ECONSTOR
Make Your Publications Visible.

A Service of

ZBW Leibniz-Informationszentrum Wirtschaft
Leibniz Information Centre for Economics

Dias, Daniel; Schmidt-Eisenlohr, Tim

Working Paper

The Uninsured Deposit Premium

CESifo Working Paper, No. 12103

Provided in Cooperation with:
Ifo Institute – Leibniz Institute for Economic Research at the University of Munich

Suggested Citation: Dias, Daniel; Schmidt-Eisenlohr, Tim (2025) : The Uninsured Deposit Premium, CESifo Working Paper, No. 12103, Munich Society for the Promotion of Economic Research - CESifo GmbH, Munich

This Version is available at:
https://hdl.handle.net/10419/327713

Standard-Nutzungsbedingungen:

Die Dokumente auf EconStor dürfen zu eigenen wissenschaftlichen Zwecken und zum Privatgebrauch gespeichert und kopiert werden.

Sie dürfen die Dokumente nicht für öffentliche oder kommerzielle Zwecke vervielfältigen, öffentlich ausstellen, öffentlich zugänglich machen, vertreiben oder anderweitig nutzen.

Sofern die Verfasser die Dokumente unter Open-Content-Lizenzen (insbesondere CC-Lizenzen) zur Verfügung gestellt haben sollten, gelten abweichend von diesen Nutzungsbedingungen die in der dort genannten Lizenz gewährten Nutzungsrechte.

Terms of use:

Documents in EconStor may be saved and copied for your personal and scholarly purposes.

You are not to copy documents for public or commercial purposes, to exhibit the documents publicly, to make them publicly available on the internet, or to distribute or otherwise use the documents in public.

If the documents have been made available under an Open Content Licence (especially Creative Commons Licences), you may exercise further usage rights as specified in the indicated licence.

WWW.ECONSTOR.EU

Mitglied der Leibniz-Gemeinschaft

CES ifo Working Papers

12103
2025

August 2025

The Uninsured Deposit Premium

Daniel Dias, Tim Schmidt-Eisenlohr

CES ifo

Imprint:

CESifo Working Papers

ISSN 2364-1428 (digital)

Publisher and distributor: Munich Society for the Promotion
of Economic Research - CESifo GmbH

Poschingerstr. 5, 81679 Munich, Germany
Telephone +49 (0)89 2180-2740

Email office@cesifo.de
https://www.cesifo.org

Editor: Clemens Fuest

An electronic version of the paper may be downloaded free of charge
· from the CESifo website: www.ifo.de/en/cesifo/publications/cesifo-working-papers
· from the SSRN website: www.ssrn.com/index.cfm/en/cesifo/
· from the RePEc website: https://ideas.repec.org/s/ces/ceswps.html

The Uninsured Deposit Premium*

Daniel A. Dias and Tim Schmidt-Eisenlohr
Federal Reserve Board

August 28, 2025

Abstract

We estimate the *uninsured deposit premium* – the difference between the rates paid on uninsured versus insured deposits – by linking observed average deposit rates to an estimated share of uninsured deposits. Using U.S. bank data from 1991 to 2025, we show that the average uninsured deposit premium rose by nearly 400 basis points over this period. This rise reflects both falling insured deposit rates and rising uninsured deposit rates. We find a strong correlation with the monetary policy cycle: a one-percentage-point increase in the Federal Funds Rate corresponds to a rise of roughly 32 basis points in the uninsured deposit premium. We develop a bargaining model between banks, insured depositors, and uninsured depositors that explains these dynamics.

JEL classification codes: E52, G21, G28.
Keywords: Uninsured Deposits; Monetary Policy; Bank Funding; Deposit Pricing

*We thank Bill Bassett, Itay Goldstein, as well as seminar participants at the Federal Reserve Board for helpful comments. The views in this paper are solely the responsibility of the authors and should not be interpreted as reflecting the views of the Board of Governors of the Federal Reserve System or of any other person associated with the Federal Reserve System. All errors are our own.

1 Introduction

Deposits are the backbone of U.S. bank funding, accounting for about 80 percent of liabilities. By the end of 2024, about 60 percent of deposits were FDIC-insured, leaving the remaining 40 percent uninsured.[1] Despite their growing importance, little is known about how much more banks pay for uninsured relative to insured deposits, how this premium evolves over time, and what factors drive its changes – largely due to limited data availability.

To address this gap, we develop a novel methodology for estimating the *uninsured deposit premium* – the gap between average deposit rates on uninsured and insured deposits. This measure helps us answer several questions: Do banks pay higher rates on uninsured deposits? How has the premium evolved over time and how is it related to the stance of monetary policy?

The uninsured deposit premium is estimated using U.S. bank-level data spanning Q1 1991 to Q1 2025. We begin with a baseline specification and subsequently account for differences in the shares of deposit categories across banks. Next, we examine how the uninsured deposit premium co-moves with the monetary policy cycle and broader macro-financial conditions. Finally, we analyze how differences in bank risk profiles, measured through balance-sheet and income statement indicators, affect the uninsured deposit premium.

Our analysis shows that the uninsured deposit premium exhibits both substantial fluctuations and a persistent upward trend. Two key facts emerge. First, the uninsured deposit premium has risen by roughly 400 basis points since the early 1990s. This increase reflects

[1] In the United States, the first $250,000 per depositor and bank are insured by the Federal Deposit Insurance Corporation (FDIC). The literature on the benefits of a capped deposit-insurance scheme is quite extensive. See, e.g., Diamond and Dybvig (1983) or Goldstein and Pauzner (2005) for theoretical motivations and (Demirgüç-Kunt and Kane, 2002) for an international review of experiences with deposit insurance schema.

both rising average uninsured deposit rates and falling average insured deposit rates. Second, the uninsured deposit premium is highly correlated with the Federal Funds Rate (FFR), and movements in the latter explain most of the premium's cyclical behavior.

The findings on the uninsured deposit premium are similar across weighted and unweighted regressions, which is notable given the concentration of assets among the largest banks. This suggests that the forces driving both the long-run increase in the uninsured deposit premium and its cyclical relationship with the federal funds rate are present in small banks, which dominate the unweighted regressions, as well as in the largest banks, which dominate the weighted regressions.

The long-term upward trend in the uninsured deposit premium also carries important distributional consequences across depositor types. Large depositors—primarily corporations—benefited from rising uninsured rates, while small retail depositors suffered from declining insured deposit rates. Although ongoing work is examining the sources of this development, our initial analysis suggests that changes in bargaining power over time played a key role. By contrast, changes in uninsured deposit risk appear to have been less important.

As guidance for our empirical analysis, we develop a theoretical framework that helps explain both the long-term trend and the cyclical dynamics of the uninsured deposit premium. In the model, banks bargain with insured and uninsured depositors over deposit rates. The bank's outside option is wholesale funding priced at the FFR. When uninsured depositors have greater bargaining power than insured depositors, and their outside option co-moves with the FFR at least as much as that of insured depositors, the model generates the observed positive relationship between the uninsured deposit premium and the FFR.

The rest of the paper is organized as follows. Section 2 reviews related literature. Section

3 outlines the methodological approach, and Section 4 describes the data. Section 5 provides supporting evidence for the methodology, and Section 6 reports the main findings. Section 7 develops a conceptual framework for the drivers of the uninsured deposit premium and tests some of its empirical implications. The final section concludes.

2 Related Literature

This paper contributes to two main strands of the literature—(i) deposit pricing and the role of deposit insurance, and (ii) deposit franchise values—by directly estimating the gap between the deposit rates uninsured and insured deposits and uncovering its magnitude and key drivers. It also documents new evidence on the time-series evolution and cross-sectional variation of the uninsured deposit premium.

Deposit Pricing The literature on banks' market power in both lending and funding dates back to Klein (1971) and Monti (1972), who were the first to formally model banks as firms with such dual market power.[2] With market power over deposits – the primary funding source for most institutions – banks can pay depositors less than they would under fully market-based funding. This literature is extensive, with prominent contributions including Hannan and Berger (1991), Neumark and Sharpe (1992), Kim et al. (2003), and Drechsler et al. (2017).[3]

More directly related to our work is the literature on the pricing of insured and uninsured

[2]The work of Klein (1971) and Monti (1972) is commonly referred to as the *Monti–Klein* model of banking, a workhorse in both applied and theoretical banking research.

[3]See also Berger and Hannan (1989); Calem and Carlino (1991); Sharpe (1997); Rosen (2007); Hannan and Adams (2011); Park and Pennacchi (2009) for additional U.S. evidence on funding-side pricing power and pass-through.

deposits. A central idea in this literature is that uninsured deposits are expected to command a premium over insured deposits because they carry additional risk. This idea, formalized in several studies (e.g., Egan et al., 2017; Drechsler et al., 2023), features prominently in the existing literature. Empirical work has also documented the rate differential between uninsured and insured deposits. Jacewitz and Pogach (2018) document a systematic uninsured deposit premium in U.S. banks. Using detailed branch-level deposit rate data, they find that large "too-big-to-fail" banks paid lower premiums than smaller banks, attributing the gap to implicit government guarantees. Maechler and McDill (2006) employ a panel of U.S. banks and show that deteriorating fundamentals lead uninsured depositors to demand higher rates or withdraw funds, confirming the existence of an uninsured deposit premium.

Cross-country analyses reinforce the view that uninsured deposits command a premium and serve as a source of market discipline. Bartholdy et al. (2003) document significantly higher deposit rate spreads in countries with limited or no deposit insurance. In a study of emerging markets, Martinez-Peria and Schmukler (2001) show that depositors demand higher rates or withdraw funds to discipline banks. For the countries they study, the presence of deposit insurance does not weaken these market discipline forces, likely because the schemes lack credibility.

We make four main contributions to this literature. First, we provide long-run estimates of the uninsured deposit premium for all U.S. banks over the past four decades. Second, we document that the uninsured deposit premium co-moves strongly with the FFR and has trended upward over the same period. Third, we develop a bargaining model that can explain these two findings. Finally, unlike much of the existing literature, we find no systematic evidence that riskier banks pay a higher uninsured deposit premium—their higher

risk shows up in higher deposit rates instead.

Deposit Franchise Values and Deposit Betas Recent studies have examined the value banks derive from deposit franchises and its relationship with interest rate risk. Drechsler et al. (2021) develop a model in which the deposit franchise is valuable because deposits provide funding at below-market rates, a benefit that grows when market rates rise, since banks pass on only part of the increase to depositors. At the same time, banks incur rate-insensitive deposit service costs, which fall in net-present-value terms as interest rates rise. As a result, the deposit franchise increases in value with market rates, behaving like an intangible asset with negative duration and serving as a hedge against positive duration risk on the asset side. In related work, DeMarzo et al. (2024) extend the analysis to total bank franchise value, arguing that it has positive duration overall because interest-insensitive spreads on deposits and loans outweigh the effect of rate-insensitive service costs.

More recently, Drechsler et al. (2025) extended the analysis of deposit franchise value to include the role of uninsured deposits. In their setup, depositor runs become more likely and more costly as market rates rise. They estimate deposit franchise values by decomposing the aggregate deposit beta into insured and uninsured components, regressing it on the respective shares of uninsured and insured deposits.[4] Their finding that the uninsured deposit beta exceeds the insured beta is consistent with our result that the uninsured deposit rate is more strongly correlated with the FFR than the insured deposit rate. While Drechsler et al. (2025) establish this relationship at three points in the 2020s, we show that it has held over the past four decades. The long-term trends we uncover in insured and uninsured deposit rates

[4]The empirical part of their analysis was conducted contemporaneously with, but independent of, our analysis.

have important implications for deposit franchise values across banks with varying reliance on uninsured deposits. In addition, we broaden the analysis by showing that the results hold when controlling for the shares of time and non-transaction deposits at the bank level and when using weighted regressions.

Overall, previous research consistently finds that uninsured deposits carry a premium reflecting depositor risk exposure and market discipline. This premium varies notably across banks with differences in size, risk profile, and macroeconomic conditions. However, most prior empirical work has typically relied on specific deposit products, limited episodes (often crises), or indirect proxies for uninsured funding costs. In contrast, our paper provides a systematic long-term measure of the uninsured deposit premium. Drawing on comprehensive U.S. bank-level data, we directly estimate the aggregate premium over four decades, documenting its substantial increase during this period and strong correlation with monetary policy cycles.

3 Methodology

In this section we first present the key idea behind our empirical methodology and then provide the precise equations we estimate.

3.1 Conceptual Approach

There are two key challenges when estimating the difference between rates paid on uninsured and insured deposits. First, information on the total amount of insured and uninsured deposits is only reported by banks above a certain asset threshold and has only been reported since 2002. Second, banks do not separately report interest expenses on insured and

uninsured deposits but only their interest expenses on total deposits.

The first challenge can be addressed by using information on the number and value of large and small deposit accounts reported by banks. The basic idea is that any funds in deposit accounts that exceed the FDIC insurance limit must be uninsured. Following this approach produces an estimate of insured (*vis-a-vis* uninsured) deposits that in the aggregate lines up well with publicly available data on amounts of uninsured and insured deposits. Addressing the second challenge is less straightforward and constitutes a key contribution of our paper.

Specifically, to overcome the issue of not observing insured and uninsured interest expenses separately, we propose a methodology that estimates the average gap between the interest rate on uninsured and insured deposits by regressing the aggregate deposit rate at the bank level on the share of uninsured deposits. Grouping deposits into two mutually exclusive groups (e.g., insured and uninsured, or time and demand deposits), we can write the average deposit rate at the bank level as the weighted average between these two groups.

$$DR_{it} = \alpha_{it}^{A} \cdot DR_{it}^{A} + (1 - \alpha_{it}^{A}) \cdot DR_{it}^{B}, \quad \text{where} \quad \alpha_{it}^{A} + \alpha_{it}^{B} = 1. \tag{1}$$

We can re-arrange equation (1) so that it relates the deposit rate at the bank level, DR_{it}, with the deposit rate in deposit type A, DR_{it}^{A}, the share of deposits type B, α_{it}^{B}, and the gap between the deposit rates in deposits type B and type A, $DR_{it}^{B} - DR_{it}^{A}$,

$$DR_{it} = DR_{it}^{A} + \alpha_{it}^{B} \cdot (DR_{it}^{B} - DR_{it}^{A}). \tag{2}$$

Letting $\overline{DR}_{t}^{A}$ and $\overline{DR}_{t}^{B}$ denote the average deposit rate of types A and B in period t, respectively, by adding and subtracting $\alpha_{it}^{B}(\overline{DR}_{t}^{B} - \overline{DR}_{t}^{A})$ and $\overline{DR}_{t}^{A}$ to equation (2), this equation can be rewritten as

$$DR_{it} \;=\; \underbrace{\overline{DR}_t^{\,A}}_{\theta_t} \;+\; \underbrace{\left(\overline{DR}_t^{\,B} - \overline{DR}_t^{\,A}\right)}_{\beta_t} \alpha_{it}^{B} \;+\; \epsilon_{it}^{pool}, \tag{3}$$

where

$$\epsilon_{it}^{pool} \;=\; \alpha_{it}^{A} \underbrace{\left(DR_{it}^{A} - \overline{DR}_t^{\,A}\right)}_{\text{insured-rate deviation}} + \alpha_{it}^{B} \underbrace{\left(DR_{it}^{B} - \overline{DR}_t^{\,B}\right)}_{\text{uninsured-rate deviation}},$$

Equation (3) can be simplified to

$$DR_{it} = \theta_t + \beta_t \cdot \alpha_{it}^{B} + \epsilon_{it}. \tag{4}$$

Parameter β_t can be interpreted as the difference between the average deposit rate on deposits type B and type A – $(\overline{DR}_t^{\,B} - \overline{DR}_t^{\,A})$. Since in the data we can observe both the aggregate deposit rate and the share of uninsured deposits for bank i in period t, we can estimate β_t with OLS.

One issue with estimating equation (4) with OLS is that α_{it}^{B} and ϵ_{it}^{B} could be correlated and the OLS estimates of β_t^B may be biased.[5] The ideal solution to this endogeneity concern would be to estimate (4) with instrumental variables. Unfortunately, at this point, we are not aware of a good instrument for α_{it}^{B}. However, we can gauge the size of the bias by comparing the estimate of β_t^B with the observed $(\overline{DR}_t^{\,B} - \overline{DR}_t^{\,A})$ in cases for which the data allow it. In particular, we directly observe interest expenses on time and demand and nontransactional and transactional deposits as well as the balances for each of these types of deposits at the bank-time level. That is, we can calculate the true rate differences in the data for these

[5] Specifically, we can derive the covariance between the share and the error term as $\text{Cov}(\alpha_{it}^{B}, \epsilon_{it}^{B}) = \text{Cov}(\alpha_{it}^{B}, DR_{it}^{A} - \overline{DR}_t^{\,A}) - \text{Cov}(\alpha_{it}^{B}, \alpha_{it}^{B} \cdot (DR_{it}^{A} - \overline{DR}_t^{\,A})) + \text{Cov}(\alpha_{it}^{B}, \alpha_{it}^{B} \cdot (DR_{it}^{B} - \overline{DR}_t^{\,B}))$

subgroups and compare them with estimates obtained from regressions analogous to equation (4).

The approach we described so far allows us to estimate the average difference in the deposit rate between two mutually exclusive types of deposits by regressing the average deposit rate at the bank level on the share of one the type of deposits (type B in the discussion above). However, we would like to be able to separate how much of the difference between two types of deposits is driven by the splitting characteristic (e.g., insured vs uninsured or time vs demand) and how much of the difference is driven by correlations with other deposit characteristics.

Consider the case of time deposits. Time deposits are more likely to be insured deposits than uninsured deposits. As time deposits pay higher interest rates than demand deposits, this will, all else equal, lead to higher average interest rates on insured deposits. In the baseline regression that only uses the uninsured deposit split, this will yield a smaller estimate of the uninsured deposit premium than when controlling for this feature of the data.

To estimate the uninsured deposit premium conditional on other deposit characteristics, we expand equation (4), adding controls for these additional shares. In the next subsection we present the specific econometric models we use in the empirical analysis.

3.2 Empirical Models

Following the discussion of the previous subsection, we estimate three econometric models. The first model is the immediate counterpart of equation (4)

$$DR_{it} = \lambda_i + \theta_t + \beta_t^j \alpha_{it}^j + \epsilon_{it}, \quad j = \{uninsured,\ time,\ non-transactional\} \tag{5}$$

In equation (5), λ_i are bank fixed effects; θ_t are time fixed effects; β_t is a time-varying coefficient associated with the share of deposits type $j = \{uninsured,\ time,\ non{-}transactional\}$, which is represented by α_{it}^j; ϵ_{it} is a zero-mean error term. Relative to equation (4), equation (5) adds bank fixed effects, λ_i. Bank fixed effects should help with controlling for cross-sectional differences in deposit rates, while time fixed effects control for changes in macroeconomic and financial conditions that are expected to affect bank's deposit rates as well as the composition of banks' deposit types.

This specification yields estimates of the difference between the average deposit rate on uninsured and insured deposits at each point in time, given by the coefficients $\beta_t, \quad t = 1, 2, \ldots, T$. We similarly obtain time-varying estimates for the differences between time and demand deposit rates, and between non-transactional and transactional deposit rates. As discussed above, comparing these estimates with observed data provides an indication of the estimator's bias. Understanding this bias is essential for interpreting our main findings on the uninsured deposit premium, which cannot be measured directly from the data.

As discussed above, estimates based on equation (5) do not control for other deposit characteristics. To address this issue, we expand equation (5) by including information on the composition of a bank's deposit base along the dimensions $j = \{uninsured,\ time,\ non-transactional\}$.

$$DR_{it} = \lambda_i + \theta_t + \beta_t^U \alpha_{it}^U + \beta_t^T \alpha_{it}^T + \beta_t^{NT} \alpha_{it}^{NT} + \epsilon_{it} \tag{6}$$

Superscripts U, T, and NT in equation (6) stand for uninsured, time, and non-transactional.

In Appendix A, we derive approximations for the bias in OLS estimates of equations 5 and 6, with results shown in Sections A.1 and A.2, respectively. A key takeaway from this analysis is that the bias is likely to be a weighted average of positive and negative components, which should attenuate its overall magnitude.

Another dimension we are interested in analyzing is whether bank-specific risk is priced in banks' average deposit rates and if bank-specific risk plays a role in the uninsured deposit premium. One challenge we face in addressing this question is that we do not observe or are able to easily estimate a bank-specific uninsured deposit premium. To overcome this challenge, we expand equation (6) to include a banks-specific risk metric and an interaction between that risk metric and the share of uninsured deposits. In the analysis we are going to consider two dimensions over which banks can be ranked from a risk point of view: the delinquency rate and the level of bank capitalization measured by the market-adjusted tangible common equity ratio (MATCE).[6]

To evaluate whether bank-specific risk affects the uninsured deposit premium at the bank

[6]We chose to use a market-adjusted measure of capital because, following the large increases in the policy rate and other interest rates that occurred between end 2021 and mid-2023, many banks had large fair-value losses on their balance sheets and only some of these losses show up in banks' book-value capital ratios. In the case of the tangible common equity (TCE) ratio, all fair-value losses (or gains) on securities in the available-for-sale (AFS) portfolio are accounted in the TCE ratio, but fair-value losses (or gains) in the hold-to-maturity (HTM) portfolio are not. These unrealized losses, together with a large dependence on uninsured deposits, were the reasons for the March 2023 banking stress episode that started with the failure of Silicon Valley Bank on March 10, 2023.

level, in each quarter, we split the sample into three groups based on their levels of the risk measure considered. Based on this split, we create indicator variables d_{it}^{mj}, with terciles $j \in \{1, 2, 3\}$, risk measure $m \in \{\text{delinquency rate, MATCE ratio}\}$ for bank i in year t. Using these indicator variables (d_{it}^{mj}), we augment equation 6 to get

$$DR_{it} = \lambda_i + \theta_t + \beta_t^U \alpha_{it}^U + \beta_t^T \alpha_{it}^T + \beta_t^{NT} \alpha_{it}^{NT} + \gamma_t^{m1} d^{m1} + \gamma_t^{m2} d^{m2} + \beta_t^{U,m1} d^{m1} \alpha_{it}^U + \beta_t^{U,m2} d^{m2} \alpha_{it}^U + \epsilon_{it} \tag{7}$$

There are two main coefficients of interest in equation 7. First, the parameter γ_t^{m1} captures how the average deposit rate paid by banks in the bottom tercile of variable m differs from that paid by banks in the top tercile in quarter t. Second, $\beta_t^{U,m1}$ measures how the average uninsured deposit premium paid by banks in the bottom tercile differs from that paid by banks in the top tercile, again with respect to variable m in quarter t.

4 Data

To estimate the models presented in the previous section, we draw on quarterly, bank-level balance-sheet data from the "Call Reports" (formally known as the *Consolidated Reports of Condition and Income*). We focus exclusively on domestic deposits at U.S. banks because only these deposits are insured by the FDIC. For each institution and quarter (1991Q1–2025Q1), among many other items, the dataset reports total deposits, time-deposit balances, and transaction-account balances—as well as the total interest paid on each deposit category.

From the Call Report data, we construct measures of average deposit rates at the bank

level and average deposit rates by deposit type as well as measures of deposit shares:

$$DR_{it}^{j} \;=\; 400 \cdot \frac{\text{Interest Expenses}_{it}^{j}}{\text{Deposits}_{it}^{j}}, \tag{8}$$

with $j \in \{\text{Total, Time, Demand, Transactional, Non-Transactional}\}$. And:

$$\alpha_{it}^{k} \;=\; 100 \cdot \frac{\text{Deposits}_{it}^{k}}{\text{Deposits}_{it}^{\text{Total}}} \tag{9}$$

with $k \in \{\text{Uninsured, Time, Non-Transactional}\}$. A key variable in our analysis is the share of insured (and uninsured) deposits. Calculating this share requires data on the breakdown between insured and uninsured deposits, which are not uniformly available across banks or over time. Under current regulations, only banks with assets above $1 billion must report uninsured deposits in their Call Reports. From 2002 to 2005, however, all banks were required to report uninsured deposits; prior to 2002 no such requirement existed. Consequently, reported uninsured-deposit data cover only part of our sample and, in recent years, only a subset of institutions (albeit those representing a large share of system assets).

To obtain a consistent, full-sample measure, we follow prior studies and impute each bank's insured and uninsured deposits using data on domestic deposit size categories (amounts and number of accounts).[7] This approach yields a methodologically comparable series spanning 1991Q1–2025Q1. When we compare imputed and reported uninsured-deposit shares

[7]For each of three categories—noninterest-bearing transaction accounts; other deposit accounts (demand, savings, and time, excluding retirement accounts); and retirement deposit accounts—the Call Reports report both the end-of-period aggregate balance and the number of accounts with balances exceeding $250K (previously $100K before 2008Q3). Using this information, we calculate uninsured deposits as the sum of balances above the threshold minus the threshold multiplied by the corresponding number of accounts; the balance below the threshold (and all other deposits) is insured. During the Transaction Account Guarantee Program (Oct. 14 2008–Dec. 31 2009), we treat all transaction accounts as fully insured regardless of size.

for reporting banks, we observe only minor differences, and, most importantly, our core results remain qualitatively and quantitatively similar whether we use reported or imputed data. Moreover, analysis of the Transaction Account Guarantee program during and after the Global Financial Crisis indicates that banks often adjusted reported uninsured deposits with a lag—suggesting that the imputed series may more accurately capture true uninsured-deposit dynamics in that period.

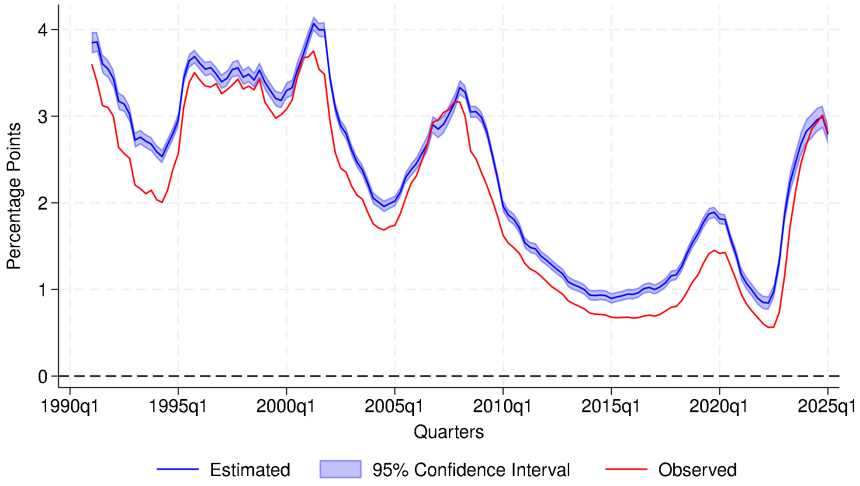
In addition to deposit rates and the share of uninsured deposits, we construct several other auxiliary variables using data from the Call Reports dataset, including the delinquency rate; tangible common equity (TCE), measured at book value; and a market-adjusted TCE (MATCE) that incorporates fair-value losses on both held-to-maturity (HTM) and available-for-sale (AFS) securities.

5 Validating the Empirical Methodology

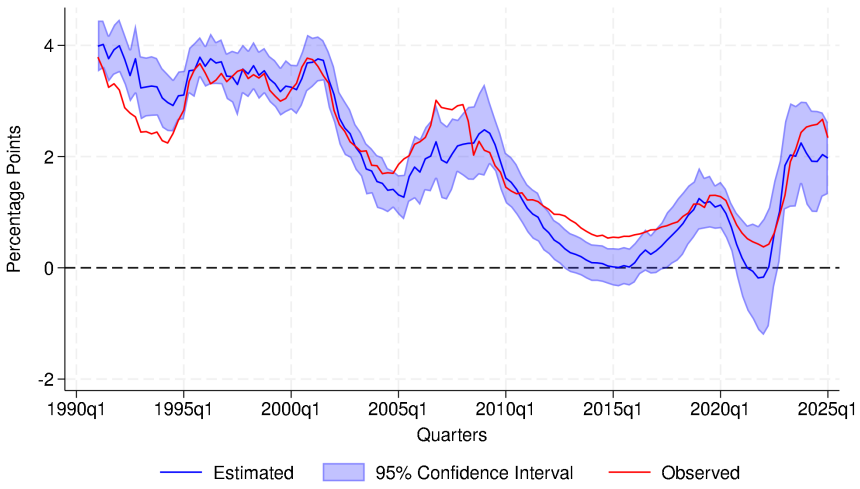
As discussed earlier, estimating interest rates by regressing the average interest rates on shares could generate biased results. Before presenting our main empirical analysis, we therefore first investigate the extent of these biases in data where we can directly compare our estimates to the observed interest rates. Specifically, we calculate the difference between the average deposit rates paid on time and demand deposits and the difference between the average deposit rates paid on non-transactional and transactional deposits and compare them to the coefficients obtained when estimating equation (5). Figures 1 and 2 show the results from this analysis, with panel a) showing results based on equally-weighted observations and panel b) showing results based on deposit-weighted observations.

Figure 1: Time Deposit Rate Premium – Estimated vs Observed

(a) Equally-weighted observations



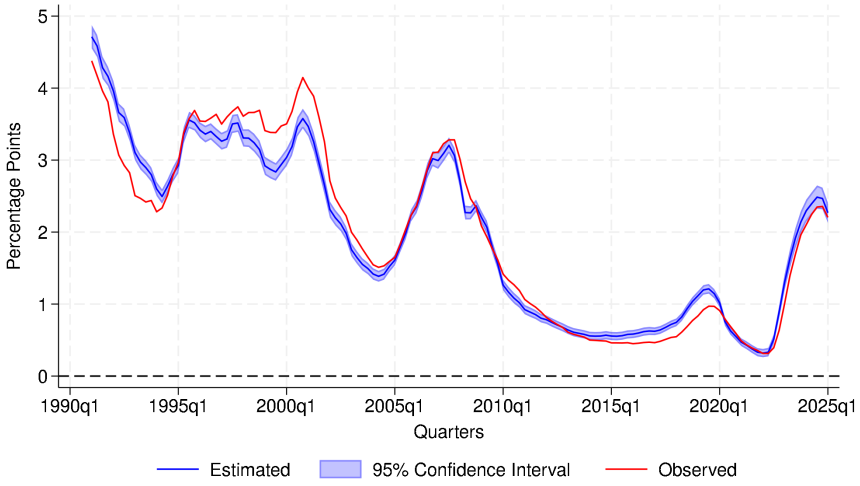
(b) Deposit-weighted observations



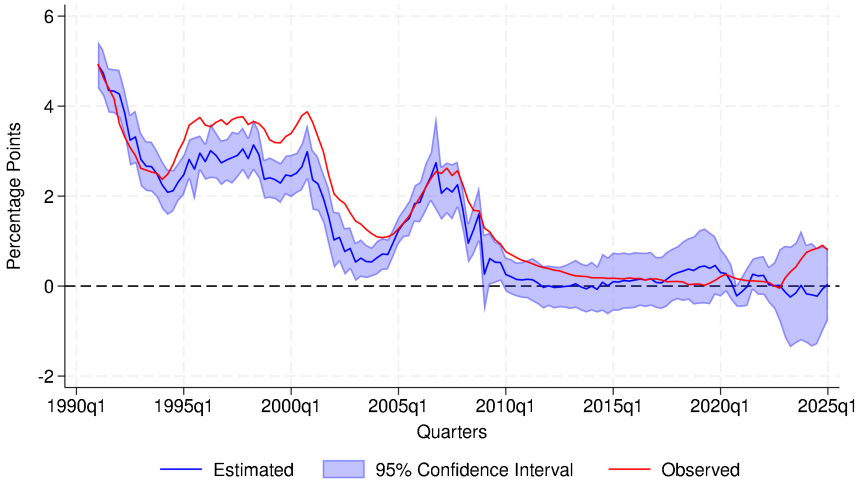
Note: Confidence intervals in shaded blue were constructed with clustered standard errors by bank.

Figure 2: Non-Transactional Deposit Rate Premium – Estimated vs Observed

(a) Equally-weighted observations



(b) Deposit-weighted observations



Note: Confidence intervals in shaded blue were constructed with clustered standard errors by bank.

The two figures show that for the case of time deposits and non-transaction accounts, the methodology can recover the interest rate premia almost perfectly, using only bank-level

average deposit rates and the share of one deposit type in total deposits. The results are consistent across specifications that assign equal weight to all observations and those that weight observations by each bank's deposit share. Although some estimation bias may be present, it appears modest and unlikely to affect the qualitative or quantitative conclusions. These findings show that the proposed method is indeed able to recover interest rates reliably in the context of bank deposits. While we cannot rule out that biases are larger in the case of the uninsured-insured deposit split, we think the robust performance in the observed categories is reassuring and supports the credibility of the main results presented in the next section.

6 Uninsured Deposit Premium Estimates

This section presents the main results from our analysis. We first present estimates of the uninsured deposit premium without controlling for deposit characteristics; then present estimates of the deposit premium that control for broad deposit characteristics; and finally discuss the role of changes in the insured and uninsured deposit rates in driving the estimated uninsured deposit premium.

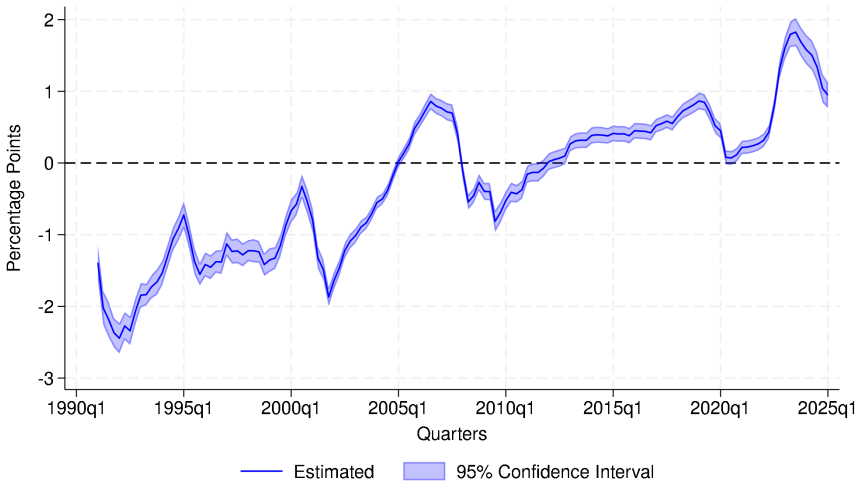
6.1 Uninsured Deposit Premium: Baseline Results

The first set of results is based on equation (5) and is summarized in Figure 3, where panel (a) presents estimates using equally weighted observations, and panel (b) presents estimates using deposit-weighted observations. We view the equally-weighted results as an estimate of the average premium on uninsured deposits paid by the average bank, regardless of size, while the deposit-weighted results are an estimate of the aggregate average uninsured deposit

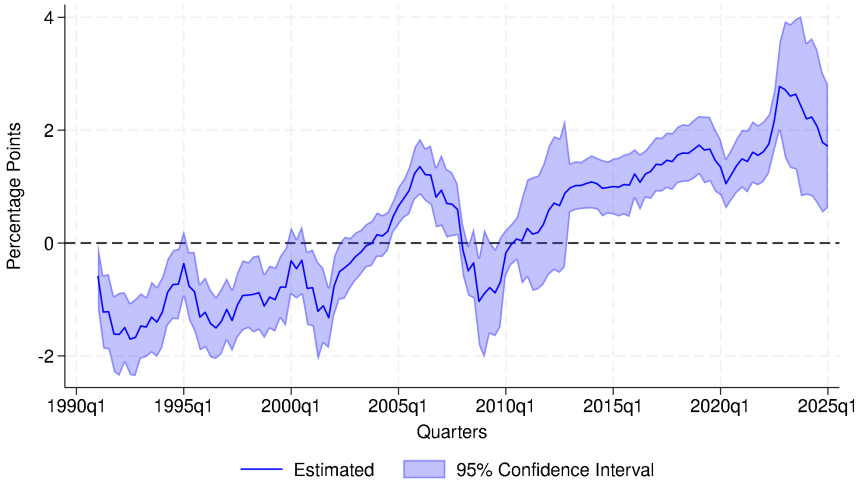
premium.

Figure 3: Uninsured Deposit Premium: Baseline

(a) Equally-weighted observations



(b) Deposit-weighted observations



Note: Confidence intervals in shaded blue were constructed with clustered standard errors by bank.

Two points are notable from Figure 3. First, there is a clear upward trend in the uninsured

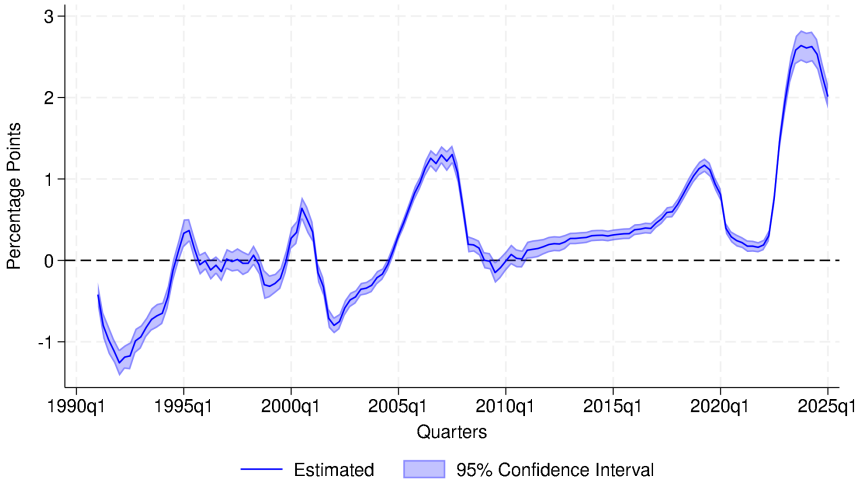
deposit premium over time, for both the equally-weighted and deposit-weighted results. The equally-weighted series starts around negative 2 percent in the early nineties and peaks around positive 2 percent in 2023, after the SVB failure. In the case of the deposit-weighted series, the series goes as low as negative 1.7 percentage points in 1992 and then reaches almost positive 2.8 percent in the first quarter of 2023, the same quarter that SVB failed. These estimates show that, without controlling for other deposit characteristics, the uninsured deposit premium increased over 400 basis points in the 34 year period covered by our sample, suggesting that financing through uninsured deposits has become much more expensive over time compared to the cost of insured deposits. Second, the uninsured deposit premium shows substantial cyclicality. Some of this cyclicality appears to be related to the monetary policy cycle, but this is something we explore in more detail in the next section.

6.2 Uninsured Deposit Premium Controlling for Deposit Characteristics

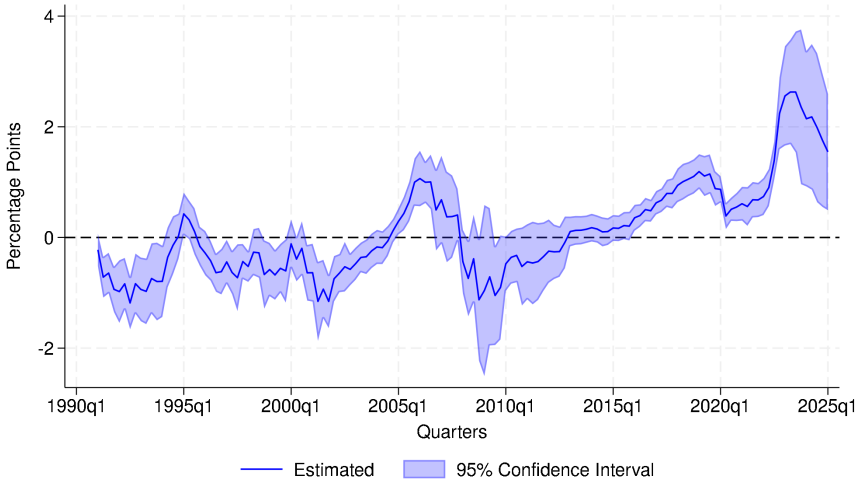
Some of the variation in the uninsured deposit premium estimates shown in Figure 3 could be caused by changes in the composition of bank deposits. To account for compositional effects, we now turn to the results that control for the share of time deposits and the share of non-transaction account deposits at the bank level as specified in equation (6). Again, panel (a) presents estimates based on equally weighted observations, while panel (b) presents estimates based on deposit-weighted observations.

Figure 4: Uninsured Deposit Premium: Controlling for Deposit Characteristics

(a) Equally-weighted observations



(b) Deposit-weighted observations



Note: Confidence intervals in shaded blue were constructed with clustered standard errors by bank.

The results shown in Figure 4 continue to display a significant upward trend in the uninsured deposit premium over the sample period. In this specification, the difference

between the equally-weighted and deposit-weighted estimates is much smaller. Both series reach a low of approximately –1 percent in the early 1990s and a peak of 2.6 percent in 2023. The peak occurs in 2023Q4 for the equally-weighted series and in 2023Q3 for the deposit-weighted series. Controlling for deposit characteristics slightly dampens the overall increase in the uninsured deposit premium, but both estimates still show an increase of just over 380 basis points from trough to peak.

Compared to the baseline results, controlling for deposit characteristics yields two notable differences. First, the post-SVB surge appears more pronounced relative to the historical trend. Second, the estimated uninsured deposit premium is less negative in the early part of the sample.

As in the baseline case, we continue to observe strong cyclical patterns in the uninsured deposit premium that appear be related to movements in U.S. monetary policy over the cycle.

6.3 Estimated Interest Rates

Before delving into what drove the uninsured deposit premium over our sample, we examine how much of its movement reflects changes in the average deposit rate on insured deposits versus the average rate on uninsured deposits.

To do so, we use the estimated time-fixed effects λ_t and the period-specific premium parameters β_t^U from equation (6) to back out the implied average rates on insured and uninsured deposits:

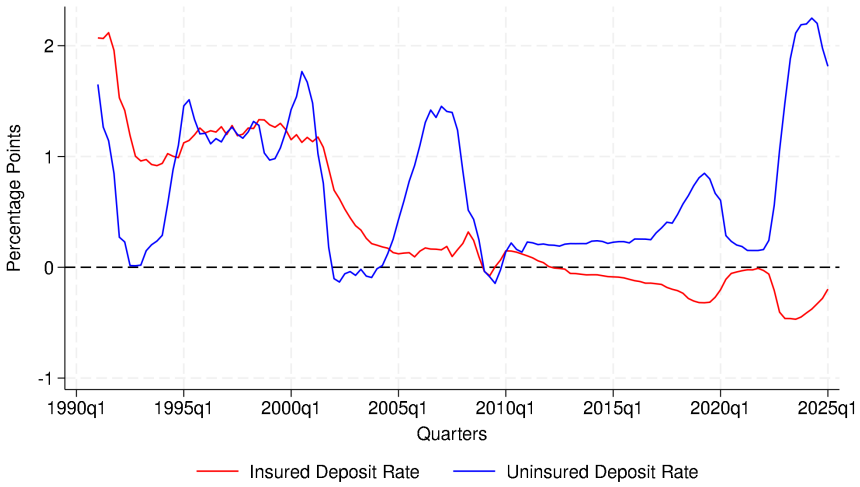
$$\widehat{r}_t^I = \hat{\lambda}_t \quad \text{and} \quad \widehat{r}_t^U = \hat{\lambda}_t + \hat{\beta}_t^U .$$

The term $\widehat{r}_t^I$ corresponds to the estimated average rate on insured deposits, while $\widehat{r}_t^U$ corre-

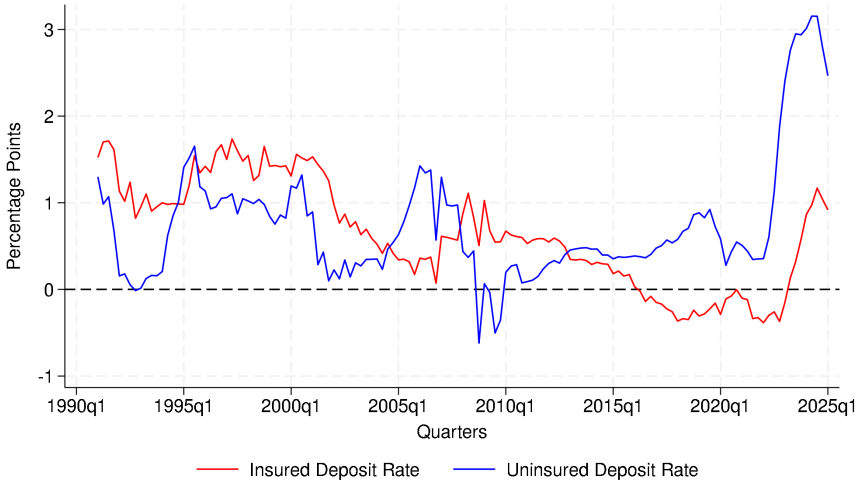
sponds to the estimated average rate on uninsured deposits. The estimated interest rates are shown in Figure 5.

Figure 5: Estimated Deposit Rates: Controlling for Deposit Characteristics

(a) Equally-weighted observations



(b) Deposit-weighted observations



In both panels of Figure 5, the average insured deposit rate falls over the sample period,

while the average uninsured deposit rate rises. By decomposing the bottom-to-top change in the uninsured deposit premium (Figure 4) into contributions from movements in the two series, we can gauge the relative importance of each. In panel (a) (equally weighted), half of the 380 basis-point increase in the uninsured premium stems from a decline in the insured rate, and half from an increase in the uninsured rate. In panel (b) (deposit weighted), only one quarter of the observed increase is driven by the insured rate's decline, while three quarters reflect the uninsured rate's rise.

While the equally-weighted and deposit-weighted regressions yield very comparable interest rate estimates over most of the sample period, there is an intriguing difference for the most recent 2022-23 rate hike cycle. Both insured and uninsured interest rates increase by more based on the weighted regressions than the unweighted regressions. This difference implies that the largest banks had a higher pass through than the average bank over the 2022-23 rate hike cycle. This finding aligns with anecdotal evidence that at least one large bank aggressively expanded its retail deposit market footprint with deposit accounts that paid interest rates close to the fed funds rate.

7 What Drives the Uninsured Deposit Premium?

Why did the uninsured deposit premium change so much over time, and what explains its cyclical behavior? To address these questions, we first develop a simple model of deposit pricing that lays out the various forces—such as depositors' alternative savings instruments, banks' other funding sources, market competition, and regulatory backstops—that can influence banks' willingness to pay for uninsured funds. We then bring these theoretical predictions to the data, testing which of these channels accounts for the observed swings in the

premium.

7.1 Conceptional Framework

Basic Setting. A bank raises one unit of deposits from depositors. An exogenous fraction of these deposits ϕ^i is insured and the remainder $\phi^u = 1 - \phi^i$ is uninsured. The bank can invest the deposits for return $\alpha(r^{FFR})$, with $\alpha(\cdot)$ being a strictly increasing function with respect to the federal funds rate (r^{FFR}). The bank pays deposit rates r^u and r^i for uninsured and insured depositors, respectively. The outside option of the bank is to borrow on the wholesale market at the federal funds rate. Let the outside option of uninsured and insured depositors be given by $out^u(r^{FFR})$, and $out^i(r^{FFR})$, respectively.

We model equilibrium deposit rates—insured and uninsured—as the outcome of a bargaining problem between depositors and banks, where β^i, β^u, and $\beta^b = 1 - \beta^i - \beta^u$ are the respective bargaining weights of insured depositors, uninsured depositors, and the bank; the corresponding surpluses (net of outside options) are

$$\begin{aligned} S_i &= \phi^i(r^i - out^i(r^{FFR})), \\ S_u &= \phi^u(r^u - out^u(r^{FFR})), \\ S_b &= \left(\alpha(r^{FFR}) - \phi^i r^i - \phi^u r^u\right) - (\alpha(r^{FFR}) - r^{FFR}) = r^{FFR} - X, \end{aligned}$$

where $X = \phi^i r^i + \phi^u r^u$ represents the average deposit rate per unit of deposits paid by the bank. The Nash-bargaining objective is

$$\mathcal{L} = S_b^{\beta^b} S_i^{\beta^i} S_u^{\beta^u}. \tag{10}$$

Equilibrium Deposit Rates Solving the Nash-bargaining problem, we can derive the optimal deposit rates as[8]

$$r^i = out^i(r^{FFR}) + \frac{\beta^i}{\phi^i}\big(r^{FFR} - out^i(r^{FFR}) - out^u(r^{FFR})\big), \tag{11}$$

$$r^u = out^u(r^{FFR}) + \frac{\beta^u}{\phi^u}\big(r^{FFR} - out^i(r^{FFR}) - out^u(r^{FFR})\big). \tag{12}$$

We summarize our findings on the deposit rates in the following proposition.

Proposition 1 (Deposit Rates)**.** *Deposit rates for uninsured and insured depositors*

i) increase with their own depositor group's normalized bargaining weights but are independent of the other group's bargaining weights.

ii) increase with their own outside option, with the strength of this relationship decreasing in their bargaining weight; and decrease with the other group's outside option.

Proof. Follows directly from equations (11) and (12). □

The Uninsured Deposit Premium We can now derive the uninsured deposit premium as

$$r^u - r^i = out^u(r^{FFR}) + \tfrac{\beta^u}{\phi^u}(r^{FFR} - O(r^{FFR})) - out^i(r^{FFR}) - \tfrac{\beta^i}{\phi^i}(r^{FFR} - O(r^{FFR})), \tag{13}$$

$$= \underbrace{out^u(r^{FFR}) - out^i(r^{FFR})}_{\text{outside-option gap}} + (r^{FFR} - O(r^{FFR})) \underbrace{\left(\frac{\beta^u}{\phi^u} - \frac{\beta^i}{\phi^i}\right)}_{\text{surplus-share gap}}. \tag{14}$$

[8] Taking logs of equation (10) delivers $\ln \mathcal{L} = \beta^b \ln S_b + \beta^i \ln S_i + \beta^u \ln S_u$. Differentiating with respect to r^i and r^u, setting equal to zero and rewriting delivers $\beta^i\, S_b = \beta^b\, S_i$ and $\beta^u\, S_b = \beta^b\, S_u$.. Then $S_i + S_u = X - O$. Summing the two FOCs, $(\beta^i + \beta^u)\, S_b = \beta^b\, (S_i + S_u) \implies (1 - \beta^b)\, S_b = \beta^b\, (X - O)$. Since $S_b = r^{FFR} - X$, substitute $r^{FFR} - X = \frac{\beta^b}{1-\beta^b}\,(X - O) \implies X = (1 - \beta^b)\, r^{FFR} + \beta^b\, O$. Hence, $S_B = r^{FFR} - X = \beta^b\, (r^{FFR} - O)$. From $\beta^i\, S_b = \beta^b\, S_i$, we get $S_i = \frac{\beta^i}{\beta^b}\, S_b = \beta^i\, (r^{FFR} - O)$. But $S_i = \phi^i(r^i - out^i(r^{FFR}))$, so $\phi^i r^i = \phi^i out^i(r^{FFR}) + \beta^i\, (r^{FFR} - O)$.

There are two elements that influence the uninsured deposit premium. First, the outside-option gap, $out^u(r^{FFR}) - out^i(r^{FFR})$, which measures how much higher (per unit) the uninsured depositor's reservation value is relative to the insured depositor's reservation value. Second, the surplus-share gap, $\frac{\beta^u}{\phi^u} - \frac{\beta^i}{\phi^i}$, which measures the difference in normalized bargaining weight per unit between the two depositor types, scaling the total surplus $r^{FFR} - O$. To shed light on the role of the FFR, take the partial derivative

$$\frac{\partial(r^u - r^i)}{\partial r^{FFR}} = out^u(r^{FFR})' - out^i(r^{FFR})' + \Big(\frac{\beta^u}{\phi^u} - \frac{\beta^i}{\phi^i}\Big)(1 - out^u(r^{FFR})' - out^i(r^{FFR})')$$

Rewriting shows that the uninsured deposit premium increases with the FFR if and only if

$$\frac{\beta^u}{\phi^u} - \frac{\beta^i}{\phi^i} > \frac{out^i(r^{FFR})' - out^u(r^{FFR})'}{1 - out^u(r^{FFR})' - out^i(r^{FFR})'} \tag{15}$$

Consider the case where the outside option of uninsured depositors changes at least as much with the FFR as the outside option of insured depositors ($out^u(r^{FFR})' \geq out^i(r^{FFR})'$). In that case, a sufficient condition for the uninsured deposit premium to co-move positively with the FFR is a positive surplus-share gap. That is, that uninsured depositors have a higher bargaining power than insured depositors. As shown in equation (15), the necessary condition for a positive co-movement between the uninsured deposit rate and the FFR is weaker. For example, the outside option of insured depositors could co-move more with the FFR, as long as this effect is compensated by a sufficiently large difference in bargaining weights. Broadly speaking, the uninsured deposit premium will co-move positively with the FFR if the bargaining weights are sufficiently higher for uninsured depositors or if the outside option of uninsured depositors co-moves more strongly with the FFR. We can derive several

predictions as summarized in the following proposition.

Proposition 2 (Uninsured Deposit Premium)**.** *The uninsured deposit premium*

i) increases (decreases) with the bargaining weight and the outside option of uninsured (insured) depositors.

ii) increases (decreases) with the FFR

i) (special case) *if uninsured depositors have a larger (smaller) normalized bargaining weight than uninsured depositors and the outside option of uninsured depositors co-moves at least (most) as much with the FFR as that of insured depositors.*

ii) (general case) *if* $\frac{\beta^u}{\phi^u} - \frac{\beta^i}{\phi^i} > \frac{out^i(r^{FFR})' - out^u(r^{FFR})'}{1 - out^u(r^{FFR})' - out^i(r^{FFR})'}$

Proof. Follows directly from equations (14) and (15). □

7.2 Empirical Results on the Drivers of the Uninsured Deposit Premium

In this subsection we empirically explore channels that may affect the uninsured deposit premium. The conceptional framework suggests a potential link between the uninsured deposit premium and the FFR. In addition, it suggests that factors that affect the outside options of insured and uninsured depositors and their respective bargaining weights should impact the uninsured deposit premium.

We start by looking at the role of monetary policy in shaping the uninsured deposit premium. We then consider three additional channels that could explain the uninsured deposit premium and its variation over time. The first is directly linked to the conceptional

framework. We test if the increasing availability of money market funds has contributed to the increase in the uninsured deposit premium in line with the idea that this development disproportionately benefited the outside option of uninsured depositors.

Going beyond the current conceptual framework, we also look at the role of risk. The price of uninsured deposits could be driven by changes in actual or perceived bank risk. This channel has received increased attention since the SVB collapse in 2023. However, bank failures are rare and uninsured depositors often have very high recovery rates in default.[9]

7.2.1 The Role of Monetary Policy

To assess the role of monetary policy in the variation of the uninsured deposit premium, we regress our estimates of the uninsured deposit premium based on the estimation of the baseline (equation 5) and the extended models (equation 6) on the FFR. For each model, we run four specifications. The first specification regresses the level of the uninsured deposit premium on the level of the FFR. The second specification adds a time trend. The third and fourth specifications estimate the relationship in first differences. The results of this exercise are presented in Table 1.

The baseline correlation shown in column (1) is negative and not significantly different from zero. However, once we add a time trend in column (2), we obtain the expected positive relationship. That is, a higher FFR is associated with a higher uninsured deposit premium. This relationship becomes even larger once we look at first differences in columns (3) and (4).

[9]The Federal Deposit Insurance Corporation (2023) reports that between 1992 and 2022, the unconditional loss rate for uninsured depositors when a bank failed was 6 percent. Conditional on being an uninsured depositor at a failing bank with uninsured depositor losses (20 percent of failures), the average conditional loss rate was 28 percent. Interestingly, the unconditional loss rate declined for the period following the Global Financial Crisis (2008-2022) compared to the period preceding it (1992-2007) from 10 percent to 3 percent.

Table 1: Uninsured Deposit Premium and Monetary Policy

	Baseline Model				Multi-Share Model			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Fed Funds Rate	-0.113	0.124***			0.0599	0.255***		
	(0.103)	(0.0363)			(0.0828)	(0.0245)		
Time Trend (dt)		0.0259***		-0.000452		0.0213***		-0.000236
		(0.00186)		(0.000290)		(0.00130)		(0.000194)
Δ Fed Funds Rate			0.275***	0.281***			0.320***	0.323***
			(0.0202)	(0.0242)			(0.0282)	(0.0306)
Constant	0.0178	-5.593***	0.0214*	0.108*	0.127	-4.492***	0.0228***	0.0682*
	(0.239)	(0.389)	(0.0125)	(0.0614)	(0.151)	(0.250)	(0.00849)	(0.0407)
Observations	137	137	136	136	137	137	136	136
R^2	0.062	0.845	0.468	0.478	0.028	0.877	0.702	0.705

Note: Standard errors robust to serial correlation and heteroskedasticity are in parentheses. *, **, and *** denote significance at the 10%, 5%, and 1% level.

Moving to the multi-share model in columns (5) through (8) strengthens the relationship further. Based on the estimate in column (7), a 100 bps increase in the FFR raises the uninsured deposit premium by 32 basis points.

The strong positive correlation between the FFR and the uninsured deposit premium confirms the prediction of our framework (Proposition 2, ii). Of note, this relationship also holds in first differences (columns (3) and (4) of the baseline model, and columns (7) and (8) of the multi-share model). The high R^2 values in these specifications point to a rapid and robust pass-through from the policy rate to the uninsured deposit premium. This stands in contrast to the broader literature, which generally finds that policy rate transmission to deposit rates is gradual.

A natural question raised by the results in Table 1 is whether the co-movement of the uninsured deposit premium with the federal funds rate is driven by changes in insured or uninsured deposit rates. To check this, we regress the estimated insured and uninsured deposit rates on the FFR, using our two preferred specifications from table 1 (Columns (6)

and (7)). Results are presented in Table 2.

Table 2: Insured and Uninsured Deposit Rates and Monetary Policy

	Insured Rate		Uninsured Rate	
	(1)	(2)	(3)	(4)
Fed Funds Rate	0.0607***		0.315***	
	(0.0220)		(0.0175)	
Time Trend (dt)	-0.0130***		0.00837***	
	(0.00105)		(0.000809)	
Δ Fed Funds Rate		0.0157		0.336***
		(0.0376)		(0.0263)
Constant	2.713***	-0.0164**	-1.780***	0.00639
	(0.218)	(0.00835)	(0.193)	(0.00528)
Observations	137	136	137	136
R^2	0.854	0.009	0.942	0.791

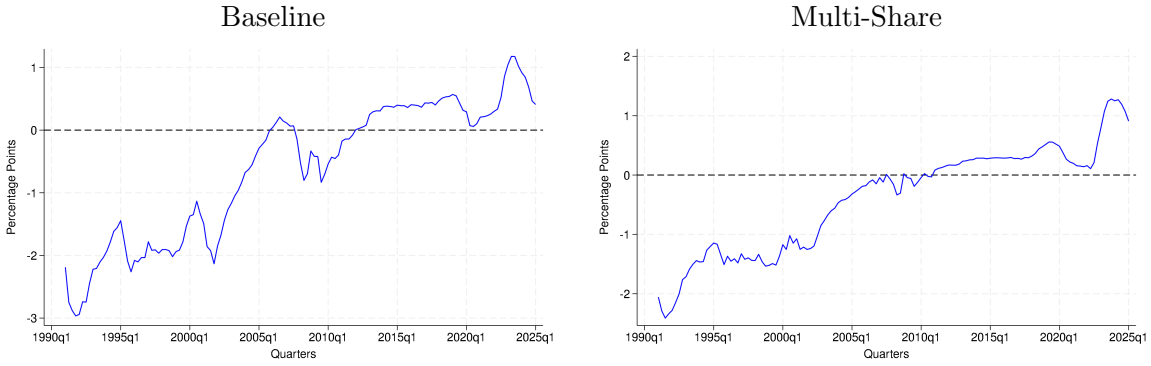
Note: Standard errors robust to serial correlation and heteroskedasticity are in parentheses. *, **, and *** denote significance at the 10%, 5%, and 1% level.

The table reveals some interesting patterns. First, the uninsured rate co-moves much more strongly with the FFR than the insured rate. This is true in the level regression with trend where the uninsured rate coefficient in Column (3) is about 5 times larger than the insured rate coefficient in Column (1). In first differences, the contrast is even starker, with changes in in the FFR only explaining the uninsured rate. Second, the time trend in the uninsured deposit premium consists of a positive time trend in the uninsured rate and a negative time trend in the insured rate. Quantitatively, the estimated downward trend in the insured rate in Column (1) is larger than the upward trend in the uninsured rate in Column (3). To summarize, the results of Table 2 suggest that the strong correlation between the uninsured deposit premium and the federal funds that the results of Table 1 show result

mostly from a strong correlation between the uninsured deposit rate and the federal funds rate.

Another intriguing question is whether there is any pattern in the uninsured deposit premium that is not explained by the federal funds rate. To answer this question, we take our uninsured deposit premium estimates and subtract the part explained by the federal funds rate using the coefficient associated with the federal funds rate in Columns (2) and (4) of Table 1 for the baseline and multi-share models, respectively. We call this object the residual uninsured deposit premium. Results are presented in Figure 6.

Figure 6: Residual Uninsured Deposit Premium



There are three main takeaways. First, there is still an upward trend in the premium that is independent of monetary policy. This result is not too surprising given the results for the time trend shown in Table 1. Second, there are periods when the premium is stable and others when it moves up significantly in discrete steps. Finally, even after controlling for the role of monetary policy, we still find a relatively substantial increase in the premium around the SVB failure in the first quarter of 2023. This result suggests that not all of the recent increase in the uninsured deposit premium can be explained by the sharp rise in the FFR

over the most recent hiking cycle. That is, the analysis suggests that some of the recent peak in the uninsured deposit premium could result from the banking sector stresses following the failure of SVB in March 2023.

7.2.2 The Role of Alternative Investment Opportunities

Measuring outside options of depositors is difficult. One key factor that affects the outside option of depositors is the availability of alternative investment opportunities. In the following, we exploit the rising relevance of money market funds as a factor that alters the outside option of depositors. Of note, while money market funds are very safe, they are not insured and are hence closer substitutes to uninsured than to insured deposits. In addition, uninsured depositors have larger deposits and are in general more sophisticated - especially if insured depositors are mostly retail whereas uninsured are mostly corporate investors. Therefore, one would expect the outside option of uninsured depositors to be more affected by the increasing availability of money market funds.

To test for the role of money market funds, we construct the ratio of overall money market fund assets in the economy over total banking sector liabilities. We then augment the regressions in Tables 1 and 2 with this ratio to explain the time-series variation in the estimated uninsured deposit premium. Results are presented in Table 3.

The coefficient in Column (2) implies that the uninsured deposit premium increases quarter-on-quarter when the size of money market investments changes relative to banking sector liabilities. As shown in Columns (4) and (6), effects are mainly driven by the estimated uninsured rate. Specifically, the estimated effect of the money market fund over bank liabilities ratio on the deposit rate is positive and significant for uninsured deposits

Table 3: Uninsured Deposit Premium and Money Market Funds

	UDP		Insured Rate		Uninsured Rate	
	(1)	(2)	(3)	(4)	(5)	(6)
Δ Fed Funds Rate	0.320***	0.324***	0.0157	0.0147	0.336***	0.339***
	(0.0282)	(0.0266)	(0.0376)	(0.0348)	(0.0263)	(0.0249)
Δ MMF Investments / Bank Liabilities		0.0228**		-0.00541		0.0174*
		(0.0102)		(0.00781)		(0.00965)
Constant	0.0228***	0.0212**	-0.0164**	-0.0160*	0.00639	0.00514
	(0.00849)	(0.00885)	(0.00835)	(0.00930)	(0.00528)	(0.00578)
Observations	136	136	136	136	136	136
R^2	0.702	0.714	0.009	0.013	0.791	0.798

Note: Standard errors robust to serial correlation and heteroskedasticity are in parentheses. *, **, and *** denote significance at the 10%, 5%, and 1% level.

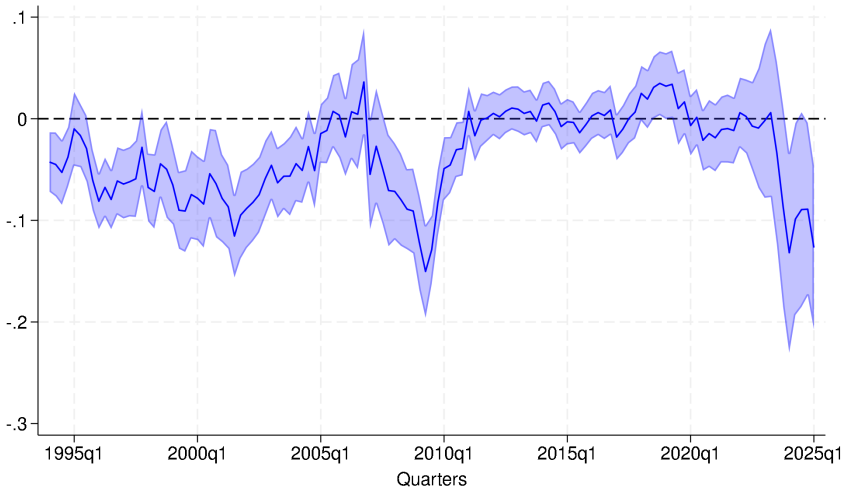
and negative but insignificant for insured deposits.

7.2.3 The Role of Bank Risk

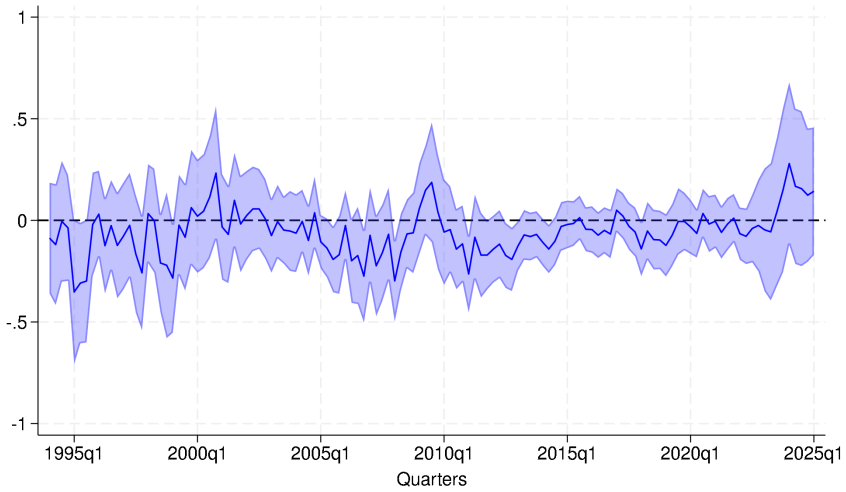
So far, we have discussed how monetary policy and alternative investment opportunities may explain the time variation in the uninsured deposit premium. We now turn to the role that bank risk may play in the uninsured deposit premium paid by banks. For this analysis we estimate equation (7) for the cases of banks grouped by delinquency rate and MATCE ratio. The estimation results for the parameters pertaining to the terciles of banks with the highest delinquency rate and lowest MATCE ratios are shown in Figures 7 and 8, respectively.

Figure 7: Effect of Delinquency Rates on Insured Deposit Rates and the Uninsured Deposit Premium

(a) Effect on insured deposit rate



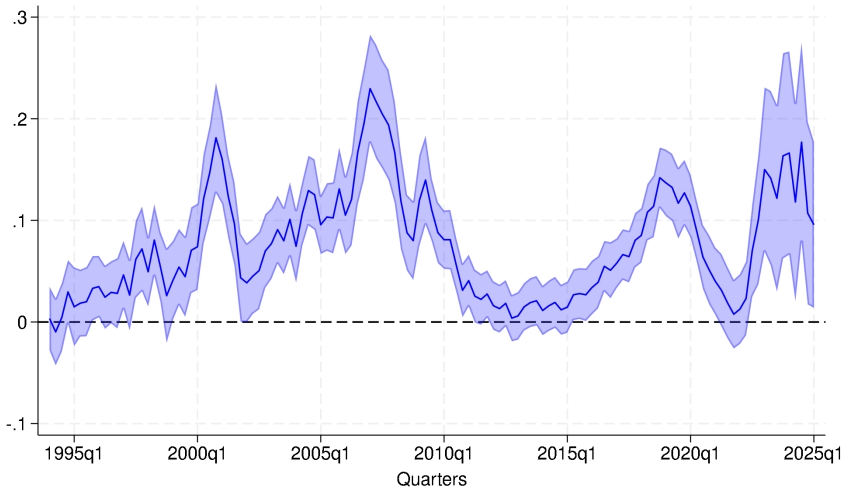
(b) Effect on uninsured deposit premium



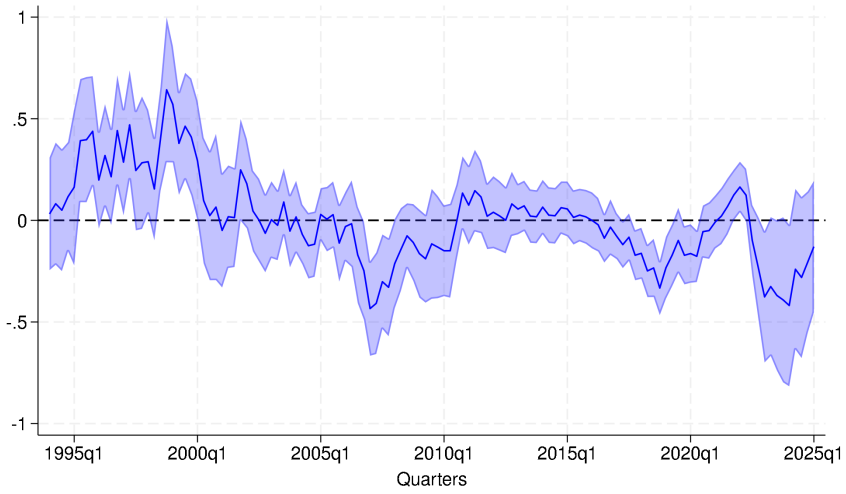
Note: Confidence intervals in shaded blue were constructed with clustered standard errors by bank.

Figure 8: Effect of Bank Capitalization on Insured Deposit Rates and the Uninsured Deposit Premium

(a) Effect on insured deposit rate



(b) Effect on uninsured deposit premium



Note: Confidence intervals in shaded blue were constructed with clustered standard errors by bank.

For both the delinquency rate and the MATCE ratio, panel a) of Figures 7 and 8 shows that, at times, riskier banks—those with higher delinquency rates or lower MATCE ra-

tios—pay higher average rates on insured deposits than less risky banks. Panel b), however, indicates that the uninsured deposit premium does not vary systematically with bank risk. This is especially clear for the delinquency rate: banks with higher delinquency rates pay roughly the same uninsured premium as safer banks. For the MATCE ratio, the estimates show more variation but still no consistent pattern.

Taken together, the results in Figures 7 and 8 suggest that riskier banks tend to offer higher overall deposit rates, but not higher uninsured deposit premiums. Put differently, bank risk is reflected in the average rate paid on deposits, but not in the differential between rates on uninsured versus insured deposits.

8 Conclusions

Data on uninsured and insured deposit rates are rarely available directly, posing a significant challenge in knowing how much banks are paying depositors for insured and uninsured deposits. To address this gap, our paper proposes a method to estimate the average difference between these two rates. In cases where deposit rates are reported by bank type, our approach is validated with empirical evidence—demonstrating its ability to recover the uninsured deposit premium despite the data limitations.

With the help of this methodology, we provide a first estimate of a time-varying uninsured deposit premium. Our results show that between 1991 and 2025, the average uninsured deposit premium rose by nearly 400 basis points, driven equally by a decline in the average insured deposit rate and a rise in the average uninsured deposit rate. Furthermore, our analysis reveals a strong link between monetary policy and deposit pricing: a one percentage point increase in the Federal Funds Rate is associated with an average increase of about 32

basis points in the uninsured premium.

We present a simple framework that rationalizes the co-movement between the uninsured deposit rate and the fed funds rate with a bargaining model. The model suggests that the upward trend in the uninsured deposit premium reflects shifts in the bargaining power in favor of uninsured depositors compared to insured depositors. Our evidence that links the growth of money market funds to the increase in the uninsured deposit premium is consistent with this mechanism.

To the best of our knowledge, both the large increase in the uninsured deposit premium over time and its strong relationship with the FFR are new facts that have not been documented in this form before. It will be important to further explore the drivers behind these facts and the uninsured deposit premium more broadly both in the time series and the cross section.

References

Bartholdy, J., Boyle, G. W., and Stover, R. D. (2003). Deposit insurance and the risk premium in bank deposit rates. *Journal of Banking & Finance*, 27(4):699–717.

Berger, A. N. and Hannan, T. H. (1989). The price-concentration relationship in banking. *The Review of Economics and Statistics*, 71(2):291–299.

Calem, P. S. and Carlino, G. A. (1991). The concentration/conduct relationship in bank deposit markets. *The Review of Economics and Statistics*, 73(2):268–276.

DeMarzo, P. M., Krishnamurthy, A., and Nagel, S. (2024). Interest rate risk in banking. Technical report, National Bureau of Economic Research.

Demirgüç-Kunt, A. and Kane, E. J. (2002). Deposit insurance around the globe: Where does it work? *Journal of Economic Perspectives*, 16(2):175–195.

Diamond, D. W. and Dybvig, P. H. (1983). Bank runs, deposit insurance, and liquidity. *Journal of Political Economy*, 91(3):401–419.

Drechsler, I., Savov, A., and Schnabl, P. (2017). The deposits channel of monetary policy. *Quarterly Journal of Economics*, 132(4):1819–1876.

Drechsler, I., Savov, A., and Schnabl, P. (2021). Banking on deposits: Maturity transformation without interest rate risk. *The Journal of Finance*, 76(3):1091–1143.

Drechsler, I., Savov, A., Schnabl, P., and Wang, O. (2023). Banking on uninsured deposits. *Available at SSRN 4411127*.

Drechsler, I., Savov, A., Schnabl, P., and Wang, O. (2025). Deposit franchise runs. mimeo, New York University.

Egan, M., Hortaçsu, A., and Matvos, G. (2017). Deposit competition and financial fragility: Evidence from the us banking sector. *American Economic Review*, 107(1):169–216.

Federal Deposit Insurance Corporation (2023). Options for deposit insurance reform. Technical Report —, Federal Deposit Insurance Corporation, Washington, D.C.

Goldstein, I. and Pauzner, A. (2005). Demand–deposit contracts and the probability of bank runs. *the Journal of Finance*, 60(3):1293–1327.

Hannan, T. H. and Adams, R. M. (2011). Consumer switching costs and firm pricing: Evidence from bank pricing of deposit accounts. *Journal of Industrial Economics*, 59(2):296–320.

Hannan, T. H. and Berger, A. N. (1991). The rigidity of prices: Evidence from the banking industry. *American Economic Review*, 81(4):938–945.

Jacewitz, S. and Pogach, J. (2018). Deposit rate advantages at the largest banks. *Journal of Financial Services Research*, 53(1):1–35.

Kim, M., Kliger, D., and Vale, B. (2003). Estimating switching costs: the case of banking. *Journal of Financial Intermediation*, 12(1):25–56.

Klein, M. A. (1971). A theory of the banking firm. *Journal of money, credit and banking*, 3(2):205–218.

Maechler, A. M. and McDill, K. M. (2006). Dynamic Depositor Discipline in U.S. Banks. *Journal of Banking & Finance*, 30(7):1871–1898.

Martinez-Peria, M. S. and Schmukler, S. L. (2001). Do depositors punish banks for bad behavior? market discipline, deposit insurance, and banking crises. *Journal of Finance*, 56(3):1029–1051.

Monti, M. (1972). *Deposit, credit and interest rate determination under alternative bank objective function*.

Neumark, D. and Sharpe, S. A. (1992). Market structure and the nature of price rigidity: Evidence from the market for consumer deposits. *Quarterly Journal of Economics*, 107(2):657–680.

Park, K. and Pennacchi, G. (2009). Harming depositors and helping borrowers: The disparate impact of bank consolidation. *Review of Financial Studies*, 22(1):1–40.

Rosen, R. J. (2007). Banking market conditions and deposit interest rates. *Journal of Banking & Finance*, 31(12):3862–3884.

Sharpe, S. A. (1997). The effect of consumer switching costs on prices: A theory and its application to the bank deposit market. *Review of Industrial Organization*, 12(1):79–94.

A Appendix: Approximating the OLS Bias (Pooled and Fixed Effects)

This appendix approximates the bias in OLS estimates for equations 4 5 (simple regression without and with bank fixed effects, respectively) and 6 (extended model). We consider four cases: pooled OLS with only time fixed effects, and OLS with both time and bank fixed effects, for each of the two models. Although our main analysis does not estimate the pooled OLS specification without bank fixed effects, we include its bias approximation here for illustrative purposes.

A.1 Simple Regression Model

A.1.1 Exact identity

Let DR_{it}^{I} and DR_{it}^{U} be the rates on insured (I) and uninsured (U) deposits for bank i in period t and let α_{it}^{U} be the uninsured share. The bank's average deposit rate satisfies the exact identity

$$DR_{it} \;=\; (1-\alpha_{it}^{U})\,DR_{it}^{I} \;+\; \alpha_{it}^{U}\,DR_{it}^{U} \;=\; DR_{it}^{I} \;+\; \alpha_{it}^{U}\big(DR_{it}^{U}-DR_{it}^{I}\big).$$

Denote simple cross–sectional means at t by $\overline{DR}_{t}^{I} \equiv \frac{1}{N}\sum_{i=1}^{N} DR_{it}^{I}$ and $\overline{DR}_{t}^{U} \equiv \frac{1}{N}\sum_{i=1}^{N} DR_{it}^{U}$.

A.1.2 Pooled OLS representation and bias

Add and subtract $\overline{DR}_t^I$ and $\overline{DR}_t^U$:

$$DR_{it} \;=\; \underbrace{\overline{DR}_t^I}_{\theta_t} \;+\; \underbrace{\left(\overline{DR}_t^U - \overline{DR}_t^I\right)}_{\beta_t}\, \alpha_{it}^U \;+\; \epsilon_{it}^{pool},$$

where

$$\epsilon_{it}^{pool} \;=\; \underbrace{\left(DR_{it}^I - \overline{DR}_t^I\right)}_{\text{insured-rate deviation}} \;+\; \underbrace{\alpha_{it}^U \Big[(DR_{it}^U - DR_{it}^I) - (\overline{DR}_t^U - \overline{DR}_t^I) \Big]}_{\text{share} \times \text{spread-deviation term}}.$$

Thus, the pooled (time–FE) regression is

$$DR_{it} \;=\; \theta_t \;+\; \beta_t\, \alpha_{it}^U \;+\; \epsilon_{it}^{pool}.$$

The period–t pooled OLS slope satisfies

$$\widehat{\beta}_t^{OLS} \;=\; \beta_t \;+\; \frac{Cov_i\big(\alpha_{it}^U,\, \epsilon_{it}^{pool}\big)}{Var_i(\alpha_{it}^U)}.$$

Linear approximation of the second term in ϵ_{it}^{pool}. Write the second (underbraced) term as $\alpha_{it}^U\, S_{it}$ with

$$S_{it} \equiv (DR_{it}^U - DR_{it}^I) - (\overline{DR}_t^U - \overline{DR}_t^I), \qquad \frac{1}{N}\sum_{i=1}^{N} S_{it} = 0.$$

Approximate this term linearly around the average share $\overline{\alpha}_t^U \equiv \frac{1}{N}\sum_{i=1}^N \alpha_{it}^U$

$$\alpha_{it}^U S_{it} \;\approx\; \overline{\alpha}_t^U\, S_{it} \;=\; \overline{\alpha}_t^U\, (DR_{it}^U - \overline{DR}_t^U) \;-\; \overline{\alpha}_t^U\, (DR_{it}^I - \overline{DR}_t^I),$$

so that, to first order (no products remain),

$$\epsilon_{it}^{pool} \;\approx\; \underbrace{(1-\overline{\alpha}_t^U)\,(DR_{it}^I - \overline{DR}_t^I)}_{\text{insured-rate component}} \;+\; \underbrace{\overline{\alpha}_t^U\,(DR_{it}^U - \overline{DR}_t^U)}_{\text{uninsured-rate component}}.$$

Therefore, the pooled bias becomes

$$\widehat{\beta}_t^{OLS} - \beta_t \;\approx\; \frac{(1-\overline{\alpha}_t^U)\,Cov_i(\alpha_{it}^U,\, DR_{it}^I) \;+\; \overline{\alpha}_t^U\,Cov_i(\alpha_{it}^U,\, DR_{it}^U)}{Var_i(\alpha_{it}^U)}.$$

Interpretation: a weighted average (weights $1-\overline{\alpha}_t^U$ and $\overline{\alpha}_t^U$) of the cross–sectional slopes of DR^I and DR^U on α^U.

A.1.3 From identity to the fixed–effects model

Add and subtract bank means to expose time–invariant heterogeneity. Let

$$\overline{DR}_i^I \equiv \frac{1}{T}\sum_{t=1}^{T} DR_{it}^I, \qquad \overline{DR}_i^U \equiv \frac{1}{T}\sum_{t=1}^{T} DR_{it}^U, \qquad \overline{\alpha}_i^U \equiv \frac{1}{T}\sum_{t=1}^{T} \alpha_{it}^U, \qquad \overline{\overline{DR}}^I \equiv \frac{1}{T}\frac{1}{N}\sum_{n=1}^{N}\sum_{t=1}^{T} DR_{it}^I,$$

and define $\lambda_i \equiv \overline{DR}_i^I - \overline{\overline{DR}}^I$. Then

$$DR_{it} \;=\; \lambda_i \;+\; \theta_t \;+\; \beta_t\,\alpha_{it}^U \;+\; u_{it},$$

with

$$u_{it} \;=\; \underbrace{\Big[(DR_{it}^I - \overline{DR}_t^I) - (\overline{DR}_i^I - \overline{\overline{DR}}^I)\Big]}_{\text{demeaned insured-rate deviation}} \;+\; \underbrace{\alpha_{it}^U\Big[(DR_{it}^U - DR_{it}^I) - (\overline{DR}_t^U - \overline{DR}_t^I)\Big]}_{\text{share} \times \text{spread-deviation term}}.$$

A.1.4 Fixed–effects bias

By Frisch–Waugh–Lovell, the time–t FE slope is the slope from regressing bank–demeaned variables

$$\widehat{\beta}_t^{FE} \;=\; \beta_t \;+\; \frac{Cov_i\big(\alpha_{it}^U - \overline{\alpha}_i^U,\; u_{it} - \overline{u}_i\big)}{Var_i\big(\alpha_{it}^U - \overline{\alpha}_i^U\big)}.$$

Linear approximation of the second term in u_{it}. As above, approximate the second (underbraced) term linearly around $\overline{\alpha}_t^U$

$$\alpha_{it}^U\Big[(DR_{it}^U - DR_{it}^I) - (\overline{DR}_t^U - \overline{DR}_t^I)\Big] \;\approx\; \overline{\alpha}_t^U\,(DR_{it}^U - \overline{DR}_t^U) \;-\; \overline{\alpha}_t^U\,(DR_{it}^I - \overline{DR}_t^I).$$

After demeaning across t within bank i (constants vanish in covariances), this yields

$$u_{it} - \overline{u}_i \;\approx\; \underbrace{(1-\overline{\alpha}_t^U)\,(DR_{it}^I - \overline{DR}_i^I)}_{\text{within insured-rate component}} + \underbrace{\overline{\alpha}_t^U\,(DR_{it}^U - \overline{DR}_i^U)}_{\text{within uninsured-rate component}}.$$

Therefore,

$$\widehat{\beta}_t^{FE} - \beta_t \;\approx\; \frac{(1-\overline{\alpha}_t^U)\,Cov_i\big(\alpha_{it}^U - \overline{\alpha}_i^U,\; DR_{it}^I - \overline{DR}_i^I\big) \;+\; \overline{\alpha}_t^U\,Cov_i\big(\alpha_{it}^U - \overline{\alpha}_i^U,\; DR_{it}^U - \overline{DR}_i^U\big)}{Var_i\big(\alpha_{it}^U - \overline{\alpha}_i^U\big)}.$$

Interpretation: the same weighted–average structure as in pooled OLS, but now entirely in terms of within–bank covariances and variance.

A.1.5 Which bias is smaller?

In general, without restrictions on how much of $Cov(\alpha^U, DR^X)$ comes from between–bank versus within–bank variation ($X \in \{I, U\}$), one cannot sign the FE bias relative to the

pooled bias. If most comovement is between banks and is therefore removed by θ_i, the FE bias tends to be smaller in magnitude; if comovement is primarily within a bank over time, the two biases can be similar, or the FE bias can be larger.

A.2 Extended Model

A.2.1 Exact identity over latent buckets

Partition deposits into the eight latent buckets

$$c \in \{I, U\} \times \{D, TD\} \times \{Tr, NTr\}.$$

Let α_{it}^c be bank i's share in c at time t (so $\sum_c \alpha_{it}^c = 1$) and let DR_{it}^c be the corresponding rate. The bank's average rate satisfies

$$DR_{it} \;=\; \sum_c \alpha_{it}^c \, DR_{it}^c.$$

The observable marginals are

$$\alpha_{it}^U = \sum_c \alpha_{it}^c I_U(c), \quad \alpha_{it}^{TD} = \sum_c \alpha_{it}^c I_{TD}(c), \quad \alpha_{it}^{NTr} = \sum_c \alpha_{it}^c I_{NTr}(c).$$

A.2.2 Additive–premia assumption (no interactions)

Index the eight latent buckets by

$$c \in \{I, U\} \times \{D, TD\} \times \{Tr, NTr\},$$

and let $I_U(c), I_{TD}(c), I_{NTr}(c) \in \{0,1\}$ be the corresponding indicators. Define the bank–time specific *baseline* bucket rate

$$DR_{it}^0 \;\equiv\; DR_{it}^{(I,D,Tr)}.$$

The *additive premia* restriction posits that, at each time t, there exist common premia $(\beta_t, \gamma_t, \delta_t)$ such that for every bank i and bucket c

$$DR_{it}^c \;=\; DR_{it}^0 \;+\; \beta_t\, I_U(c) \;+\; \gamma_t\, I_{TD}(c) \;+\; \delta_t\, I_{NTr}(c) \;+\; \eta_{it}^c,$$

where η_{it}^c captures bucket–specific idiosyncrasies (e.g., measurement noise or residual cell heterogeneity). This restriction rules out $U \times TD$, $U \times NTr$, and $TD \times NTr$ *interaction premia*: only the U, TD, and NTr premia appear, and they are the same regardless of the other two attributes. It does *not* assume that DR_{it}^c are constant across banks; DR_{it}^0 is allowed to vary with (i,t).

Substituting into the exact identity $DR_{it} = \sum_c \alpha_{it}^c DR_{it}^c$ and collecting by the *observable* marginals α_{it}^U, α_{it}^{TD}, and α_{it}^{NTr} yields

$$DR_{it} \;=\; DR_{it}^0 \;+\; \beta_t\, \alpha_{it}^U \;+\; \gamma_t\, \alpha_{it}^{TD} \;+\; \delta_t\, \alpha_{it}^{NTr} \;+\; \sum_c \alpha_{it}^c\, \eta_{it}^c.$$

With time fixed effects λ_t (and, if included, bank fixed effects θ_i), set $\lambda_t = \overline{DR}_t^0$ and write the regression error as

$$\epsilon_{it} \;=\; (DR_{it}^0 - \overline{DR}_t^0) \;+\; \sum_c \alpha_{it}^c\, \eta_{it}^c,$$

which makes explicit that it is the *premia* $(\beta_t, \gamma_t, \delta_t)$ that are common across buckets, while the baseline level DR_{it}^0 and the bucket idiosyncrasies η_{it}^c remain bank–time specific.

A.2.3 Pooled time–FE representation and residual

Under the additive–premia restriction,

$$DR_{it} \;=\; \theta_t \;+\; \beta_t\,\alpha_{it}^{U} \;+\; \gamma_t\,\alpha_{it}^{TD} \;+\; \delta_t\,\alpha_{it}^{NTr} \;+\; \epsilon_{it}^{pool},$$

with (period–t)

$$\epsilon_{it}^{pool} \;=\; \underbrace{(DR_{it}^{0} - \overline{DR}_{t}^{0})}_{\text{baseline dev.}} \;+\; \underbrace{\alpha_{it}^{U}\Big[(DR_{it}^{U} - DR_{it}^{0}) - (\overline{DR}_{t}^{U} - \overline{DR}_{t}^{0})\Big]}_{U \text{ prem. dev. } \times \text{ share}}$$
$$+\; \underbrace{\alpha_{it}^{TD}\Big[(DR_{it}^{TD} - DR_{it}^{0}) - (\overline{DR}_{t}^{TD} - \overline{DR}_{t}^{0})\Big]}_{TD \text{ prem. dev. } \times \text{ share}} \;+\; \underbrace{\alpha_{it}^{NTr}\Big[(DR_{it}^{NTr} - DR_{it}^{0}) - (\overline{DR}_{t}^{NTr} - \overline{DR}_{t}^{0})\Big]}_{NTr \text{ prem. dev. } \times \text{ share}}.$$

Linear approximation of the three underbraced terms. For compactness define the premium deviations (all mean zero across i at t):

$$S_{it}^{U} \equiv (DR_{it}^{U} - DR_{it}^{0}) - (\overline{DR}_{t}^{U} - \overline{DR}_{t}^{0}),$$
$$S_{it}^{TD} \equiv (DR_{it}^{TD} - DR_{it}^{0}) - (\overline{DR}_{t}^{TD} - \overline{DR}_{t}^{0}),$$
$$S_{it}^{NTr} \equiv (DR_{it}^{NTr} - DR_{it}^{0}) - (\overline{DR}_{t}^{NTr} - \overline{DR}_{t}^{0}).$$

Let $\overline{\alpha}_t^k \equiv E_i[\alpha_{it}^k]$ for $k \in \{U, TD, NTr\}$ and set[10]

$$w_t^0 \equiv 1 - \overline{\alpha}_t^{U} - \overline{\alpha}_t^{TD} - \overline{\alpha}_t^{NTr}.$$

[10] w_t^0 arises from the first–order linearization; it is not a share and has no standalone economic interpretation. Because the three dimensions along which deposits are defined overlap, w_t^0 can be positive or negative. Its sign only indicates whether the baseline component $(DR_{it}^0 - \overline{DR}_t^0)$ enters the first–order residual with

Apply a first–order linear approximation to each share–premium product:

$$\alpha_{it}^{k} S_{it}^{k} \;\approx\; \overline{\alpha}_{t}^{k}\, S_{it}^{k}, \qquad k \in \{U, TD, NTr\}.$$

Substituting gives a product–free residual:

$$\begin{aligned}\epsilon_{it}^{pool} \;\approx\; & w_t^0\,(DR_{it}^0 - \overline{DR}_t^0) \;+\; \overline{\alpha}_t^U\,(DR_{it}^U - \overline{DR}_t^U) \\ & +\; \overline{\alpha}_t^{TD}\,(DR_{it}^{TD} - \overline{DR}_t^{TD}) \;+\; \overline{\alpha}_t^{NTr}\,(DR_{it}^{NTr} - \overline{DR}_t^{NTr}).\end{aligned}$$

A.2.4 Pooled OLS bias (three shares)

Within period t, obtain $\tilde{\alpha}_{it}^U$ as the residual from regressing α_{it}^U on a constant, α_{it}^{TD}, and α_{it}^{NTr} (Frisch–Waugh–Lovell); define $\tilde{\alpha}_{it}^{TD}$ and $\tilde{\alpha}_{it}^{NTr}$ analogously. Then

$$\begin{aligned}\widehat{\beta}_t^{OLS} - \beta_t \;\approx\; & \frac{w_t^0\, Cov_i(\tilde{\alpha}_{it}^U,\, DR_{it}^0) \;+\; \overline{\alpha}_t^U\, Cov_i(\tilde{\alpha}_{it}^U,\, DR_{it}^U)}{Var_i(\tilde{\alpha}_{it}^U)} \\ & +\; \frac{\overline{\alpha}_t^{TD}\, Cov_i(\tilde{\alpha}_{it}^U,\, DR_{it}^{TD}) \;+\; \overline{\alpha}_t^{NTr}\, Cov_i(\tilde{\alpha}_{it}^U,\, DR_{it}^{NTr})}{Var_i(\tilde{\alpha}_{it}^U)}.\end{aligned}$$

By symmetry,

$$\begin{aligned}\widehat{\gamma}_t^{OLS} - \gamma_t \;\approx\; & \frac{w_t^0\, Cov_i(\tilde{\alpha}_{it}^{TD},\, DR_{it}^0) \;+\; \overline{\alpha}_t^U\, Cov_i(\tilde{\alpha}_{it}^{TD},\, DR_{it}^U)}{Var_i(\tilde{\alpha}_{it}^{TD})} \\ & +\; \frac{\overline{\alpha}_t^{TD}\, Cov_i(\tilde{\alpha}_{it}^{TD},\, DR_{it}^{TD}) \;+\; \overline{\alpha}_t^{NTr}\, Cov_i(\tilde{\alpha}_{it}^{TD},\, DR_{it}^{NTr})}{Var_i(\tilde{\alpha}_{it}^{TD})},\end{aligned}$$

$$\begin{aligned}\widehat{\delta}_t^{OLS} - \delta_t \;\approx\; & \frac{w_t^0\, Cov_i(\tilde{\alpha}_{it}^{NTr},\, DR_{it}^0) \;+\; \overline{\alpha}_t^U\, Cov_i(\tilde{\alpha}_{it}^{NTr},\, DR_{it}^U)}{Var_i(\tilde{\alpha}_{it}^{NTr})} \\ & +\; \frac{\overline{\alpha}_t^{TD}\, Cov_i(\tilde{\alpha}_{it}^{NTr},\, DR_{it}^{TD}) \;+\; \overline{\alpha}_t^{NTr}\, Cov_i(\tilde{\alpha}_{it}^{NTr},\, DR_{it}^{NTr})}{Var_i(\tilde{\alpha}_{it}^{NTr})}.\end{aligned}$$

the same or opposite sign.

A.2.5 Fixed–effects bias (three shares)

Include bank fixed effects θ_i and work with demeaned shares and rates:

$$\alpha_{it}^{k} - \overline{\alpha}_i^{k}, \quad DR_{it}^{X} - \overline{DR}_i^{X}, \qquad k \in \{U, TD, NTr\},\ X \in \{0, U, TD, NTr\}.$$

Let $\tilde{\alpha}_{it}^{U,w}$ be the residual from regressing $(\alpha_{it}^{U} - \overline{\alpha}_i^{U})$ on a constant, $(\alpha_{it}^{TD} - \overline{\alpha}_i^{TD})$, and $(\alpha_{it}^{NTr} - \overline{\alpha}_i^{NTr})$ (within period t). Define $\tilde{\alpha}_{it}^{TD,w}$ and $\tilde{\alpha}_{it}^{NTr,w}$ analogously.

The linearized FE residual is

$$\begin{aligned} u_{it} - \overline{u}_i \;\approx\; & w_t^0 \,(DR_{it}^0 - \overline{DR}_i^0) \;+\; \overline{\alpha}_t^U \,(DR_{it}^U - \overline{DR}_i^U) \\ & +\; \overline{\alpha}_t^{TD} \,(DR_{it}^{TD} - \overline{DR}_i^{TD}) \;+\; \overline{\alpha}_t^{NTr} \,(DR_{it}^{NTr} - \overline{DR}_i^{NTr}). \end{aligned}$$

Hence, the first–order FE bias approximations for all three slopes are

$$\begin{aligned} \widehat{\beta}_t^{FE} - \beta_t \;\approx\; & \frac{w_t^0 \, Cov_i\big(\tilde{\alpha}_{it}^{U,w},\, DR_{it}^0 - \overline{DR}_i^0\big) \;+\; \overline{\alpha}_t^U \, Cov_i\big(\tilde{\alpha}_{it}^{U,w},\, DR_{it}^U - \overline{DR}_i^U\big)}{Var_i(\tilde{\alpha}_{it}^{U,w})} \\ & +\; \frac{\overline{\alpha}_t^{TD} \, Cov_i\big(\tilde{\alpha}_{it}^{U,w},\, DR_{it}^{TD} - \overline{DR}_i^{TD}\big) \;+\; \overline{\alpha}_t^{NTr} \, Cov_i\big(\tilde{\alpha}_{it}^{U,w},\, DR_{it}^{NTr} - \overline{DR}_i^{NTr}\big)}{Var_i(\tilde{\alpha}_{it}^{U,w})}, \end{aligned}$$

$$\begin{aligned} \widehat{\gamma}_t^{FE} - \gamma_t \;\approx\; & \frac{w_t^0 \, Cov_i\big(\tilde{\alpha}_{it}^{TD,w},\, DR_{it}^0 - \overline{DR}_i^0\big) \;+\; \overline{\alpha}_t^U \, Cov_i\big(\tilde{\alpha}_{it}^{TD,w},\, DR_{it}^U - \overline{DR}_i^U\big)}{Var_i(\tilde{\alpha}_{it}^{TD,w})} \\ & +\; \frac{\overline{\alpha}_t^{TD} \, Cov_i\big(\tilde{\alpha}_{it}^{TD,w},\, DR_{it}^{TD} - \overline{DR}_i^{TD}\big) \;+\; \overline{\alpha}_t^{NTr} \, Cov_i\big(\tilde{\alpha}_{it}^{TD,w},\, DR_{it}^{NTr} - \overline{DR}_i^{NTr}\big)}{Var_i(\tilde{\alpha}_{it}^{TD,w})}, \end{aligned}$$

$$\begin{aligned} \widehat{\delta}_t^{FE} - \delta_t \;\approx\; & \frac{w_t^0 \, Cov_i\big(\tilde{\alpha}_{it}^{NTr,w},\, DR_{it}^0 - \overline{DR}_i^0\big) \;+\; \overline{\alpha}_t^U \, Cov_i\big(\tilde{\alpha}_{it}^{NTr,w},\, DR_{it}^U - \overline{DR}_i^U\big)}{Var_i(\tilde{\alpha}_{it}^{NTr,w})} \\ & +\; \frac{\overline{\alpha}_t^{TD} \, Cov_i\big(\tilde{\alpha}_{it}^{NTr,w},\, DR_{it}^{TD} - \overline{DR}_i^{TD}\big) \;+\; \overline{\alpha}_t^{NTr} \, Cov_i\big(\tilde{\alpha}_{it}^{NTr,w},\, DR_{it}^{NTr} - \overline{DR}_i^{NTr}\big)}{Var_i(\tilde{\alpha}_{it}^{NTr,w})}. \end{aligned}$$

Interpretation. Each bias is a weighted average (weights w_t^0, $\overline{\alpha}_t^U$, $\overline{\alpha}_t^{TD}$, $\overline{\alpha}_t^{NTr}$) of simple covariances between the residualized share and the component rates. Signs will typically differ across components; offsets can make the net bias small. The pooled vs. FE comparison again depends on whether these covariances are driven mainly by between–bank or within–bank variation.