Bond Liquidity Premia

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Theory predicts that funding conditions faced by financial intermediaries are an important limit to arbitrage. We identify and measure the value of funding liquidity from the cross-section of Treasury securities. To validate our interpretation, we establish linkages with funding conditions in the repo market, the shadow banking sector, and the overall economy. Looking at asset pricing implications, we find that increases in funding liquidity predict lower risk premia for all Treasury securities but higher risk premia on LIBOR loans, swap contracts, and corporate bonds. The impact of funding conditions on interest rates is large and pervasive throughout crises and normal times. (*JEL* E43, H12)

> "... a part of the interest paid, at least on long-term securities, is to be attributed to uncertainty of the future course of interest rates." (p. 163)
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> "... the imperfect 'moneyness' of those bills which are not money [...] causes the trouble of investing in them and [causes them] to stand at a discount." (p. 166)
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> "... In practice, there is no rate so short that it may not be affected by speculative elements; there is no rate so long that it may not be affected by the alternative use of funds in holding cash." (p. 166)
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> —John R. Hicks, *Value and Capital*, 2nd edition, 1948

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Bond traders know very well that liquidity affects asset prices. One prominent example is the on-the-run liquidity premium, whereby the most recently issued (on-the-run) bonds sell at a premium relative to seasoned (off-the-run) bonds with similar coupons and maturities.[1] Moreover, systematic variations in liquidity drive interest rates across markets. A case in point took place around the Federal Open Market Committee (FOMC) decision, on October 15, 1998, to lower the federal funds rate by 25 basis points (bps). In the opening of the meeting, Vice-Chairman McDonough of the New York district bank stressed that the spreads were increasing in several fixed-income markets, a sign that a run to quality and a serious drying up of liquidity were occurring.[2] The recent financial crisis provides another example. Facing sharp increases of interest rate spreads in most markets, the Committee reduced its target for the federal funds rate, approved novel policy instruments, and expanded its balance sheet substantially. These events attest to the occasional dramatic impact of liquidity seizures.

An appealing explanation for these seizures rests on a common wealth shock to intermediaries or speculators (e.g., Shleifer and Vishny 1997; Kyle and Xiong 2001; Gromb and Vayanos 2002; He and Krishnamurthy 2008; Lagos 2006). Intuitively, lower wealth hinders their ability to pursue quasi-arbitrage opportunities and to provide liquidity. Therefore, across markets in which they operate, risk premia will share a component linked to the shadow price of the capital constraints faced by these intermediaries. Recently, Brunnermeier and Pedersen (2008) have highlighted the importance of funding markets for intermediaries who want to leverage scarce capital. In practice, Adrian and Shin (2009) show that repo markets are the key markets where investment banks, hedge funds, and other speculators obtain the marginal funds for their activities.[3]

Our main contribution is to show that the value of funding liquidity is an aggregate risk factor driving a substantial share of risk premia across fixed-income markets. We document that tight funding conditions lower substantially the risk premium on U.S. Treasury bonds but raise the risk premium implicit in LIBOR rates, swap rates, and corporate bond yields. This pattern is consistent with accounts of flight-to-quality, but the relationship is pervasive even in normal times. Jointly, the evidence across markets is hard to reconcile

[1] The on-the-run liquidity premium was first documented by Warga (1992) when he measured the impact of age and maturity on the average premium. Several theoretical and empirical papers have since linked the on-the-run premium to liquidity and often focus on the price difference with the just-off-the-run security. Specific references will be found in the core of the text.

[2] Specifically, he noted increases in the spread between the on-the-run and the most recent off-the-run 30-year Treasury bonds (0.05% to 0.27%), the spreads between swaps and Treasury notes with two years and ten years to maturity (0.35% to 0.70%, and 0.50% to 0.95%, respectively), the spreads between Treasuries and investment-grade corporate securities (0.75% to 1.24%), and between Treasuries and mortgage-backed securities (1.10% to 1.70%). Minutes of the FOMC, October 15, 1998, conference call.

[3] They document the positive relationship between annual growth in total assets, growth in repo positions, and growth of other collateral financing by the six primary dealers. See their Figure 3.5.

with theories based on variations of default probability, inflation, real interest rates, and their associated risk premia. In contrast, we add considerable evidence for the importance of intermediation frictions in asset pricing. Beyond the difference between their cash flows, different securities serve to fulfill investors' uncertain future needs for cash. Therefore, funding liquidity and its risk are an important component of observed risk premia.

We extend the standard no-arbitrage dynamic term structure model of Christensen, Diebold, and Rudebusch (2011) [hereafter CDR] to allow for a liquidity factor, L_t, that affects coupon bond prices.[4] We measure this latent liquidity premium by estimating a term structure model from a panel of pairs of U.S. Treasury securities, where each pair has similar cash flows but different ages. This strategy is consistent with the existence of an on-the-run premium in the short run but also with the evidence that older bonds are less liquid. Therefore, estimates of the liquidity factor will be obtained through price differentials that can be attributed to differences in age. We use a sample of end-of-month bond prices running from December 1985 until the end of 2007.[5] Hence, our results cannot be attributed to the extreme influence of the recent financial crisis. In a concluding section, we repeat the estimation including 2008 and 2009 to gauge the magnitude of the increase in liquidity value.

The core of the article aims at demonstrating that L_t, our age-based measure, can be interpreted as a measure of the value of funding liquidity. We present evidence from different sources, at three successive levels of aggregation, explicitly linking our measure to funding conditions. First, we relate the liquidity value to the expected benefits of holding a more liquid security. We measure these benefits by a common component in repo spreads,[6] that is, the differences between the general collateral and the special collateral rates, the latter being associated with recently issued securities. We find that as much as 20% of the variations in the liquidity factor can be linked to future variations of that systemic component in repo spreads.

Second, we trace the linkages of funding liquidity to the shadow banking sector, a large non-bank component of the financial intermediation system that relies heavily on short-term funding to finance long-lived illiquid assets. We document the price-quantity relationship between the value of liquidity and the quantity of funding liquidity supplied by shadow banks. A one-standard-deviation increase in the value of liquidity is associated with a 7.2% increase of total assets supported in the shadow banking sector. The impact is large relative

4 This underlying CDR term structure model captures parsimoniously the usual level, slope, and curvature factors, while delivering good in-sample fit and forecasting power. Moreover, the smooth shape of Nelson-Siegel curves helps identify small deviations, relative to an idealized curve, which may be caused by variations in market liquidity.

5 Note that using Treasury securities sidesteps credit risk issues. Nonetheless, a significant tax premium is entangled with the liquidity premium in the earlier period.

6 We consider maturities of 2, 5, 10, and 30 years.

to the impact of changes in interest rates, or measures of flight-to-liquidity based on the flows into and within Money Market Mutual Funds (MMMF).

Third, we study the relationship between the value of funding liquidity and broader measures of funding conditions. We find that variations of non-borrowed reserves of commercial banks at the Federal Reserve are negatively related to variations of the liquidity factor. Similarly, increases in the rate of money supply growth, as measured by $M2$, are associated with decreases of the liquidity factor. Therefore, the value of liquidity decreases when the supply of funds to intermediaries is ample. These results are robust to the inclusion of information from a broad range of financial and economic variables. Overall, these three sources of evidence vindicate our interpretation of the extracted liquidity factor as a measure of the value of funding liquidity.

We then test whether the tightness of funding conditions affects risk premia across several fixed-income securities. We first consider seasoned U.S. Treasury bonds. The evidence is decisive. Controlling for the level and shape of the term structure, an increase in the value of funding liquidity predicts lower excess bond returns for all maturities in the future. For a two-year bond, a one-standard-deviation change predicts that annualized returns in excess of one-year yields decrease by 85 bps. This compares to an average excess return of 69 bps. This is consistent with viewing Treasury bonds as hedges against funding liquidity shocks. All Treasury bonds can be converted into cash, via the funding market, quickly and at low costs relative to other asset classes. In practice, investors see them as liquid substitutes so that bond prices rise and their risk premium declines following funding shocks.

Next, we find that variations of LIBOR rolling excess returns are positively linked to variations in the value of funding liquidity. A one-standard-deviation change predicts an increase of 34 bps in excess returns from rolling a three-month interbank loan over a year.[7] The relationship is significant, both statistically and economically, and the marginal contribution to the predictive power is high. Then, an increase in the value of funding liquidity is associated with a higher spread between three-month LIBOR and T-Bill rates (TED spread) via its effect on the Treasury curve *and* on the LIBOR curve. The sample correlation is 0.63. Our results support the recent interpretation of the TED spread as an aggregate measure of funding risk. The effect of funding liquidity on LIBOR rates extends to swap rates. It predicts an increase of 6 bps of the five-year swap relative to a par Treasury yield. This is economically significant given the higher sensitivity of this contract value to changes in yields.

We also consider a sample of corporate bond spreads from the National Association of Insurance Commissioners (NAIC). We find that the impact of funding liquidity is significant and follows a flight-to-quality pattern across

7 Rolling excess returns are the returns from rolling a short-term interbank loan computed in excess of the known (longer-term) yield on the Treasury market for the same investment horizon. For the example in the text, it is the returns from rolling a 3-month loan over four periods computed in excess of the one-year Treasury yield.

ratings. For bonds of the highest credit quality, spreads decrease, on average, following a shock to funding liquidity value.[8] In contrast, spreads increase for bonds with lower ratings. We reach a similar conclusion using excess returns from the AAA, AA, A, BBB, and High Yield Merrill Lynch corporate bond indices.

In a final section, we reestimate the model with data up to December 2009. Perhaps unsurprisingly, most of the above results are strengthened and we only report results that highlight the distinctive features of this episode. In particular, we document changes in the funding exposures of swap contracts, high-grade corporate bonds, and U.S. agency bonds. In each case, the direction of risk premium variations predicted by a higher value of funding liquidity changes sign during the financial crisis. The relationship with monetary aggregates is also reversed due to the endogenous response of the Fed to funding conditions. Hence, the funding market provides a propagation and amplification mechanism for shocks to the financial system. The linkages between the value of funding liquidity, the shadow banking sector, and monetary aggregates suggest an important channel between monetary policy and financial asset prices. The impact on risk premia may arise whether the Fed affects funding conditions indirectly through its endogenous response to inflation and real activity or directly through its support to the financial system.

Vayanos and Weill (2006) make explicit the mechanisms linking price differences between two securities with identical cash flows to short-sale frictions and collateral constraints in the repo market (see also Duffie 1996). An investor cannot choose which bond to deliver to unwind a repo position; she must find and deliver the same security she had originally borrowed. These constraints, combined with search frictions on the repo market, imply that the repo rate is lower for the more liquid issue to provide an incentive for bondholders to bring their bonds to the repo market. Typically, recent issues benefit most from the lower financing costs and the greater liquidity, leading to the on-the-run premium. These bonds also offer lower bid-ask spreads, adding to the wedge between asset prices (Amihud and Mendelson 1986). Empirically, both channels seem to be at work, although the effect of funding conditions appears more important.[9] At a broader level, frictions in the repo markets introduce the possibility that markets for bonds with similar cash flows and maturities can be segmented. The preferred-habitat model of Vayanos and Vila (2009) suggests that risk-averse or capital-constrained arbitrageurs may not successfully align the yields of similar bonds with different ages in the case

8 This corresponds to an average effect throughout our sample, although recent events suggest that this is not always the case.

9 Amihud and Mendelson (1991) and Goldreich, Hanke, and Nath (2005) consider transaction costs. Jordan and Jordan (1997); Krishnamurthy (2002); and Cheria, Jacquier, and Jarrow (2004) consider funding costs. Buraschi and Menini (2002) consider funding costs in Germany. Graveline and McBrady (2006) and Banerjee and Graveline (forthcoming) provide evidence that both buy-and-hold investors and short-sellers value liquid securities.

where market habitats are segmented along the age dimension. In this case, the wealth and the risk aversion of arbitrageurs would affect the observed liquidity premium associated with time since issuance.

A few empirical papers document the effects of intermediation constraints on risk premium in specific markets.[10] Instead, we measure the effect of intermediation constraints directly from observed prices rather than quantities. Prices aggregate information and anticipations about intermediaries' wealth, their portfolios, and the margins they face. Prices also aggregate information about market liquidity such as depth and transaction costs. These variables can be difficult to observe or to model. We also study a cross-section of money-market and fixed-income securities, providing evidence that funding constraints should be thought of as an aggregate risk factor driving liquidity premia across markets. Recent empirical investigations are limited to a single market, and none considers the role of funding constraints or funding liquidity.[11]

This article departs from the modern term structure literature in two significant ways. First, the latter focuses almost exclusively on zero-coupon yields.[12] This approach is convenient because a large family of models delivers zero-coupon yields that are linear in the state variables (see Dai and Singleton 2000). However, we argue that pre-processing the data wipes out the most accessible evidence on liquidity, which is the on-the-run or age premium, and turns the funding liquidity factor into a "hidden" factor (Duffee 2011). Therefore, we use coupon bond prices directly. As a consequence, the state space is no longer linear and we handle nonlinearities with the Unscented Kalman Filter (UKF) introduced in Julier, Uhlmann, and Durrant-Whyte (1995). We first estimate a model without liquidity and, notwithstanding differences in data and filtering methodologies, our results are consistent with CDR. Pricing errors in this standard term structure model reveal systematic differences within pairs, correlated with age. Estimation of the model with liquidity produces a persistent factor that captures the differences between the prices of recently issued bonds and the prices of older bonds. The liquidity premium increases with maturity but decays with the age of a bond.

This article is also distinct from the recent literature that uses a reduced-form approach to model a convenience yield in interest rate markets (Duffie and Singleton 1997; Grinblatt 2001; Liu, Longstaff, and Mandell 2006; Fedlhütter and Lando 2008). Jump risk (Tauchen and Zhou 2006) or the debt-GDP ratio (Krishnamurthy and Vissing-Jorgensen 2007) have also been proposed

10 See Froot and O'Connell (2008) for insurance, Gabaix, Krishnamurthy, and Vigneron (2007) for mortgage-backed securities, Gârleanu, Pedersen, and Pothesman (2009) for options, and Hameed, Kang, and Vishnawathan (2010) for equities.

11 See Longstaff (2004) for U.S. Treasury bonds and Collin-Dufresne, Goldstein, and Spencer Martin (2001); Longstaff, Mithal, and Neis (2005); and Ericsson and Renault (2006) for corporate bonds.

12 The CRSP dataset, based on the bootstrap method of Fama and Bliss (1987), is the most common.

to explain the non-default component of corporate spreads. Finally, Pastor and Stambaugh (2003) and Acharya and Pedersen (2005) provide evidence of a liquidity risk factor in expected stock returns.

The rest of the article is organized as follows. Section 1 introduces the term structure model with liquidity, while Section 2 describes the data used to estimate the model. Estimation results are collected in Section 3. In Section 4, we relate the extracted liquidity factor to the tightness of funding conditions. Section 5 measures the impact of funding liquidity on risk premia in several fixed-income markets. In Section 6, we provide a narrative analysis of funding conditions during the 2007–2009 financial crisis. Section 7 concludes.

1. A Term Structure Model with Liquidity

1.1 Term structure model

To build our model, we add a liquidity factor to the Arbitrage-Free Extended Nelson-Siegel (AFENS) model introduced in CDR where the zero-coupon yield at maturity m is given by

$$y(F_t, m) = a(m) + F_{1,t} b_1(m) + F_{2,t} b_2(m) + F_{3,t} b_3(m), \tag{1}$$

and where the latent variables, $F_{i,t}$, have the usual interpretations in terms of level, slope, and curvature.[13] The loadings, $b_i(m)$, are smooth functions of a single parameter, λ, as in the static Nelson-Siegel representation of yields (Nelson and Siegel 1987; hereafter NS). The NS representation is parsimonious, robust to over-fitting, and in line with, or better than, other methods for fitting bond prices out-of-sample in the cross-section of maturities.[14] Its smooth shape is useful to identify deviations of observed yields from an idealized curve.

The AFENS model belongs to the affine family (Duffie and Kan 1996) of term structure models. Intuitively, it corresponds to a canonical affine model (Dai and Singleton 2000) where the loading shapes have been restricted through over-identifying assumptions on the parameters governing the risk-neutral factor dynamics. For our purpose, imposing the absence of arbitrage prevents the model from fitting price differences that are not matched by differences in cash flows. In practice, CDR combine an AFENS model with restrictions on the historical dynamics of factors and document improvements in yield forecasts with respect to the preferred essentially affine model of Duffee (2002).[15]

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13 Diebold and Li (2006) first introduced the Extended NS model (ENS), a dynamic extension of the NS model. We detail the AFENS model in the Online Appendix available at http://jean-sebastienfontaine.com/.

14 See Bliss (1997) and Anderson et al. (1996) for an evaluation of yield curve estimation methods.

15 Joslin, Singleton, and Zhu (2011) show that, absent any restrictions on the historical dynamics, the AFENS model cannot improve forecasts of future term structure factors, F_t, relative to other unrestricted Gaussian DTSM.

1.2 Coupon bonds

Term structure models are usually not estimated from observed prices. Rather, coupon bond prices are converted to forward rates using the bootstrap method. This is convenient since affine models deliver forward rates that are linear in state variables. It is also thought to be innocuous because bootstrapped forward rates achieve near-exact pricing of the original sample of bonds. Unfortunately, this extreme fit means that a naïve application of the bootstrap pushes any liquidity effects and other price idiosyncracies into forward rates. Fama and Bliss (1987) handle this sensitivity to over-fitting by excluding bonds with "large" price differences relative to their neighbors.[16] This approach is certainly justified for many of the questions addressed in the literature, but it removes any evidence of large liquidity effects. Moreover, the FB dataset focuses on prices of discount bonds at annual maturity intervals. This smooths away remaining evidence of liquidity effects in intermediate forward maturities.

We proceed directly from observed coupon bonds with maturity M, say, and with coupons at maturities $m = m_1, \ldots, M$. To value intermediate payoffs, we use the price of a discount bond with maturity $m > 0$, $D_t(m)$, given by

$$D_t(m) = \exp\left(-m(a(m) + b(m)^\top F_t)\right),$$

which follows directly from Equation 1, with the obvious vector notation. In a frictionless economy, the absence of arbitrage implies that the price of a coupon bond equals the sum of discounted coupons and principal. That is, the frictionless price is

$$P^*(F_t, Z_{n,t}) = \sum_{m=m_1}^{M} D_t(m) \times C_t(m), \tag{2}$$

where $Z_{n,t}$ includes (deterministic) characteristics relevant for pricing a bond such as the schedule of coupons and principal payments, $C_t(m)$.

1.3 Coupon bonds in an economy with frictions

The arbitrage restrictions given by Equation (2) do not hold with equality in an economy with frictions. Simultaneously selling the relatively expensive and buying the relatively cheap Treasury bonds is not feasible. An investor cannot issue new U.S. Treasury securities to establish a short position. Instead, she must borrow the bond on the repo market through a collateralized loan and,

[16] See the CRSP documentation. Briefly, a first filter includes a quote if its yield to maturity falls within a range of 20 bps from one of the moving averages on the three longer or the three shorter maturity instruments *or* if it falls between the two averages. Precedence is given to bills to compute averages to exclude the impact of liquidity on notes and bonds with a maturity of less than one year. Amihud and Mendelson (1991) document that yield differences between notes and adjacent bills are 43 bps on average. The second filter excludes observations that cause reversals of 20 pbs in the bootstrapped discount yield function. To our knowledge, the impact of these filters has not been studied.

in a separate transaction, she must use some of her own capital and fund the purchase of the other bond via the repo market. Furthermore, the investor must find and deliver the same security she had originally borrowed to unwind her position. Luttmer (1996) shows that, in this case, the set of stochastic discount factors consistent with the absence of arbitrage satisfies $P \geq P^*$. Therefore, we model the price, $P(F_t, L_t, Z_t)$, of any coupon bond with characteristics Z_t as the sum of discounted coupons to which we add a liquidity term:

$$P(F_t, L_t, Z_{n,t}) = \sum_{m=1}^{M_n} D_t(m) \times C_{n,t}(m) + \zeta(L_t, Z_{n,t}),$$

which must remain non-negative.[17] Grouping observations together and adding an error term gives the measurement equation,

$$P(F_t, L_t, Z_t) = C_t D_t + \zeta(L_t, Z_t) + \Omega \nu_t, \tag{3}$$

where C_t is the matrix of payoffs, D_t is a vector of discount bond prices, $\zeta(L_t, Z_t)$ is a vector of liquidity premia, and ν_t is a Gaussian white-noise vector uncorrelated with innovations in state variables.[18] The matrix Ω is assumed diagonal, and its elements are a linear function of maturity, $\Omega_n = \Omega_0 + \Omega_1 M_n$.[19]

The liquidity premium applies to all bonds, old and new. Our specification is based on a latent factor that drives the common dynamics but with loadings varying with the maturity and the age of each bond. The liquidity premium is given by

$$\zeta(L_t, Z_{n,t}) = L_t \times \beta_{M_n} \exp\left(-\frac{1}{\kappa}\, age_{n,t}\right), \tag{4}$$

where $age_{n,t}$ is the age, in years, of the bond at time t and β_M controls the average premium at each fixed maturity M. We estimate β for a fixed set of maturities but leave the shape of β unrestricted between these maturities.[20] The parameter κ controls the liquidity premium's decay with age. For instance, immediately following its issuance (i.e., $age = 0$), the loading on the liquidity factor is $\beta_M \times 1$. Taking $\kappa = 0.5$, the loading decreases by half after four months: $\zeta(L_t, 4) \approx \frac{1}{2}\zeta(L_t, 0)$).

[17] Section F of the Online Appendix shows how to generalize the AFENS Nelson-Siegel to accomodate additional variables in the state vector and maintain the no-arbitrage restriction.

[18] C_t is an $(N \times M_{max})$ matrix obtained from stacking the N row vectors of individual bond payoffs, and M_{max} is the longest maturity in the sample. Shorter payoff vectors in C_t are completed with zeros. Similarly, D_t is a $(M_{max} \times 1)$ vector, while $\zeta(L_t, Z_t)$ and ν_t are $N \times 1$ vectors.

[19] This reduces substantially the dimension of the estimation problem. We verified that leaving the diagonal elements of Ω unrestricted did not affect materially our results.

[20] But note that we must have $\beta_0 = 0$.

The specification above reflects our priors about the impact of age and maturity. The scale parameters are left unrestricted at estimation, and we allow for a continuum of shapes for the decay of liquidity. Equation 3 shows that omitting the last term pushes the impact of liquidity into pricing errors, possibly leading to biased estimators and large filtering errors. Alternatively, adding a liquidity term amounts to filtering a latent factor present in pricing errors. But Equation 3 shows that this factor can only capture that part of pricing errors common across bonds and correlated with age. We have motivated earlier why age can be a good variable to capture a liquidity premium similar to an on-the-run premium for recent issues. We show below that residuals from a model without a liquidity factor are correlated with age and display a factor structure. The impact of age on the price of a bond can hardly be rationalized in a frictionless economy. Price differences due to funding costs and funding liquidity are real arbitrage opportunities unless we explicitly consider the costs and the risks involved when shorting the more expensive bond and buying the cheaper bond. A structural approach should deliver a joint model of the bond and repo markets and include variations in funding rates, funding liquidity, arbitrage capital, or other frictions.[21] However, search frictions and collateral constraints on the repo markets are either difficult to observe or available for limited subsamples. We check the validity of our empirical strategy in Section 4.

2. Data

We use end-of-month prices of U.S. Treasury securities from the CRSP dataset. Our sample covers the period from January 1986 to December 2009. However, most of our estimation results exclude the 2008–2009 period to avoid tilting the findings in favor of our model. We include the 2008–2009 data in a last section to assess the robustness of our results and discuss the unusual actions of the Federal Reserve in this exceptional period. Before 1986, interest income had a favorable tax treatment compared to capital gains and investors favored high-coupon bonds. In that period, interest rates rose steadily and recently issued bonds had relatively high coupons and were priced at a premium both for their liquidity and for their tax benefits. The resulting tax premium cannot be disentangled from the liquidity premium using bond ages. Green and Ødegaard (1997) confirm that the tax premium mostly disappeared when the asymmetric treatment of interest income and capital gains was eliminated following the 1986 tax reform.

The CRSP dataset provides quotes on all outstanding U.S. Treasury securities. We construct bins around maturities of 3, 6, 9, 12, 18, 24, 36, 48, 60, 84, and 120 months. Then, at each date, and for each bin, we choose a pair of securities to identify the liquidity premium. First, we pick the most recently

[21] These features are absent from modern term structure models, with the notable exception of Cheria, Jacquier, and Jarrow (2004), who allow for a convenience yield accruing to bondholders.

issued security in each maturity bin. Second, we pick the security that most closely matches the bin's maturity.[22] By construction, securities within each pair have very similar durations and exposure to interest rate risk, the same credit quality, and offer the same compensation for inflation and inflation risk. Hence, the most important aspect of our sample is that whenever a security trades at a premium relative to its pair companion, any large price difference cannot be rationalized from small coupon or maturity differences under the no-arbitrage restriction of the CDR term structure model. On the other hand, price differences that are common across maturities and correlated with age will be attributed to liquidity. Note that pinning the older securities at fixed maturities ensures a stable coverage of the term structure of interest rates for the purpose of estimation. This also introduces variability in age differences which, in turn, identifies how the liquidity premium varies with age. Finally, the most recent issue can at times be somewhat old due to a quarterly issuance pattern, to the absence of new issuance in some maturity bins throughout the whole sample (e.g., 18 months to maturity) or within some subperiods (e.g., 84 months to maturity). Far from a drawback, this provides further variability in age differences and helps identify the relationship with price differences.

We now investigate some features of our sample of 5,830 observations.[23] The first two columns of Table 1 present means and standard deviations of age for each liquidity-maturity category. Typically, the old security has been in circulation for more than a year. In contrast, the recent security is typically a few months old and only a few weeks old in the 6- and 24-month categories, indicating a regular issuance pattern. On the other hand, the relatively high standard deviations of ages in the 36- and 84-month categories reflect the decision by the U.S. Treasury to stop the issuance cycles at these maturities. Table 1 also presents means and standard deviations of duration. Average duration is almost linear in maturity and, as expected, pairs have very similar duration. The last columns of Table 1 show that the term structure of coupons is upward sloping, on average. The high standard deviations are in part due to the general decline of interest rates throughout the sample. More importantly, coupon differences within pairs are small, on average. To summarize, differences in duration and coupon rates are kept small within each pair, but differences of ages are magnified so that we can identify any effect of liquidity on prices that is correlated with age.

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22 See Elton and Green (1998) and Piazzesi (2005) for discussions of the CRSP dataset. See the Online Appendix for more details on data filters. In particular, we exclude securities issued more than five years before the observation date, in effect excluding very old bond issues with 20-year or 30-year maturities at the time of issuance. Unreported results based on subsamples restricted to recent issues only or to old issues only show that combining information from the broader cross-section, as we do here, provides a better measurement of L_t and better predictability results.

23 Our sample is 265 months, times 11 maturities, times 2 bonds, one recently issued and one older bond.

Table 1
Summary statistics of bond characteristics

Maturity	Age				Duration				Coupon			
	Old		New		Old		New		Old		New	
3	12.01	9.31	1.64	0.09	3.01	0.03	4.38	0.09				
6	16.93	6.27	0.12	0.11	6.00	0.10	5.90	0.11				
9	14.45	6.05	4.42	4.88	8.89	0.11	10.00	0.40				
12	13.11	5.78	2.51	3.90	11.77	0.23	12.14	1.08				
18	28.29	11.92	6.74	0.62	17.14	0.50	16.81	0.59	7.12	2.81	6.84	3.06
24	22.90	13.45	0.33	0.52	22.56	0.59	22.68	0.72	7.11	2.82	6.74	3.13
36	24.64	10.17	4.61	6.74	32.56	1.42	32.75	2.63	7.49	2.96	7.10	2.82
48	18.42	9.57	4.42	3.00	41.95	2.30	44.17	4.40	7.38	3.03	7.25	2.86
60	29.06	21.58	2.29	3.85	50.41	3.09	51.36	3.02	7.72	2.80	7.13	2.90
84	34.41	8.61	12.51	11.82	65.85	5.04	68.71	8.45	7.74	2.56	7.55	2.63
120	14.91	18.59	4.02	7.56	84.43	8.34	85.55	9.16	7.15	2.24	7.44	2.71

We present summary statistics of age (in months), duration (in months), and coupon (in %) for each maturity and liquidity category. New refers to the low-age securities, and Old refers to the high-age securities (see text for details). In each case, the first column gives the sample means and the second column gives the sample standard deviations. Coupon statistics are not reported for maturity categories of 12 months and less since T-bills do not pay coupons. End-of-month data from CRSP (1985:12–2007:12).

3. Estimation Results

The price function in (3) together with an autoregressive dynamics for the factors $X_t \equiv [F_t, L_t]$ define a state-space system,

$$(X_t - \bar{X}) = \Phi_X(X_{t-1} - \bar{X}) + \Sigma_X \epsilon_t,$$
$$P_t = \Psi(X_t, C_t, Z_t) + \Omega \nu_t, \tag{5}$$

where Ψ is the (nonlinear) mapping of cash flows C_t, bond characteristics, Z_t, and current states, X_t, into prices, P_t. A filtering algorithm allows us to build the following log-likelihood function,

$$L(\theta) = \sum_{t=1}^{T} l(P_t; \theta) = \sum_{t=1}^{T} \left[\log \Phi(\hat{P}_{t+1|t}, R_{t+1|t})\right], \tag{6}$$

where $\Phi(\cdot, \cdot)$ is the multivariate Gaussian density and $\hat{P}_{t+1|t}$ and $R_{t+1|t}$ are the predicted bond prices and the corresponding mean square pricing errors, respectively. When Ψ is linear, the Kalman filter is optimal and ML estimators are feasible. Instead, we use the Unscented Kalman Filter algorithm that provides approximations of these conditional moments. Nonetheless, the measurement equation is nonlinear and parameter estimates of θ are based on a Quasi-Maximum Likelihood Estimator (QMLE). From standard results (White 1982), this estimator is asymptotically normal and converges to a well-defined limit. The variance-covariance matrix is based on both the Hessian and the outer product of the scores evaluated at the optimum. This holds even when the model is misspecified and, in particular, when we employ an approximate filter.[24]

3.1 Results for the benchmark model without liquidity

First, we estimate the benchmark model without liquidity and verify that we obtain similar risk factors than CDR since our data and estimation methods differ. They use smoothed zero-coupon yields, while we use observed coupon bond prices. They use a linear Kalman filter to build their likelihood function, while we use a nonlinear filter. Notwithstanding these potentially important differences, we obtain very similar filtered risk factors, given in Figure 1A. The level factor is very persistent and declines slowly throughout the sample period. The slope factor is slightly less persistent and exhibits the usual association with business cycles. Its sign changes before the recessions of 1990 and 2001 and during the so-called "conundrum" episode. Finally, the curvature factor is closely related to the slope factor.

[24] For space considerations, all details about the state-space system, the filtering algorithm, estimation, and inference are provided in the Online Appendix.

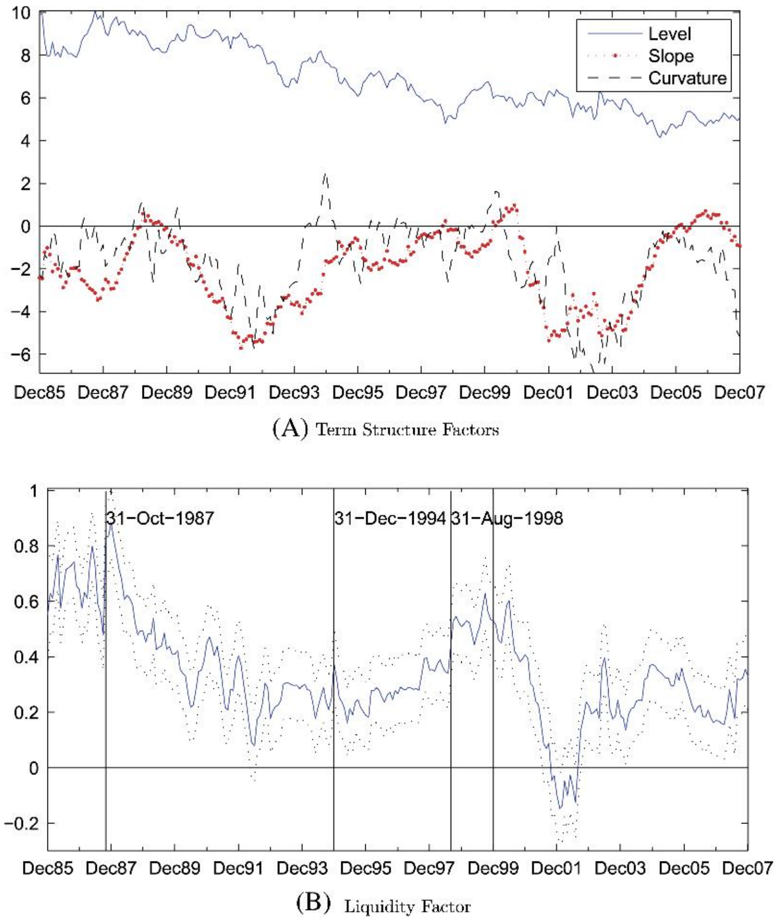

Figure 1
Funding liquidity and term structure factors
Factors from the AFENS model with liquidity. Panel A displays the term structure factors. The scale is in percentages. Panel B displays the funding liquidity factor. The scale is in dollars, and the dotted line provides 95% confidence intervals around the filtered factor at each date based on the mean squared errors estimates from the Kalman filter. End-of-month data from CRSP (1985:12–2007:12).

Panel A of Table 2 gives more information on the fit of the benchmark model. Root Mean Squared Errors (RMSE) increase from \$0.047 and \$0.046 for recent and older 3-month securities, respectively, to \$0.35 and \$0.39 for 10-year securities. The monotonous increase of RMSE with maturity is standard. It may reflect the higher sensitivity of longer maturity bonds to interest rates. It may also be due to higher uncertainty surrounding the true prices, as signaled by wider bid-ask spreads. The RMSE is higher for recent bonds, and it is \$0.188 in the entire sample. Panel A also confirms that Mean Pricing Errors (MPE) are systematically higher for bonds with lower ages. Their price is

Table 2
Pricing errors in the benchmark model without liquidity

Panel A: Pricing error statistics

	MPE		RMSE	
Maturity	Old	New	Old	New
3	0.015	0.026	0.047	0.046
6	−0.002	0.020	0.035	0.041
9	−0.031	0.026	0.054	0.062
12	−0.041	0.037	0.072	0.078
18	−0.064	−0.061	0.092	0.090
24	−0.028	0.000	0.063	0.085
36	0.005	0.069	0.102	0.139
48	−0.008	0.079	0.171	0.183
60	0.006	0.250	0.226	0.318
84	−0.167	−0.041	0.327	0.298
120	−0.239	0.070	0.353	0.394
All	−0.050	0.043	0.177	0.197

Panel B: Age differences and price differences

Age Difference	<3 mths	3-15	15-30	30-42	> 42 mths
Mean Age Difference	1.62	8.97	20.68	35.28	55.40
Mean Residual Difference	0.0418	0.0709	0.1191	0.1667	0.2407
Standard Errors	0.0064	0.0050	0.0111	0.0177	0.0262

Panel C: Principal components of residual differences

	Factor Loadings		
Maturity	PC1	PC2	PC3
3	0.0090	−0.0149	−0.0030
6	0.0124	−0.0186	0.0008
9	0.0339	−0.0620	0.0003
12	0.0360	−0.0641	−0.0222
18	0.0222	0.0037	0.0057
24	0.0675	−0.0963	−0.0327
36	0.2081	−0.0002	−0.0772
48	0.2856	−0.3192	−0.3833
60	0.4855	−0.5619	−0.3064
84	0.7527	0.6347	0.0757
120	0.2550	−0.4021	0.8636
	Percentage of Variance Explained		
R^2	65.6%	14.1%	10.1%

Panel A presents Mean Pricing Errors (MPE) and Root Mean Squared Errors (RMSE) (in \$) within each Age-Maturity subgroup. Panel B presents averages of residual differences (in \$) between *New* and *Old* issues for all maturities but in different age-difference categories (in month). Panel C presents a Principal Components Analysis of residual differences (in \$) between *New* and *Old* issues for different maturity groups. Residuals computed from AFENS model without liquidity. End-of-month data from CRSP (1985:12–2007:12).

higher than what can be rationalized in a no-arbitrage framework, relative to their older counterparts. For a recent 12-month Treasury bill, the average difference is close to \$0.08. Similarly, the pricing error on a recently issued 5-year bond is \$0.25 higher, on average, than a similar but older issue.[25]

25 Consistent with Amihud and Mendelson (1991), the price impact of liquidity increases with maturity.

3.2 Relation between pricing errors and age

Panel B of Table 2 presents the average price-residuals difference for pairs in age-difference bins of 3 months, 1 year, 2 years, 3 years, and 4 years or more. Again, residuals are computed relative to the ENS model with no liquidity term. Hence, this reflects the price difference between similar bonds unaccounted for by a no-arbitrage term structure model. In the 3-month category, the older bond is worth 4.18 cents less than the more recent bond and has been issued 1.62 months before, on average. In the second category, the price difference is 7.09 cents and the age difference is 9 months, on average. The premium then increases to 11.91 cents, 16.67 cents, and 24.07 cents in older categories. The results clearly show that the premium rises gradually as we consider longer age differences and that the relationship extends far into the age spectrum. The high average premium *level* when the age difference is large should not be confounded with the large *rate* at which the premia decay for small age differences. Clearly, the age-premium relationship is steepest when a new note or bond is issued and replaces the older, now just-off-the-run issue. However, this does not preclude that the prices of off-the-run bonds keep decreasing with age, even at a low rate.

Next, we want to show that there exists a common and significant factor driving the variations of the liquidity premium across maturities. Panel C of Table 2 presents the first three principal components of residual differences, which explain 66%, 14%, and 10% of total variations, respectively. Factor loadings confirm that the first component is the most economically meaningful. The premium at *every* maturity increases whenever this component increases. Such a large common component can hardly be induced by measurement errors. Moreover, it is striking that a common factor emerges so clearly even when not controlling for bond ages. For example, the 7-year maturity category has had regular issuance, and a large liquidity premium, in about half our sample but the premium disappears when issuance stops. The second component seems to pick this effect. It affects all premia in the same direction but with the opposite sign for the 7-year maturity category. Much of these component variations will be captured by a specification controlling for age. The last component seems to capture variations specific to the 10-year maturity group.[26]

3.3 Results for the liquidity model

Estimation of the unrestricted model leads to a substantial increase of the log-likelihood. The benchmark model is nested with 15 parameter restrictions, and the improvement in likelihood is such that the LR test-statistic leads to a p-value that is essentially zero.[27] Parameter estimates imply average short- and

[26] This is consistent with Fleming (2003), who finds that the 10-year on-the-run premium has a large idiosyncratic component.

[27] The benchmark model reached a maximum at 1,998.6, while the liquidity model reached a maximum at 3,482.6. The liquidity factor is available at http://jean-sebastienfontaine.com.

long-term discount rates of 4.09% and 5.76%, respectively, compared to 3.73% and 5.45%, respectively, for the model without liquidity. Therefore, the yield curve is on average between 35 bps and 31 bps higher than estimates from a model without liquidity. Figure 1B presents filtered values of the liquidity factor. It exhibits significant variations through normal and crisis periods. In particular, the stock crash of 1987, the Mexican peso devaluation of December 1994, and the Long Term Capital Management (LTCM) failure of 1998 are associated with peaks in L_t. Estimates of the level of the liquidity premium for the various maturities, $\hat{\beta}$, are given in Table 3. They increase with maturity,[28] and the pattern accords with the residuals from the model without liquidity. Table 3 shows that the residual differences are reduced considerably by the introduction of the liquidity factor.

A formal test rejects the null hypothesis of zero-mean residual differences in the 24-month group. Nonetheless, the liquidity premium is smaller relative to the surrounding maturities. This small premium may appear surprising given the regular issuance cycle, and we can only conjecture as to its causes. Interestingly, Jordan and Jordan (1997) show that 2-year notes remain "special" for shorter periods of time and, moreover, conclude against any specialness effect on prices at that maturity.[29] Similarly, Goldreich, Hanke, and Nath (2005) find that the on-the-run premium on 2-year Treasury notes decreases faster than other maturities, on average. This, together with its short issuance cycle, suggest a lower role for the special status of the recent issue

Table 3
Liquidity premium

	Residual Difference			
Maturity	Benchmark	Liquidity	$\hat{\beta}$	Standard Error
3	0.0111	−0.0053	0.2642	0.03041
6	0.0221	−0.0295	0.2837	0.03261
9	0.0566	0.0202	0.3158	0.03709
12	0.0783	0.0396	0.3026	0.03622
18	0.0025	−0.0036	0.0428	0.02481
24	0.0280	−0.0117	0.2005	0.03207
36	0.0644	−0.0260	0.5325	0.07391
48	0.0892	0.0165	0.7446	0.09452
60	0.2477	0.0102	1.2270	0.13694
84	0.1250	−0.0509	1.2174	0.10268
120	0.3106	0.2640	1	–

Each line corresponds to a maturity category (months). The first two columns provide average residual differences between *old* and *new* securities for the AFENS model with and without liquidity, respectively. The last two columns display estimates and QML standard errors of the liquidity level parameters, $\hat{\beta}$. End-of-month data from CRSP (1985:12–2007:12).

28 We fix $\beta_{10} = 1$ to identify the level of the liquidity factor with the average premium of a just-issued 10-year bond. The estimated average level is lower in the 10-year group relative to the 5-year and 7-year groups. This is due to the lower average age of bonds in these groups.

29 See Table I of Jordan and Jordan (1997), p. 2057, as well as p. 2061.

in the premium variation. Moreover, the magnitude of the premium also depends on the relative benefits offered by the recent issues. Fleming (2003) documents the larger volume of transactions and the higher liquidity of securities around this key maturity. One hypothesis is that the two-year mark serves as a focus point for buyers and sellers whereas holders of long-term bonds reallocate funds from their now short-term maturity bonds into newly issued longer-term securities. Ultimately, these two effects may reduce the liquidity gap across securities with different ages.

4. The Value of Funding Liquidity

The previous section points toward an important factor related to age behind the residuals of a usual three-factor term structure model in the cross-section of U.S. Treasury securities. We have argued, based on the previous theoretical and empirical literature, and the observed price differences in the repo market, that the observed premia between younger and older bonds are liquidity premia. Brunnermeier and Pedersen (2008) emphasize the importance of collateral constraints and funding liquidity as a determinant of market prices. In this section, we provide compelling evidence linking our extracted liquidity factor to measures of funding conditions at three successive levels of aggregation: at the level of the Treasury spot and funding markets, at the level of the shadow banking sector and, finally, at the level of monetary aggregates. In the following, and throughout the remainder of the article, we refer to L_t as the value of funding liquidity or as the liquidity factor interchangeably.

4.1 Funding liquidity and funding costs

The higher prices of the more liquid Treasury securities reflect the benefits, $b_{n,t}$, offered by these bonds over some holding horizon (Goldreich, Hanke, and Nath 2005; Krishnamurthy 2002). Then, to the extent that the liquidity factor reflects these future benefits, we can write

$$L_t = E_t\left[\delta_{t,h} b_{t,t+h}\right] = \delta_{t,h} b_{t,t+h} + \epsilon_{t,t_h}, \tag{7}$$

where $\delta_{t,h}$ is a discount factor and $b_{t,t+h}$ aggregates future benefits over a horizon, h. In practice, we use different proxies for these benefits. First, we obtain daily general collateral (GC) and special collateral (SC) overnight repo rates from GovPX between November 1995 and December 2009. The difference between GC and SC rates measures the funding cost advantage of special, typically recently issued, securities in the overnight market. SC rates are available without interruption for bonds with 2-year, 5-year, 10-year, and 30-year maturities. For each maturity, we compute its spread relative to the GC rate and label it the repo spread. Second, we run a principal component analysis to extract a repo spread factor common to all maturities. The first component loads positively on each repo spread and explains 56% of total variation. Finally, we compute from the CRSP dataset the difference between the median and the minimum bid-ask spreads across all bonds, labeled the BA spread. This controls for differences across the transaction costs of the most liquid bond and median bonds. The benefits of lower funding costs and lower transaction costs are particularly important in the case of on-the-run or just-off-the-run securities. However, the liquidity factor captures price variations across a broad cross-section of bonds where older issues offer similar transaction and funding costs. In particular, Bartolini et al. (2011) show that Treasury securities that are not "special" do offer substantial funding advantages relative to other bonds (i.e., agency bonds) and relate these advantages to higher prices for these off-the-run bonds. Therefore, we expect that only part of the liquidity factor will be explained by future variations in trading and funding costs of securities with a "special" status.[30]

Hence, each month, we match the current funding liquidity factor with average BA spread, the averages of each repo spread, and the average of the repo spread factor computed over the following three months.[31] We then model the common component as $b_{t,t+h} = a + b_h' x_{t,t+h} + u_t$, where $x_{t,t+h}$ includes different proxies. This implies the following regression for L_t:

$$L_t = \alpha + \beta' x_{t,t+h} + \tilde{\epsilon}_t, \tag{8}$$

where $\tilde{\epsilon}_t$ captures the variation of L_t that is unrelated to $x_{t,t+h}$. We expect each of these proxies to be positively related to the liquidity factor.

Table 4 presents regression results for different combinations of proxies from 1995 to 2007. The first line (Model A) shows that future transaction costs have weak explanatory power for the value of funding liquidity, with an R^2 of 1.6%. However, Figure 2A shows that a break occurred in the behavior of bid-ask spreads following the advent of the GovPX platform. A second break matches the introduction of the eSpeed electronic trading platform. The R^2 increases to 37.7% when we include data prior to 1995 in the regression. Transaction costs contribute to variations of the on-the-run premium (see, e.g., Goldreich, Hanke, and Nath 2005) but play a lesser role in the evolution of the value of funding liquidity since these breaks. Model B in Table 4 shows that the expected repo spread factor accounts for 21% of liquidity factor variations. The coefficient is highly significant and implies that a one-standard-deviation change increases the liquidity factor by half a standard deviation. The remaining models (i.e., C to F) combine the repo spread factor, the BA spread, and each of the individual repo spreads, one at a time, and ask whether

[30] Also, the error term, ϵ_{t,t_h}, includes any risk adjustment induced by the covariance between $\delta_{t,h}$ and $b_{t,t+h}$. Therefore, empirical estimates will combine the effect of funding liquidity and its risk premium. We found no significant relationship with the Pastor-Stambaugh measure of stock market liquidity risk. Note that this measure is related to spot market liquidity, while our measure is related to funding market liquidity.

[31] A three-month window is a reasonable holding period over which the benefits of holding a liquid bond accrue to the investor. The results are robust to the choice of window. Moreover, $D_t(3)$ is very close to one and is neglected for simplicity.

Table 4
Market determinants of funding liquidity

Model	*cst*	*BA*	Spread Factor	RSP2	RSP5	RSP10	RSP30	R^2
A	0.28	0.020						1.6
	(12.8)	(1.74)						
B	0.28		0.07					21.0
	(12.9)		(4.09)					
C	0.28	−0.005	0.08	−0.01				21.4
	(14.7)	(0.43)	(4.20)	(0.58)				
D	0.28	−0.006	0.08		−0.02			22.5
	(14.8)	(0.59)	(3.95)		(0.96)			
E	0.28	−0.001	0.07			−0.01		21.1
	(14.6)	(0.09)	(2.48)			(0.21)		
F	0.28	−0.05	0.05				−0.02	21.6
	(14.7)	(0.50)	(2.10)				(0.83)	

Results from regressions of the funding liquidity factor, L_t, on selected market variables. BA is the 3-month rolling average of the difference between the minimum and the median bid-ask spreads in the cross-section of maturity at each date. RSP2, RSP5, RSP10, and RSP30 are 3-month averages of daily repo spreads for 2-year, 5-year, 10-year, and 30-year Treasury securities. The spread factor is the first Principal Component from these average repo spreads. Regressors are demeaned and divided by their standard deviations. Newey-West *t-statistics* (3 lags) in parentheses. End-of-month data (1995:11–2007:12).

idiosyncratic variations of individual spread increase our ability to explain the liquidity factor. The unambiguous answer is no. Individual spread coefficients are not significant, have the wrong sign, and do not increase the R^2. In each case, the repo spread factor coefficient is stable and significant. We conclude that the common spread factor is an important determinant of the liquidity factor.

4.2 Funding liquidity and shadow banking

Adrian et al. (2010) identify the shadow banking sector: a large non-bank component of the intermediation system that relies heavily on short-term funding.[32] Its key economic function is to channel short-term financing toward long-lived illiquid assets but outside the traditional banking sector. The shadow banking sector was at the heart of the securitization boom and bust. We adopt a simple price-quantity interpretation of the liquidity factor and of the quantity of funding liquidity supplied by market-based intermediaries that compose the shadow banking sector. Our hypothesis is that the demand and supply of funding liquidity vary with our measure of funding liquidity value. Positive demand shocks should tighten conditions and increase funding liquidity value along the funding supply curve. Conversely, positive supply shocks should be associated with a decline of funding liquidity value along the funding demand curve.

We measure the supply of funding liquidity via the log-growth of financial assets held in the shadow banking sector (see Figure 3A) and test the prediction that supply increases with the price of funding liquidity. The supply curve

[32] Adrian, Moench, and Shin (2010) aggregate balance sheet data from agency and GSE-backed mortgage pools, issuers of asset-backed securities, finance companies, and funding corporations available quarterly from the Federal Reserve Board's Flow of Funds Accounts.

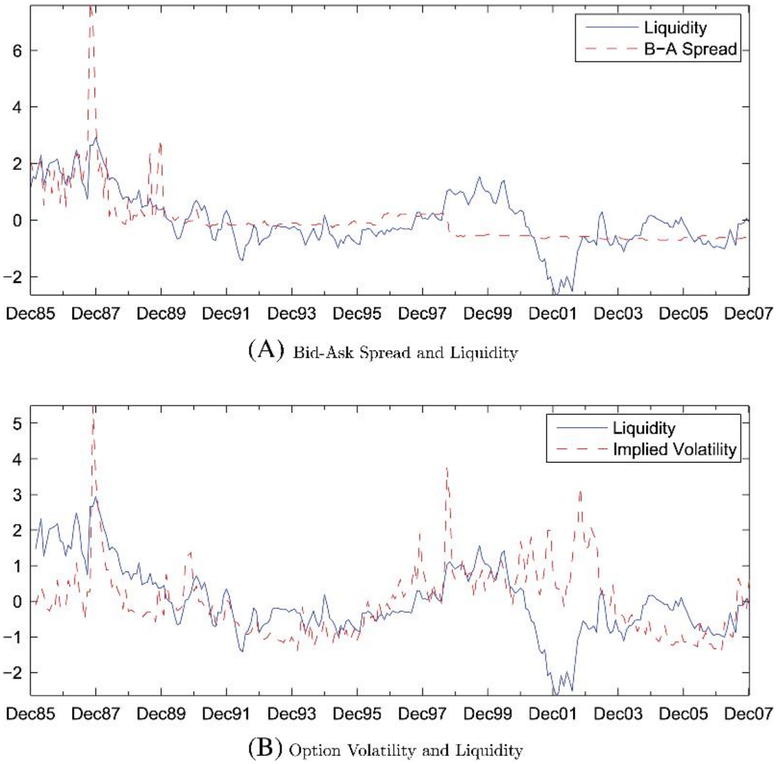

Figure 2
Funding liquidity, market liquidity, and option volatility
Panel A traces the funding liquidity factor against the difference between the median and the minimum bid-ask spreads across maturities for a given date. Panel B traces the funding liquidity factor against the VXO index of implied volatility from S&P 500 call options. The funding liquidity factor is obtained from the AFENS model with liquidity. End-of-month data from CRSP (1985:12–2007:12).

describing the relationship between the log-growth of assets held in the shadow banking sector, q_t^s, and a measure of its price, the value of funding liquidity, is given by

$$q_t^s = \alpha + \beta_l L_t + \beta_x' X_t + u_t, \tag{9}$$

and we should have $\beta_l > 0$. The vector X_t contains other predetermined economic or financial indicators that shift the supply curve. These variables must affect the terms at which different quantities of funding liquidity are provided but, in turn, remain relatively unaffected by the resulting equilibrium in the funding market. First, we use the level, slope, and curvature of the term structure. The configuration of short- and long-term interest rates is determined by conditions in the broader economy but affects the trade-offs relevant in the funding supply decision. Lower interest rates or a steeper term structure increase the profitability of leveraging long-term assets. Second,

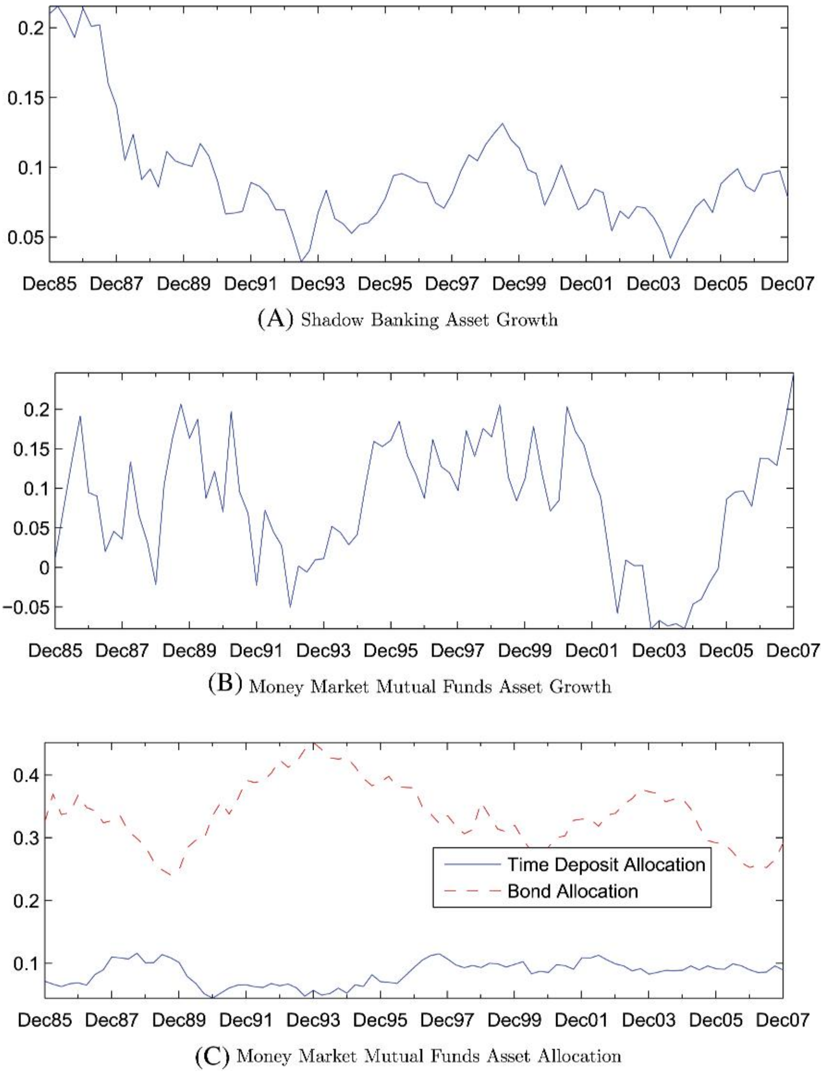

Figure 3
Shadow banking asset growth
Asset growth in the shadow banking sector (Panel A), asset growth of money market mutual funds (MMMF) (Panel B), and allocation to time deposit as well as allocation to Treasury, agency, and municipal bonds (i.e., Bond Allocation) by MMMR (Panel C). Quarterly data (Q1:1986–Q4:2007).

we include measures of flight-to-liquidity based on flows into and within the MMMF sector. MMMF provide the main channel through which the shadow banking sector obtains short-term funding. Hence, the ability of the shadow banking sector to supply funding varies with the size of these funds and with their allocation to risky assets. From the Flow of Funds Accounts we construct a quarterly series of MMMF asset growth, labeled MMG, the proportion of assets allocated to time deposits, $MMA1$, and the proportion of assets allocated to Treasury, agency, or municipal bonds, $MMA2$. These are displayed in the two lower panels of Figure 3.

Equation 9 represents the supply schedule. The coefficient β_l in the supply equation cannot be estimated consistently via OLS since we do not observe q_t^s, but, rather, the equilibrium asset growth in the shadow banking sector. Clearly, q_t and L_t are jointly endogenous and L_t is correlated, via the demand equation, with the residuals of a regression where we substitute q_t as the dependent variable. The intuition from the simultaneous equation literature suggests that a valid instrument for price in the supply equation can be found among the predetermined variables that are unique to the *demand* equation. This guarantees that the instrument is both valid and relevant. It is relevant since it is a determinant of demand and, therefore, correlated with price. It is valid since it affects the equilibrium price and quantity only via demand shifts and, therefore, is uncorrelated with the residuals from the supply equation. We argue that variations in the aggregate quantities of residential and commercial mortgages in the U.S. economy provide a good instrument. Mortgage growth increases the total amount of illiquid assets in the economy, yet shifts the demand for funding liquidity while leaving the supply curve unaffected. The growth of mortgages in this period can be attributed to some mix of government policies favoring homeownership, a low-interest-rate environment, a degradation in lending practices, innovations in retail mortgage products, poor securitization practices, or sheer exuberance about the future path of house prices. On the other hand, we argue that, in our pre-crisis sample, mortgage growth did not affect the trade-offs a shadow bank faces when drawing its supply schedule. In other words, variations in the aggregate stock of mortgages are a valid instrument since they are uncorrelated with shocks to the supply schedule. Of course, the value of funding liquidity and the quantity of funding supplied are correlated with mortgage growth in equilibrium, but this follows from the effect of mortgage growth on the demand curve.

Results from OLS and IV regressions are presented in Table 5. Reduced-form coefficients are counter intuitive. They imply that funding liquidity value, mortgage growth, the slope of the term structure, and flows into the MMMF sector have no impact on the shadow banking sector's size. Moreover, results suggest that asset growth *increases* with the level of interest rates and with the proportion of MMMF assets allocated to time deposits. But it is well known that estimates of β_l based on the reduced-form regression are a weighted average of the (negative) coefficient from the demand curve and the (positive) coefficient from the supply curve. Moreover, the presence of one endogenous relationship biases the estimates of all the reduced-form regression coefficients.

Using mortgage growth as an instrument leads to much more plausible results. Strikingly, the IV regression indicates that the liquidity factor has the

Table 5
Supply and price of funding liquidity

	Regressors									
	cst	*L*	Level	Slope	Curv.	MMG	MMA1	MMA2	Mrtg.	R^2
OLS	0.091	0.001	0.024	0.011	−0.002	0.009	0.028	−0.003	0.014	
	(19.7)	(0.15)	(2.98)	(1.26)	(−0.46)	(1.79)	(3.27)	(−0.56)	(1.40)	60.0
2SLS (1)	0.099		0.25	0.43	0.70	−0.01	0.24	0.62	−0.25	
	(17.84)		(4.43)	(5.99)	(7.38)	(−0.14)	(3.92)	(6.56)	(−4.25)	69.1
2SLS	0.091	0.072	−0.025	−0.069	0.027	0.010	−0.030	−0.057		
	(19.70)	(3.41)	(−1.92)	(−3.05)	(3.24)	(1.89)	(−2.77)	(−3.68)		61.9

Results from OLS and 2SLS regressions (first and second stages) of shadow banking asset growth on the funding liquidity factor (*L*), the Level, Slope, and Curvature factors from the term structure of interest rates, Money Market Mutual Funds (MMMF) asset growth (MMG), MMMF allocation to time deposits (MMA1), and MMMF allocation to Treasury, agency, and municipal bonds (MMA2). Aggregate mortgage growth is the instrument (Mrtg). Regressors are demeaned and divided by their standard deviations. Standard instrumental variables *t-statistics* in parentheses. End-of-month data (1985:12–2007:12).

highest impact on shadow banking. Asset growth increases by 7.2% when the value of funding liquidity increases by one standard deviation. This is consistent with the quantity of funding supplied rising with marginal revenue. Other implications from the IV regressions are also economically intuitive. Asset growth in the shadow banking sector decreases by 2.5% following a one-standard-deviation increase in the level of interest rates, by 6.9% when the term structure slope flattens by one standard deviation, and by 2.7% following a one-standard-deviation decrease of its curvature. This is consistent with reductions of the profitability of funding operations inducing shifts to the supply schedule. Moreover, asset growth increases by 1.0% following inflows into MMMF but decreases by 3.0% when these funds increase their allocations to time deposits and by a stunning 5.7% when they increase their allocations to Treasury, agency, and municipal debt. This is consistent with flights to the safety and liquidity of these assets reducing the amount of funds available to the shadow banking sector and shifting the funding supply schedule. Overall, the results support our choice of instrument, as well as our interpretation of the age-based liquidity factor as a funding liquidity factor.

4.3 Funding liquidity and the macroeconomy

This section explores the linkages between the value of funding liquidity and broader measures of funding conditions. Figure 2B compares the liquidity factor with the VXO index of aggregate uncertainty.[33] This broad indicator of uncertainty, or fear, is often cited as a measure of liquidity shocks. In particular, note that peaks in volatility are associated with peaks of the liquidity factor and provide some justification for considering broader economic indicators. Nonetheless, the evidence is mitigated by the period around 2001–2003 when monetary policy was particularly accommodative.

33 The VXO is a close analogue to the VIX measure of volatility. The VXO is available for a longer sample period.

Table 6
Economic determinants of funding liquidity

Model	cst	$Reserves$	$\Delta M0$	$\Delta M1$	$\Delta M2$	VXO	BA	R^2
A	0.343	−0.077						17.5%
	(19.73)	(−5.31)						
B	0.343		−0.064	0.019	−0.019			14.0%
	(17.10)		(−1.97)	(0.81)	(−1.05)			
C	0.342					0.05		7.9%
	(18.52)					(2.42)		
D	0.343	−0.121	−0.009	−0.003	−0.098			41.4%
	(23.83)	(−7.31)	(−0.34)	(−0.15)	(−5.15)			
E	0.343	−0.110	−0.014	0.002	−0.097	0.021		42.4%
	(23.83)	(−6.91)	(−0.56)	(0.10)	(−5.22)	(1.40)		
F	0.343	−0.098	0.005	−0.012	−0.072	−0.005	0.056	46.7%
	(25.41)	(−6.01)	(0.20)	(−0.74)	(−3.51)	(−0.27)	(2.78)	

Results from regressions of the funding liquidity factor on selected economic variables. $Reserves$ is the aggregate amount of non-borrowed reserves at the Fed. $\Delta M0$, $\Delta M1$, and $\Delta M2$ are annual changes of the corresponding monetary aggregate measures. VXO is the implied volatility from S&P 500 call options. BA is the difference between the minimum and the median bid-ask spreads in the cross-section of maturity at each date. Regressors are demeaned and divided by their standard deviations. Newey-West *t-statistics* (6 lags) in parentheses. End-of-month data (1985:12–2007:12) except when VXO is included, in which case it is (1986:01–2007:12).

Figure 2B suggests a special role for monetary aggregates. In the following, we consider the relative role of the VXO index, standard monetary aggregates (i.e., $M0$, $M1$, and $M2$), and non-borrowed bank reserves at the Federal Reserve.[34] First consider the relationship between funding liquidity value, bank reserves, and monetary aggregates. The first row of Table 6 (Model A) presents results from a regression of funding liquidity with bank reserve growth. Reserve variations alone explain 17.5% of funding liquidity variations with a negative coefficient. Model B presents results from a regression on the monetary aggregates. Together, they explain 14.0% of funding liquidity variations. Increases in the monetary base are associated with substantially lower values of funding liquidity value. Model C shows that the VXO index has a positive coefficient and explains 7.9% of funding liquidity variations.

Next, Model D combines bank reserves with monetary aggregates. The results in Table 6 provide evidence of strong interactions between these variables since their combined explanatory power is almost one-third higher (41.4% vs. 31.5%) than the sum of their individual R^2s. The impact of higher reserves is magnified. Higher reserves decrease funding liquidity value by two-thirds of a standard deviation. The impact of $M2$ growth is also magnified and has a similar impact but, conditional on reserves, monetary base variations are no longer significant. Interestingly, reserves and $M2$ growth reveal margins of choice located near the funding market. Non-borrowed reserves are a substitute to providing funding liquidity in secured or unsecured overnight money markets, at least before 2008. All else equal, high reserve growth is

34 $M0$ includes physical currencies and accounts at the Fed; $M1$ adds demand accounts; and $M2$ adds most savings accounts, money-market accounts, retail money-market mutual funds, and small-denomination time deposits.

associated with better funding conditions presumably because a larger pool of funds is available. On the other hand, for a given level of bank reserves and monetary base, $M2$ growth is primarily driven by flows into MMMF, which increases the supply of available funds. These results are robust to the inclusion of measures of aggregate uncertainty and transaction costs (i.e., Models E and F, respectively). Interestingly, aggregate uncertainty plays no marginal role when we controlled for monetary aggregates. Finally, unreported results show that our conclusions are robust to the inclusion of latent macro-economic factors derived from a large set of macro-economic and financial indicators (Ludvigson and Ng 2009).

These three sources of evidence, from funding costs, from the shadow banking sector, and from broad measures of funding conditions, together provide strong support for our interpretation of the extracted liquidity factor as a measure of funding liquidity conditions. Brunnermeier and Pedersen (2008) also suggest that the value of funding liquidity induces a common liquidity premium across markets. In the next section, we explore this implication by measuring the impact of our liquidity factor on several fixed-income markets: Treasury securities, LIBOR loans, swap contracts, and corporate bonds.

5. Funding Liquidity Across Fixed-income Markets

5.1 U.S. Treasury bonds

5.1.1 Funding liquidity predicts excess returns. We first consider the market for U.S. Treasury securities where the impact of liquidity on prices is generally considered to be negligible or purely transitory. Figure 4A compares the value of liquidity with annual excess returns on old 2-year Treasury bonds. A negative relationship is visually apparent throughout the sample with peaks at the crash of October 1987, the Mexican peso crisis late in 1994, the LTCM crisis in August 1998, and the end of the millennium. We test this hypothesis formally through predictive regressions of off-the-run bond excess returns on the liquidity factor including term structure factors to control for the information content of forward rates (Fama and Bliss 1987; Campbell and Shiller 1991; Cochrane and Piazzesi 2005). We consider (annualized) excess returns from holding off-the-run bonds with maturities of 2, 3, 4, 5, 7, and 10 years and for investment horizons of 1, 3, 6, 12, and 24 months.[35] We find that the risk premium of all Treasury bonds decreases when the value of funding liquidity increases.[36]

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[35] Here and in the following sections, and unless otherwise stated, we use results from the model to compute excess returns where appropriate. Results are robust to the choice of the off-the-run yield curve used to compute spreads or excess returns. Similarly, for ease of interpretation, we standardize each regressor to zero mean and unit variance.

[36] This accords with Longstaff (2004), who documents price differences between off-the-run Treasury and Refcorp bonds with similar cash flows. These bonds' principals backed with U.S. Treasury bonds and their coupons are explicitly guaranteed by the U.S. Treasury. He argues that discounts on Refcorp bond are due to "... the liquidity of Treasury bonds, especially in unsettled markets." This also accords with theoretical models of the liquidity

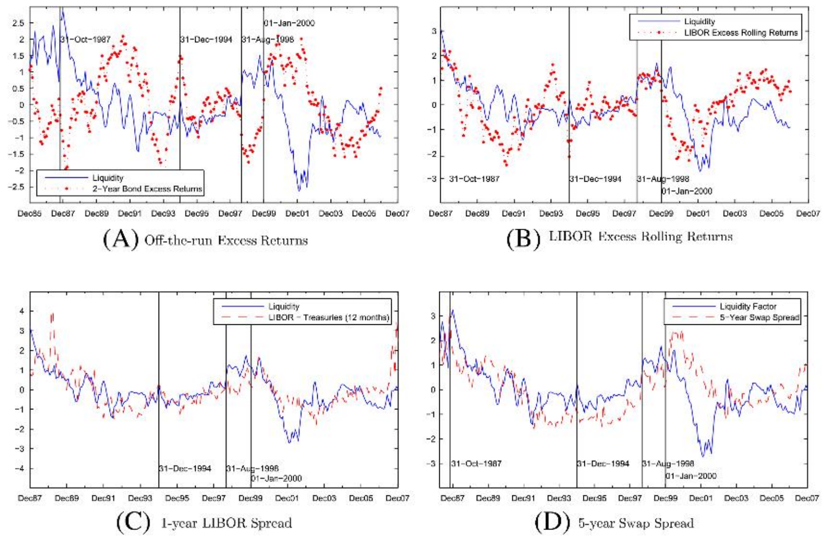

Figure 4
Excess returns and funding liquidity
Funding liquidity and the risk premia across different markets. Panel A displays annual excess returns on 2-year off-the-run U.S. Treasury bonds. Panel B displays annual excess rolling returns on a 3-month LIBOR loan. Panel C displays the spread of the 12-month LIBOR rate above the off-the-run 1-year zero-coupon yield from the model. Panel D displays the spread of the 5-year swap rate of the off-the-run par yield curve from the model. End-of-month data from CRSP (1985:12–2007:12).

Panel A of Table 7 reports average risk premia. These range from 153 bps to 471 bps at a one-month horizon and from 69 bps to 358 bps at an annual horizon. These large excess returns are consistent with an average positive term structure slope and with a period of declining interest rates. Panel B presents estimates of the liquidity coefficients. The results are conclusive. Estimates are negative and significant at all horizons and maturities, and the impact of funding liquidity on future excess returns is economically significant. At a one-month horizon, a one-standard-deviation shock lowers expected excess returns by 139 bps and 542 bps (annualized) for 2-year and 10-year zero-coupon bonds, respectively. At this horizon, R^2 statistics range from 4.7% to 3.6% (see Panel C). Annual excess returns exhibit substantial predictability, with R^2s ranging from 36% to 41%. Of course, these R^2s measure the joint explanatory power of all regressors. Panel C presents, in brackets, the R^2s of regressions excluding the liquidity factor. The value of funding liquidity accounts for more or less half of the total predictive power. Its effect on annual excess returns is also economically significant. A one-standard-deviation shock decreases expected 1-year excess returns by 85 bps and by as much as 408 bps for 2-year and 10-year zero-coupon bonds, respectively.

value of Treasury bonds (e.g., Svensson 1985; Bansal and Coleman 1996; Krishnamurthy and Vissing-Jorgensen 2007).

Table 7
Treasury bond excess returns and funding liquidity

Panel A: Average risk premium

	Bond Maturity											
Horizon	2		3		4		5		7		10	
1	1.53	(7.07)	2.09	(11.17)	2.59	(15.00)	3.03	(18.53)	3.80	(24.86)	4.71	(33.49)
3	1.36	(4.17)	1.90	(6.64)	2.39	(8.89)	2.83	(10.89)	3.57	(14.36)	4.44	(18.90)
6	1.10	(2.67)	1.63	(4.38)	2.10	(5.89)	2.51	(7.22)	3.21	(9.53)	3.99	(12.53)
12	0.69	(1.37)	1.21	(2.59)	1.66	(3.62)	2.07	(4.50)	2.78	(6.03)	3.58	(8.07)
24	0.00	(0.00)	0.61	(0.96)	1.11	(1.67)	1.56	(2.20)	2.34	(2.94)	3.26	(3.78)

Panel B: Liquidity coefficients

	Bond Maturity											
Horizon	2		3		4		5		7		10	
1	−1.39	(−2.49)	−2.27	(−2.53)	−3.01	(−2.47)	−3.61	(−2.39)	−4.52	(−2.27)	−5.42	(−2.07)
3	−1.35	(−3.28)	−2.12	(−3.14)	−2.74	(−2.97)	−3.23	(−2.84)	−3.98	(−2.64)	−4.70	(−2.34)
6	−1.25	(−4.67)	−2.00	(−4.51)	−2.59	(−4.29)	−3.07	(−4.09)	−3.84	(−3.75)	−4.69	(−3.26)
12	−0.85	(−5.47)	−1.63	(−5.63)	−2.24	(−5.63)	−2.73	(−5.57)	−3.44	(−5.18)	−4.08	(−4.15)
24	0.00	(0.00)	−0.53	(−3.24)	−0.91	(−3.23)	−1.17	(−3.27)	−1.51	(−3.29)	−1.75	(−2.91)

(*continued*)

Table 7
Continued

Panel C: R^2

	Bond Maturity											
Horizon	2		3		4		5		7		10	
1	4.74	[2.28]	4.65	[2.02]	4.51	[1.95]	4.34	[1.93]	4.03	[1.92]	3.55	[1.89]
3	13.56	[6.84]	13.33	[6.83]	13.07	[7.03]	12.78	[7.18]	12.07	[7.17]	10.52	[6.57]
6	24.23	[10.34]	24.50	[11.21]	24.57	[12.26]	24.61	[13.11]	24.44	[14.10]	22.92	[14.02]
12	35.36	[11.23]	37.71	[12.66]	39.24	[14.96]	40.32	[17.20]	41.46	[21.00]	40.54	[24.42]
24	0.00	[0.00]	35.53	[16.92]	31.91	[13.91]	29.46	[11.92]	26.56	[10.32]	25.82	[12.69]

Results from the predictive regressions,

$$xr_{t+h}^{(m)} = \alpha_h^{(m)} + \delta_h^{(m)} L_t + \beta_h^{(m)T} F_t + \epsilon_{(t+h)}^{(m)},$$

where L_t is the funding liquidity factor, F_t are term structure factors from the AFENS model, and $xr_{t+h}^{(m)}$ are the excess returns at horizon h (months) on an off-the-run zero-coupon bond with maturity m (years) computed from the model. Regressors are demeaned and divided by their standard deviations. Panel A displays estimates of α. Panel B displays estimates of δ with *t-statistics* based on Newey-West standard errors (h+3 lags) in parentheses. Panel C displays R^2s obtained when including or excluding [in brackets] the funding liquidity factor. End-of-month data from CRSP(1985:12–2007:12).

5.1.2 Robustness. One potential issue is that zero-coupon bond prices, which are used to compute excess returns and forward rates, are computed from the same model as the funding liquidity factor. Model misspecification could yield estimates of term structure factors that do not correctly capture the information content of forward rates or that induce spurious correlations between excess returns and funding liquidity. As a robustness check against both possibilities, we reexamine the predictability regressions but use excess returns and forward rates computed from bootstrapped zero-coupon data available from CRSP.[37] From this alternative dataset, we compute annual excess returns on zero-coupon bonds with maturity from two to five years. As regressors, we include annual forward rates from CRSP at horizon from one to five years, along with the funding liquidity factor from the model. Panel A of Table 8 presents the results. Estimates of the liquidity coefficients are close to our previous results (see Table 7) and are highly significant. We conclude that the predictability power of the liquidity factor is robust to how we compute excess returns and forward rates.

Furthermore, this alternative set of returns allows us to check whether the AFENS model captures important aspects of observed excess returns. Panel B of Table 8 provides results for the regressions of CRSP excess returns on CRSP forward rates, excluding the liquidity factor. This is a replication of the unconstrained regressions in Cochrane and Piazzesi (2005), but for our shorter sample period. This exercise confirms their stylized predictability results in this sample. The predictive power of forward rates is substantial, and we recover a similar, tent-shaped pattern of coefficients across maturities, revealing that a single linear combination of forward rates captures the predictability. Next, Panel C provides results of a similar regression with CRSP forward rates but using excess returns computed from the model. Comparing Panels B and C, we see that average excess returns, forward rate coefficients, as well as R^2s are similar across datasets. The AFENS model captures the stylized facts of bond risk premia, which is an important measure of success for term structure models and provides further support for our results.[38]

Finally, another potential issue is the use of a short-term Treasury yield in the computation of excess returns. The positive liquidity coefficients could be due to a negative correlation of the liquidity factor with variations in short-term yields. But this is not the case. The direct impact of funding liquidity is purged from these yields since we used off-the-run yields computed from the model but shutting off the impact of liquidity. Moreover, we find that the liquidity factor has little explanatory power for the residuals from a projection of short-term rates on the term structure factors.

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37 Using zero-coupon data based on Gurkaynak, Sack, and Wright (2006) and available from the Board of Governors leads to similar results.

38 Fama (1984) originally identified this modeling challenge, but see also Dai and Singleton (2002). Conversely, our results show that the empirical results of Cochrane and Piazzesi (2005) are not an artifact of the bootstrap method. See the discussion in Dai, Singleton, and Yang (2004).

Table 8
Treasury excess returns and funding liquidity—Alternate dataset

Panel A: Excess returns and forward rates from Fama-Bliss data with the liquidity factor

Maturity	cst	$f_t^{(1)}$	$f_t^{(2)}$	$f_t^{(3)}$	$f_t^{(4)}$	$f_t^{(5)}$	L_t	R^2
2	0.72	0.29	−1.31	1.88	0.93	−0.95	−0.78	41.65
	(3.49)	(0.49)	(−1.18)	(1.50)	(1.04)	(−1.60)	(−5.97)	
3	1.31	0.15	−2.26	4.32	0.76	−1.49	−1.55	41.66
	(3.41)	(0.14)	(−1.13)	(1.89)	(0.48)	(−1.27)	(−5.93)	
4	1.79	−0.51	−1.74	4.58	1.53	−1.85	−2.18	42.82
	(3.53)	(−0.35)	(−0.66)	(1.51)	(0.75)	(−1.13)	(−6.07)	
5	1.98	−1.51	−0.24	4.57	0.36	−0.81	−2.66	40.87
	(3.23)	(−0.84)	(−0.07)	(1.24)	(0.15)	(−0.39)	(−5.83)	

Panel B: Excess returns and forward rates from Fama-Bliss data

Maturity	cst	$f_t^{(1)}$	$f_t^{(2)}$	$f_t^{(3)}$	$f_t^{(4)}$	$f_t^{(5)}$	L_t	R^2
2	0.72	−0.43	−1.34	2.66	0.99	−1.53		21.04
	(2.95)	(−0.57)	(−1.06)	(1.50)	(0.95)	(−2.13)		
3	1.31	−1.27	−2.33	5.86	0.88	−2.64		19.29
	(2.87)	(−0.87)	(−1.04)	(1.77)	(0.46)	(−1.86)		
4	1.79	−2.52	−1.83	6.74	1.70	−3.46		19.86
	(2.95)	(−1.26)	(−0.62)	(1.51)	(0.67)	(−1.76)		
5	1.98	−3.96	−0.35	7.20	0.56	−2.79		18.27
	(2.71)	(−1.65)	(−0.10)	(1.35)	(0.19)	(−1.14)		

Panel C: Excess returns from the model and forward rates from Fama-Bliss data

Maturity	cst	$f_t^{(1)}$	$f_t^{(2)}$	$f_t^{(3)}$	$f_t^{(4)}$	$f_t^{(5)}$	L_t	R^2
2	0.66	−0.13	−1.91	2.97	0.93	−1.51		21.10
	(2.71)	(−0.17)	(−1.53)	(1.69)	(0.91)	(−2.09)		
3	1.27	−1.15	−2.04	4.97	1.19	−2.43		18.19
	(2.82)	(−0.79)	(−0.90)	(1.50)	(0.63)	(−1.73)		
4	1.74	−2.46	−1.26	6.09	1.18	−2.92		17.22
	(2.83)	(−1.24)	(−0.41)	(1.34)	(0.46)	(−1.47)		
5	2.09	−3.86	0.00	6.62	1.06	−3.12		17.15
	(2.80)	(−1.61)	(0.00)	(1.20)	(0.34)	(−1.26)		

Results from the regressions,

$$xr_{t+12}^{(m)} = \alpha^{(m)} + \delta^{(m)} L_t + \beta^{(m)T} f_t + \epsilon_{(t+12)}^{(m)},$$

where $xr_{t+h}^{(m)}$ are the annual excess returns on an off-the-run zero-coupon bond with maturity m (years) computed from the model, f_t is a vector of annual forward rates $f_t^{(h)}$ from one to five years, and L_t is the liquidity factor. Regressors are demeaned and divided by their standard deviations. Panel A presents results using returns and forward rates obtained directly from CRSP data but with the liquidity factor from the model. Panel B excludes the liquidity factor. Panel C excludes the funding liquidity factor and uses excess returns from the model. Newey-West *t-statistics* (in parentheses) with 15 lags. End-of-month data from CRSP (1985:12–2007:12).

5.2 LIBOR loans

This section considers returns obtained from rolling over a short-term lending position at the LIBOR rate and funding this position at the risk-free rate. In contrast with the government bond market, we find that higher valuation of funding liquidity predicts higher excess returns. Figure 4B highlights the positive correlation between liquidity and rolling excess returns. Note the spikes in 1987, 1994, 1998, and around the end of the millennium. Figure 5

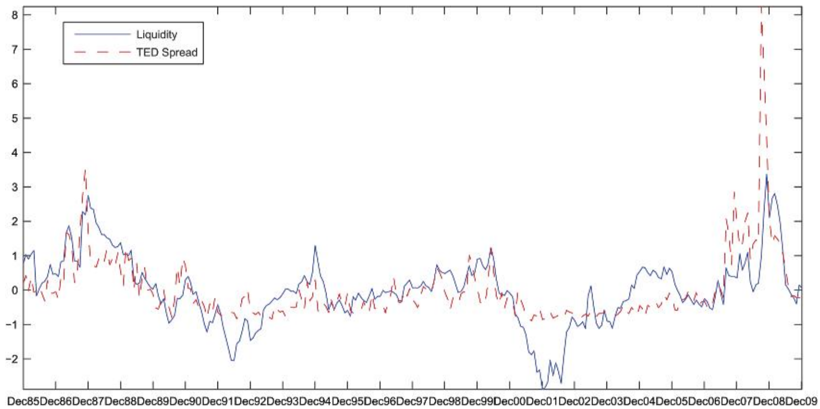

Figure 5
Funding liquidity and the 3-month TED spread
The value of funding liquidity (left scale) and the spread between the 3-month LIBOR rate and the 3-month Treasury Bill yield (right scale). End-of-month data (1985:12–2009:12).

compares the spread between the 3-month LIBOR rate and the 3-month T-bills yield (TED spread) with the value of funding liquidity. The strong positive relationship is visually clear. The sample correlation is 0.63.

This does not preclude that part of the LIBOR spread is due to the higher default risk of the average issuer relative to that of the U.S. government. Indeed, the TED spread is often interpreted as a combination of the credit risk of participating banks, which affects the LIBOR curve, and the refuge value of Treasury securities, which affects the T-bills curve. Therefore, a traditional risk-based explanation calls for the synchronous manifestation of different risks, regularly over time. In contrast, an explanation based on intermediation frictions appeals to only one channel. It also explains and justifies the growing importance attributed to the TED spread as an indicator of systemic risk. The reward for providing liquidity in the inter-bank market is higher when the value of funding liquidity increases. Interbank loans are poor substitutes to U.S. Treasury securities in times of funding stress. Hence, part of the spreads of LIBOR rates in excess of Treasury yields reflect the opportunity costs, in terms of future liquidity, of an interbank loan compared to the liquidity of a Treasury bond. Indeed, in order to convert a loan back to cash, a bank must enter into a new bilateral contract to borrow money. The search costs of this transaction depend on the number of willing counterparties in the market. In particular, the spread also reflects the risk that it may be difficult at critical times, when funding conditions are tight, to convert a LIBOR position back to cash. A Treasury bond can be converted into cash on the repo or the secondary markets.

We test this hypothesis formally through predictive regressions of excess returns. We compute returns from rolling short-term LIBOR loans with 1, 3, 6, or 12 months to maturity. Each investment strategy consists of rolling over one of the short-maturity loans as many times as required to reach the end of the (longer) investment horizon. We consider investment horizons between 1, 3, 6, 12, and 24 months from January 1987 to December 2007. We then compute excess returns of these rolling strategies over the zero-coupon corresponding to the investment horizons computed from the model. Term structure factors are included in the regressions to control for the information content of the term structure. Results are reported in Table 9. For each loan maturity, the average excess return is around 25 bps for the shortest possible rolling horizon. Returns then decrease with the horizon and become negative at the longest horizons. This reflects the average positive slope of the term structure. Funding rolling short-term investments at a fixed rate incurs a negative carry and does not produce positive returns, on average.

The impact of liquidity is unambiguously positive for all horizons and maturities with most *t-statistics* above 5. Interestingly, its impact increases with the horizon. A one-standard-deviation shock to the value of liquidity increases returns on a rolling investment in one-month LIBOR loans by 16 bps and 73 bps at horizons of 3 and 24 months, respectively. Results are similar for other maturities. In fact, the impact is sufficiently large to compensate for the negative carry and reflects the persistence of the liquidity premium. The R^2s from these regressions range from 30% to 50%, and the marginal contribution of funding liquidity is substantial. In the case of annual excess rolling returns from 3-month loans, the predictive power increases from 10.8% to 43.2% when we include the funding liquidity factor.

As an alternative test, we use LIBOR spreads on loans with maturities of 1, 3, 6, and 12 months as ex ante measures of risk premium. Figure 4C shows the positive relationship between liquidity and the 12-month LIBOR spread. Panel A of Table 10 presents the regression results. A one-standard-deviation shock to liquidity is associated with concurrent increases of 18, 15, 11, and 8 bps for loans with maturity of 1, 3, 6, and 12 months, respectively.

5.3 Interest rate swaps

The impact of funding liquidity extends to the swap market. We use a sample of U.S. dollar interest-rate swap rates from April 1987 to December 2007, available in Datastream. We focus on maturities of 2, 5, 7, and 10 years and compute their spreads above the yield to maturity of the corresponding off-the-run par yield. Figure 4D compares the funding liquidity factor with the 5-year swap spread. The positive relationship is apparent. Panel B of Table 10 reports results from regressions of swap spreads on funding liquidity and the term structure factors as conditioning information. First, the average spread rises with maturity, from 38 bps to 43 bps, and extends the pattern of LIBOR risk premia. Next, estimates of the liquidity coefficients imply that, controlling for term structure factors, a one-standard-deviation shock to liquidity raises swap

Table 9
LIBOR rolling excess returns and funding liquidity

Panel A: Average excess returns

	Loan Maturity							
Horizon	1		3		6		12	
1	0.277	(0.347)	0.000	(0.000)	0.000	(0.000)	0.000	(0.000)
3	0.183	(0.248)	0.265	(0.245)	0.000	(0.000)	0.000	(0.000)
6	0.062	(0.322)	0.144	(0.264)	0.239	(0.165)	0.000	(0.000)
12	−0.153	(0.615)	−0.070	(0.560)	0.029	(0.439)	0.253	(0.151)
24	−0.537	(1.120)	−0.453	(1.079)	−0.351	(0.985)	−0.120	(0.743)

Panel B: Liquidity coefficients

	Loan Maturity							
Horizon	1		3		6		12	
1	0.184	(7.837)	0.000	(0.000)	0.000	(0.000)	0.000	(0.000)
3	0.162	(7.853)	0.149	(6.364)	0.000	(0.000)	0.000	(0.000)
6	0.193	(6.139)	0.173	(6.985)	0.101	(5.699)	0.000	(0.000)
12	0.360	(5.700)	0.340	(6.364)	0.277	(7.329)	0.076	(3.695)
24	0.732	(5.578)	0.715	(5.909)	0.664	(6.395)	0.526	(7.366)

Panel C: R^2

	Loan Maturity							
Horizon	1		3		6		12	
1	46.4	[28.0]	0.0	[0.0]	0.0	[0.0]	0.0	[0.0]
3	44.7	[16.8]	50.6	[26.5]	0.0	[0.0]	0.0	[0.0]
6	24.7	[1.4]	30.7	[2.9]	44.8	[20.4]	0.0	[0.0]
12	29.2	[7.1]	30.3	[6.6]	32.3	[6.7]	35.2	[18.6]
24	38.8	[12.3]	38.9	[11.7]	39.4	[11.2]	41.2	[10.1]

Results from the regressions,

$$xr_{t+h}^{(m)} = \alpha_h^{(m)} + \delta_h^{(m)} L_t + \beta_h^{(m)T} F_t + \epsilon_{(t+h)}^{(m)},$$

where $xr_{t+h}^{(m)}$ are the returns on rolling investments in a loan with maturity m (months) over a horizon $h \geq m$ in excess of the zero-coupon rate for that horizon computed from the model, L_t is the funding liquidity factor, and F_t is the vector of term structure factors. Regressors are demeaned and divided by their standard deviations. Panel A contains estimates of average returns. Panel B contains estimates of $\delta_h^{(m)}$ and Newey-West *t-statistics* (h+3 lags) in parentheses. Panel C presents R^2 from the regressions including and excluding [in brackets] the funding liquidity factor. End-of-month data from CRSP (1985:12–2007:12).

spreads from 9 bps to 11 bps across maturities. The estimates are significant, both statistically and economically, given the higher price sensitivities of swaps to changes in yields. For a 5-year swap with a duration of 4.5, say, the price impact of a 6 bps change is $0.27 for a notional of $100. This translates into substantial returns given the leveraged nature of swap positions.[39] Finally, the explanatory power of liquidity is high and increases with maturity.

[39] We do not use returns on swap investment to measure expected returns. Swap investment requires zero initial investment. Determining the proper capital-at-risk to use in returns computation is somewhat arbitrary.

Table 10
LIBOR spreads, swap spreads, and funding liquidity

Panel A: LIBOR spreads

	Maturity							
	1		3		6		12	
Avg Spread	0.423	(0.027)	0.422	(0.023)	0.406	(0.019)	0.429	(0.019)
$\delta_m^{(h)}$	0.183	(6.463)	0.153	(5.939)	0.106	(5.166)	0.080	(4.410)
R^2	58.4	[44.9]	59.4	[47.8]	53.2	[42.2]	53.9	[37.7]

Panel B: Swap spreads

	Maturity							
	24		60		84		120	
Avg. Spread	0.384	(0.016)	0.483	(0.018)	0.477	(0.019)	0.432	(0.020)
$\delta_m^{(h)}$	0.094	(4.556)	0.104	(4.525)	0.107	(4.395)	0.095	(3.917)
R^2	37.8	[35.4]	38.0	[34.2]	45.5	[38.6]	51.7	[38.5]

Results from the regressions,

$$sprd_t^{(m)} = \alpha^{(m)} + \delta^{(m)} L_t + \beta^{(m)T} F_t + \epsilon_{(t)}^{(m)},$$

where $sprd_t^{(m)}$ is the spread at time t and for maturity m (months), L_t is the liquidity factor, and F_t is the vector of term structure factors. Spreads in excess of the zero-coupon curve (LIBOR spreads) and to the par yield curve (swap spreads) computed from the model. Regressors are demeaned and divided by their standard deviations. Panel A presents results for LIBOR spreads. Panel B presents results for swap spreads. Newey-West *t-statistics* (3 lags) are in parentheses. R^2 are from regressions including and excluding [in brackets] the funding liquidity factor. End-of-month data from CRSP (1985:12–2007:12).

Interestingly, funding liquidity affects swap spreads and LIBOR spreads similarly. This suggests that compensation for funding liquidity risk arising from the exposure to future LIBOR rates is the main driver behind the liquidity component of swap spreads. Moreover, variations in funding liquidity may affect the swap market directly since the same intermediaries operate in the Treasury and the swap markets. This is consistent and supports previous literature (Duffie and Singleton 1997; Grinblatt 2001; Liu, Longstaff, and Mandell 2006; Fedlhütter and Lando 2008) pointing toward the convenience yield from holding Treasury bonds as the most important driver of swap spreads.[40] However, we also show that the effect of funding liquidity reaches across markets, not only by pushing the Treasury yield down, as the convenience yield increases, but also by pushing the LIBOR and swap curves up.

5.4 Corporate bonds

This section examines the impact of liquidity on corporate bonds. We first consider end-of-month data from December 1988 to December 2007 on five

[40] Counterparty risk arising from swap contracts is mitigated by netting agreements and posting of collateral. In practice, its effect on swap spread is small (e.g., Duffie and Huang 1996).

Merrill Lynch indices with credit ratings of AAA, AA, A, BBB, and High Yield (HY) ratings (i.e., HY Master II index), respectively. For each index, and each month, we compute returns in excess of the off-the-run zero coupon yield for investment horizons of 1, 3, 6, 12, and 24 months.

Results from predictive regressions of excess returns on the liquidity and term structure factors are reported in Panel A of Table 11. As expected, average excess returns are higher for lower credit ratings. Next, estimates of the liquidity coefficients are negative for the AAA, AA, and A credit ratings and become positive for BBB and HY credit ratings. A one-standard-deviation increase of funding liquidity decreases expected returns by 1.78% for the AAA index but increases expected returns by 3.12% for the HY index. For comparison, a similar shock would decrease expected returns on Treasury bonds with 7 and 10 years to maturity by 4.52% and 5.42%, respectively. Thus, on average, high-quality bonds were considered substitutes, albeit imperfect, to U.S. Treasuries as a hedge against variations in funding conditions.

We consider an alternative dataset, based on individual bond transaction data from the NAIC.[41] This sample covers a shorter period, from February 1996 until December 2001, but provides actual transaction data and a better coverage of the rating spectrum. Once restricted to end-of-month observations, the sample includes 2,171 transactions over 71 months. To preserve parsimony, we group ratings into five categories.[42] We consider regressions of NAIC corporate spreads on the liquidity and term structure factors, but we also include the control variables used by Ericsson and Renault (2006). These are the VIX index, the returns on the S&P 500 index, the level and slope of the term structure of interest rates, the difference between Moody's Baa and Aaa-rated bond yield indices, and an on-the-run dummy signaling whether that particular bond was on-the-run at the time of the transaction. The panel regressions of credit spreads for bond i at date t are given by

$$sprd_{i,t} = \alpha + \beta_1 L_t I(G_i = 1) + \cdots + \beta_5 L_t I(G_i = 5) + \gamma_h^T X_t + \epsilon_{i,t}, \qquad (10)$$

where L_t is the liquidity factor and $I(G_i = j)$ is an indicator function equal to one if the credit rating of bond i belongs in group $j = 1, \ldots, 5$. Control variables are grouped in the vector X_{t+h}.

A flight-to-quality pattern emerges clearly in Panel B of Table 11. An increase in funding liquidity value of one standard deviation decreases spreads by 31 bps and 20 bps in groups 1 and 2, respectively. The effect is smaller and statistically undistinguishable from zero for group 3 but becomes positive for

41 We thank Jan Ericsson for providing the NAIC transaction data and control variables. See Ericsson and Renault (2006) for a discussion of this dataset.

42 Group 1 includes ratings from AAA to A+, group 2 from A to A-, group 3 includes BBB+, BBB, and BBB-, group 4 includes CCC+, CCC, and CCC-, and group 5 includes the remaining ratings down to C-.

Table 11
Corporate bond excess returns and funding liquidity

Panel A: Merrill lynch indices excess returns

	Rating									
	AAA		AA		A		BBB		HY	
Avg.	3.162	(15.502)	3.130	(15.291)	3.162	(15.618)	3.204	(16.196)	3.785	(23.400)
$\delta_h^{(G)}$	−1.775	(−1.396)	−1.626	(−1.341)	−1.154	(−0.913)	0.073	(0.057)	3.117	(1.461)
R^2	4.5	[3.7]	4.9	[4.2]	4.4	[4.1]	3.4	[3.4]	6.3	[5.2]

Panel B: NAIC corporate spreads

	Rating Group									
	G1		G2		G3		G4		G5	
Avg.	1.51	(0.19)	1.65	(0.21)	2.25	(0.30)	3.38	(0.59)	3.70	(0.54)
$\delta_m^{(G)}$	−0.31	(−2.98)	−0.20	(−1.96)	−0.04	(−0.34)	0.25	(2.29)	0.26	(2.47)
R^2	3.9	[2.0]	5.7	[2.0]	6.5	[2.0]	7.0	[2.0]	7.5	[2.0]

(*continued*)

Table 11
Continued

Panel C: Merrill lynch spread indices

	Rating									
	Aaa		Aa		A		Baa		HY	
Avg.	0.930	(0.036)	0.976	(0.049)	1.227	(0.046)	1.856	(0.077)	5.385	(0.270)
$\delta_m^{(h)}$	0.065	(2.294)	0.060	(1.188)	0.073	(1.268)	0.119	(1.379)	0.334	(1.168)
R^2	59.5	[55.5]	31.4	[29.6]	39.6	[34.9]	49.7	[42.7]	39.2	[29.9]

Results from the regressions,

$$y_t = \alpha_h^{(r)} + \delta_h^{(r)} L_t + \beta_h^{(r)T} F_t + \epsilon_{(t+h)}^{(r)},$$

where y_t is either a spread, $sprd_t^r$, observed at time t for rating r or an excess return, $xr_{t+h}^{(r)}$ over a horizon h (months) on an investment in the corporate index with rating r, L_t is the funding liquidity factor, and F_t is the vector of term structure factors. See Equation (10) for the panel specification in the case of spreads. Panel A presents results for excess returns. Panel B presents results for corporate spreads. Corporate bond yields are obtained at the security level from NAIC. Corporate bond returns are computed using Merrill Lynch indices obtained from Bloomberg. Spreads and excess returns are computed relative to off-the-run zero-coupon rates obtained from the model. Regressor are demeaned and divided by their standard deviations. Newey-West *t-statistics* in parentheses and R^2 are from regressions including and excluding [in brackets] the funding liquidity factor. Results from Merrill Lynch indices cover the entire sample [1985:12–2007:12] but not results from NAIC corporate bond yields [1996:02–2001:12].

groups 4 and 5, implying increases of 25 bps and 26 bps, respectively. This is an average effect through time and across ratings within each group.[43]

The pattern of liquidity coefficients obtained from excess returns computed from Merrill Lynch indices and spreads computed from NAIC transactions differ. While results from Merrill Lynch were statistically inconclusive, estimates of liquidity coefficients obtained from NAIC data confirm that a shock to funding liquidity leads to lower corporate spreads in the highest rating groups but higher corporate spreads in the lowest rating groups. Three important differences between samples may explain the results. First, the composition of the index is different from the composition of NAIC transaction data. The impact of liquidity on corporate spreads may not be homogeneous across issues. For example, the maturity or the age of a bond, the industry of the issuer, and security-specific option features may introduce heterogeneity. Second, the Merrill Lynch indices cover a much longer time span and the pattern of liquidity premia across the quality spectrum may be time-varying, a topic we return to in Section 6. Third, ex post excess returns are a noisier proxy of risk premium than ex ante spreads.

Our results are consistent with Collin-Dufresne, Goldstein, and Spencer Martin (2001), who find that most of the variations of the large non-default component in corporate spreads are driven by a single latent factor. We formally link this factor with funding liquidity risk. It is also consistent with the differential impact of liquidity across ratings in Ericsson and Renault (2006). However, we link the flight-to-quality pattern to a systemic factor in the liquidity risk premium instead of bond-specific measures of liquidity.

5.5 Discussion

Funding liquidity is an aggregate risk factor. When the value of the most easily funded collateral rises relative to other securities, we observe variations in risk premia for off-the-run U.S. government bonds, eurodollar loans, swap contracts, and corporate bonds. Empirically, the funding liquidity premium appears strongly during crises, and the pattern is suggestive of the well-known flight-to-quality behavior, but its effect is pervasive even in normal times.

Using observed bond prices instead of a smoothed zero-coupon curve highlights important cross-sectional variations along the yield curve. Also, focusing on the common component of liquidity premia across maturities filters out local or idiosyncratic demand and supply effects. Duffee (2011) argues that some term structure factors may be hard to measure due to distortions or measurement issues on the bond markets. In particular, he suggests that small, transitory, and idiosyncratic deviations may be important. In contrast, we show

[43] We do not report other coefficients. Briefly, the coefficient on the level factor is negative and significant. All other coefficients are insignificant, but these results are not directly comparable with Ericsson and Renault (2006) due to differences of models [we do not include individual bond fixed effects as our sample is small relative to the number (998) of securities] and sample frequencies.

the importance of small but *persistent* and *common* deviations relative to an idealized yield curve. This factor is "hidden" from low-dimensional frictionless no-arbitrage models where estimation is based on zero-coupon curves and where pricing errors are assumed to be transitory and uncorrelated. We show how to measure this factor using the cross-section of Treasury bonds, notes, and bills.

The assets we consider have payoffs that are unrelated to the liquidity of Treasury securities. Therefore, the evidence is hard to reconcile with theories based on default, inflation, or interest rate risk. On the other hand, any attempt to arbitrage these yield differences involves financial exposures (i.e., borrowing and shorting a liquid asset while holding and funding an illiquid asset) that become riskier and costlier in tight funding conditions.[44] These funding risks not only reveal the liquidity premium but also inhibit investors' ability to profit from predictable returns and explain the robustness of long-horizon regressions. Hence, the evidence supports theories emphasizing the role of financial intermediaries (Kyle and Xiong 2001; Gromb and Vayanos 2002; He and Krishnamurthy 2008; Vayanos and Vila 2009) and, in particular, those highlighting the role of funding markets and collateral margin for financial intermediation (Brunnermeier and Pedersen 2008; Gârleanu and Pedersen 2011). In this context, beyond differences between their cash flows, different securities serve, in part, and to varying degrees, to fulfill uncertain future needs for cash.

6. The 2007–2009 Financial Crisis

We have limited our investigation of funding liquidity to the 1985–2007 period to avoid the inclusion of the major financial crisis that affected the economy over the years 2007–2009. Our objective was to show that the value of funding liquidity varies also in normal times and that its effect is pervasive. In this section, we show that the extracted funding liquidity factor would have provided a good measure of the extreme funding tensions that arose during the crisis by associating peaks in the value of funding liquidity with major events during this period, as well as with major interventions of the Federal Reserve. In terms of returns predictability, most results are magnified when we add data up to the end of 2009.[45] We show that during the 2007–2009 period, the relationship between funding liquidity and compensation for risk changed considerably in several key markets.

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[44] Only the U.S. Treasury can issue new securities and take advantage of this price differential. "In addition, although it is not a primary reason for conducting buy-backs, we may be able to reduce the government's interest expense by purchasing older, 'off-the-run' debt and replacing it with lower-yield 'on-the-run' debt." (Treasury Assistant Secretary for Financial Markets Lewis A. Sachs, Testimony before the House Committee on Ways and Means.)

[45] Results available from the authors.

6.1 Federal Reserve interventions

Figure 6A shows the evolution of the funding liquidity factor together with the Fed funds rate, while Figure 6B shows the evolution of Federal Reserve system assets, subdivided into different categories, from December 2006 until December 2009.[46] The components of total assets are (i) Treasury securities; (ii) agency bonds and mortgage-backed securities; and (iii) facilities and other assets.[47] This provides key measures of the Fed's responses as the crisis unfolded. It shows that while ample bank reserves or monetary aggregates were associated to a low value of funding liquidity in the main sample (see Section 4.3), this relationship is reversed during the crisis due to the endogenous response of the Fed.

At the end of 2009, the value of the funding liquidity factor was about the same as at the beginning of 2007. However, it went through a series of steep increases over the first two years as several credit events agitated the financial markets, to culminate in November 2008. The two peaks in 2007 can be associated with Bear Stearns' losses (1.18 at the end of May) and BNP Paribas's freezing of some of its investments (1.41 at the end of August). The two increases at the beginning of 2008 indicate that funding conditions remained tense. Eventually, the failure of Lehman Brothers in September, the rescue of AIG, and the difficulties of the Reserve Primary Fund led the way into the peak of the crisis. The value of funding liquidity value shot up in September 2008 and eventually peaked at 3.08 in November, breaching every level seen in the previous episodes of financial stress.

From August 2007, the Federal Reserve Board approved several cuts in the primary credit rate from 5.75% to 2% by the end of April 2009 to relieve the deepening of the housing contraction and considerable stress in financial markets. Concurrently, the Fed crept up its supply of funding liquidity to the market via a panoply of new facilities, some of them targeted at different segments of the shadow banking sector.[48] The credit extended by the Fed stabilized in November, standing at $1.4 trillion. Conditions remained tense, as indicated by a funding liquidity value rising from 2.31 in December to 2.74 in February. In March 2009, the Fed accelerated its purchases to increase its balanced sheet around levels similar to December 2008, where it stabilized. Funding conditions improved over the following months. Clearly, direct interventions in the capital of some systemic institutions helped ease funding conditions. Nonetheless, the unfolding of the crisis put funding markets, and

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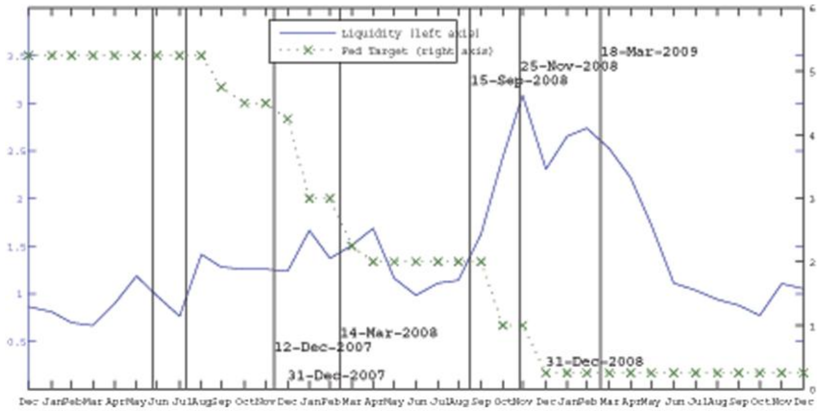
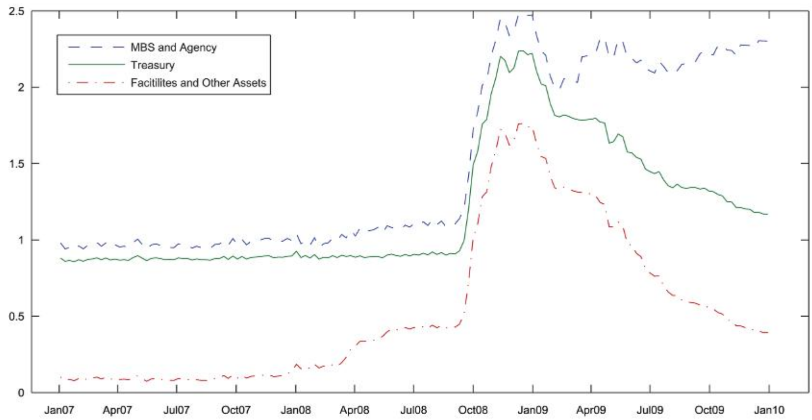
[46] We reestimate the model with data up to December 2009. Keeping parameters at their values estimated with data as of December 2007 yields filtered series with very similar dynamics and a correlation of 0.85.

[47] This component comprises central bank liquidity swaps and the standard primary, secondary, and seasonal credit facilities. It also includes the new facilities created as the crisis unfolded.

[48] The Term Auction Facility, the swap agreements with the ECB, the SNB, and other central banks, the Term Securities Lending Facility, the Asset Backed Commercial Paper Money Market Mutual Fund Liquidity Facility, the Commercial Paper Funding Facility, and the Money Market Investor Funding Facility.



(A) Liquidity Factor



(B) Federal Reserve Assets

Figure 6
Funding liquidity and the federal reserve balance sheet—Updated parameters
Panel A displays the value of funding liquidity using data up to 12:2009 and the Federal Reserve target rate between December 2007 and December 2009. The following dates are highlighted: Bear Stearns reorganizes MBS hedge funds (June 23, 2007), FMOC responds to BNP-Paribas to funds freeze (August 10, 2007), the Fed initiate TAF and swap agreements (December 12, 2007), J. P. Morgan acquires Bear Stearns (March 14, 2008), Lehman Brothers fails (September 15, 2008), the Fed announces purchases of agency bonds (November 25, 2008), the Fed increases agency bond purchases and announces Treasury bond purchases (March 18, 2009). Panel B displays components of the Federal Reserve system assets, in trillions, weekly, between December 2007 and December 2009. See text for components of facilities and other assets category.

funding conditions, at the center stage of the amplification and propagation of financial shocks.

6.2 Time-varying funding risk

Figure 7 displays the evolution of 10-year and 30-year swap spreads during the financial crisis. The negative correlation with the value of funding liquidity

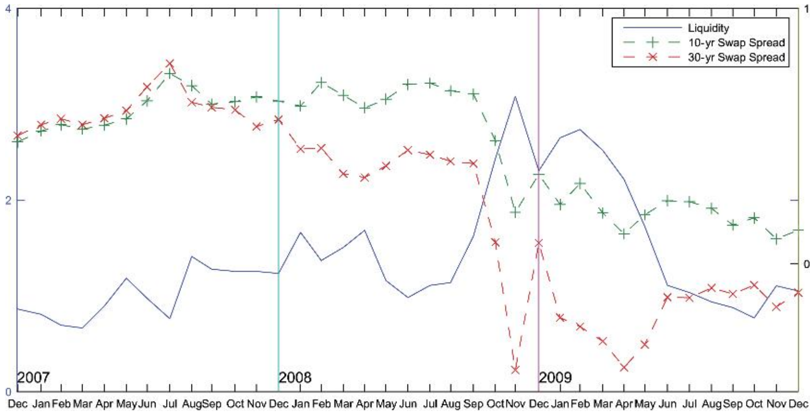


Figure 7
Funding liquidity and swap spreads
The value of funding liquidity (left scale) with the 10-year and 30-year swap spreads, respectively, relative to the corresponding off-the-run Treasury par yield from the model (right scale). End-of-month data (2006:12–2009:12).

is striking and contrasts with the positive correlation found in the rest of the sample. Moreover, the 30-year rate breached below the U.S. Treasury curve in November 2008. Presumably, swap positions designed to hedge against interest-rate and pre-payment risks had to be unwound following the announcements of large-scale purchases of GSE and MBS securities. In any case, the quasi-arbitrage strategy that it opened[49] would require access to unsecured funding for as long as swap spreads did not revert to a level consistent with expectations of future floating rates, as well as capital to meet margin requirements. Hence, the rising value of funding liquidity signals the funding risk faced by would-be arbitrageurs.

Swaps are unfunded assets. U.S. agency bonds are funded securities and provide another interesting case where the relation with funding liquidity risk varied through time. Figure 8A displays the funding liquidity factor against annual excess returns on an index of U.S. agency bonds with 10 years to maturity. In the first half of the sample, up until 1998, investors saw them as substitutes to Treasury securities and required a lower risk premium when the value of funding liquidity increased. In contrast, perhaps with the hindsight of the 1998 financial crisis, it was no longer the case in the second half of the sample. The liquidity risk premium of agency bonds rose when funding liquidity became more valuable. This relationship seems to break again in fall 2008 when the Federal Reserve announced outright purchases of agency bonds and the agencies were placed in federal conservatorship, clarifying the risk of an eventual default.

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[49] One could benefit from entering the fixed side of a swap, buying a Treasury bond that matches the fixed payments, and financing this purchase in the unsecured markets using the swap floating payments.

Variations in the exposure to funding liquidity may help explain the weak statistical evidence in the case of corporate bonds. The last two Panels of Figure 8 compares the liquidity factor with the spread of Merrill Lynch indices with different credit ratings. In the sample excluding 2008–2009, the estimated average impact of a shock to funding liquidity was negative for AAA bonds and positive for BBB. The large and positively correlated shock in the crisis reverses this conclusion for AAA bonds. AAA spreads and funding liquidity value were also positively correlated in 1998. Overall, the evidence from swap spreads, agency bonds, and corporate bonds suggests that funding risk is not stable and depends on the nature and size of shocks to funding liquidity, even in funded markets. Nonetheless, it does not affect our conclusion that funding liquidity risk permeates across bond and money market instruments.

7. Conclusion

We augment the AFENS term structure model of Christensen, Diebold, and Rudebusch (2011) by allowing for a liquidity factor. We identify the common liquidity factor from a panel of Treasury bond pairs at different maturity points where elements of each pair differ by their age. Estimation of the model proceeds directly from coupon-bond prices using a nonlinear filter. The liquidity factor is found to affect all bonds, and its effect increases with maturity but decreases with age. We interpret this factor as a measure of funding liquidity based on its relationship with funding costs, funding supply in the shadow banking sector, and with broad measures of funds availability based on monetary aggregates. We find that funding liquidity predicts a substantial share of the risk premium of Treasury bonds. It also predicts LIBOR spreads, swap spreads, agency spreads, and corporate bond spreads. The pattern across interest rate markets and credit ratings is consistent with accounts of flight-to-liquidity events, but the effect is pervasive in normal times. The evidence points toward the importance of the funding market for the intermediation mechanism and, hence, for asset pricing, even for government bonds.

Our results raise important research questions. What is the impact of the Fed on the liquidity premium across markets? This follows since the Fed affects funding liquidity either indirectly through interventions related to monetary policy or directly when providing support to the financial system. Does the impact of funding liquidity risk extend to markets for equities or foreign exchange? In other words, are intermediaries or levered speculators involved in these other markets exposed to funding liquidity risk? Why are some assets at different times perceived as hedges against funding liquidity risk? There are no truly risk-free assets for the purposes of meeting uncertain liquidity needs other than cash and, perhaps, very short maturity Treasury bills. Nonetheless, investors still succeed to coordinate around long-duration bonds in equilibrium and to make them substitutes for money. In this context, the real-time measure

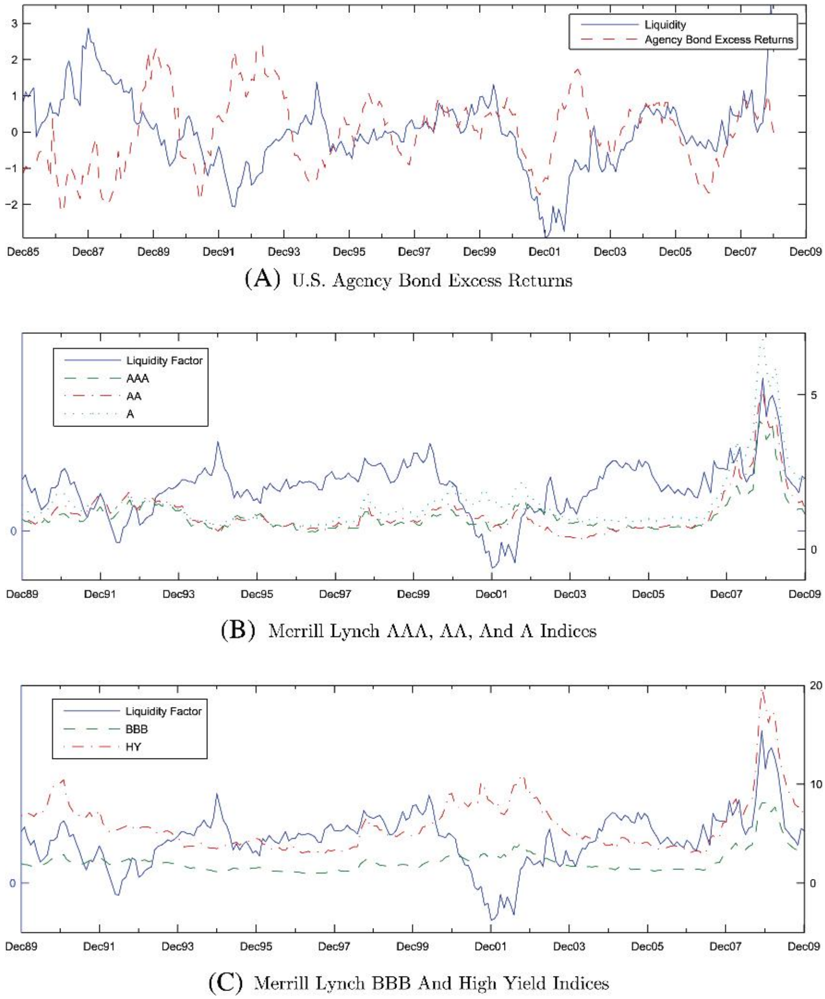

Figure 8
Funding liquidity, agency bond, and Corporate Bonds, including 2007–2009
Panel A compares the value of funding liquidity with the annual excess returns on an index of U.S. agency bonds with 10 years to maturity (1985:12–2009:12), Panel B compares with the spreads of Merrill Lynch indices of AAA, AA, and A ratings (1989:12–2009:12), and Panel C compares with the spread of BBB and high-yield bond indices (1989:12–2009:12). Spreads and excess returns are computed with respect to off-the-run 10-year Treasury par yield from the model.

proposed here is useful to gauge an asset's exposure to funding liquidity risk. We leave these questions for future research.

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