Financial Conditions in the Real Yield Curve

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Abstract

I show that a positive average real term premium can be rationalized by transitory shocks to financial conditions affecting the real economy. Tight financial conditions—defined as states in which financial intermediaries' risk-bearing capacity is low and investors have an unusually strong demand for highly liquid nominal assets—tend to raise the level of real yields, flatten the real yield curve, and increase expected excess returns on real bonds. I rationalize this evidence with a parsimonious consumption-based term structure model in which the endowment is driven by permanent shocks, while transitory shocks to financial conditions distort the consumption process.

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> *The above eight changes cause (9)* Complicated disturbances in the rates of interest*, in particular, a fall in the nominal, or money, rates and a rise in the real, or commodity, rates of interest.*
>
> Fisher (1933, p. 342), describing the nine points that characterize financial crises.

Nearly a century ago, Irving Fisher characterized financial crises as episodes in which the short-term real interest rate (i.e., the interest rate denominated in goods or "commodities" units) would rise, even though the short-term nominal interest rate and inflation would fall. I begin this paper by providing evidence that verifies Fisher's prediction linking the dynamics of the short-term real interest rate to financial conditions. I then argue that this prediction is helpful for understanding the more recent experience observed in the U.S. real yield curve. My main finding is that tight financial conditions—defined as states in which financial intermediaries' risk-bearing capacity is impaired and agents in the economy have an unusual preference for highly liquid nominal financial instruments instead of physical goods/commodities—are associated with transitory increases in real rates, higher real term premia, and a flatter real yield curve. Based on this finding, the central idea of this paper is that a positive real term premium (i.e., an average upward-sloping real yield curve) can be rationalized by the presence of shocks to financial conditions having a transitory effect on the real economy.

Figure 1 shows the mean paths of the short-term real interest rates around financial crises (panel A) and around normal recessions (panel B). I consider 14 advanced economies from 1870 to 2020 using data from Jordà, Schularick and Taylor (2017). Section 1 provides more details about the data, as well as a more formal statistical analysis using local projections. The key takeaway from the figure is that the short-term real rate experiences a short-lived increase during financial crises that is clearly not observed in normal recessions, in line with Fisher (1933)'s predictions.[1] As discussed in Section 1 below, the transitory increase in the real rate is statistically significant and different from the path observed during normal recessions for about two years. This evidence is similar in nature to the more pronounced increase in the dividend-price ratio over financial recessions than in normal recessions, also for about 2 years (see Figure II in Muir (2017)).

[1]The ex-ante real rate is constructed using inflation expectations from an autoregressive model for inflation, given the absence of long-run market- or survey-based expectations. The ex-ante real rate increases for a few more periods than the ex-post real rate because of the persistence implied by the autoregressive model for inflation.

Motivated by the evidence about the dynamics of the real interest rate around periods of tight financial conditions, I then zoom in on higher-frequency evidence in a more recent sample, using prices of Treasury Inflation-Protected Securities (TIPS), the inflation-linked bonds issued by the U.S. Treasury. The objective is to use real yields to test the correlations predicted by Fisher beyond periods of extremely tight financial conditions (such as financial crises). I construct the first principal component of a measure that captures intermediaries' risk-bearing capacity and a measure that captures investors' desire to hold highly liquid nominal assets. My measure is intended to capture Fisher's idea that tight financial conditions are events in which there is a contraction in the flow of credit among agents—as intermediaries' balance sheets are impaired—together with an increased preference toward liquid nominal financial instruments. As noted in Fisher (1933), the abnormal demand for nominal liquid assets can be related to the deflationary forces in periods of financial stress.

To capture intermediaries' risk-bearing capacity, I use the market value of primary dealers' capital ratio, proposed by He, Kelly and Manela (2017). As discussed in He et al. (2017), this measure is tightly connected to theories in which financial intermediaries distort the allocation of risk and capital in the economy (see, e.g., He and Krishnamurthy (2013), Brunnermeier and Sannikov (2016), Gertler and Kiyotaki (2015), among others). To capture investors' desire for liquid nominal financial instruments, in my baseline estimation, I use the changes in the 5- and 10-year on-the-run/off-the-run (on/off) spreads in nominal bonds. I then construct my empirical measure as the first principal component of intermediaries' risk-bearing capacity and the on/off spread from nominal yields. I show that this principal component can be interpreted as the fraction of variation in intermediaries' risk-bearing capacity that is spanned by the on/off spreads and their lags. I use my empirical measure as a proxy for states in which intermediaries' risk-bearing capacity is low and investors display a relatively higher preference for highly liquid nominal assets. I then conduct several empirical exercises that I summarize next.

First, I regress the first and second principal components (PC) of the TIPS yield curve, in levels and in differences, on my empirical proxy for financial conditions.[2] I find that tighter financial conditions are associated with a higher level of real yields and a flatter real yield curve, even after removing the 2008–2009 global financial crisis from the sample. I show that these effects are not driven by proxies of TIPS liquidity (such as bid-ask spreads or fitting errors from parametric yield

[2]The first and second PC of the TIPS yield curve are the level and slope factors, respectively. They explain 96% of TIPS variation.

curve models).[3] I then regress individual TIPS yields onto my empirical measure of financial conditions and find that the front-end of the curve is significantly more sensitive to financial conditions than the long-end of the curve. In particular, the coefficient associated with financial conditions is twice as large for the 2-year yield as for the 10-year yield. This evidence illustrates how financial conditions move the level and the slope of the curve: They are mostly transitory shocks and, as a result, have a relatively larger effect in the front-end of the curve. Motivated by the evidence that financial conditions move the front-end of the curve relatively more (and hence the slope), I regress TIPS excess returns onto my proxy of financial conditions. Tight financial conditions are associated with higher expected excess returns on TIPS, and the effect is more pronounced in long-term bonds (because they carry a higher duration risk). As a result, long-term rates move less than short rates when financial conditions tighten because the relatively larger increases in the term premium are offset by the expectations that the short rate will decline in the near term. Again, this evidence can be rationalized by the fact that shocks to financial conditions are mostly transitory.

I conclude the empirical analysis by studying how financial conditions correlate with the real economy. In particular, I focus on understanding how my measure of financial conditions (which I denote by x_t throughout the text) correlates with expected growth and the conditional volatility of consumption growth. I find that a tightening of financial conditions is associated with higher conditional volatility as well as higher expected consumption growth. Intuitively, this evidence is consistent with financial conditions having a transitory effect on consumption: a tightening reduces consumption but this reduction is short-lived and, as a result, the level of consumption is expected to recover in the subsequent months.

I rationalize the empirical analysis using a parsimonious consumption-based term structure model. The central assumption in the model is that consumption is distorted by the presence of transitory changes in financial conditions. I model such distortions in reduced form, but I follow closely the discipline imposed by the market-clearing conditions that prevail in general equilibrium models featuring financial intermediaries—see, e.g., Brunnermeier and Sannikov (2014), Gertler and Kiyotaki (2015) and Schneider (2026). More precisely, I assume that the endowment follows a random walk but consumption is distorted by transitory shocks to financial conditions. Tight financial conditions translate into higher deadweight losses in consumption, which manifest in a

[3] Pflueger and Viceira (2016), D'Amico, Kim and Wei (2018), Andreasen, Christensen and Riddell (2021) estimate a time-varying premium in the relative value of TIPS and nominal Treasuries. In this paper, I do not focus on the relative valuation of TIPS and their nominal counterparts. Instead, I focus on controlling for TIPS liquidity in an absolute sense, using measures such as the bid-ask spread in TIPS securities.

lower consumption level and a higher conditional volatility of consumption growth. Because shocks to financial conditions are transitory, the low level of consumption is expected to last for some time, which means agents expect consumption to grow back to a higher level in the future, once financial conditions normalize. As a result, shocks to financial conditions produce a decline in consumption and an increase in *expected* consumption growth as well as in the conditional volatility of consumption growth, and these two forces push the real interest rate in opposite directions. On the one hand, the increase in expected consumption growth pushes the real rate up through the intertemporal smoothing force: agents see that the low level of consumption will be higher in the near horizon once financial conditions loosen, and therefore have a preference to smooth the shock by borrowing from the future, exerting upward pressure on the real rate. On the other hand, the increase in the volatility of consumption growth pushes the real rate down through the precautionary savings motive: agents seek to increase their savings when facing a more uncertain consumption path. On net, the interest rate rises, together with the price of risk, when the intertemporal smoothing force dominates over the precautionary savings. The increase in the real rate in states in which financial conditions are tight (thus, the price of risk is high) causes real bonds to be risky assets, thus carrying a premium in equilibrium.

Related Literature. The findings in this paper relate to two broad strands of literature. First, many papers have underscored, both empirically and theoretically, the role of financial intermediation for understanding fluctuations in aggregate asset prices and macroeconomic quantities (Brunnermeier and Sannikov, 2014; He and Krishnamurthy, 2013; Adrian, Etula and Muir, 2014; Gertler and Kiyotaki, 2015; He et al., 2017; Haddad and Muir, 2021; Elenev, Landvoigt and Nieuwerburgh, 2021; Haddad and Muir, 2025). I contribute to this literature by underscoring the importance of shocks to financial intermediaries' risk-bearing capacity for understanding the real yield curve. In particular, I show that transitory episodes of tight financial conditions are an important driver of the real term premium. The key assumption behind my results is that there are deadweight losses in consumption when financial conditions are tight—an assumption that is in line with the literature. In addition, I propose an empirical measure that correlates with key metrics of the real yield curve (first and second PC, as well as expected excess returns on real bonds). The empirical measure is motivated by Fisher (1933)'s characterization of periods of tight financial conditions, in which intermediaries' risk-bearing capacity is impaired and investors display an abnormal demand for highly liquid nominal assets.

Second, this paper contributes to another body of literature, traced back to Backus, Gregory

and Zin (1989), which has shown the difficulties in generating a positive real term premium in representative agent economies. One possible solution for generating a positive real term premium is to introduce exogenous fluctuations in the representative agent's preferences (Wachter, 2006). Instead, in this paper I assume that the endowment follows a random walk but consumption is distorted by transitory shocks to financial conditions. Thus, this paper provides a simple and parsimonious way to link financial conditions to an otherwise standard consumption-based asset pricing model, which can be extended in multiple ways. In fact, I show that a simplification of the model can be cast as a quadratic term structure model, which is a framework that has been widely studied in the reduced-form term structure literature (see, e.g., Ahn, Dittmar and Gallant (2002)). Other papers in the representative agent literature have underscored the role of transitory shocks to consumption as a plausible explanation for a positive real term premium, but my paper connects those transitory shocks to financial conditions.[4] I make the connection between financial conditions and the real term premium in a parsimonious but disciplined way, which is by using the standard market-clearing conditions that prevail in more elaborate general equilibrium models featuring financial intermediaries and savers such as Brunnermeier and Sannikov (2014), Gertler and Kiyotaki (2015), and Schneider (2026).

The paper is organized as follows. Section 1 presents the motivating evidence that supports Fisher (1933)'s predictions. Section 2 presents the empirical exercises using the TIPS yield curve. Section 3 presents a representative-agent model to rationalize the connection between financial conditions and the yield curve. Section 4 concludes. The Appendix contains proofs, robustness checks, and data description.

1 Motivating Evidence

In this section I provide evidence about the dynamics of the short-term real interest rate around financial crises and compare it against normal recessions. The source of the data is Jordà et al. (2017) and it spans from 1870 to 2020 across 14 advanced economies. The countries I consider are Australia, Canada, Denmark, France, Germany, Italy, Japan, the Netherlands, Norway, Spain, Sweden, Switzerland, the United Kingdom, and the United States. The event dates for financial crises are in the dataset Jordà et al. (2017), which defines financial crises as events "characterized by major

[4] See, e.g., Labadie (1994), Alvarez and Jermann (2005), Bletzinger, Lemke and Renne (2025), Chernov, Lochstoer and Song (Forthcoming).

bank failures, banking panics, substantial losses in the banking sector, significant recapitalization, and/or significant government intervention." I use the list of normal recessions from Table 1 in Jordà, Schularick and Taylor (2013), who calculate business cycles for these 14 advanced economies and characterize the recessions as "normal" or "financial." In the appendix, I show the results are robust to using the event dates from Baron, Verner and Xiong (2021), for "bank equity crises," "panic banking crises," or either of the two.

As discussed in the introduction, Figure 1 plots the raw data for the short-term real interest rate. In particular, the figure shows mean paths for the ex-post and ex-ante real interest rate around financial crises (in Panel A) and around normal recessions (in Panel B) between 1870 and 2020. To construct the ex-ante real rate, because there is no data available for inflation expectations in such a long period, I construct expected inflation for each country using an autoregressive model for inflation and compute the expectations based on such model. I focus on narrow windows around the events (two years before and after), to illustrate the short-run dynamics of the real rate at the height of the episodes, especially movements that may be transitory. Paths are averaged with equal weight across countries over the window $t-2$ to $t+2$. I extend the window when conducting local projections below. The salient feature of the chart is that the real rate behaves very differently across the two types of recessions. In the year after a financial crisis begins, the real rate increases temporarily; the increase is especially pronounced for the ex-post rate, reflecting the sharp decline in realized inflation, and is also present, though milder, for the ex-ante rate. In contrast, during normal recessions the real rate declines.

To formally test for the patterns documented in the previous paragraph, I run

$$\Delta^{(s)} r_{i,t} = \alpha_i + \beta_F^{(s)} 1_{F,i,t} + \beta_N^{(s)} 1_{N,i,t} + c_i 1_{(t>1971)} + \varepsilon_{i,t}^{(s)} \tag{1}$$

where $r_{i,t}$ is the real rate in country i at time t, and $\Delta^{(s)} r_{i,t} = r_{i,t+s} - r_{i,t-1}$, for different s. I use country-specific fixed-effects (α_i) and add a dummy for the sample starting in 1972 (denoted by $1_{(t>1971)}$). The coefficients of interest are β_F and β_N, which load on the dummy associated with financial crises and normal recessions, respectively.

Figure 2 shows the results. The left panel shows the responses of the short-term real rate around financial crises (in blue) and around normal recessions (in red dashed). As shown, the real rate has a transitory increase of about 3 percentage points following the financial crises (and declines slightly after normal recessions). The increase in the short-term real rate following financial crises lasts for

about 2 years. The right panel in the Figure shows the difference in the response of the short-term real rate during financial crises against normal recessions, showing that the cumulative responses are statistically different for about two years.

All in all, the evidence validates Fisher (1933)'s hypothesis regarding the transitory increases in the real rate around episodes with tight financial conditions. These episodes are characterized by deflationary forces that may manifest as agents' increased desire for nominal instruments.[5]

2 Empirical Analysis

In this section I focus on a more recent period than in the previous section, to exploit the evidence related to real bonds (Treasury inflation-protected securities or TIPS) in the United States. The objective is to understand how real rates correlate with financial conditions, not to establish causality between them. Indeed, in general equilibrium models, real rates and financial conditions would be pinned down endogenously in equilibrium. To study this correlation, I begin by constructing a measure that captures both intermediaries' risk-bearing capacity (which tends to be low in episodes of financial stress) and investors' desire for liquid nominal assets (which tends to be high in episodes of financial stress). These two features—intermediaries with low risk-bearing capacity and investors having high demand for liquid nominal instruments—are the core of Fisher's prediction discussed in the previous section.

To gauge intermediaries' risk-bearing capacity, I use the shocks to the market value of financial intermediaries' capital ratios. This measure was proposed by He et al. (2017) (HKM hereafter).[6] As described in HKM, the market value of intermediaries' capital ratio is tightly connected to general equilibrium models in which the wealth of financial intermediaries is a key state variable driving aggregate asset prices and quantities (see, e.g., Brunnermeier and Sannikov (2014); He and Krishnamurthy (2013); Di Tella (2017); Schneider (2026)). In such models, when intermediaries' wealth is low, intermediation capacity declines and economic activity contracts. As a result, through the lens of those theoretical models, the HKM measure is ideal to understand the extent to which credit is flowing throughout the economy.

To gauge investors' appetite to hold highly liquid nominal assets, in my baseline, I use the spread

[5] I show in the Appendix the mean paths around financial crises and normal recessions for consumption, inflation, and nominal rates.

[6] HKM defines financial intermediaries as the primary dealers, counterparties of the New York Federal Reserve.

between on-the-run and off-the-run (on/off) nominal Treasuries. In particular, I focus on changes in the on/off spreads in the 5-year and 10-year nominal Treasuries, although I extend the analysis to include short-term maturities (or "near-money" assets) below. The on/off spreads represent the difference in the yield between most recently issued government nominal bonds (i.e., on-the-run) and those that have been previously issued and have the exact same cash flows (i.e., off-the-run). The on/off spreads have been widely documented by prior studies and several papers have provided different explanations for this phenomenon (see, e.g., Krishnamurthy (2002) or Pasquariello and Vega (2009)). In this paper, I do not take a stand on the sources of the spread. Instead, I only use it as an observable variable that captures investors' desire to hold liquid nominal assets. In this sense, I interpret a widening in the on/off spreads as investors having a relatively larger desire to hold liquid nominal assets.

I compute the first principal component of HKM and changes in the 5- and 10-year on-the-run/off-the-run spreads, and denote the resulting series by x_t. Since principal components are identified only up to sign, I normalize x_t so that higher values correspond to looser financial conditions: an improvement in intermediaries' risk-bearing capacity and/or a narrowing of on/off spreads raises x_t. With this normalization, the loadings on HKM, the 5-year on/off spread, and the 10-year on/off spread are 0.58, -0.55, and -0.60, respectively. The first principal component explains 50% of the variation in the three series. The blue line in panel A of Figure 3 shows the time series. I also show the fitted values of HKM using the on/off spreads and their lags. More precisely, I project HKM onto the on/off spreads and their three lags, and I show such projection in the red-dashed line. The correlation between the projection and the PC is 81%. Hence, one could interpret the PC as if it were the fraction of intermediaries' risk-bearing capacity that is spanned by variation in the on/off spreads. Panel B compares the baseline x_t with an alternative version that replaces the 5-year on/off spread with the spread between the 3-month uncollateralized repo rate and the 3-month T-bill. The spread between the 3-month uncollateralized repo rate and the 3-month Treasury bill was proposed by Nagel (2016) as a proxy to capture abnormal demand for highly liquid "money-like" assets. Thus, the interpretation of x_t is similar to the baseline construction but broadened to include "money-like" financial instruments that are concentrated in the front-end of the maturity spectrum of nominal assets.

Real Yields, Forwards, and x_t. I begin by regressing TIPS yield-curve PCs on x_t. I summarize TIPS variation using the first and second PCs of the cross-section of yields. As is well known, the first PC of the yield curve reflects the so-called "level factor" (i.e., the overall level of real yields),

while the second PC of the yield curve reflects the so-called "slope factor" (i.e., the slope of the yield curve, or the difference between long- and short-term real rates, orthogonal to the level). The first and second PCs are, by definition, orthogonal. Table 1 shows the results from the following regression:

$$z_t = \alpha + \beta_0 x_t + \sum_{h=1}^{3} \beta_h x_{t-h} + \sum_{h=1}^{3} \gamma_h z_{t-h} + \varepsilon_t, \tag{2}$$

where z_t is the outcome variable, namely either the first or second PC of TIPS yields, in levels or differences. The focus of the table is primarily β_0, which captures how the TIPS PCs correlate with x_t. I add three lags of x_t and z_t as controls. The key takeaway from the table is that a lower x_t is associated with a higher first PC of TIPS and a lower second PC. Put differently, tight financial conditions—impaired intermediaries' risk-bearing capacity and investors displaying abnormal demand for highly liquid nominal assets—are associated with higher real rates and a flatter real yield curve. Note that the correlations are largely robust to excluding the 2008–2009 global financial crisis (GFC) from the sample, although their magnitude and statistical significance weaken somewhat relative to the full sample, as expected. This is unsurprising, since the GFC is a key episode with a sharp tightening of financial conditions; removing it from the sample naturally weakens the estimated correlation. In Table A.1 in the Appendix, I report the results of regression (2) using each component of x_t separately. In short, the 10-year on/off spread is mostly associated with PC1, while HKM and the 5-year on/off spread are mostly associated with PC2. Taken together, the first principal component of these measures—that is, the common variation among the three series—is associated with movements in both the level and the slope of real yields. As discussed above, this principal component can be interpreted as the component of HKM that is spanned by the on/off spreads.

To further investigate the dynamic relationship between x_t and the PCs of TIPS, I estimate (2) using local projections. That is, I estimate

$$z_{t+s} = \alpha^{(s)} + \beta_0^{(s)} x_t + \sum_{h=1}^{3} \beta_h^{(s)} x_{t-h} + \sum_{h=1}^{3} \gamma_h^{(s)} z_{t-h} + \varepsilon_t^{(s)}, \tag{3}$$

for different horizons $s = 0, \ldots, 12$. Figure 4 shows the local projection estimates for the coefficient of interest, $\beta_0^{(s)}$. The left (right) plot shows the results when z_t is the first (second) PC. A reduction in x_t (tighter financial conditions) increases the first PC while it reduces the second PC. In other words, tighter financial conditions are associated with higher real yields and a flatter real yield curve.

Note that the effects are statistically significant for approximately 3 months in both cases, and thus are relatively transitory in nature.

Controlling for liquidity. Prior literature has studied the liquidity properties of TIPS relative to nominal Treasuries (see, e.g., D'Amico et al., 2018; Pflueger and Viceira, 2016; Andreasen et al., 2021, among others). From such studies, it is hard to identify whether TIPS are becoming illiquid in certain states or if, instead, the nominal Treasuries are actually becoming special at the same time. Rather than using measures of relative liquidity, I focus on measures of TIPS liquidity that do not depend on the dynamics of nominal Treasuries. In particular, I next run the specification (2), including TIPS bid-ask spreads to control for potential episodes in which the TIPS market became illiquid.[7] As shown in Table 2, x_t remains strongly negatively correlated with the level of real rates and positively correlated with the slope of the real yield curve—tight financial conditions are correlated with a higher level of real rates and a flatter real yield curve.

One additional observation is worth noting regarding TIPS liquidity. TIPS at shorter maturities are typically less liquid than TIPS at longer maturities. All else equal, this maturity pattern should push short-maturity TIPS yields up relative to long-maturity TIPS yields, biasing the slope downward. Therefore, if TIPS illiquidity were the main driver of the real yield curve, it would tend to work against the observed positive average slope.

Individual yields. I next regress individual yields and forwards onto x_t.[8] That is, I run the same specification (2) as I used for the PCs, but using the yields and forwards of different maturities as the outcome variable. Figure 5 shows the results. As was the case with the PCs, a tightening of financial conditions (lower x_t) increases real yields across the curve. The results are more pronounced in yields with shorter maturities: The coefficient is approximately twice as big in the 2-year TIPS as in the 10-year TIPS. This result shows that a lower x_t correlates with an overall increase in the level of real yields but a flattening of the real yield curve. The right panel shows that the results are qualitatively similar for forwards: A reduction in x_t correlates with higher overall level of forward rates and the effect is smaller for longer-dated forwards. The evidence on yields and forwards indicates that changes in x_t are transitory and, as a result, they have a relatively smaller effect on longer-dated bonds.

[7] The availability of bid-ask spreads starts in May 2008. Therefore, the sample is somewhat smaller than the full sample. Using fitting errors from a parametric yield curve model as a proxy for illiquidity gives similar results to the bid-ask spread.

[8] The log forward rate between periods $t+\tau-1$ and $t+\tau$ is $f_t^{(\tau)} \equiv p_t^{(\tau-1)} - p_t^{(\tau)}$, where $p_t^{(\tau)}$ is the log of the zero-coupon bond price.

TIPS excess returns and x_t. I construct excess returns on TIPS by computing the 12-month holding period return on TIPS of different maturities. That is,

$$rx_{t,t+12}^{(\tau)} = p_{t+12}^{(\tau-12)} - p_t^{(\tau)} - y_t^{(12)}.$$

Because excess returns are proportional to changes in yields, I then regress excess returns onto changes in x_t,

$$rx_{t,t+12}^{(\tau)} = \alpha^{(\tau)} + \beta_0^{(\tau)}\Delta x_t + \sum_{h=1}^{3}\beta_h^{(\tau)}\Delta x_{t-h} + \sum_{h=0}^{3}\gamma_h^{(\tau)}\Delta pc_{1,t-h} + \sum_{h=0}^{3}\delta_h^{(\tau)}\Delta pc_{2,t-h} + \varepsilon_{t+12}^{(\tau)}. \quad (4)$$

Figure 6 shows the results. The left panel shows $\beta_0^{(\tau)}$ for different τ. As shown, a reduction in x_t (a tightening of financial conditions) is correlated with higher expected excess returns. This effect is relatively more pronounced for longer maturities, because longer bonds have higher duration risk. The right panel shows how the level of x_t changes the expected excess returns for bonds with 5- and 10-year maturities. This figure is computed by estimating regression (4) in levels instead of changes. The intuition is similar to the one in the left panel, but it is shown in a different way, with x_t on the horizontal axis (which is the same way I show excess returns in the model). A decline in x_t increases expected excess returns on real bonds, and the effect is relatively more pronounced for longer-maturity bonds.

An alternative measure. As elaborated in the previous subsections, the main idea behind the construction of the empirical measure x_t is to capture states in which financial intermediaries' risk-bearing capacity is low and there is an abnormal demand for highly liquid nominal instruments. In the subsections above, I used the HKM measure (market value of intermediaries' capital ratio) and the changes in the 5-year and 10-year on/off spreads in nominal bonds. I now construct an alternative x_t by substituting the 5-year on/off spread in nominal bonds with the spread between the 3-month uncollateralized repo rate and the 3-month T-bill. This measure, proposed by Nagel (2016), captures abnormal demand in "money-like" nominal financial instruments. The key idea is the same as in the previous subsection and is motivated by the same prediction in Fisher (1933), but focuses on spreads that capture abnormal demand for "money-like" assets. Table 3 shows the results from the same specification (2), controlling for TIPS bid-ask spreads. The results are qualitatively similar to those in the baseline specification, with the exception that the effect of x_t on the second PC of TIPS is not statistically significant after the GFC is removed (although it has a positive sign).

A comparison with the NFCI. The Chicago Fed produces a well-known financial conditions index, the National Financial Conditions Index (NFCI). The NFCI is not motivated by any particular theory and, instead, is constructed as a weighted average of 105 time series associated with many different asset prices.[9] Because the NFCI is a widely followed indicator of financial conditions, in this subsection I compare it with my baseline measure for financial conditions, x_t.

The correlation between my baseline x_t and changes in the NFCI is -56%—negative because a higher NFCI indicates tighter financial conditions, whereas a higher x_t indicates looser conditions—while the resulting R^2 of regressing the latter on the former is 0.33.[10] As a result of this high correlation between my baseline x_t and the fluctuations in the NFCI, it is reasonable to expect that the key correlations between TIPS and x_t would hold when using the NFCI as x_t in regression (2). Table 4 shows the results. The results are qualitatively similar to those in Table 1, where I estimated regression (2) using my baseline x_t.[11] Note, however, that the effect of NFCI on TIPS PC1 is statistically insignificant once the GFC is removed. That said, when the GFC period is excluded, the positive sign remains. The sensitivity of TIPS PC2 is strong and statistically significant even after removing the GFC. In particular, the sensitivity of PC2 to NFCI is about 10 times as large as the sensitivity of PC2 to x_t. This evidence suggests that the relationship between real yields and financial conditions is not simply driven by the definition of x_t.

Consumption dynamics and financial conditions. I conclude the empirical analysis by studying how financial conditions affect the real economy, in particular real consumption growth. This analysis is relevant because consumption growth will be a central ingredient for understanding the yield curve dynamics through the lens of a consumption-based asset pricing framework in Section 3. I proceed with two pieces of evidence. First, following the analysis throughout this section, I study how expected consumption growth and the conditional volatility of consumption growth move with x_t. Second, I zoom out from the recent evidence and conclude by providing evidence for consumption growth using the historical panel of Jordà et al. (2017), presented in Section 1.

To study how expected consumption growth and its conditional volatility move with x_t, I run

$$\Delta c_{t+1} = \mu_t + \sigma_t \varepsilon_{t+1}, \tag{5}$$

[9] See https://www.chicagofed.org/research/data/nfci/current-data for more details.

[10] Table 3 in He et al. (2017) shows that the shocks to intermediaries' capital ratio and changes in the NFCI have a correlation of approximately 40% in their sample (1970:Q1-2012:Q4).

[11] Note that, following the definition of the NFCI, a tightening of financial conditions means a higher NFCI. Instead, in my construction of x_t, a tightening of financial conditions means a lower x_t. Thus, the sensitivity of TIPS PCs to x_t and NFCI have opposite signs.

with

$$\mu_t = \beta_0 + \beta_1 x_t + \phi_0 \Delta c_t + \sum_{j=1}^{3} \phi_j \Delta c_{t-j}, \tag{6}$$

$$\sigma_t^2 = \exp(\gamma_0 + \gamma_1 x_t)\,. \tag{7}$$

Equations (6) and (7) are the expected consumption growth and conditional variance, respectively. The estimates of interest are β_1 and γ_1, which show how expected consumption growth and the conditional volatility of consumption growth move with x_t. I control for the lagged consumption growth in the conditional mean, to remove the persistence from the ε_{t+1} shocks. I estimate the parameters in the model (5)-(7) via maximum likelihood, using the monthly growth rate of personal consumption expenditures for Δc_t, reported by the Bureau of Economic Analysis, and the same x_t discussed in the previous subsections.

Figure 7 shows the results. The left panel shows how μ_t changes with x_t (holding the lags of consumption growth at their unconditional means) and the right panel shows how σ_t changes with x_t. Notably, both μ_t and σ_t increase as financial conditions tighten (i.e., lower x_t). The intuition is that, while the consumption level declines when financial conditions are tight (because credit flows are suddenly interrupted and agents seek to hold highly liquid nominal assets), the effect on the level of economic activity is expected to be transitory. As a result, expected consumption growth increases, because the transitory nature of tight financial conditions implies that the level of economic activity will recover in the near term.[12] At the same time, because tight financial conditions imply heightened uncertainty, the conditional volatility of consumption growth also increases when x_t is low.

I conclude by studying the fluctuations of consumption growth around financial crises, using the historical data presented in Section 1 for 14 advanced economies over 1870–2020. On the country-year panel, I estimate

$$\Delta c_{i,t+1} \;=\; \alpha_i \;+\; \left(b + \theta_F\, D^F_{i,t} + \theta_N\, D^N_{i,t}\right) \Delta c_{i,t} \;+\; g_F\, D^F_{i,t} + g_N\, D^N_{i,t} \;+\; \delta\, \mathbf{1}\{t \geq 1972\} \;+\; \varepsilon_{i,t+1}, \tag{8}$$

where α_i are country fixed effects and $D^F_{i,t}$, $D^N_{i,t}$ are the financial crisis and normal recession indicators (which are zero in tranquil years), respectively. The dummies are equal to one in the financial-crisis

[12] In contrast, permanent shocks shift the level of consumption without changing expected consumption growth: after a negative permanent shock, consumption is not expected to recover, so expected consumption growth does not increase.

and normal-recession events listed in Jordà et al. (2017) plus in their following three years, to test the dynamics of consumption growth right after the events. Because no financial-conditions measure like x_t is available over this long historical sample, I proxy tight conditions with the crisis dates and read off the transitory dynamics from the autocorrelation of consumption growth—the same mechanism, with the state observed in the modern sample and latent in the historical one. In particular, I am interested in testing the autocorrelation of consumption growth around financial crises, in order to understand the evolution of expected consumption growth in such events. That is, the object of interest is the state-specific autocorrelation of consumption growth, i.e. the slope of Δc_{t+1} on Δc_t around different episodes (b in tranquil times, $b + \theta_F$ during financial crisis and $b + \theta_N$ in normal recessions).

Figure 8 shows the results. The solid blue line shows that there is a negative relationship between consumption growth and expected consumption growth during financial crises (i.e., $b + \theta_F$ is negative). Consistent with the analysis done with x_t in a much more recent sample and at higher frequency, the evidence from the historical panel suggests consumption growth exhibits a negative autocorrelation during financial crises. That is, a tightening of financial conditions is associated with a short-run increase in expected consumption growth due to the transitory nature of the consumption decline.[13] Instead, in tranquil times (i.e., when neither normal nor financial recessions are hitting the economy) the autocorrelation of consumption growth is positive, and in normal recessions it is statistically indistinguishable from zero. That is, only financial crises are associated with a switch in the sign of the autocorrelation of consumption growth.

In sum, the evidence discussed in this section shows that a tightening of financial conditions (i.e., a low x_t) is associated with higher real rates (a higher first PC of TIPS), a flatter real yield curve (a lower second PC of TIPS), higher expected excess returns on TIPS, an increase in the conditional volatility of consumption growth and in the expected consumption growth. In the next section, I present a parsimonious consumption-based term structure model to rationalize the evidence. The key feature of the model is that financial frictions cause deadweight losses in consumption, along the lines of Brunnermeier and Sannikov (2014), Gertler and Kiyotaki (2015), and Schneider (2026).

[13] Importantly, this evidence does not rule out the presence of permanent shocks hitting consumption simultaneously. I show the results from the coefficients in Table A.2 in the Appendix.

3 Model

In this section I present a representative-agent endowment-economy model in which transitory shocks to financial conditions distort consumption in the economy. I model financial frictions in reduced form, borrowing the market-clearing conditions that prevail in general equilibrium models with financial intermediaries (e.g., Gertler and Kiyotaki (2015), Brunnermeier and Sannikov (2014), Schneider (2026)).

There is a representative household in an endowment economy with complete markets. Time is continuous and denoted by $t \geq 0$. The endowment is given by

$$\frac{\mathrm{d}y_t}{y_t} = \mu \, \mathrm{d}t + \sigma \, \mathrm{d}W_{1,t}, \tag{9}$$

where W_1 is a standard Brownian motion. The key assumption in the model is that consumption may not be equal to the endowment at all times. Instead, I assume that consumption, c_t, is distorted by the presence of financial frictions, which I model in reduced form. In particular, I assume that c_t is given by

$$c_t = f\left(x_t\right) y_t, \tag{10}$$

where $f(x) \in [0, 1]$ captures the fact that a fraction $1 - f$ of the endowment may be wasted due to the misallocation of resources associated with the presence of financial frictions. The function $f(x)$ is motivated by models like Brunnermeier and Sannikov (2014), in which the misallocation of resources is associated with the presence of relatively unproductive savers holding risky assets when intermediaries' constraints are binding. As a result, when financial conditions are tight (intermediaries have low capital), consumption declines. Similarly, in models along the lines of Gertler and Kiyotaki (2015), the misallocation of resources is associated with households having to pay a cost to hold risky assets and such cost causes deadweight losses in consumption.

I assume

$$f(x_t) = \omega + (1 - \omega)x_t, \tag{11}$$

with $\omega < 1$ and $x_t \in [0, 1]$. One could think of the functional form (11) as the market-clearing condition of the goods market in an economy in which there are two types of dividends. Dividends received from assets held by financial intermediaries would be $D_{f,t} = \omega y_t$ and dividends paid when savers hold risky assets would be $D_{s,t} = x_t \left(1 - \omega\right) y_t$. The fraction $x_t < 1$ captures the idea that

savers are inefficient in handling risky assets. In this interpretation, given that consumption would be equal to total dividends, the market clearing condition for goods (assuming intermediaries do not consume but rebate profits to savers), $c_t = D_{s,t} + D_{f,t}$, gives the same expression as combining (10) and (11).

I interpret states in which $x_t < 1$ as "tight financial conditions" because low x_t captures the fact that there are inefficiencies causing consumption to be lower than the endowment. When $x_t = 1$, there are no inefficiencies and consumption equals the endowment. In general equilibrium models with financial intermediaries, x_t would be an endogenous state variable, the wealth share of intermediaries. In this paper, for the sake of tractability, I model x_t as an exogenous state variable following

$$\mathrm{d}x_t = \kappa\left(\overline{x} - x_t\right)\mathrm{d}t + \sigma_x\sqrt{\left(1 - x_t\right)x_t}\mathrm{d}W_{2,t}, \tag{12}$$

where W_2 is a standard Brownian motion with correlation ϕ_{12} with W_1. Importantly, note that the shocks W_2 have a transitory effect on consumption while shocks W_1 have a permanent effect.

The representative agent has recursive preferences as in Duffie and Epstein (1992)

$$U = E_t\int_t^\infty u\left(c_s, U_s\right)\mathrm{d}s,$$

with

$$u\left(c, U\right) = \frac{\rho}{1 - \frac{1}{\psi}}\left\{\frac{c^{1-\frac{1}{\psi}}}{\left[\left(1-\gamma\right)U\right]^{\left(\gamma-1/\psi\right)/\left(1-\gamma\right)}} - \left(1-\gamma\right)U\right\}, \tag{13}$$

where γ is the risk aversion (RA) and ψ is the elasticity of intertemporal substitution (EIS). When $\gamma = 1/\psi$, I recover time-additive, constant-relative-risk-aversion (CRRA) preferences.

Equilibrium. The equilibrium in this model can be characterized using a single state variable, x_t. Given the consumption process, I first compute the value function $U(c, x)$ that characterizes the stochastic differential utility. I then compute the stochastic discount factor (SDF), to obtain the prices of risk and the short-term interest rate. Finally, I compute the term structure of interest rates in the economy. The following proposition characterizes the key elements of the solution, which can be simplified in the case of constant-relative risk aversion (CRRA) utility, as I discuss below.

Proposition 1 *I guess and verify the solution of the value function*

$$U^* = \frac{\left(c_t\xi\left(x_t\right)\right)^{1-\gamma}}{1-\gamma},$$

where ξ is a function that solves the following ordinary differential equation

$$0 = \frac{\rho}{1-\frac{1}{\psi}}\left\{\xi\left(x\right)^{\frac{1}{\psi}-1}-1\right\}+E_t\left[\frac{dc_t}{c_t}\right]-\frac{\gamma}{2}\left(\frac{dc}{c}\right)^2+\frac{\xi_x}{\xi}E_t\left[dx\right]+\frac{1}{2}\left[\frac{\xi_{xx}}{\xi}-\gamma\left(\frac{\xi_x}{\xi}\right)^2\right](dx)^2+(1-\gamma)\,E_t\left[\frac{d\xi}{\xi}\frac{dc}{c}\right].$$

Using Itô's lemma on (10), the consumption process is

$$\frac{dc_t}{c_t} = \mu_c\left(x_t\right)dt + \sigma dW_{1,t} + \sigma\left(x_t\right)dW_{2,t},$$

with

$$\mu_c(x_t) = \mu + \frac{1-\omega}{f(x_t)}\kappa(\overline{x}-x_t) + \phi_{12}\frac{(1-\omega)\,\sigma\,\sigma_x\sqrt{(1-x_t)x_t}}{f(x_t)}, \tag{14}$$

$$\sigma(x_t) = \frac{1-\omega}{f(x_t)}\,\sigma_x\sqrt{(1-x_t)x_t}. \tag{15}$$

The SDF, denoted by $m_t > 0$, is given by

$$\frac{dm_t}{m_t} = \mu_{m,t}\,dt \;-\; \lambda_{m1,t}\,dW_{1,t} \;-\; \lambda_{m2,t}\,dW_{2,t}. \tag{16}$$

where u_U and u_c are the derivatives of (13) with respect to U and c, respectively. Then,

$$E_t\left[\frac{dm_t}{m_t}\right] = \begin{bmatrix} -\gamma E_t\left[\frac{dc_t}{c_t}\right] + \frac{1}{2}\left(1+\gamma\right)\gamma\left(\frac{dc_t}{c_t}\right)^2 + \\ \left(\frac{1}{\psi}-\gamma\right)\frac{\xi_x}{\xi}E\left[dx_t\right] + \frac{1}{2}\left(\left(\frac{1}{\psi}-\gamma\right)\left(\frac{1}{\psi}-\gamma-1\right)\left(\frac{\xi_x}{\xi}\right)^2 + \left(\frac{1}{\psi}-\gamma\right)\frac{\xi_{xx}}{\xi}\right)(dx_t)^2 \\ -\gamma\left(\frac{\xi_x}{\xi}\right)\left(\frac{1}{\psi}-\gamma\right)E\left[\frac{dc_t}{c_t}dx_t\right] - \frac{\rho(1-\gamma)}{1-\frac{1}{\psi}} - \frac{\rho\left(\gamma-\frac{1}{\psi}\right)}{1-\frac{1}{\psi}}\xi\left(x\right)^{\frac{1}{\psi}-1}, \end{bmatrix}$$

and

$$\begin{aligned} \lambda_{m1} &= \gamma\sigma, \\ \lambda_{m2} &= \left[\gamma\frac{1-\omega}{f\left(x_t\right)} - \left(\frac{1}{\psi}-\gamma\right)\left(\frac{\xi_x}{\xi}\right)\right]\sigma_x\sqrt{\left(1-x_t\right)x_t}. \end{aligned}$$

Proof. See appendix.

Given the equilibrium SDF in (16), the time-t price of a zero-coupon bond maturing τ periods

ahead is

$$P_t^{(\tau)} \;=\; \mathbb{E}_t\left[\frac{m_{t+\tau}}{m_t}\right].$$

The corresponding continuously compounded yield is

$$y_t^{(\tau)} \;=\; -\frac{1}{\tau}\,\log P_t^{(\tau)}.$$

Simplified case: CRRA with autoregressive x_t. I present a simplified version of the model to help build the intuition behind the more general case. In Proposition 2 below, I show that, in this simplified case, the model can be expressed as a quadratic term structure model—a framework that has been extensively studied in the reduced-form affine term structure literature (see, e.g., Duffie and Kan (1996) and Ahn et al. (2002)). I make two simplifying assumptions. First, I assume $\gamma{=}1/\psi$, so I obtain the constant relative risk aversion case. Second, I assume the state variable x_t follows an autoregressive process that is not bounded between zero and one as was the case in (12), but can theoretically be any real number.[14] That is, I simplify x_t to follow

$$\mathrm{d}x_t = \kappa(\bar{x} - x_t)\,\mathrm{d}t + \sigma_x\,\mathrm{d}W_{2,t}. \tag{17}$$

Then, assuming for tractability that shocks are uncorrelated (i.e., $\phi_{12} = 0$), I can write the real rate as[15]

$$r_t = \rho + \gamma\,\mu_c(x_t) - \frac{1}{2}\gamma(\gamma+1)\big(\sigma^2 + \sigma(x_t)^2\big). \tag{18}$$

where μ_c and $\sigma\,(x_t)$ are (14) and (15) but with $\phi_{12} = 0$, respectively.

The second term in (18) is the so-called intertemporal smoothing force. It captures the fact that agents have a desire to smooth consumption over time: when expected consumption growth increases, agents would like to borrow to smooth consumption across time, which pushes the real rate up. The third term captures the precautionary savings force. As consumption becomes more volatile, agents have a preference to save, which pushes the real rate down. It can be clearly seen from expression (18) that the intertemporal smoothing and precautionary saving forces would move the short rate in opposite directions, because a decline in x_t pushes both μ_c and $\sigma(x_t)$ higher. Hence, how much $r(x_t)$ changes with x_t depends on the net effect of x_t on the intertemporal smoothing and

[14] Assuming unbounded x_t is inconsistent with the consumption being a fraction f of the endowment, but I relax this for the sake of tractability in the CRRA example.

[15] I show this in the Appendix, as part of the Proposition 1.

precautionary savings force. This is the main reason why using recursive preferences is key: It helps balance the relative strength of the intertemporal smoothing and precautionary saving forces, so that the real rate has sufficient nonlinear dynamics consistent with the evidence.

In this simplified model, the term structure can be solved in closed-form. The next proposition shows that, after making an additional assumption of linearizing $g(x) = \frac{1-\omega}{\omega+(1-\omega)x_t}$ around the unconditional mean of x_t, the model can be written as a quadratic term structure model, whose solution can be found in Ahn et al. (2002).

Proposition 2 *Assuming that: 1) the representative agent has CRRA preferences (i.e., $\gamma = 1/\psi$); 2) the variable x_t follows the process (17); 3) the function $g(x) = \frac{1-\omega}{\omega+(1-\omega)x_t}$ is linearized around $\overline{x}$, the model can be written as a quadratic term structure model. The short rate can be expressed as*

$$r(x_t) = r_0 + r_1 x_t + r_2 x_t^2,$$

the prices of risk can be expressed as

$$\begin{aligned} \lambda_{m1} &= \lambda_{m1}^0, \\ \lambda_{m2,t} &= \lambda_{m2}^0 - \lambda_{m2}^1 x_t, \end{aligned}$$

where parameters $r_0, r_1, r_2, \lambda_{m1}^0, \lambda_{m2}^0, \lambda_{m2}^1$ are shown in the Appendix. Then, bond prices have the form

$$P_t^{(\tau)} = \exp\left(A^{(\tau)} + B^{(\tau)} x_t + C^{(\tau)} x_t^2\right),$$

where $A^{(\tau)}$, $B^{(\tau)}$, and $C^{(\tau)}$ are functions of maturity and are shown in the Appendix.

The expected excess return of a bond with maturity τ can be written as

$$E_t\left[\frac{dP_t^{(\tau)}}{P_t^{(\tau)}} - r_t dt\right] = \lambda_{m2,t}\sigma_x\left(B^{(\tau)} + 2C^{(\tau)} x_t\right)$$

For a maturity sufficiently small, the term $\left(B^{(\tau)} + 2C^{(\tau)} x_t\right)$ is (minus) the derivative of the short rate with respect to x_t, namely, $-r_x$. Given that $\lambda_{m2,t} > 0$, the sign of the expected excess return is pinned down by the term related to r_x.

The key idea in Proposition 2 is that the sign of the term premium in the model reduces to understanding r_x—the derivative of the short rate with respect to the state variable x. When r_x is

negative, the term premium is positive. A negative r_x means that, after a decline in x (i.e., tight financial conditions and high deadweight losses), the intertemporal smoothing force is stronger than the precautionary savings force, causing the short rate to increase.

I next present the general case with recursive preferences, which I solve numerically.

Numerical solution. I solve the model with recursive preferences using numerical methods. With CRRA preferences, the model can be solved in closed-form for the short-term real rate and the prices of risk, which I do in the next subsection. I calibrate the model at monthly frequency. Table 5 shows the parameters.

For the preferences, I set a risk aversion coefficient, γ of 5, a value within the range commonly used in the literature. I tried alternative values of the EIS, and I show that it is a critical parameter to capture the term structure dynamics. In the baseline, I set an EIS equal to 1/8. I calibrate the parameters associated with the process (12), namely κ and σ_x, to match the statistical properties of the empirical measure of x_t. The chart below table 5 shows the invariant distribution of x.[16] I set $\overline{x}$=0.9, to obtain a negatively skewed invariant distribution of x, which is qualitatively consistent with the evidence. However, the diffusion process for x cannot capture substantial negative skewness that could be attributed to jumps or extreme events. I set the financial wedge to 0.4. On very rare occasions, x_t would drop below 0.85 (only 5% of the time). In such cases, consumption losses would be 9% of the endowment (i.e., $\omega + (1-\omega) \times 0.85 = 0.91$), which is approximately consistent with the peak-to-trough losses during financial crises (Muir (2017)). Finally, I calibrate σ to obtain a volatility of consumption of 2.5%, which is in line with the most recent evidence of consumption dynamics.[17] I set ϕ_{12}=0.75 to capture the idea that shocks to the endowment and shocks to financial conditions are positively correlated most of the time. Under this calibration, the implied autocorrelation of consumption growth is approximately zero. In the sample period I consider in the empirical analysis (i.e., the sample period for which TIPS data is available), the autocorrelation of monthly consumption growth is 0.09, with a 95% confidence interval between -0.2 and 0.2.

Figure 9 shows the solution. The top-left and top-right panels show the real rate and the price of risk, respectively, across the state space (x_t). The gray area shows the invariant distribution of x_t in the background. The central feature of the model is that the real rate and the price of risk increase as x_t declines. Intuitively, this is because a decline in x causes a decline in the consumption

[16] The invariant distribution of x is Beta, as I show in the Appendix.

[17] Real personal consumption expenditures.

level. However, given the transitory nature of x, the disruption in consumption is expected to be transitory. As a result, consumption is expected to increase in the near horizon, pushing the real rate up through the so-called intertemporal substitution force (i.e., the household has a preference to borrow from the future in order to smooth the decline in consumption). The strength of such force is governed by the EIS. On the other hand, the price of risk also increases, because the quantity of x-risk (i.e., the diffusion component in consumption associated with W_2 shocks) rises as x declines. The increase in the conditional volatility of consumption pushes the real rate down, through the so-called precautionary savings channel (i.e., the household has a preference to save into the future once the consumption path becomes more uncertain). In the baseline calibration, the intertemporal smoothing force dominates over precautionary savings and, as a result, the interest rate increases as x declines. In the next subsection I discuss the CRRA case, to further illustrate this feature of the model.

The increase in the real rate when the price of risk increases as x declines produces a positive real term premium, as shown in the lower-left panel. Intuitively, real bond prices decline in states in which financial conditions are temporarily tight. As a result, they command a premium in order to incentivize the household to hold such bonds. The premium is more pronounced for bonds with a longer horizon, because they carry higher duration risk. As a result of this positive real term premium, the average real term structure is upward sloping, as shown in the lower-right panel. The model can capture the slope of the real term structure observed in the 1999-2025 period, as shown by the black dots in the lower-right figure.

The CRRA case. When $\gamma=1/\psi$, the recursive preferences collapse to the time-additive CRRA case. I show the expressions for the real rate and the prices of risk in the Appendix.

Figure 10 shows the model solution under two alternative calibrations: CRRA and a high EIS (i.e., $\psi=1/3$ instead of the $\psi=1/8$ used in the baseline[18]). As the EIS increases, from the baseline 1/8, to CRRA of 1/5 and then to a high EIS of 1/3, the real interest rate becomes less sensitive to x and the price of risk becomes smaller. This is shown in the upper panels of Figure 10. As a result, the term premium on long-term bonds is smaller when the EIS is higher (bottom-left panel), resulting in a flatter average real yield curve (bottom-right panel).

[18]With CRRA, the EIS is the inverse of the RA, so it is 1/5.

4 Conclusion

This paper argues that financial conditions are a fundamental macro source of risk behind the real term premium. I motivate the analysis by providing evidence that supports Fisher (1933)'s prediction that the real rate increases in periods of extremely tight financial conditions (e.g., financial crises). I then zoom in on more recent and higher-frequency evidence using U.S. TIPS and study how real rates correlate with financial conditions. For this, I construct an empirical measure that is the first principal component of shocks to financial intermediaries' capital ratios and changes in the on-the-run/off-the-run spreads of 5- and 10-year nominal bonds. The empirical measure is intended to capture the idea in Fisher (1933) that tight financial conditions (in the extreme, financial crises) are characterized by a reduction in credit availability (which tends to happen when intermediaries' risk-bearing capacity is low) and an abnormal demand for liquid nominal assets. I show that when financial conditions are tight, the level of the real yield curve increases, the slope of the curve flattens, and excess returns on real bonds increase. These correlations are valid even after removing the Global Financial Crisis (which was an event of extremely tight financial conditions) from the sample and after controlling for TIPS bid-ask spreads (or related liquidity measures). Importantly, tight financial conditions have, on impact, a transitory effect on the real economy: in both the historical country-panel and the more recent, higher-frequency sample, I find that expected consumption growth increases when financial conditions are tight (i.e., the autocorrelation of consumption growth becomes sharply negative shortly after the onset of such events).

I propose a parsimonious consumption-based term structure model in which transitory shocks to financial conditions generate distortions in the consumption process. I model these distortions in reduced form, but I show they are in line with the role financial frictions play in more elaborate models featuring financial intermediaries. When financial conditions are tight, the level of consumption in the economy declines while the volatility of consumption increases. The decline in consumption is expected to reverse in the near future, as financial conditions normalize. Thus, shocks to financial conditions create two opposing forces in the dynamics of the real rate. On the one hand, the representative agent has an increased preference to borrow from the future to smooth the decline in consumption, pushing the real rate up. On the other hand, the representative agent has an increased preference for higher savings as its consumption path becomes more uncertain. On net, under plausible calibrations of the EIS and risk aversion, I show that the real rate increases when financial conditions are tight, because the preference to smooth the decline in consumption more than offsets

the desire to save. The increase in the real interest rate causes real bonds to decline in value in states in which the price of risk is high (due to elevated consumption volatility), thus generating a positive real term premium on average, which increases when financial conditions tighten.

5 Figures and Tables

TABLE 1. x_t and the principal components of TIPS

	PC_1		ΔPC_1		PC_2		ΔPC_2	
	Full sample	ex-GFC	Full sample	ex-GFC	Full sample	ex-GFC	Full sample	ex-GFC
x_t	-0.030**	-0.018*	-0.030**	-0.019**	0.042**	0.028	0.044**	0.031*
	(0.014)	(0.010)	(0.014)	(0.010)	(0.019)	(0.019)	(0.020)	(0.019)
3 Lags x_t	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
3 Lags z_t	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
N	314	296	314	296	314	296	314	296
Adj.-R^2	0.9840	0.9864	0.1294	0.0772	0.9379	0.9511	0.1710	0.1409

NOTE: This table shows the estimates from regression (2). PC1 and PC2 are the first and second principal components of TIPS. The variable x_t, shown in Figure 3, is defined as the first principal component of shocks to primary dealers' capital ratio (He et al. (2017)) and the changes in the 5- and 10-year on-the-run/off-the-run spread in nominal Treasuries. All variables are standardized so coefficients are comparable. Sample period 1999:01-2025:06. HAC standard errors in parentheses. Significance at 1% is ***, at 5% is **, and at 10% is *.

TABLE 2. x_t and the principal components of TIPS, controlling for bid-ask spreads

	PC_1		ΔPC_1		PC_2		ΔPC_2	
	Full sample	ex-GFC	Full sample	ex-GFC	Full sample	ex-GFC	Full sample	ex-GFC
x_t	-0.056***	-0.040***	-0.058***	-0.044***	0.054**	0.042	0.068***	0.053**
	(0.018)	(0.013)	(0.018)	(0.013)	(0.022)	(0.028)	(0.020)	(0.025)
Bid-Ask$_t$	-0.045	-0.024	-0.055	-0.031	-0.041	-0.019	-0.033	-0.048
	(0.032)	(0.034)	(0.034)	(0.036)	(0.082)	(0.069)	(0.078)	(0.074)
3 Lags x_t	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
3 Lags Bid-Ask	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
3 Lags z_t	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
N	204	192	204	192	204	192	204	192
Adj.-R^2	0.9713	0.9700	0.2401	0.1191	0.9369	0.9484	0.3033	0.2006

NOTE: This table shows the estimates from regression (2). PC1 and PC2 are the first and second principal components of TIPS. The variable x_t, shown in Figure 3, is defined as the first principal component of shocks to primary dealers' capital ratio (He et al. (2017)) and the changes in the 5- and 10-year on-the-run/off-the-run spread in nominal Treasuries. $Bid-Ask_t$ is the first principal component of changes in the 5- and 10-year TIPS bid-ask spread. All variables are standardized so coefficients are comparable. Sample period 1999:01-2025:06. HAC standard errors in parentheses. Significance at 1% is ***, at 5% is **, and at 10% is *.

TABLE 3. Money-like-based x_t and the principal components of TIPS

	PC_1		ΔPC_1		PC_2		ΔPC_2	
	Full sample	ex-GFC	Full sample	ex-GFC	Full sample	ex-GFC	Full sample	ex-GFC
x_t	-0.065***	-0.048***	-0.066***	-0.051***	0.043*	0.021	0.057**	0.030
	(0.016)	(0.014)	(0.016)	(0.015)	(0.024)	(0.029)	(0.022)	(0.026)
Bid-Ask$_t$	-0.059**	-0.040	-0.071**	-0.049	-0.034	-0.032	-0.024	-0.062
	(0.029)	(0.033)	(0.032)	(0.034)	(0.081)	(0.069)	(0.079)	(0.076)
3 Lags x_t	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
3 Lags Bid-Ask	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
3 Lags z_t	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
N	204	192	204	192	204	192	204	192
Adj.-R^2	0.9731	0.9710	0.2902	0.1468	0.9373	0.9483	0.3105	0.1936

NOTE: The table shows the estimates from regression (2). PC1 and PC2 are the first and second principal components of TIPS. The variable x_t, shown in Figure 3, is defined as the first principal component of shocks to primary dealers' capital ratio (He et al. (2017)), the spread between the 3-month uncollateralized repo rate and the 3-month T-bill, and changes in the 10-year on-the-run/off-the-run spread in nominal Treasuries. $Bid - Ask_t$ is the first principal component of changes in the 5- and 10-year TIPS bid-ask spread. All variables are standardized so coefficients are comparable. Sample period 1999:01-2025:06. HAC standard errors in parentheses. Significance at 1% is ***, at 5% is **, and at 10% is *.

TABLE 4. NFCI as x_t and the principal components of TIPS

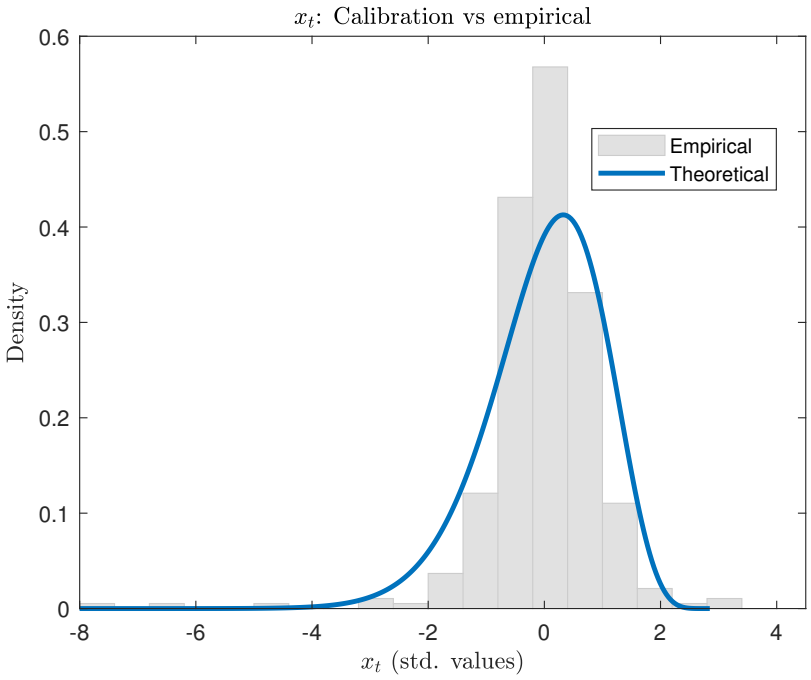
	PC_1		ΔPC_1		PC_2		ΔPC_2	
	Full sample	ex-GFC	Full sample	ex-GFC	Full sample	ex-GFC	Full sample	ex-GFC
$\Delta NFCI_t$	0.119*	0.033	0.122**	0.029	-0.436***	-0.411**	-0.449***	-0.410**
	(0.061)	(0.083)	(0.060)	(0.082)	(0.106)	(0.166)	(0.107)	(0.175)
Lags $\Delta NFCI$	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
3 Lags z_t	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
N	315	297	315	297	315	297	315	297
Adj.-R^2	0.9838	0.9861	0.1214	0.0590	0.9476	0.9537	0.2998	0.1805

NOTE: This table shows the estimates from regression (2). PC1 and PC2 are the first and second principal components of TIPS. NFCI is the Chicago Fed National Financial Condition Index. All variables are standardized so coefficients are comparable. Sample period 1999:01-2025:06. HAC standard errors in parentheses. Significance at 1% is ***, at 5% is **, and at 10% is *.

TABLE 5. Calibration

Calibration (monthly)

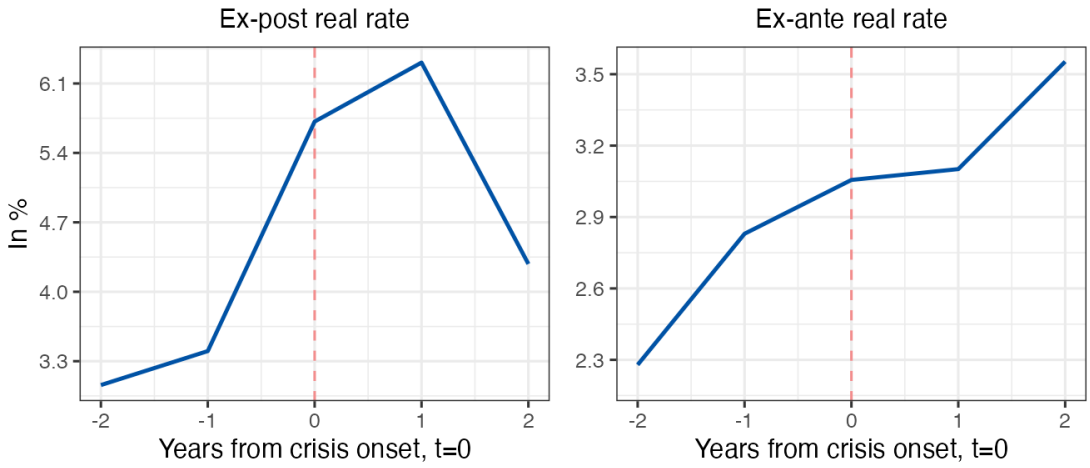
Parameter	Value	Description
γ	5	Risk aversion
ψ	1/8	EIS
σ	0.003	Conditional vol. y_t
μ	0.001	Drift of y_t
ϕ_{12}	0.75	Correlation W_1, W_2
$\overline{x}$	0.9	Mean x_t
σ_x	0.029	Conditional vol. x_t
κ	0.03	Persistence x_t
ω	0.4	Financial wedge



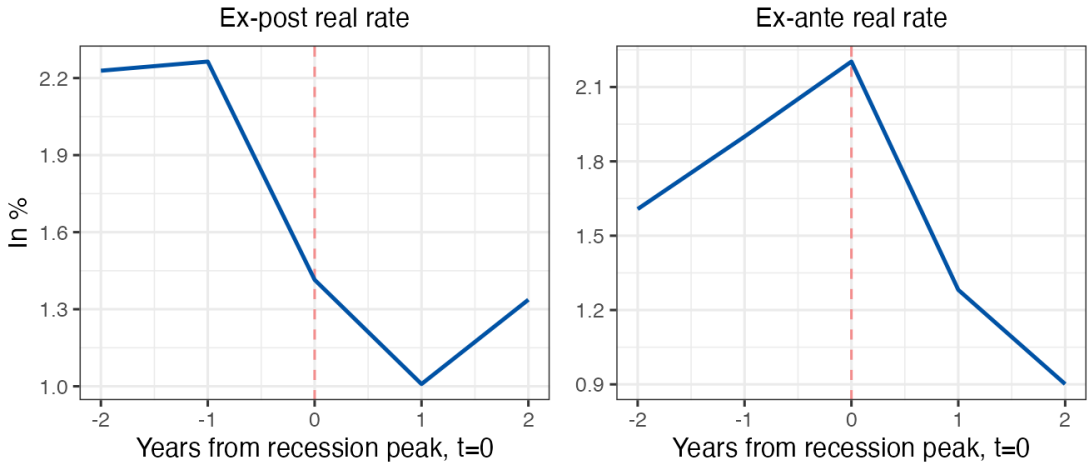
NOTE: The table shows the model calibration at monthly frequency. The figure shows the distribution of x_t implied by the parameters in the table (in blue) and the histogram of the realization of x_t shown in Figure 3.

FIGURE 1. Mean paths around financial crises and normal recessions

Panel A. Mean paths around financial crises, 1870–2020

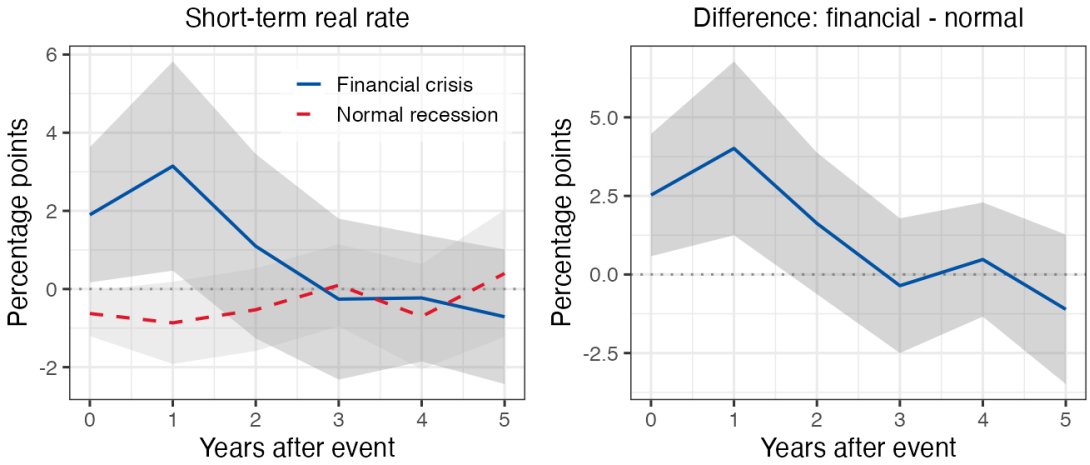


Panel B. Mean paths around normal recessions, 1870–2020



NOTE: The figure shows the mean path of the short-term real interest rate—ex-post (using realized inflation) and ex-ante (using expected inflation from a country-by-country AR(1) model)—around financial crises (Panel A) and normal recessions (Panel B), for 14 advanced economies over 1870–2020. The mean path is computed first averaging across episodes within each country, and then averaging across countries. The sample includes 69 financial crises and 227 normal recessions. The dates of financial crises and normal recessions come from Jordà et al. (2017) and Jordà et al. (2013), respectively; the list of countries and event dates are in the Appendix.

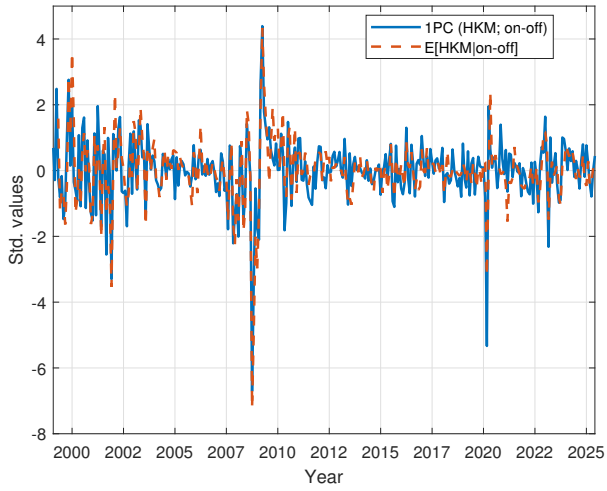
FIGURE 2. Local projections: Financial crises and normal recessions



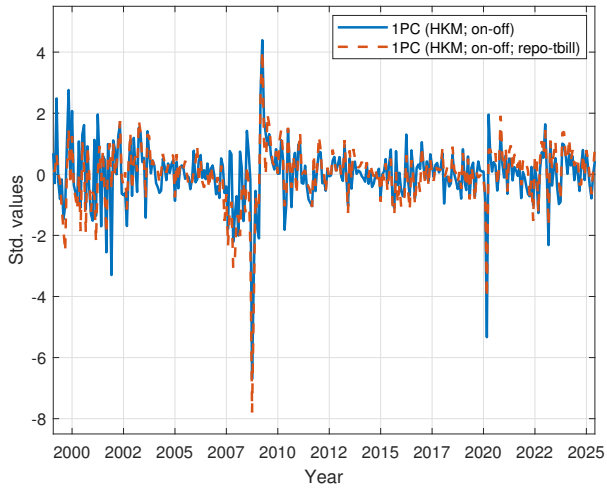
NOTE: The left panel shows the responses for a financial crisis (solid blue lines) and for a normal recession (dashed-red lines) from specification (1). The x-axis is in years. The right panel shows the difference between the responses. Gray areas are the 90% confidence bands. Standard errors are clustered by year and country.

FIGURE 3. x_t: Principal component and projection.

Panel A.

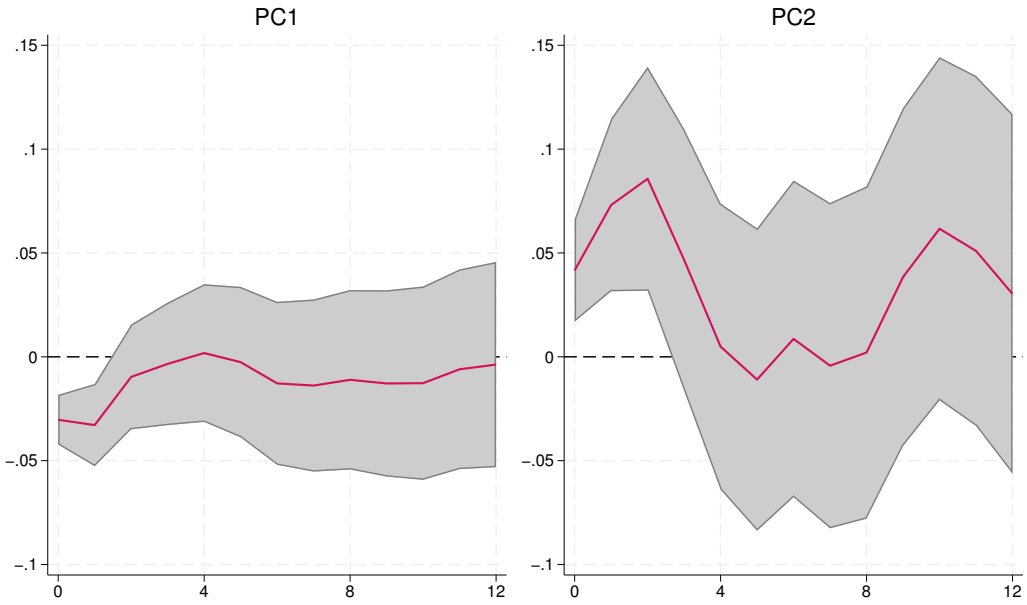


Panel B.



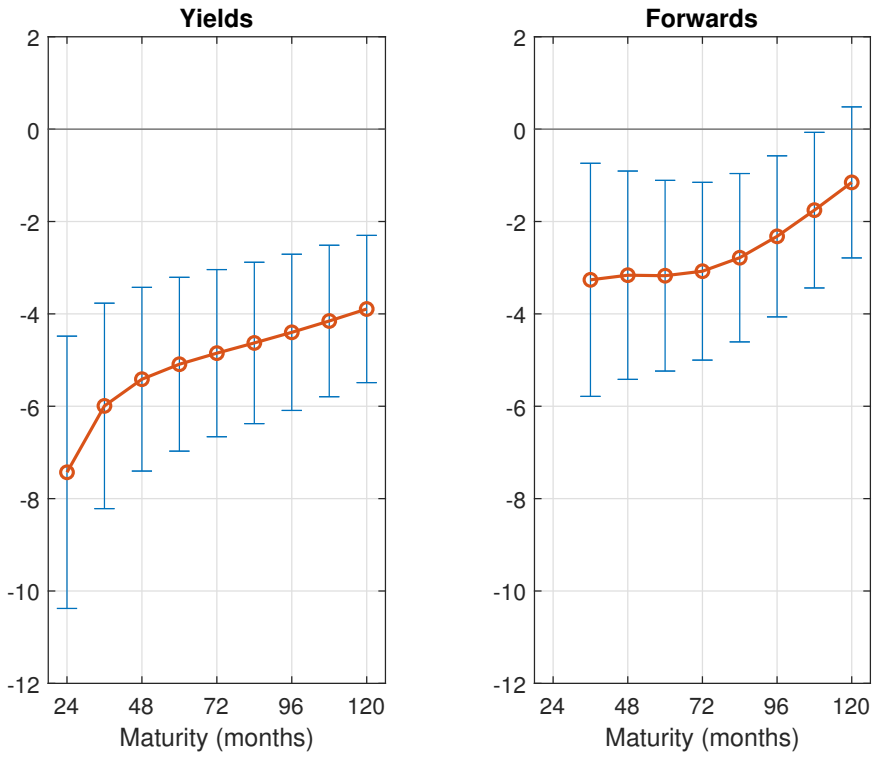
NOTE: This figure shows, in blue, the first principal component of shocks to the capital ratio of primary dealers (He et al., 2017) and the changes in the 5- and 10-year nominal on-the-run/off-the-run spread. In panel A, the red dashed line shows the projection of the shocks to the capital ratio onto the changes in the 5- and 10-year nominal on-the-run/off-the-run spread and their three lags. The red dashed line in panel B is the first principal component of shocks to the capital ratio of primary dealers (He et al., 2017), the change in the 10-year nominal on-the-run/off-the-run spread, and the spread between the 3-month uncollateralized repo and the 3-month T-bill.

FIGURE 4. Local projections: x_t and the principal components of TIPS.



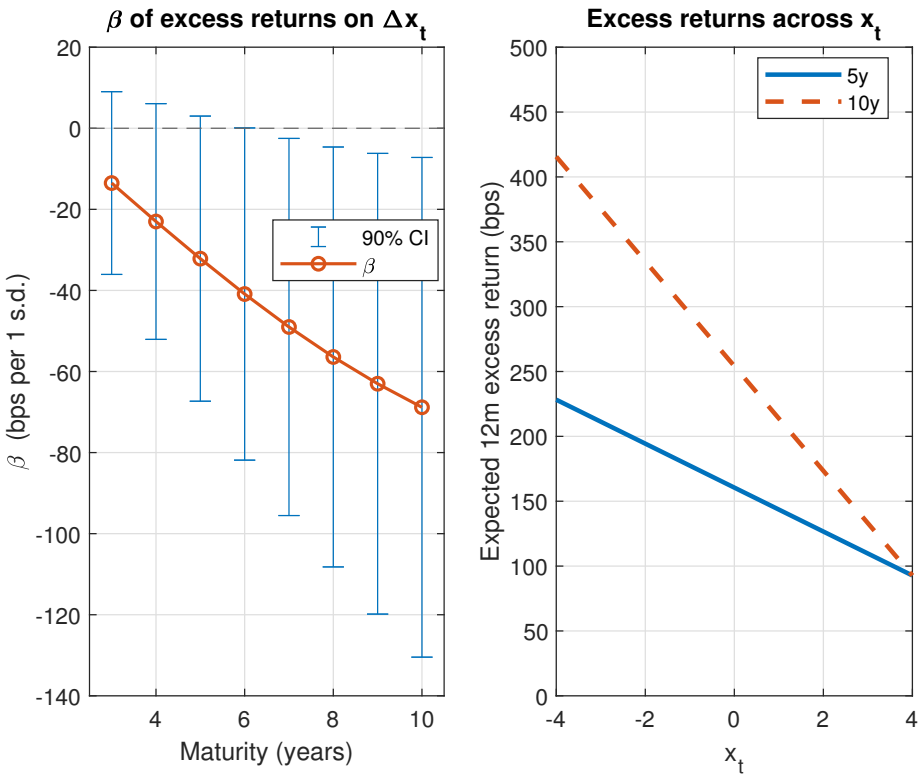
NOTE: This figure shows the local projections from specification (3). The x-axis is months. The y-axis shows the coefficients of interest. The left (right) panel shows the results for the first (second) principal component of TIPS. Gray areas are the 90% confidence band with HAC standard errors. Sample is 1999:01-2025:06 in monthly frequency. See Appendix for more details about the data.

FIGURE 5. Real yields and forwards onto x_t.



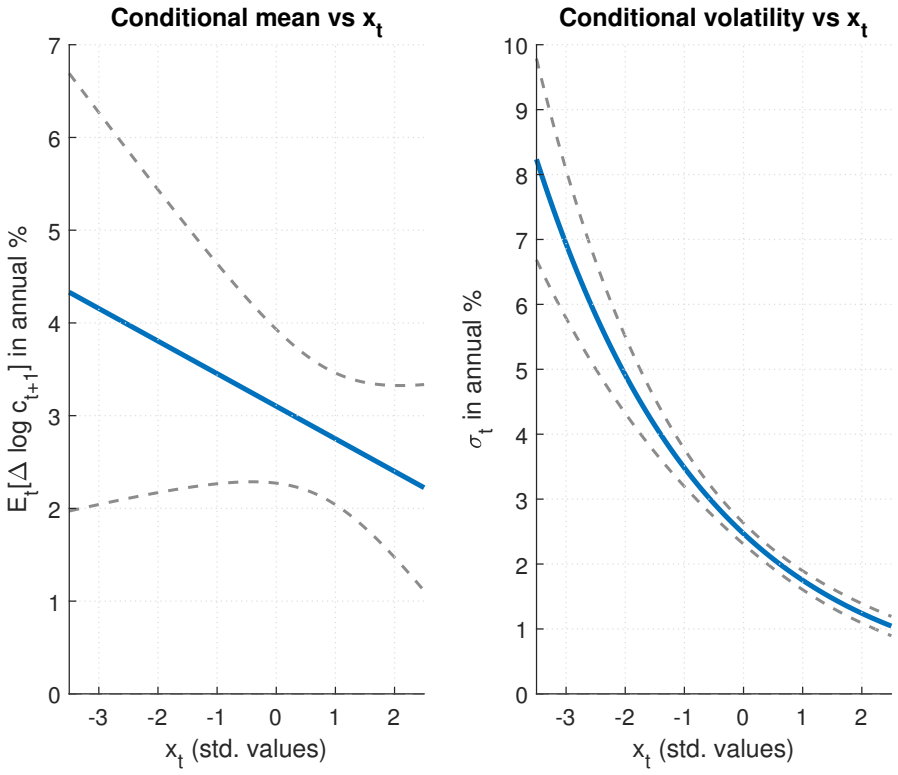
NOTE: This figure shows the betas associated with x_t, for TIPS of different maturities, in specification (2). The error bands are 90% HAC standard errors. Sample is 1999:01-2025:06 in monthly frequency. See Appendix for more details about the data.

FIGURE 6. Excess returns on TIPS and x_t.



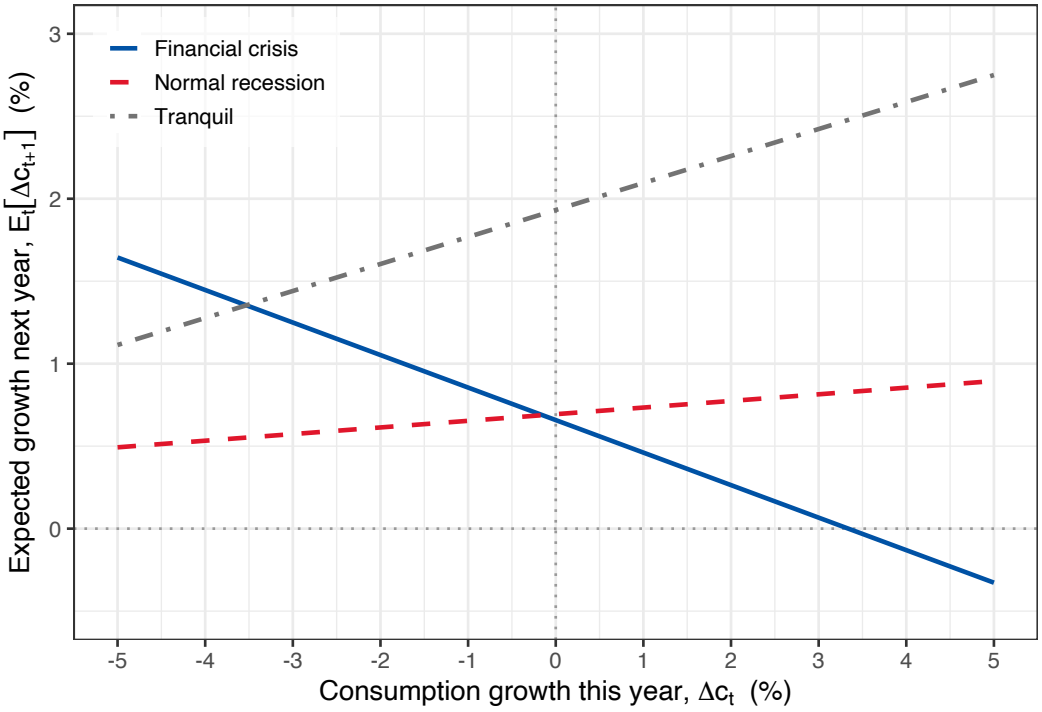
NOTE: The left panel shows the β_0 in regression (4), for TIPS with maturities 2- to 10-year. The right panel shows how excess returns vary across x_t, which is computed after running regression (4) in levels instead of differences. Confidence bands at 90% using HAC standard errors.

FIGURE 7. Consumption and x_t.



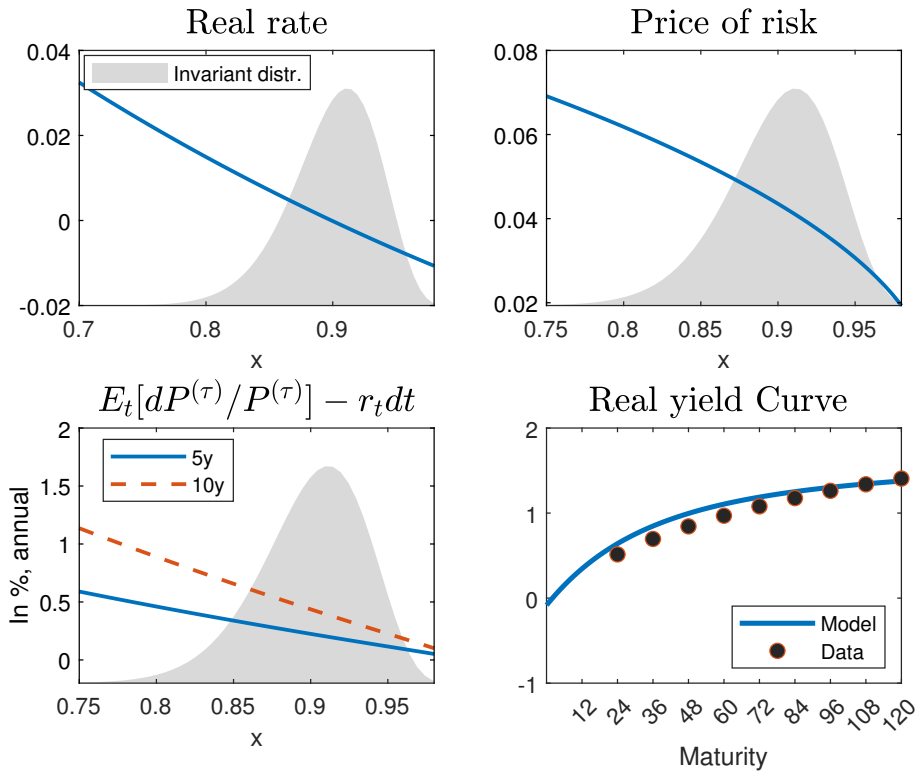
NOTE: This figure shows the results from the heteroskedastic regression (5). The left panel shows $\mu_t = E_t[\Delta c_{t+1}]$, expected consumption growth, for different values of x_t. The right panel shows σ_t, the conditional volatility of consumption growth, for different values of x_t. The gray dashed lines are 90% confidence bands.

FIGURE 8. Expected consumption growth during financial crises and normal recessions, country-panel 1870-2020



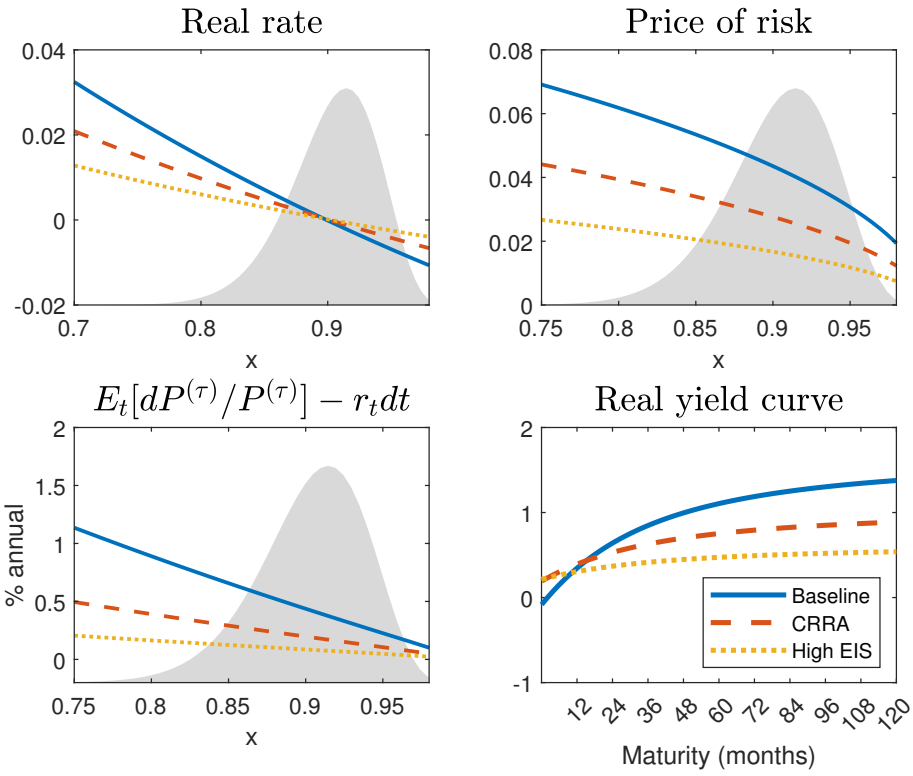
NOTE: This figure plots the relationship between expected consumption growth and current consumption growth during financial crises (in solid blue), normal recessions (in dashed red) and tranquil times (in dotted gray), implied by the country-panel regression (8). Each line shows $E_t[\Delta c_{t+1}]$ as a function of Δc_t; its slope is the autocorrelation of consumption growth in that state—b in tranquil years (gray, dot-dashed), $b + \theta_F$ in financial crises (blue, solid), and $b + \theta_N$ in normal recessions (red, dashed). A downward slope indicates mean reversion, i.e., a negative autocorrelation: a decline this year predicts a rebound (higher expected growth) next year. The sample comprises 14 advanced economies over 1870–2020 using real consumption per capita from Jordà et al. (2017); crisis and normal-recession dates are those in Section 1.

FIGURE 9. Model solution.



NOTE: This figure shows the model solution.

FIGURE 10. CRRA vs recursive preferences.



NOTE: This figure shows the model solution with CRRA preferences (red-dashed line) and with a high EIS (dotted-orange line). The solid blue line shows the solution Figure 9, which uses the calibration in Table 5.

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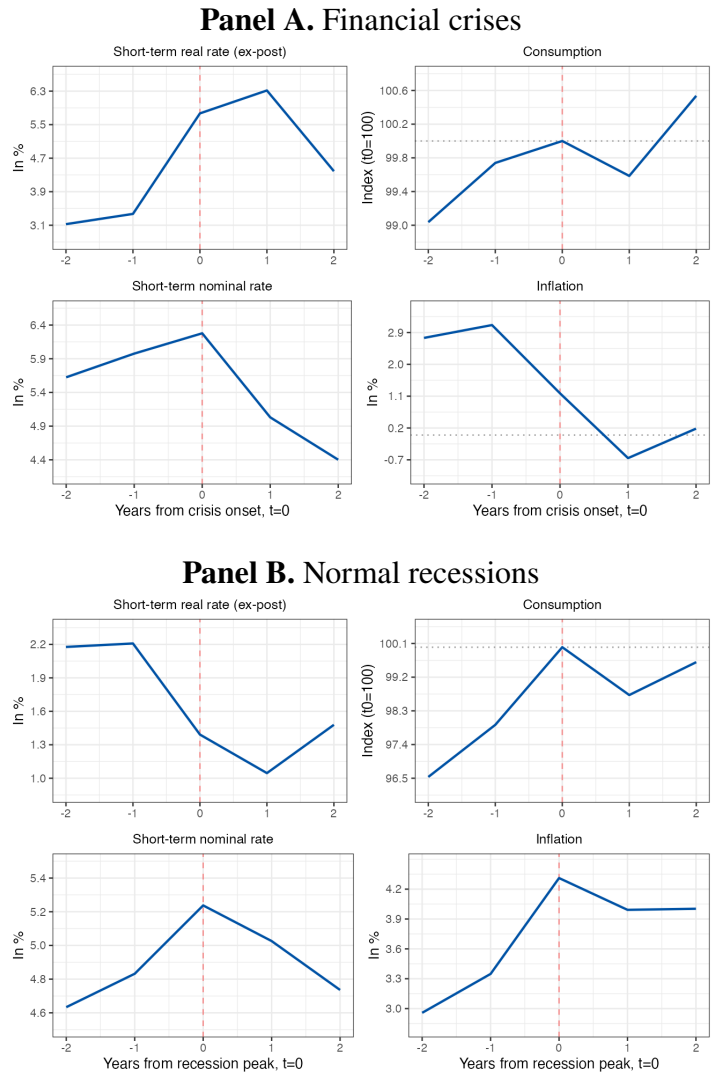
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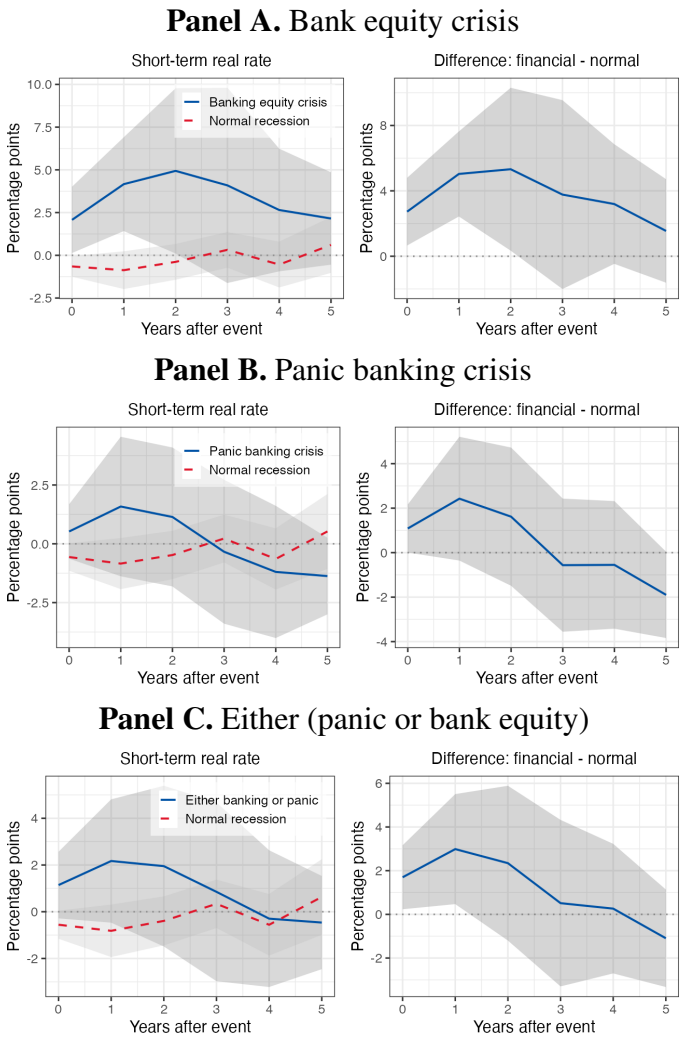
Appendix

FIGURE A.1. Mean paths around financial crises and normal recessions, 1870–2020



NOTE: This figure shows the mean path of consumption (index, $t_0 = 100$), the short-term ex-post real rate, the nominal rate, and inflation around financial crises (Panel A) and normal recessions (Panel B), for 14 advanced economies over 1870–2020. Each series is averaged with equal weight across countries (after averaging episodes within country) over the window $t - 2$ to $t + 2$; the ex-post real-rate panel corresponds to the series in Figure 1. Crisis and normal-recession dates come from Jordà et al. (2017) and Jordà et al. (2013).

FIGURE A.2. Short-term real rate around events in Baron et al. (2021)



NOTE: This figure shows the responses of the local projection in (1) for the short-term (ex-post) real rate around different types of financial crises defined in Baron et al. (2021). Each panel uses a different crisis definition: bank equity crises (Panel A), panic banking crises (Panel B), and either a panic or a bank equity crisis (Panel C). In each panel, the left plot shows the response around financial crises (blue, solid) and normal recessions (red, dashed), and the right plot shows the difference between the two (financial minus normal); the horizontal axis is years after the event. The sample covers 14 advanced economies; normal-recession dates come from Jordà et al. (2017) and Jordà et al. (2013). Shaded areas are 90% confidence bands based on standard errors clustered by country and year.

TABLE A.1. The components of x_t and the principal components of TIPS

A. Full sample

	PC_1					PC_2				
	(1)	(2)	(3)	(4)	(5)	(1)	(2)	(3)	(4)	(5)
HKM	-0.005			-0.004	0.008	0.034**			0.028*	0.036**
	(0.011)			(0.011)	(0.008)	(0.017)			(0.016)	(0.016)
On/off 5y		0.004		0.005			-0.034**		-0.028*	
		(0.009)		(0.009)			(0.017)		(0.015)	
On/off 10y			0.050***		0.053***			-0.008		0.002
			(0.013)		(0.013)			(0.028)		(0.027)
Lags z_t	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Lags RHS	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
N	315	315	315	314	314	315	315	315	314	314
Adj.-R^2	0.9831	0.9828	0.9855	0.9828	0.9856	0.9372	0.9373	0.9372	0.9383	0.9385

B. Ex-GFC

	PC_1					PC_2				
	(1)	(2)	(3)	(4)	(5)	(1)	(2)	(3)	(4)	(5)
HKM	0.003			0.004	0.007	0.033**			0.028*	0.034**
	(0.008)			(0.008)	(0.007)	(0.016)			(0.015)	(0.016)
On/off 5y		0.005		0.005			-0.043**		-0.038**	
		(0.010)		(0.010)			(0.017)		(0.016)	
On/off 10y			0.040***		0.041***			0.029*		0.037**
			(0.013)		(0.013)			(0.018)		(0.018)
Lags z_t	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Lags RHS	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
N	297	297	297	296	296	297	297	297	296	296
Adj.-R^2	0.9863	0.9861	0.9873	0.9860	0.9872	0.9518	0.9523	0.9520	0.9527	0.9532

NOTE: This table shows the estimates from regression (2). PC_1 and PC_2 are the first and second principal components of TIPS. The variables HKM, On/off 5y and On/off 10y are the components of x_t and are the shocks to primary dealers' capital ratio (He et al., 2017) and changes in the 5- and 10-year on-the-run/off-the-run spread in nominal Treasuries, respectively. All variables are standardized so coefficients are comparable. The sample period is 1999:01–2025:06. HAC standard errors in parentheses. Significance at 1% is ***, at 5% is **, and at 10% is *.

TABLE A.2. Expected consumption growth during financial crises and normal recessions, country-panel 1870-2020

	Estimate
Autocorrelation, tranquil (b)	0.164***
	(0.021)
Autocorrelation, financial crisis ($b + \theta_F$)	-0.197**
	(0.089)
Autocorrelation, normal recession ($b + \theta_N$)	0.040
	(0.088)
Extra slope, financial (θ_F)	-0.361***
	(0.089)
Extra slope, normal (θ_N)	-0.123*
	(0.068)

NOTE: This table reports the state-specific autocorrelation of consumption growth from regression (8), estimated on the country-year panel from JST (14 advanced economies, 1870–2020). The dummies $D^F_{i,t}$ and $D^N_{i,t}$ equal one in the year of a financial crisis (respectively, normal recession) and in the following three years. The slope of next-year on this-year consumption growth is b in tranquil years, $b + \theta_F$ in financial crises, and $b + \theta_N$ in normal recessions. The total number of country-year observations is 2,080. Standard errors clustered by country and year in parentheses. Significance at 1% is ***, at 5% is **, and at 10% is *.

Proof of Proposition 1. Apply Itô's lemma on the guessed value function,

$$U = \frac{(c\xi(x))^{1-\gamma}}{1-\gamma},$$

and take the derivatives

$$\begin{aligned}
U_c &= c^{-\gamma}\xi(x)^{1-\gamma}, \\
U_{cc} &= -\gamma c^{-\gamma-1}\xi(x)^{1-\gamma}, \\
U_x &= c^{1-\gamma}\xi(x)^{-\gamma}\xi_x, \\
U_{xx} &= c^{1-\gamma}\xi(x)^{-\gamma}\xi_{xx} - \gamma c^{1-\gamma}\xi(x)^{-\gamma-1}(\xi_x)^2, \\
U_{cx} &= (1-\gamma)c^{-\gamma}\xi(x)^{-\gamma}\xi_x,
\end{aligned}$$

thus $\mathrm{d}U^*$ can be written as

$$\mathrm{d}U^* = c^{1-\gamma}\xi(x)^{1-\gamma}\left[\frac{\mathrm{d}c}{c} - \frac{\gamma}{2}\left(\frac{\mathrm{d}c}{c}\right)^2 + \frac{\xi_x}{\xi}\mathrm{d}x + \frac{1}{2}\left[\frac{\xi_{xx}}{\xi} - \gamma\left(\frac{\xi_x}{\xi}\right)^2\right](\mathrm{d}x)^2 + (1-\gamma)\left(\frac{\xi_x}{\xi}\right)\frac{\mathrm{d}c}{c}\mathrm{d}x\right].$$

Then, insert the guessed value function on (13), to get

$$\begin{aligned}
u(c,U^*) &= \frac{\rho}{1-\frac{1}{\psi}}\left\{\frac{c^{1-\frac{1}{\psi}}}{\left[(1-\gamma)\frac{(c\xi(x))^{1-\gamma}}{1-\gamma}\right]^{(\gamma-1/\psi)/(1-\gamma)}} - (1-\gamma)\frac{(c\xi(x))^{1-\gamma}}{1-\gamma}\right\}, \\
&= \frac{\rho}{1-\frac{1}{\psi}}c^{1-\gamma}\xi^{1-\gamma}\left\{\xi^{\frac{1}{\psi}-1} - 1\right\}.
\end{aligned}$$

Thus, in equilibrium,

$$0 = u(c,U^*) + E\left[\mathrm{d}U^*\right].$$

Substituting U^* on the previous expression, I obtain the ordinary differential equation for $\xi(x)$ shown in the Proposition 1:

$$0 = \frac{\rho}{1-\frac{1}{\psi}}\left\{\xi^{\frac{1}{\psi}-1} - 1\right\} + E_t\left[\frac{\mathrm{d}c}{c}\right] - \frac{\gamma}{2}\left(\frac{\mathrm{d}c}{c}\right)^2 + \frac{\xi_x}{\xi}E_t\left[\mathrm{d}x\right] + \frac{1}{2}\left[\frac{\xi_{xx}}{\xi} - \gamma\left(\frac{\xi_x}{\xi}\right)^2\right](\mathrm{d}x)^2 + (1-\gamma)E_t\left[\frac{\mathrm{d}\xi}{\xi}\frac{\mathrm{d}c}{c}\right].$$

Equations (14) and (15) in Proposition 1 follow from applying Itô's lemma on (10) and using the laws of motion for x_t and y_t, which are equations (9) and (12), respectively.

To obtain equation (16), follow Duffie and Epstein (1992), and write the SDF as

$$\frac{\mathrm{d}m_t}{m_t} = \frac{\mathrm{d}u_c}{u_c} + u_U \mathrm{d}t.$$

with partial derivatives of $u(c, U)$ being

$$u_c = \rho c^{-\frac{1}{\psi}} \left[(1-\gamma)\, U\right]^{\frac{\frac{1}{\psi}-\gamma}{1-\gamma}},$$

and

$$u_U = -\frac{\rho\,(1-\gamma)}{1-\frac{1}{\psi}} - \frac{\rho\left(\gamma-\frac{1}{\psi}\right)}{1-\frac{1}{\psi}} c^{1-\frac{1}{\psi}} \left((1-\gamma)\, U\right)^{\frac{\frac{1}{\psi}-1}{1-\gamma}}.$$

Using the guessed value function and applying Itô's lemma, the SDF is given by

$$\begin{aligned}
\frac{\mathrm{d}m_t}{m_t} &= \begin{bmatrix} -\gamma E\left[\frac{\mathrm{d}c}{c}\right] + \frac{1}{2}(1+\gamma)\,\gamma\sigma_c^2 + \\ \left(\frac{1}{\psi}-\gamma\right)\frac{\xi_x}{\xi} E\left[\mathrm{d}x\right] + \frac{1}{2}\left(\left(\frac{1}{\psi}-\gamma\right)\left(\frac{1}{\psi}-\gamma-1\right)\left(\frac{\xi_x}{\xi}\right)^2 + \left(\frac{1}{\psi}-\gamma\right)\frac{\xi_{xx}}{\xi}\right)\sigma_x^2\,(1-x_t)\,x_t \\ -\gamma\left(\frac{\xi_x}{\xi}\right)\left(\frac{1}{\psi}-\gamma\right) E\left[\frac{\mathrm{d}c}{c}\mathrm{d}x\right] - \frac{\rho(1-\gamma)}{1-\frac{1}{\psi}} - \frac{\rho\left(\gamma-\frac{1}{\psi}\right)}{1-\frac{1}{\psi}}\xi\,(x)^{\frac{1}{\psi}-1} \end{bmatrix} \\
&\quad -\gamma\sigma\mathrm{d}W_1 - \gamma\sigma_c\,(x)\,\mathrm{d}W_2 + \left(\frac{1}{\psi}-\gamma\right)\left(\frac{\xi_x}{\xi}\right)\sigma_x\sqrt{(1-x_t)\,x_t}\mathrm{d}W_2
\end{aligned}$$

which are the expressions at the end of Proposition 1. Note that the real rate can be simplified in the CRRA case (i.e., when $\gamma = 1/\psi$),

$$\begin{aligned}
r_t &= -E_t\left[\frac{\mathrm{d}m_t}{m_t}\right], \\
&= \rho + \gamma\mu_c\,(x_t) - \frac{1}{2}(1+\gamma)\,\gamma\left(\sigma^2 + \sigma_c\,(x)^2\right).
\end{aligned}$$

Proof of Proposition 2. Assume for tractability that shocks $\mathrm{d}W_1$ and $\mathrm{d}W_2$ are uncorrelated.

Assuming (1) CRRA preferences; (2) x_t follows an Ornstein-Uhlenbeck process, the real rate is

$$r = \rho + \gamma\mu_c(x_t) - \frac{1}{2}(1+\gamma)\gamma\left(\sigma^2 + \sigma(x)^2\right),$$

with

$$\begin{aligned}\mu_c(x) &= \mu + \frac{1-\omega}{\omega + (1-\omega)x_t}\kappa(\overline{x} - x_t),\\ \sigma(x) &= \frac{1-\omega}{\omega + (1-\omega)x_t}\sigma_x.\end{aligned}$$

I next linearize $g(x) = \frac{1-\omega}{\omega+(1-\omega)x_t}$ around $\overline{x}$:

$$g(x) \approx g_0 - g_1 x_t$$

with

$$\begin{aligned}g_0 &= \frac{1-\omega}{\omega + (1-\omega)\overline{x}} + g_1\overline{x},\\ g_1 &= \frac{(1-\omega)^2}{(\omega + (1-\omega)\overline{x})^2}.\end{aligned}$$

Then, using this approximation, we can write $\mu_c(x)$ and $\sigma(x)$ as

$$\begin{aligned}\mu_c(x) &= \mu + \kappa(g_0 - g_1 x_t)(\overline{x} - x_t),\\ \sigma(x) &= (g_0 - g_1 x_t)\sigma_x,\end{aligned}$$

so

$$\mu_c(x) = \mu + \kappa\left(g_0\overline{x} - g_0 x_t - g_1 x_t\overline{x} + g_1 x_t^2\right), \tag{A-1}$$

$$\sigma(x) = g_0\sigma_x - g_1\sigma_x x_t. \tag{A-2}$$

Plugging (A-1) and (A-2) in the real rate, the real rate can be written as

$$r(x_t) = r_0 + r_1 x_t + r_2 x_t^2,$$

with

$$\begin{aligned} r_0 &= \rho + \gamma\mu + \gamma\kappa g_0\overline{x} - \frac{1}{2}\gamma\left(\gamma+1\right)\left(\sigma^2 + g_0^2\sigma_x^2\right), \\ r_1 &= \gamma\left[\kappa\left(-g_0 - g_1\overline{x}\right) + \left(1+\gamma\right)g_0 g_1\sigma_x^2\right], \\ r_2 &= \gamma\left[\kappa g_1 - \frac{1}{2}\left(1+\gamma\right)g_1^2\sigma_x^2\right]. \end{aligned}$$

Using the prices of risk shown in Proposition 1, together with $\gamma = 1/\psi$ and the linearized $g(x)$, the prices of risk are

$$\begin{aligned} \lambda_{m1} &= \gamma\sigma = \lambda_{m1}^0, \\ \lambda_{m2,t} &= \gamma\left(g_0\sigma_x - g_1\sigma_x x_t\right) = \lambda_{m2}^0 - \lambda_{m2}^1 x_t. \end{aligned}$$

Then, the zero-coupon bond prices are given by

$$P_t^{(\tau)} = P\left(x_t, \tau\right) = E_t\left[\frac{m_{t+\tau}}{m_t}\right], \tag{A-3}$$

where

$$\frac{dm_t}{m_t} = -r_t\mathrm{d}t - \lambda_{m1}\mathrm{d}W_{1,t} - \lambda_{m2,t}\mathrm{d}W_{2,t}.$$

The ordinary differential equation solving (A-3) is

$$-\frac{\partial P\left(x,\tau\right)}{\partial\tau} - r_t P\left(x,\tau\right) + \frac{\partial P\left(x,\tau\right)}{\partial x}E\left[\mathrm{d}x_t\right] + \frac{1}{2}\frac{\partial^2 P\left(x,\tau\right)}{\partial x^2}E\left[\mathrm{d}x_t^2\right] + E_t\left[\frac{\mathrm{d}m}{m}\mathrm{d}P_t^{(\tau)}\right] = 0,$$

with the boundary condition $P(x,0) = 1$. The solution of the partial differential equation is affine in x_t and x_t^2, the bond price can be written as

$$P\left(x_t,\tau\right) = \exp\left(A^{(\tau)} + B^{(\tau)}x_t + C^{(\tau)}x_t^2\right),$$

where functions $A^{(\tau)}$, $B^{(\tau)}$ and $C^{(\tau)}$ solve the system of ordinary differential equations

$$\begin{aligned}
C'^{(\tau)} &= 2\left(-\kappa + \sigma_x \lambda^1_{m2}\right) C^{(\tau)} + 2\sigma_x^2 \left(C^{(\tau)}\right)^2 - r_2, \\
B'^{(\tau)} &= 2\left(\kappa\overline{x} - \sigma_x \lambda^0_{m2}\right) C^{(\tau)} + \left(-\kappa + \sigma_x \lambda^1_{m2}\right) B^{(\tau)} + 2\sigma_x^2 B^{(\tau)} C^{(\tau)} - r_1, \\
A'^{(\tau)} &= \left(\kappa\overline{x} - \sigma_x \lambda^0_{m2}\right) B^{(\tau)} + \frac{1}{2}\sigma_x^2 \left(B^{(\tau)2} + 2C^{(\tau)}\right) - r_0,
\end{aligned}$$

with boundary conditions $A^{(0)} = B^{(0)} = C^{(0)} = 0$. Then expected excess returns on bonds with maturity τ are

$$\begin{aligned}
E_t\left[\frac{\mathrm{d}P_t^{(\tau)}}{P_t^{(\tau)}} - r_t \mathrm{d}t\right] &= -E_t\left[\frac{\mathrm{d}m}{m}\mathrm{d}P_t^{(\tau)}\right], \\
&= \lambda_{m2,t}\left(B^{(\tau)} + 2C^{(\tau)} x_t\right)\sigma_x \mathrm{d}t, \\
&= \underbrace{\left(\lambda^0_{m2} - \lambda^1_{m2} x_t\right)}_{>0}\underbrace{\left(B^{(\tau)} + 2C^{(\tau)} x_t\right)}_{\approx -r_x}\sigma_x \mathrm{d}t.
\end{aligned}$$

Thus, the sign of the expected excess returns is proportional to the sign of $-r_x$. If $r_x < 0$ (the intertemporal smoothing dominates the precautionary savings), the expected excess returns on real bonds are positive.

Data description. The historical cross-country analysis uses the Jordà–Schularick–Taylor Macrohistory Database for 14 advanced economies over 1870–2020, from which I use real consumption per capita, short-term nominal interest rates, and inflation. Financial-crisis dates are those defined in Jordà et al. (2017), and normal-recession dates are from Jordà et al. (2013); for robustness I use the bank-equity and panic-crisis dates of Baron et al. (2021). The modern U.S. analysis is at the monthly frequency over 1999:01–2025:06. Intermediaries' risk-bearing capacity is measured by shocks to the market value of primary dealers' capital ratio from He et al. (2017). Investors' demand for liquid nominal assets is measured by changes in the 5- and 10-year on-the-run/off-the-run spreads in nominal Treasuries and, in the alternative measure, by the spread between the 3-month uncollateralized repo rate and the 3-month Treasury bill (Nagel (2016)). TIPS zero-coupon yields and their bid-ask spreads (available from May 2008) are used to construct the real yield curve and its principal components. Monthly real personal consumption expenditures from the Bureau of Economic Analysis are used for Δc_t. The National Financial Conditions Index is from the Federal Reserve Bank of Chicago.