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ABSTRACT

We use the term structure of bank CD rates to examine whether maturity-transformation risk is priced into the rates banks offer customers. We find that depositors pay a significant cost for the liquidity provided by bank deposits. This cost is strongly related to the amount of maturity-transformation risk that these deposit accounts create. The cost is also negatively correlated with the convenience premia in Treasury markets, which suggests that households do not view deposit liquidity and Treasury liquidity as perfect substitutes. The results have important implications about the role of deposit franchises and market power in banking markets.

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A data appendix is available at http://www.nber.org/data-appendix/w32724

1. INTRODUCTION

Modern banking theory focuses on the key role that banks play by creating deposit accounts that provide households with safe liquid investment opportunities (Diamond and Dybvig (1983), Gorton and Pennacchi (1990), Diamond and Rajan (2000, 2001)). In doing this, banks follow a maturity-transformation strategy in which they invest in longer-term assets while issuing shorter-term deposits. This maturity-transformation strategy, however, can expose banks to substantial interest rate risk because of the inherent duration mismatch (Diamond and Dybvig (1983), Goldstein and Pauzner (2005), Segura and Suarez (2017), Hoffman, Langfield, Pierobon, and Vuillemey (2019), Drechsler, Savov, and Schnabl (2021)).

The literature considers several approaches that banks may follow in addressing their maturity-transformation risk. The first is the traditional view that banks simply accept the inherent interest rate risk and are compensated for their exposure by the term premium (via the net interest margin).[1] A second approach is that banks may choose to hedge their maturity-transformation risk, either directly or indirectly. For example, banks can hedge their risk directly using interest rate swaps or other types of fixed income derivatives.[2] Alternatively, Drechsler, Savov, and Schnabl (2021) discuss how banks can create deposit franchises giving them the market power to pay deposit rates that are relatively insensitive to changes in market rates. This means that the deposit franchise essentially functions as a synthetic interest rate swap in converting the bank's deposits into the equivalent of long-term fixed rate debt. Thus, the deposit franchise may serve as an indirect hedge for the bank's maturity-transformation risk.

In this paper, we consider a third approach that banks could use to address the maturity-transformation issue. In particular, banks could choose to actively price their maturity-transformation risk by offering proportionally lower rates

[1]This perspective is described in Borio, Gambacorta, and Hofmann (2017), Di Tella and Kurlat (2021), Drechsler, Savov, and Schnabl (2021), Paul (2023), Minoiu, Schneider, and Wei (2023), and many others.

[2]Diamond (1984), Gorton and Rosen (1995), Purnanandam (2007), Begenau, Piazzesi, and Schneider (2015), Hoffman, Langfield, Pierobon, and Vuillemey (2019), Vuillemey (2019), McPhail, Schnabl, and Tuckman (2023), and others discuss the use of derivatives by banks.

on deposit accounts requiring more maturity transformation. The rationale for doing this is that in taking on maturity-transformation risk, banks may face significant additional costs such as higher regulatory capital costs, risk overhang costs, equity issuance costs, risk-related agency costs, etc.[3] Since these costs are likely directly related to the amount of maturity-transformation risk banks absorb, pricing this risk into the rates banks offer would parallel the approach used by other financial institutions in pricing their balance-sheet-related costs into the quotes they provide for securities and derivative contracts.[4] The spreads between fair market rates and the deposit rates offered by banks could be viewed as an implicit type of seigniorage compensating them for the maturity-transformation costs of creating money-like financial products.

There are several possible mechanisms that could allow banks to price their maturity-transformation risk. One would be if banks had some form of blanket market power over depositors, but were then to choose to exercise that market power in this maturity-specific way. A more plausible way is suggested by the rapidly-growing literature on the convenience premia associated with safe assets. If households value the unique liquidity/convenience features of bank deposits, they can induce banks to offer these accounts by compensating them for the costs/inconvenience they face because of the maturity-transformation process. In equilibrium, deposit spreads would adjust to reflect the maturity-transformation costs of banks and could also be interpreted as a type of convenience premium similar to that described in DeMarzo, Krishnamurthy, and Nagel (2024).

To examine whether maturity-transformation risk is priced into bank deposit rates, we extend the literature in a new direction by making use of the term structure of rates for bank certificates of deposit (bank CDs). In particular, we estimate the cost (or spread) that CD investors pay for liquidity for every tenor

[3]For example, the Net Stable Funding Ratio, the Supplementary Leverage Ratio, and the Liquidity Coverage Ratio of the Basel III framework can significantly impact the capital requirements of banks with maturity-transformation risk. As another example, Bolton, Li, Wang, and Yang (2023) argue that offering deposit accounts exposes banks to costly equity issuance since banks cannot perfectly control deposit flows. As a third example, DeYoung, Gron, Torna, and Winton (2015) describe how the risk overhang associated with their long-term illiquid assets may result in risk-constrained banks having to forego more-profitable lending opportunities (see also Gron and Winton (2001)).

[4]For recent evidence on the pricing of balance-sheet costs, see Duffie (2016), Anderson, Duffie, and Song (2019), Fleckenstein and Longstaff (2020), Lewis, Longstaff, and Petrasek (2021), van Binsbergen, Diamond, and Grotteria (2022), and Du, Hébert, and Li (2023).

across the term structure. The variation in these spreads over time and across maturities provides a natural way of identifying the relation between deposit pricing and maturity-transformation risk.

In measuring the cost of liquidity for bank CDs, however, we cannot simply use the conventional deposit spread based on the difference between CD rates and matched-maturity riskless rates such as Treasury rates. Unlike Treasury securities, CDs provide depositors a continuously-exercisable option to put the CD back to the bank at its accrued value minus an early withdrawal penalty. This option can be very valuable, especially when interest rates (and, therefore, early withdrawal penalties) are relatively low. Accordingly, we use a standard no-arbitrage valuation framework to solve for the option-adjusted deposit spread which more accurately measures the actual cost of liquidity to CD investors. We note that the value of the embedded option is also impacted by the early withdrawal strategies followed by households. This parallels how household prepayment behavior impacts the pricing of mortgage-backed securities. In valuing the option, we take into account that households may be subject to liquidity shocks, causing them to withdraw early, and that they may forego early withdrawal opportunities as a result of being inattentive.

The estimation results indicate that bank depositors pay a significant cost for the liquidity provided by bank CDs. Using weekly data from 2001 to 2023, we find that the average option-adjusted deposit spreads range from 39.54 basis points for six-month CDs to 7.37 basis points for five-year CDs. These average spreads, however, are smaller than those associated with short-term deposits such as checking and savings accounts. We contrast the estimated spreads with several measures of the convenience premia in longer-term Treasury bonds discussed in the literature. We find that there are significant correlations between the spreads and these measures of Treasury convenience premia. This lends support for interpreting these deposit spreads as a type of convenience premium. We note, however, that the correlations are negative in sign, suggesting that households may not view bank deposits and Treasuries as perfect substitutes in providing liquidity/convenience services.[5]

Having estimates of the option-adjusted deposit spreads now allows us to examine directly the relation between deposit pricing and maturity-transformation risk. To measure the maturity-transformation risk banks face in creating a deposit account, we use the difference between the average maturity of the assets held by banks and the tenor of the CD which we refer to as the maturity mis-

[5]These results are consistent with those in Krishnamurthy and Li (2023) and Acharya and Laarits (2024) implying that the convenience premia for imperfect substitutes could move in opposite directions.

match. We explore the relation between deposit pricing and maturity mismatching in three different ways.

First, we examine the time-series relation between the option-adjusted deposit spreads and maturity mismatch for each tenor along the term structure. The results indicate that there is a strong correlation between the spreads and maturity mismatch over time. In particular, the correlation between the spreads and maturity mismatch ranges from about 65 to 70 percent across the different tenors on the term structure. The positive relation between the spreads and maturity mismatch is significant both in levels and in differences.

Second, we use a panel regression framework to identify the relation between the option-adjusted deposit spreads and maturity mismatch from the cross-sectional information in the term structure. The results confirm that there is a strong positive relation between the spreads and maturity mismatch, which is significant from both a statistical and an economic perspective. In particular, the regression estimates imply that increasing the maturity mismatch by one year maps into a 5.45 basis-point increase in the spread. This suggests that the spread compensating banks for maturity-transformation risk could represent a substantial portion of their net interest margin, particularly for deposits with shorter maturities.

Third, we make use of a natural experiment made possible by the implementation of the Net Stable Funding Ratio (NSFR) capital requirement on banks on July 1, 2021. This capital requirement imposes significant capital costs on banks with asset/funding mismatches. Using a panel regression framework, we find evidence that the pricing of maturity-transformation risk increased significantly following the implementation of the NSFR. In particular, the impact on the spread of an increase in the maturity mismatch by one year more than triples following the NSFR implementation.

Taken together, these results make a strong case that maturity-transformation risk is priced into the CD rates that banks offer their customers. These results are also consistent with the view that households value the unique liquidity/convenience services that bank deposit accounts provide and are willing to accept lower interest rates in order to gain access to them.

These results also have implications for several current issues in the banking literature. First, there is an ongoing debate about how much of the interest rate risk faced by banks is hedged by their deposit franchise (Haddad and Sraer (2020), Drechsler, Savov, and Schnabl (2021), Drechsler, Savov, Schnabl, and Wang (2023), Begenau and Stafford (2023), Emin, James, and Li (2023)). Finding that banks price their maturity-transformation risk suggests that banks may not view their interest risk as fully hedged. Second, there is a rapidly-growing literature about the value of the deposit franchise itself (Drechsler, Savov, and Schnabl (2021), Begenau and Stafford (2019), Bolton, Li, Wang, and Yang (2023)). Our results suggest that if deposit spreads simply offset the costs of maturity-transformation activity, then their net effect on the value of the deposit franchise may be modest. Finally, a number of recent papers focus on the role of banking market power on deposit pricing (Drechsler, Savov, and Schnabl (2017), Begenau and Stafford (2021, 2023), Wang, Whited, Wu, and Xiao (2022)). Bank deposit accounts, however, have unique liquidity features not shared by other types of investments. For example, bank deposits can be converted into cash without having to execute a transaction in a secondary market, which is not the case even for Treasury securities. Our results imply that household demand for the special characteristics of bank deposits could provide an alternative non-market-power-based explanation for the existence of deposit spreads (Krishnamurthy and Li (2022), d'Avernas, Eisfeldt, Huang, Stanton, and Wallace (2023)).

Related Literature

This paper is directly related to the extensive literature on bank deposit pricing. Two key findings in this literature are that bank deposit rates are lower than comparable market interest rates, and that banks are slow to adjust deposit rates when market interest rates increase. Examples of this literature include Diebold and Sharpe (1990), Neumark and Sharpe (1992), Hutchison (1995), Driscoll and Judson (2013), and Yankow (2023).

One strand of this literature explains these findings by focusing on the role of banking market power in setting deposit rates. Important recent examples include Hutchison and Pennacchi (1996), Drechsler, Savov, and Schnabl (2017, 2021), Begenau and Stafford (2021, 2023), Wang, Whited, Wu, and Xiao (2022), and Whited, Wu, and Xiao (2021). Other recent papers focus on institutional differences at the bank level and characteristics of local deposit markets. These include Heitfield and Prager (2004), Park and Pennacchi (2008), Egan, Hortaçsu, and Matvos (2017), Haendler (2022), Kundu, Park, and Vats (2021), d'Avernas, Eisfeldt, Huang, Stanton, and Wallace (2023), Koont (2023), and Koont, Santos, and Zingales (2023).

This paper extends this literature by studying the pricing of maturity-transformation risk. In doing this, our approach differs from earlier work in the area in several important ways. First, rather than focusing on the pricing of individual short-term deposit products, we make us of the entire term structure of CD rates. Second, we use an option pricing framework to account for the continuously-exercisable put option feature that CDs provide depositors.

This paper is also related to the growing literature on the valuation of safe assets. A key theme in this literature is that investors are often willing to pay a convenience premium above and beyond the present value of an asset's explicit cash flows for characteristics such as safety and liquidity. Examples of this literature documenting convenience premia in Treasury securities include Longstaff (2004), Krishnamurthy and Vissing-Jørgensen (2012), Nagel (2016), Du, Im, and Schreger (2018), Fleckenstein and Longstaff (2020b), Lewis, Longstaff, and Petrasek (2021), van Binsbergen, Diamond, and Grotteria (2022), He, Nagel, and Song (2022), Fleckenstein and Longstaff (2024), and Acharya and Laarits (2024). Our results suggest that investors may also be willing to pay a premium for the unique convenience features that bank deposits provide by accepting below-market interest rates.

2. BANK CDs

Bank certificates of deposit (CDs) are savings certificates where the principal amount deposited is held in a bank account for a fixed period of time. The term of these time deposits typically ranges from one month to five years or more. The holder of a CD accrues interest at a specified fixed rate (the CD rate) and receives a single cash flow in the amount of the principal plus accrued interest at maturity. Bank CDs effectively have the same credit risk as Treasury bonds since they are guaranteed by the Federal Deposit Insurance Corporation (FDIC) (up to a specified limit). CDs represent an important investment class for U.S. households. For example, the total amount of CDs (with notional amounts of $100,000 or less) was more than $900 billion as of the end of July 2023.

A unique feature of bank CDs that distinguishes them from conventional fixed income instruments such as Treasury bonds is that the depositor has the right to redeem a CD at any time prior to its stated maturity at its accrued value minus a withdrawal penalty. We refer to this feature as the early withdrawal option and note that it is similar in nature to a put option on a bond. The early withdrawal penalty is typically assessed in terms of a certain number of days of interest.[6] To illustrate, suppose that a depositor owns a $10,000 CD with a one-year term and a CD rate of 4.00 percent. Assume that the early withdrawal penalty is three months of interest. This means that if the depositor redeems the $10,000 anytime before the one-year term is over, the penalty would be $3/12 \times 0.0400 \times \$10,000 = \$100$. Furthermore, if the depositor were to redeem the CD early at any time during the first three months of the one-year term, the

[6]Investors can redeem a CD at any time without prior notice. Federal regulations require a minimum withdrawal penalty of seven days of simple interest on early withdrawals during the first six days after investing in a CD. There is no rule limiting the maximum withdrawal penalties.

depositor would actually incur a loss of principal.

Our focus in this paper is exclusively on the plain-vanilla type of FDIC-insured bank CDs described above. We observe, however, that there are other types of investments in the markets that are often referred to as CDs. These include brokered CDs and negotiable CDs. These types of investments, however, are fundamentally different in nature from bank CDs. For example, these CDs have no early redemption feature since holders can sell them in the secondary market at any time. Furthermore, these investments may be callable, or may not be covered by FDIC insurance. Accordingly, we exclude these types of CDs from the analysis.

3. THE VALUATION MODEL

To estimate the option-adjusted deposit spreads in bank CD rates, we compare market CD rates to the fair market rates implied by a no-arbitrage fixed-income modeling framework that values bank CDs with embedded optionality. This framework allows for the possibility that households may experience exogenous liquidity shocks and need to make early withdrawals, or that households may be inattentive and fail to exercise the early withdrawal option in a timely manner.

3.1 The No-Arbitrage Fixed-Income Valuation Framework

As the underlying valuation framework, we use the standard Black, Derman, and Toy (BDT 1990) model throughout the analysis. The BDT model is a popular one-factor no-arbitrage model discussed in many fixed-income textbooks and is widely used by practitioners. The BDT model is implemented using a simple binomial-tree setup to specify the dynamics of the short-term riskless interest rate. The model is calibrated such that it matches exactly the term structure of riskless rates for maturities of up to five years and the market prices of at-the-money interest rate caps with maturities of one, two, three, and five years. The calibration process is documented in the Internet Appendix.

The binomial-tree structure of the BDT model allows us to incorporate the actual early withdrawal behavior of households into the valuation process. To illustrate this, it is useful to first introduce some notation. Let $r_{i,t}$ denote the value of the riskless rate at node i and time t on the BDT binomial tree, where the nodes $i = 1, 2, 3, \ldots, t + 1$ are ordered from the highest to the lowest values. Recall that the binomial tree begins with the current value of the riskless rate at time zero. After one time step, the binomial process results in two possible values for the riskless rate which we denote as $r_{1,1}$ and $r_{2,1}$. After two time steps, the riskless rate now has three possible values, which are designated as $r_{1,2}$, $r_{2,2}$, and $r_{3,2}$, and so forth. We continue the process until the binomial tree is extended out over 60 monthly time steps.

Let $P_{i,t}$ denote the value of the bank CD at node i and time t. Let $\text{Exer}_{i,t}$ denote the cash flow received by the household if the early withdrawal option is exercised at node i and time t. This cash flow is simply the principal value of the CD plus accrued interest, minus the early withdrawal penalty,

$$\text{Exer}_{i,t} = \text{Principal} + \text{Accrued}_\text{t} - \text{Penalty}. \tag{1}$$

Let $\text{Cont}_{i,t}$ denote the value of the CD if the household does not exercise the early withdrawal option at node i and time t, and continues to hold the CD until time $t+1$. This continuation value is given by simply present-valuing the two subsequent values of the CD on the binomial tree at time $t+1$ back to time t,

$$\text{Cont}_{i,t} = \frac{\frac{1}{2}\left(P_{i,t+1} \ + \ P_{i+1,t+1}\right)}{1 + r_{i,t}\ \Delta t}, \tag{2}$$

where Δt is the time step.

In theory, an optimizing household would make an early withdrawal decision by comparing the immediate exercise value $\text{Exer}_{i,t}$ with the continuation value $\text{Cont}_{i,t}$ and taking the action that maximizes the value of the CD at that node, which would imply

$$P_{i,t} = \max(\text{Exer}_{i,t},\ \text{Cont}_{i,t}). \tag{3}$$

In reality, however, a household might face a liquidity shock and be forced to make an early withdrawal. In this case, the value of the CD at that node would simply be

$$P_{i,t} = \text{Exer}_{i,t} \leq \max(\text{Exer}_{i,t},\ \text{Cont}_{i,t}). \tag{4}$$

Similarly, if the household is inattentive, the household passively continues to hold the CD even if it might be optimal to make an early withdrawal. In this case, the value of the CD at that node becomes

$$P_{i,t} = \text{Cont}_{i,t} \leq \max(\text{Exer}_{i,t},\ \text{Cont}_{i,t}). \tag{5}$$

Thus, to value bank CDs based on the actual early withdrawal behavior of households, we need to specify a model of the actions (or inactions) followed by the household at each mode. The value of the CD is then obtained by discounting the resulting cash flows backwards through the binomial tree in the usual way.

3.2 Modeling Liquidity-Based Withdrawals

Bank deposit accounts can represent an important source of liquidity to households in adverse states of the world. Accordingly, we allow for the possibility that a household may face an exogenous shock that requires it to liquidate a CD by making an early withdrawal. In doing this, we assume that liquidity shocks are triggered by the realization of a Poisson process. Let λ denote the arrival probability of a Poisson event over the next time step. If the Poisson event occurs at time step t, then the household makes an early withdrawal, and the value of the CD at each node at time t is $\text{Exer}_{i,t}$, which is simply the principal value of the CD plus accrued interest minus the early withdrawal penalty. Note that the possibility of an exogenous liquidity-based withdrawal has the effect of reducing the value of a CD (holding all else fixed). Intuitively, this is because the liquidity shock forces the household to make an early withdrawal even if exercising the implicit put option is not optimal, or worse, when the put option is actually out-of-the-money.

3.3 Modeling Household Inattention

There is an extensive literature documenting that individual investors may not participate continuously in financial markets. This might occur if there are ongoing informational or search costs associated with being present in a market, or if investors lack financial sophistication. A well-known example of this is the tendency of some homeowners to delay refinancing fixed-rate mortgages when mortgage rates decline. This tendency directly impacts prepayment behavior, which in turn has major implications for the valuation of mortgage-backed securities. To allow for the possibility of inattentive household behavior in the bank CD market, we use the following model. Let γ denote the probability that a household is attentive to the early withdrawal decision at time t. To keep things simple, we assume that γ is constant across nodes and times. If the household is attentive at node i at time t, then the household compares the value of exercising the early withdrawal option $\text{Exer}_{i,t}$ with the continuation value of not exercising the option $\text{Cont}_{i,t}$. If the value of immediate exercise is greater, then the household makes an early withdrawal. If the value of continuation is greater, then the household does not make an early withdrawal. In contrast, if the household is not attentive, it simply continues to hold the CD by default and does not make an early withdrawal even if it is optimal to do so.

3.4 The Model Specification

Given the assumptions about liquidity-based withdrawals and household attentiveness, we can now specify the value of the CD at each binomial node. Taking expectations over Poisson arrivals and household attentiveness, the value of a bank CD at node i and time t can be represented as

$$\begin{aligned} P_{i,t} = \; & \lambda \; \text{Exer}_{i,t} \\ & + \; (1-\gamma)\,(1-\lambda)\;\text{Cont}_{i,t} \\ & + \; \gamma\,(1-\lambda)\;\max(\text{Exer}_{i,t},\;\text{Cont}_{i,t}). \end{aligned} \tag{6}$$

The first term in this expression is the cash flow resulting from a liquidity-based withdrawal times the probability that a liquidity event occurs. The second term is the value of passively continuing to hold the CD times the joint probability that the household is inattentive and does not experience a liquidity shock. The third term is the value of making an optimal early withdrawal decision times the joint probability that the household is attentive and does not experience a liquidity shock.

4. THE DATA

This section provides a brief description of the primary data sets used throughout the paper. The Internet Appendix provides complete details about the data and methodology used in the analysis.

4.1 Bank CD Rates

S&P RateWatch collects weekly branch-level data on interest rates from over 96,000 branch locations in the U.S. for a wide variety of products such as checking, savings, and money market accounts, and CDs. We obtain weekly data on CD rates from S&P RateWatch for the period from January 5, 2001 to June 30, 2023 for six-month, one-year, two-year, three-year, four-year, and five-year tenors. To ensure that the rates are for CDs that are fully insured by the FDIC, we restrict the sample to CDs for account sizes less than or equal to $100,000.

As discussed, previous research on deposit pricing typically focuses on rates at the individual branch or bank level and studies their cross-sectional variation over geographical location, bank size, deposit betas, etc. In contrast, this paper takes the novel approach of focusing specifically on the term structure of CD rates. Accordingly, our analysis will be done at the aggregate market level by taking weekly averages of CD rates by tenor across all observations. On average, about 6,400 branches/banks provide quotes each week for individual CD tenors. Table 1 provides summary statistics for these average CD rates by tenor. Table 1 also reports the average values of the corresponding matched-maturity Treasury spot rates which are higher than the average CD rates by roughly 20 to 35 basis points. Recall, however, that the direct comparison of CD rates to riskless rates is misleading since it does not control for the early withdrawal option.

4.2 Early Withdrawal Penalties

We obtain data on early withdrawal penalties for CDs from two sources. First, we collect quarterly interest rate risk exposure reports from the Office of Thrift Supervision (now merged with the Comptroller of the Currency) for the period from Q1 2001 to Q4 2011 and compute annual averages of the reported withdrawal penalties for CDs with original maturity T for the categories $T \leq 12$ months, $12 < T \leq 36$ months, and $T > 36$ months.[7] Second, we collect early withdrawal penalties from the RateWatch database for the period from January 2, 2012 to June 30, 2023 and compute annual averages across all observations for individual CD tenors of up to 60 months.

By combining the OTS and RateWatch series, we obtain an annual time series of early withdrawal penalties measured in terms of days of foregone interest for the period from 2001 to 2023. We then compute weekly measures of the cost of making an early withdrawal by multiplying the annual averages by the weekly CD rate for each tenor. Table 2 provides summary statistics for the withdrawal penalties, both in terms of the number of days of foregone interest, as well as a percentage of par value.

4.3 Valuation Framework Data

The valuation framework for bank CDs requires having a calibrated BDT binomial tree for each valuation date, as well as estimates of the parameters λ and γ used in modeling the early withdrawal behavior of households.

4.3.1 Riskless discount factors

A key input for the BDT model is the vector of prices for zero-coupon riskless bonds (discount factors) with maturities ranging from 1 to 60 months. To obtain

[7] See https://www.occ.gov/news-events/newsroom/news-issuances-by-year/ots-issuances/ots-aggregate-irr-exposure-and-cmr-reports.html.

these riskless discount factors, we follow the approach used in Fleckenstein and Longstaff (2024). This approach treats the repo overnight index swap (OIS) curve as the riskless term structure. We use a standard bootstrapping algorithm to obtain monthly riskless zero-coupon bond prices from the repo OIS curve for the period from January 5, 2001 to June 30, 2023. As a later robustness exercise, we also use zero-coupon bond prices bootstrapped from the Treasury constant maturity (CMT) term structure provided by the Federal Reserve.

4.3.2 Interest rate caps

Constructing the BDT binomial tree also requires specifying the volatility of the binomial process at each time step. In doing this, we use market quotes for the implied volatilities of at-the-money one-year, two-year, three-year, and five-year interest rate caps for the period from January 5, 2001 to June 30, 2023. The data are obtained from the Bloomberg system.

4.3.3 Early withdrawal activity

Obtaining data on early withdrawal activity is difficult since it is not publicly reported. Fortunately, we were able to find a source of summary information about early withdrawals covering a substantial portion of the sample period. The Office of Thrift Supervision published a quarterly interest rate risk exposure report from 1998 to 2011 that includes data on the amount of early withdrawals for CDs with remaining maturity t for the categories $0 < t \leq 3$ months, $3 < t \leq 12$ months, $12 < t \leq 36$ months, and $t > 36$ months. This source provides us with a time series of 39 quarterly early withdrawal rates for each of these four maturity categories.

Table 3 reports summary statistics for the early withdrawal rates by maturity category. As shown, the average early withdrawal rate ranges from 4.944 percent for the shortest maturity to 6.128 percent for the longest. The early withdrawal rates, however, vary significantly through time, ranging from about three to ten percent or more. This evidence suggests that early withdrawal is a significant factor in the CD market and may play a major role in determining CD rates in competitive markets. These estimates are also broadly consistent with previous research. Artavanis, Paravisini, Robles-Garcia, Seru, and Tsoutsoura (2022) estimate an early withdrawal rate of roughly ten percent per year for a sample of short-term time deposits in a major Greek bank during the year 2014. These estimates are also consistent with early withdrawal rates from tax-deferred retirement accounts such as IRAs and 401(k)s. Argento, Bryant, and Sabelhaus (2015) find that roughly five percent of taxpayers under the age of 55 received tax-penalized early distributions from retirement accounts each year during their

2004–2010 sample period. Amromin and Smith (2003) report similar percentages for an earlier sample period.

5. MODEL CALIBRATION

This section provides a brief overview of how the valuation model is calibrated. The Internet Appendix provides a complete discussion of the calibration methodology.

5.1 Calibrating the BDT Binomial Tree

The algorithm for calibrating a BDT binomial tree for the short-term interest rate is well documented in standard textbooks such as Hull (2021). To calibrate the binomial tree using monthly time steps out to five years, we need as inputs a vector of 60 discount factors and a vector of 60 volatilities for the short rate. As discussed above, the vector of discount factors is obtained by bootstrapping the riskless curve. To calibrate the volatility function in a parsimonious way, we make the simplifying assumption that it is piecewise constant with value σ_1 for horizons up to one year, σ_2 for horizons between one and two years, σ_3 for horizons between two and three years, and σ_4 for longer horizons. To identify these volatilities, we use a numerical search procedure and iterate the BDT calibration algorithm until the resulting binomial tree is able to exactly match the vector of discount factors and the market prices of one-year, two-year, three-year, and five-year at-the-money interest rate caps.

5.2 Calibrating the Early Withdrawal Model

To fully specify the model, we need the parameter λ that determines the frequency at which exogenous liquidity shocks occur, as well as the parameter γ representing the probability that the household is attentive. In doing this, we make use of the early withdrawal activity data from the Office of Thrift Supervision described above.

First, we make the identifying assumption that early withdrawals for CDs with remaining maturities of three months or less are driven solely by liquidity shocks. This assumption is a safe one since the probability of finding it optimal to exercise the early withdrawal option for a CD with such a short remaining maturity is infinitesimal.

Next, we use the following regression approach to decompose the observed early withdrawal rate for longer-maturity CDs into a liquidity-based component and a strategic interest-rate-related component. Specifically, we regress changes in the early withdrawal rate for CDs with maturities of one year or more on changes in the early withdrawal rate for CDs with a maturity of less than or equal to three months. Since the early withdrawal rate for CDs with a maturity of three months or less is due solely to liquidity shocks, this measure serves as an exogenous instrument for liquidity-based shocks. The regression specification is

$$\Delta \text{Withdrawal}_{\text{LT,t}} = c_0 + c_1 \, \Delta \text{Withdrawal}_{\text{ST,t}} + \epsilon_t, \tag{7}$$

where $\Delta \text{Withdrawal}_{\text{LT,t}}$ and $\Delta \text{Withdrawal}_{\text{ST,t}}$ denote the quarterly change in the withdrawal rates for the longer-term and shorter-term CDs, respectively. Table 4 reports the results from this regression.

As shown, the early withdrawal rate for the short-term CDs is significantly positively related to the early withdrawal rate for the longer-maturity CDs. This implies that the early withdrawal rates for longer-term CDs contain a liquidity-based component. We note, however, that the slope coefficient for the short-term early withdrawal rate is 0.2749 which is much less than one. This result makes intuitive sense since when a liquidity shock occurs and households need to liquidate some of their CDs, they have strong incentives to liquidate the CDs with the lowest early withdrawal penalties first. These are typically the shortest-maturity CDs. Because of this, we would expect the liquidity-related early withdrawal rate for longer-term CDs to be significantly lower than that for short-term CDs.

Given these regression results, we adopt the following parsimonious approach to calibrate the parameters λ and γ of the early withdrawal model. In the context of this model, λ represents the probability that longer-term CDs experience a liquidity-based early redemption. As the point estimate of λ, we simply multiply the 4.944 percent average value of the early withdrawal rate for CDs with maturities of three months or less by the 0.2749 slope coefficient from the regression, giving a value of 1.359 percent (on an annualized basis). The remainder of the average early withdrawal rate of longer-maturity CDs can now be directly attributed to the strategic exercise of the early withdrawal option.[8]

To solve for the value of γ, we note that the average early withdrawal rate for the longest-tenor category is 6.128 percent. Applying this average rate to a five-year CD implies that the probability of an early withdrawal during the

[8]This interpretation is consistent with the fact that the difference between the early withdrawal rates for short-term and longer-term CDs is highly correlated with the level of interest rates. For example, the correlation of this difference with the three-year Treasury rate is nearly 60 percent.

lifetime of the CD is 30.640 percent. Substituting the value of 1.359 percent for λ into the valuation framework and then numerically searching for the value of γ that results in the probability of an early withdrawal during the lifetime of a five-year CD being 30.640 percent results in a point estimate for γ of 1.825 percent.

6. OPTION-ADJUSTED DEPOSIT SPREADS

Using the calibrated model, we can now obtain weekly estimates of the option-adjusted deposit spreads incorporated into bank CD rates. The option-adjusted deposit spread is measured as the difference between the CD rate implied by the calibrated model and the actual market CD rate.

6.1 Summary Statistics

Table 5 presents summary statistics for the option-adjusted deposit spreads by tenor for the CDs. The average values of the spreads are positive for each of the individual tenors and range from a high of 39.54 basis points for the six-month tenor to 7.37 basis points for the five-year tenor. The t-statistics (based on Newey-West estimates of the standard errors) are significant for all but the five-year tenor. These results indicate that bank depositors face substantial liquidity costs when investing in time deposits. This is consistent with current banking theory which implies that households may be willing to accept lower rates on bank deposits in exchange for the safety and liquidity/convenience these types of accounts provide.

To provide perspective, Figure 1 plots the time series of the model-implied CD rates and the market CD rates for each tenor. As shown, the market and model-implied CD rates generally track each other closely. In fact, market CD rates resemble a smoothed version of the model-implied CD rates. The largest divergences occur around the times when the model-implied rates attain their lowest or highest values.

Figure 2 plots the time series of the option-adjusted deposit spreads for each CD tenor. As illustrated, there is considerable variation in the spread estimates over time with a range that typically exceeds 500 basis points over the sample period. The highest values tend to occur during the 2022–2023 period during which the Federal Reserve increased rates dramatically in an effort to address inflation concerns. One interesting aspect of the spreads is that they frequently take on negative values. Furthermore, these negative values can persist over extended multi-year horizons. In particular, the spreads for all tenors are negative during much of the 2001–2003 period. Similarly, the spreads are generally negative throughout much of the 2007–2012 financial-crisis and post-crisis periods. Finally, the spreads again become generally negative during the early stages of the Covid-19 pandemic.[9]

Figure 2 also shows that the spreads tend to move together. The pairwise correlations across tenors in the weekly changes of the spreads range from about 88 percent to more than 99 percent. Furthermore, a simple principal components analysis shows that the spreads are driven primarily by a single common factor that explains nearly 97 percent of the variation in the estimates.

6.2 Comparison to Other Deposit Spreads

It is also useful to contrast the option-adjusted deposit spreads for bank CDs with the corresponding spreads for short-term bank deposits. Using the same RateWatch data, methodology, and sample period as for the CDs, we obtain weekly averages of the rates for checking, savings, and money market accounts. Given the short-term nature of these accounts, we do not attempt to model any implicit optionality and simply estimate their deposit spreads as the difference between the one-month riskless rate and the average rates for these accounts.

Table 5 reports summary statistics for the checking, savings, and money market deposit spreads. As shown, the average spreads for these accounts are much larger than those for the CDs. In particular, the average spreads for checking, savings, and money market accounts are 122.92, 109.08, and 87.50 basis points, respectively. These estimates are consistent with those from the prior literature.

6.3 Comparison to Treasury Convenience Premia

As discussed, much of the previous literature points to banking market power as the underlying reason for the deposit spreads we observe in the market. It is important to recognize, however, that there may be other potential explanations for the existence of deposit spreads. For example, households may be willing to accept below-market rates on deposits to induce banks to offer these types of accounts. If so, then deposit spreads could be similar in nature to the convenience premia in Treasury markets.

To explore this possibility, we compare the option-adjusted deposit spreads with several widely-cited measures of the convenience premia associated with longer-term Treasury securities. These measures are the AAA-Treasury spread used in Krishnamurthy and Vissing-Jørgensen (2012), the Treasury richness mea-

[9]This pattern of negative deposit spreads is consistent with the evidence of periodic negative convenience premia in Treasury markets presented in He, Nagel, and Song (2022) and Fleckenstein and Longstaff (2024).

sure used in Fleckenstein and Longstaff (2024), and the Refcorp-Treasury spread used in Longstaff (2004).

To provide a general overview, Figure 3 presents scatterdiagrams of the option-adjusted deposit spreads and convenience premia measures for some selected tenors. As shown, there is a a substantial amount of correlation between the spreads and the premia. The correlations are particularly strong with the AAA-Treasury spread, but are also significant for the other two measures. These correlations lend support to the view that the deposit spreads may represent a type of convenience premium similar to those in the Treasury market.

One striking aspect of the scatterdiagrams shown in Figure 3 is that the correlations tend to be negative in sign. This implies that deposit spreads tend to be larger when Treasury convenience premia are smaller, and vice versa. This is particularly true during periods when Treasury convenience premia are actually negative. This pattern could be consistent with a scenario in which the asset that households view as the most convenient switches back and forth between deposits and Treasuries. This is also consistent with the recent model presented in Krishnamurthy and Li (2022) in which deposits and Treasuries are imperfect substitutes. Similarly, these results are consistent with the implications in Acharya and Laarits (2024) that the convenience yields for assets that are imperfect substitutes could move in opposite directions. Since our primary focus in this paper is on the pricing of maturity-transformation risk, we leave a more-detailed analysis of the relation between deposit spreads and Treasury convenience premia to future research. These results, however, raise intriguing questions about the role that bank deposits may play as the ultimate "liquidity" reserve asset in the financial markets.

7. IS MATURITY-TRANSFORMATION RISK PRICED?

An important advantage of the term structure of CD rates is that it provides us natural ways of exploring the relation between maturity-transformation risk and deposit pricing. If the option-adjusted deposit spread compensates banks for the costs or inconvenience of creating safe/liquid assets for households, then there should be a direct relation between spreads and measures of maturity-transformation risk.

7.1 Measuring Maturity-Transformation Risk

To measure maturity-transformation risk, we begin by collecting data on bank balance sheets from the quarterly Reports of Condition and Income (Call Reports), which we obtain from Wharton Research Data Services (WRDS).[10] We follow English, Van den Heuvel, and Zakrajšek (2018) and Drechsler, Savov, and Schnabl (2021) and calculate the repricing maturity of bank assets at the individual bank level. We then take a simple cross-sectional average of the bank-level asset durations each quarter, and then take the difference between these quarterly averages and the individual CD tenors to estimate the maturity mismatch of a given tenor.

7.2 Time-Series Tests

A direct way of exploring the relation between the option-adjusted deposit spreads and maturity mismatch is by comparing their time-series behavior. Figure 4 plots the time series of the spreads by tenor and the corresponding maturity mismatch. As shown in Figure 4, there is a strong relation between the deposit spreads and the maturity mismatch over time. In particular, both series tend to increase and decrease at about the same time. The simple correlations between the spreads and the maturity mismatch are all on the order of 70 percent.

To examine the relation more formally, we regress the option-adjusted deposit spreads for individual tenors on the corresponding maturity mismatch. We estimate this regression both in levels and in quarterly changes. Table 6 reports the regression results. Panel A presents the results based on the level of the spreads. The regression specification is

$$\text{Spread}_t = c_0 + c_1 \text{ Mismatch}_t + \epsilon_t, \tag{8}$$

where Spread_t and Mismatch_t denote the option-adjusted deposit spread and the maturity mismatch for the indicated tenor at date t, respectively. As shown, there is a strong positive relation between the spread and the maturity mismatch for all six tenors. The slope coefficients range from a high of 1.931 for the six-month tenor to a low of 1.139 for the five-year tenor. This implies that an increase in the average maturity mismatch of one year translates into an increase in the spread of roughly one to two basis points. The adjusted R^2s indicate that nearly half of the variation in spreads can be explained in terms of the variation in the maturity mismatch. As another way of illustrating the results, Figure 5 plots the fitted regression line on the scatterdiagram of the spreads and maturity mismatch for each tenor.

Panel B in Table 6 presents the results based on quarterly changes in the

[10] Call reports offer information on balance sheet and income statement items for the majority of FDIC-insured institutions.

option-adjusted deposit spreads. The corresponding regression specification is

$$\Delta\,\text{Spread}_t = c_0 + c_1\,\Delta\,\text{Mismatch}_t + \epsilon_t. \tag{9}$$

As shown, the changes in the spreads are significantly positively related (at either the five- or ten-percent level) to changes in the maturity mismatch for five of the six tenors (the six-month tenor is the only exception). These results corroborate those shown in Panel A and in Figures 4 and 5.

7.3 Cross-Sectional Tests

The term structure provides us a natural experiment for testing whether there is a relation between option-adjusted deposit spreads and maturity mismatch. Specifically, we use a panel regression framework to test whether maturity-transformation risk is priced in the cross-section.

In this panel framework, we pool the data for all tenors and regress the option-adjusted deposit spreads on the corresponding maturity mismatch. To control for the variation in the level of the spreads over time, we include fixed effects for each quarter. This ensures that the regression captures the pure cross-sectional relation between spreads and maturity mismatch. The regression specification is

$$\text{Spread}_{i,t} = c_1\ \text{Mismatch}_{i,t} + \text{FE}_t + \epsilon_{i,t}, \tag{10}$$

where $\text{Spread}_{i,t}$ and $\text{Mismatch}_{i,t}$ denote the option-adjusted deposit spread and the maturity mismatch for the ith tenor at date t, respectively, and FE_t denotes the quarterly fixed effects. Table 7 reports the results from the panel regression (robust standard errors are clustered by tenor). As shown, there is a highly significant positive relation between the option-adjusted deposit spreads and the corresponding maturity mismatch. The t-statistic for the maturity mismatch is 3.74. The relation is also significant in economic terms. In particular, the slope of the coefficient implies that the spread increases by 5.45 basis points as the maturity mismatch increases by one year. Thus, a maturity mismatch of five years would map into an additional 27.25 basis points of deposit spread. Given the relatively flat term structure of rates during much of the sample period, this amount could potentially represent a significant fraction of the net interest margin earned by a bank for that type of deposit account. These results provide strong support for the hypothesis that maturity-transformation risk is priced into the rates banks offer their CD customers.

7.4 Net Stable Funding Ratio Tests

The Net Stable Funding Ratio (NSFR) was first introduced by the Basel Committee on Banking Supervision (BCBS) in 2010 as part of the Basel III capital framework (Basel Committee on Banking Supervision (2010)). The NSFR is designed to reduce the funding risks stemming from maturity mismatches between bank assets and liabilities. The NSFR represents the ratio of available stable funding to required stable funding, and banks subject to the NSFR are required to maintain a ratio greater than one. Available stable funding measures the portion of bank funding that is stable and includes FDIC-insured retail deposits. Required stable funding measures the amount of stable funding a bank is required to hold, where long-maturity loans require more funding than unencumbered liquid short-term investments, for instance. By linking available stable funding to required stable funding, the NSFR imposes costs on banks bearing maturity-transformation. The NSFR was widely believed to trigger fundamental changes in business models and product pricing (see Standard & Poors (2010)). The U.S. implementation of the NSFR took effect on July 1, 2021.

The NSFR capital requirement also provides us with a natural experiment in which we can identify the deposit-pricing effects of a major exogenous shock in the costs associated with maturity-transformation activity. If banks price their maturity-transformation costs into deposit rates, we would expect to see a stronger relation between spreads and maturity mismatch after the NSFR capital requirement took effect. To test this, we again use a panel regression framework in which we regress spreads on the corresponding maturity mismatch. In this regression, however, we allow the slope coefficient for the maturity mismatch to differ for the period after July 1, 2021, when the NSFR capital requirement took effect. Specifically, we regress the spreads on the maturity mismatch and on the maturity mismatch interacted with an indicator variable that takes value one for observations after July 1, 2021 and zero otherwise. The regression specification is

$$\text{Spread}_{i,t} = \; c_1 \, \text{Mismatch}_{i,t} \; + \; c_2 \, \text{Mismatch}_{i,t} \times I_{\text{NSFR}} \; + \; \text{FE}_t \; + \; \epsilon_{i,t}, \quad (11)$$

where I_{NSFR} is the indicator variable, and the other terms have the same interpretation as in Equation (10). We then test if there is a change in the pricing of maturity-transformation risk by simply examining the significance of the coefficient for the interaction term. As before, the panel regression includes quarterly fixed effects to control for changes in the level of deposit spreads.

Table 8 reports the results from the panel regression (robust standard errors are clustered by tenor). As shown, there is a strong positive relation between option-adjusted deposit spreads and maturity mismatch during the pre-NSFR period. In particular, the coefficient for the maturity mismatch variable implies that the spread increases by 4.61 basis points for each additional year of maturity mismatch.

Turning now to the question of whether the imposition of the NSFR capital requirement impacted deposit pricing, Table 8 shows that the relation between spreads and maturity mismatch becomes significantly stronger after July 1, 2021. In particular, the slope coefficient increases by 0.0945 with a corresponding t-statistic of 4.82. This increase in the slope coefficient is significant from both a statistical and economic perspective. Adding together the slope coefficients implies that the spread now increases by $4.61 + 9.45 = 14.06$ basis points for each additional year of maturity mismatch following the implementation of the NSFR requirement in 2021. These results provide direct support for the hypothesis that deposit spreads are impacted by the costs that banks face in engaging in maturity-transformation activity.[11]

8. ROBUSTNESS RESULTS

In this section, we test the robustness of the main results with respect to an alternative choice of the discounting function used to compute option-adjusted deposit spreads. We also test the robustness of the results with respect to alternative choices of the parameters of the early withdrawal model. Finally, we examine how the results are impacted if we assume that households do not exercise their early withdrawal option strategically.

8.1 Treasury Discounting Function

We begin by redoing the analysis in Tables 5 through 8 using the discounting function based on Treasury rates (rather than the repo OIS discounting function). Tables A5 through A8 in the Internet Appendix present the results corresponding to those in Tables 5 through 8. As shown, the results obtained by using the discounting function based on Treasury rates are very similar to those based on the repo OIS discounting function. For instance, the average spreads for the six-month, one-year, two-year, three-year, four-year, and five-year CDs are 35.08, 21.02, 16.43, 11.27, 12.78, and 5.73 basis points, respectively, compared

[11] We note that if banks adjusted deposit spreads in anticipation of the NSFR taking effect, the coefficient estimate on the NSFR indicator variable would actually represent an underestimate of the total effect the NSFR had on deposit spreads.

to 39.54, 24.12, 19.79, 15.56, 16.58, and 7.37 basis points in Table 5. The coefficient estimates in the regressions presented in Tables 6, 7, and 8 are likewise very similar to those obtained using the Treasury discounting function. These robustness tests provide assurance that the results are robust to the choice of the discounting function.

8.2 Alternative Parameters for the Withdrawal Model

In this subsection, we redo the analysis using alternative values for the λ and γ parameters of the early withdrawal model. Specifically, we set γ and λ to twice their baseline values (resulting in values for γ and λ of 3.650 and 2.718 percent, respectively). Tables A9 through A12 in the Internet Appendix present the results corresponding to those in Tables 5 through 8. As shown, the results obtained by using these alternative values for λ and γ are very similar to those from the baseline parameterization.

8.3 The Early Withdrawal Option

As described earlier, there is strong evidence that some households exercise the early withdrawal option strategically. To highlight the role that the early withdrawal option plays in our results, however, we redo the analysis using the assumption that households do not exercise the early withdrawal option strategically. Under this counterfactual assumption, early withdrawal occurs only in response to exogenous liquidity shocks. This scenario can be nested within the early withdrawal model by setting γ equal to zero. Tables A13 through A16 in the Internet Appendix again present the results corresponding to those in Tables 5 through 8. While the results of this robustness test for Tables 5 and 6 are qualitatively similar, there are differences in the robustness tests for Tables 7 and 8. Specifically, in Table A15 the coefficient on the maturity mismatch variable is insignificant. In Table A16, only the interaction of the maturity mismatch variable with the Net Stable Funding Ratio indicator is significant. These results show that the early withdrawal option is important for understanding the economics of deposit spreads. They also provide assurance that the key results are qualitatively robust to specific parameter choices.

9. CONCLUSION

The term structure of bank CD rates provides a natural setting for testing whether maturity-transformation risk is priced into the rates that banks offer their customers. To do this, we first estimate the liquidity costs for each tenor using a valuation framework that takes into account the value of the early withdrawal option associated with bank CDs. We then examine the time-series and cross-sectional relations between these option-adjusted deposit spreads and the degree of maturity mismatch between bank assets and the tenor of the respective CDs.

We find that the option-adjusted spreads are strongly related to the maturity mismatch both in the time series and in the cross-section. The results suggest that an increase in maturity mismatch by one year translates into a 5.25 basis point increase in the deposit spread. Furthermore, this effect becomes much larger after the implementation of the Net Stable Funding Ratio capital requirement in 2021. These results imply that maturity-transformation risk is priced into bank CD rates, and that the resulting effect on deposit spreads can represent a significant portion of bank net interest margins.

Our results also have several implications for the ongoing debate about whether maturity-transformation risk is fully hedged by the deposit franchise. Similarly, the results raise the possibility that deposit spreads may be a reflection of depositors' willingness to pay for the liquidity/convenience provided by these bank products, rather than being an artifact of some broad type of market power that banks may have over their customers.

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Yankov, Vladimir, 2023, In Search of a Risk-Free Asset: Search Costs and Sticky Deposit Rates, *Journal of Money, Credit and Banking*, https://doi.org/10.1111/jmcb.13040.

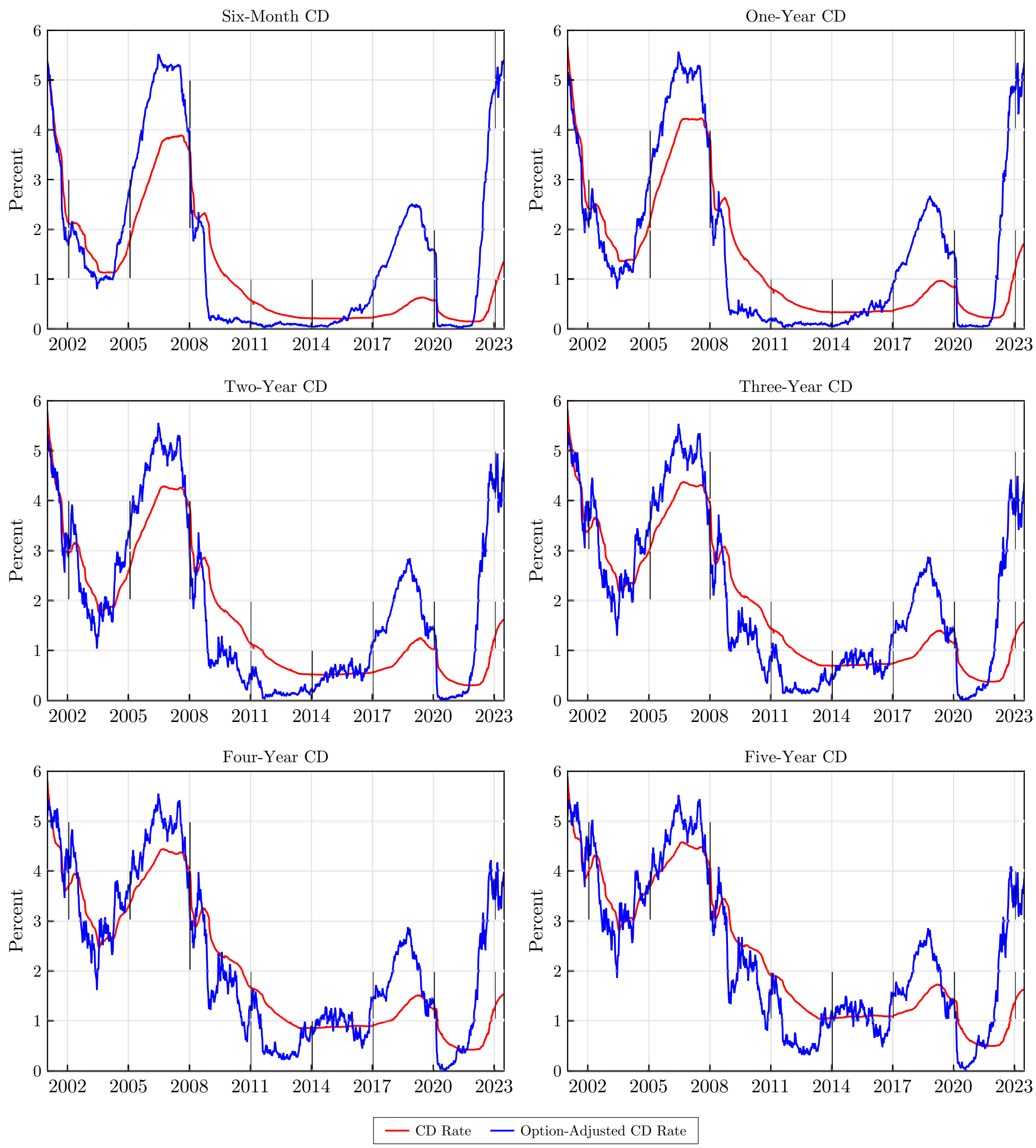

Figure 1. This graph plots the time series of the model-implied option-adjusted CD rate and the market CD rate by tenor.

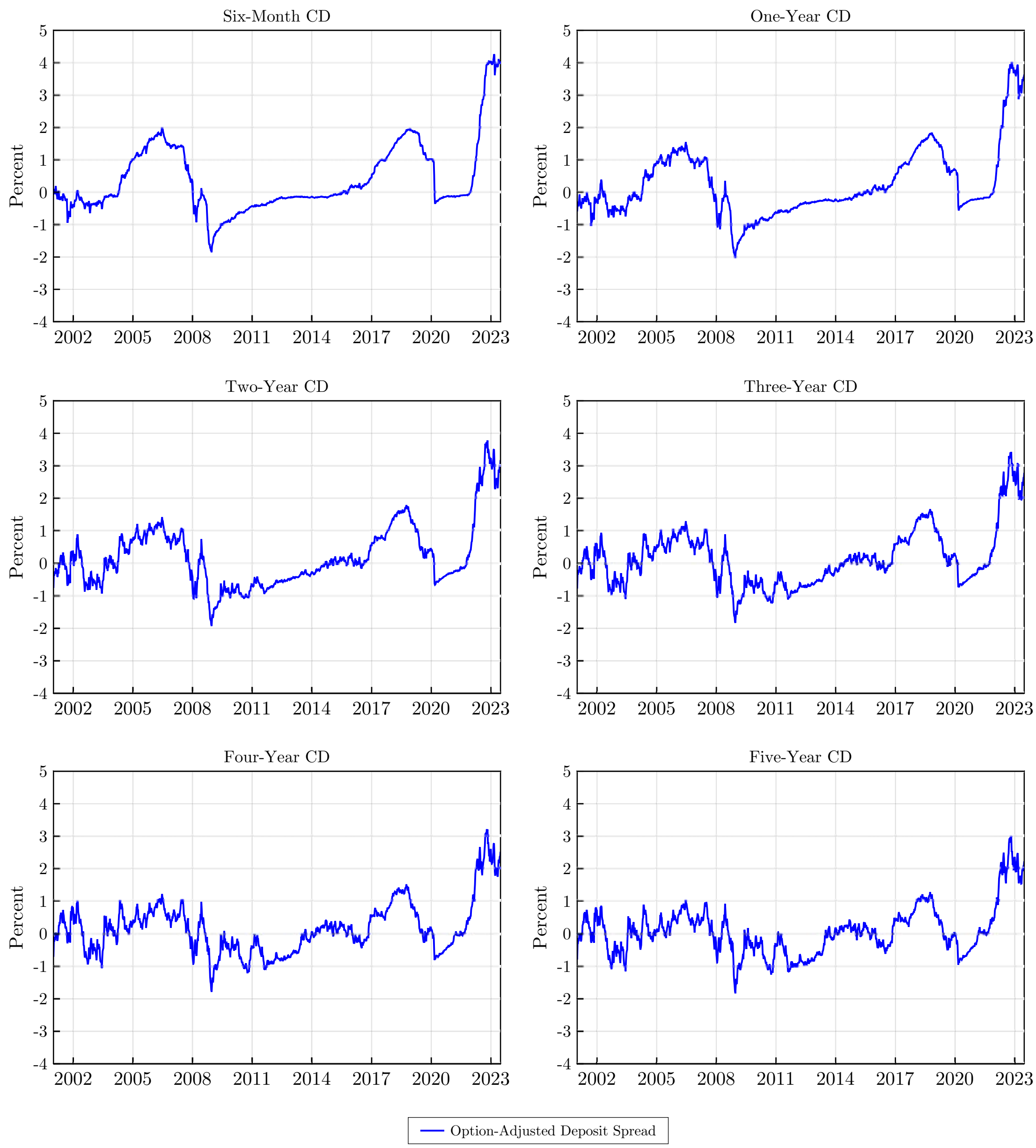

Figure 2. This graph plots the time series of the option-adjusted deposit spread by tenor.

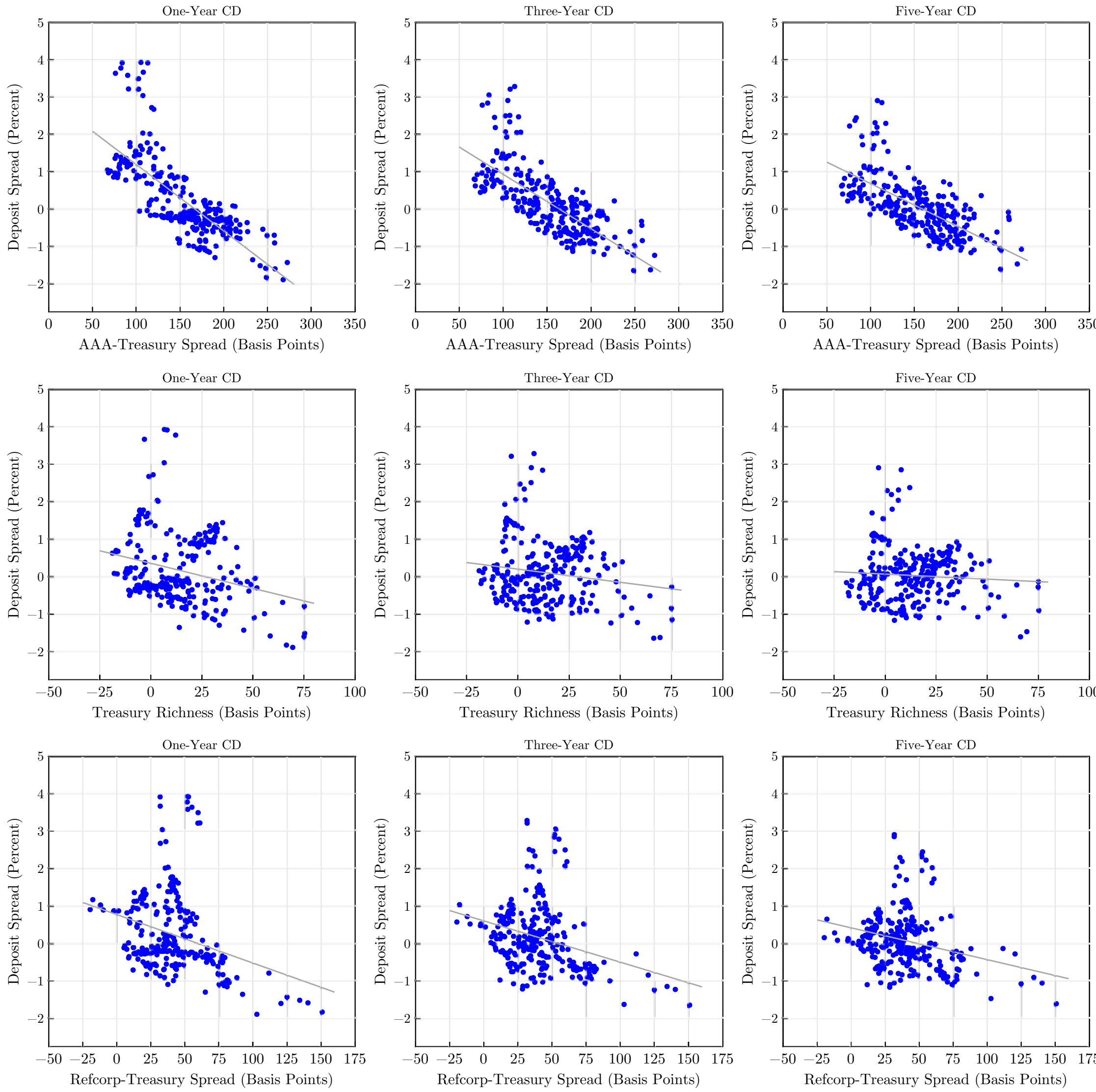

Figure 3. These scatterdiagrams plot the option-adjusted deposit spread for the indicated tenors against the Treasury convenience premium measures. The upper three panels plot the spreads against the AAA-Treasury spread. The middle three panels plot the spreads against the three-year Treasury richness spread. The lower three panels plot the spreads against the three-year Refcorp-Treasury spread.

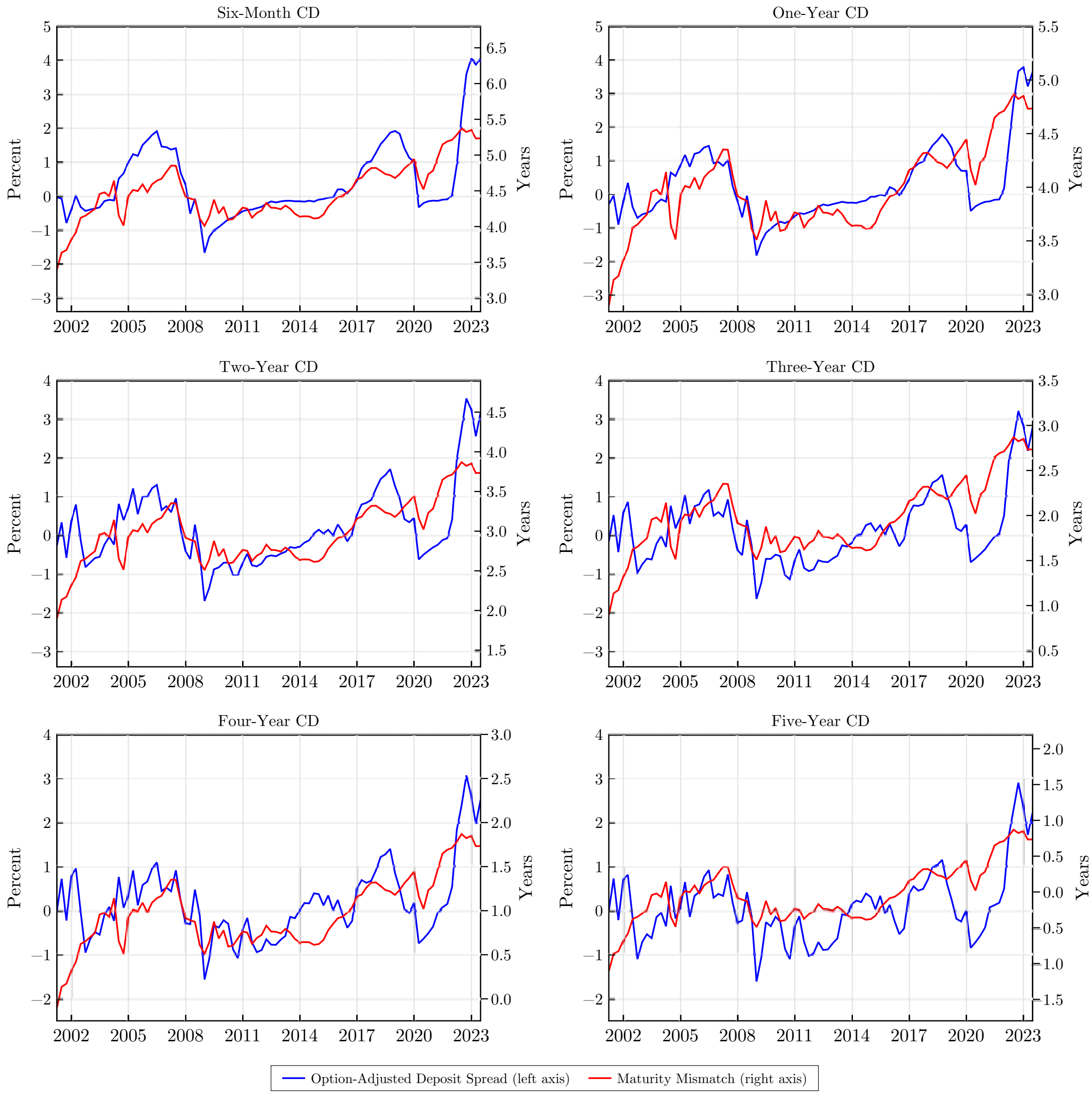

Figure 4. This graph plots the time series of the option-adjusted deposit spread and the maturity mismatch by tenor.

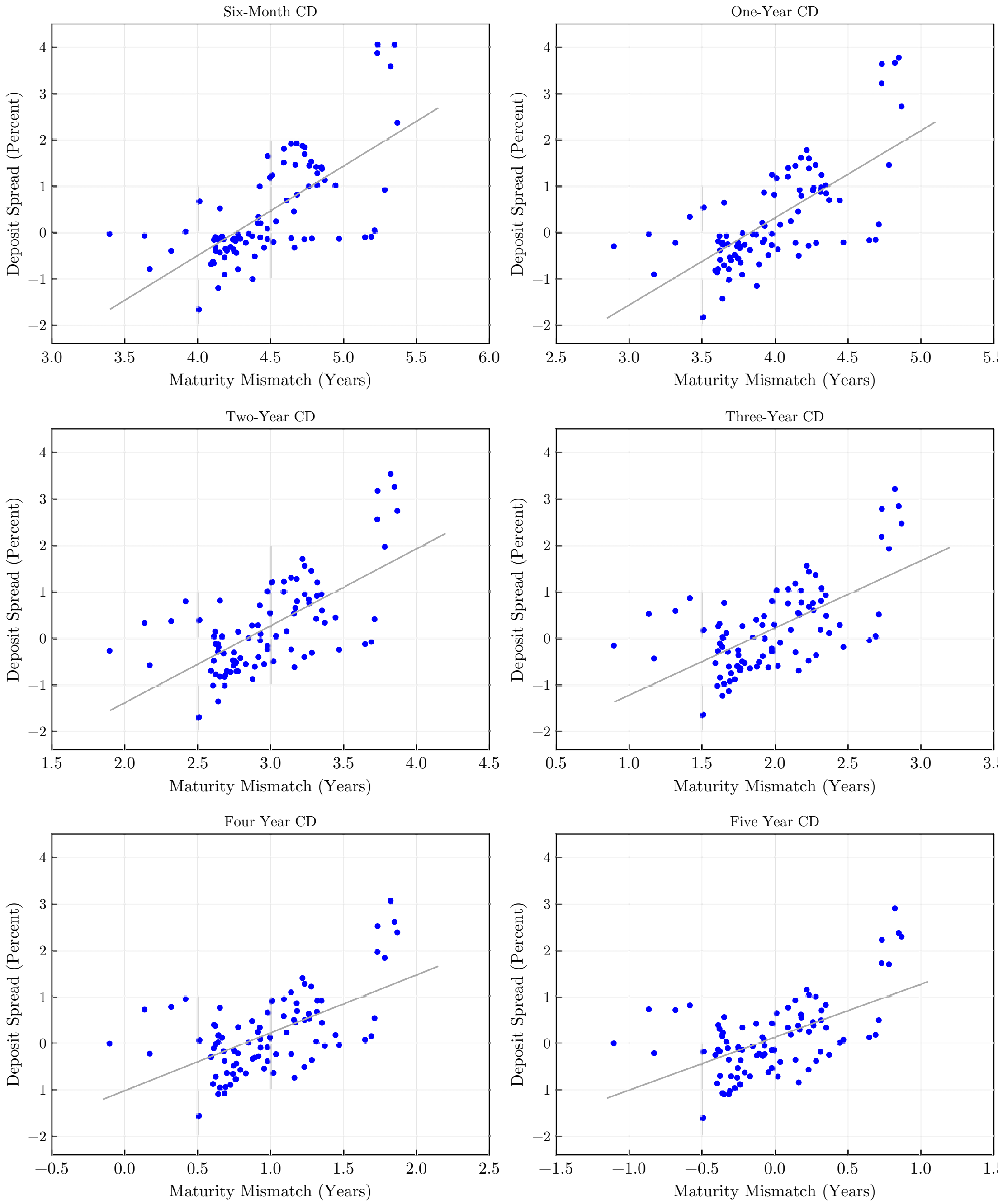

Figure 5. This graph plots the option-adjusted deposit spread against the maturity mismatch by tenor.

Table 1

Summary Statistics for CD Rates by Tenor. This table presents summary statistics for CD and riskless rates with the indicated tenors. The data on CD rates are furnished by S&P RateWatch and consist of CD rates quoted by banks for principal amounts less than or equal to $100,000. In the first panel, Mean presents the average CD rate for the indicated tenors where the average is taken across all weekly observations for a given tenor. Min, Med, and Max present the minimum, median and the maximum of weekly CD rates for the indicated tenors. CD rates are expressed as percentages. In the second panel, Mean presents the average Treasury spot rate for the indicated tenors. Treasury rates are semi-annually compounded and expressed as percentages. Min, Med, and Max present the minimum, median and the maximum riskless rate for the indicated tenors. Spread presents the average difference between the Treasury rate and the CD rate for the indicated tenors and is expressed in basis points. N presents the number of weekly observations for the indicated tenors. The sample period is weekly from January 5, 2001 to June 30, 2023.

	CD Rates				Treasury Rates					
	Mean	Min	Med	Max	Mean	Min	Med	Max	Spread	N
Six-Month CD	1.191	0.146	0.595	5.380	1.549	0.020	1.030	5.500	35.782	1,171
One-Year CD	1.431	0.221	0.889	5.686	1.651	0.040	1.220	5.399	22.008	1,171
Two-Year CD	1.675	0.300	1.141	5.779	1.868	0.090	1.504	5.270	19.222	1,171
Three-Year CD	1.888	0.369	1.349	5.826	2.079	0.110	1.655	5.228	19.012	1,171
Four-Year CD	2.048	0.411	1.472	5.810	2.295	0.143	1.953	5.214	24.690	1,171
Five-Year CD	2.251	0.485	1.670	5.883	2.503	0.210	2.326	5.207	25.157	1,171

Table 2

Summary Statistics for CD Early Withdrawal Penalties. This table presents summary statistics for early withdrawal penalties for CDs with the indicated tenors. The left panel presents summary statistics for early withdrawal penalties expressed in terms of the number of days of foregone interest for the indicated tenors. The right panel presents summary statistics for early withdrawal penalties expressed as a percentage of par value, calculated by multiplying withdrawal penalties expressed as the number of days of foregone interest with weekly deposit rates for the indicated tenors. The data on early withdrawal penalties are furnished by the Office of Thrift Supervision for the period from 2001 to 2012 and by S&P RateWatch for the period from 2013 to 2023. Withdrawal penalties for 2012, 2013, and 2014, are linearly interpolate between the annual averages for 2011 and 2015. Mean presents the average withdrawal penalty for the indicated tenors where the average is taken across all observations for a given tenor. Min, Med, and Max present the minimum, median and the maximum of early withdrawal penalties for the indicated tenors. N presents the number of weekly observations for the indicated tenors. The sample period is weekly from January 5, 2001 to June 30, 2023.

	Withdrawal Penalties as Number of Days of Interest				**Withdrawal Penalties as Percentage of Par Value**				
	Mean	Min	Med	Max	Mean	Min	Med	Max	N
Six-Month CD	86.31	73.30	87.20	99.76	0.29	0.03	0.13	1.39	1,171
One-Year CD	103.09	86.29	103.41	116.20	0.38	0.07	0.27	1.47	1,171
Two-Year CD	180.46	162.31	181.33	192.81	0.81	0.16	0.58	2.57	1,171
Three-Year CD	198.14	162.31	191.27	235.49	0.95	0.23	0.79	2.59	1,171
Four-Year CD	243.90	225.09	243.52	257.96	1.35	0.29	1.02	3.58	1,171
Five-Year CD	256.90	225.09	247.70	290.10	1.52	0.38	1.24	3.63	1,171

Table 3

Summary Statistics for CD Early Withdrawal Rates. This table presents summary statistics for early withdrawal rates for CDs with the indicated remaining maturities. Early withdrawal rates are calculated by dividing the total amount withdrawn prior to maturity of the CD to the total balance subject to early withdrawal penalties for the indicated maturity categories. Mean presents the average withdrawal rate where the average is taken across all observations for a given maturity category. Min, Med, and Max present the minimum, median, and the maximum early withdrawal rate for the indicated maturity categories. N presents the number of observations. The data are furnished by the Office of Thrift Supervision (OTS). The sample period is quarterly from September 2001 to December 2011.

Maturity	Mean	Min	Med	Max	N
$t \leq 3$ Months	4.944	2.631	4.078	10.279	39
$3 < t \leq 12$ Months	5.074	3.443	4.469	9.908	39
$12 < t \leq 36$ Months	3.739	2.692	3.348	8.194	39
$t > 36$ Months	6.128	2.363	4.704	23.096	39

Table 4

Results from the Regression of Changes in Early Withdrawal Rates for Long-Term CDs on Changes in Early Withdrawal Rates for Short-Term CDs. This table presents the results from the regression of quarterly changes in early withdrawal rates of CDs with a maturity greater than 12 months and up to 60 months ($\Delta\text{Withdrawal}_{\text{LT}}$) on quarterly changes in early withdrawal rates for CDs with a maturity of up to three months ($\Delta\text{Withdrawal}_{\text{ST}}$). Early withdrawal rates for CDs are quarterly for the period from Q3 2001 to Q4 2011 and are furnished by the Office of Thrift Supervision (OTS). Adj. R^2 and N denote the adjusted regression R-squared and the number of observations, respectively. Standard errors are based on Newey and West (1987). The superscripts * and ** denote significance at the ten-percent and five-percent levels, respectively. The regression is

$$\Delta\text{Withdrawal}_{\text{LT},t} = c_0 + c_1\ \Delta\text{Withdrawal}_{\text{ST},t} + \epsilon_t.$$

	Coeff	t-Stat
Intercept	−0.0016	−0.77
$\Delta\text{Withdrawal}_{\text{ST},t}$	0.2749	2.23^{**}
Adj. R^2		0.063
N		38

Table 5

Summary Statistics for Option-Adjusted Deposit Spreads. This table presents summary statistics for the option-adjusted deposit spreads for CDs with the indicated tenors and summary statistics for the deposit spreads for checking, savings, and money market account rates. Deposit spreads for checking, savings, and money market accounts are calculated as the difference between checking, savings, and money market account rates and the one-month riskless rate. All spreads are expressed in basis points. Mean, Min, Med, and Max present the average, minimum, median, and maximum of the spreads. The column t-Stat shows the Newey and West (1987) t-Statistic associated with the average premium reported in the column Mean. N presents the number of weekly observations. The sample period is weekly from January 5, 2001 to June 30, 2023.

	Mean	t-Stat	Min	Med	Max	N
Six-Month CD	39.54	4.71	−186.10	−9.41	426.25	1,171
One-Year CD	24.12	2.99	−202.74	−12.37	402.04	1,171
Two-Year CD	19.79	2.67	−193.53	1.15	377.75	1,171
Three-Year CD	15.56	2.32	−183.76	3.84	342.17	1,171
Four-Year CD	16.58	2.72	−179.67	8.87	321.29	1,171
Five-Year CD	7.37	1.30	−184.30	3.24	300.44	1,171
Checking	122.92	10.41	−21.26	62.75	515.09	1,171
Savings	109.08	9.56	−40.88	42.77	504.02	1,171
Money Market	87.50	8.66	−80.70	26.05	470.82	1,171

Table 6

Results from Regressions of Option-Adjusted Deposit Spreads on Maturity Mismatch. This table reports the results from the regressions of option-adjusted deposit spreads (Spread) for the indicated tenors on the maturity mismatch of the banking sector in Panel A. Panel B presents the results from the regressions of changes in option-adjusted deposit spreads (ΔSpread) for the indicated tenors on changes in maturity mismatch of the banking sector, where the changes are calculated over two quarters. Maturity mismatch is defined as the difference between the Drechsler, Savov, and Schnabl (2021) measure of asset repricing maturity of the banking sector and the tenor of the CD. Spreads are expressed in basis points and maturity mismatch is expressed in years. Adj. R^2 and N denote the adjusted regression R-squared and the number of observations, respectively. Standard errors are based on Newey and West (1987). The superscripts * and ** denote significance at the ten-percent and five-percent levels, respectively. The sample period is quarterly from Q1 2001 to Q2 2023. The regressions are

$$\text{Spread}_t = c_0 + c_1\,\text{Mismatch}_t + \epsilon_t \text{ (Panel A)} \quad \text{and} \quad \Delta\text{Spread}_t = c_0 + c_1\,\Delta\text{Mismatch}_t + \epsilon_t \text{ (Panel B)}.$$

	Six-Month CD		One-Year CD		Two-Year CD		Three-Year CD		Four-Year CD		Five-Year CD	
	Coeff	t-Stat	Coeff	t-Stat	Coeff	t-Stat	Coeff	t-Stat	Coeff	t-Stat	Coeff	t-Stat
Panel A - Levels												
Intercept	−8.223	−3.38**	−7.211	−3.46**	−4.701	−3.11**	−2.673	−2.82**	−1.021	−2.22**	0.130	1.05
Mismatch	1.931	3.44**	1.881	3.50**	1.656	3.20**	1.446	2.98**	1.245	2.65**	1.139	2.50**
Adj. R^2		0.461		0.473		0.423		0.382		0.333		0.312
N		90		90		90		90		90		90
Panel B - Differences												
Intercept	0.028	0.50	0.021	0.41	0.006	0.13	−0.003	−0.07	−0.010	−0.23	−0.014	−0.33
ΔMismatch	0.487	1.59	0.539	1.79*	0.672	2.17**	0.742	2.32**	0.783	2.43**	0.794	2.46**
Adj. R^2		0.045		0.049		0.061		0.070		0.077		0.079
N		88		88		88		88		88		88

Table 7

Results from Panel Regressions of Option-Adjusted Deposit Spreads on Maturity Mismatch. This table presents results from the panel regressions of option-adjusted deposit spreads (Spread) for CDs with tenors $i = 0.5, 1, 2, 3, 4, 5$ years on the maturity mismatch of the banking sector (Mismatch) and on quarterly fixed effects (FE). Maturity mismatch is defined as the difference between the Drechsler, Savov, and Schnabl (2021) measure of asset repricing maturity of the banking sector and the tenor of the CD. Adj. R^2 and N denote the adjusted regression R-squared and the number of observations, respectively. The superscripts * and ** denote significance at the ten-percent and five-percent levels, respectively. Robust standard errors are clustered by CD tenor. The data are quarterly from Q1 2001 to Q2 2023. The regression is

$$\text{Spread}_{i,t} = c_1 \,\text{Mismatch}_{i,t} + FE_t + \epsilon_{i,t}.$$

	Coeff	t-Stat
Mismatch	0.0545	3.74**
Quarterly Fixed Effects		Yes
Adj. R^2		0.918
N		540

Table 8

Results from Panel Regressions of Option-Adjusted Deposit Spreads on Maturity Mismatch and an NSFR Indicator Variable. This table presents results from the panel regressions of option-adjusted deposit spreads (Spread) for CDs with tenors $i = 0.5, 1, 2, 3, 4, 5$ years on the maturity mismatch of the banking sector (Mismatch) and on quarterly fixed effects (FE). The regression includes an interaction of the maturity mismatch with an indicator variable I_{NSFR} that takes the value one for quarterly dates after July 1, 2021, when the Net Stable Funding Ratio (NSFR) capital requirement became effective. Maturity mismatch is defined as the difference between the Drechsler, Savov, and Schnabl (2021) measure of asset repricing maturity of the banking sector and the tenor of the CD. Adj. R^2 and N denote the adjusted regression R-squared and the number of observations, respectively. The superscripts * and ** denote significance at the ten-percent and five-percent levels, respectively. Robust standard errors are clustered by CD tenor. The data are quarterly from Q1 2001 to Q2 2023. The regression is

$$\text{Spread}_{i,t} = c_1 \, \text{Mismatch}_{i,t} + c_2 \, \text{Mismatch}_{i,t} \times I_{\text{NSFR}} + FE_t + \epsilon_{i,t}.$$

	Coeff	t-Stat
Mismatch	0.0461	2.83^{**}
Mismatch $\times$ I_{NSFR}	0.0945	4.82^{**}
Quarterly Fixed Effects		Yes
Adj. R^2		0.920
N		540