MBS Liquidity: Drivers and Risk Premiums[1]

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Abstract

This paper examines determinants of liquidity in the agency MBS market, measured by dollar-roll implied financing rates (IFRs). The IFRs can be viewed as measures of collateral scarcity and funding liquidity in the MBS market. A decline (rise) in the IFR can be interpreted as a signal that the MBS collateral is more scarce (readily available) and can be funded at a lower (higher) cost. Our results suggest that factors representing higher net supply are generally associated with lower implied financing rates. In addition, we find that liquidity risk is compensated in the cross-section of expected returns—agency MBS that have a higher beta with respect to liquidity shocks on average deliver higher returns—and that liquidity risk premiums are separate from the premiums associated with prepayment and agency credit risks.

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1. Introduction

With theoretical advances on links among funding liquidity, market liquidity, and the behavior of asset prices, a growing empirical literature has been trying to provide empirical evidence on such relationships. One approach adopted in this literature is to construct an aggregate measure of deviations of asset prices from their "no-arbitrage" counterparts and to relate such measure to overall levels of pressures in funding markets, as well as variables aiming to capture balance sheet constraints of financial intermediaries, and then examine if such measures predict returns.[2]

In this paper we propose an alternative empirical approach that relies on measures of funding liquidity that directly reflect funding costs at the individual security level rather than on indirect aggregate proxies of funding liquidity that, in addition, are sometimes constructed from prices of instruments different from those whose returns are examined and from balance sheets of investors and intermediaries that are not guaranteed to be marginal price-setters in the particular market under investigation. In particular, we use the market for mortgage-backed securities (MBS)—in which the collateralized funding rates are available at security level and reflect both the cost of funding of these securities and, thus, funding liquidity, and their scarcity and, thus, market liquidity—as a laboratory and show that that exposures to liquidity risks measured from such rates are priced in the cross-section of MBS excess returns even after controlling for prepayment risk.

In addition to shedding light on links among funding liquidity, market liquidity and asset prices, our paper contributes to understanding of liquidity in the MBS market and determinants of MBS spreads—the topics that are of great importance in light of the MBS and mortgage markets' size, as well as the potential interactions among MBS spreads, mortgage rates, investors' demand for MBS and lenders' propensity to originate mortgages, but that received a fairly limited attention in the literature.

MBS are financed and borrowed through dollar rolls and repurchase agreements. In particular, dollar rolls are used to finance securities in the to-be-announced (TBA) market, a very active and liquid forward market of agency mortgage-backed securities. Prices of dollar rolls can be used to

[2] For example, see Brunnermeir and Pedersen (2009) for a theoretical model and Fontaine and Garcia (2012), Hu, Pan and Wang (2013) for empirical studies.

derive the so-called dollar roll implied financing rates (IFRs). The IFRs can be thought of as indicators of liquidity in the TBA market; they reflect the scarcity of a given MBS security and the cost of its financing in that market. A decline in the IFR can be interpreted as a signal that the MBS collateral is more scarce or special and can be funded at a lower cost. In this paper, we first investigate the determinants of this indicator of MBS liquidity, with a particular focus on the effects of supply-demand factors, and then examine whether liquidity shocks are priced in the cross-section of expected returns on MBS.

As we discuss further, in certain ways dollar rolls are similar to repo transactions, possessing the features of both general and special collateral repos. The findings suggesting that in the Treasury collateral market net supply affects repo rates lead us to expect that similar factors may be relevant for the dollar-roll-implied IFRs.[3] Indeed, we find that supply-demand factors are important determinants of IFRs. In particular, we document that a lower private supply of MBS, higher agency CMO production, and higher volume of MBS transactions by primary dealers are associated with lower IFRs or increased dollar roll specialness. Similarly, an increase in the Federal Reserve's MBS holdings and outright purchases are associated with dollar roll specialness of some coupons. This effect is not surprising, given the large size of the Federal Reserve's purchases and the role that movements in IFRs play in alleviating short-term imbalances between the supply of and demand for collateral by incentivizing holders of MBS to lend securities.

There is an extensive literature showing that investors in various markets demand compensation for holding assets with lower liquidity and for risk associated with changes in liquidity. More recently the literature has focused on funding liquidity—an aspect of liquidity captured in the TBA MBS market by the dollar-roll IFRs—and its interplay with market liquidity and asset prices.[4] Naturally, the question arises whether liquidity risk is priced in the TBA market. To measure liquidity risk, we first construct MBS liquidity shocks by computing the components of IFRs that cannot be explained by our reduced-form model. We then capture liquidity risk by estimating sensitivities of individual MBS returns to these shocks (liquidity betas) and show that

[3] Papers studying effects of supply-demand factors on Treasury security-specific repo rates include Duffie (1996). Jordan and Jordan (1997), Krishnamurthy (2002), Moulton (2004), Graveline and McBrady (2011) and D'Amico, Fan and Kitsul (2014).
[4] See, for example, Brunnermeier and Pederesen (2009), Garcia and Fontaine (2012).

MBS issues with higher exposures to liquidity shocks on average provide higher excess returns. As proxies for MBS excess returns, we use securities' option adjusted spreads (OAS), following Gabaix, Krishnamurthy and Vigneron (2007) who argue that using OAS in place of the actual returns reduces the measurement error at the cost of higher potential of a model misspecification for the prepayment option embedded in the MBS.

Next, we estimate a time-varying market price of liquidity risk and document its rich time-series dynamics. In particular, we find that the market price of liquidity risk jumped at the end of 2008, while remaining elevated through the end of the first quarter of 2009, that it trended up from the end of 2011 through December of 2013, and that it started reverting towards lower values in 2014. Interestingly, the timing of the increases in the price of the MBS liquidity risk aligns with the timing of severe market stress during the financial crisis, as well as the timing of early stages of the Federal Reserve's purchases of agency MBS under the reinvestment and the flow-based asset purchase programs.

Since our indicator of liquidity risk is derived using prices of MBS dollar rolls, there is a possibility that it contains information about prepayment risk. To control for this, we run panel regressions of individual securities' OAS on the compensation for liquidity risk together with compensation for the prepayment risk. We find that compensation for the liquidity risk remains a significant component of MBS risk premiums even after controlling for compensation for the prepayment risk.

From a broader perspective, our findings shed light on the question of whether supply-demand imbalances in collateral markets can have implications for collateral rental rates and broader asset prices, which has become a subject of increased attention in light of recent regulatory developments that could potentially boost demand for high-quality liquid assets (HQLA) and large asset purchases by central banks, and can be also viewed as a part of the literature studying asset price responses to expected and unexpected changes in the net supply of securities.[5,6] Most

[5] See, for example, Committee on the Global Financial System report for a general discussion of factors influencing supply and demand for HQLA and D'Amico et al. (2014), which examines the effects of such factors on Treasury collateral special repo rates.

[6] Studies of such effects in equity markets include Shleifer (1986), Kaul et al. (2000), Wurgler and Zhuravskaya (2002), Greenwood (2005) and those in the bond markets include Brandt and Kavajecz (2004), Lou et al. (2013), D'Amico and King (2013).

of the papers exploring the impact of supply-demand factors on the agency MBS market focus on the effect of the Federal Reserve's MBS purchases on MBS prices in the TBA market.[7] Our study is the first to explore the effects of an extensive array of drivers of dollar-roll IFRs, including several agency MBS supply-demand factors, and to document that liquidity risk embedded in the IFRs is priced in the cross-section of agency MBS returns.

Our paper is related to the study by Boyarchenko, Fuster and Lucca (2014) who find that cross-section of OAS of MBS sorted on their moneyness is explained by pre-payment risk, while time-series variation in OAS is mostly due to non-prepayment risk factor, which could potentially reflect liquidity and supply-demand imbalances. In our work, we employ an observable proxy of liquidity in the MBS market and find that, after controlling for prepayment risk, variance of innovations in this measure of liquidity, as well as exposure to these innovations, is priced in the cross-section of OAS.

The rest of the paper proceeds as follows. Section 2 explains the mechanics of a dollar roll transaction and the interpretation of IFRs. Then, it explores the response of dollar roll IFRs to private and public supply-demand of agency MBS. Section 3 quantifies the compensation for liquidity risk in the TBA market. Section 4 concludes.

2. Liquidity in the agency MBS market

2.1. A brief introduction to dollar rolls and IFRs

In some respects, dollar roll transactions are very similar to MBS repo transactions. In both transaction types one party agrees to sell securities to another (the front leg) and purchase them back at a later point (the back leg). Two important differences are that (1) in a dollar roll transaction the repurchased security can be "substantially similar" [8] to the one sold originally, as opposed to exactly the same, as in the repo transaction; (2) the ownership of the security sold in

[7] See, for example, Krishnamurthy and Vissing-Jorgensen (2013), and Stroebel and Taylor (2012). Kandrac (2014) and Song and Zhu (2014) consider the impact of Federal Reserve Purchases on IFRs. In addition, Song and Zhu (2014) also studies the impact of mortgage rates and MBS issuance on IFRs.

[8] "Substantially similar" means that the security needs to have the same basic characteristics, including the issuing agency, term, and coupon.

the dollar roll transaction is transferred to the purchaser, who receives the intervening cash flows such as (scheduled and unscheduled) principal and coupon payments.

Dollar rolls can be also viewed as combinations of forward sale and purchase transactions in the to-be-announced (TBA) market; both the front- and back-leg transactions are forward-settled. The front-leg price is often above the back-leg price and, therefore, the spread between the two is often referred to as the "drop". The drop compensates the roll seller for the lost "carry" (coupon and principal payments) and the risk of being delivered a less desirable security at the back leg, while also reflecting net funding and collateral demands in the MBS market. Given a projected prepayment speed and the information on scheduled interest and principal payments, it is possible to calculate the drop that would be due only to payments foregone by its seller (everything else equal). By comparing this projected drop to the observed drop one can compute the dollar-roll IFR. As dollar rolls are often used for financing long positions in MBS securities, the IFRs can be thought of as a rough gauge of expected funding pressures in the dollar roll market, conceptually similar to how General Collateral (GC) repo rate is used in the repo market. This is in contrast to special collateral (SC) repo in Treasury markets, which is mostly used to obtain securities for the purposes of subsequent short selling rather than to fund long positions. At the same time, dollar rolls possess features similar to those of special collateral (SC) repo, as the dollar roll's underlying collateral includes MBS satisfying specific albeit not exhaustive characteristics, such as coupon, maturity and issuing agency, rather than just belongs to a broad asset class. IFRs tend to decline when underlying MBS collateral is scarce and rise when collateral is readily available.

Another way to visualize the concept of the IFR, is to consider an investor who is scheduled to take delivery of an MBS with a particular set of characteristics and a nominal value of L_t dollars in the TBA market. This investor has two options. First, she could postpone the delivery and roll this position from month t to month t+1 (and reinvest the proceeds of the sale at the rate r_t). Alternatively, she could hold the MBS over the same period. The IFR is the rate of return under which the investor receives the same expected cash flows under these two choices. That is, given assumptions about the expected prepayment rate, the IFR must satisfy the following equality,

$$\underbrace{(1+r_t)P_tL_t - P_{t+1|t}E(L_{t+1})}_{Cash\ flow\ from\ dollar\ roll\ plus\ reinvesment} = \underbrace{PR_t + I_t + E(PP_{t+1})}_{Cash\ flow\ from\ MBS} \qquad (1)$$

where P_t is the front leg MBS price for month-*t* settlement, $P_{t+1|t}$ is the agreed repurchase MBS price for month-*t+1* settlement, $E(L_{t+1})$ is the expected remaining principal after the scheduled principal payments and its prepayments, PR_t is scheduled balance payment, I_t is the interest payment, and $E(PP_{t+1})$ is the expected principal prepayment. The IFR for the month *t/t+1* dollar roll is denoted r_t.[9]

In contrast to the case of the MBS GC repo rate, for each TBA contract there is a separate IFR series that may differ in lengths depending on when the MBS of a given coupon was issued. Figure 1 displays the IFRs for 30-year Fannie Mae MBS across coupons ranging from 3.0% through 6.5%, while Figure 2 displays the spreads of the same IFRs in excess of the 1-month General Collateral (GC) agency MBS repo rate. Selected descriptive statistics for these series are presented in Table 1. The IFRs and their spreads over the MBS repo rates exhibit some notable fluctuations. For example, during the recent financial crisis the IFRs increased, and these increases at times exceeded those of other financing rates. In particular, during the second half of 2008 the spread of the IFRs over the MBS repo rates reached levels exceeding 200 basis points.[10] Aside from these episodes, the IFR has remained below the MBS repo rate across the different coupons (see Table 1). Among other things, generally negative spreads between IFRs and the MBS GC repo rate reflect dollar roll sellers' compensations for providing the option of being delivered a "substantially similar" security on the back leg and for the lost carry.

2.2 Empirical model: explanatory variables, hypotheses and results

[9] In practice, investors also compare the exact dates during the month when interest and principal payments are received with the front- and back-leg delivery dates to account for accrued interest when computing the IFRs.

[10] We have heard several explanations for the sharp rise in the spread during this period. First, some financial firms, faced with funding pressures, sold MBS holdings. Second, rumors of a government-sponsored refinancing program created risks of faster-than-expected prepayment speeds. Lastly, the reduction of balance sheet capacity of primary dealers likely prevented them from arbitrating the spread away by buying dollar rolls and funding these purchases in the MBS repo market.

To better understand the drivers of IFRs, we run the following monthly time-series regressions for the IFR of individual dollar roll contracts, using 30-year Fannie Mae MBS with coupons of 3 to 6.5%:

$$r_{i,t} = \alpha_i + \boldsymbol{z}_{i,t}\beta_{i2} + \boldsymbol{x}_t\beta_{i1} + \varepsilon_{it} \tag{2}$$

where $r_{i,t}$ is the IFR for the month *t*/*t+1* dollar roll, and $\boldsymbol{z}_{i,t}$ and $\boldsymbol{x}_t$ denote contract-level and aggregate explanatory variables, respectively. We define the IFR for the dollar roll with front-month *t* and back-month *t+1* as the average of the IFR from the day after the notification day for the month *t-1* through the notification day for the month *t*. We omit the first six months for each of the coupons because according to anecdotal reports trading tends to be scarce over the first few months after the newly-produced coupon is introduced into the market. We present results from time-series regressions for individual contracts, as well as from an unbalanced panel data regression that pools together the data across contracts and allows for contract-level fixed effects. Our sample starts in January 2003 and ends in July 2014. Next, we describe the explanatory variables that we include into our empirical model and discuss their hypothesized and estimated effects on the IFRs. The baseline results are presented in Table 2.

To capture the private supply of MBS, we include the total **stock of MBS outstanding** (net of SOMA holdings) for each of the coupons.[11] Our null hypothesis is that a larger private supply makes the security less scarce in the dollar roll market and, therefore, should be associated with a higher IFR. Similarly, an increase in SOMA holdings would decrease the outstanding private supply of MBS leading to a reduction in the IFR.[12] Indeed, the coefficients on our supply variable are positive and highly significant in all of our regressions except one. This result is consistent with the findings in Kandrac (2013).

We also explore the impact of the **Federal Reserve's agency MBS outright purchases** on IFRs. We include to our regression the face value (in US$ billions) of outright SOMA purchases. Our hypothesis is that the coefficient on this variable is negative, because larger outright SOMA

[11] The availability of securities in the TBA market is determined by the stock of cheapest-to-deliver securities as participants tend to deliver the most economical, or "cheapest-to-deliver," securities. We approximate this stock with the total MBS outstanding, although it is important to keep in mind that this proxy is not perfect.
[12] Although this effect may be offset to the extent Federal Reserve purchases of MBS lower primary mortgage rates and lead to an increase in mortgage origination.

purchases might push up the front-month MBS prices, leading to increases in the "drop" and decreases in the IFRs. Anecdotal evidence suggests that large amounts of expected or actual SOMA purchases in a given contract can create temporary scarcity, which is generally resolved when IFRs decline, as lower IFRs provides an additional incentive for holders of MBS to lend their collateral. The estimated impact of outright SOMA purchases is in fact negative and statistically significant for the 3.5% and 4.5% coupons, implying that according this specification MBS purchases had some impact on those securities' IFRs. The coefficient on these terms imply that a $1 billion increase in SOMA monthly purchases lower IFRs by between 1.2 and 2.4 basis points.

IFRs are computed under particular assumptions on the expected speed of prepayments and, therefore, changes in expected prepayment speeds will have an impact on the breakeven financing rate of roll transactions, unless the price of the drop fully adjusts to reflect the change in expected prepayments (see Equation 1). For example, for a given price drop, an increase in the expected prepayment rate will lower (increase) the IFRs for securities purchased at a discount (premium) to their face values. The rationale is that faster prepayments will increase (decrease) yield-to-maturity for discount (premium) securities and make holding (rolling) them more desirable. As a result, investor's "compensation" (captured by the "reinvestment" rate or IFR, $r_{i,t}$) from rolling the security should increase. To test whether changes in prepayment expectations affect discount and premium securities, we use the **median prepayment speed (PSA) forecast of the major Wall Street dealers** interacted with the **spread between a security's price in the TBA market over its face value of 100** (we refer to this spread as "premium"). The coefficient is negative and statistically different from zero for all coupons except one. This result is somewhat surprising, given that in an efficient market, the drop would be expected to adjust to reflect the new information about anticipated prepayment speeds. Our findings may indicate frictions in the market or, possibly, prepayment model misspecification. Additional work on this topic could be useful.

To capture the general level of MBS financing rates, we use **MBS repo rates**. Even though the two funding instruments are used for different purposes and financing rates implied by dollar rolls are typically modestly lower than the MBS repo rates, the two rates are closely linked as some investors have an ability to switch between the two sources of financing. We hypothesize the coefficient of repo rate to have a positive sign. The hypothesis is confirmed by our estimation results in the 4.0%- through 6.5%-coupon regressions, as well as in the pooled regression. As expected, the coefficient is positive and we cannot reject the null hypothesis that the coefficient is equal to one. Surprisingly, the coefficients for the lower coupons suggest that the IFR fluctuations do not seem to mirror those of the MBS repo rate over that period. This may partially be due to the smaller number of observations for those coupons, as they were not issued earlier in the sample. It is also possible that SOMA purchases of MBS in the TBA market over this period weakened the traditional relationship between dollar rolls and MBS repo.

Higher agency CMO production during the front month could increase demand in the TBA market during that month and, as a result, decrease the IFRs by driving the front-month MBS prices higher relative to their back-month counterparts. In the regression we use the data for the overall production of CMOs that use 30-year Fannie Mae MBS of various coupons as underlying cash-flow source. The estimated CMO coefficients are indeed negative in most cases but are significant only in the specifications with the 3.0% and 5.5% coupon IFRs, as well as in the pooled regression.

Large **transactions by dealers (dealer volumes)** in agency MBS could be associated with higher IFRs if dealers finance their positions in the dollar roll market (and thus sell dollar rolls) or with lower IFRs if the dealers are purchasing the securities through this market. Data are collected from the FR-2004 and correspond to inter-dealer MBS transactions. The coefficient on **dealer volume growth** is negative and statistically significant for the specifications with the 4.5% and 5.5% through 6.5% coupons as well as in the pooled regression. This finding is consistent with reports that MBS dealers primarily buy dollar rolls and use them to cover short positions that are hedges against other MBS products. For example, dealer positions in Treasury securities and Treasury GC repo rates appeared to line up after they had to absorb a large amount of short-term Treasury securities during the Maturity Extension Program.

To check robustness of our results, we also estimate our model with the spread of the IFR over the MBS GC repo rate as a dependent variable. The IFR over the MBS repo is less persistent than the IFR, so any econometric issues arising from using very persistent independent variables should be mitigated in specifications using this spread (see Table 1). Table 3 presents the estimated coefficients for coupon-specific and pooled regressions. In general, all our conclusions remain valid.

3. Is there a premium for liquidity in the agency MBS market?

Our results show that fluctuations in private and public supply and demand of agency MBS have an impact on the funding liquidity of agency mortgage securities. In this section we explore if investors require a higher compensation for mortgage securities with a higher uncertainty about funding liquidity conditions.

To proxy for the agency MBS risk premium we use each security's option-adjusted spread (OAS). The OAS captures the valuation of the MBS relative to Treasuries and explicitly takes into account the value of prepayment option. In particular, in the spirit of the CAPM framework, we posit the following expression for the *OAS*,

$$OAS_{i,t} = \alpha_{i,t} + \beta^{i}_{liq}\lambda_t \quad (3)$$

where $OAS_{i,t}$ is the option-adjusted spread at time *t* for an MBS with coupon *i*, β^{i}_{liq} is the liquidity beta, λ_t is the market price of liquidity risk, $\alpha_{i,t}$ and captures other risk factors that are unrelated to liquidity risk.

3.1 Estimating the funding liquidity beta

As a first step, we rely on the model presented in Section 2.2 to recover an estimate of β^{i}_{liq}. As in Gabaix, Krishnamurthy, and Vigneron (2007), we assume that the error term from our model of the IFR for coupon *i* follows a first-order autoregressive process:

$$\varepsilon_{i,t} = \rho\varepsilon_{i,t-1} + u_{i,t} \quad (4)$$

Then, we use the standard deviation of $u_{i,t}$ for each coupon as our proxy for the liquidity beta of each security, $\hat{\beta}^{i}_{liq} = \sigma(u_{i,t})$. Our measure of β^{i}_{liq} is intended to capture the uncertainty about funding liquidity conditions across different securities.[13] We also perform the following robustness check. We compute the liquidity betas using the error term from the regression of the spread between the IFR and one-month MBS repo rate. Our results are robust to this alternative measure of liquidity beta.

3.2 Market price of funding liquidity risk

As a first exercise we compute the market price of funding liquidity risk using an unconditional version of the OAS model,

$$\overline{OAS}_i = \overline{\alpha}_i + \hat{\beta}^i_{liq}\overline{\lambda} + \epsilon_i \tag{5}$$

Figure 3 shows a scatter plot of the liquidity beta and the average OAS for generic Fannie Mae MBS with coupons from 3.0% to 6.5%. The figure shows that securities with a higher funding liquidity beta also have a higher average OAS. This suggests that the market price of liquidity risk in MBS markets is positive. We run a cross-sectional regression and find that indeed the market price of liquidity risk is positive and statistically significant.

To explore the time-series variation of the market price of liquidity risk, we estimate λ_t each month in our sample running the following cross-sectional regression,

$$OAS_{i,t} = \alpha_{i,t} + \hat{\beta}^i_{liq}\lambda_t + \epsilon_{i,t} \tag{6}$$

where $\hat{\beta}^i_{liq}$ is our proxy of the liquidity beta for coupon *i*. Figure 4 displays the estimates of λ_t along with 95% confidence bands. The time series plot of λ_t reveals substantial time variation in the market price of liquidity. In particular we find that the market price of liquidity risk jumped at the end of 2008 and remained elevated for a few months until the end of the first quarter of 2009. This episode coincides with the Federal Reserve's QE1 program in which the Federal Reserve purchased agency MBS and agency debt. We also find that λ_t trended up since the end of 2011 until December of 2013, and started reverting towards lower values in 2014. The increase in λ_t coincides with the subsequent expansion of the Federal Reserve's holdings of

[13] We use the term "liquidity beta" to refer to the measure of liquidity risk as the matter of convenience. Our terminology is analogues to that of Gabaix, Krishnamurthy, and Vigneron (2007) who use standard deviations of residuals from a prepayment model as a one of their proxies for prepayment betas.

agency MBS under reinvestment program announced in September of 2011 and the QE3 program. Interestingly, the decline in λ_t coincides with the tapering of the Federal Reserve's MBS purchases under QE3 in December of 2013.

3.3 The importance of liquidity risk: Panel data evidence

There is the possibility that our model is capturing other potential determinants of the risk premium on agency mortgage securities, in particular, prepayment risk. To test for the robustness of our results to other variables driving the MBS OAS, we use the entire panel to estimate the following regression,

$$OAS_{i,t} = \alpha_i \; + \theta_t + \mathbf{A}\hat{\beta}^i_{liq}\lambda_t + \mathbf{B}\,\hat{\beta}^i_{prepayment}(\overline{c}_t - r_t)\big(c_{i,t} - r_t\big) + \phi GSE\_SPREAD_t + \epsilon_{i,t} \quad (7)$$

where we control for both year (θ_t) and security (α_i) fixed-effects, and the spread between 10-year Fannie Mae debt and U.S. Treasury securities. This spread intends to capture the default risk of agency MBS. Stroebel and Taylor (2013) show that this spread is closely related to credit default swap spreads on debt issued by GSEs. Most importantly, we add as control variable the risk compensation for prepayment risk of Gabaix, Krishnamurthy, and Vigeron (2007). Gabaix et al. (2007) shows that the OAS is proportional to $\hat{\beta}^i_{prepayment}(\overline{c}_t - r_t)\big(c_{i,t} - r_t\big)$, and this term captures the compensation for prepayment risk. As in their paper, we compute $\overline{c}_t$ as the average coupon of outstanding Fannie Mae securities weighted by the amount outstanding of the security. We set r_t as the current primary mortgage rate. Finally, we compute the prepayment beta $\hat{\beta}^i_{prepayment}$ as the standard deviation of the residual from a model of observed prepayment rates and the median forecast from Wall Street dealers.

Column 1 of Table 4 shows the estimates of **A** without prepayment risk, and column 2 shows the estimates of **A** and **B**. As expected, **A** is positive and statistically significant even after controlling for default risk and prepayment risk. Also, the coefficients **B** and **C** are also significant and statistically different from zero. These results suggest that our measure of liquidity risk compensation is not capturing these measures of default or prepayment risk.

Finally, we also take advantage of the full panel data to estimate the model of OAS in differences. Column 3 of the same table presents the results. Again, the coefficient on A is still positive and significant which suggests our results are not spurious.

4. Alternative definition of liquidity risk

One might argue that our results may be influenced by our choice of empirical specification for IFRs in equation (2) and the way we compute liquidity betas. Therefore, in this section we employ an alternative approach for constructing liquidity shocks and corresponding exposures. Under this approach, we assume that the IFRs across all the coupons are driven by an unobservable factor that follows an AR-1 process. We also assume the shocks in this model to be normal and iid and estimate the following system of equations using Kalman filter:

$$L_{t+1} = c + FL_t + v_{t+1} \tag{8}$$

$$r_{i,t} = HL_t + w_t \tag{9}$$

After we estimate the parameters, we calculate conditional expectations for the latent factor and compute liquidity shock as $Liq = L_{t+1|t} - L_{t|t}$. One advantage of this approach is that we obtain a single liquidity factors that spans the sample over which at least one security comprising our unbalanced panel is available; the sample employed for analysis presented in this section spans the period of January 1998 through August 2015. The evolution of the liquidity shock obtained with this method is shown in Figure 5. A negative shock should be interpreted as exerting downward pressure on IFRs, leading to an increase in MBS specialness or scarcity, but lower funding cost for an investor willing to finance a long MBS through the dollar roll.

Next, we obtain liquidity betas from the covariance of OAS with liquidity shocks and plot coupon-specific OAS (averaged over time) versus the corresponding betas in Figure 6. As can be seen in that figure, assets can have positive and negative liquidity betas. On one hand, negative shocks to IFRs mean that, since securities became scarcer, they can be funded at lower IFRs, leading to higher prices and lower OAS. On the other hand, more scarce securities are less liquid, and lower liquidity should exert downward pressure on prices and push OAS higher. The ultimate sign of beta will depend on which effect dominates. The assets with larger negative liquidity betas tend to be the MBS with higher coupons, which were originated earlier in the sample period. In addition, it can be seen that the relationship between OAS and liquidity betas obtained with this method is negative. In other words, assets whose OAS are more negatively correlated with liquidity shocks command a higher average return. In other words, our results suggest that investors require a return premium to hold those assets whose prices drop and OAS increase more per unit of liquidity decline than those of other assets. The similar logic also works for positive betas: the assets whose prices increase and OAS decrease more per unit of funding/financing cost decrease should, on average command higher prices and lower OAS. Indeed, we observe the negative relationship between OAS and betas for observations with both negative and positive betas in Figure 6. These results are broadly consistent with those presented in Figure 3 after taking into account the fact that liquidity betas presented in that figure are computed as volatility of liquidity shocks rather than covariance between shocks and OAS (so in that case larger positive beta means higher liquidity risk).

To examine how liquidity risk interacts with prepayment risk, we repeat the exercise for average OAS of MBS portfolios containing securities whose coupons lie within some fixed distance from the production coupon (for OAS of individual MBS). In particular, we form four portfolios, with portfolio 1 containing coupons below the lowest of the main production coupons, portfolio 2 containing main production coupons, portfolio 3 containing coupons up to 100 basis points above the main production coupon, and portfolio 4 containing coupons that more than 100 basis points above the production coupon. The rationale for forming such portfolios is that loans underlying securities with coupons above the production coupons have a higher probability of prepayment. Figure (6) plots average OAS of these portfolios against their liquidity betas. It can be seen that the higher the portfolio's distance is from the production coupon, the lower (more negative) is the portfolio's liquidity beta and the higher is its time-averaged OAS.

Figures 7 and 8 display time-series of market prices of liquidity risk based on portfolios sorted on coupon and distance to production coupon. The prices of risk are produced with regressions analogous to those we used to obtain prices of risk shown in Figure 4, but obtained with the new liquidity shock and beta measures. Again, taking into account that more negative betas now imply higher liquidity risk, the dynamics of market prices of risk is broadly consistent with that in Figure 4.

5. Conclusions

In this paper we explore the determinants of and risk compensation for liquidity in the mortgage-backed securities market. Using the implied financing rates (IFRs) as indicators of liquidity in the MBS market, we provide empirical evidence of that supply-demand factors are important determinants of IFRs. We document that a lower private supply of MBS, higher agency CMO production, and higher volume of MBS transactions by primary dealers are associated with lower IFRs or increased dollar roll specialness. We also show that an increase in the Federal Reserve's MBS holdings and outright purchases are associated with increased scarcity of some securities.

We use our reduced-form model of the IFR to extract a liquidity beta for each mortgage security in our sample and show that liquidity risk is priced. In particular, we show that that MBS securities with higher exposures to liquidity shocks (higher liquidity betas) on average provide higher excess returns. This result is robust even after we control for the measure of prepayment risk of Gabaix, Krishnamurthy and Vigneron (2007) and agency-related default risk.

Our results contribute to understanding of the links between funding and cash markets, as well as between liquidity and expected returns.

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Table 1: Summary Statistics

	Mean	Std. Dev.	Min	Max	Autocorrelation
Coupon	Implied Financing Rate (IFR)				
3.0%	-0.47	0.65	-2.15	0.21	0.642
3.5%	-0.48	0.50	-1.57	0.27	0.841
4.0%	-0.10	0.29	-0.78	0.53	0.438
4.5%	1.37	2.21	-2.22	5.37	0.979
5.0%	1.00	2.45	-3.56	5.56	0.978
5.5%	1.85	2.57	-2.74	6.82	0.978
6.0%	2.16	2.52	-3.13	6.70	0.962
6.5%	2.19	2.53	-6.33	6.64	0.889
	Spread of the IFR over MBS GC repo				
3.0%	-0.64	0.65	-2.30	0.01	0.659
3.5%	-0.66	0.52	-1.74	0.04	0.852
4.0%	-0.29	0.30	-1.03	0.17	0.463
4.5%	-0.32	0.60	-2.47	1.90	0.732
5.0%	-0.66	0.95	-3.69	2.23	0.856
5.5%	-0.39	0.76	-2.88	2.45	0.823
6.0%	-0.24	0.75	-3.23	2.65	0.735
6.5%	-0.21	0.84	-6.47	2.68	0.388

This table presents summary statistics of the implied financing rates (IFRs) for Fannie Mae 30-year MBS securities. The IFR for the dollar roll with front-month t and back-month $t+1$ is computed using daily data as the average of the IFR from the day after the notification day for the month $t-1$ through the notification day for the month t.

Table 2: Drivers of the Implied Financing Rate

	(1) 3.0%	(2) 3.5%	(3) 4.0%	(4) 4.5%	(5) 5.0%	(6) 5.5%	(7) 6.0%	(8) 6.5%	(9) All
Security specific variables									
ln MBS Outstanding	0.173***	0.370***	0.131**	0.439***	0.678**	0.940***	0.842***	-0.223	0.277***
	(5.68)	(11.67)	(2.16)	(3.12)	(2.58)	(5.95)	(4.28)	(-1.26)	(3.90)
Prepayment forecast x Premium	-0.0369*	-0.0640***	-0.0204***	-0.0517***	-0.0643***	-0.0328***	-0.0290	-0.0614**	-0.0446***
	(-2.00)	(-20.48)	(-4.17)	(-5.03)	(-5.94)	(-3.33)	(-1.42)	(-2.56)	(-5.78)
SOMA MBS purchases	0.00556	-0.0237*	0.00173	-0.0116**					-0.00877
	(0.49)	(-1.96)	(0.44)	(-2.23)					(-1.07)
Aggregate market variables									
MBS repo rate 1m	7.182**	3.849***	1.178**	1.026***	0.873***	0.896***	0.912***	0.909***	0.959***
	(2.51)	(5.04)	(2.47)	(29.46)	(17.19)	(29.85)	(9.81)	(8.81)	(26.05)
ln CMO FNMA coll.	-0.437*	0.210***	0.0704	-0.142	0.0310	-0.186**	-0.139	-0.149	-0.109**
	(-1.87)	(2.88)	(1.36)	(-1.01)	(0.15)	(-2.07)	(-0.77)	(-1.30)	(-2.43)
Growth of dealers MBS volume	-0.000929	0.00338*	-0.00273	-0.00317**	-0.000546	-0.00322**	-0.00762**	-0.00667***	-0.00402***
	(-0.53)	(1.95)	(-1.63)	(-2.04)	(-0.30)	(-2.38)	(-2.58)	(-2.91)	(-3.61)
Constant	-2.472***	-5.172***	-1.841**	-4.823***	-8.407**	-11.50***	-9.824***	3.686*	-2.985***
	(-4.50)	(-14.65)	(-2.33)	(-2.91)	(-2.49)	(-5.57)	(-3.99)	(1.74)	(-3.93)
T	42	61	66	128	136	138	138	138	847
Adj. R-squared	0.494	0.735	0.239	0.962	0.949	0.967	0.914	0.859	-

This table presents the estimated coefficients of a regression of the IFR of individual dollar roll contracts on several supply-demand factors. The sample covers 2003m1 to 2014m7. The last column presents the results of a pooled regression using all coupons and allowing for contract level fixed-effects. The table reports in parenthesis Newey-West *t*-statistics for the time series regressions and *t*-statistics using clustered standard errors by coupon for the panel regression. ***, **, and * denote significant at the 1%, 5% and 10%, respectively.

Table 3: Drivers of the Spread between the IFR and the MBS GC Repo Rate

	(1) 3.0%	(2) 3.5%	(3) 4.0%	(4) 4.5%	(5) 5.0%	(6) 5.5%	(7) 6.0%	(8) 6.5%	(9) All
Security specific variables									
ln MBS Outstanding	0.125***	0.339***	0.122**	0.407***	0.616**	0.943***	0.860***	-0.182	0.265***
	(2.83)	(9.91)	(2.62)	(3.11)	(2.30)	(5.52)	(4.08)	(-1.01)	(4.37)
Prepayment forecast x Premium	-0.00257	-0.0526***	-0.0197***	-0.0530***	-0.0497***	-0.0172*	-0.0130	-0.0430***	-0.0398***
	(-0.30)	(-19.16)	(-4.55)	(-5.32)	(-5.81)	(-1.83)	(-1.39)	(-3.54)	(-8.44)
SOMA MBS purchases	-0.000206	-0.0257**	0.00145	-0.0123**					-0.00743
	(-0.03)	(-2.29)	(0.41)	(-2.36)					(-1.11)
Aggregate market variables									
ln CMO FNMA coll.	-0.315	0.302***	0.0680	-0.133	0.00588	-0.194*	-0.155	-0.186	-0.118**
	(-1.32)	(4.59)	(1.36)	(-0.93)	(0.03)	(-1.91)	(-0.87)	(-1.33)	(-2.65)
Growth of dealers MBS volume	-0.00236	0.00305	-0.00278*	-0.00306*	-0.00122	-0.00358**	-0.00804***	-0.00713***	-0.00418***
	(-0.98)	(1.64)	(-1.69)	(-1.97)	(-0.63)	(-2.48)	(-2.76)	(-3.12)	(-3.71)
Constant	-1.169	-4.556***	-1.705***	-4.426***	-7.904**	-11.85***	-10.36***	2.850	-2.928***
	(-1.37)	(-13.19)	(-3.24)	(-3.02)	(-2.26)	(-5.33)	(-3.88)	(1.44)	(-4.15)
T	42	61	66	128	136	138	138	138	847
Adj. R-squared	0.428	0.706	0.269	0.473	0.631	0.708	0.402	0.174	-

This table presents the estimated coefficients of a regression of the spread between the IFR and the MBS GC repo rate of individual dollar roll contracts on several supply-demand factors. The last column presents the results of a pooled regression using all coupons and allowing for contract level fixed-effects. The table reports in parenthesis Newey-West *t*-statistics for the time series regressions and *t*-statistics using clustered standard errors by coupon for the panel regression. ***, **, and * denote significant at the 1%, 5% and 10%, respectively.

Table 4: Liquidity risk and prepayment risk

	(1)	(2)	(3)
$\hat{\beta}^i_{liq}\lambda_t$	0.633***	0.531***	
	(0.055)	(0.071)	
$\hat{\beta}^i_{prepayment}(\overline{c}_t - r_t)(c_{i,t} - r_t)$		0.00414***	
		(0.001)	
GSE_SPREAD_t	0.605***	0.713***	
	(0.114)	(0.107)	
$\Delta\,\hat{\beta}^i_{liq}\lambda_t$			0.381***
			(0.053)
T	979	979	976
Adj. R-squared	0.741	0.750	0.332

This table presents the estimated coefficients from the following panel regression,

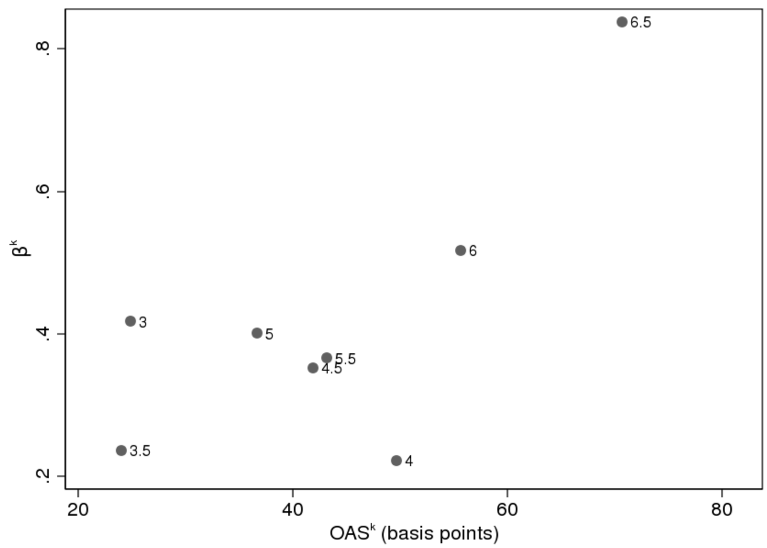
$$OAS_{i,t} = \alpha_i + \theta_t + \mathbf{A}\hat{\beta}^i_{liq}\lambda_t + \mathbf{B}\,\hat{\beta}^i_{prepayment}(\overline{c}_t - r_t)(c_{i,t} - r_t) + \phi GSE\_SPREAD_t + \epsilon_{i,t}$$

where we control for both year (θ_t) and security (α_i) fixed-effects, the spread between 10-year Fannie Mae debt and U.S. Treasury securities. and the risk compensation for prepayment risk of Gabaix, Krishnamurthy, and Vigeron (2007), $\hat{\beta}^i_{prepayment}(\overline{c}_t - r_t)(c_{i,t} - r_t)$. $\overline{c}_t$ is the average coupon of outstanding Fannie Mae securities weighted by the amount outstanding of the security. We set r_t as the current primary mortgage rate. The prepayment beta $\hat{\beta}^i_{prepayment}$ is the standard deviation of the residual from a model of observed prepayment rates and the median forecast from the most important Wall Street dealers. Column 1 sets $\mathbf{B} = 0$, and column 3 estimates the model in differences. Clustered standard errors by coupon and month are shown in parenthesis. The sample covers 2003m1 to 2014m7. ***, **, and * denote significant at the 1%, 5% and 10%, respectively.

Figure 1: Implied Financing Rate

Figure 2: Spread of IFR over MBS GC Repo Rate

Figure 3: The cross-section of MBS OAS and liquidity risk



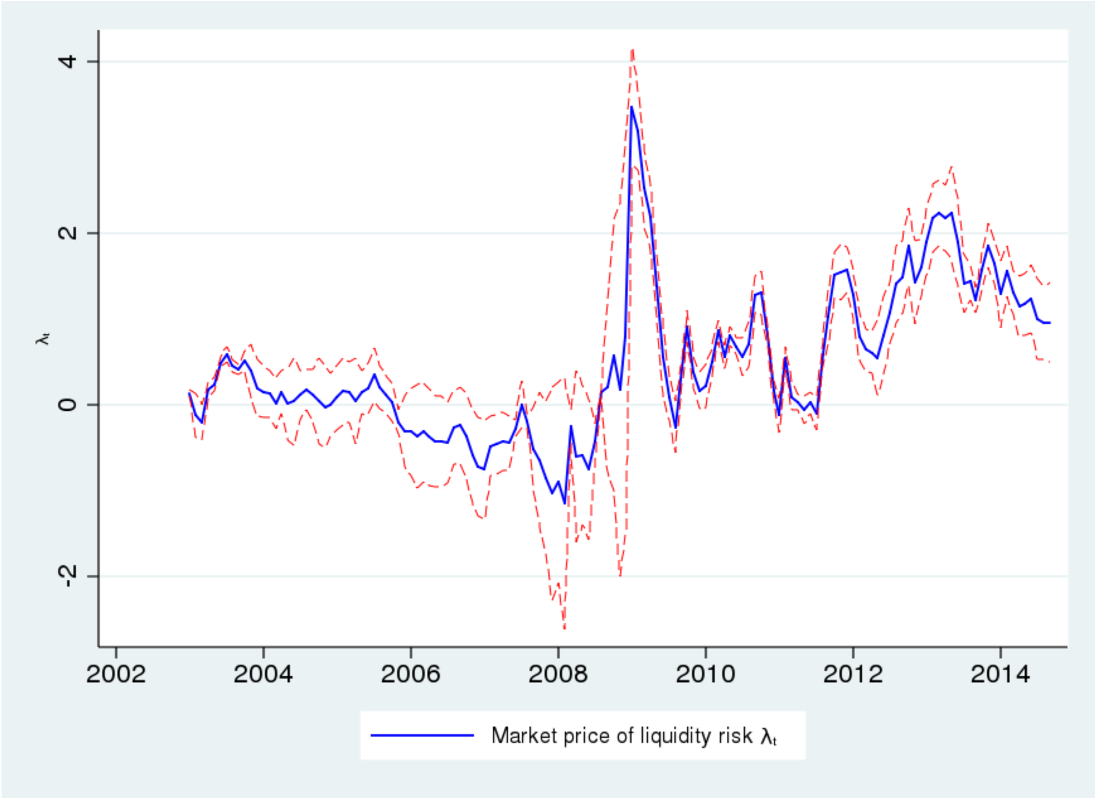
$\overline{\lambda}$	0.575**
	(0.180)
$\overline{\alpha}_i$	0.192
	(0.105)
T	8
Adj. R-squared	0.433

This figure shows a scatter plot of the liquidity beta $\hat{\beta}_{liq}^{i}$ and the average OAS for generic Fannie Mae MBS with coupons from 3.0% to 6.5%. Each point has a label identifying the coupon. The table presents the estimated market price of risk and $\overline{\alpha}_i$ from the following cross-sectional regression,

$$\overline{OAS}_i = \overline{\alpha}_i + \hat{\beta}_{liq}^{i}\overline{\lambda} + \epsilon_i$$

The OAS is measured in percentage points. The sample covers 2003m1 to 2014m7. ***, **, and * denote significant at the 1%, 5% and 10%, respectively.

Figure 4: Market Price of Liquidity Risk

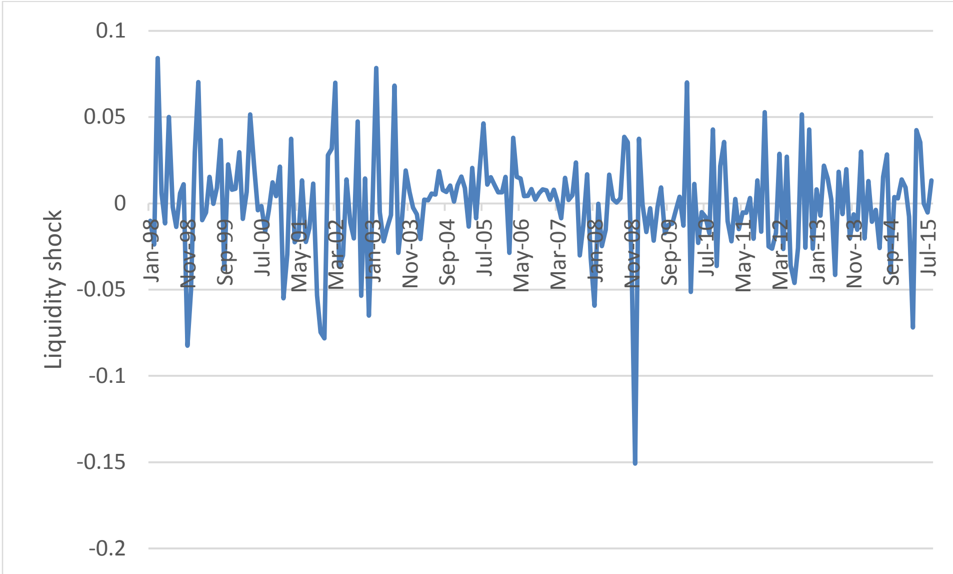


This figure shows the estimates of the market price of liquidity risk λ_t which is running the following cross-sectional regression each month in our sample,

$$OAS_{i,t} = \alpha_{i,t} + \hat{\beta}^i_{liq}\lambda_t + \epsilon_{i,t}$$

where $\hat{\beta}^i_{liq}$ is our proxy of the liquidity beta for coupon *i*, based on volatility of liquidity shocks The dotted lines are 95% confidence bands. The sample covers 2003m1 to 2014m7.

Figure 5. An Alternative Series of Liquidity Shocks



This figure shows a series of liquidity shock obtained by extracting a latent AR-1 common factor from the unbalanced panel of IFRs.

Figure 6. The cross-section of MBS OAS and liquidity risk obtained with alternative liquidity factor

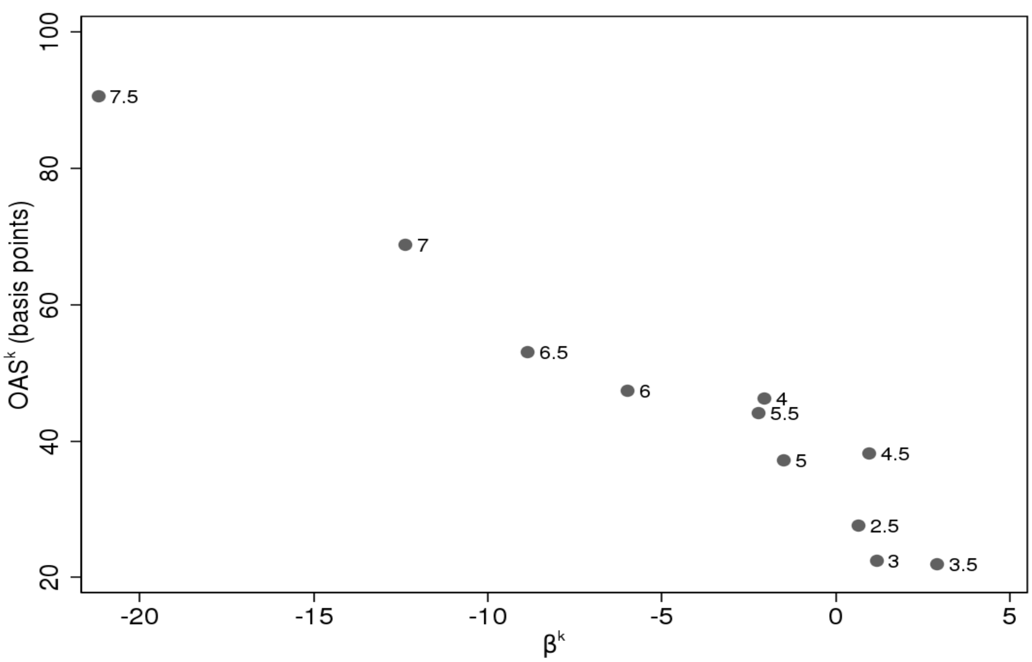
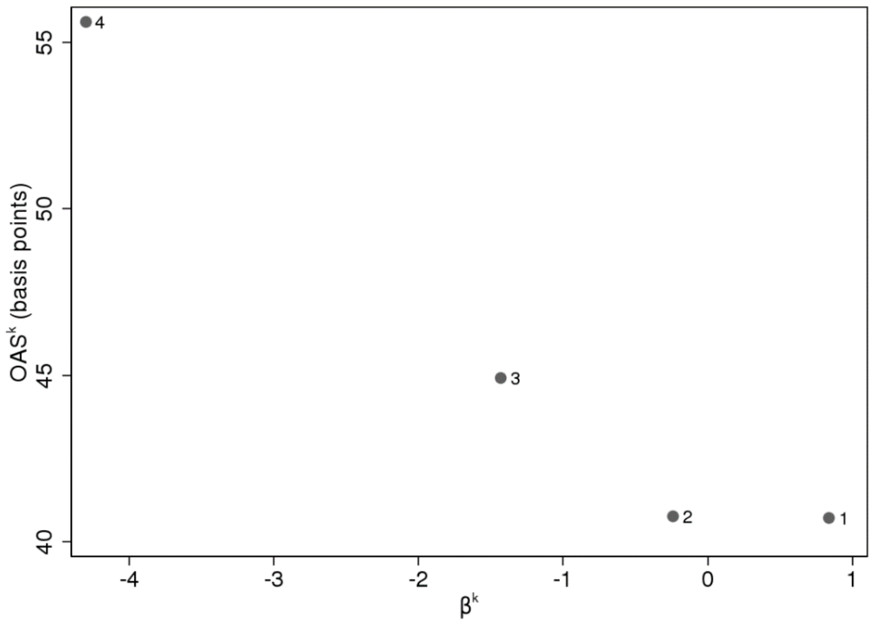


Figure 7. The cross-section of MBS OAS and liquidity risk obtained with alternative liquidity factor and portfolios formed by distance from production coupon



	(2) OAS (coupon)	(3) OAS (distance prod coupon)
$\overline{\lambda}$	-2.735***	-3.089**
	(0.156)	(0.453)
$\overline{\alpha}_i$	33.205***	41.534***
	(2.125)	(1.357)
T	11.000	4.000
Adj. R-squared	0.934	0.928

This figure shows a scatter plot of the liquidity beta $\hat{\beta}^i_{liq}$ and the average OAS for portfolios formed on distance to production coupon, with portfolio 1 containing securities with coupons below current production coupon, portfolios 2 containing securities with production coupons, portfolio 3 containing coupons up to 100 basis points above the production coupon, and portfolio 4 containing securities paying coupons that are more than 100 basis points above the production coupon. Each point has a label identifying the coupon. The table presents the estimated market price of risk and $\overline{\alpha}_i$ from the following cross-sectional regression,

$$\overline{OAS}_i = \overline{\alpha}_i + \hat{\beta}^i_{liq}\overline{\lambda} + \epsilon_i$$

The OAS is measured in percentage points. The sample covers 1998m1 to 2015m8. ***, **, and * denote significant at the 1%, 5% and 10%, respectively.

Figure 8. Market price of liquidity risk based on cross-sectional regression of portfolios sorted by coupon.

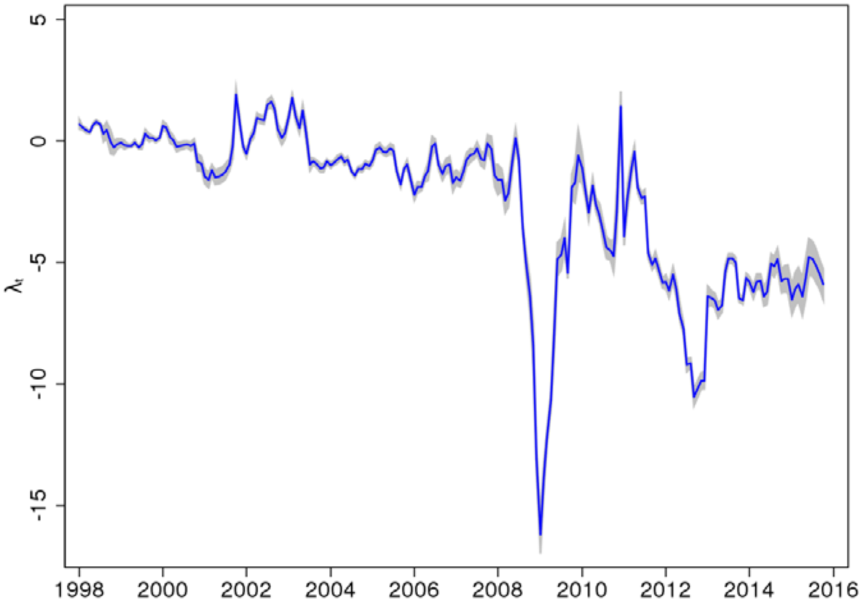


Figure 9. Market price of liquidity risk based on cross-sectional regression of portfolios sorted on distance to production coupon.