Inflation and Treasury Convenience

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August 2024

Abstract

Using a century of data, we show that Treasury convenience yield and inflation comove positively during the inflationary 1970s-1980s, but negatively pre-WWII and post-2000. An inflation decomposition reveals that higher supply inflation predicts higher convenience, while lower demand inflation follows higher convenience. In our model, inflationary cost-push shocks raise the opportunity cost of holding money and money-like assets, inducing higher convenience, as in 1970s-1980s. Conversely, liquidity demand shocks drive up convenience but lower consumption demand and inflation in the model, as pre-WWII and post-2000. By linking the evidence to macroeconomic drivers, our results challenge the notion that inflation directly depresses Treasury convenience.

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).

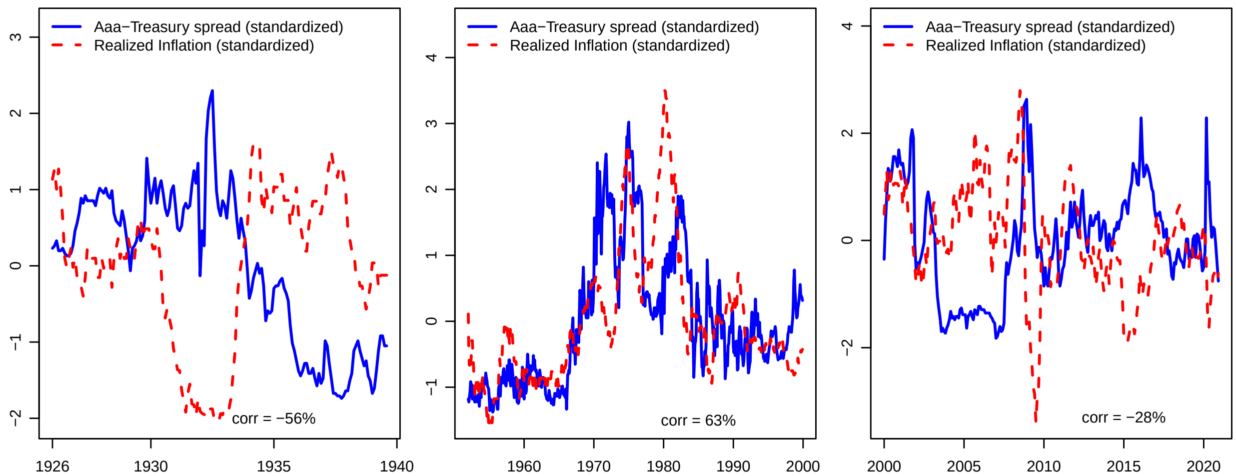
Using a century of U.S. data, we document secular shifts in the relationship between Treasury convenience and inflation, and link them to the macroeconomic drivers of inflation. Figure 1 illustrates a striking fact: Convenience comoves positively with inflation during the Great Inflation episode, but not before or after.[1] In periods typically associated with supply-side shocks, such as the 1970s and 1980s, 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 and 1980s. 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

[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.



lowers aggregate 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, 2023). 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. 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, 2023). 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. (2022), 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 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 interpretation 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.

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

[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.

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

[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

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 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 the default risks 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. Over longer horizons, Aaa-rated bonds may transition to lower rating categories, but even the 5-year cumulative credit loss rate for Aaa-rated bonds is only 0.02% for the period 1982-2008. By contrast, Baa-rated bonds have a cumulative credit loss rate of 1.1% at the 5-year horizon, corresponding to an annualized credit loss rate of 0.22%. 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.

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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).

(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. Again, this suggests that credit risks are negligible for Aaa bonds but an important component for Baa bonds. 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 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.

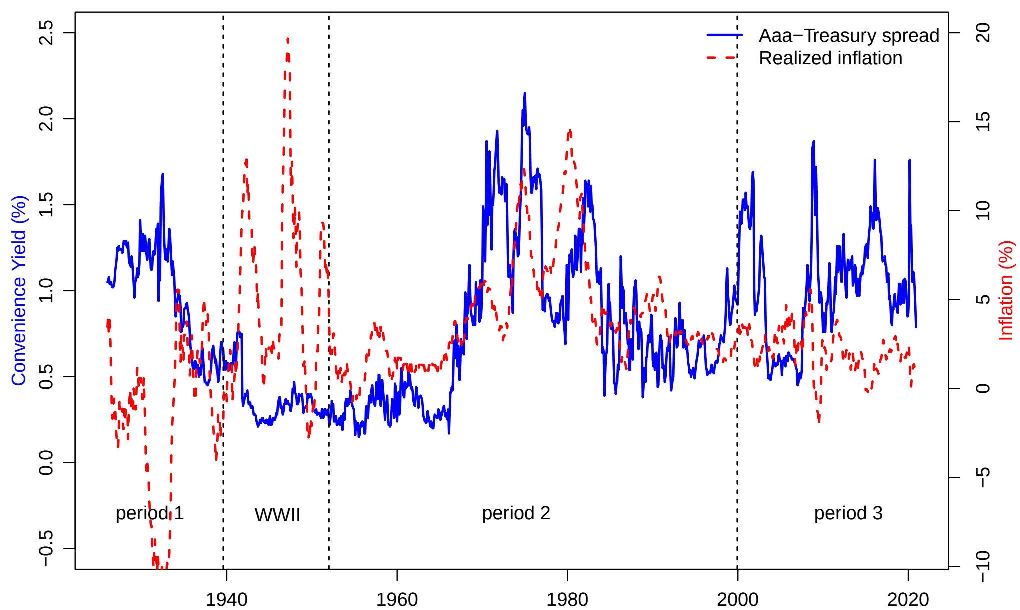
Variable	Obs	Mean	Std. Dev.	Mean		
	(1926-2020)			(1926-1939)	(1952-1999)	(2000-2020)
Aaa-Tsy spread (%)	992	0.87	0.42	0.98	0.78	1.03
T-bill convenience yield (%)	992	0.43	0.46	0.28	0.56	0.23
Inflation (%)	992	2.60	3.64	-1.58	3.99	2.13
VIX (%)	992	19.64	8.16	27.39	17.30	19.96
Baa-Aaa spread (%)	992	1.14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

[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.

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 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,

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

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 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 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 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

[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 detrended debt/GDP growth.

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

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 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

[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

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 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

[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 influenced 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).

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 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 indistinguish-

[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.

able 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 (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.

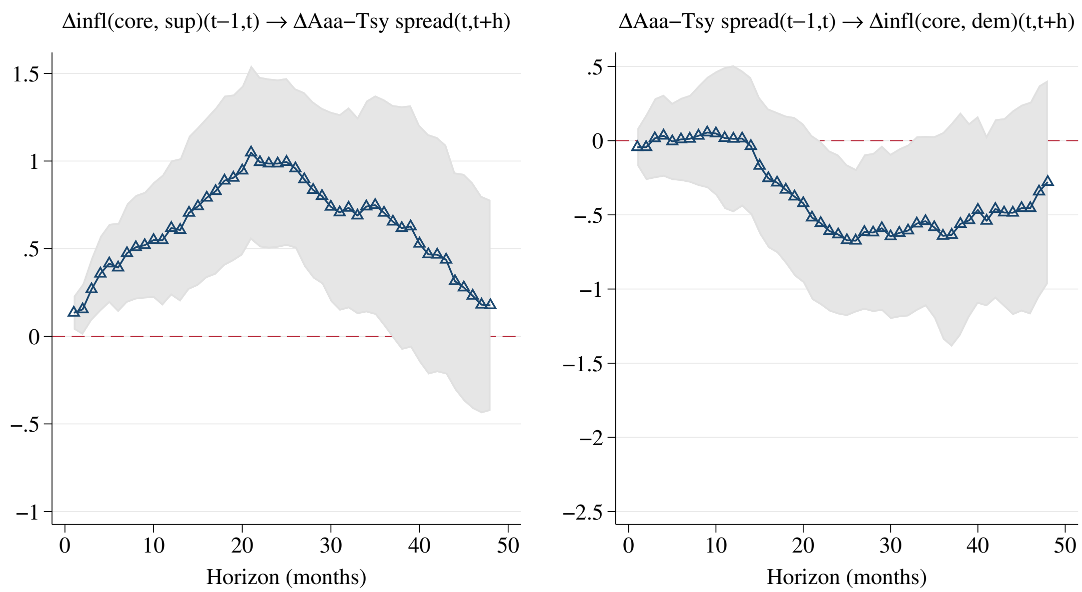
[17] In terms of magnitudes, the estimates for the 1969–1999 period imply that a one-standard-deviation positive monthly change in the supply inflation (or 13 bps) predicts a 0.3 standard deviations (or 13 bps) increase in the Aaa-Treasury spread over the next 24 months (top left panel in Figure 3).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

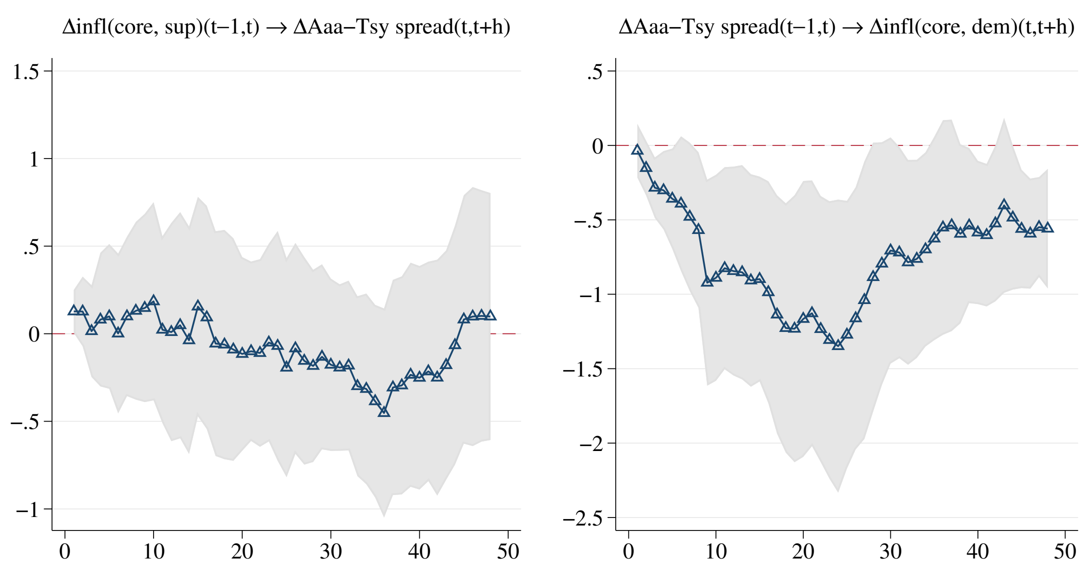
	Long-term convenience: Aaa-Treasury spread					
	(1)	(2)	(3)	(4)	(5)	(6)
	Core decomposition			Headline decomposition		
Infl (sup)	0.25***	0.24***	0.22***	0.15***	0.18***	0.18***
	(5.62)	(4.25)	(4.44)	(3.44)	(3.92)	(3.94)
Infl (dem)	-0.29***	-0.17	-0.11	-0.17**	-0.11	-0.076
	(-4.13)	(-1.12)	(-0.85)	(-2.18)	(-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

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



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., 2017, 2020; Ermolov, 2022; Pflueger, 2023), 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

[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

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 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

Our main conclusions are not sensitive to the specific break dates chosen for our analysis, as we show in Appendix A.5. We find that Aaa-Treasury spread loading on the second-period inflation interaction is significantly positive and the loading on the third-period inflation interaction is significantly negative, provided that the second period starts in 1950s or 1960s and the third period starts between 1995 and 2005. The economic mechanism driving the inflation-convenience relationship hence changed once sometime in the middle of the 20th century and then again around 2000.

Another important consideration for understanding the inflation-convenience relationship in the 1970s and 1980s relates to the regulatory constraints faced by banks. The Great Inflation is usually thought to have been triggered by recurring cost-push shocks. Drechsler et al. (2023) show that interest rate caps on deposit rates, mandated by Regulation Q, amplified the effect of monetary tightening on credit and endogenously exacerbated inflationary pressures. Consistent with Drechsler et al. (2023), we find that the interaction of the fed funds rate with a Regulation Q dummy explains substantial variation in short-term Treasury convenience and somewhat weakens the positive relationship between short-term convenience and inflation during the 1952-1999 period. However, the relationship between long-term Treasury 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 a Regulation Q dummy. Details are provided in Appendix A.4. We conclude that while monetary policy interacted with deposit rate pass-through affects the short-term Treasury convenience, long-term inflation expectations separately drive the money channel for the long-term Treasury convenience, in line with the model that we present next.

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.[19] 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.

[19] Bianchi et al. (2022) 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 Model Setup

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.[20] 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. We refer to λ_t as Treasury liquidity.

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

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

where

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

[20]The assumption of perfect substitutability makes the model as simple as possible while illustrating our main points. Appendix B.3 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.

Here, N_t denotes market labor supplied outside the home, and $H_t = hC_{t-1}$ denotes external consumption habit (Fuhrer (2000), Christiano et al. (2005)), i.e., consumers do not internalize the effects of their choices on future habit. External habit H_t serves to generate a backward-looking term in the Euler equation and slows down the output response to monetary policy.[21] The shifter Θ_t represents a taste shock increasing the utility that households derive from consumption today.

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

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

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:

$$I^d_t = \delta I^l_t, \qquad (8)$$

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.[22] 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

[21] We follow the macroeconomics literature in our specification of habit because we are not solving for risk premia here. Campbell et al. (2020) show that a somewhat more complicated habit specification can also explain salient features of asset prices while preserving the same macroeconomic equilibrium.

[22] 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)).[23] 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 iid, 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.[24] 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.[25] 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 the Appendix. 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). Such a setup generates a standard log-linearized Phillips curve with an extra backward-looking term. Since the model does not have real investment, the aggregate resource constraint implies that con-

[23] Appendix B.4 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.

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

[25] 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.

sumption 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. Details are in Appendix B.

4.1.4 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

$$\lambda_t = \bar{u} - b\pi_t + u_t \tag{10}$$

$$u_t = \rho^{\lambda} u_{t-1} + v_{\lambda,t}. \tag{11}$$

Here, ρ^{λ} is the persistence of Treasury liquidity. 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 $\theta_t \equiv \log \Theta_t$ is also normally distributed and follows an AR(1) process with the same autocorrelation coefficient ρ^{λ}. The cost-push supply shock is assumed to be log-normal and iid.

4.2 Asset Pricing Euler Equations

The nominal consumption-based stochastic discount factor equals

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

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

$$E_t\left[M^{\$}_{t+1}\left(1 + I^l_t\right)\right] \;=\; 1. \tag{13}$$

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

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

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. (2022) 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

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

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 Log-Linearized Model Dynamics

We log-linearize the model around the flexible-price steady-state $\bar{c} = \bar{y}$, $\bar{\pi}$, $\bar{i}^l$, $\bar{i}^b$, $\bar{\theta}$, and $\bar{\lambda}$. 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. Because potential output is constant, the log output gap x_t equals log output up to a constant, i.e., $x_t = y_t = c_t$. We focus on the first-order effects of liquidity. A first-order approximation abstracts from second-order terms, such as liquidity risk premia, which are treated in complementary papers by Du et al. (2023) and Acharya and Laarits (2023).

We start with the log-linearized expressions for bond yields and convenience spreads. Log-linearizing the expression (16) gives the following log-linear expression 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 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

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

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

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 log-linearized intertemporal first-order condition takes the standard form

$$
x_t = \rho^x x_{t-1} + (1 - \rho^x) E_t x_{t+1} - \psi \left(i^l_t - E_t \pi_{t+1} \right) + v_{x,t}, \quad (21)
$$

where the backward-looking coefficient equals $\rho^x = \frac{h}{1+h}$, and the elasticity of intertemporal substitution is given by $\psi = \gamma^{-1} \frac{1-h}{1+h}$. The demand shock equals $v_{x,t} = \psi \left(\theta_t - E_t \theta_{t+1} \right)$ and 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:

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

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 $f^\lambda \lambda_t$ 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 $f^i > 1$.[26]

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

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

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

$$Z_t = P^Z Z_{t-1} + Q^Z v_t, \quad (24)$$

[26]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.

where the state vector equals $Z_t = [x_t, \pi_t, i_t^b, u_t]$ and the vector of exogenous iid shocks is given by $v_t = [v_{\lambda,t}, v_{\pi,t}, v_{i,t}]$. To compare the effects of the convenience-driven demand, we also consider a standard New Keynesian benchmark without Treasury convenience where demand shocks originate solely from preference shocks, $v_{x,t}$. 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 Quantitative Illustration

We illustrate the properties of the model using standard parameter values, with parameters listed in Appendix Table A2. We will 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. 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 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 treats liquidity shocks λ_t as completely exogenous and, thus, sets $b = 0$ in equation (10).

The parameters for the New Keynesian block of the model are set to standard values, following the literature as detailed in Appendix Table A2. 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. We consider versions of the model with only one shock switched on at a time, so the impulse responses and regression results are invariant to the magnitude of the shock volatilities. Hence, we need to specify which shock is active and the size of the impulse, but not the equilibrium volatilities of shocks.

4.5 Baseline Model Impulse Responses

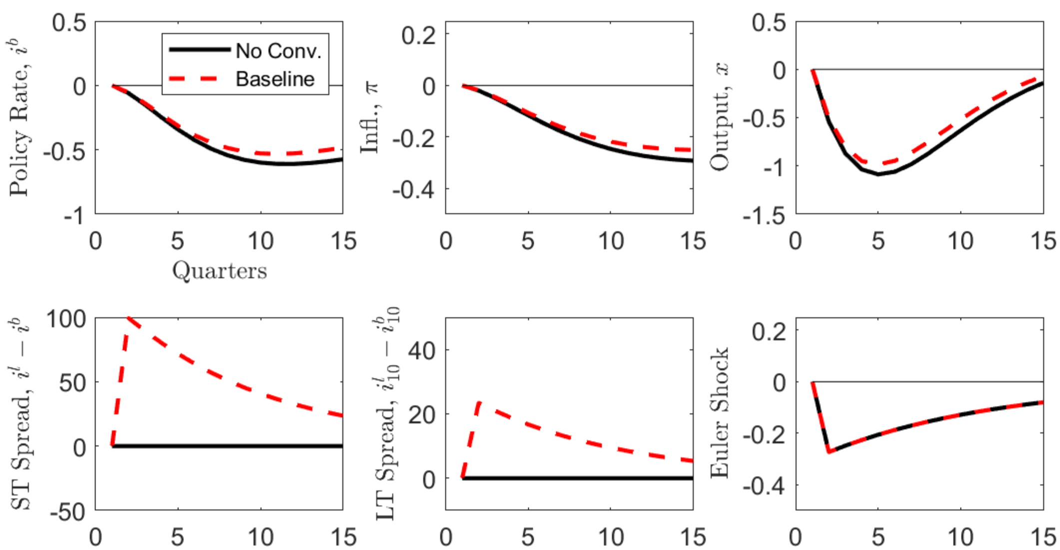
Figure 4 traces out the effect of a negative liquidity demand shock $v_{\lambda,t}$ for inflation, convenience spreads, the nominal policy rate i_t^b, and the output gap via impulse responses. The $v_{\lambda,t}$ shock is scaled so that its on-impact effect on the short-term convenience spread is 100 bps, implying a negative effect on aggregate demand in equation (22).

Figure 4 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 almost indistinguishable from an analogous sized traditional aggregate demand shock in a model without convenience (denoted by "No Conv."), showing the analogy between traditional demand-driven inflation and inflation movements due to the liquidity demand channel. The liquidity demand channel further indicates that higher convenience spreads should predict a lower demand-driven inflation, consistent with the empirical lead-lag relationship we find.

Figure 5 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).[27] Monetary policy is slightly amplified in the presence of Treasury convenience, deepening the recession but mitigating inflation in response to a positive cost-push

[27] 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 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.

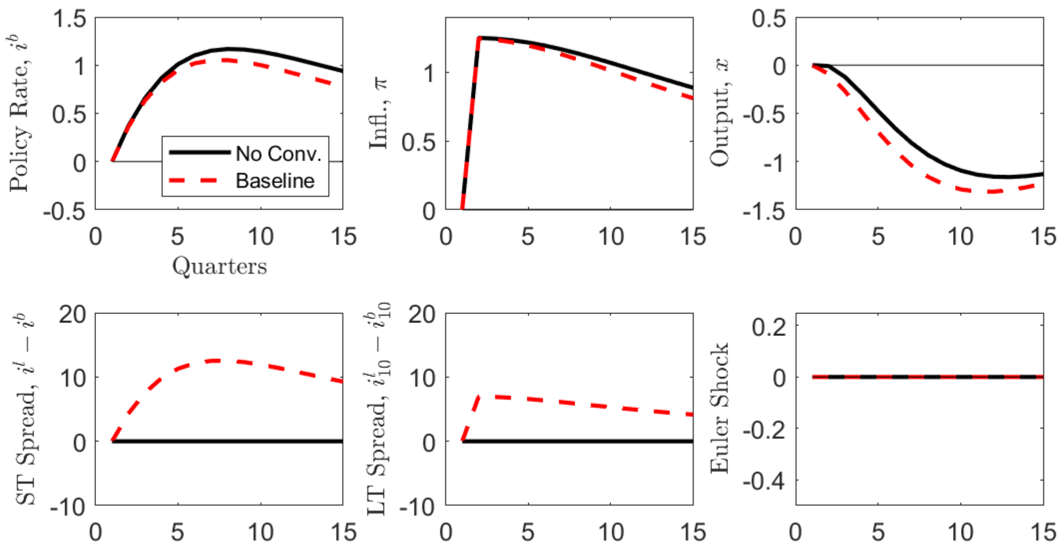
Figure 4. Baseline model responses to liquidity demand shock. This figure shows impulse responses to a liquidity demand shock to $v_{\lambda,t}$ for our baseline model with $b = 0$. The shock is scaled so that in the model with convenience it corresponds to a 100 bps increase in the convenience yield spread $i_t^l - i_t^b$. The black line reports the impulse response to an aggregate demand shock to the Euler equation, $v_{x,t}$, with identical AR(1) coefficient as the liquidity demand shock while setting Treasury convenience to zero ($\bar{\lambda} = 0$). 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.



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 component of inflation.[28]

[28]While the impulse responses in Figure 5 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.

Figure 5. 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, showing that the response of long-term Treasury convenience to a surprise interest rate hike is close to zero (see Appendix Figure A5). As such, 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). The small and ambiguous long-term convenience response to a monetary policy shock arises in the model because a surprise monetary policy tightening slightly lowers long-term inflation expectations while raising the short-term nominal policy rate, which has opposing effects on long-term convenience through equation (20). 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 expectations. Liquidity demand shocks are hence necessary to explain why in the data the long-term convenience-inflation comovement is negative while controlling for the federal funds rate during the pre-WWII and post-2000 periods.

4.6 Alternative Model Impulse Responses with Direct Inflation-Convenience Link

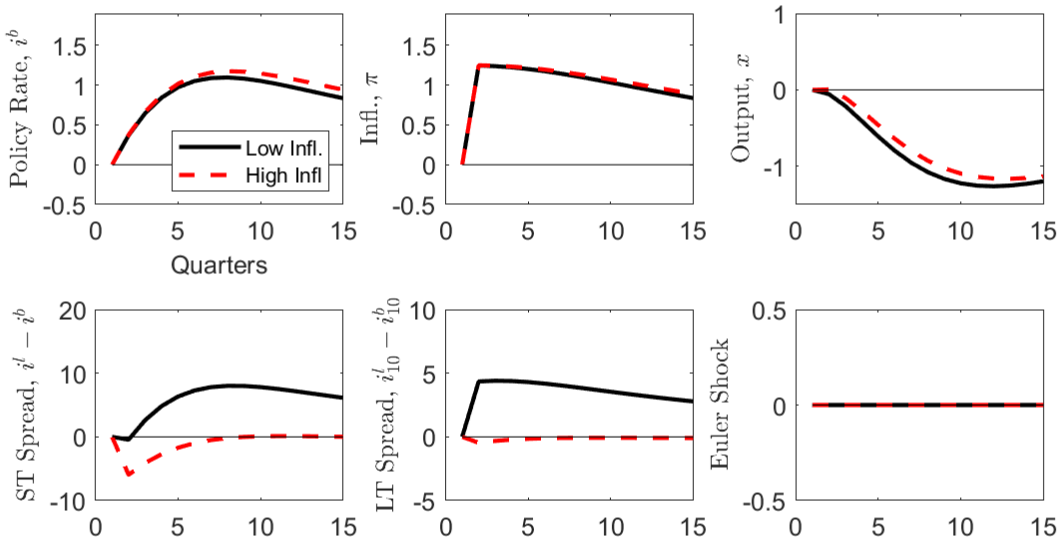
Finally, we investigate the implications of an alternative model allowing for a direct inflation-Treasury liquidity link, and show that it cannot explain our empirical findings. Specifically, by setting $b > 0$ in equation (10), we assume that inflation can directly undermine the liquidity benefits of 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 6 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 6 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.[29] 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 6 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-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

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

Figure 6. 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.



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.

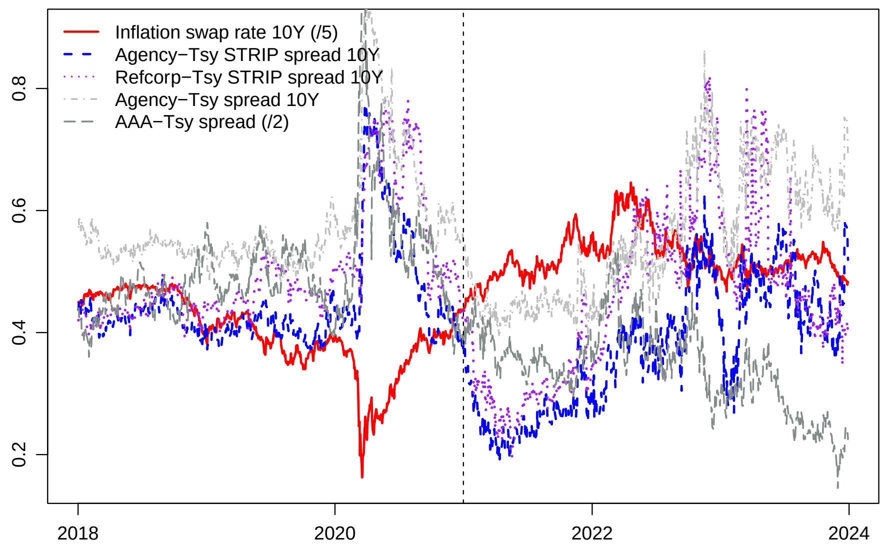
Figure 7 superimposes various measures of long-term convenience against the 10-year inflation swap rate during the period from 2018 to 2023. While the correlation between Treasury convenience and market-implied inflation expectations was strongly negative during and immediately following the initial 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.[30] Even though the levels of convenience differ, Figure 7 shows that their negative comovement with the inflation swap rate largely disappears from 2021 onward across all proxies.

In Table 8, 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-

[30]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.

Figure 7. 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[31] 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

[31]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.

Table 8. 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, $N = 36$								
$\mathbb{E}_{1q}^{\text{BC}}$(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

6 Conclusion

This paper argues that two competing mechanisms driving Treasury bond convenience – the "money channel" and the "liquidity demand channel" – dominated over distinct historical periods, leading to secular shifts in the comovement between Treasury convenience and inflation. Using a century of data, we show that during the 1970s and 1980s, higher inflation was robustly correlated 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 in 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 empirical long-term convenience-inflation relationship during the second half of the 20th century. 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 experience of the early 20th century and the 2000s.

While intuition might suggest that episodes of high inflation deplete the convenience benefits of Treasuries, this intuition does not accurately describe the historical experience to date. Our results highlight a more complex link between Treasury convenience and the macroeconomy through the interplay of money and liquidity demand channels, with substantial periods dominated by a positive inflation-convenience relationship through the money channel.

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Online Appendix to
"Inflation and Treasury Convenience"

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A Data and Robustness of Empirical Results

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

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 $\mathbf{E}$[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

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

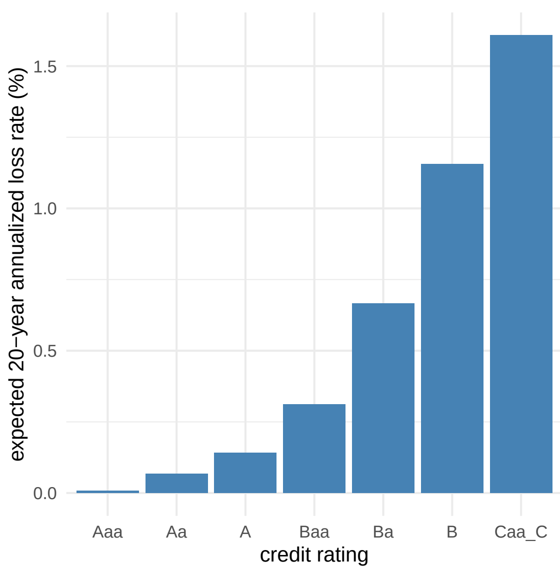
Then the annualized loss rate in a 20-year horizon is

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

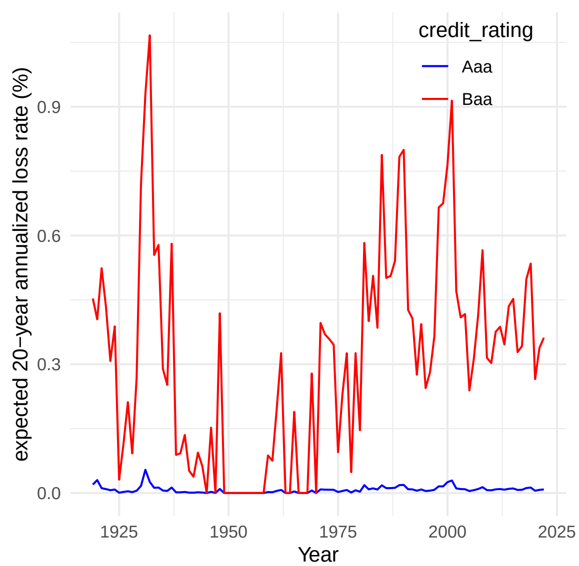
In the data, the realized default rate varies a lot across different years, while loss given default and the transition probabilities across ratings are more stable. The change in realized default rate could influence the market perception of default risks. To reflect such dynamics, we use the realized 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.

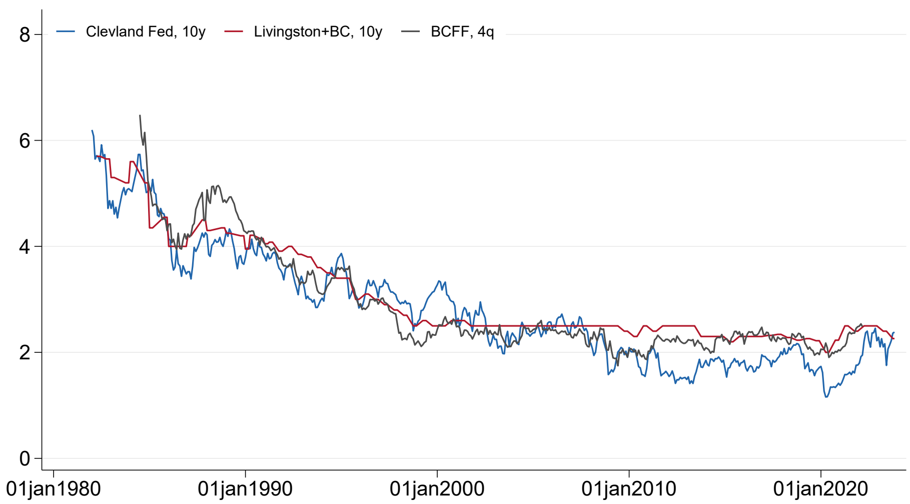
A.3 The Cleveland Fed index, inflation expectations, and term premia

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.

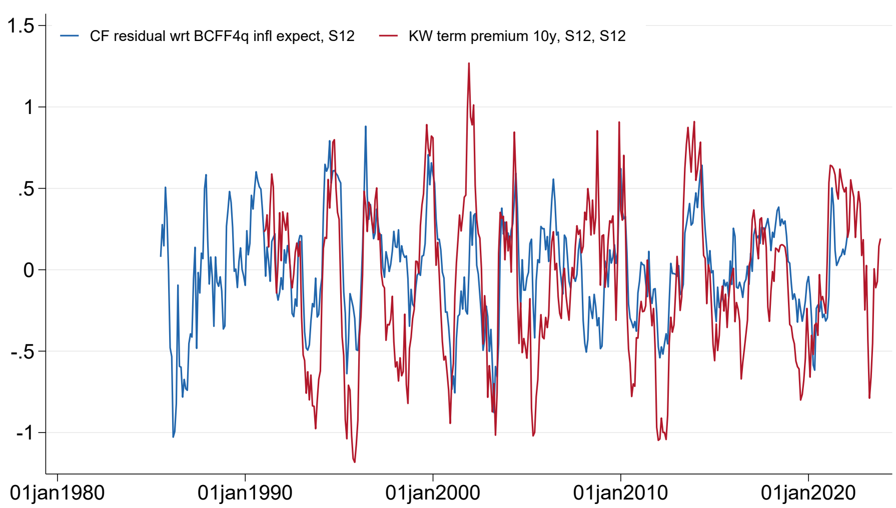
Panel A of Figure A2 shows 10-year inflation expectations from the Cleveland Fed against long-term and 4-quarter consensus inflation expectations from surveys. It is clear that the Cleveland Fed inflation expectations roughly move similarly at lower frequencies, but are substantially more volatile, raising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

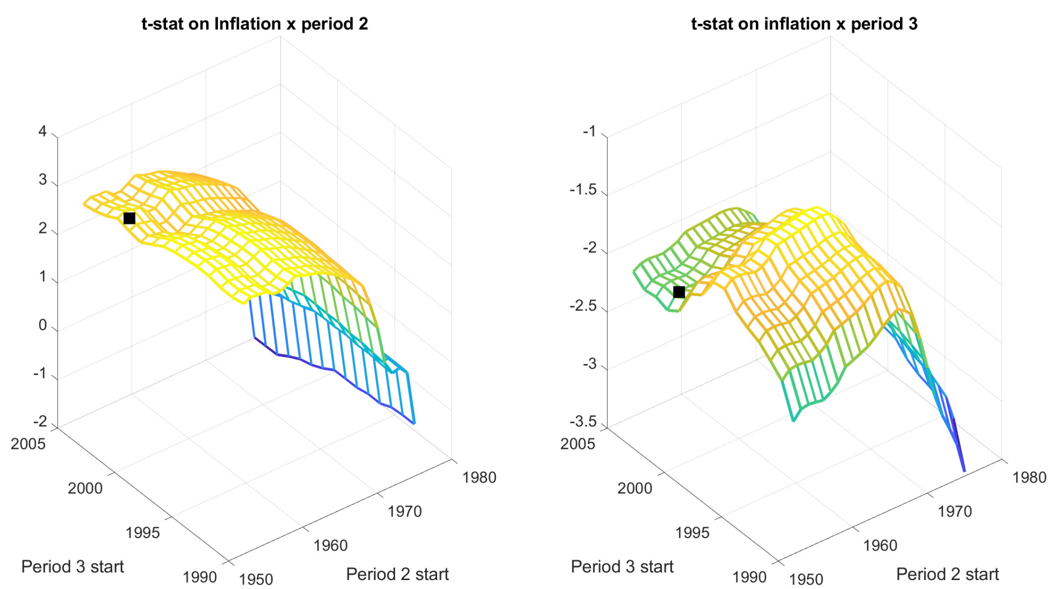
	Tbill spread			AAA-Tsy spread		
	(1)	(2)	(3)	(4)	(5)	(6)
Inflation	-0.029***	-0.027***	-0.027***	-0.038***	-0.038***	-0.038***
	(-4.69)	(-5.95)	(-5.83)	(-5.15)	(-5.31)	(-5.24)
Inflation x $I_{1952-1999}$	0.15***	0.082***	0.045*	0.13***	0.12***	0.088**
	(9.04)	(3.95)	(1.91)	(6.45)	(3.87)	(2.42)
Inflation 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

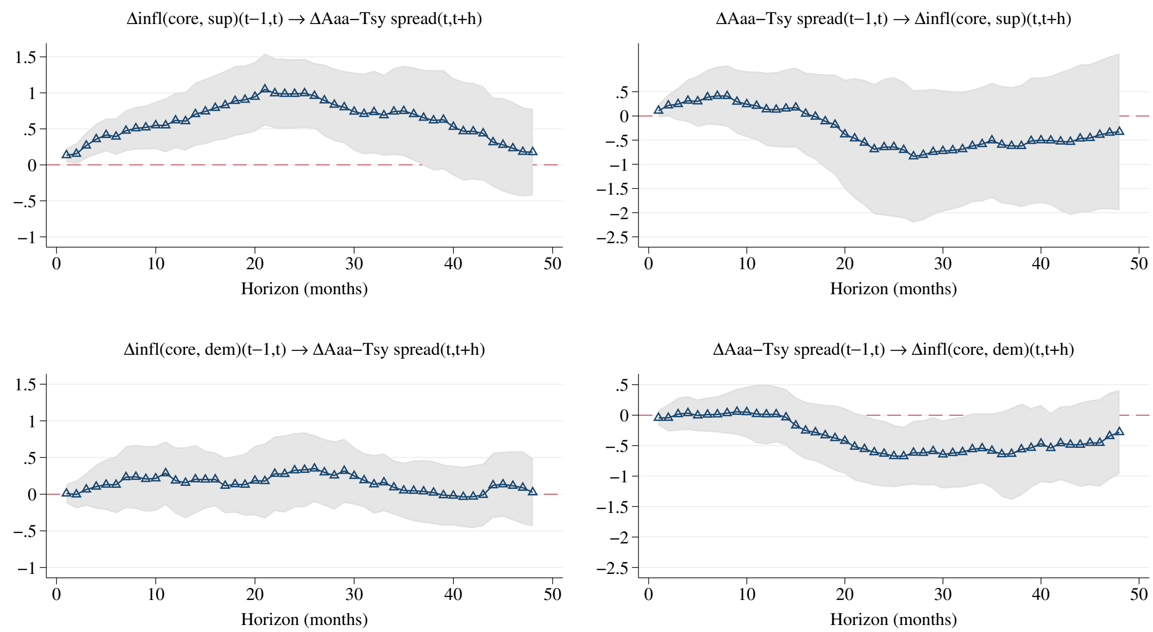
In this section, we will show the full set of estimates on predictive regressions between Treasury convenience yield and the supply and demand components of inflation in Figure A4, which encompasses subplots in Figure 3 of the main text. The main message is that convenience yields only predict the demand 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.

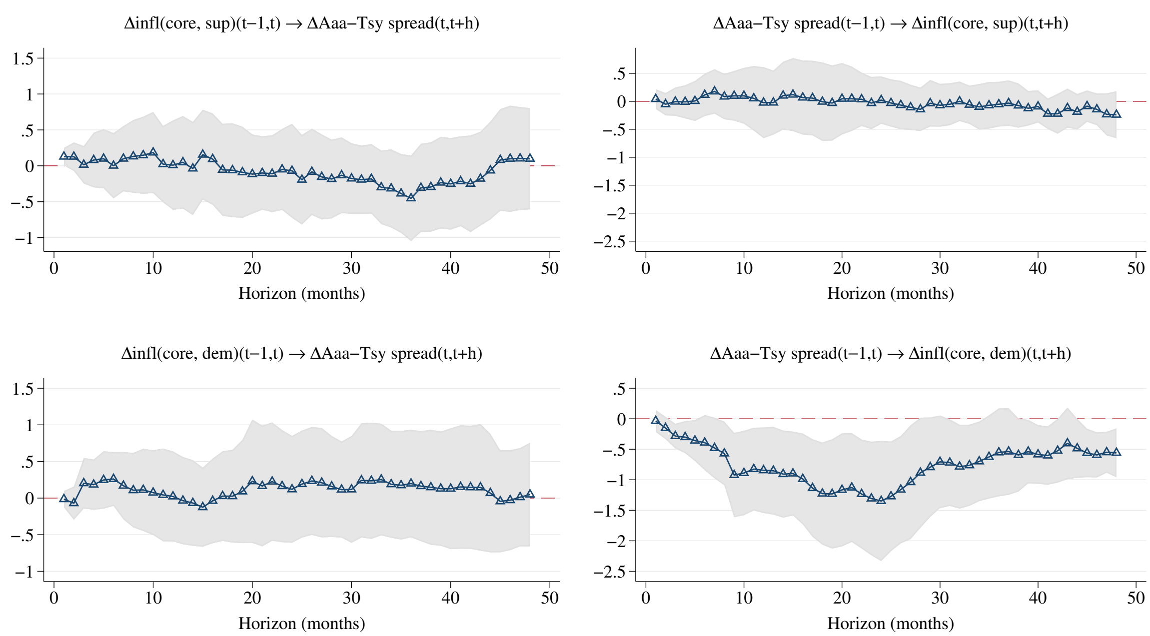
In Panel B of Figure 3, we find that the only significant relationship is that higher convenience yield leads to a lower demand component of inflation, consistent with the liquidity demand channel. The insignificance of the supply component of inflation leading to higher convenience indicates that there 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. We start the baseline model, and then provide details on the extension model that directly introduces the inflation-convenience linkage. Appendix B.3 solves a model extension, where Treasury bonds and deposits are imperfect substitutes, rather than perfect substitutes as in our baseline.

B.1 Supply Side and Price-Setting Frictions

The consumption aggregate is given by

$$C_t = \left(\int_0^1 C_{jt}^{(\sigma_t - 1)/\sigma_t} \right)^{\frac{\sigma_t}{\sigma_t - 1}}, \tag{A2}$$

where C_{jt} denotes the quantity consumed of consumption good of variety j, and σ_t is the potentially time-varying elasticity of substitution across varieties, which will give rise to supply-type shocks in the log-linearized Phillips curve (Woodford (2003a), p. 451). Household optimization then implies that demand for variety j is downward-sloping in its price P_{jt}

$$C_{jt} = \left(\frac{P_{jt}}{P_t} \right)^{-\sigma} C_t, P_t = \left(\int_0^1 P_{jt}^{1-\sigma_t} \right)^{1/(1-\sigma_t)}. \tag{A3}$$

We set up the firm's problem as simple and standard as possible. Firms face price-setting frictions in the manner of Calvo (1983). We assume that there is a unit mass of firms producing consumption good j. Firms of type j have a constant returns to scale production technology and use labor as their only input

$$Y_{jt} = N_{jt}. \tag{A4}$$

Each period a random fraction $1 - \omega$ of firms is allowed to adjust prices, while the remaining fraction ω of firms have a price that is automatically indexed to lagged inflation. That is, the time $t + \tau$ price of a firm that last re-set its product price to P_t^* at time t equals $P_t^* \left(\frac{P_{t-1+\tau}}{P_{t-1}} \right)^{\zeta}$, where ζ is an indexation parameter as in Christiano et al. (2005) and leads to a backward-looking term in the Phillips curve. There is no real investment in the model, so consumption must equal output for each variety

$$C_{jt} = Y_{jt}. \tag{A5}$$

B.2 Model Derivations

B.2.1 Liquidity Spread

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

$$\frac{I_t^l - I_t^b}{1 + I_t^l} = \frac{\frac{\alpha}{Q_t}\lambda_t}{U_c\left(C_t, H_t\right)} \tag{A6}$$

$$\frac{I_t^l - I_t^d}{1 + I_t^l} = \frac{\frac{\alpha}{Q_t}(1 - \lambda_t)}{U_c\left(C_t, H_t\right)}. \tag{A7}$$

Substituting in (8) into (A7) then gives (16) in the main text. A simple rearrangement gives that

$$I_t^l = \frac{1}{1 - \frac{\lambda_t}{1-\lambda_t}(1 - \delta)} I_t^b \tag{A8}$$

B.2.2 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. Before we can log-linearize we need to solve for the flexible-price steady-state. With flexible prices, profit-maximization implies all firms optimally choose to charge a constant markup

$$\frac{P_t^*}{P_t} = \frac{\sigma}{\sigma - 1}\frac{W_t}{P_t}. \tag{A9}$$

The representative household's optimal labor-leisure choice implies that the real wage must satisfy

$$\frac{\chi N_t^\eta}{(C_t - hC_{t-1})^{-\gamma}} = \frac{W_t}{P_t}. \tag{A10}$$

In the flexible-price equilibrium, we must have $P_t^* = P_t$. Substituting in good markets clearing ($Y_t = C_t$) implies that the steady-state flexible price output is constant and equals

$$\bar{Y} = \left(\frac{(1-h)^{\gamma}(\sigma-1)}{\chi\sigma}\right)^{1/(\gamma+\eta)}. \tag{A11}$$

In the steady-state interest rates must satisfy:

$$\left(1+\bar{I}^l\right)\beta E\left[\frac{U_c(C_{t+1})}{U_c(C_t)}\frac{P_t}{P_{t+1}}\right] = 1, \tag{A12}$$

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:

$$1+\bar{I}^l = \frac{1+\bar{\Pi}}{\beta}. \tag{A13}$$

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

$$\bar{I}^b = \bar{I}^l\left(1-\frac{\bar{\lambda}}{1-\bar{\lambda}}(1-\delta)\right). \tag{A14}$$

B.2.3 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

$$\phi(\lambda_t) = \frac{1}{1-\frac{\lambda_t}{1-\lambda_t}(1-\delta)}. \tag{A15}$$

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

$$\begin{aligned} \phi(\lambda_t) &\approx \phi\left(\bar{\lambda}\right) + \phi'\left(\bar{\lambda}\right)\hat{\lambda}_t, && \text{(A16)} \\ \phi\left(\bar{\lambda}\right) &= \frac{1}{1 - \frac{\bar{\lambda}}{1-\bar{\lambda}}(1-\delta)}, && \text{(A17)} \\ \phi'\left(\bar{\lambda}\right) &= \left(\frac{1}{1 - \frac{\bar{\lambda}}{1-\bar{\lambda}}(1-\delta)}\right)^2 (1-\delta)\frac{1}{\left(1-\bar{\lambda}\right)^2} && \text{(A18)} \end{aligned}$$

We can then re-write expression (A8)

$$\exp(\hat{i}^l_t + \bar{i}^l) - 1 = \phi\left(\lambda_t\right)\left(\exp(\hat{i}^b_t + \bar{i}^b) - 1\right) \quad \text{(A19)}$$

Substituting in the log-linear approximation for ϕ,

$$(1+\bar{I}^l)\hat{i}^l_t + \bar{I}^l \approx \left(\phi(\bar{\lambda}) + \phi'(\bar{\lambda})\hat{\lambda}_t\right)\left((1+\bar{I}^b)\hat{i}^b_t + \bar{I}^b\right) \quad \text{(A20)}$$

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

$$\begin{aligned} \hat{i}^l_t &\approx \phi(\bar{\lambda})\frac{1+\bar{I}^b}{1+\bar{I}^l}\hat{i}^b_t + \phi'(\bar{\lambda})\frac{\bar{I}^b}{1+\bar{I}^l}\hat{\lambda}_t, && \text{(A21)} \\ &= f^i \hat{i}^b_t + f^\lambda \hat{\lambda}_t, && \text{(A22)} \end{aligned}$$

where the coefficients are given by

$$\begin{aligned} f^i &= \phi(\bar{\lambda})\frac{1+\bar{I}^b}{1+\bar{I}^l}, && \text{(A23)} \\ &= \frac{1}{1 - \left(\frac{\bar{\lambda}}{1-\bar{\lambda}}(1-\delta)\right)}\frac{1+\bar{I}^l\left(1 - \left(\frac{\bar{\lambda}}{1-\bar{\lambda}}(1-\delta)\right)\right)}{1+\bar{I}^l} && \text{(A24)} \\ &= \frac{1}{1 - \left(\frac{\bar{\lambda}}{1-\bar{\lambda}}(1-\delta)\right)}\left(1 - \frac{\bar{I}^l}{1+\bar{I}^l}\left(\frac{\bar{\lambda}}{1-\lambda}(1-\delta)\right)\right), && \text{(A25)} \\ f^\lambda &= \phi'(\bar{\lambda})\frac{\bar{I}^b}{1+\bar{I}^l}. && \text{(A26)} \end{aligned}$$

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

$$0 = \beta E_t \left[U_c\left(C_{t+1}\right) \left(\exp(i^b_t) - \frac{\exp\left(-(n-1)i^b_{n-1,t+1}\right)}{\exp\left(-ni^b_{n,t}\right)} \right) \right]. \tag{A27}$$

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:

$$i^b_{n,t} = \frac{1}{n} E_t \left[\sum_{i=0}^{n-1} i^b_{t+i} \right]. \tag{A28}$$

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:

$$i^l_{n,t} = \frac{1}{n} E_t \left[\sum_{i=0}^{n-1} i^l_{t+i} \right]. \tag{A29}$$

We derive log-linearized Euler equation (21) following standard steps. The representative household's intertemporal first-order condition is

$$\Theta_t \left(C_t - hC_{t-1}\right)^{-\gamma} = \beta\left(1 + I^l_t\right) E_t \left[\Theta_{t+1} \frac{P_t}{P_{t+1}} \left(C_{t+1} - hC_t\right)^{-\gamma} \right] \tag{A30}$$

Log-linearizing around the flexible-price steady-state $\bar{C}$ gives up to a constant

$$\log\left(C_t - hC_{t-1}\right) \approx \frac{1}{1-h}(c_t - hc_{t-1}) \tag{A31}$$

The log-linearized consumption Euler equation then equals (up to constant)

$$\left(\theta_t - E_t\theta_{t+1}\right) - \frac{\gamma}{1-h}\left(c_t - hc_{t-1}\right) = i_t^l - E_t\pi_{t+1} - \frac{\gamma}{1-h}\left(E_t c_{t+1} - hc_t\right). \tag{A32}$$

A simple re-arrangement then gives

$$c_t = \frac{h}{1+h}c_{t-1} + \frac{1}{1+h}E_t c_{t+1} - \gamma^{-1}\frac{1-h}{1+h}\left(i_t^l - E_t\pi_{t+1}\right) + \gamma^{-1}\frac{1-h}{1+h}\left(\theta_t - E_t\theta_{t+1}\right) \tag{A33}$$

Equation (21) then follows from $x_t = c_t$ with the demand shock taking the form $v_{x,t} = \gamma^{-1}\frac{1-h}{1+h}\left(\theta_t - E_t\theta_{t+1}\right)$.

Because the labor-leisure trade-off is standard, the firm's problem is also entirely standard. Walsh (2017) provides a detailed derivation of the log-linearized New Keynesian Phillips curve (23).

The standard textbook treatment of firm decision problem will give rise to the log-linearized Phillips curve,

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

B.2.4 Solution Details

Denote scaled liquidity shock by $\xi_t \equiv -\psi f^\lambda u_t$, so that

$$\xi_t = \rho^\xi \xi_{t-1} + v_{\xi,t}, \tag{A34}$$

with $v_{\xi,t} = -\psi f^\lambda v_{\lambda,t}$ iid and serially uncorrelated and $\rho^\xi = \rho^\lambda$.

The log-linearized dynamics for the state vector $Z_t = \left[x_t, \pi_t, i_t^b, \xi_t\right]$ can then be summarized

$$x_t = (1-\rho^x)E_t x_{t+1} + \rho^x x_{t-1} - \psi f^i i_t^b + \psi E_t\pi_{t+1} + \psi b f^\lambda \pi_t + \xi_t + v_{x,t}, \tag{A35}$$

$$\pi_t = (1-\rho^\pi)E_t\pi_{t+1} + \rho^\pi\pi_{t-1} + \kappa x_t + v_{\pi,t}, \tag{A36}$$

$$i_t^b = (1-\rho^i)\left(\gamma^x x_t + \gamma^\pi \pi_t\right) + \rho^i i_{t-1}^b + v_{i,t}, \tag{A37}$$

$$\xi_t = \rho^\xi \xi_{t-1} + v_{\xi,t}. \tag{A38}$$

We only need to solve the model with either the liquidity shock ξ_t or the demand shock $v_{x,t}$. We start with the solution for the model liquidity shock ξ_t, setting the demand shock to zero. In matrix form, the model can be written as

$$0 \quad = \quad FE_t[Z_{t+1}] + GZ_t + HZ_{t-1} + Mv_t, \tag{A39}$$

where the matrices are given by

$$F \quad = \quad \begin{bmatrix} 1-\rho^x & \psi & 0 & 0 \\ 0 & 1-\rho^\pi & 0 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 \end{bmatrix} \tag{A40}$$

$$G \quad = \quad \begin{bmatrix} -1 & \psi b f^\lambda & -\psi f^i & 1 \\ \kappa & -1 & 0 & 0 \\ (1-\rho^i)\gamma^x & (1-\rho^i)\gamma^\pi & -1 & 0 \\ 0 & 0 & 0 & -1 \end{bmatrix} \tag{A41}$$

$$H \quad = \quad \begin{bmatrix} \rho^x & 0 & 0 & 0 \\ 0 & \rho^\pi & 0 & 0 \\ 0 & 0 & \rho^i & 0 \\ 0 & 0 & 0 & \rho^\xi \end{bmatrix} \tag{A42}$$

$$M \quad = \quad \begin{bmatrix} 0 & 0 & 0 \\ 0 & 1 & 0 \\ 0 & 0 & 1 \\ 1 & 0 & 0 \end{bmatrix} \tag{A43}$$

and the vector of exogenous shocks is given by

$$v_t \quad = \quad [v_{\xi,t}, v_{\pi,t}, v_{i,t}]. \tag{A44}$$

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

The convenience spread of maturity then has the following log-linearized expression:

$$spread_{n,t} \equiv i^l_{n,t} - i^b_{n,t} = \frac{1}{n}\left(\left(f^i - 1\right)e_3 - \psi^{-1}e_4 - bf^{\lambda}e_2\right)\left(I - B\right)^{-1}\left(I - B^n\right)Z_t \qquad \text{(A45)}$$

We then show impulse responses to shocks of the size $v_{\xi,t} = \frac{\psi}{400}$, $v_{\pi,t} = \frac{1}{400}$, and $v_{i,t} = \frac{1}{400}$ in natural units.

To solve the model with a generic demand shock but no liquidity shocks, note that if θ_t follows an AR(1) with autoregression coefficient ρ^{ξ} then $\theta_t - E_t\theta_{t+1} = (1-\rho^{\xi})\theta_t$ also follows an AR(1) process with the same AR(1) coefficient. The same solution then goes through, except we need to set $f^i = 1$ and $b = 0$ to obtain the impulse responses to a generic demand shock when Treasury bonds yield no liquidity.

B.3 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)

$$Q_t = \left((1-\lambda_t)D_t^{\rho} + \lambda_t B_t^{\rho}\right)^{1/\rho}, \qquad \text{(A46)}$$

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

$$I^l_t - I^b_t = \frac{\lambda_t}{1-\lambda_t}\left(\frac{B_t}{D_t}\right)^{\rho-1}(1-\delta)I^l_t, \qquad \text{(A47)}$$

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

$$\hat{i}^l_t \approx f^i \hat{i}^b_t + f^\lambda \hat{\lambda}_t - f^b \left(\hat{b}_t - \hat{d}_t\right), \tag{A48}$$

$$f^b = \frac{1}{\left(1 - \frac{\bar{\lambda}}{1-\bar{\lambda}}(1-\delta)\left(\bar{B}/\bar{D}\right)^{\rho-1}\right)^2}(1-\rho)(1-\delta)\frac{\bar{\lambda}}{1-\bar{\lambda}}\left(\bar{B}/\bar{D}\right)^{\rho-1}\frac{\bar{I}_b}{1+\bar{I}_l} \geq 0, \tag{A49}$$

$$f^i = \frac{1}{1 - \left(\frac{\bar{\lambda}}{1-\lambda}(1-\delta)\left(\bar{B}/\bar{D}\right)^{\rho-1}\right)}\left(1 - \frac{\bar{I}^l}{1+\bar{I}^l}\left(\frac{\bar{\lambda}}{1-\lambda}(1-\delta)\left(\bar{B}/\bar{D}\right)^{\rho-1}\right)\right), \tag{A50}$$

$$f^\lambda = \left(\frac{1}{1 - \frac{\bar{\lambda}}{1-\bar{\lambda}}(1-\delta)\left(\bar{B}/\bar{D}\right)^{\rho-1}}\right)^2(1-\delta)\left(\bar{B}/\bar{D}\right)^{\rho-1}\frac{1}{\left(1-\bar{\lambda}\right)^2}\frac{\bar{I}^b}{1+\bar{I}^l} \tag{A51}$$

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

$$x_t = \rho^x x_{t-1} + (1-\rho^x)E_t x_{t+1} - \psi\left(f^i i^b_t - E_t \pi_{t+1}\right) - \psi\left(f^\lambda \hat{\lambda}_t - f^b\left(\hat{b}_t - \hat{d}_t\right)\right) + v_{x,t} \tag{A52}$$

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.4 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

$$E_t\left[M_{t+1}^{\$}\right]\left(I_t^l - I_t^b\right) = \frac{\alpha_t/Q_t\lambda_t}{U_c\left(C_t, Q_t, H_t, N_t, \Theta_t\right)}. \tag{A53}$$

Combining equations (13) and (15) gives

$$E_t\left[M_{t+1}^{\$}\right]\left(I_t^l - I_t^d\right) = \frac{\alpha_t/Q_t(1-\lambda_t)}{U_c\left(C_t, Q_t, H_t, N_t, \Theta_t\right)}. \tag{A54}$$

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 (A53) into (A54) 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 (A53) and (A54) 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.5 Details on Model Calibration

Details about model calibration are shown in Table A2.

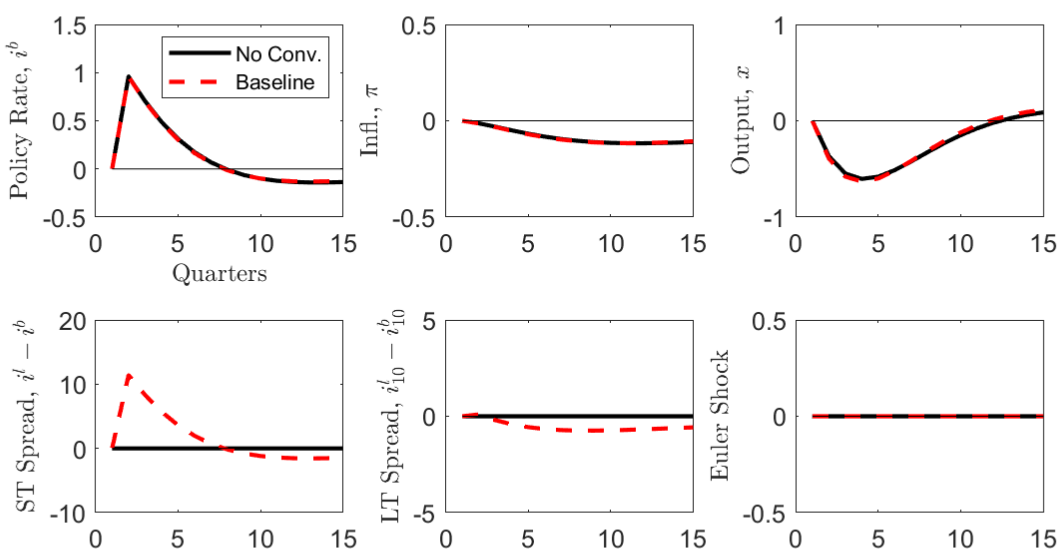
Table A2. Model Calibration. This table contains the calibration parameters for the New Keynesian model with convenience yields. Parameters are reported in units corresponding to inflation and interest rates in annualized percent, and output gap in percent, that is we report $\frac{\psi}{4}$, 4κ and $4\gamma^x$ 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 (1997)
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

B.6 Additional Model Results

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.

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.



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 6 in the main paper. The black 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 black 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. Our key empirical result is 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 6 in the main paper but sets $b = 0.01$.

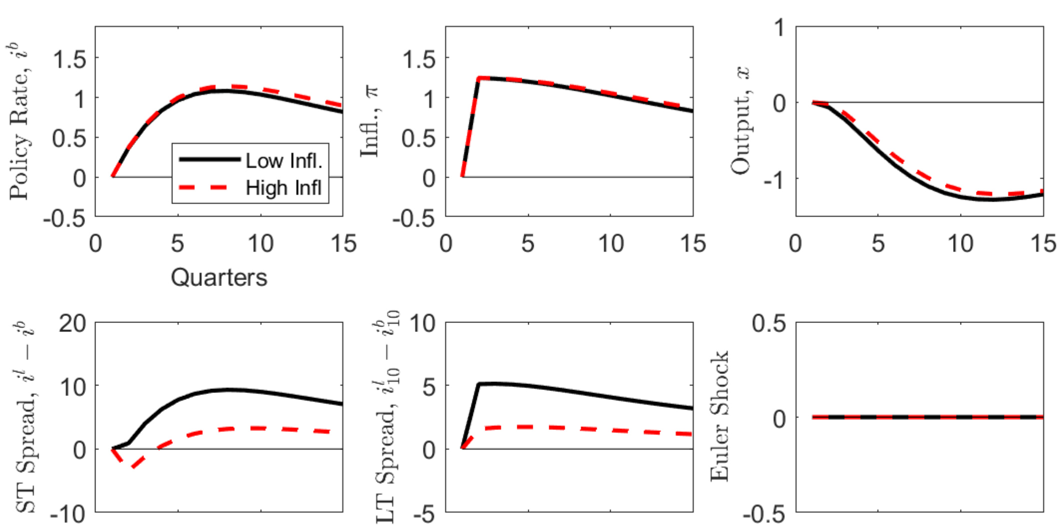


Figure A7. Alternative Model Responses to Cost-Push Shock with $b = 0.02$. This figure is identical to Figure 6 in the main paper but sets $b = 0.02$.