supply shocks and persistent inflation, whereas the negative Treasury convenience-inflation relationship during the post-2000 period was associated with demand-driven inflation.

3.4 Robustness

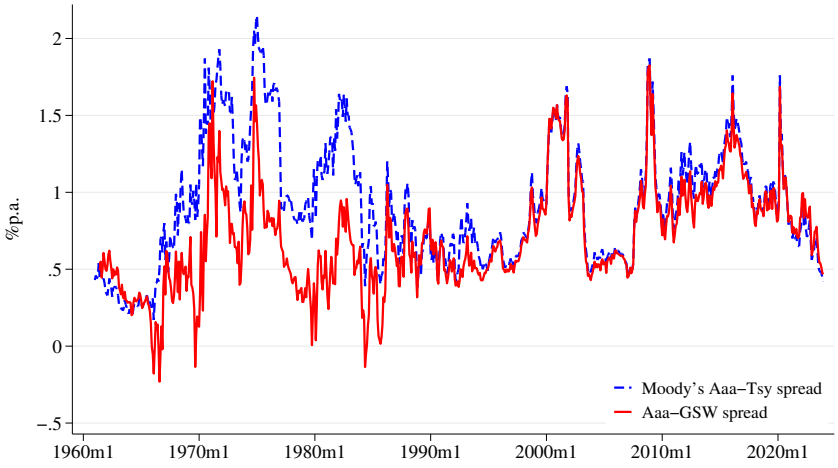
In the analysis so far we follow the literature using the Moody's Aaa-Treasury spread as a measure of long-term convenience, in the absence of alternatives available over the long historical sample (Krishnamurthy and Vissing-Jorgensen, 2012). However, the Moody's Aaa-Treasury spread can be confounded by other factors unrelated to convenience. The specific 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. Below, we discuss the adjustments we undertake to address these points and how they affect our results.

As recently highlighted by Lehner et al. (2025), the benchmark long-term Treasury yield used by Moody's includes the so-called flower bonds.[19] Flower bonds offer bondholders additional benefits similar to life insurance as they could be redeemed at par to cover the payment of estate taxes upon the holder's death. They were issued before 1966 with coupon rates below 4.5% and were effectively the only US long-term government bonds available in the early part of the sample. The analysis of Mayers and Smith (1987) indicates that the option became especially valuable when interest rose in the second half of the 1970s. This effect depresses the Treasury yield used by Moody's, and therefore overstates the value of Treasury convenience. Accordingly, using a carefully constructed Treasury benchmark that excludes flower bonds, Lehner et al. (2025) show that the long-term Aaa-Treasury convenience spread was significantly lower from the mid-1970s through the early 1980s than the Moody's Aaa spread would imply.

To study how this issue affects our analysis, 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

[19]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 4. 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.



years; it extends to 10 years in August 1971, to 15 years in November 1971, and to 20 years in July 1981. Since Moody's Aaa index is based on bonds with a remaining maturity of at least 20 years, we approximate the long-term government counterpart with a 20-year GSW par yield when it becomes available. For the prior years, we use the longest available maturity in GSW, assuming a flat par yield curve beyond the last available maturity knot.[20]

We denote the spread between the Moody's Aaa yield and the long-term GSW par yield as the Aaa–GSW spread. Figure 4 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.

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

[20] 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.

the constituents of the Moody's Aaa index, we rely on the duration estimates from van Binsbergen et al. (2025) in the combined Lehman/Warga and Bank of America Merill Lynch datasets. This data becomes available from March 1974. The time-series correlation between the average duration of Aaa corporate bonds with at least 20 years to maturity and the 20-year par bond duration estimated using the long-term GSW par yield[21] 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.[22]

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).[23] 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.[24] 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.[25]

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[26] and use the fitted value as a proxy for the variation in the call, which we subtract from

[21] 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.

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

[23] 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.

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

[25] 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.

[26] We include the one-year yield as level, the spread between the 7-year yield and the one-year yield as slope, and

the Aaa–GSW spread. We denote this variable as "Aaa–GSW (ca) spread."

The adjustments above highlight significant limitations in the long-term Treasury and corporate bond data in the early parts of our sample. At the same time, the short-term convenience spreads should be less subject, if at all, to the confounding issues that pervade the long-term yields. As an alternative approach, we therefore construct a fitted long-term convenience spread by projecting Moody's Aaa spread onto the short-term T-bill convenience spread and volatility proxies (constructed VIX and its square) using only data prior to 1961, covering the period where Lehner et al. (2025) document relatively close overlap between their measure and the Moody's Aaa spread. We call this variable "Moody's Aaa–Treasury (fit) spread."

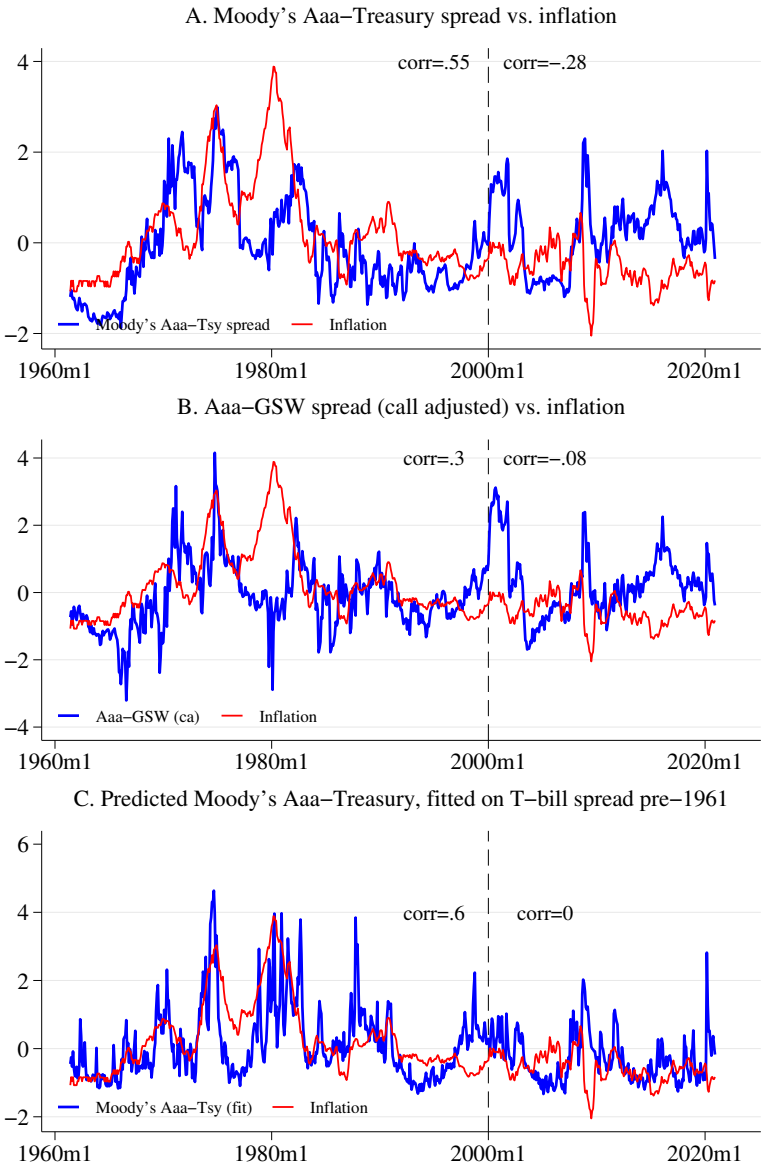
Figure 5 plots the different constructed spreads against inflation in the post-1961 sample and reports correlations before and after 2000. The comovement between inflation and convenience remains consistently positive before 2000, but is weakened for the GSW-based spread. The graph shows a robust relationship between the increase in inflation and Treasury convenience from the 1960s to the mid-1970s, and again as convenience declines with inflation after the early 1980s. However, there is substantial data uncertainty in the late 1970s, during the so-called monetarist experiment. This is perhaps not surprising given the multifaceted nature of shocks and the unprecedented interest rate volatility in this period,[27] which may have affected safety benefits of long-term Treasuries separately from the level of inflation considered throughout the paper. Importantly, the Moody's Aaa-spread fitted to the T-bill convenience displays a strong positive relationship with inflation throughout this volatile period.

In Table 8, we report regressions of various spreads on inflation and controls corresponding to column (4) in Table 2. We also include a separate dummy variable for the 1979–1980 period. The regressions are estimated over the post-1961 sample when GSW data become available. Column (1) presents the specification with the original Moody's Aaa spread for this shorter sample. Columns (2) through (4) present results based on the adjusted spreads discussed above. Higher inflation in the pre-2000 period is associated with higher adjusted spreads, with a magnitude of the coefficient of about 40% of that using the Moody's spread in column (1). In standardized units, this implies a one-standard-deviation inflation increase in the pre-2000 period is associated with approximately one-third standard deviation spread increase. Those magnitudes are consistent between the GSW-based spreads and the long-term spread predicted using the T-bill spread. Fi-

the realized monthly volatility of the 7-year yield changes. All variables are constructed using GSW data.

[27]The realized yield volatility between 1979–1980 was 3.4 times higher than in the prior years from 1961.

Figure 5. Proxies for long-term convenience spread vs. inflation. The spreads are (1) the original Moody's Aaa-Treasury spread; (2) Aaa-GSW spread using long-term par yield; (3) Aaa-GSW spread adjusted for time-varying moneyness of call options and (4) the predicted Moody's Aaa-Treasury spread obtained by fitting Moody's Aaa-Treasury spread with the T-bill spread and volatility using the 1925:01–1960:12 sample. All variables are standardized. The sample period is 1961:06 to 2020:12.



nally, for comparison, column (5) displays estimates for the T-bill convenience, which confirms a positive inflation-convenience link in the high inflation episode without being subject to various adjustments in the long-term spread.

Table 8. Regressions with different measures of long-term convenience spread. The table reports regressions of different spreads on inflation and controls corresponding to column (4) in Table 2. Additionally, $I_{1979-80}$ is a dummy variable taking a value of one in 1979 and 1980. In columns, the spreads are (1) the original Moody's Aaa-Treasury spread, (2) the Aaa-GSW spread based on the long-term GSW par yield, (3) the Aaa-GSW with adjustment for the moneyness of the call option, (4) Moody's Aaa-Treasury spread predicted by the T-bill spread and volatility, with coefficients fitted on the pre-1961 sample, and (5) T-bill liquidity spread. The sample period is 1961:06–2020:12. Newey-West t-statistics with 12 lags are shown in parentheses. * p <0.10, ** p <0.05, *** p <0.01.

	(1) Moody's Aaa-Tsy	(2) Aaa-GSW	(3) Aaa-GSW (ca)	(4) Moody's Aaa-Tsy (fit)	(5) T-bill spr
Inflation	0.10***	0.040**	0.040**	0.036***	0.081***
	(5.59)	(2.26)	(2.15)	(3.23)	(3.22)
Inflation x $I_{\geq 2000}$	-0.12***	-0.038	-0.033	-0.042**	-0.096**
	(-3.02)	(-1.10)	(-0.93)	(-2.53)	(-2.52)
FFR	-0.035**	-0.029*	-0.0023	0.038***	0.084***
	(-2.10)	(-1.83)	(-0.13)	(6.37)	(5.89)
Debt/GDP	-0.28	0.20	0.031	0.13	0.31
	(-0.79)	(0.74)	(0.11)	(1.49)	(1.35)
VIX	0.0068	0.014***	0.0089**	0.026***	0.014***
	(1.49)	(3.29)	(2.11)	(13.67)	(3.26)
Baa spread	0.32***	0.14**	0.14*	-0.055	-0.12
	(4.03)	(2.01)	(1.85)	(-1.25)	(-1.23)
$I_{1979-80}$	-0.68***	-0.41***	-0.52***	-0.18*	-0.39*
	(-5.03)	(-3.44)	(-3.25)	(-1.86)	(-1.82)
$I_{\geq 2000}$	0.47***	0.34***	0.23*	0.14**	0.30**
	(3.33)	(2.88)	(1.84)	(2.53)	(2.23)
Constant	0.32	0.16	0.39**	-0.15**	-0.48***
	(1.60)	(1.12)	(2.53)	(-2.50)	(-3.45)
$\bar{R}^2$	0.48	0.49	0.26	0.77	0.64
N	715	715	715	715	715

Overall, the analysis supports a positive comovement between inflation and convenience spreads during a significant part of the high-inflation period. It also highlights the data challenges of measuring the convenience benefits of long-term Treasuries and the distinct roles of monetary versus cost-push shocks as convenience drivers during this period. We use the model below to illustrate those distinct mechanisms.

4 Model of Convenience Yields and Inflation Drivers

This section formalizes the two competing channels of inflation and Treasury convenience – the "money channel" and the "liquidity demand channel" – in a New Keynesian model with demand for deposits and Treasuries that are substitutes in providing liquidity (Sidrauski (1967), Friedman (1969), Nagel (2016)). To keep the exposition simple, we build on a three-equation New Keynesian model of inflation and monetary policy (e.g., Galí (2008)). We focus on the new implications for the changing relationship between inflation and convenience yields.[28] We solve for log-linear dynamics for inflation, the output gap, interest rates, and, importantly, the convenience spread between illiquid loans and liquid bond rates in the model. We illustrate the effects of supply or cost-push and liquidity demand shocks on convenience yields and inflation via impulse responses.

There are three different short-term interest rates in our model, corresponding to different rates in practice. We use I_t^l to denote the interest rate on illiquid loans. In practice, households and firms cannot directly borrow and lend at T-bill rates, instead relying on less liquid bank loans, credit cards, student loans, mortgages etc. In order to capture lower liquidity inherent in these markets, we proxy for I_t^l using high-grade corporate bond yields or commercial paper rates in our empirical analysis. We denote I_t^b as the short-term interest rate on Treasuries, which are highly liquid. The liquid one-period interest rate I_t^b is also the instrument of monetary policy. Finally, I_t^d denotes the interest rate on liquid deposits, representing the interest rate that consumers and households can earn by depositing money with a bank, i.e., the most liquid and money-like asset in this model. Deposits can be interpreted as cash in the special case when the deposit rate equals zero. Each of these interest rates are available at longer maturities and we denote the n-period zero-coupon interest rates by $I_{n,t}^l$, $I_{n,t}^b$, $I_{n,t}^d$. Log interest rates are related to level interest rates via $i_t^l = \log\left(1 + I_t^l\right)$ etc. We use lowercase letters to denote logs throughout.

4.1 Model Setup

The model combines preferences over liquidity as in Krishnamurthy and Vissing-Jorgensen (2012) with a New Keynesian model of monetary policy.

[28] Bianchi et al. (2022b) also feature a convenience yield shock as a driver of business cycles and asset prices but do not focus on the changing inflation-convenience relationship, which we have documented in Section 3.

4.1.1 Preferences, Consumption, and Liquidity

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

$$Q_t = D_t + \frac{\lambda_t}{1-\lambda_t} B_t. \tag{4}$$

Here, $D_t = D_{1,t} + D_{2,t} + ...$ and $B_t = B_{1,t} + B_{2,t} + B_{3,t} + ...$ denote real balances of zero-coupon bank deposits and Treasury bonds of various maturities. While we consider the case of perfect substitutability between Treasury bonds and deposits for our calibration, this assumption is not crucial and the qualitative results are similar if Treasuries and deposits are not perfect substitutes.[29] The parameter λ_t controls the relative contribution of government bonds to the liquidity aggregate. A spike in λ_t can be interpreted as heightened uncertainty in the economy (Caballero and Krishnamurthy (2008)), tightened collateral constraints (Del Negro et al. (2017)), or a liquidity shock in the financial sector (Li (2024)), all of which would increase the preference for government debt relative to less liquid instruments. We refer to λ_t as Treasury liquidity.

We combine preferences for liquidity with habit preferences over consumption and labor/leisure. A representative household is assumed to maximize

$$E_0 \sum_{t=0}^{\infty} \beta^t U\left(C_t, Q_t, H_t, \Theta_t\right), \tag{5}$$

where

$$U\left(C_t, Q_t, H_t, \Theta_t\right) = \frac{\Theta_t \left(C_t - H_t\right)^{1-\gamma}}{1-\gamma} + \alpha \log Q_t. \tag{6}$$

Here, H_t denotes external consumption habit, which is a function of past aggregate consumption, i.e., consumers do not internalize the effects of their choices on future habit. We follow the highly

[29] The assumption of perfect substitutability makes the model as simple as possible while illustrating our main points. Appendix B.7 considers an extension to imperfect substitutability, and shows that in that case, liquidity demand shocks need to be interpreted more broadly as incorporating shocks to the quantity of Treasury bonds. We include the debt/GDP ratio as a measure of Treasury debt quantity, which tends to mostly move at a lower frequency than inflation, to control for this possibility throughout our empirical analysis.

nonlinear habit specification of Campbell et al. (2020) and Pflueger and Rinaldi (2022), which can jointly match the hump-shaped output response to monetary policy in the data (Fuhrer (2000), Christiano et al. (2005)) and volatile and time-varying risk premia in asset prices. For our main results, which rely on log-linearized macroeconomic dynamics and bond yields, the precise habit specification is not crucial. Details of the habit dynamics are in Appendix B. The shifter Θ_t represents a taste shock increasing the utility that households derive from consumption today and giving rise to a non-liquidity demand shock in the macroeconomic Euler equation.

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

$$\begin{aligned} & D_t + B_t - L_t + C_t \\ = \; & \frac{W_t}{P_t} N_t + \Pi_t + \frac{P_{t-1}}{P_t} D_{1,t-1}(1 + I^d_{t-1}) + \frac{P_{t-1}}{P_t} B_{1,t-1}(1 + I^b_{t-1}) - \frac{P_{t-1}}{P_t} L_{1,t-1}(1 + I^l_{t-1}) \\ & + \frac{P_{t-1}}{P_t} \sum_{i=2}^{\infty} \left(B_{i,t-1}(1 + R^b_{i,t}) + D_{i,t-1}(1 + R^d_{i,t}) - L_{i,t-1}(1 + R^l_{i,t}) \right] \Big), \end{aligned} \tag{7}$$

where P_t is the aggregate price level in the economy at time t, $L_{i,t}$ denotes the real quantity of zero-coupon loans of maturity i, $L_t = \sum_{i=1}^{\infty} L_{i,t}$ is the total real quantity of loans, Π_t is the sum of firm and bank profits remitted to the household sector, and $R^b_{i,t}$, $R^d_{i,t}$ and $R^l_{i,t}$ denote the nominal returns from buying an i-period bond, deposit or loan at time $t-1$ and selling it again at time t.

4.1.2 Deposits and Monetary Policy

We assume that the deposit rate I^d_t equals a fraction of the illiquid loan rate I^l_t:

$$I^d_t = \delta I^l_t, \tag{8}$$

where the constant δ is generally less than 1 to reflect banks' market power and ability to raise deposit rates less than one-for-one with market rates.[30] In the special case where $\delta = 0$, deposits in the model can be interpreted as cash that carries a liquid benefit but earns no interest. The Fed implements policy by affecting the liquid bond rate, I^b_t. Intuitively, the Fed is not allowed to operate directly through private loan markets, as deciding which borrower is creditworthy would be

[30] A long-standing and growing literature has documented the role of bank market power, see e.g. Barro and Santomero (1972); Startz (1979); Drechsler et al. (2017); Egan et al. (2022); Wang et al. (2022). The same functional form is endogenously generated in the model of Drechsler et al. (2017) and also assumed by Nagel (2016).

considered fiscal policy and hence outside the purview of the central bank. We assume that the target policy rate follows a log-linear Taylor (1993)-type rule. Theoretical and empirical research has documented the relevance of interest-rate smoothing and policy inertia (Woodford (2003b), Taylor and Williams (2010), Bernanke (2004), Stein and Sunderam (2018)).[31] Therefore, we include an inertial term in the policy rule:

$$i_t^b = (1-\rho^i)\left(\gamma^x x_t + \gamma^\pi \pi_t\right) + \rho^i i_{t-1}^b + v_{i,t}, \tag{9}$$

where the monetary policy shock $v_{i,t}$ is assumed to be i.i.d. with standard deviation of σ_i, i_t^b is the one-period log liquid bond rate, π_t is log inflation, and x_t is the log output gap, or the difference between log real output and its natural level in the absence of price-setting frictions. Specifying monetary policy in terms of an interest rate target is consistent with how monetary policy was conducted throughout almost all of our sample.[32] We assume that the central bank conducts monetary policy by choosing an interest rate target and then setting the amount of deposits to the implicit value satisfying households' money demand function to ensure the interest rate target is met.[33] The rule (9) says that the central bank raises the policy rate when the output gap or inflation are higher, though it does so gradually over time as captured by the parameter ρ^i. A higher inertia parameter ρ^i also implies that monetary policy may raise interest rates slowly in response to an increase in inflation, as the short-term response to inflation $(1-\rho^i)\gamma^\pi$ may be substantially smaller than the long-term response γ^π.

4.1.3 Firms

The supply side of the economy is standard and we relegate the details to Appendix B. Partially monopolistic firms are assumed to set product prices but can adjust their product prices only in some periods according to Calvo (1983) with inflation indexation (Christiano et al., 2005) and partly backward-looking price-setter inflation expectations (Fuhrer, 1997b). Following Pflueger

[31] Appendix B.8 shows that shocks to overall liquidity preference, α, cannot be interpreted as a shock to the overall demand for liquidity as long as assumption (8) is assumed to hold without shocks. However, disturbances to the deposit rate pass-through (8) would act analogously to λ_t. We therefore refer to λ_t broadly as liquidity demand shocks.

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

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

and Rinaldi (2022), we model households' labor-leisure trade-off from the consumption of home-produced goods in the manner of Greenwood et al. (1988), with an external home consumption habit exactly offsetting the equilibrium level of home goods. These assumptions ensure that we can abstract from the labor-leisure trade-off in solving for equilibrium asset prices. Since the model does not have real investment, the aggregate resource constraint implies that consumption equals output, $C_t = Y_t$. Supply shocks in the model are defined as cost-push shocks, and formally arise as a markup shock to firms' market power over the variety they produce.

4.1.4 Liquidity Shock Processes

In our baseline model, we assume that the liquidity demand shock λ_t follows a simple AR(1) process, similar to Anzoategui et al. (2019). We also consider an extension where Treasury liquidity is allowed to depend directly on inflation. Encompassing both of these cases, we assume

$$\begin{aligned} \lambda_t &= \bar{u} - b\pi_t + u_t \\ u_t &= \rho^{\lambda} u_{t-1} + v_{\lambda,t}. \end{aligned} \tag{10}$$

Here, ρ^{λ} is the persistence of Treasury liquidity, and $v_{\lambda,t}$ is i.i.d. normally distributed with volatilty σ_{λ}. Our baseline analysis sets $b = 0$, so inflation does not have a direct effect on Treasury liquidity. Our alternative model considers $b > 0$, which captures the intuition that low and stable inflation may be important for the safety and convenience of nominal Treasury bonds as discussed in the introduction. The steady-state Treasury bond liquidity weight equals $\bar{\lambda} = \bar{u} - b\bar{\pi}$, where $\bar{\pi}$ is steady-state log inflation.

To allow for a clear comparison between liquidity and taste shocks, we assume that the log taste shifter is also normally distributed and follows an AR(1) process. To save notation, we define non-liquidity demand shock as $\theta_t = \psi \left(\log(\Theta_t) - E_t \log(\Theta_{t+1})\right)$, with a process

$$\theta_t = \rho^{\theta} \theta_{t-1} + v_{\theta,t}, \tag{11}$$

where $v_{\theta,t}$ is i.i.d. normal with standard deviation σ_{θ}. Finally, the cost-push supply shock $v_{\pi,t}$, which will be formalized in equation (23), is assumed to be log-normal and i.i.d. with volatility σ_{π}.

4.2 Asset Pricing Euler Equations

The nominal consumption-based stochastic discount factor equals

$$M_{t+1}^{\$} = \beta \frac{U_c\left(C_{t+1}, Q_{t+1}, H_{t+1}, \Theta_{t+1}\right)}{U_c\left(C_t, Q_t, H_t, \Theta_t\right)} \frac{P_t}{P_{t+1}}. \tag{12}$$

This stochastic discount factor prices all nominal assets that have no special liquidity benefits, such as illiquid loans, giving the standard asset pricing Euler equation for the one-period loan rate

$$E_t\left[M_{t+1}^{\$}\left(1+I_t^l\right)\right] = 1. \tag{13}$$

In equilibrium, the representative household must be indifferent between marginally increasing Treasury bond holdings while decreasing consumption subject to the budget constraint (7), giving the Treasury bond Euler equation

$$E_t\left[M_{t+1}^{\$}\left(1+I_t^b\right)\right] = 1 - \underbrace{\frac{\frac{\alpha}{Q_t}\frac{\lambda_t}{1-\lambda_t}}{U_c\left(C_t, Q_t, H_t, \Theta_t\right)}}_{\zeta_t^b}. \tag{14}$$

Note that the Euler equation (14) for liquid Treasury bonds takes exactly the form as in models with a reduced-form Treasury convenience benefit ζ_t^b, which has proven useful in understanding global currency fluctuations (Jiang et al. (2021)) and international business cycles (Jiang et al. (2023), Kekre and Lenel (2021)). Bianchi et al. (2022b) introduce a similar wedge between the household and financial market Euler equations in their model of high-frequency market responses to monetary policy. We provide a new connection between this increasingly successful financial market shock and the real economy.

The analogous Euler equation for deposits is given by

$$E_t\left[M_{t+1}^{\$}\left(1+I_t^d\right)\right] = 1 - \underbrace{\frac{\frac{\alpha}{Q_t}}{U_c\left(C_t, Q_t, H_t, \Theta_t\right)}}_{\zeta_t^d}. \tag{15}$$

Equation (14) shows that Treasury bond convenience ζ_t^b increases with the liquidity weight of Treasury bonds, $\lambda_t/(1-\lambda_t)$, and also with the marginal consumption value of liquidity $\frac{\alpha/Q_t}{U_{c,t}}$.

Equation (15) shows that the convenience of deposits ζ_t^d increases with the liquidity value of deposits measured in marginal consumption units. Combining the first-order conditions for Treasury bonds (14) and deposits (15) with assumption (8) linking the deposit and loan rates, delivers the central equation for the Treasury bond convenience spread:

$$I_t^l - I_t^b = \frac{\lambda_t}{1-\lambda_t}(1-\delta)I_t^l. \tag{16}$$

To interpret equation (16) note that in the special case where deposits are simply liquid cash ($\delta = 0$) the nominal loan rate I_t^l is the cost of holding non-interest bearing cash, and $\lambda_t/(1-\lambda_t)$ is the liquidity value of Treasuries relative to cash.

4.3 Model Solution

We now describe the log-linear solution for macroeconomic dynamics and convenience yields in detail. Macroeconomic dynamics and the dynamics for the short-term and long-term convenience spread are derived by log-linearizing the model around the flexible-price steady-state $\bar{c} = \bar{y}$, $\bar{\pi}$, $\bar{i}^l$, $\bar{i}^b$, $\bar{\theta}$, and $\bar{\lambda}$. This solution is fast and allows for efficient GMM estimation of key model parameters. We thereby focus on the first-order effects of convenience and abstract from convenience risk premia, which are treated in complementary papers by Du et al. (2023) and Acharya and Laarits (2023). For ease of notation, we use c_t, y_t, π_t, i_t^l, i_t^b, and i_t^d to denote log deviations from these steady-state values.[34]

We start by log-linearizing the expression (16) for the illiquid loan rate

$$i_t^l \;=\; f^i i_t^b + f^\lambda \lambda_t, \tag{17}$$

where the log-linearization constant

$$f^i \;=\; \frac{1}{1-\frac{\bar{\lambda}}{1-\bar{\lambda}}(1-\delta)}\frac{1+\bar{I}^b}{1+\bar{I}^l}, \tag{18}$$

can be shown to be strictly greater than one ($f^i > 1$), provided that $0 < \bar{\lambda} < \frac{1}{2}$ and $\delta < 1$. The condition on $\bar{\lambda}$ means that Treasury bonds offer positive liquidity but less than liquidity than liquid

[34] In a future extension, we plan to implement a a highly non-linear solution for bond-stock betas.

deposits in steady-state. The condition on δ implies that the liquid deposit rate increases less than one-for-one with the illiquid loan rate. The constant f^{λ} is a log-linearization constant linking the magnitude of the liquidity demand shock λ_t to its impact on illiquid loan rates.

The intuition for $f^i > 1$ is that a higher policy rate i_t^b increases the cost of holding liquid deposits, which is the foregone interest of holding liquid deposits rather than less liquid interest-bearing loans. Since Treasury bonds act as substitutes in the liquidity aggregate (4), the cost of holding Treasuries, the convenience yield $i_t^l - i_t^b$, also increases. This is the money channel for short-term bonds. The log-linearized one-period convenience yield spread then equals

$$\underbrace{i_t^l - i_t^b}_{\text{Convenience yield}} = \underbrace{\left(f^i - 1\right) i_t^b}_{\text{Money channel}} + \underbrace{f^{\lambda}\lambda_t}_{\text{Liquidity demand}} . \tag{19}$$

The convenience spread for an n-period Treasury bond, to first order, equals the expected short-term Treasury convenience over the lifetime of the bond

$$\begin{aligned} i_{n,t}^l - i_{n,t}^b &= \frac{1}{n} E_t \left[\sum_{h=0}^{n-1} \left(i_{t+h}^l - i_{t+h}^b \right) \right], \\ &= \underbrace{\frac{f^i - 1}{n} E_t \sum_{h=1}^{n-1} \pi_{t+h+1}}_{\text{Expected inflation}} + \underbrace{\frac{f^{\lambda}}{n} E_t \sum_{h=0}^{n-1} \lambda_{t+h}}_{\text{Liquidity demand}} + \underbrace{\frac{f^i - 1}{n} i_t^b}_{\text{Policy rate}} + \underbrace{\frac{f^i - 1}{n} E_t \sum_{h=1}^{n-1} r_{t+h}^b}_{\text{Long-term real rate}} . \end{aligned} \tag{20}$$

Here, we used the Fisher equation $i_t^b = E_t[\pi_{t+1}] + r_t^b$ to replace the nominal rate i_t^b with the sum of expected inflation and the real rate r_t^b.

Expression (20) is the model analogue of the Aaa-Treasury spread at the center of our empirical analysis. It consists of four terms. The first term equals long-term inflation expectations over the lifetime of the bond, capturing the money channel for the long-term Treasury convenience. Intuitively, higher long-term inflation expectations raise the opportunity cost of holding money and deposits in the long term, making interest-paying long-term substitutes, such as long-term Treasuries, more attractive. The convenience yield that investors are willing to pay to hold long-term Treasuries is therefore predicted to increase with long-term inflation expectations.

The second term captures the liquidity demand channel, averaged over the expected lifetime of the bond. The third term is the current monetary policy rate. The last term reflects the expected long-term real rate, which in the limit is invariant to the monetary policy rate and inflation, since

equilibrium real rates are pinned down by productive capacity in the long term and outside the traditional scope of monetary policy (Fisher, 1930; Friedman, 1968). Overall, equation (20) shows that at the long end of the term structure the money channel drives a positive relationship between the persistent component of inflation and long-term Treasury bond convenience, controlling for liquidity demand shocks and the current nominal policy rate.

The representative household's intertemporal first-order condition takes the exactly log-linear form

$$x_t \;=\; \rho^x x_{t-1} + (1-\rho^x) E_t x_{t+1} - \psi \left(i_t^l - E_t \pi_{t+1} \right) + \theta_t, \tag{21}$$

where the parameters ρ^x and ψ are functions of the representative agent's preference parameters as in Campbell et al. (2020). The non-liquidity demand shock θ_t captures the typical New Keynesian demand shifter arising from preference or taste shocks, unrelated to Treasury liquidity (e.g., Galí, 2008).

Substituting the log-linearized convenience yield (17) into the macroeconomic Euler equation yields the Euler equation with liquidity:

$$x_t = \rho^x x_{t-1} + (1-\rho^x) E_t x_{t+1} - \psi(\underbrace{f^i i_t^b}_{\text{money channel}} - E_t \pi_{t+1}) - \underbrace{\psi f^\lambda \lambda_t}_{\text{liquidity demand}} + \theta_t. \tag{22}$$

For better interpretability, we scale the liquidity demand as $\xi_t \equiv -\psi f^\lambda u_t$ and define the corresponding volatility of $v_\xi = -\psi f^\lambda v_{\lambda,t}$ as $\sigma_\xi = \psi f^\lambda \sigma_\lambda$ (see equation (10)). In this way, the scaled liquidity demand shock volatility σ_ξ is directly comparable with non-liquidity demand shock volatility σ_θ. In what follows, we will use σ_ξ rather than the parameter σ_λ and simply refer to it as the liquidity-demand shock volatility (in output units).

Demand for liquid assets enters the macroeconomic Euler equation (22) in two ways. First, an increase in government bond convenience acts just like a negative aggregate demand shock via the $\psi f^\lambda \lambda_t$ term, reflecting the liquidity demand channel, and thereby provides an alternative microfoundation for macroeconomic demand shocks. When the Treasury bond convenience increases due to a shift in λ_t, households face a higher loan rate for a given Treasury bond rate, increasing their incentive to save and decreasing the incentive to consume this period. Second, the effect of the nominal rate on consumption and output is amplified by a factor f^i, which arises from the money channel of bond convenience. This amplification implies that a rise in inflation that is accompanied

by the same rise in the nominal rate is contractionary, as $f^i > 1$.[35]

Because utility is separable in consumption and liquidity, the standard log-linearization of the firm's optimal price setting problem gives the log-linearized Phillips curve

$$\pi_t \quad = \quad \rho^\pi \pi_{t-1} + (1-\rho^\pi) E_t \pi_{t+1} + \kappa x_t + v_{\pi,t}, \tag{23}$$

where ρ^π and κ are log-linearization constants, and the cost-push supply shock $v_{\pi,t}$ arises from deviations in the markup from its steady-state value (see Appendix B for details). The slope parameter κ reflects the rise in marginal costs of production when output is running above potential, leading firms to optimally raise prices.

We use Blanchard and Kahn (1980) algorithm to solve the equilibrium dynamics (22), (23) and (9) for an equilibrium of the form

$$Z_t = P^Z Z_{t-1} + Q^Z v_t. \tag{24}$$

where the state vector equals $Z_t = [x_t, \pi_t, i_t^b, \xi_t, \theta_t]$ and the vector of exogenous i.i.d. shocks is given by $v_t = [v_{\xi,t}, v_{\pi,t}, v_{i,t}, v_{\theta,t}]$.[36] Our quantification features a single non-explosive equilibrium of the form (24). The short- and long-term convenience spreads can then be solved by substituting (10) into the log-linear expressions (17) and (20).

4.4 Estimation Methodology

We set model parameters in two steps. First, we set most parameters using standard values in the literature, as shown in Appendix Table A2. We express all rates in annualized units, although the model implementation treats one quarter as the time unit so additional conversion is needed for the model. In particular, we set the pass-through of loan rates to deposit rates to $\delta = 0.34$, within the range of $1/3$ to $1/2$ suggested by Nagel (2016). The steady-state discount rate is set to $\beta = 0.98$ and inflation to $\bar{\Pi} = 2\%$ in annual units, so the steady-state illiquid loan rate equals

[35]This channel is complementary to Drechsler et al. (2023), who argue that an increase in inflation affects firms directly, thereby amplifying the increase in inflation. We abstract from that channel, which would tend to amplify but not change the sign of the inflation-convenience relationship. Binsbergen and Grotteria (2024) estimate the transmission of long-term Treasury bond yields around monetary policy announcements to long-term private borrowing rates and find a strong transmission mechanism.

[36]The surplus consumption ratio, s_t, would be a state variable for bond-stock betas, for which we plan to solve in future versions of the paper.

4.03% annualized. Setting $\bar{\lambda} = 0.14$ and substituting into equation (16) then implies a steady-state Treasury liquidity spread of $\bar{i}^l - \bar{i}^b = \frac{0.14}{1-0.14} \times (1 - 0.34) \times 4.03\% = 43$ bps, matching the average T-bill convenience yield spread in the data as reported in Table 1. We set the autocorrelation of the liquidity demand shock to 0.91, matching the quarterly AR(1) coefficient over our sample for the Aaa-Treasury spread, which is less likely to pick up on potentially transitory monetary policy shocks than T-bill convenience. Our baseline scenario does not allow inflation to directly affect convenience and thus sets $b = 0$ in equation (10). The parameters for the New Keynesian block of the model are set to standard values. Broadly, empirically plausible inertia parameters in the model help because they ensure that monetary policy, inflation, and the output gap are not perfectly correlated in the case of one single macroeconomic shock.

Second, for each period, we use data moments to estimate shock volatilities, as well as how Aaa-Treasury spread in the data maps to the Euler-equation term $i^l_{n,t} - i^b_{n,t}$ in the model. The Aaa-Treasury spread may over- or under-state the contribution of convenience yield to the Euler equation, given the difficulty in measuring long-term convenience. We therefore define the mapping as

$$i^l_{n,t} - i^b_{n,t} = \text{(scaling factor)} \times \text{(Aaa-Treasury spread)}, \tag{25}$$

where we use $n = 40$ quarters (i.e., 10 years) for long-term convenience yield. We estimate the scaling factor together with four types of volatilities, including the scaled liquidity demand shock volatility σ_ξ, cost-push supply shock volatility σ_π, monetary policy shock volatility σ_i, and non-liquidity demand shock volatility σ_θ.

Table 9 shows the model moments versus data. Comparing data moments with model moments, we find that the model does a good job matching both volatilities and correlations across all three periods. In particular, the signs of all correlations are consistent with the data. Moreover, the regression coefficient of the AAA-Treasury spread on inflation is also similar between the model and data across all three periods. Intuitively, the three moments on volatilities discipline the overall scale of shocks and the scaling factor, while the relative importance of different shocks is disciplined by the correlations. For example, in the model, a liquidity demand shock drives a negative convenience and inflation correlation, but non-liquidity demand shocks, monetary policy shocks, and supply shocks all drive a positive convenience and inflation correlation. Therefore, the correlation between convenience and inflation is highly informative about the presence of liquidity-demand shocks. Moreover, only a supply shock can generate a negative inflation-output gap correlation, so the distinct sign of inflation-output gap in the second period is highly informative on

the presence of supply shocks.

Table 10 report estimated parameters. We find that liquidity demand shocks σ_λ play a more important role in the first and last period, while inflation shocks only play a role in the middle period. The ratio between liquidity demand shock volatility and supply shock volatility clearly shows the dichotomy between the second period versus the other two periods, consistent with our main empirical findings. The estimated volatility monetary policy shocks σ_i plays no significant role in our estimation across all periods and is imprecisely estimated. Non-liquidity demand shocks are estimated to be much less volatile than liquidity demand shocks with tight standard errors. Overall, the inflation-convenience relationship is an informative moment for the drivers of the economy, and helps identify that liquidity demand shocks were a substantial driver of economic fluctuations during the Great Depression and again after 2000.

The estimated scaling factor shows that it is almost undistinguishable from one during the last period, suggesting that the Aaa-Treasury spread accurately measures the scale of Treasury convenience during this period. However, the Moody's Aaa-Treasury spread may understate the size of Treasury convenience in the Great Depression period, and overstate it in the middle period. In particular, its value of 0.43 estimated for the 1952–1999 period aligns with our findings in Section 3.4 based on alternative convenience proxies, which indicate that the long-term convenience spread may have been lower though still highly correlated with the Moody's Aaa-Treasury spread. A higher scaling factor during the pre-WWII period, as estimated by the model, is plausible if Aaa bonds had significant liquidity during the liquidity disruptions of the Great Depression, and the Aaa-Treasury spread hence understated true Treasury convenience during this period.

4.5 Impulse Responses

In this subsection, we illustrate impulse responses. We note that given all other parameters in the first step of our parameter setting, volatility parameters (estimated in the second step) do not affect state-variable dynamics due to the linearity structure in (24). As a result, across all three periods, model-implied impulse responses to the same underlying shock are the same across three periods, although the relative importance of different shocks and how convenience yield in the model maps to AAA-Treasury spread could vary across periods.

Impulse responses show why the inflation-convenience relationship is an informative moment for the volatility of the liquidity demand shock vs. non-liquidity demand shocks or supply shocks.

Table 9. Empirical and Model Moments. This table reports period-by-period estimation of our model comparing key model-implied moments with the data. Aaa-Treasury spread, inflation, and output gap are all expressed in percentage points.

		1920–1939		**1952–1999**		**2000–2020**	
		Data	Model	Data	Model	Data	Model
	Vol(Aaa-Treasury)	0.27	0.27	0.45	0.44	0.33	0.33
		(0.02)		(0.02)		(0.03)	
Volatilities	Vol(Inflation)	6.27	5.09	3.09	2.07	1.25	1.08
		(0.51)		(0.16)		(0.10)	
	Vol(Output Gap)	9.28	10.50	2.42	2.91	2.01	2.22
		(0.75)		(0.12)		(0.16)	
	Corr(Inflation, Output Gap)	0.34	0.14	-0.21	-0.56	0.53	0.14
		(0.10)		(0.07)		(0.08)	
Correlations	Corr(Aaa-Treasury, Inflation)	-0.37	-0.18	0.64	0.60	-0.29	-0.15
		(0.10)		(0.04)		(0.10)	
	Corr(Aaa-Treasury, Output Gap)	-0.20	-0.46	-0.38	-0.31	-0.23	-0.42
		(0.11)		(0.06)		(0.11)	
Reg Coefs	Aaa-Treasury $\sim \beta\cdot$ Inflation	-0.016	-0.010	0.094	0.129	-0.074	-0.045
		(0.009)		(0.022)		(0.043)	

Table 10. Estimated Model Parameters. This table reports period-by-period estimation of model parameters, including volatilities of four shocks in the model, and the scaling factor of how to map Aaa-Treasury spread into $i^l_{n,t} - i^b_{n,t}$ in equation (20). In each period, the second column report standard deviations of estimated parameters in bracket.

parameters (yearly values)	1920–1939		1952–1999		2000–2020	
σ_ξ (liquidity demand)	0.768	(0.033)	0.067	(0.002)	0.156	(0.007)
σ_π (cost-push)	0.000	(0.930)	0.413	(0.008)	0.000	(0.186)
σ_i (monetary policy)	0.010	(9.665)	0.001	(2.803)	0.019	(2.180)
σ_x (non-liquidity demand)	0.192	(0.007)	0.045	(0.001)	0.042	(0.001)
spread scaling	5.549	(0.202)	0.426	(0.010)	0.940	(0.035)

Figure 6 traces out the effect of a negative liquidity demand shock $v_{\lambda,t}$ for inflation, convenience spreads, the nominal policy rate i_t^b, and the output gap via impulse responses. The $v_{\lambda,t}$ shock is scaled so that its on-impact effect on the short-term convenience spread is 100 bps, implying a negative effect on aggregate demand in equation (22).

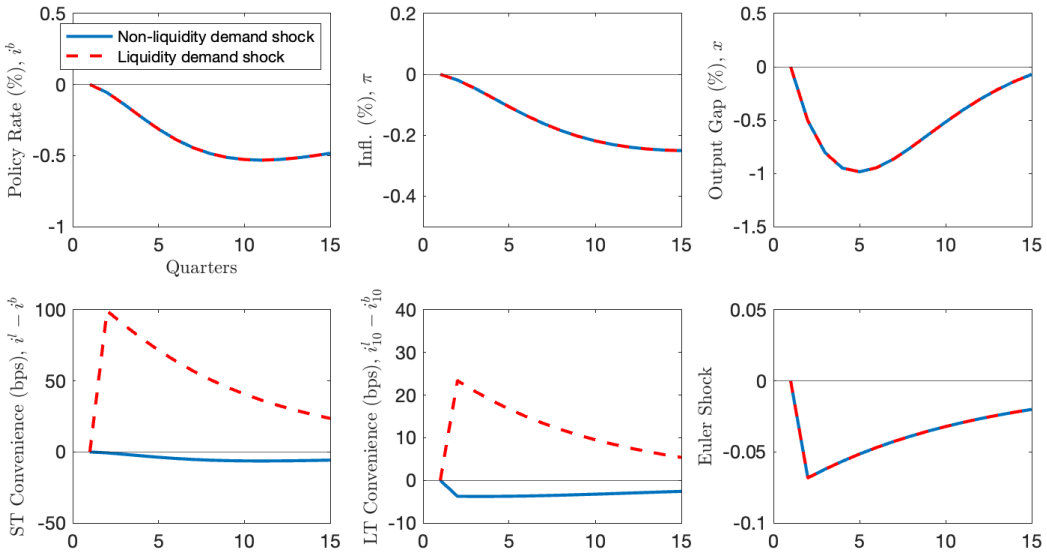
Figure 6 shows that the liquidity demand channel in the model ties together several of our empirical findings in Section 3. The liquidity demand channel moves inflation and Treasury convenience in opposite directions, explaining the negative empirical relationship between convenience and headline inflation in the pre-WWII and post-2000 periods. This behavior arises in the model because a positive convenience shock acts similarly to a negative demand shock to the Euler equation, decreasing the output gap, inflation, and the Treasury bond rate. The intuition follows directly from equation (17): households face a higher illiquid loan rate i_t^l at a given policy rate i_t^b, decreasing their demand to borrow and consume. Firms meet this weaker demand, reducing price pressure through the Phillips curve (23).

The macroeconomic responses to a liquidity demand shock are indistinguishable from an equally-sized traditional aggregate demand shock in a model without convenience (top panels in Figure 6), showing the analogy between traditional demand-driven inflation and inflation movements due to the liquidity demand channel. However, the inflation-convenience comovement reveals a crucial distinction between the two types of demand shock. The liquidity demand channel implies that higher convenience spreads should predict lower demand-driven inflation, consistent with the empirical lead-lag relationship we find. In contrast, the non-liquidity demand channel drives a positive and overall much weaker relationship between inflation and convenience. The comovement between inflation-convenience is therefore a critical new moment allowing us to distinguish the nature of the demand shocks driving economic fluctuations.

Figure 7 shows that the money channel explains the empirical evidence for the 1952–1999 period, as cost-push supply shocks move inflation and convenience – especially long-term convenience – in the same direction. The impulse response for the long-term convenience spread follows most closely the inflation response, while the short-term convenience spread follows most closely the policy rate response to a cost-push supply shock. The intuition is that long-term Treasury convenience moves with long-term inflation expectations, even controlling for the current policy rate as in equation (20).[37] Monetary policy is slightly amplified in the presence of Treasury

[37] As shown by equation (20), we do not need restrictions on the monetary policy rule coefficients or active monetary policy for the money channel. If inflation was subject to sunspot fluctuations and the inflation coefficient in the

Figure 6. Baseline model responses to liquidity shocks. This figure shows impulse responses to a liquidity demand shock to $v_{\lambda,t}$ versus a non-liquidity demand shock to $v_{x,t}$ for our baseline model. The shock is scaled so that the liquidity demand shock corresponds to a 100 bps increase in the convenience yield spread $i_t^l - i_t^b$, and the non-liquidity demand shock has the same magnitude. Responses for inflation, π and the Treasury rate i^b are in annualized percent units. The response for the convenience spread $i^l - i^b$ is in annualized basis points units. The response for the output gap x is in percent units. The Euler shock shows ξ_t and ζ_t, i.e., the persistent components of liquidity demand and non-liquidity demand shocks. Quarters are shown on the x-axis.

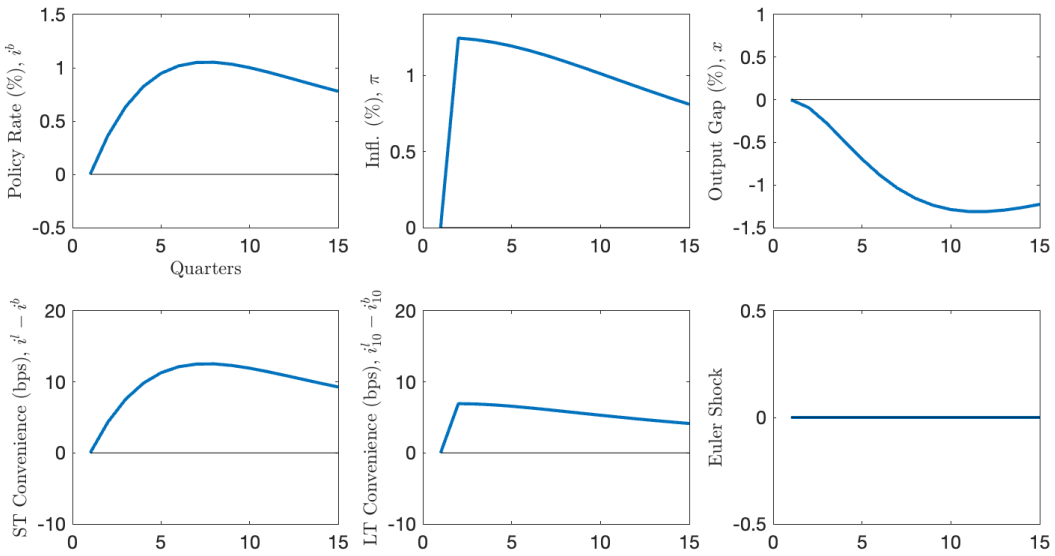


convenience, deepening the recession but mitigating inflation in response to a positive cost-push supply shock compared to a model without Treasury convenience. Overall, the money channel ties together our empirical evidence for a positive long-term convenience-inflation correlation in the second half of the 20th century and during periods when supply-shock inflation is dominant according to bond-stock betas. As in the data, this positive relationship in the model is not explained by the contemporaneous nominal policy rate. The money channel in the model also predicts that this positive relationship should be particularly pronounced for the persistent components of inflation – as it is long-term inflation expectations that enter in equation (20) – and for the supply

monetary policy rule was less than one during the 1970s, as argued by Clarida et al. (2000), this would further drive a wedge between the short-term nominal rate and long-term expectations for inflation. Such sunspot fluctuations would, therefore, act similarly to cost-push supply shocks for our purposes, generating a close relationship between long-term inflation expectations and long-term Treasury convenience that is not explained by the short-term policy rate.

component of inflation.[38]

Figure 7. Baseline model impulse responses to cost-push supply shock. This figure shows impulse responses to a cost-push supply shock for our baseline model. The shock is a positive 100 bps shock to the Phillips curve. Responses for inflation, π and the Treasury rate i^b are in annualized percent units. The response for the convenience spread $i^l - i^b$ is in annualized basis points units. The response for the output gap x is in percent units. Quarters are shown on the x-axis.



The model also delivers predictions for the effect of monetary policy shocks on convenience (see Appendix Figure A5), showing that monetary policy shocks in the Taylor rule cannot explain our main empirical findings, an implication consistent with the empirical evidence of Binsbergen and Grotteria (2024). Upon a monetary tightening shock, the long-term convenience declines because a surprise tightening in the model lowers long-term inflation expectations while raising the short-term nominal policy rate. This depresses the long-term convenience through equation (20), despite increasing short-term convenience. Short-term convenience has a shorter maturity and is thus more closely linked to the current monetary policy rate and less to long-run inflation expec-

[38]While the impulse responses in Figure 7 do not exhibit a clear lead-lag pattern between inflation and the long-term convenience spread because expectations are perfectly forward-looking, a slight model extension with sluggish expectations formation (Malmendier and Nagel (2016)) could plausibly explain the empirical finding that supply-driven inflation tends to predict higher long-term convenience spreads. We do not pursue this extension here in the interest of parsimony.

tations. This mechanism speaks to the evidence in Section 3.4 that the relationship of inflation with long-term convenience changed in the Volcker episode of aggressive monetary tightening compared to the earlier part of the 1970s, while that of short-term convenience remained stable.

4.6 Alternative Model Impulse Responses with Direct Inflation-Convenience Link

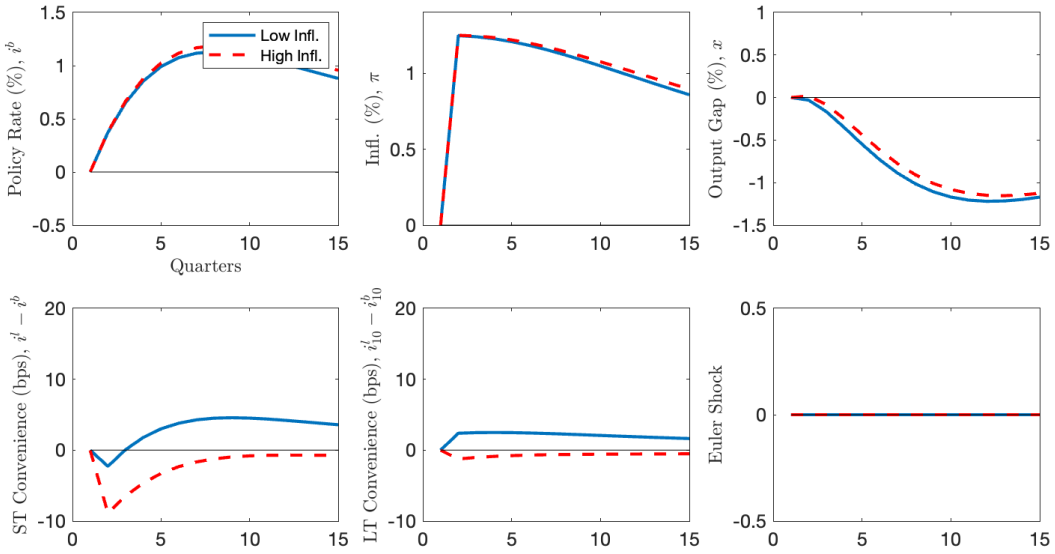
Finally, we investigate the implications of an alternative model allowing for a direct inflation-Treasury liquidity link, and show that it cannot explain our empirical findings. Specifically, by setting $b > 0$ in equation (10), we assume that inflation can directly undermine the liquidity benefits of Treasuries. To illustrate how such a direct inflation-liquidity link would have plausibly altered the relationship between inflation and long-term convenience spreads in different subperiods, we consider two levels for steady-state inflation in this alternative model. The high-inflation steady state considers $\bar{\Pi}^{high} = 4.01\%$ matching average inflation in the data for the 1952-1999 period, while the low-inflation steady state uses $\bar{\Pi}^{low} = 0.65\%$ to match inflation averaged across the pre-WWII and post-2000 periods in the data. We calibrate $b = 0.014$. This value means that at an annualized steady-state inflation rate of 10% the liquidity value of Treasury bonds would be eliminated because $\bar{u} - 10 \times b = 0.14 - 10 \times 0.014 = 0$. All other parameters are held constant across the two quantifications and are as in Table A2.

Figure 8 shows that this alternative model implies a *more negative* convenience-inflation relationship when steady-state inflation is high than when it is low. This is in stark contrast to the data where we found a *more positive* inflation-convenience relationship during the high-inflation 1970s and 1980s than during the low-inflation pre-WWII and 2000s periods in the data. Figure 8 focuses on the impulse responses to a cost-push supply shock, as impulse responses to the other two shocks are very similar to our baseline model.[39] To the extent that higher inflation reduces the attractiveness of Treasury bonds as a convenient asset, one might expect a decrease in convenience spreads following an inflationary cost-push supply shock. The bottom panels of Figure 8 show that this pattern emerges only for the high-inflation steady state. In the data, the inflation-convenience relationship is instead more positive during the high-inflation 1970s and 1980s.

Why does the alternative model imply a more negative relationship between the Treasury convenience yield and inflation when inflation is high? The intuition is simply that with higher steady-

[39] We show that these implications are robust to alternative parameter values in the Appendix.

Figure 8. Alternative model impulse responses to cost-push supply shock. This figure shows impulse responses to a cost-push supply shock for the model with a direct inflation-convenience link ($b = 0.014$ in equation (10)). The high-inflation equilibrium assumes $\bar{\Pi}^{high} = 4.01\%$ and the low-inflation equilibrium assumes $\bar{\Pi}^{low} = 0.65\%$. The shock is a positive 100 bps shock to the Phillips curve. Responses for inflation, π and the Treasury rate i^b are in annualized percent units. The response for the convenience spread $i^l - i^b$ is in annualized basis points units. The response for the output gap x is in percent units. Quarters are shown on the x-axis.



state inflation, Treasuries have less steady-state convenience because $\bar{\lambda} = \bar{u} - b\bar{\pi}$ is lower, implying that the convenience response to inflation is weaker. Therefore, the money channel is weaker. In this case, the response of the Treasury convenience yield is dominated by the direct effect on λ_t. By contrast, when steady-state inflation is low, the money channel is more important and a cost-push supply shock has a more pronounced positive effect on the Treasury convenience yield through the money channel.

Overall, the model shows that low-frequency shifts between the liquidity demand channel and the money channel of Treasury convenience can explain the patterns documented in our empirical analysis for aggregate inflation and inflation components, while a direct link between inflation and Treasury convenience cannot.

5 Evidence from the Post-COVID Period

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

Figure 9 superimposes various measures of long-term convenience against the 10-year inflation swap rate during the period from 2018 to 2023. While the correlation between Treasury convenience and market-implied inflation expectations was strongly negative during and immediately following the initial COVID-19 shock in March 2020, it became slightly positive just as inflationary pressures emerged in 2021. This fact is robust to using various convenience yields available in the post-2000 sample, including the agency-Treasury bond spread, agency-Treasury STRIPs spread, and the Refcorp-Treasury STRIPs spread. These measures offer some advantages over the Aaa-Treasury spread for capturing the long-term Treasury convenience because agency bonds and Refcorp bonds are guaranteed by the U.S. government and the STRIP-based proxies have exactly matched maturities. However, they are only available over a shorter sample.[40] Even though the levels of convenience differ, Figure 9 shows that their negative comovement with the inflation swap rate largely disappears from 2021 onward across all proxies.

In Table 11, we analyze the relation between convenience yield and inflation for the recent period more formally by projecting spreads on inflation and inflation expectations. In all regressions, we control for the federal funds rate to capture changes in the monetary policy stance. We supplement the analysis with direct measures of inflation expectations (the one-quarter-ahead CPI inflation forecast from the Blue Chip Financial Forecasts, available monthly) due to concerns about inflation risk premia in inflation swaps. The common pattern across the different measures of inflation and convenience yields is that the inflation-convenience correlation is negative before 2021 but positive afterward. Importantly, in line with the earlier findings over the long sample, we recover the relationship between inflation and long-term convenience while controlling for the monetary policy stance.

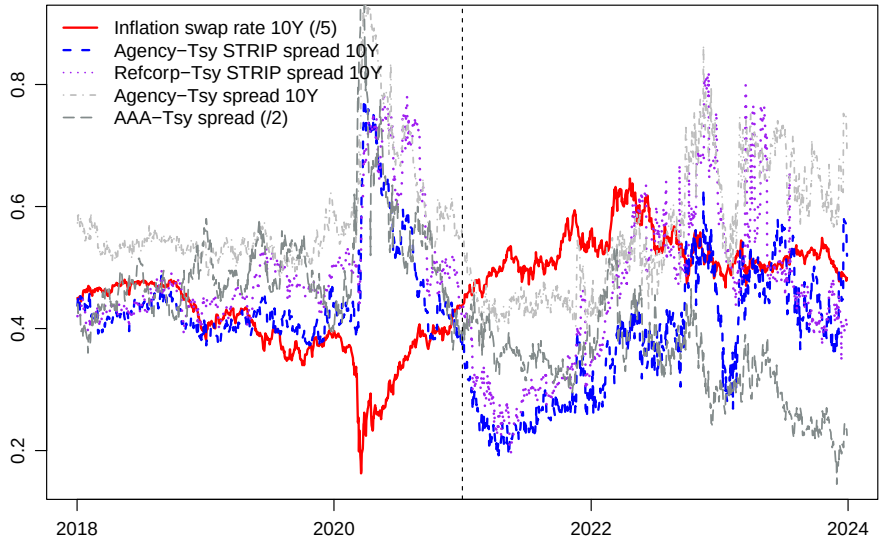
The change in correlation is consistent with our model mechanism. The year 2020 witnessed a sharp drop in inflation due to the initial negative demand shock induced by the COVID-19 pandemic and then a gradual recovery from that shock. While the initial demand shock did not origi-

[40] For our main analysis over the long sample, we necessarily rely on Aaa-Treasury spread where other measures are not consistently available. However, among the available proxies in more recent years, the agency-Treasury STRIP spread should most closely reflect the notion of long-term convenience.

Table 11. Post-COVID inflation shocks and long-term convenience spreads. The table presents estimates that regress various measures of long-term Treasury convenience onto realized inflation, expected inflation, and inflation swap rates. 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. Realized inflation is the 12-month CPI inflation rate. Expected inflation is the forecast of one-quarter-ahead inflation from the Blue Chip Financial Forecasts. Agency-Tsy spread is the yield spread between 10-year agency bonds and 10-year Treasury notes. Agency-Tsy STRIP spread is the yield spread between 10-year agency STRIPs (zero-coupon) and 10-year Treasury STRIPs (zero-coupon). Refcorp represents 10-year Refcorp STRIPs (zero-coupon). All regressions cover a 36-month period ($N = 36$). Newey-West t-statistics with 12 lags are reported in parentheses. Constant terms are included in regressions but not reported for conciseness. * p <0.10, ** p <0.05, *** p <0.01.

	Aaa-Tsy		Agency-Tsy		Agency-Tsy STRIP		Refcorp-Tsy STRIP	
	(1) pre	(2) post	(3) pre	(4) post	(5) pre	(6) post	(7) pre	(8) post
Panel A: 12-Month Inflation, $N = 36$								
Inflation	-0.15***	0.016	-0.047**	0.012***	-0.041	0.015***	-0.075***	0.043***
	(-4.15)	(1.62)	(-2.19)	(7.27)	(-1.41)	(4.97)	(-4.19)	(5.30)
FFR	-0.0055	-0.037***	-0.051*	0.047***	-0.034	0.045***	-0.037*	0.043***
	(-0.15)	(-4.69)	(-2.01)	(20.75)	(-1.36)	(14.86)	(-1.80)	(5.94)
$\bar{R}^2$	0.28	0.58	0.47	0.80	0.33	0.72	0.53	0.60
Panel B: Blue Chip Inflation Expectations, $N = 36$								
$\mathbb{E}_{1q}^{\mathrm{BC}}$(CPI infl)	-0.47***	0.060***	-0.16**	0.037***	-0.15*	0.049***	-0.28***	0.13***
	(-5.08)	(3.10)	(-2.11)	(3.59)	(-1.85)	(4.51)	(-4.39)	(8.93)
FFR	0.058	-0.045***	-0.028	0.042***	-0.011	0.037***	0.0074	0.023***
	(1.24)	(-6.97)	(-0.80)	(14.96)	(-0.30)	(12.60)	(0.27)	(3.38)
$\bar{R}^2$	0.26	0.61	0.47	0.79	0.34	0.73	0.57	0.61
Panel C: Zero-Coupon 10YR Inflation Swap, $N = 36$								
Inflation Swap 10Y	-0.37***	0.080	-0.14***	0.10***	-0.12**	0.13***	-0.13***	0.27***
	(-6.02)	(1.52)	(-2.92)	(8.87)	(-2.20)	(9.20)	(-6.78)	(9.44)
FFR	-0.0052	-0.036***	-0.045***	0.051***	-0.030*	0.050***	-0.048***	0.050***
	(-0.31)	(-3.86)	(-2.80)	(15.00)	(-1.70)	(13.46)	(-2.99)	(4.02)
$\bar{R}^2$	0.69	0.54	0.71	0.81	0.56	0.74	0.61	0.46

Figure 9. Daily time series of various measures of long-term convenience and 10-year breakeven inflation after 2018. We plot various Treasury convenience yield measures and 10-year inflation swap rates from Bloomberg (USSWIT10) 2018:01 to 2023:12. All convenience yield measures are maturity-matched spreads of 10-year maturity. In particular, agency-Tsy STRIP spread is the yield spread between 10-year agency STRIPs (zero coupon) and 10-year Treasury STRIPs (zero coupon). Refcorp represents 10-year Refcorp STRIPs (zero coupon). Agency-Tsy spread is the yield spread between 10-year agency bonds and 10-year Treasury notes. The vertical dashed line marks the end of 2020.



nate as primarily a liquidity or banking crisis, the pandemic shock was associated with heightened liquidity demand. With widespread lock-downs, households faced sudden income declines, necessitating immediate access to cash or liquid savings to cover essential expenses. Accordingly, different convenience yields spiked at the onset of the pandemic[41] and then recovered during the rest of 2020. We interpret the negative comovement as reflecting the liquidity demand-driven convenience-inflation relationship. Starting from 2021, inflation shocks pushed the inflation rate above 2%, at the same time when the inflation-convenience correlation turns from negative to slightly positive. This change is consistent with the money mechanism in our model, whereby supply shocks generate a positive comovement between long-term convenience and inflation.

Throughout our analysis, we emphasize the distinct impacts of liquidity demand and supply

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

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

Several historical fiscal episodes are pertinent in this context. The fiscal dynamics underlying the Great Inflation from the 1960s to the early 1980s have long been debated. If fiscal policy contributed to the rise in trend inflation (e.g., Bianchi and Ilut, 2017), it may have also driven the positive comovement between inflation and convenience through the money channel, as highlighted in our model and supported by empirical evidence.

More recently, the COVID-19 pandemic prompted a substantial fiscal expansion. Early pandemic fiscal stimulus (e.g., the March 2020 CARES Act) helped to reverse the initial liquidity demand shock and facilitated economic recovery. This demand-driven interpretation aligns with the observed negative relationship between inflation and convenience documented above. Conversely, subsequent fiscal measures, implemented when the economy was nearing full potential—especially the American Rescue Plan Act in March 2021—propelled short-term inflation (Bianchi et al., 2023) and might have simultaneously reduced the Treasury convenience. This direct fiscal channel could partially explain the negative comovement between inflation and convenience in the first half of 2021. However, the subsequent shift to a positive comovement suggests a significant influence of the money channel due to changing inflation expectations. The money channel might have been driven by narrowly constructed supply shocks or by shifters to the Phillips curve more broadly, which could also include fiscal factors. Understanding the interplay between fiscal policy, money, and liquidity demand channels is an important area for future research.

6 Conclusion

This paper argues that two competing mechanisms driving Treasury bond convenience – the "money channel" and the "liquidity demand channel" – dominated over distinct historical periods, leading to secular shifts in the comovement between Treasury convenience and inflation. Using a century of data, we show during a large part of the inflationary 1970s and 1980s, higher inflation coincided with higher Treasury convenience. By contrast, high Treasury convenience was associated with low inflation during the first half of the 20th century and again during the post-2000 period.

An empirical decomposition of inflation into components, lead-lag patterns, and a sample split by financial market indicators for supply vs. demand shocks paint a coherent picture whereby the positive correlation during the 1970s and 1980s was due to persistent inflation driven by aggregate cost-push supply shocks, while the negative correlation during the post-2000 period was associated with aggregate liquidity demand shocks to the economy.

We explain these findings within a New Keynesian model that embeds the money channel of Treasury convenience along with liquidity demand shocks. The model predicts that an inflationary cost-push supply shock raises long-term expectations of future inflation, the long-term opportunity cost of holding money, and hence, the price of holding long-term convenient assets, including Treasuries. Therefore, the money channel explains the positive convenience-inflation relationship that emerged during the 1970s. The model also predicts that a higher liquidity value of Treasuries increases the incentive to save and reduces consumption, lowering demand and hence inflation. A negative inflation-convenience relationship ensues from the liquidity demand channel, explaining the experience of the early 20th century and the 2000s.

Our results underscore that the relationship between Treasury convenience and inflation is informative about the sources of macroeconomic fluctuations – particularly, the nature of demand shocks. While inflationary cost-push shocks can induce a positive relationship between inflation and Treasury convenience, demand shocks for Treasury convenience act as significant macroeconomic demand shocks. Distinct from non-liquidity demand shocks, they drive a negative convenience-inflation relationship as observed during the Great Depression and the post-2000 period.

References

Joseph Abadi, Markus Brunnermeier, and Yann Koby. The reversal interest rate. *American Economic Review*, 113(8):2084–2120, 2023.

Viral V. Acharya and Toomas Laarits. When do Treasuries earn the convenience yield? - A hedging perspective. NBER Working Paper wp31863, 2023.

Diego Anzoategui, Diego Comin, Mark Gertler, and Joseba Martinez. Endogenous technology adoption and R&D as sources of business cycle persistence. *American Economic Journal: Macroeconomics*, 11(3): 67–110, 2019.

Andrew Atkeson and Lee E Ohanian. Are Phillips curves useful for forecasting inflation? *Federal Reserve Bank of Minneapolis Quarterly Review*, 25(1):2–11, 2001.

Robert J Barro and Anthony M Santomero. Household money holdings and the demand deposit rate. *Journal of Money, Credit and Banking*, 4(2):397–413, 1972.

Christiane Baumeister, Dimitris Korobilis, and Thomas K Lee. Energy markets and global economic conditions. *Review of Economics and Statistics*, 104(4):828–844, 2022.

Geert Bekaert, Eric Engstrom, and Andrey Ermolov. Macro risks and the term structure of interest rates. *Journal of Financial Economics*, (141):479–504, 2021.

Ben Bernanke. Gradualism, 2004. URL `https://www.federalreserve.gov/boarddocs/speeches/2004/200405202/default.htm`. Remarks at an economics luncheon co-sponsored by the Federal Reserve Bank of San Francisco and the University of Washington, Seattle, Washington, on May 20.

Ben S Bernanke. Nonmonetary effects of the financial crisis in the propagation of the great depression. *The American Economic Review*, 73(3):257–276, 1983.

Harjoat S Bhamra, Christian Dorion, Alexandre Jeanneret, and Michael Weber. High inflation: Low default risk and low equity valuations. *Review of Financial Studies*, 36(3):1192–1252, 2023.

Francesco Bianchi and Cosmin Ilut. Monetary/fiscal policy mix and agents' beliefs. *Review of Economic Dynamics*, 26:113–139, 2017.

Francesco Bianchi, Sydney C Ludvigson, and Sai Ma. Monetary-based asset pricing: A mixed-frequency structural approach. NBER Working Paper wp30072, 2022b.

Francesco Bianchi, Renato Faccini, and Leonardo Melosi. A fiscal theory of persistent inflation. *The Quarterly Journal of Economics*, 138(4):2127–2179, 2023.

Jules H van Binsbergen and Marco Grotteria. Monetary policy wedges and the long-term liabilities of households and firms. *National Bureau of Economic Research wp32137*, 2024.

Jules H van Binsbergen, William F Diamond, and Marco Grotteria. Risk-free interest rates. *Journal of Financial Economics*, 143(1):1–29, 2022.

Olivier Jean Blanchard and Charles M Kahn. The solution of linear difference models under rational expectations. *Econometrica*, 48(5):1305–1311, 1980.

Scott Brown and Eric Powers. The life cycle of make-whole call provisions. *Journal of Corporate Finance*, 65:101772, 2020.

Markus K Brunnermeier, Sebastian A Merkel, and Yuliy Sannikov. Debt as safe asset. NBER Working Paper wp29626, 2022b.

Markus K Brunnermeier, Sergio Correia, Stephan Luck, Emil Verner, and Tom Zimmermann. The debt-inflation channel of the German hyperinflation. NBER Working Paper wp31298, 2023.

Ricardo J Caballero and Arvind Krishnamurthy. Collective risk management in a flight to quality episode. *Journal of Finance*, 63(5):2195–2230, 2008.

Ricardo J Caballero and Alp Simsek. A risk-centric model of demand recessions and speculation. *Quarterly Journal of Economics*, 135(3):1493–1566, 2020.

Ricardo J Caballero and Alp Simsek. A monetary policy asset pricing model. 2022b.

Phillip Cagan. The demand for currency relative to the total money supply. *Journal of Political Economy*, 66(4):303–328, 1958.

Guillermo A Calvo. Staggered prices in a utility-maximizing framework. *Journal of Monetary Economics*, 12(3):383–398, 1983.

John Y Campbell and John H Cochrane. By force of habit: A consumption-based explanation of aggregate stock market behavior. *Journal of Political Economy*, 107:205–251, 1999.

John Y Campbell, Andrew WenChuan Lo, and Archie Craig MacKinlay. *The Econometrics of Financial Markets*. Princeton University Press Princeton, NJ, 1997.

John Y Campbell, Adi Sunderam, and Luis M Viceira. Inflation bets or deflation hedges? The changing risks of nominal bonds. *Critical Finance Review*, 6(2):263–301, 2017.

John Y Campbell, Carolin Pflueger, and Luis M Viceira. Macroeconomic drivers of bond and equity risks. *Journal of Political Economy*, 128(8):3148–3185, 2020.

Joshua CC Chan. Specification tests for time-varying parameter models with stochastic volatility. *Econometric Reviews*, 37(8):807–823, 2018.

Mikhail Chernov and Philippe Mueller. The term structure of inflation expectations. *Journal of financial economics*, 106(2):367–394, 2012.

Lawrence H Christiano, Martin Eichenbaum, and Charles L Evans. Nominal rigidities and the dynamic effects of a shock to monetary policy. *Journal of Political Economy*, 113(1):1–45, 2005.

Anna Cieslak and Carolin Pflueger. Inflation and asset returns. *Annual Review of Financial Economics*, 15, 2023.

Richard Clarida, Jordi Galí, and Mark Gertler. The science of monetary policy: A New Keynesian perspective. *Journal of Economic Literature*, 37(4):1661–1707, 1999.

Richard Clarida, Jordi Galí, and Mark Gertler. Monetary policy rules and macroeconomic stability: Evidence and some theory. *Quarterly Journal of Economics*, 115(1):147–180, 2000.

John H Cochrane. Long-term debt and optimal policy in the fiscal theory of the price level. *Econometrica*, 69(1):69–116, 2001.

Alexandre Corhay, Thilo Kind, Howard Kung, and Gonzalo Morales. Discount rates, debt maturity, and the fiscal theory. *Journal of Finance*, 78(6):3561–3620, 2023.

Vasco Curdia and Michael Woodford. Credit spreads and monetary policy. *Journal of Money, Credit and Banking*, 42:3–35, 2010.

Adrien d'Avernas, Quentin Vandeweyer, and Damon Petersen. The central bank's balance sheet and treasury market disruptions. *Available at SSRN 4826919*, 2024.

Marco Del Negro, Gauti Eggertsson, Andrea Ferrero, and Nobuhiro Kiyotaki. The great escape? A quantitative evaluation of the Fed's liquidity facilities. *American Economic Review*, 107(3):824–857, 2017.

Sebastian Di Tella. Risk premia and the real effects of money. *American Economic Review*, 110(7):1995–2040, 2020.

William Diamond and Peter Van Tassel. Risk-free rates and convenience yields around the world. Working Paper, University of Pennsylvania, 2023.

Itamar Drechsler, Alexi Savov, and Philipp Schnabl. The deposits channel of monetary policy. *Quarterly Journal of Economics*, 132(4):1819–1876, 2017.

Itamar Drechsler, Alexi Savov, and Philipp Schnabl. A model of monetary policy and risk premia. *Journal*

of Finance, 73(1):317–373, 2018.

Itamar Drechsler, Alexi Savov, and Philipp Schnabl. Credit crunches and the great stagflation. Working Paper, New York University and University of Pennsylvania, 2023.

Wenxin Du, Joanne Im, and Jesse Schreger. The US Treasury premium. *Journal of International Economics*, 112:167–181, 2018a.

Wenxin Du, Alexander Tepper, and Adrien Verdelhan. Deviations from covered interest rate parity. *Journal of Finance*, 73(3):915–957, 2018b.

Wenxin Du, Benjamin M Hébert, and Wenhao Li. Intermediary balance sheets and the Treasury yield curve. *Journal of Financial Economics*, 150(3), 2023.

Gregory R Duffee. The relation between Treasury yields and corporate bond yield spreads. *Journal of Finance*, 53(6):2225–2241, 1998.

Darrell Duffie. Resilience redux in the us treasury market. In *Jackson Hole Symposium*. Graduate School of Business, Stanford University, 2023.

Darrell Duffie, Nicolae Gârleanu, and Lasse Heje Pedersen. Valuation in over-the-counter markets. *Review of Financial Studies*, 20(6):1865–1900, 2007.

Mark Egan, Stefan Lewellen, and Adi Sunderam. The cross-section of bank value. *The Review of Financial Studies*, 35(5):2101–2143, 2022.

Kenneth Emery, Sharon Ou, Jennifer Tennant, Adriana Matos, and Richard Cantor. Corporate default and recovery rates, 1920–2008. *Moody's Investors Service, Special Comment, February*, 2009.

Charles Engel and Steve Pak Yeung Wu. Liquidity and exchange rates: An empirical investigation. *Review of Economic Studies*, 90(5):2395–2438, 2023.

Andrey Ermolov. Time-varying risk of nominal bonds: How important are macroeconomic shocks? *Journal of Financial Economics*, 145:1–28, 2022.

Irving Fisher. *The Theory of Interest*. London: Macmillan, 1930. URL `https://books.google.com/books?id=Dhf9MAEACAAJ`.

Milton Friedman. The role of monetary policy. *American Economic Review*, 58(1), 1968.

Milton Friedman. *The optimum quantity of money*. Aldine, Chicago, 1969.

Milton Friedman and Anna J Schwartz. The effect of term structure of interest rates on the demand for money in the united states. *Journal of Political economy*, 90(1):201–212, 1982.

Milton Friedman and Anna Jacobson Schwartz. *A monetary history of the United States, 1867-1960*. Princeton University Press, 1963.

Zhiyu Fu, Jian Li, and Yinxi Xie. Convenience yield, inflation expectations, and public debt growth. Working Paper, Columbia University and University of Chicago, 2023.

Jeffrey C Fuhrer. The (un)importance of forward-looking behavior in price specifications. *Journal of Money, Credit and Banking*, 29(3):338–350, 1997a.

Jeffrey C Fuhrer. The (un) importance of forward-looking behavior in price specifications. *Journal of Money, Credit, and Banking*, pages 338–350, 1997b.

Jeffrey C Fuhrer. Habit formation in consumption and its implications for monetary-policy models. *American Economic Review*, 90(3):367–390, 2000.

Masao Fukui, Emi Nakamura, and Jón Steinsson. The macroeconomic consequences of exchange rate depreciations. NBER Working Paper wp31279, 2023.

Jordi Galí. *Monetary policy, inflation, and the business cycle*. Princeton University Press, 2008.

Mark Gertler and Peter Karadi. A model of unconventional monetary policy. *Journal of Monetary Economics*, 58(1):17–34, 2011a.

Simon Gilchrist and Egon Zakrajšek. Credit spreads and business cycle fluctuations. *American Economic Review*, 102(4):1692–1720, 2012.

Jeremy Greenwood, Zvi Hercowitz, and Gregory W Huffman. Investment, capacity utilization, and the real business cycle. *American Economic Review*, pages 402–417, 1988.

Refet S Gürkaynak, Brian Sack, and Jonathan H Wright. The US Treasury yield curve: 1961 to the present. *Journal of Monetary Economics*, 54(8):2291–2304, 2007.

Jonathan S Hartley and Urban J Jermann. The pricing of US treasury floating rate notes. *Journal of Financial Economics*, 155:103833, 2024.

Zhiguo He, Stefan Nagel, and Zhaogang Song. Treasury inconvenience yields during the covid-19 crisis. *Journal of Financial Economics*, 143(1):57–79, 2022.

Sebastian Di Tella Benjamin Hébert, Pablo Kurlat, and Qitong Wang. The zero-beta rate. Working Paper, Stanford University and USC, 2023.

Bengt Holmström and Jean Tirole. Private and public supply of liquidity. *Journal of Political Economy*, 106 (1):1–40, 1998.

Jing-Zhi Huang, Yoshio Nozawa, and Zhan Shi. The global credit spread puzzle. *Journal of Finance*, 80(1): 101–162, 2025.

Oleg Itskhoki and Dmitry Mukhin. Exchange rate disconnect in general equilibrium. *Journal of Political Economy*, 129(8):2183–2232, 2021.

Urban Jermann. Negative swap spreads and limited arbitrage. *The Review of Financial Studies*, 33(1): 212–238, 2020.

Zhengyang Jiang, Arvind Krishnamurthy, and Hanno Lustig. Foreign safe asset demand and the dollar exchange rate. *Journal of Finance*, 76(3):1049–1089, 2021.

Zhengyang Jiang, Arvind Krishnamurthy, and Hanno Lustig. Dollar safety and the global financial cycle. 2023.

Alejandro Justiniano and Giorgio E Primiceri. The time-varying volatility of macroeconomic fluctuations. *American Economic Review*, 98(3):604–641, 2008.

Johnny Kang and Carolin Pflueger. Inflation risk in corporate bonds. *Journal of Finance*, 70(1):115–162, 2015.

Rohan Kekre and Moritz Lenel. The flight to safety and international risk sharing. NBER Working Paper wp29238, 2021.

John Maynard Keynes. The general theory of employment. *Quarterly Journal of Economics*, 51(2):209–223, 1937.

Lutz Kilian. Not all oil price shocks are alike: Disentangling demand and supply shocks in the crude oil market. *American Economic Review*, 99(3):1053–1069, 2009.

Don H Kim and Jonathan H Wright. An arbitrage-free three-factor term structure model and the recent behavior of long-term yields and distant-horizon forward rates. 2005.

Arvind Krishnamurthy and Wenhao Li. The demand for money, near-money, and Treasury bonds. *Review of Financial Studies*, 36(5):2091–2130, 2023.

Arvind Krishnamurthy and Annette Vissing-Jorgensen. The aggregate demand for Treasury debt. *Journal of Political Economy*, 120(2):233–267, 2012.

Clemens Lehner, Jonathan Payne, and Bálint Szőke. Historical us funding cost advantage: 1860–2024. Presented at the BSE 2024 session on Safe Assets and the 2024 Richmond Fed US Government Debt Conference, 2025.

Wenhao Li. Public liquidity and financial crises. *American Economic Journal: Macroeconomics*, 2024.

Yang Liu, Lukas Schmid, and Amir Yaron. The risks of safe assets. Working Paper, University of Hong Kong and USC, 2021.

Francis Longstaff. The flight-to-liquidity premium in U.S. Treasury bond prices. *Journal of Business*, 77 (3):511–526, 2004.

Robert E. Jr. Lucas. On the mechanics of economic development. *Journal of Monetary Economics*, 22: 3–42, 1988.

Ulrike Malmendier and Stefan Nagel. Learning from inflation experiences. *Quarterly Journal of Economics*, 131(1):53–87, 2016.

David Mayers and Clifford Smith. Death and taxes: The market for flower bonds. *Journal of Finance*, 42 (3):685–698, 1987.

Robert Mundell. Inflation and real interest. *Journal of Political Economy*, 71(3):280–283, 1963.

Stefan Nagel. The liquidity premium of near-money assets. *Quarterly Journal of Economics*, 131(4):1927–1971, 2016.

Carolin Pflueger. Back to the 1980s or not? The drivers of inflation and real risks in Treasury bonds. *Journal of Financial Economics*, forthcoming, 2025.

Carolin Pflueger and Gianluca Rinaldi. Why does the Fed move markets so much? A model of monetary policy and time-varying risk aversion. *Journal of Financial Economics*, 146(1):71–89, 2022.

Monika Piazzesi, Ciaran Rogers, and Martin Schneider. Money and banking in a New Keynesian model. *Working Paper, Stanford University*, 2019.

Julio J Rotemberg and Michael Woodford. An optimization-based econometric framework for the evaluation of monetary policy. *NBER Macroeconomics Annual*, 12:297–346, 1997.

Andreas Schrimpf, Ilhyock Shim, and Hyun Song Shin. Liquidity management and asset sales by bond funds in the face of investor redemptions in march 2020. *BIS Bulletin*, (39), 2021.

Adam Hale Shapiro. Decomposing supply and demand driven inflation. *Journal of Money, Credit and Banking*, forthcoming, 2024.

Robert J Shiller. Irrational exuberance. Princeton University Press, 2016.

Miguel Sidrauski. Rational choice and patterns of growth in a monetary economy. *American Economic Review*, 57(2):534–544, 1967.

Richard Startz. Implicit interest on demand deposits. *Journal of Monetary Economics*, 5(4):515–534, 1979.

Jeremy C Stein and Adi Sunderam. The Fed, the bond market, and gradualism in monetary policy. *Journal of Finance*, 73(3):1015–1060, 2018.

James H Stock and Mark W Watson. Why has U.S. inflation become harder to forecast? *Journal of Money, Credit and Banking*, 39:3–33, 2007.

John B Taylor. Discretion versus policy rules in practice. *Carnegie-Rochester Conference Series on Public Policy*, 39:195–214, 1993.

John B. Taylor and John C. Williams. Chapter 15 - Simple and robust rules for monetary policy. In *Handbook of Monetary Economics*, volume 3, pages 829–859. Elsevier, 2010.

James Tobin. A general equilibrium approach to monetary theory. *Journal of Money, Credit and Banking*, 1 (1):15–29, 1969.

Harald Uhlig. A toolkit for analysing nonlinear dynamic stochastic models easily. In Ramon Marimon and Andrew Scott, editors, *Computational Methods for the Study of Dynamic Economics*, pages 30–61. Oxford University Press, 1999.

Jules H van Binsbergen, Yoshio Nozawa, and Michael Schwert. Duration-based valuation of corporate bonds. *The Review of Financial Studies*, 38(1):158–191, 2025.

Annette Vissing-Jorgensen. The treasury market in spring 2020 and the response of the federal reserve. *Journal of Monetary Economics*, 124:19–47, 2021.

Olivier Wang. Banks, low interest rates, and monetary policy transmission. Working Paper, NYU Stern School of Business, 2022.

Yifei Wang, Toni M Whited, Yufeng Wu, and Kairong Xiao. Bank market power and monetary policy transmission: Evidence from a structural estimation. *Journal of Finance*, 77(4):2093–2141, 2022.

Michael Woodford. Optimal interest-rate smoothing. *Review of Economic Studies*, 70(4):861–886, 2003b.

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.

A.1 Data Sources

Our main measure of inflation is the annual change in the consumer price index (CPI) for all urban consumers. Data after 1947 can be easily downloaded from FRED. Since we need a longer horizon, we use CPI data from Shiller (2016), who reports the data starting from the late 1800s. Data are downloaded using *"ie_data"* link on the website https://shillerdata.com/.

Our measure of Aaa-Treasury spread directly replicates Krishnamurthy and Vissing-Jorgensen (2012). From 1924 until 1999, we use the average yield on long-term government bonds (LTGVTBD) from the St. Louis Fed's FRED. From 2000 onwards, we use the yield on 20-year maturity Treasury bonds (GS20) from FRED. The monthly Moody's seasoned Aaa corporate bond index (Aaa) is also from FRED. Similarly, we obtain Moody's seasoned Baa corporate bond index from FRED.

We additionally construct the T-bill convenience following Nagel (2016). Specifically, we directly use the main T-bill convenience series in Nagel (2016), downloaded 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 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.

For the decomposition of inflation into supply and demand components, we rely on the data constructed by Adam Shapiro, downloaded from https://www.frbsf.org/research-and-insights/data-and-indicators/supply-and-demand-driven-pce-inflation/. This data is available at monthly frequency from 1969 onwards.

For the analysis in Section 5, we also use many other measures of convenience yields, including agency-Treasury spread, agency-Treasury STRIP spread, and Refcorp-Treasury STRIP spread. Agency-Treasury spread is the yield spread between matched-maturity agency bonds and Treasuries. Data are from Bloomberg, with tickers "H15T[maturity]" for Treasuries and "C090[maturity]" for agency bonds, where "[maturity]" can be "3M", "6M", "1Y", "5Y", etc. Agency-Treasury STRIP spread is the yield spread between matched-maturity agency-bond STRIPs and Treasury STRIPs (STRIPs are zero-coupon bonds derived stripped from principal and coupon payments), with tickers "C094[maturity]" for agency-bond STRIPs and "C079[maturity]" for Treasury STRIPs. Refcorp-Treasury STRIP spread is the yield spread between Refcorp STRIPs and Treasury STRIPs. Refcorp STRIPs are zero-coupon bonds stripped from Refcorp bonds, which are 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 STRIP spread is free from default risks. Tickers for Refcorp STRIPs are "C091[maturity]".

We also use both survey-based inflation measures and inflation-swap rates in Section 5. The survey inflation expectation, denoted by $\mathbb{E}$[inflation], is the one-year (four quarters) forecast of inflation from the Survey of Professional Forecasters. This expectation data are available at Philadelphia Fed, https://www.philadelphiafed.org/surveys-and-data/data-files. Inflation swap rates are available from Bloomberg, tickers "USSWIT[maturity]".

We control for other well-known drivers of Treasury convenience, in particular, market volatility, the total government debt supply, and monetary policy. For market volatility, we use the VIX index. The VIX data are only available since 1990 (ticker "VIXCLS" on FRED). For the period before 1990, we use a linear projection of VIX on monthly realized volatility of the S&P 500 index returns (calculated as the standard deviation of daily index returns at each month), where the projection coefficients are estimated on the post-1990 data. S&P 500 index return data can be obtained from WRDS. We use the variable "VW_return", which is a value-weighted return including dividends.

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".

A.2 Expected Credit Risks in Aaa and Baa Corporate Bonds

The Aaa- and Baa-corporate bonds in Moody's index have long maturities around 20 years. One concern about the Aaa index is that it contains non-negligible credit risks over that longer horizon. In this subsection, we provide a way to quantify that risks and we show that even at 20-year horizon, after accounting for transition dynamics across different ratings, the credit risk in Aaa index is still negligible compared to the Baa index.

Denote the one-year default probability for various credit ratings as π^d and loss given default as L_{loss}, which are both column vectors. We assume that the expectation of default probability at any future year is still π^d.

Let the one-year transition probability across ratings be Q. Then the probability of not defaulting in the next year is simply $1 - \pi^d$ for the vector of ratings. The probability of not defaulting in the next two years for rating bucket i is

$$(1 - \pi^d(i)) \sum_j Q_{i,j}(1 - \pi^d(j)).$$

Denote $\odot$ as element-wise multiplication and $*$ as matrix multiplication. In matrix form, the vector

of probabilities for not defaulting in the next two years is

$$(1-\pi^d)\odot Q*(1-\pi^d)$$

Similarly, the probability of not defaulting in the next three years for rating bucket i is

$$(1-\pi^d(i))\sum_j Q_{i,j}(1-\pi^d(j))\sum_k Q_{j,k}(1-\pi^d(k)),$$

To generalize the matrix notation, we denote the transition probability accounting for not defaulting as $\hat{Q}$, defined as

$$\hat{Q}_{i,j}\equiv q_{i,j}(1-\pi^d(j)).$$

Then the probability of not defaulting in the next three years in matrix notation is

$$(1-\pi^d)\odot\hat{Q}*Q*(1-\pi^d).$$

More generally, the probability of not defaulting in n year is

$$(1-\pi^d)\odot\underbrace{\hat{Q}*\cdots*\hat{Q}}_{n-2}*Q*(1-\pi^d).$$

Then the annualized loss rate in a 20-year horizon is

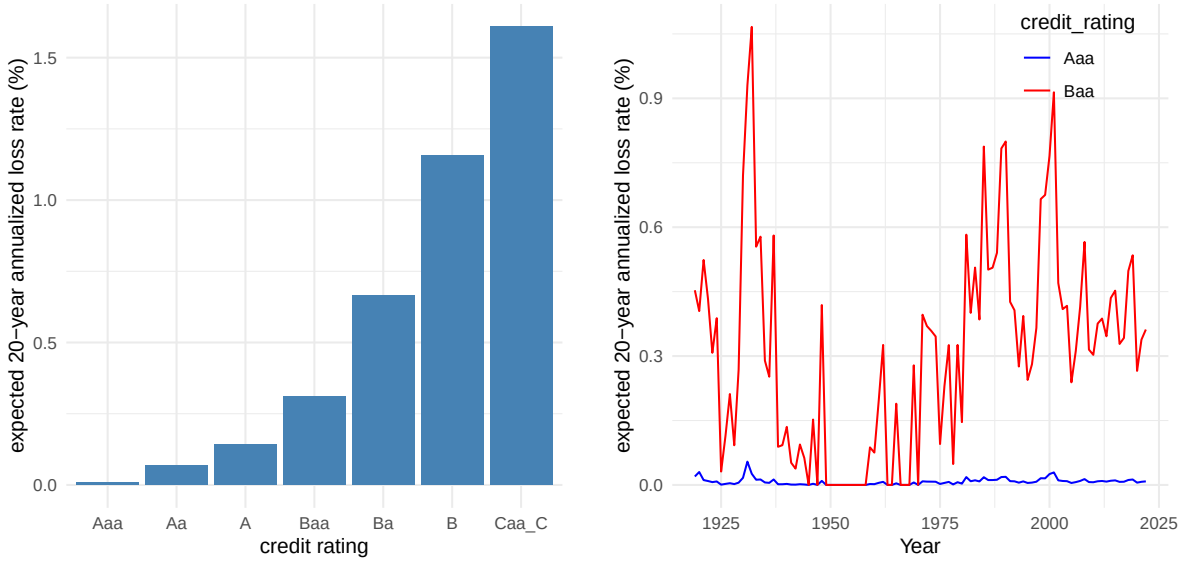
$$\frac{1}{20}\left(1-(1-\pi^d)\odot\underbrace{\hat{Q}*\cdots*\hat{Q}}_{18}*Q*(1-\pi^d)\right)\odot L_{\text{loss}}. \tag{A1}$$

In the data, the realized default rate varies a lot across different years, while loss given default and the transition probabilities across ratings are more stable. The change in realized default rate could influence the market perception of default risks. To reflect such dynamics, we use the realized default rate in each year as an approximation for expected default rates π_d, so we can uncover dynamics of the expected long-term credit loss rate for each year in the data. We use Moody's Default & Recovery Database for this purpose, and our sample are from 1920 to 2022. To match the Moody's Baa and Aaa index construction, we consider U.S. domiciled firms in the industrial and public utilities sectors. That is, different from the report, international issuers and

financial issuers are excluded. For each year, the realized default rates for this domestic set of firms are used to represent π_d.

To calculate the expected loss rate in the data, we use the average credit migration matrix Q reported by Exhibit 31 in Moody's investor service report (Emery et al., 2009). This migration matrix is estimated on data from 1920–2008. We extract the transition-without-default probabilities by conditioning the transition on no default, since the model deals with default separately. Then, we use the recovery rates L_{loss} reported by the last column of Exhibit 27 in the same report.

Figure A1. Expected Annualized Credit Loss for a 20-year Horizon. The graphs present the expected annualized credit loss rate for a 20-year horizon, calculated using equation (A1) and data from Moody's investor service report (Emery et al., 2009). To match the Moody's Baa and Aaa index construction, we consider U.S. domiciled firms in the industrial and public utilities sectors. That is, different from the report, international issuers and financial issuers are excluded. To reflect the dynamics of belief updating, we use the realized default ratio in each year from 1920 to 2022 for each credit rating, i.e., π_d is updated each year. In panel (a), we illustrate the average expected loss rate by credit rating over the whole sample. In panel (b), we plot the dynamics of expected loss for Aaa- and Baa-rated bonds.



(a) Average expected credit loss rate

(b) Dynamics of expected credit loss rate

In Figure A1, we illustrate both the dynamics of the expected loss rate (only for Aaa and Baa) and the average loss rate across ratings. In panel (a), we find an extremely low average of the expected 20-year annualized loss rate for Aaa-rated bonds, which is 0.007%. In contrast, the number for Baa-rated bonds is 0.31%, which is more than 40 times that of the Aaa-rated bonds.

We therefore conclude that for a 20-year horizon, on average, the credit risk of Aaa-rated bonds is negligible. In panel (b), we further probe if the fluctuations of the expected credit loss rate are of concern. We find that the maximum expected credit loss rate for Aaa-rated bonds is 0.05% over the entire sample, but the maximum expected credit loss rate for Baa-rated bonds is 1%. Again, we find that fluctuations of the expected loss rate are negligible for Aaa-rated bonds but not the case for Baa-rated bonds.

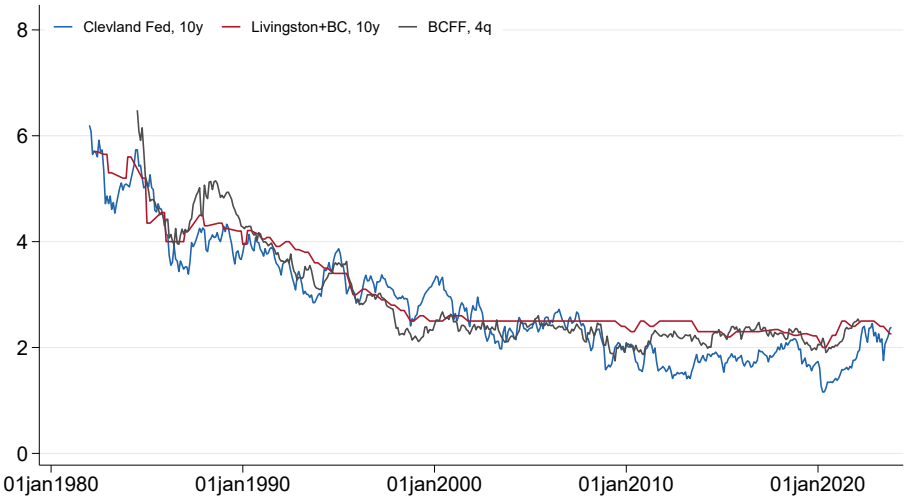
A.3 The Cleveland Fed index, inflation expectations, and term premia

We next investigate the suitability of the popular Cleveland Fed inflation expectations measures for analyzing convenience yields. This index is used, for example, in Acharya and Laarits (2023) and Fu et al. (2023). 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.

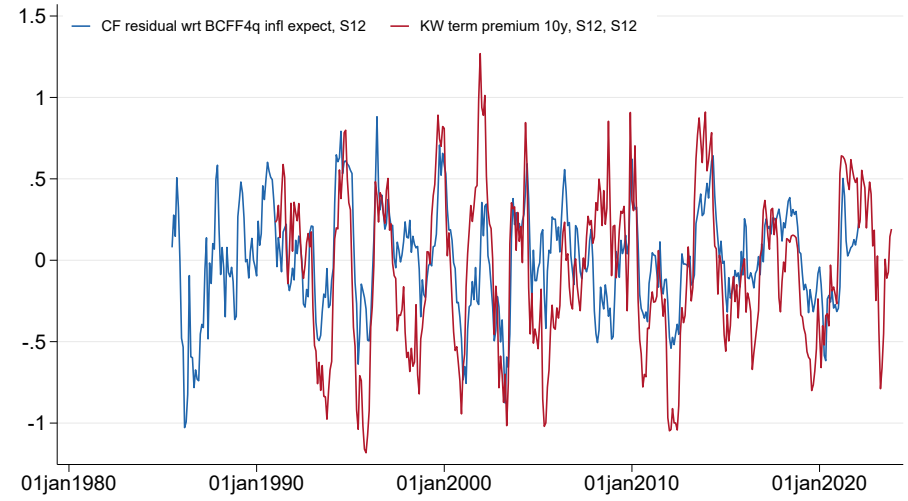
Panel A of Figure A2 shows 10-year inflation expectations from the Cleveland Fed against long-term and 4-quarter consensus inflation expectations from surveys. It is clear that the Cleveland Fed inflation expectations roughly move similarly at lower frequencies, but are substantially more volatile, raising 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 (2005). 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 (2005). 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.4 Regulation Q period

A deposit cap mandated by law might lower the pass-through from loan rates to deposit rates, and increase the pass-through to short-term convenience, similarly to a lower δ in our model.

Table A1 confirms this prediction in the data by interacting the fed funds rate with the Regulation Q (Reg Q) period dummy. The Reg Q dummy takes the value of one from 1966 to 1982, when Drechsler et al. (2023) document the constraint was binding, and zero otherwise. Columns (1)–(3) show that Reg Q amplifies the comovement between the short-term Tbill convenience and the fed funds rate, in line with Drechsler et al. (2023), but leaves our broad conclusions unchanged. Column (3) shows that including the fed funds rate interacted with the Reg Q dummy reduces the positive loading on inflation in the 1952-1999 period and makes it somewhat less statistically significant, though the interaction between inflation and the 1952-1999 dummy remains statistically significant at the 10% level. This is as expected if short-term convenience is driven by monetary policy interacted with the constraints on deposit-taking institutions, as formalized in our model.

Moreover, columns (4) through (6) show that the relationship between long-term convenience and inflation – our main object of interest – remains economically and statistically unchanged in every subperiod even when we control for the fed funds rate interacted with the Reg Q dummy. Different from the short-term convenience spread in columns (1)–(3), we see that the fed funds rate interaction with the Reg Q dummy is only weakly significant in column (6). At the same time, the coefficient on inflation interacted with the 1952-1999 dummy remains positive and economically and statistically highly significant in column (6). Column (6) also shows that the baseline coefficient on inflation and on the interaction between inflation and the post-2000 dummy remain unchanged compared to column (4). These results make sense, if investors expected that deposit caps would eventually be phased out over the lifetime of the long-term bonds, which was on average 20-30 years. Overall, these results are in line with our model where monetary policy interacted with deposit rate pass-through is a key driver of short-term Treasury convenience, but long-term inflation expectations drive the money channel for long-term Treasury convenience.

Table A1. Baseline results with regulation Q period. Monthly data runs from 1926:01 through 2020:12, excluding the 1939:09–1951:12 period. Inflation is interacted with the three subperiod dummies defined as in the baseline specifications (see Table 2). The period 1926:01–1939:08 acts as the omitted period. I_{regQ} is the dummy variable taking the value of one over the 1966-1982 period. Newey-West t-statistics with 12 lags are shown in parentheses. * p <0.10, ** p <0.05, *** p <0.01.

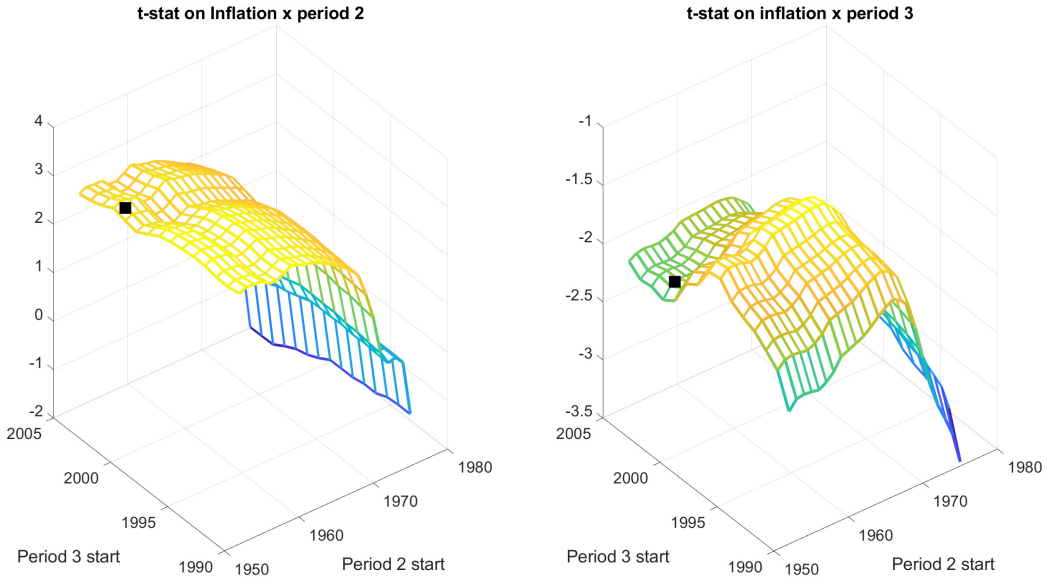
	Tbill spread			AAA-Tsy spread		
	(1)	(2)	(3)	(4)	(5)	(6)
Inflation	-0.029***	-0.027***	-0.027***	-0.038***	-0.038***	-0.038***
	(-4.69)	(-5.95)	(-5.83)	(-5.15)	(-5.31)	(-5.24)
Inflation x $I_{1952-1999}$	0.15***	0.082***	0.045*	0.13***	0.12***	0.088**
	(9.04)	(3.95)	(1.91)	(6.45)	(3.87)	(2.42)
Inflation x $I_{\geq 2000}$	0.070***	0.0040	0.012	-0.034	-0.049	-0.043
	(4.30)	(0.16)	(0.48)	(-0.93)	(-1.33)	(-1.17)
FFR		0.075***	0.066***		0.017	0.0097
		(6.00)	(5.41)		(0.86)	(0.49)
FFR x I_{regQ}			0.038***			0.030*
			(3.00)			(1.95)
$I_{1952-1999}$	-0.15**	-0.15**	-0.074	-0.52***	-0.51***	-0.46***
	(-2.12)	(-2.15)	(-1.09)	(-5.12)	(-5.30)	(-4.13)
$I_{\geq 2000}$	-0.091	0.11	0.082	0.27***	0.31***	0.29***
	(-1.48)	(1.48)	(1.19)	(2.71)	(2.89)	(2.70)
Constant	0.24***	0.046	0.068	0.92***	0.88***	0.89***
	(4.87)	(0.90)	(1.35)	(13.20)	(10.56)	(10.48)
$\bar{R}^2$	0.49	0.58	0.60	0.39	0.39	0.41
N	992	992	992	992	992	992

A.5 Robustness for Different Break Dates

In this subsection, we present robustness to the specification of the break dates in our baseline regressions. We also include plots of the moving averages of Treasury convenience and inflation.

Figure A3 plots the t-statistics for the interaction coefficients from the baseline regression (1) over a range of dates demarcating the starts of the second and third periods. The specification is identical to column (1) of Table 2, except that we vary the start of the second period between 1952 and 1975 and the start of the third period between 1990 and 2005. The left panel shows that the $\pi_t \times I_{period2,t}$ loading is positive and significant for a broad range of start dates for the second period and almost completely insensitive to the start of the third period. The right panel shows that the $\pi_t \times I_{period3,3}$ loading is negative with a t-statistic exceeding -2 in absolute value if we allow the second period to start in the 1950s or 1960s and the third period to start any time between 1995 and 2005.

Figure A3. T-statistics for different period start dates. This figure reports results for the baseline regression in column (1) of Table 2 using different start dates for periods 2 and 3. The first break date (period 2 start) ranges from 1952 to 1975 and is shown on the x-axis. The second break date (period 3 start) ranges from 1990 to 2005 and is shown on the y-axis. Our baseline break dates – 1st break in 1952 and 2nd break in 2000– are indicated with black squares. The t-statistics for the interaction coefficients between period dummies with inflation are shown on the z-axis and are based on Newey-West standard errors with 12 lags.



A.6 Supply/Demand Decompositions

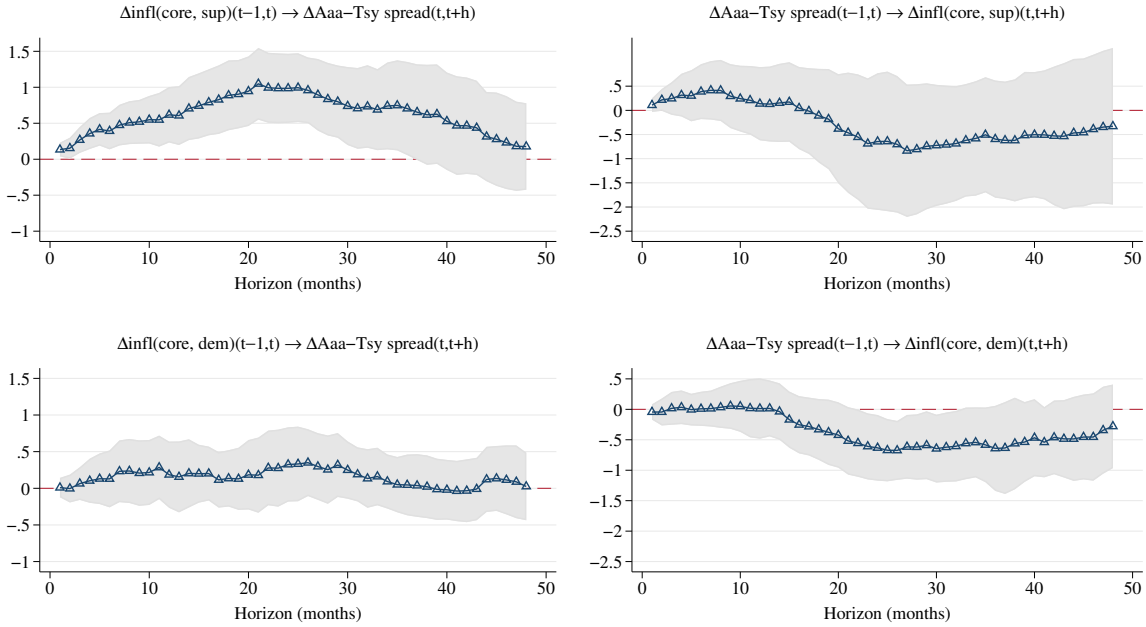
In this section, we will show the full set of estimates on predictive regressions between Treasury convenience yield and the supply and demand components of inflation in Figure A4, which encompasses subplots in Figure 3 of the main text. The main message is that convenience yields only predict the demand component of inflation, but not the supply component. Conversely, only the supply component of inflation predicts convenience, but demand inflation does not.

In Panel A of Figure 3, we find that the supply component of inflation leads to significantly higher convenience yield during the high-inflation period 1969-1999, but not the other way around. Moreover, higher convenience yield leads to a significantly lower demand component of inflation, not the other way around. These results support the presence of the money channel and the liquidity demand channel. Moroever, they together imply that supply shocks dominate demand shocks during this inflationary period.

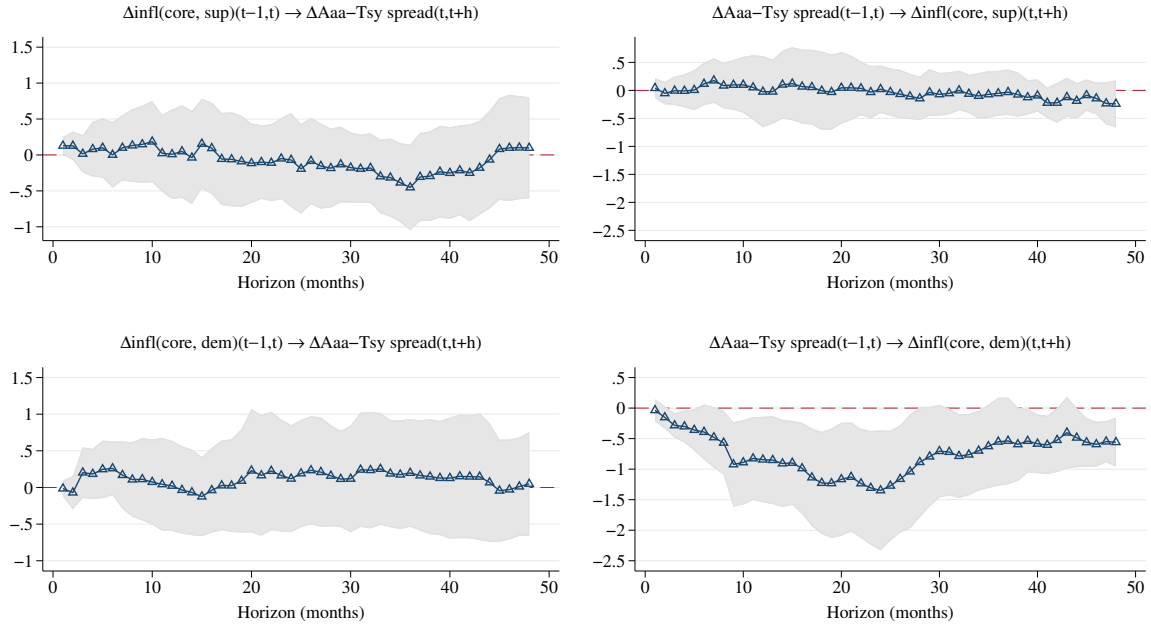
In Panel B of Figure 3, we find that the only significant relationship is that higher convenience yield leads to a lower demand component of inflation, consistent with the liquidity demand channel. The insignificance of the supply component of inflation leading to higher convenience indicates that there are few supply shocks post-2000.

Figure A4. Predictive regressions with demand- vs. supply-driven inflation components. This figure presents the full set of estimates to accompany results in Figure 3 in the main text.

Panel A: 1969–1999



Panel B: 2000-2020



B Model Appendix

In this appendix, we provide detailed model solutions. Appendix B.7 solves a model extension, where Treasury bonds and deposits are imperfect substitutes, rather than perfect substitutes as in our baseline.

B.1 Habit dynamics

In this model, relative risk aversion equals $-U_{CC}C/U_C = \gamma/S_t$, where surplus consumption is the share of market consumption available to generate utility:

$$S_t = \frac{C_t - H_t}{C_t}. \tag{A2}$$

The stochastic discount factor (SDF) M_{t+1} used to price all financial assets and to derive the macroeconomic consumption Euler equation equals:

$$M_{t+1} = \beta \frac{\frac{\partial U_{t+1}}{\partial C}}{\frac{\partial U_t}{\partial C}} = \beta \exp\left(-\gamma(\Delta s_{t+1} + \Delta c_{t+1})\right). \tag{A3}$$

As equation (A2) makes clear, a model for market habit implies a model for surplus consumption and vice versa. Following Campbell et al. (2020), we model market habit implicitly by assuming that log surplus consumption, s_t, satisfies:

$$\begin{aligned} s_{t+1} &= (1-\theta_0)\bar{s} + \theta_0 s_t + \theta_1 x_t + \theta_2 x_{t-1} + \lambda(s_t)\varepsilon_{c,t+1}, & \text{(A4)} \\ \varepsilon_{c,t+1} &= c_{t+1} - \mathrm{E}_t c_{t+1}. & \text{(A5)} \end{aligned}$$

The sensitivity function $\lambda(s_t)$ takes the form first proposed by Campbell and Cochrane (1999):

$$\begin{aligned} \lambda(s_t) &= \begin{cases} \frac{1}{\bar{S}}\sqrt{1-2(s_t-\bar{s})} - 1 & s_t \leq s_{max} \\ 0 & s_t > s_{max} \end{cases}, & \text{(A6)} \\ \bar{S} &= \sigma_c \sqrt{\frac{\gamma}{1-\theta_0}}, & \text{(A7)} \\ \bar{s} &= \log(\bar{S}), & \text{(A8)} \\ s_{max} &= \bar{s} + 0.5(1-\bar{S}^2). & \text{(A9)} \end{aligned}$$

Here, σ_c denotes the standard deviation of the consumption surprise $\varepsilon_{c,t+1}$ and $\bar{s}$ is the steady-state value for log surplus consumption. The consumption surprise is an equilibrium object depending on fundamental shocks, and we will verify in our solution that it is conditionally homoskedastic and lognormal.

Here, x_t equals stochastically detrended consumption (up to a constant):

$$x_t = c_t - (1-\phi)\sum_{j=0}^{\infty}\phi^j c_{t-1-j}, \tag{A10}$$

where ϕ is a smoothing parameter. For the production side described below, x_t is also the log output gap, where potential output is defined as the level of output that would be achieved with flexible prices and steady-state markups, taking aggregate productivity as given.

B.2 Production Side

B.2.1 Final good

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

$$Y_t = \left(Y_{i,t}^{\frac{\epsilon_{p,t}-1}{\epsilon_{p,t}}}\right)^{\frac{\epsilon_{p,t}}{\epsilon_{p,t}-1}}. \tag{A11}$$

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 $P_{i,t}$:

$$Y_{i,t} = Y_t\left(\frac{P_{i,t}}{P_t}\right)^{-\epsilon_{p,t}}. \tag{A12}$$

The aggregate price level is given by

$$P_t = \left(\int_0^1 P_{i,t}^{-(\epsilon_{p,t}-1)}di\right)^{-\frac{1}{\epsilon_{p,t}-1}}. \tag{A13}$$

B.2.2 Intermediate good producers

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

$$Y_{i,t} = A_t N_{i,t}^{1-\tau}, \tag{A14}$$

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 (A13) and may therefore choose a different amount of the aggregate labor index. With the final good equation (A11) aggregate output equals

$$Y_t = A_t N_t^{1-\tau}, \tag{A15}$$

where aggregate labor is defined:

$$N_t \equiv \left[\int_0^1 N_{i,t}^{\frac{(\epsilon_{p,t}-1)(1-\tau)}{\epsilon_{p,t}}} di \right]^{\frac{\epsilon_{p,t}}{(\epsilon_{p,t}-1)(1-\tau)}}. \tag{A16}$$

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) we assume that productivity depends on past skills gained by all agents, and depends on past market labor, n_{t-1}:

$$a_t = \nu + a_{t-1} + (1-\phi)(1-\tau)n_{t-1}, \tag{A17}$$

where $0 \leq \phi < 1$ and $\nu > 0$ are constants. The assumption (A17) ensures that potential output increases with past output and implies (A10).

B.2.3 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.

B.3 Price-setters' inflation expectations

Price-setters are assumed to have partially backward-looking inflation expectations

$$\tilde{E}_t\pi_{t+1} = (1-\zeta)E_t\pi_{t+1} + \zeta\pi_{t-1}, \tag{A18}$$

B.4 Model Derivations

B.4.1 Liquidity Spread

Taking the difference between (14) vs. (13) and (15) vs. (13) gives the following expressions

$$\frac{I_t^l - I_t^b}{1+I_t^l} = \frac{\frac{\alpha}{Q_t}\lambda_t}{U_c\left(C_t, H_t\right)} \tag{A19}$$

$$\frac{I_t^l - I_t^d}{1+I_t^l} = \frac{\frac{\alpha}{Q_t}(1-\lambda_t)}{U_c\left(C_t, H_t\right)}. \tag{A20}$$

Substituting in (8) into (A20) then gives (16) in the main text. A simple rearrangement gives that

$$I_t^l = \frac{1}{1-\frac{\lambda_t}{1-\lambda_t}(1-\delta)} I_t^b \tag{A21}$$

B.4.2 Flexible Price Steady-State

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. In the steady-state interest rates must satisfy:

$$\left(1+\bar{I}^l\right)\beta E\left[\frac{U_c(C_{t+1})}{U_c(C_t)}\frac{P_t}{P_{t+1}}\right] = 1, \tag{A22}$$

where we suppress the non-consumption arguments in the utility function to simplify notations. In the nonstochastic steady-state consumption and habit are constant, so:

$$1 + \bar{I}^l = \frac{1 + \bar{\Pi}}{\beta}. \tag{A23}$$

The steady-state government bond yield (in levels) then satisfies

$$\bar{I}^b = \bar{I}^l \left(1 - \frac{\bar{\lambda}}{1 - \bar{\lambda}}(1 - \delta)\right). \tag{A24}$$

B.4.3 Log-Linearization

We define the log steady-state interest rates by $\bar{i}^l = \log(1 + \bar{I}^l)$, and $\bar{i}^b = \log(1 + \bar{I}^b)$. Also define the log interest rates by $i_t^l = \log(1 + I_t^l)$ and $i_t^b = \log(1 + I_t^b)$. The deviations are expressed as $\hat{i}_t^l = i_t^l - \bar{i}^l$ and $\hat{i}_t^b = i_t^b - \bar{i}^b$. For conciseness we define the function

$$\phi(\lambda_t) = \frac{1}{1 - \frac{\lambda_t}{1-\lambda_t}(1 - \delta)}. \tag{A25}$$

The function ϕ has the first-order Taylor approximation around $\bar{\lambda}$ in terms of $\hat{\lambda}_t \equiv \lambda_t - \bar{\lambda}$:

$$\phi(\lambda_t) \approx \phi\left(\bar{\lambda}\right) + \phi'\left(\bar{\lambda}\right)\hat{\lambda}_t, \tag{A26}$$

$$\phi\left(\bar{\lambda}\right) = \frac{1}{1 - \frac{\bar{\lambda}}{1-\bar{\lambda}}(1 - \delta)}, \tag{A27}$$

$$\phi'\left(\bar{\lambda}\right) = \left(\frac{1}{1 - \frac{\bar{\lambda}}{1-\bar{\lambda}}(1 - \delta)}\right)^2 (1 - \delta)\frac{1}{\left(1 - \bar{\lambda}\right)^2} \tag{A28}$$

We can then re-write expression (A21)

$$\exp(\hat{i}_t^l + \bar{i}^l) - 1 = \phi(\lambda_t)\left(\exp(\hat{i}_t^b + \bar{i}^b) - 1\right) \tag{A29}$$

Substituting in the log-linear approximation for ϕ,

$$(1 + \bar{I}^l)\hat{i}_t^l + \bar{I}^l \approx \left(\phi(\bar{\lambda}) + \phi'(\bar{\lambda})\hat{\lambda}_t\right)\left((1 + \bar{I}^b)\hat{i}_t^b + \bar{I}^b\right) \tag{A30}$$

Solving out for $\hat{i}_t^l$ and dropping second-order terms gives the first-order Taylor expansion

$$\begin{aligned}\hat{i}_t^l &\approx \phi(\bar{\lambda})\frac{1+\bar{I}^b}{1+\bar{I}^l}\hat{i}_t^b + \phi'(\bar{\lambda})\frac{\bar{I}^b}{1+\bar{I}^l}\hat{\lambda}_t, &\text{(A31)}\\ &= f^i\hat{i}_t^b + f^\lambda\hat{\lambda}_t, &\text{(A32)}\end{aligned}$$

where the coefficients are given by

$$\begin{aligned}f^i &= \phi(\bar{\lambda})\frac{1+\bar{I}^b}{1+\bar{I}^l}, &\text{(A33)}\\ &= \frac{1}{1-\left(\frac{\bar{\lambda}}{1-\bar{\lambda}}(1-\delta)\right)}\frac{1+\bar{I}^l\left(1-\left(\frac{\bar{\lambda}}{1-\bar{\lambda}}(1-\delta)\right)\right)}{1+\bar{I}^l} &\text{(A34)}\\ &= \frac{1}{1-\left(\frac{\bar{\lambda}}{1-\bar{\lambda}}(1-\delta)\right)}\left(1-\frac{\bar{I}^l}{1+\bar{I}^l}\left(\frac{\bar{\lambda}}{1-\bar{\lambda}}(1-\delta)\right)\right), &\text{(A35)}\\ f^\lambda &= \phi'(\bar{\lambda})\frac{\bar{I}^b}{1+\bar{I}^l}. &\text{(A36)}\end{aligned}$$

As long as $\frac{\bar{I}^l}{1+\bar{I}^l} < 1$, $\bar{\lambda} > 0$ and $\delta < 1$ the second expression for f^i makes clear that $f^i > 1$. Alternatively, f^i can be written as $f^i = \phi(\bar{\lambda})\frac{1+\bar{I}^b}{1+\bar{I}^l} = \frac{\phi(\bar{\lambda})+\bar{I}^l}{1+\bar{I}^l}$, which can be easily used to see that $f^i > 1$.

We then derive the relationship between convenience spreads across the term structure. Investing one dollar into an n-period zero coupon government bonds at time t and selling it at time $t+1$ generates a return $R_{n,t+1} = \frac{\exp\left(-(n-1)i^b_{n-1,t+1}\right)}{\exp\left(-ni^b_{n,t}\right)}$. Since government bonds are assumed to provide the same liquidity services at time t irrespective of bond maturity, the first-order condition between investing in an n-period vs. 1-period bond becomes

$$0 = \beta E_t\left[\frac{U_c\left(C_{t+1}\right)}{U_c\left(C_t\right)}\left(\exp(i_t^b) - \frac{\exp\left(-(n-1)i^b_{n-1,t+1}\right)}{\exp\left(-ni^b_{n,t}\right)}\right)\right]. \quad \text{(A37)}$$

Log-linearizing gives the long-term liquid government bond yield in terms of the expected short-term government bond yields according to the expectations hypothesis:

$$i^b_{n,t} = \frac{1}{n}E_t\left[\sum_{i=0}^{n-1} i^b_{t+i}\right]. \quad \text{(A38)}$$

Since short- and long-term illiquid loans also generate the same liquidity value at time t, their yields up to first-order also satisfy an expectations hypothesis:

$$i^l_{n,t} = \frac{1}{n} E_t \left[\sum_{i=0}^{n-1} i^l_{t+i} \right]. \tag{A39}$$

We derive log-linearized Euler equation (21) following standard steps. As shown in Pflueger and Rinaldi (2022), the representative household's first order-condition for an illiquid one-period bond is exactly log-linear in the current, expected, and lagged output gaps.

$$x_t = f^x E_t x_{t+1} + \rho^x x_{t-1} - \psi r^l_t + \theta_t, \tag{A40}$$

$$f^x = \frac{1}{\phi - \theta_1},\ \rho^x = \frac{\theta_2}{\phi - \theta_1},\ \psi = \frac{1}{\gamma(\phi - \theta_1)}, \tag{A41}$$

with the demand shock taking the form

$$\theta_t = \psi \left(\log(\Theta_t) - E_t \log(\Theta_{t+1}) \right). \tag{A42}$$

Because the labor-leisure trade-off is standard, the firm's problem is also entirely standard. The first-order condition for firms' price-setting behavior leads to the following log-linearized Phillips curve

$$\pi_t = f^\pi E_t \pi_{t+1} + \rho^\pi \pi_{t-1} + \kappa x_t + v_{\pi,t}, \tag{A43}$$

$$\rho^\pi = \frac{1}{1+\beta_g} + \frac{\beta_g}{1+\beta_g}\zeta,\ f^\pi = 1 - \rho^\pi,\ \beta_g = \beta exp(-(\gamma - 1)g). \tag{A44}$$

where 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.

B.5 Fisher equation

Following Campbell et al. (2020) we approximate the one-period log nominal yields for liquid and illiquid bonds by assuming that the Fisher equation holds for one-period, i.e.

$$r_t^l = i_t^l - E_t\pi_{t+1}, \tag{A45}$$
$$r_t^b = i_t^b - E_t\pi_{t+1}. \tag{A46}$$

These approximations are appropriate if inflation risk premia on one-quarter bonds are small. When solving for bond prices with risk premia, we do not make these assumptions for two- and longer-term bonds, instead solving for inflation risk premia.

B.5.1 Solution algorithm details

Replacing u_t with ξ_t in the state vector, we will use the following state vector $Z_t = \left[x_t, \pi_t, i_t^b, \xi_t, \theta_t\right]$ and the vector of exogenous shocks $v_t = [v_{\xi,t}, v_{\pi,t}, v_{i,t}, v_{\theta,t}]$. The log-linearized dynamics for the state vector can then be summarized as

$$x_t = (1-\rho^x)E_t x_{t+1} + \rho^x x_{t-1} - \psi f^i i_t^b + \psi E_t\pi_{t+1} + \psi b f^\lambda \pi_t + \xi_t + \theta_t, \tag{A47}$$
$$\pi_t = (1-\rho^\pi)E_t\pi_{t+1} + \rho^\pi \pi_{t-1} + \kappa x_t + v_{\pi,t}, \tag{A48}$$
$$i_t^b = (1-\rho^i)\left(\gamma^x x_t + \gamma^\pi \pi_t\right) + \rho^i i_{t-1}^b + v_{i,t}, \tag{A49}$$
$$\xi_t = \rho^\xi \xi_{t-1} + v_{\xi,t}, \tag{A50}$$
$$\theta_t = \rho^\theta \theta_{t-1} + v_{\theta,t}. \tag{A51}$$

In matrix form, the model can be written as

$$0 = FE_t[Z_{t+1}] + GZ_t + HZ_{t-1} + Mv_t, \tag{A52}$$

where the matrices are given by

$$F = \begin{bmatrix} 1-\rho^x & \psi & 0 & 0 & 0 \\ 0 & 1-\rho^\pi & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 \end{bmatrix} \tag{A53}$$

$$G = \begin{bmatrix} -1 & \psi b f^\lambda & -\psi f^i & 1 & 1 \\ \kappa & -1 & 0 & 0 & 0 \\ (1-\rho^i)\gamma^x & (1-\rho^i)\gamma^\pi & -1 & 0 & 0 \\ 0 & 0 & 0 & -1 & 0 \\ 0 & 0 & 0 & 0 & -1 \end{bmatrix} \tag{A54}$$

$$H = \begin{bmatrix} \rho^x & 0 & 0 & 0 & 0 \\ 0 & \rho^\pi & 0 & 0 & 0 \\ 0 & 0 & \rho^i & 0 & 0 \\ 0 & 0 & 0 & \rho^\xi & 0 \\ 0 & 0 & 0 & 0 & \rho^\theta \end{bmatrix} \tag{A55}$$

$$M = \begin{bmatrix} 0 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ 0 & 0 & 1 & 0 \\ 1 & 0 & 0 & 0 \\ 0 & 0 & 0 & 1 \end{bmatrix} \tag{A56}$$

We use Uhlig (1999)'s formulation of Blanchard and Kahn (1980) to solve for an equilibrium of the form (24).

To express the dependence of convenience yield on the state vector, we define e_i as an indicator row vector of the same size as Z_t with a 1 in the i-th position and 0 elsewhere. Using equation

(10), (19), and the definition of ξ_t, we obtain

$$\begin{aligned} i_t^l - i_t^b &= (f^i - 1)i_t^b + f^\lambda(\bar{u} - b\pi_t + u_t) \\ &= (f^i - 1)i_t^b + f^\lambda \bar{u} - bf^\lambda \pi_t + f^\lambda u_t \\ &= (f^i - 1)i_t^b + f^\lambda \bar{u} - bf^\lambda \pi_t + \psi^{-1}\xi_t \\ &= f^\lambda \bar{u} + \left((f^i - 1)e_3 - bf^\lambda e_2 - \psi^{-1}e_4\right) Z_t \end{aligned} \tag{A57}$$

Then using the first-order relation between long-term convenience and short-term convenience in equation (20), we get the following expression:

$$spread_{n,t} \equiv i_{n,t}^l - i_{n,t}^b = f^\lambda \bar{u} + \frac{1}{n}\left(\left(f^i - 1\right)e_3 - \psi^{-1}e_4 - bf^\lambda e_2\right)\left(I - P^Z\right)^{-1}\left(I - (P^Z)^n\right) Z_t \tag{A58}$$

B.6 Illustrative Risk Premia

We can derive illustrative expressions for the risk premia on two-period liquid and illiquid bonds. Equation (A37) gives the standard expression for the expected excess return on a two-period liquid bond. Let $r_{2,t+1}^b$ denote the log nominal return on the two-period liquid bond, and π_{t+1} log inflation from time t to time $t+1$. Letting m_{t+1} denote the log real SDF so the log nominal SDF equals $m_{t+1}^\$ - \pi_{t+1}$, the bond pricing equation for the liquid two-period bond becomes

$$E_t exp\left(m_{t+1} + r_{2,t+1}^b - \pi_{t+1}\right) \;=\; E_t exp\left(m_{t+1} + i_t^b - \pi_{t+1}\right) \tag{A59}$$

Assuming that all variables are log-normal, the left-hand side gives

$$E_t\left(m_{t+1} + r_{2,t+1}^b - \pi_{t+1}\right) + \frac{1}{2}Var_t\left(r_{2,t+1}^b - \pi_{t+1}\right) + \frac{1}{2}Var_t m_{t+1} + Cov_t\left(r_{2,t+1}^b - \pi_{t+1}, m_{t+1}\right)$$

The right-hand-side gives

$$E_t(m_{t+1} - \pi_{t+1}) + i_t^b + \frac{1}{2}Var_t\left(\pi_{t+1}\right) + \frac{1}{2}Var_t m_{t+1} + Cov_t\left(-\pi_{t+1}, m_{t+1}\right)$$

Canceling terms gives

$$0 \;=\; E_t\left(r^b_{2,t+1} - i^b_t\right) + \frac{1}{2}Var_t\left(r^b_{2,t+1} - \pi_{t+1}\right) - \frac{1}{2}Var_t\pi_{t+1} + Cov_t\left(r^b_{2,t+1}, m_{t+1}\right). \quad \text{(A60)}$$

We can then use that the innovation in the log real SDF is proportional to consumption and hence output gap innovations via $m_{t+1} - E_t m_{t+1} = -\gamma\left(x_{t+1} - E_t x_{t+1}\right)$ and substitute in for the log liquid bond return to obtain the expected real excess return on a two-period liquid bond. Using that the log nominal return on the two-period liquid bond equals $r^b_{2,t+1} = -i^b_{t+1} + 2i^b_{2,t}$

$$\begin{aligned} E_t\left[r^b_{2,t+1} - i^b_t\right] + \frac{1}{2}Var_t\left[r^b_{2,t+1} - \pi_{t+1}\right] - \frac{1}{2}Var_t\pi_{t+1} &= -Cov_t\left(m_{t+1}, r^b_{2,t+1}\right), \quad \text{(A61)} \\ &= \gamma Cov_t\left(x_{t+1}, r^b_{2,t+1}\right), \quad \text{(A62)} \\ &= -\gamma Cov_t\left(x_{t+1}, i^b_{t+1}\right). \quad \text{(A63)} \end{aligned}$$

Because i^b_t equals the sum of the real liquid short-rate and expected inflation, this is the sum of a real rate risk premium plus an expected inflation risk premium.

Analogously, for the two-period illiquid bond offers the identical liquidity services as the one-period illiquid bond, so

$$E_t exp\left(m_{t+1} + r^l_{2,t+1} - \pi_{t+1}\right) \;=\; E_t exp\left(m_{t+1} + i^l_t - \pi_{t+1}\right). \quad \text{(A64)}$$

Analogous algebra then implies

$$E_t\left[r^l_{2,t+1} - i^l_t\right] + \frac{1}{2}Var_t\left[r^l_{2,t+1} - \pi_{t+1}\right] - \frac{1}{2}Var_t\pi_{t+1} \;=\; -\gamma Cov_t\left(x_{t+1}, i^l_{t+1}\right) \quad \text{(A65)}$$

Substituting in the log-linearized dynamics used for the macroeconomic solution (16) gives

$$\begin{aligned} E_t\left[r^l_{2,t+1} - i^l_t\right] + \frac{1}{2}Var_t\left[r^l_{2,t+1} - \pi_{t+1}\right] - \frac{1}{2}Var_t\pi_{t+1} &= -\gamma Cov_t\left(x_{t+1}, f^i i^b_{t+1} + f^\lambda \lambda_{t+1}\right), \\ &= \underbrace{-\gamma f^i Cov_t\left(x_{t+1}, i^b_{t+1}\right)}_{\text{nominal rate RP}} \\ &\quad \underbrace{+ -\gamma f^\lambda Cov_t\left(x_{t+1}, \lambda_{t+1}\right)}_{\text{liquidity RP}} \quad \text{(A66)} \end{aligned}$$

The expected return on a two-period illiquid minus that on a two-period liquid bond then equals

$$\begin{aligned} & E_t\left[r^l_{2,t+1} - r^b_{2,t+1}\right] + \frac{1}{2}\left(Var_t\left[r^l_{2,t+1} - \pi_{t+1}\right] - Var_t\left[r^b_{2,t+1} - \pi_{t+1}\right]\right) \\ = \; & \underbrace{-\gamma\left(f^i - 1\right)Cov_t\left(x_{t+1}, i^b_{t+1}\right)}_{\text{nominal rate RP}} + \underbrace{-\gamma f^\lambda Cov_t\left(x_{t+1}, \lambda_{t+1}\right)}_{\text{liquidity RP}} \\ & \underbrace{+(f^i - 1)i^b_t + f^\lambda \lambda_t}_{\text{Expectations Hypothesis}}. \end{aligned} \tag{A67}$$

B.7 Extension to 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 Fu et al. (2023)

$$Q_t = \left((1-\lambda_t)D_t^\rho + \lambda_t B_t^\rho\right)^{1/\rho}, \tag{A68}$$

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

$$I^l_t - I^b_t = \frac{\lambda_t}{1-\lambda_t}\left(\frac{B_t}{D_t}\right)^{\rho-1}(1-\delta)I^l_t, \tag{A69}$$

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 $\hat{d}_t \equiv \log D_t - \log \bar{D}$

$$\hat{i}_t^l \approx f^i \hat{i}_t^b + f^\lambda \hat{\lambda}_t - f^b \left(\hat{b}_t - \hat{d}_t \right), \tag{A70}$$

$$f^b = \frac{1}{\left(1 - \frac{\bar{\lambda}}{1-\bar{\lambda}}(1-\delta)\left(\bar{B}/\bar{D}\right)^{\rho-1}\right)^2} (1-\rho)(1-\delta) \frac{\bar{\lambda}}{1-\bar{\lambda}} \left(\bar{B}/\bar{D}\right)^{\rho-1} \frac{\bar{I}_b}{1+\bar{I}_l} \geq 0, \tag{A71}$$

$$f^i = \frac{1}{1 - \left(\frac{\bar{\lambda}}{1-\bar{\lambda}}(1-\delta)\left(\bar{B}/\bar{D}\right)^{\rho-1}\right)} \left(1 - \frac{\bar{I}^l}{1+\bar{I}^l} \left(\frac{\bar{\lambda}}{1-\bar{\lambda}}(1-\delta)\left(\bar{B}/\bar{D}\right)^{\rho-1}\right)\right), \tag{A72}$$

$$f^\lambda = \left(\frac{1}{1 - \frac{\bar{\lambda}}{1-\bar{\lambda}}(1-\delta)\left(\bar{B}/\bar{D}\right)^{\rho-1}}\right)^2 (1-\delta)\left(\bar{B}/\bar{D}\right)^{\rho-1} \frac{1}{\left(1-\bar{\lambda}\right)^2} \frac{\bar{I}^b}{1+\bar{I}^l} \tag{A73}$$

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.

Substituting into the Euler equation (21) gives

$$x_t = \rho^x x_{t-1} + (1-\rho^x) E_t x_{t+1} - \psi \left(f^i i_t^b - E_t \pi_{t+1} \right) - \psi \left(f^\lambda \hat{\lambda}_t - f^b \left(\hat{b}_t - \hat{d}_t \right) \right) + v_{\theta,t} \tag{A74}$$

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 negative 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.8 Shocks to Overall Liquidity Demand

A simple extension considers shocks to the overall liquidity weight in the utility function, α. Combining equations (13) and (14) gives

$$E_t\left[M_{t+1}^{\$}\right]\left(I_t^l - I_t^b\right) = \frac{\alpha_t/Q_t\lambda_t}{U_c\left(C_t, Q_t, H_t, N_t, \Theta_t\right)}. \tag{A75}$$

Combining equations (13) and (15) gives

$$E_t\left[M_{t+1}^{\$}\right]\left(I_t^l - I_t^d\right) = \frac{\alpha_t/Q_t(1-\lambda_t)}{U_c\left(C_t, Q_t, H_t, N_t, \Theta_t\right)}. \tag{A76}$$

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 (8) this possibility is precluded, as α_t is not allowed to enter into the deposit spread by assumption in equilibrium. Substituting (A75) into (A76) then gives equation (16) 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 (8) is assumed to hold.

Different assumptions are of course possible. For example, one could replace (8) 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 (A75) and (A76) 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.

B.9 Details on Model Parameters

Our model parameters are set in two steps. First, we use typical values from the literature for standard parameters. Second, in each period, targeting data moments, we estimate shock volatility parameters and how Aaa-Treasury spread maps to convenience yield in the model. This two-step approach allows us to identify the dominant economic forces in each period without overly complicating the process.

First, basic parameters of the model are calibrated and shown in Table A2.

Table A2. 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 values for δ and ρ^λ in the extension with direct liquidity-inflation link are identical to the baseline model and therefore not repeated.

Panel A: Inflation 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	0.80	Fuhrer (1997a)
Monetary Policy			
MP inertia	ρ^i	0.8	Clarida et al. (2000)
Output gap weight	γ^x	0.5	Taylor (1993)
Inflation weight	γ^π	1.5	Taylor (1993)
Panel B: Interest Rates and Liquidity			
Discount rate	β	0.98	Average nominal policy rate = 4%
Steady-state inflation	$\bar{\Pi}$	2%	Fed inflation target
Deposit rate pass-through	δ	0.34	Nagel (2016)
Bond liquidity weight	$\bar{u}$	0.14	Level T-bill Convenience Spread
Persistence liquidity	ρ^λ	0.91	AR(1) Aaa-Tsy Spread

Second, we use data moments to identify shock volatilities across in each period, as well as how Aaa-Treasury spread in the data maps to convenience yield in the model. Since Aaa-Treasury spread may over- or under-state the contribution of convenience yield to the Euler equation, we define the mapping as

$$i^l_{n,t} - i^b_{n,t} = \chi * \text{(Aaa-Treasury spread)}, \tag{A77}$$

where we use $n = 40$ quarters (i.e., 10 years) for long-term convenience yield. We will estimate the scaling factor together with 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_θ. 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_θ. 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.

For each period, we estimate the scaling factor χ and the four volatility parameters by targeting seven data moments: volatilities and the pairwise correlations among convenience yield (AAA-Treasury spread), yearly inflation, and output gap, as well as the regression coefficient of AAA-Treasury spread on yearly inflation. Denote these data moments as m^{data}. For each parameter vector (σ, χ), 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 $\hat{m}(\sigma, \chi)$. We estimate parameters to minimize the weighted squared diviations from the data moments,

$$\min_{\sigma,\chi}(m^{\text{data}} - \hat{m}(\sigma,\chi))' W^{-1}(m^{\text{data}} - \hat{m}(\sigma,\chi)) \tag{A78}$$

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 9 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. We estimate the parameters using the fminunc solver in Matlab. In Table 10 of the main text, we report estimate parameters.

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 $g(\hat{\sigma}, \hat{\chi}) = m - \hat{m}(\hat{\sigma}, \hat{\chi})$, 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 devided by T^{data} as S. Second, we numerically compute the Jacobian matrix that represents how sensitive the moments are to parameter changes,

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

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

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

where as explained before, S denotes the estimated covariance matrix of moment rsiduals, 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 numberes in parentheses in second columns of each period in Table 10.

Next, to illustrate how the scaling factor χ in equation (A77) affects estimation results, in Table A4 and Table A3, we show estimated parameters and moments when we set the scaling factor to one, so that we enforce the Aaa-Treasury spread to be one to one mapping to the Euler-equation gap $i^l_{n,t} - i^b_{n,t}$. In this case, the estimation problem becomes

$$\min_{\sigma}(m^{\text{data}} - \hat{m}(\sigma; \chi = 1))'W^{-1}(m^{\text{data}} - \hat{m}(\sigma; \chi = 1)). \tag{A79}$$

And we adjust the standard error calculations for parameters accordingly.

In Table A4, we find that the model matches data moments well in the second and third periods even though there is no flexibility of the scaling factor. This is consistent with Table 9 that the estimated scaling factor is not far from 1 in both periods. Indeed, the estimated parameters as reported in Table A4 are similar to those in Table 10. However, in the pre-WWII period, absent from scaling, the model implies a much larger monetary policy shock with high confidence. Intuitively, since there is no scaling and AAA-Treasury spread is small compared to output gaps, other sources of shocks are needed to generate volatile output. However, voltile monetary policy shokcs generate positive correlation between convenience and inflation, to the contrary of the data moment. Our estimation is thus informative about the classical debate on whether the great depression is driven

Table A3. Estimated Parameters (without Scaling the Spread). This table report period-by-period estimation of model parameters, including volatilities of four shocks in the model, and the scaling factor of how to map Aaa-Treasury spread into $i^l_{n,t} - i^b_{n,t}$ in equation (20). We set the scaling factor in (A77) to one so there is no scaling on the spread.

parameters (yearly values)	1920–1939		1952–1999		2000–2020	
σ_ξ (liquidity demand)	0.181	(0.016)	0.103	(0.002)	0.162	(0.007)
σ_π (cost-push)	0.000	(0.407)	0.425	(0.009)	0.000	(0.191)
σ_i (monetary policy)	3.958	(0.101)	0.001	(2.299)	0.014	(2.168)
σ_x (non-liquidity demand)	0.001	(0.106)	0.053	(0.001)	0.042	(0.001)
spread scaling	1.000		1.000		1.000	

by liquidity demand shocks or policy mistakes (Friedman and Schwartz, 1963; Bernanke, 1983). Contrasting Table 9 with Table A4, we find that the liquidity demand shock is a more plausible main driver of the great Depression while monetary policy shocks play a secondary role. Our conclusion highly relies on the novel moment that we bring to the model – the correlation between inflation and Treasury convenience yield, which was not systematically analyzed by the literature.

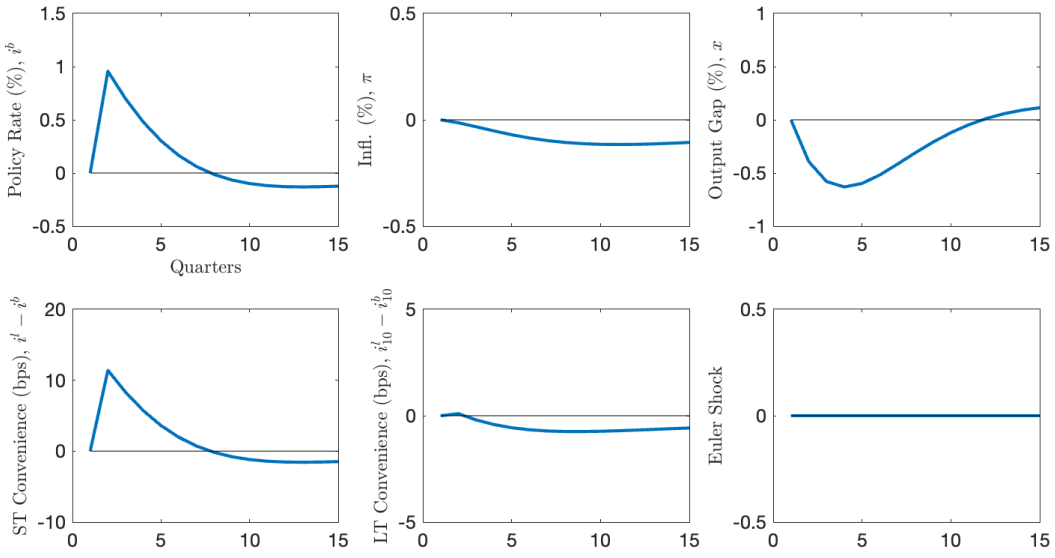
Table A4. Target Moment: Data v.s. Model (without Scaling the Spread). In this table, we report data moments and their associated standard deviations (in brackets), as well as the simulated model moments with estimated parameters using SMM estimation in (A78). We set the scaling factor in (A77) to one so there is no scaling on the spread.

		1920–1939		**1952–1999**		**2000–2020**	
		Data	Model	Data	Model	Data	Model
Volatilities	Vol(Aaa-Treasury)	0.27	0.36	0.45	0.25	0.33	0.32
		(0.02)		(0.02)		(0.03)	
	Vol(Inflation)	6.27	2.10	3.09	2.21	1.25	1.09
		(0.51)		(0.16)		(0.10)	
	Vol(Output Gap)	9.28	6.03	2.42	3.24	2.01	2.25
		(0.75)		(0.12)		(0.16)	
Correlations	Corr(Inflation, Output Gap)	0.34	0.18	-0.21	-0.48	0.53	0.14
		(0.10)		(0.07)		(0.08)	
	Corr(Aaa-Treasury, Inflation)	-0.37	0.10	0.64	0.44	-0.29	-0.17
		(0.10)		(0.04)		(0.10)	
	Corr(Aaa-Treasury, Output Gap)	-0.20	-0.16	-0.38	-0.30	-0.23	-0.44
		(0.11)		(0.06)		(0.11)	
Reg Coefs	Aaa-Treasury $\sim \beta \cdot$ Inflation	-0.02	0.02	0.09	0.05	-0.07	-0.05
		(0.01)		(0.02)		(0.04)	

B.10 Impulse response to monetary policy shock

Figure A5 shows the impulse responses to a monetary policy shock in our baseline model. We see that long-term convenience and inflation both decline in response to a monetary policy shock, whereas short-term convenience and the nominal policy rate increase. This happens because the monetary policy shock first drives up the policy rate, but then eventually causes overshooting in the policy rate, as inflation declines following the contractionary shock. The short-term spread increases, similarly to the increase in the policy rate. The long-term spread, which is forward-looking declines similarly to inflation. Monetary policy shocks hence cannot explain the negative relationship between inflation and long-term convenience during the Great Depression or the post-2000 period, though they may help understand why short- and long-term convenience diverged during the Volcker years.

Figure A5. Baseline Model Responses to a Monetary Shock. This figure shows impulse responses to a monetary policy shock for our baseline model. The monetary policy shock is a positive 100 bps shock to the 3-month T-bill. Responses for inflation, π and the Treasury rate i^b are in annualized percent units. The response for the convenience spread $i^l - i^b$ is in annualized basis points units.The response for the output gap x is in percent units. Quarters are shown on the x-axis.



B.11 Robustness for alternative model

We next provide robustness for the alternative model, varying the magnitude of the direct inflation-liquidity link, b. Figures A6 and A7 report alternative versions of Figure 8 in the main paper. The solid line shows the impulse response for the low inflation equilibrium (similar to the pre-WWII and post-2000s periods in the data), while the red dashed line shows the high-inflation equilibrium (similar to the second half of the 20th century in the data). Figure A6 sets a lower direct inflation-liquidity link parameter $b = 0.01$, while Figure A7 sets a higher direct inflation-liquidity link parameter $b = 0.02$. We see that the LT spread responses are robustly above the red dashed LT spread responses even if we vary the direct inflation-liquidity link parameter b to be either higher or lower than the values shown in the main paper. This means that the alternative model with a direct inflation-convenience link implies a *more negative* inflation-convenience relationship when inflation is high in steady-state, regardless of the strength of the direct inflation-convenience link parameter b. Of course, a stronger direct impact of inflation on liquidity, as in Figure A7, exacerbates those differences, but the qualitative ordering of the LT spread responses across high- and low-inflation equilibria is the same for both Figures A6 and A7. We find in the data that long-term Treasury convenience was significantly more positively related to inflation during the high-inflation 1970s and 1980s than during the lower-inflation pre-WWII and post-2000 periods. These additional plots show that the alternative model with $b > 0$ implies the opposite robustly across different numerical values for b.

Figure A6. Alternative Model Responses to Cost-Push Shock with $b = 0.01$. This figure is identical to Figure 8 in the main paper but sets $b = 0.01$.

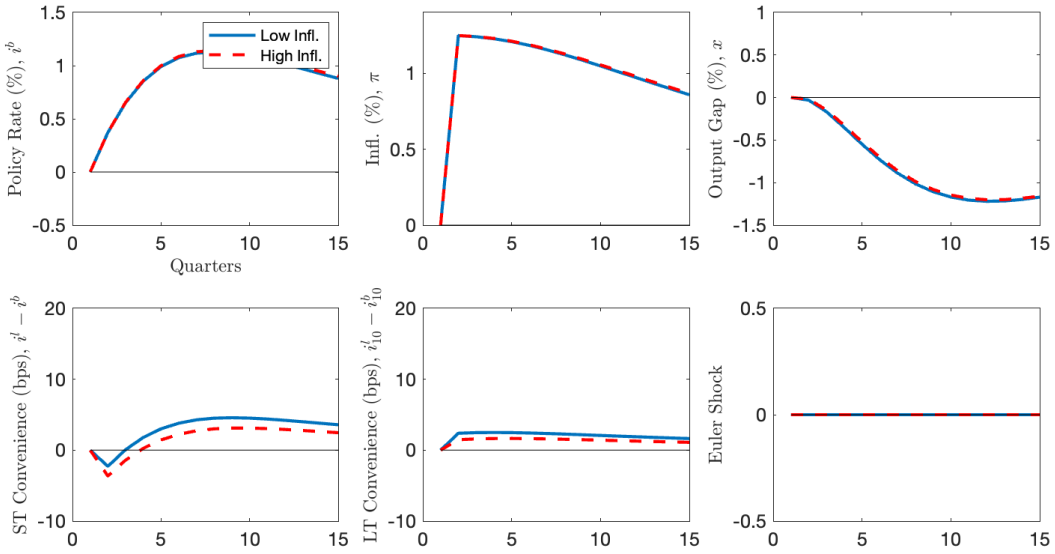


Figure A7. Alternative Model Responses to Cost-Push Shock with $b = 0.02$. This figure is identical to Figure 8 in the main paper but sets $b = 0.02$.

