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The Effect of State Solvency on Bank Values and Credit Supply: Evidence from State Pension Cut Legislation

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Abstract

We find the financial condition of states impacts bank credit supply through their municipal bond holdings. In particular, we treat sudden political and statutory actions during the 2011 union bargaining rights debates in Wisconsin and Ohio as exogenous shocks to state solvency. We show bank valuations and municipal bond spreads adjust to the announcements, and, over longer horizons, a new lending channel linked to state solvency emerges, whereby banks supply credit as municipal bond appreciations free up capital.

I. Introduction

In the late 2000s, the United States, and indeed the world, experienced the worst financial crisis since the 1930s and the Great Depression. As financial markets crumbled, jobs were lost and consumer spending slowed to a crawl. With the stagnation of consumer spending, tax revenues at the state levels experienced their steepest decline ever. Without these tax revenues, state budgets experienced massive deficits; fiscal year 2011 budget deficits were estimated to exceed $100 billion.[1] Further, total outstanding state debt surpassed $4.2 trillion in 2011.[2] As large as they were, these budget deficits were overshadowed in size and scope by

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[1]"State Spending Restraint: An Analysis of the Path Not Taken," Mercatus Center at George Mason University, Aug. 2010.

[2]"State Budget Solutions, Real Solutions for Real Budget Problems," Oct. 25, 2011.

unfunded liabilities in state pension and health care systems for public employees. On average, in fiscal 2010, state pension plans had 75% of the assets needed to cover long-term benefits owed to government workers. These fiscal 2010 deficits came after even greater deterioration directly after the financial crisis. The average state pension fund was 78% funded in fiscal 2009 and 84% funded in fiscal 2008.[3]

These pension funding gaps spurred many states to pursue changes to their retirement systems. Wisconsin and Ohio were the first states to see pension reform legislation signed into law (Wisconsin on Mar. 11, 2011 and Ohio on Mar. 31, 2011).[4] By the end of Mar. 2011, 13 other states had introduced legislation to reduce pension benefits. Associated with changes in pension benefits were changes in some collective bargaining rights of state employees: Most government employees are union organized. Pension contributions and benefits are a part of the collective bargaining process between states and their unionized employees. Governors stated that they needed changes in collective bargaining rights in order to cut pension obligations. Thus, these pension reform bills became synonymous with the abolition of public union bargaining rights and were termed "anti-union bills."

It is well documented that changes in regulations affect bank values. (See Akhigbe and Martin (2006), Cornett, Ors, and Tehranian (2002), and Cornett, Razaee, and Tehranian (1996) for just three examples.) However, the effect of state pension cuts and the diminishment or abolishment of union activities on bank values in the states has not been examined. This is particularly important, as banks play a vital role in economic growth in the states in which they operate, at the very least through lending channels. This paper explores the link between states' attempts to improve state fiscal conditions through pension cuts and the value and activities of banks operating in those states. Specifically, we analyze the impact of Wisconsin and Ohio pension cut legislation on values and activities of banks operating in Wisconsin and Ohio relative to similar publicly traded banks operating in other Midwestern states.

Legislative efforts aimed at reducing pension guarantees to public employees and at reducing public employees' collective bargaining rights are seen as a way of addressing states' huge budget deficits. Successful passage of this type of legislation arguably improves states' budget situations. Further, while public employees would experience direct adverse financial effects of pension reform, a shift in employee pension contributions away from states' balance sheets could stimulate economic activity by lowering state borrowing costs and offsetting tax revenues. Banks, in their role as lenders to municipalities, businesses, and consumers in the states, and as investors in municipal securities issued by the state, would directly feel the effects of improved budget conditions. However, it is also possible that with the passage of this legislation, pension cuts would mean that public employees and retirees have less income and less financial capacity to borrow from banks. To that extent, banks would be hurt as lending falls and bank values would shrink with the passage of pension cut legislation. Hence, at face

[3]"State Pension Plans Lose Ground," The Pew Center on States, Apr. 2011.

[4]However, Ohio's reform was repealed in Nov. 2011.

value, the impact of pension cut legislation on bank values and operations is not clear-cut. Regardless, it is the case that bank values and operations would be affected by public-sector pension policy and the solvency conditions of the states in which they operate.

We find that banks doing business in Wisconsin and Ohio experience positive (negative) stock price reactions to announcements that indicate an increased (decreased) probability of pension cut legislation. The stock price reactions are positively related to the extent to which banks operate in Wisconsin and Ohio and to the value of municipal securities held in the portfolios of banks operating heavily in Wisconsin and Ohio. We also find that Wisconsin and Ohio municipal bond spreads decrease (increase) significantly in reaction to announcements that indicate an increased (decreased) probability of pension cut legislation.

These effects signify a special connection between bank values and state solvency through bank municipal bond holdings.[5] Specifically, state solvency fundamentally interacts with bank liquidity and credit supply as municipal bond revaluations create balance sheet slack.[6] We find that total lending by banks operating in Wisconsin increases over and after the period in which pension cut legislation is enacted, while total lending by banks operating in Ohio decreases over and after the period in which pension cut legislation is enacted and, eventually, overturned. We find these different lending patterns are largely a function of bank municipal bond holdings. Finally, we find that the increased lending by Wisconsin banks comes from riskier portfolios with lower loan loss provisioning. This suggests a lending channel linked to state solvency, whereby capital-constrained banks disproportionately supply credit as municipal bond value appreciations free up capital. Thus, the paper documents a new channel of transmission of shocks from banks to borrowers through state reforms.

A number of papers have examined how bank capital shocks affect lending. For example, Chava and Purnanandam (2011) use an exogenous shock to the U.S. banking system during the Russian crisis of 1998 to provide causal evidence that adverse capital shocks to banks affect their borrowers' performance negatively. Consistent with an adverse shock to the supply of credit, crisis-affected banks decreased the quantity of their lending and increased loan interest rates in the

[5]We examine the effect of pension cut legislation on industrial firms and find no evidence that the increased probability of pension legislation passed in Wisconsin and Ohio impacts stock prices of industrial firms headquartered in these two states. The stock price reactions, thus, appear unique to banks operating in the two affected states.

[6]Recent policy revisions support this notion. In Oct. 2014, the U.S. Treasury published rule 79 FR 61439, which increased liquidity coverage requirements across all federal banks and restricted the eligible set of high-quality liquid assets (HQLAs) available for banks to satisfy the new requirements. The exclusion of bank municipal bond holdings from HQLA status elicited particular criticism during a commenting period before the final rule became effective on Jan. 1, 2015 (OCC 2014, http://www.occ.gov/news-issuances/news-releases/2014/nr-ia-2014-120.html, p. 14). In Feb. 2016, The U.S. House of Representatives passed the Messer–Maloney Bill (H.R.2209), intended to reduce the liquidity coverage requirements for all federal banks by "[requiring] the appropriate Federal banking agencies to treat certain municipal obligations as level 2A liquid assets," a classification in line with other HQLAs such as U.S. agency securities. Further congressional voting has not occurred to date. However, Federal Reserve regulators began to specifically relax the HQLA restrictions on member-bank municipal bond holdings on July 1, 2016 (Federal Reserve 2016, https://www.federalreserve.gov/newsevents/press/bcreg/20160401a.htm).

post-crisis period significantly more than the unaffected banks. Similarly, Schnabl (2012) exploits the 1998 Russian default as a negative liquidity shock to international banks and analyzes its transmission to Peru. After the shock, international banks reduce bank-to-bank lending to Peruvian banks and Peruvian banks reduce lending to Peruvian firms. The results suggest that international banks transmit liquidity shocks across countries and that negative liquidity shocks reduce bank lending in affected countries. Adelino and Ferreira (2016) study the causal effect of bank credit rating downgrades on the supply of bank lending. They examine the asymmetric impact of sovereign downgrades on ratings of banks bounded by the sovereign rating of their home country relative to banks that are not bound as a result of rating agencies' sovereign ceiling policies. This asymmetric effect leads to greater reductions in rating-sensitive funding and lending of banks at the bound relative to other banks.

Our results complement Abowd and Lemieux (1993), who exploit exogenous shocks to firms' product markets to isolate the impact of firms' financial situations on union wage bargaining power. Just as their paper addresses endogeneity in the relationship between profitability and negotiated wages, we exploit exogenous variations in pension reform and union bargaining power to highlight a reverse causality effect: Union wage bargaining power and a reduction in pension benefits can impact financial situations in budget constrained states, particularly through the banking channel.

The remainder of the paper is organized as follows. Section II recaps the pension cut legislative process in Wisconsin and Ohio and presents hypotheses. Section III describes the data and methodology used in the analysis. Section IV discusses the results of the analysis. Finally, Section V concludes the paper.

II. Pension Cut Legislation

The financial crisis and the resulting economic recession of the late 2000s caused the steepest decline in state tax receipts on record. As of the third quarter of 2010, state tax collections, adjusted for inflation, were 11% below pre-recession levels. However, the need for state-funded services did not decline. As a result, even after making very deep spending cuts, states continued to face large budget gaps. Forty-four states and the District of Columbia were projecting budget shortfalls for fiscal year 2012, which began July 1, 2011. These deficits followed large shortfalls in fiscal years 2009–2011. The budget difficulties led at least 46 states to reduce services to their residents and over 30 states to raise taxes to at least some degree, in some cases quite significantly.[7]

As projected revenue remained depressed at low levels, many states looked to additional spending and service cuts as well as further increases in taxes. However, states recognized that spending cuts are problematic during an economic downturn because they reduce overall demand and can result in the downturn getting even deeper. Spending cuts result in, among other things, employee layoffs, cancellations of contracts with vendors, and cutting of benefit payments to

[7]"States Continue to Feel Recession's Impact," Center on Budget and Policy Priorities, Mar. 9, 2011.

individuals, all of which directly remove demand from the state's economy. Tax increases also remove demand from the economy by reducing the amount of money people have to spend. Thus, increasing spending cuts and taxes to reduce state budget shortfalls places a considerable number of jobs at risk.

As the recession dragged on, states, struggling to further cut spending and increase taxes to address budget deficits, also began to seriously consider reducing state employee benefits. Approximately 80% of state, county, and city budget deficits are the result of employee costs.[8] Further, the Bureau of Labor Statistics at the U.S. Department of Labor estimates that, as of 2009, state and local government employees received benefits that were 69% higher than those in the private sector (this on top of wages that topped their private sector counterparts).[9] Without reform, these disparities were projected to increase. For example, in the late 2000s, the lifetime employment costs of a government worker in the state of Ohio were 221% higher than that for a private sector employee. In 2011, Ohio had an $8 billion budget gap. Data showed that realigning Ohio state worker compensation packages to match those of their private sector peers would save taxpayers over $2.1 billion in 2011 and 2012 (or 28% of the $8 billion deficit).[10] Thus, as states looked for ways to reduce budget deficits (particularly ways that would not put jobs at risk), many states turned to pension reform. Since most state workers are unionized, states needed to reduce or remove collective bargaining rights of these state employees to accomplish the desired pension reform. By Mar. 2011, 15 states had proposed legislation that would remove some collective bargaining powers from unionized state workers and allow states to reduce their contributions to public employee pension and health care plans. Wisconsin and Ohio were the first two states to actually pass reform legislation (Wisconsin on Mar. 11, 2011 and Ohio on Mar. 31, 2011).

Public employees saw these changes in their collective bargaining rights as an assault on their rights as union workers. Thus, these pension reform bills became associated with anti-union movements intended to weaken the power of labor unions. In Feb. 2011, a series of public employee protests began. A number of Wisconsin legislators and judges were sympathetic to the unions' arguments. These politicians and judges took actions to prevent passage and enactment of the Wisconsin reform bill: most notably was the walkout of 14 Wisconsin Senate Democrats who fled the state to deprive Republicans of the three-fifths majority needed to pass a reform bill. Further, while Wisconsin's reform bill was signed into law on Mar. 11, 2011, just 5 days later, the Dane County District Attorney filed a legal challenge to the bill, stating that Republican lawmakers violated Wisconsin's open meetings law (by not giving the proper public notice that the committee planned to meet) when they amended the plan. The challenge requested that a Dane County Circuit Court judge void the law and issue an emergency order blocking the secretary of state from publishing the law. One week later, the judge temporarily blocked the law from taking effect. Then, on Mar. 31, the law was put on indefinite hold by the same judge until the case could be heard by

[8]"Property Tax Woes Mean More Local Budget Pain," Public Sector, Inc., Mar. 14, 2012.
[9]U.S. Bureau of Labor Statistics, Employer Costs for Employee Compensation survey, Dec. 2009.
[10]The Buckeye Institute for Public Policy Solutions, press release, July 2010.

the Wisconsin Supreme Court. Thus, at this point, the pension reform bill would not go into effect. On May 27, 2011, the Dane County Judge issued a permanent injunction that effectively threw out the pension cut bill. Despite these actions to overturn the bill, on June 15, 2011, the Wisconsin Supreme Court rejected the ruling from the county court that invalidated the bill, and 2 weeks later, on June 29, 2011, the pension cut bill officially took effect.

Ohio's pension cut bill moved quickly through the legislative process. On Feb. 1, 2011, Bill 5 (the pension cut/anti-union bill) was introduced in the Ohio Senate, and the bill was signed into law on Mar. 31, 2011. However, after this relatively swift passage, opponents of the bill started the repeal process, and on Nov. 8, 2011, the bill was overwhelmingly repealed: by a vote of 61% in favor of repeal to 39% against repeal.

Despite union protests and legislative actions to prevent changes in collective bargaining rights and pension reform from enactment, the soaring levels of state employee benefits were seen as unsustainable cost drivers that threatened the financial solvency of many states. Thus, states continued their pursuit of pension reform. The reform in Wisconsin and Ohio requires state workers to contribute a larger share of their pension and health care costs. In Wisconsin, state workers would make a 5% contribution to their pensions and increase their share of health insurance costs up to 12%. Most state workers contributed nothing to their pensions and paid between 4% and 6% of their health insurance costs prior to passage of the reform bill. The reform was expected to save the state $1.5 billion. The state had a budget deficit of $3.6 billion at the time. In Ohio, the bill bans strikes by public workers and establishes penalties for those who participate in walkouts. Unionized workers can negotiate wages, hours, and certain work conditions but not health care, sick time, or pension benefits. The reform was expected to produce savings of $2.1 billion dollars. The state had a budget deficit of $8 billion at the time.

Given the expected costs savings, changes in union employees' bargaining rights and pension reform in Wisconsin and Ohio should result in a decrease in the strain on the states' budgets. Thus, successful passage of this type of legislation arguably improves the general economic conditions in the states. Banks, in their role as lenders to municipalities, businesses, and consumers in the states, would directly feel the effects of changes in the economic condition of the states in which they operate.

For example, this legislation should increase the probability that states can make their promised payments on municipal securities issued by the state. This, in turn, should result in municipal securities issued by the states of Wisconsin and Ohio being less risky. Thus, through holding state-issued bonds, banks would directly feel the effects of changes in the fiscal solvency of the states. Accordingly, events that signal an increased (decreased) probability of passage of pension reform (anti-union) legislation in Wisconsin and Ohio should result in positive (negative) value changes for banks that operate in these states and hold municipal securities issued by them. Likewise, improved state solvency prospects may also result in greater long-term capacity to lend for banks holding state-issued bonds. This would be particularly the case for more capital-constrained banks where municipal bond appreciations can relax constraints on reserves. Altogether, we test the following hypotheses:

Hypothesis 1. Bank stock return adjustments to news of Wisconsin's and Ohio's pension cut legislation prospects reflect banks' extent of business activity in these pension cut legislation states.

Hypothesis 2. Municipal bond spreads adjust negatively (positively) to news that conveys an increased (decreased) probability of Wisconsin's and Ohio's passage of pension cut legislation.

Hypothesis 3. Increased state solvency prospects resulting from pension cut legislation support bank credit supply through banks' municipal bond holdings.

III. Data and Methodology

A. Data

The sample examined in this study starts with the set of all publicly traded banks headquartered in the United States and operating during the first quarter of 2011. All accounting data used throughout the study are obtained from Federal Financial Institutions Examination Council (FFIEC) call reports databases found on the Chicago Federal Reserve's Web site (https://www.chicagofed.org/banking/financial-institution-reports/index). Data on banks are collected at the holding company level. That is, based on the highest holding company number of the bank, we collect and combine data for all banks with the same highest holding company number. Thus, we treat the bank holding companies as if they have only one bank, by combining their subsidiaries into one (consolidated) statement. Bank stock return data are collected from the Center for Research in Security Prices (CRSP) data tapes. Our initial sample includes a total of 329 banks.

Because a bank is more likely to be doing business in its headquarter state, to test our hypotheses, we ideally would like to compare banks that are headquartered in Wisconsin or Ohio with similar banks headquartered nearby in other Midwestern states. However, we find that only 2 U.S. banks are headquartered in Wisconsin and only 14 banks are headquartered in Ohio. Next, we rationalize that banks headquartered in a particular state can do business in many states. For example, banks headquartered in Illinois are likely to be issuing assets, collecting deposits, and buying municipal securities issued by the state of Wisconsin. Thus, we next attempt to collect the dollar value of assets of each U.S. bank located in Wisconsin or Ohio. However, the dollar value of assets (e.g., loans) issued by a bank in a particular state is also not available in the call reports. Finally, rather than using assets to measure the extent to which a bank conducts business in Wisconsin or Ohio, we examine the amount of deposits issued in Wisconsin or Ohio to the total deposits of the bank. Data on deposit holdings of banks by state are listed in the Federal Deposit Insurance Corporation (FDIC) Summary of Deposits Web site (https://www.fdic.gov/sod). Thus, our measure of the extent to which a bank conducts business in Wisconsin or Ohio is the amount of deposits held in the state to the total deposits of the bank. Any bank with a Wisconsin deposit ratio exceeding 1 standard deviation from the mean Wisconsin deposit ratio across all banks is classified as a "Wisconsin bank." Likewise, any bank with an Ohio deposit ratio exceeding 1 standard deviation from the mean Ohio deposit ratio across all banks is classified as an "Ohio bank." This classification identifies 5 Wisconsin banks and 15 Ohio banks for our treatment group.[11]

Having measured the extent of bank operations in Wisconsin and Ohio, we would next like to incorporate banks' holdings of municipal securities issued by Wisconsin and Ohio into the analysis. However, this level of detail regarding banks' municipal securities portfolios is not available in call reports or elsewhere, to our knowledge. Thus, we collect municipal security holdings at the bank level. In particular, we use the component of each bank's total municipal holdings designated as nonfiduciary and tradable.[12] It is likely that a bank that conducts a substantial degree of its business in a state also holds a substantial proportion of its municipal securities from that state or one of its subdivisions. We contacted about a dozen banks in Wisconsin and Ohio to inquire about the composition of their municipal securities portfolios. They all indicated that they hold state/local debt instruments of their own state and no other state. They also stated that this is not a home state bias. Rather, they hold local municipal securities to develop relations with their own localities with the hope of getting lending and deposit-taking business from the municipalities in return.

Finally, we construct a control group by drawing on the set of banks headquartered in the remaining five Midwestern states: Indiana, Illinois, Iowa, Michigan, and Minnesota. Specifically, we collect data on banks that are similar in asset size, total loans, and total deposit. Banks are then propensity score matched to link each Wisconsin and Ohio sample bank to its most similar Midwest bank as of year end 2009, as public pension reform discussions began in 2010.[13]

Table 1 lists the number of banks in the sample across each treatment group and matched control group, along with each group's proportion of deposits from Wisconsin and Ohio, tradable municipal bond holdings, and tier I capital ratios. The share of Wisconsin-based deposits is distinctly higher for Wisconsin banks relative to all other bank categories, including Midwest banks matched to Wisconsin banks, with an average of 41.1% of Wisconsin-bank deposits coming from Wisconsin. Wisconsin banks are also more heavily invested in municipal securities. However, the Wisconsin bank mean of $1.5 billion is highly skewed by the maximum municipal bond holding of $6.5 billion within the sample. On average, all groups of banks are financially healthy, as measured by the tier I capital ratio: averaging 11.9% for Wisconsin banks (ranging from 9.2% to 15.1%), 13.5%

[11]We examine treatment banks across 20 events and a subset of 8 core events, comparing them with other Midwest banks in a matched sample to improve identification (detailed subsequently). However, overall sample restrictions on treatment and control group banks may still limit the generality of our tests. In unreported tests, we examine a broader unmatched sample, covering *all* publicly traded banks, with relaxed deposit classification thresholds of 1% for Wisconsin and Ohio banks (thereby yielding 8 Wisconsin banks and 26 Ohio banks). We note this alternative sample yields very similar announcement effects. In subsequent bank lending tests (see Section IV.B), the matched sample of Midwest banks further encompasses all public and private banks, yielding 28 and 22 Wisconsin and Ohio treatment banks, respectively.

[12]We collect security data under item number 8499 in the FFIEC call reports, defined as the "fair value of available-for-sale securities issued by states and political subdivisions in the U.S." These municipal securities are nonfiduciary, not designated as being held to maturity, and available for trade.

[13]The geographical restriction to Midwest banks helps control for regional economic conditions. We then match bank characteristics within the region using one-to-one nearest neighbor propensity score matching with replacement.

TABLE 1

Descriptive Statistics on Key Variables for Wisconsin, Ohio, and Midwest Banks Matched to Wisconsin and Ohio Banks

Table 1 presents descriptive statistics for the sample banks at year end 2010. Wisconsin and Ohio banks are defined as publicly traded bank holding companies with 2010 deposit shares falling at least 1 standard deviation above the mean deposit shares in those states relative to the general population of publicly traded bank holding companies. Midwest banks are publicly traded bank holding companies headquartered in Indiana, Illinois, Iowa, Michigan, or Minnesota but not designated as Wisconsin or Ohio banks. The final Midwest bank matching samples are propensity score matched to Wisconsin and Ohio banks using total assets, total lending, and total deposits in 2009 as matching characteristics. Data on deposit holdings of banks by state are from the FDIC Summary of Deposits data. Data on other bank characteristics are from call reports provided by the Federal Reserve Bank of Chicago.

	Bank Holding Company Categories			
Variable	Wisconsin Banks	Midwest Banks Matched to Wisconsin Banks	Ohio Banks	Midwest Banks Matched to Ohio Banks
No. of obs. (no. of banks)	5 (5)	5 (3)	15 (15)	15 (9)
DEPOSITS_IN_WI/TOTAL_DEPOSITS (%)				
Mean	41.1	0.2	0	0.4
Minimum	8.5	0	0	0
Median	14.1	0.4	0	0
Maximum	100	0.4	0	5.2
DEPOSITS_IN_OH/TOTAL_DEPOSITS (%)				
Mean	3.5	1.2	75.3	0.1
Minimum	0	0	31.7	0
Median	0	0.1	82.9	0
Maximum	12.8	5.6	100	1.1
TRADABLE_MUNICIPAL_BOND_HOLDINGS (in USD millions)				
Mean	1,513	137	105	39
Minimum	0	52	0	0
Median	244	52	47	22
Maximum	6,303	300	456	228
TIER_I_CAPITAL_RATIO (%)				
Mean	11.9	13.5	11.5	10.2
Minimum	9.2	13.1	7.2	2.4
Median	11.5	13.1	11.6	11.9
Maximum	15.1	14.4	14.9	14.1

for Midwest banks matched to Wisconsin banks (ranging from 13.1% to 14.4%), 11.5% for Ohio banks (ranging from 7.2% to 14.9%), and 10.2% for Midwest banks matched to Ohio banks (ranging from 2.4% to 14.1%). Note that the vast majority of the capital ratio numbers are well above the minimum required tier I capital ratio needed for "adequately" capitalized banks (which was 4% during the sample period).

Table 2 lists descriptive statistics for the matching characteristics in 2009 and their growth in 2010 across the sample of banks. From Table 2, Wisconsin and Ohio banks are similar in size (as measured by total assets) to their matched banks in the region. The greatest difference is seen for the largest Wisconsin bank, which held total assets of $562 billion in 2009, relative to the largest matched Midwest bank, which held total assets of $151 billion in 2009. However, the value of total assets from the smallest Wisconsin bank ($2.2 billion) is similar to that of total assets from the smallest matched Midwest bank (1.8 billion), and the overall distribution of total assets across Wisconsin banks and their matched Midwest banks looks similar, with Wisconsin banks moderately larger, on average (mean = $128 billion), than their matched Midwest banks (mean = $92 billion). Thus, the treatment (Wisconsin and Ohio) and control (Midwest) banks appear reasonably matched by geographical region and size. We see similar distributional patterns

TABLE 2

Descriptive Statistics on the Levels and Growth of Matching Variables across Wisconsin, Ohio, and Midwest Banks Matched to Wisconsin and Ohio Banks

Table 2 presents descriptive statistics on matching variable levels and growth determined at year end 2009 (2010). Wisconsin and Ohio banks are defined as publicly traded bank holding companies with 2010 deposit shares falling at least 1 standard deviation above the mean deposit shares in those states relative to the general population of publicly traded bank holding companies. Midwest banks are publicly traded bank holding companies headquartered in Indiana, Illinois, Iowa, Michigan, or Minnesota but not designated as Wisconsin or Ohio banks. The final Midwest bank matching samples are propensity score matched to Wisconsin and Ohio banks using total assets, total lending, and total deposits in 2009 as matching characteristics. Data on deposit holdings of banks by state are from the FDIC Summary of Deposits data. Data on other bank characteristics are from call reports provided by the Federal Reserve Bank of Chicago.

	Bank Holding Company Categories			
Variable	Wisconsin Banks	Midwest Banks Matched to Wisconsin Banks	Ohio Banks	Midwest Banks Matched to Ohio Banks
TOTAL_ASSETS in 2009 (in USD billions)				
Mean	128	92	38.7	34.1
Minimum	2.2	1.8	0.89	0.90
Median	23.6	151	4.0	2.0
Maximum	562	151	225	151
TOTAL_DEPOSITS in 2009 (in USD billions)				
Mean	36.2	14.8	14.0	6.1
Minimum	0.93	0.69	0.35	0.88
Median	9.09	23.6	1.4	2.0
Maximum	152	23.6	78.3	151
TOTAL_LOANS in 2009 (in USD billions)				
Mean	88.3	32.5	27.1	13.5
Minimum	1.7	0.8	0.52	0.65
Median	16	52.6	2.1	1.7
Maximum	391	52.6	158	52.6
Growth in TOTAL_ASSETS during 2010 (%)				
Mean	−3.4	−3.0	1.2	−5.6
Minimum	−17.8	−5.3	−8.3	−17.1
Median	−3.3	−5.3	−0.6	−5.3
Maximum	9.3	5.5	34.2	5.5
Growth in TOTAL_DEPOSITS during 2010 (%)				
Mean	3.1	0.7	8.7	2.0
Minimum	−7.8	−1.3	−8.4	−11.4
Median	3.6	−1.3	1.7	1.8
Maximum	11.4	5.8	80.4	15.6
Growth in TOTAL_LOANS during 2010 (%)				
Mean	−5.0	−1.8	−0.3	−7.2
Minimum	−21.7	−10.8	−10.6	−35.9
Median	0.1	0	−1.6	−5.6
Maximum	3.4	2	28.6	2.0

across our other matching characteristics. For example, both total deposits and total loans are highest, on average, for Wisconsin banks, driven by the Wisconsin bank maximum for each variable. However, differences between treatment banks and their respective Midwest control banks across means, minimums, and medians appear economically moderate.

The 2010 growth distributions across each of the matching variables compare similarly between the treatment banks and their control group. We generally find growth in matching characteristics across each treatment and control group minimum, median, and maximum follows the same signs, overall reflecting similar distributional patterns and means. For example, the mean total asset growth by Wisconsin banks (−3.4%) is similar to their Midwest-matched banks (−3.0%). Moreover, we consistently observe the same signs for growth when comparing means for each characteristic across the sample banks and their control groups. One exception is the mean total asset growth by Ohio banks (1.2%), which exceeds that of their Midwest-matched banks (−5.6%). This difference is driven by the Ohio bank with maximum growth (34.2%). Yet the minimums, medians, and maximums across the comparison groups follow the same signs, reflecting a similar distributional pattern.

B. Methodology

The tests performed on the sample require us to identify dates on which important new information about pension reform became publicly available. News items pertaining to the pension cut bills are compiled by examining articles retrieved from a Google search using the keywords Wisconsin union (anti-union/pension cut) bill, Ohio union (anti-union/pension cut) bill, states with anti-union/pension cut bills, and anti-union/pension cut legislation. As of Mar. 2011, 15 states had introduced some type of anti-union/pension cut legislation. However, it is the legislation in Wisconsin and Ohio that received the earliest and most widespread attention in the press. Thus, we focus our analysis on announcements associated with passage of legislation in Wisconsin and Ohio. While our search produces several events associated with the passage of Wisconsin and Ohio pension cut (union rights) bills, we require that there be at least 5 items listed about an event to be included. This leaves us with 20 events relating to major announcements. Table 3 lists the event dates and a short description of each (events pertaining to the Ohio pension cut legislation are italicized, while events pertaining to Wisconsin pension cut legislation are not). For each event, the table also lists our anticipated announcement period stock price reaction for the sample banks. Details underlying the events are further discussed subsequently.

To measure the stock market effect of announcements associated with pension cut legislation, we estimate cumulative abnormal returns (CARs) for each sample bank and matched control bank at each event date using a market model framework. Since not all public banks trade regularly, we add additional lead and lag market excess return factors to control for nonsynchronous trading effects (Dimson (1979)). In particular, the following model is estimated over the last half of 2010:

$$(1) \qquad \begin{aligned} r_{i,t} - r_{f,t} \quad = \quad & \beta_{i,1}\left(r_{m,t} - r_{f,t}\right) + \beta_{i,2}\left(r_{m,t-1} - r_{f,t-1}\right) \\ & + \beta_{i,3}\left(r_{m,t+1} - r_{f,t+1}\right) + \varepsilon_{i,t}, \end{aligned}$$

where i and t index banks and events, respectively, and $r_{m,t} - r_{f,t}$ is the value-weighted excess market return relative to the 1-month Treasury bill. We require each regression to cover at least 25 observations. We, then, use the parameter estimates of equation (1) to estimate CARs by compounding residual estimates over the 3-day window centered around each announcement:[14]

$$(2) \qquad \mathrm{CAR}_{i,t} \quad = \quad \prod_{\tau=t-1}^{t+1}\left(1 + \hat{\varepsilon}_{i,\tau}\right) - 1,$$

where

$$\hat{\varepsilon}_{i,t} \quad = \quad r_{i,t} - \hat{\beta}_{i,1}\left(r_{m,t} - r_{f,t}\right) - \hat{\beta}_{i,2}\left(r_{m,t-1} - r_{f,t-1}\right) - \hat{\beta}_{i,3}\left(r_{m,t+1} - r_{f,t+1}\right),$$

[14]Because we have three events over the period Mar. 29–31, we also examine CARs over a 5-day window centered around Mar. 30, 2011.

TABLE 3

Major Announcements and Announcement Dates Associated with Wisconsin's and Ohio's Pension Cut Bill

Table 3 lists major dates related to the passage and implementation of Wisconsin's and Ohio's pension cut bill. News items pertaining to the changes in regulations are compiled by examining articles retrieved from a Google search using the keywords Wisconsin union (anti-union/pension cut) bill, Ohio union (anti-union/pension cut) bill, states with anti-union/pension cut bills, and anti-union/pension cut legislation. For each event, the table also lists our anticipated announcement period stock price reaction for the sample banks.

Event	Date	Description	Hypothesized Announcement Effect
1	Dec. 7, 2010	Governor-elect Walker raises possibility of changing state law to decertify unions and cut state employee benefits	+
2	Feb. 1, 2011	*Bill 5 introduced in the Ohio Senate*	+
3	Feb. 2, 2011	Governor Walker targets state worker benefits in speech	+
4	Feb. 17, 2011	Bill clears committee/Democratic lawmakers leave state	+/−
5	Feb. 24, 2011	Assembly ready to vote	+
6	Mar. 2, 2011	*Ohio Senate approves pension cut bill*	+
7	Mar. 9, 2011	Assembly passes bill	+
8	Mar. 11, 2011	Governor Walker signs bill into law	+
9	Mar. 16, 2011	Law challenge filed	−
10	Mar. 18, 2011	Wisconsin state judge puts law on hold	−
11	Mar. 25, 2011	Wisconsin pension reform law published despite court order	+
12	Mar. 29, 2011	*Ohio pension reform bill approved by House panel*	+
13	Mar. 30, 2011	*Ohio pension reform bill approved by House and Senate*	+
14	Mar. 31, 2011	Wisconsin judge rules pension cut bill not in effect	−
		Ohio pension reform bill signed into law	+
15	Apr. 15, 2011	*Ohio Attorney General certified summary language for a referendum seeking repeal of Senate Bill 5*	−
16	May 27, 2011	Wisconsin pension cut legislation struck down	−
17	June 15, 2011	Wisconsin Supreme Court rejects ruling that invalidates pension cut law	+
18	June 29, 2011	Wisconsin pension cut law goes into effect	+/−
		Ohio opponents submit signatures needed to get repeal of bill on November ballot	
19	July 21, 2011	*Ohio puts repeal of pension cut bill on November ballot*	−
20	Nov. 8, 2011	*Ohio voters repeal new law that would cut pensions*	−

which makes use of the parameter estimates produced by equation (1), denoted $\hat{\beta}$.[15]

To test Hypothesis 1, we run cross-sectional regressions to gauge each event's influence in the underlying data. That is, for each event t and each bank i, we run various forms of the following regression:

$$\text{(3)}\quad \begin{aligned} \text{CAR}_i = {} & \alpha + \delta_1 \ln(1 + \text{TRADABLE\_MUNICIPAL\_BOND\_HOLDINGS}_i) \\ & + \delta_2\, \text{TIER\_I\_CAPITAL\_RATIO}_i \\ & + \delta_3\, \text{DEPOSITS\_HELD\_IN\_WI\_TO\_TOTAL\_DEPOSITS}_i \\ & + \delta_4\, \text{DEPOSITS\_HELD\_IN\_OH\_TO\_TOTAL\_DEPOSITS}_i + \varepsilon_i. \end{aligned}$$

Since state-specific banking laws may create correlations within announcement returns, standard error estimates from equation (3) are clustered by bank headquarter state. The independent variables are lagged by 1 year.

Equation (3) tests Hypothesis 1 across banks *within* each event. However, segmenting the data this way reduces degrees of freedom and, importantly, ignores potentially correlated errors due to event-level clustering. We, therefore, test our hypotheses more formally by pooling the data and additionally

[15]In unreported results, we find our results are robust to modeling event day abnormal stock returns using a Fama–French 3-factor model (Fama and French (1993)), as well as to using raw returns. The results are available from the authors.

clustering by event. Since different events associate with varying signed predictions, we sign our dependent variable in the pooled regressions, multiplying it by a value of −1 for each fully negative hypothesized event date within the sample (shown in Table 3).[16] The pooled model is

$$\text{(4)} \quad \mathrm{CAR}_{i,t} = \alpha + \delta_1 \ln(1 + \mathrm{TRADABLE\_MUNICIPAL\_BOND\_HOLDINGS}_{i,t}) \\ + \delta_2\, \mathrm{TIER\_I\_CAPITAL\_RATIO}_{i,t} \\ + \delta_3\, \mathrm{DEPOSITS\_HELD\_IN\_WI\_TO\_TOTAL\_DEPOSITS}_{i,t} \\ + \delta_4\, \mathrm{DEPOSITS\_HELD\_IN\_OH\_TO\_TOTAL\_DEPOSITS}_{i,t} + \varepsilon_i,$$

where the i and t index banks and events and the independent variables are again lagged by 1 year. Standard error estimates from equation (4) are clustered by bank headquarter state and event date.

IV. Results

A. Announcement Period Abnormal Returns

1. Regression Results by Event

To identify major events empirically within our set of 20 events, we examine regression results one event at a time using equation (3). Since the matched sample allows only 40 observations per event, for these tests, we include the full sample of publicly traded banks to preserve degrees of freedom. Eight events produce a significant relationship between abnormal announcement period returns and the individual characteristics of the sample banks with the expected sign (consistent with Hypothesis 1), which we discuss next. Table 4 reports regressions for these 8 events, while Appendix A discusses and reports results for the other 12 events that do not produce significant results consistent with Table 3.

On Feb. 1, 2011, Bill 5 was introduced in the Ohio Senate; the next day, Feb. 2, 2011, Wisconsin Governor Walker targeted state worker benefits in a speech. Both of these events signal an increased probability of the passage of pension cut legislation. The regression for Feb. 1, 2011 in column 1 of Table 4 reports that the greater the extent to which a bank does business in either Wisconsin or Ohio the more positive is the bank's stock price reaction to the news (coefficients on Wisconsin and Ohio deposits to total deposits are 0.0277 and 0.0193, respectively, both significant at 1%). Consistent with Hypothesis 1, the increased probability of pension cut legislation being passed in Wisconsin and Ohio appears to be positive news for banks doing business in these states.

On Feb. 17, 2011, the Wisconsin bill cleared the legislature's budget writing committee. After this vote, Democratic lawmakers left the state in an attempt to stop a vote on the bill. Regression 2 in Table 4 reports that banks doing more business in Wisconsin react more positively to the news (the coefficient on Wisconsin deposits to total deposits is 0.0289, significant at 1%). Thus, even though the walkout by lawmakers was intended to stop a subsequent vote on the bill, the clearing of the bill by the budget writing committee appears to be positive

[16]This results in signed adjustments to CARs over the following event dates in 2011: Mar. 16, 18, and 31; Apr. 15, May 27; July 21; and Nov. 9.

TABLE 4

Regression Results of Stock Price Reactions to Eight Major News Announcements Associated with the Passage of Pension Cut Legislation

Table 4 presents regression results for announcements associated with the passage of pension cut legislation in Wisconsin and Ohio. Cumulative abnormal returns (CARs) for each event are estimated from a market model. Since not all public banks trade regularly, additional lead and lag market excess return factors are added to control for nonsynchronous trading effects (Dimson (1979)). For each event t, the following regressions are run across all publicly traded bank holding companies (with each bank denoted by subscript i):

$$\begin{aligned}\text{CAR}_i \quad = \quad & \alpha+\delta_1 \ln(\text{ASSETS}_i)+\delta_2 \ln(1+\text{TRADABLE\_MUNICIPAL\_BOND\_HOLDINGS}_i)+\delta_3 \text{TIER\_I\_CAPITAL\_RATIO}_i+\delta_4 \text{DEPOSITS\_HELD\_IN\_WI\_TO\_TOTAL\_DEPOSITS}_i \\ & +\delta_5 \text{DEPOSITS\_HELD\_IN\_OH\_TO\_TOTAL\_DEPOSITS}_i+\varepsilon_i.\end{aligned}$$

Independent variable levels are lagged by 1 year (determined at year end 2010). Data on deposit holdings of banks by state are from the FDIC Summary of Deposits data. Data on other bank characteristics are from call reports provided by the Federal Reserve Bank of Chicago. Standard errors are clustered by bank headquarter state. t-statistics are reported in parentheses. *, **, and *** indicate significance at the 10%, 5%, and 1% levels, respectively. †Dependent Variable: 5-day CAR used.

	Dependent Variable: 3-Day CAR (market model)								
	Sample: Publicly Traded Bank Holding Companies								
	Feb. 1, 2011	Feb. 17, 2011	Mar. 2, 2011	Mar. 9, 2011	Mar. 25, 2011	Mar. 30, 2011	Mar. 30, 2011†	May 27, 2011	June 29, 2011
Variable	1	2	3	4	5	6	7	8	9
ln(ASSETS)	−0.0029	−0.0042**	−0.0018**	0.0059***	−0.0003	0.0007	0.0034*	0.0002	−0.0032*
	(−1.670)	(−2.360)	(−2.216)	(5.320)	(−0.262)	(0.419)	(1.971)	(0.117)	(−1.732)
TIER_I_CAPITAL_RATIO	0.4238***	0.1694**	0.1628**	0.2391***	−0.0186	0.0884	0.1566	0.0025	−0.2227
	(2.895)	(2.026)	(2.659)	(2.765)	(−0.171)	(0.678)	(1.325)	(0.028)	(−1.332)
ln(1 + TRADABLE_MUNICIPAL_ BOND_HOLDINGS)	0.0010	0.0005	0.0011*	−0.0003	−0.0003	−0.0005	0.0008	−0.0004	−0.0024***
	(1.152)	(0.933)	(2.007)	(−0.577)	(−0.499)	(−0.523)	(1.285)	(−0.754)	(−2.710)
DEPOSITS_HELD_IN_WI_ TO_TOTAL_DEPOSITS	0.0277***	0.0289***	0.0308**	0.0079	−0.0000	−0.0186**	0.0056	−0.0131**	0.1056***
	(3.222)	(7.858)	(2.216)	(1.231)	(−0.006)	(−2.531)	(0.627)	(−2.205)	(17.097)
DEPOSITS_HELD_IN_OH_ TO_TOTAL_DEPOSITS	0.0193***	−0.0085*	0.0051**	0.0238***	0.0164***	0.0000	0.0128***	0.0012	−0.0116*
	(3.204)	(−1.869)	(2.153)	(7.178)	(4.605)	(0.007)	(2.794)	(0.300)	(−1.876)
Intercept	−0.0294	0.0335	−0.0085	−0.1140***	0.0018	−0.0150	−0.0804**	−0.0013	0.0908*
	(−0.737)	(1.122)	(−0.633)	(−4.725)	(0.065)	(−0.444)	(−2.200)	(−0.047)	(1.938)
R^2	0.0551	0.0480	0.0434	0.0922	0.00657	0.00550	0.0426	0.00154	0.0812
No. of obs.	319	319	319	318	318	318	318	311	308

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news for banks in the state. Interestingly, Ohio banks show a weak negative overall reaction to the news (the Ohio deposit ratio coefficient is 0.0085, significant at 10%). The negative overall reaction by Ohio banks is interpreted as evidence that the walkout by Democratic lawmakers in Wisconsin signified the potential for greater political tension in the legislative process for Ohio banks, despite the bill having cleared the legislature's budget writing committee in Wisconsin.

On Mar. 2, 2011, the Ohio Senate approved the pension cut bill. Regression 3 in Table 4 reports that the greater the extent to which a bank does business in either Wisconsin or Ohio the greater the bank's stock price reaction to the news (the Wisconsin deposit ratio coefficient is 0.0308 and the Ohio deposit ratio coefficient is 0.0051 (both significant at 5%)). Thus, the increased probability of passage of pension cut legislation in Ohio appears to be positive news for banks in both Wisconsin and Ohio.

On Mar. 9, 2011, the Wisconsin assembly passed the pension cut bill despite the absence of the 14 Democratic senators. This event appears to have a positive but statistically weak impact on banks doing business in Wisconsin. Further, the greater the extent to which a bank does business in Ohio the greater the bank's stock price reacts to the news (the Ohio deposit ratio coefficient in column 4 of Table 4 is 0.0238, significant at 1%). Thus, this event (signaling positive news about the passage of pension cut legislation in Wisconsin) appears to be positive news for banks in Ohio (also going through the process of pension cut legislation).

On Mar. 25, 2011, the Legislative Reference Bureau published the Wisconsin pension cut law (the last step before the law goes into effect), despite a court order blocking its publication, while challenges to the law were being considered. While bank stock returns for this event are not related to the extent to which banks do business in Wisconsin, banks doing more business in Ohio react more positively to the news (the Ohio deposit ratio coefficient is 0.0164 in column 5 of Table 4, significant at 1%). Thus, this event (which signals positive news about the passage of pension cut legislation in Wisconsin) also appears to be positive news for banks in Ohio.

On Mar. 29, 2011, Ohio's pension cut bill was approved by a House panel; on Mar. 30, the bill was approved by the full House and Senate in Ohio; and on Mar. 31, the bill was signed into law. The three consecutive events signal an increased probability of the passage of pension cut legislation. However, on Mar. 31, the Wisconsin law was put on indefinite hold by a Dane County Circuit Court judge until the case could be heard by the Wisconsin Supreme Court. Thus, at this point, the pension cut bill would not go into effect. Further, the Supreme Court had not indicated whether it would even take the case. This event signals a decreased probability of the passage of Wisconsin's pension cut legislation. Because we have three similar Ohio-based events over the period Mar. 29–31, we examine both a 3-day (column 6 of Table 4) and a 5-day (column 7) window centered around Mar. 30, 2011. Using the 3-day event window, we find that only banks doing business in Wisconsin are impacted by the events. Specifically, over the 3-day window, stock prices of banks doing more business in Wisconsin react more negatively to the negative news of the law being put on indefinite hold (the Wisconsin deposit ratio coefficient is −0.0186 in column 6, significant at 5%). However, using the 5-day event window, only banks doing business in Ohio are impacted by the events. Specifically, stock prices of banks doing more business in Ohio react more positively to the positive news of the pension cut bill's approval by a House panel than by the full House and Senate in Ohio (the Ohio deposit ratio coefficient is 0.0128 in column 7, significant at 1%).

On May 27, 2011, a Dane County judge issued a permanent injunction that effectively threw out Wisconsin's pension cut bill, concluding that Republicans passed the bill by violating the state's strong "open meeting" law, and that the law was thus invalid. Column 8 of Table 4 reports that the greater the extent to which a bank operates in Wisconsin the more the bank's stock price reacts to the news (the Wisconsin deposit ratio coefficient −0.0131, significant at 5%). Ohio banks do not react to this event, consistent with the state-specific nature of this open meeting injunction eliciting a state-specific bank response.

Finally, on June 29, 2011, pension cut legislation became reality as Wisconsin's pension cut law went into effect. Also, on this day, Ohio opponents submitted enough signatures needed to get a repeal of Bill 5 on the November ballot. Column 9 of Table 4 reports that the greater the extent to which a bank does business in Wisconsin the greater the bank's stock price reacts to the news (the Wisconsin deposit ratio coefficient is 0.1056, significant at 1%). Thus, the implementation of pension cut legislation in Wisconsin appears to be positive news for banks in these states. Further, the greater the extent to which a bank does business in Ohio the more negative is the bank's stock price reaction to the news (the Ohio deposit ratio coefficient is −0.0116, significant at 10%). Thus, news that Ohio's pension cut legislation would be up for a vote on the November ballot appears to be bad news for banks operating in Ohio. Despite the enactment of pension cut legislation and the resulting easing of the strain on Wisconsin's state budget, the increased possibility of repeal of the Ohio pension cut bill dominates the stock price reactions to these events for Ohio banks.

2. Pooled Regression Results

Table 5 reports results of the pooled cross-sectional model of equation (4) for the matched sample of treatment (Wisconsin and Ohio) and control (Midwest) banks. Regressions in columns 1 and 2 use Wisconsin and Ohio bank indicators to isolate the differential impact of pension cut legislation on Wisconsin and Ohio banks, with the column 1 regression pooling across the 20 event dates in Table 3 and the column 2 regression pooling across the 8 major event dates in Table 4. Regressions in columns 3 and 4 show these effects again using as test variables the continuous Wisconsin and Ohio deposit ratios across the sample of banks.

Pooling across all 20 events, column 1 of Table 5 documents that pension cut legislation generates stronger stock price reactions in banks doing more business in Wisconsin and Ohio. These results are consistent with the signed predictions of Table 3 and with Hypothesis 1. The Wisconsin indicator coefficient indicates that, on average, Wisconsin bank abnormal stock returns respond by 1.09% (significant at 5%) across the set of 20 events in Table 3, while the Ohio indicator coefficient shows a similar average abnormal market response of 1.26% (significant at 1%) across the 20 events in Table 3 for Ohio banks. Column 2 shows these Wisconsin and Ohio bank average effects increase substantially across the set of 8 events from Table 4, with Wisconsin and Ohio bank abnormal returns now responding on average by 2.78% and 1.85%, respectively (both significant at 1%).

TABLE 5

Pooled Regression Results of Stock Price Reactions to Major News Announcements Associated with the Passage of Pension Cut Legislation

Table 5 presents regression results for announcements associated with the passage of pension cut legislation in Wisconsin and Ohio. Hypothesis 1 is tested by pooling cross-sectional event data. We estimate 3-day cumulative abnormal returns (CARs) for each event from a market model. For Mar. 30, 2011, a 5-day market CAR is used to span 3 consecutive event days. Since not all public banks trade regularly, we add additional lead and lag market excess return factors to control for nonsynchronous trading effects (Dimson (1979)). Across events, we run the following regression using bank holding companies with abnormal deposit shares in Wisconsin and Ohio in 2010 and their matched set of Midwest banks:

$$\begin{aligned}\text{CAR}_{i,t} \quad = \quad & \alpha + \delta_1 \ln(1 + \text{TRADABLE\_MUNICIPAL\_BOND\_HOLDINGS}_{i,t}) + \delta_2 \,\text{TIER\_I\_CAPITAL\_RATIO}_{i,t} \\ & + \delta_3 \,\text{DEPOSITS\_HELD\_IN\_WI\_TO\_TOTAL\_DEPOSITS}_{i,t} \\ & + \delta_4 \,\text{DEPOSITS\_HELD\_IN\_OH\_TO\_TOTAL\_DEPOSITS}_{i,t} + \varepsilon_{i,t},\end{aligned}$$

where the i and t index banks and events and the independent variables are lagged by 1 year. Since different events associate with varying signed predictions, we sign our dependent variable in the pooled regressions, multiplying it by a value of −1 for each fully negative hypothesized event date within the sample, as shown in Table 3. Wisconsin and Ohio banks are defined as publicly traded bank holding companies with 2010 deposit shares falling at least 1 standard deviation above the mean deposit shares in those states relative to the general population of publicly traded bank holding companies. Midwest banks are publicly traded bank holding companies headquartered in Indiana, Illinois, Iowa, Michigan, or Minnesota but not designated as Wisconsin or Ohio banks. The final Midwest bank matching samples are propensity score matched to Wisconsin and Ohio banks using total assets, total lending, and total deposits in 2009 as matching characteristics. Data on deposit holdings of banks by state are from the FDIC Summary of Deposits data. Data on other bank characteristics are from call reports provided by the Federal Reserve Bank of Chicago. Standard errors are clustered by event dates and bank headquarter state. t-statistics are reported in parentheses. *, **, and *** indicate significance at the 10%, 5%, and 1% levels, respectively.

	Dependent Variable: 3-Day CAR (market model)			
	Sample: Matched Banks across Event Dates			
	All 20 Events	8 Major Events	All 20 Events	8 Major Events
Variable	1	2	3	4
ln(1 + TRADABLE_MUNICIPAL_BOND_HOLDINGS)	−0.0016 (−1.629)	−0.0025** (−2.360)	−0.0013 (−1.322)	−0.0017 (−1.547)
TIER_I_CAPITAL_RATIO	0.2770* (1.769)	0.3943** (1.963)	0.2587* (1.711)	0.3354* (1.688)
WI_INDICATOR	0.0109** (1.963)	0.0278*** (3.594)		
OH_INDICATOR	0.0126*** (3.014)	0.0185*** (2.785)		
WI_DEPOSIT_RATIO			0.0102** (2.279)	0.0319*** (5.273)
OH_DEPOSIT_RATIO			0.0127*** (2.853)	0.0191*** (2.719)
Intercept	−0.0295** (−2.111)	−0.0408** (−2.045)	−0.0284** (−2.116)	−0.0382* (−1.954)
R^2	0.0349	0.0726	0.0295	0.0578
No. of obs.	790	316	790	316

Columns 3 and 4 of Table 5 show the Wisconsin and Ohio bank announcement effects also scale in proportion to the extent to which banks operate in those states. For the full sample of 20 events, the deposit ratio coefficients in column 3 indicate that a 1% increase in a bank's Wisconsin-based deposit share generates a 1.02 basis point increase in its market response to the sequence of events associated with passage of pension cut laws (significant at 5%), while a 1% increase in a bank's Ohio-based deposit share increases its market response by 1.27 basis points (significant at 1%). The effects are again more significant over the set of 8 major events from Table 4. Column 4 of Table 5 shows that a 1% increase in bank Wisconsin-based deposit shares strengthens market responses to the pension cut news by 3.19 basis points (significant at 1%), while a 1% increase in bank Ohio-based deposit shares strengthens market responses by 1.91 basis points (significant at 1%).

Overall, we find that the probability of passage of pension cut legislation in Wisconsin and Ohio is more value relevant to banks doing a greater extent of business in these states.

Table 6 further examines the connection between bank valuations and the probability of pension cut legislation by considering banks' potential direct economic connection to state solvency through their municipal bond holdings. Pooling over the 20 event dates listed in Table 3, column 1 of Table 6 shows market responses to pension cut news are generally not significantly related to banks' holdings of tradable municipal bonds or banks' tier 1 capital ratios for the matched sample of banks. However, in column 2 within the matched sample, banks'

TABLE 6

Municipal Bond Holdings, Capital Adequacy, and Pooled Stock Price Reactions to News Announcements Associated with the Passage of Pension Cut Legislation

Table 6 presents regression results for announcements associated with the passage of pension cut legislation in Wisconsin and Ohio by pooling cross-sectional event data. We estimate 3-day cumulative abnormal returns (CARs) for each event from a market model. For Mar. 30, 2011, we use a 5-day market CAR to span 3 consecutive event days. Since not all public banks trade regularly, additional lead and lag market excess return factors are added to control for non-synchronous trading effects (Dimson (1979)). Since different events associate with varying signed predictions, we sign the dependent variable in the pooled regressions, multiplying it by a value of −1 for each fully negative hypothesized event date within the sample, as shown in Table 3. Across events, we run several regression models over bank holding companies with abnormal deposit shares in Wisconsin and Ohio in 2010 and their matched set of Midwest banks, with all independent variables lagged by 1 year. Wisconsin and Ohio banks are defined as publicly traded bank holding companies with 2010 deposit shares falling at least 1 standard deviation above the mean deposit shares in those states relative to the general population of publicly traded bank holding companies. Midwest banks are publicly traded bank holding companies headquartered in Indiana, Illinois, Iowa, Michigan, or Minnesota but not designated as Wisconsin or Ohio banks. The final Midwest bank matching samples are propensity score matched to Wisconsin and Ohio banks using total assets, total lending, and total deposits in 2009 as matching characteristics. Data on deposit holdings of banks by state are from the FDIC Summary of Deposits data. Data on other bank characteristics are from call reports provided by the Federal Reserve Bank of Chicago. Standard errors are clustered by event dates and bank headquarter state. t-statistics are reported in parentheses. *, **, and *** indicate significance at the 10%, 5%, and 1% levels, respectively.

	Dependent Variable: 3-Day CAR (market model)							
	Sample: Matched Banks across Event Dates							
	All 20 Events				8 Major Events			
Variable	1	2	3	4	5	6	7	8
ln(1 + TRADABLE_MUNICIPAL_BOND_HOLDINGS)	−0.0013	−0.0026**	−0.0137***	−0.0138***	−0.0019	−0.0038**	−0.0132***	−0.0135***
	(−1.428)	(−1.978)	(−4.074)	(−4.225)	(−1.550)	(−2.042)	(−3.770)	(−3.870)
TIER_1_CAPITAL_RATIO	0.2654	0.3438	0.0373*	−0.0040	0.3739	0.4929	0.3311	0.1082
	(1.455)	(1.586)	(1.700)	(−0.043)	(1.372)	(1.533)	(1.633)	(0.443)
ln(1 + TRADABLE_MUNICIPAL_BOND_HOLDINGS) × WI_INDICATOR		0.0013*	0.0131***	0.0126***		0.0024**	0.0147***	0.0123***
		(1.759)	(3.094)	(2.943)		(2.444)	(3.256)	(2.628)
ln(1 + TRADABLE_MUNICIPAL_BOND_HOLDINGS) × OH_INDICATOR		0.0015***	0.0131***	0.0128***		0.0019**	0.0144***	0.0124***
		(2.657)	(3.566)	(3.643)		(2.033)	(3.424)	(2.779)
ln(1 + TRADABLE_MUNICIPAL_BOND_HOLDINGS) × WI_INDICATOR ×TIER_1_CAPITAL_RATIO			−0.0940***	−0.0953***			−0.0969***	−0.1016***
			(−3.074)	(−3.184)			(−2.705)	(−2.836)
ln(1 + TRADABLE_MUNICIPAL_BOND_HOLDINGS) × OH_INDICATOR ×TIER_1_CAPITAL_RATIO			−0.0912***	−0.0907***			−0.1001***	−0.0979***
			(−3.361)	(−3.328)			(−2.941)	(−2.887)
ln(1 + TRADABLE_MUNICIPAL_BOND_HOLDINGS) × TIER_1_CAPITAL_RATIO			0.0930***	0.0962***			0.0749**	0.0923***
			(4.253)	(4.155)			(2.500)	(2.914)
WI_INDICATOR				0.0080*				0.0347
				(1.747)				(1.591)
OH_INDICATOR				0.0023				0.0190
				(0.340)				(1.140)
Intercept	−0.0245	−0.0294*	−0.0011	0.0004	−0.0327	−0.0393	−0.0226	−0.0150
	(−1.550)	(−1.655)	(−0.29)	(0.09)	(−1.324)	(−1.416)	(−1.287)	(−0.830)
R^2	0.0189	0.0405	0.0823	0.0828	0.0315	0.0685	0.102	0.110
No. of obs.	790	790	790	790	316	316	316	316

tradable municipal bond holdings are associated with significantly stronger announcement returns for banks with greater operations in Wisconsin and Ohio (the coefficient on Wisconsin bank municipal bonds is 0.0013, significant at 10%, and the coefficient on Ohio bank municipal bonds is 0.0015, significant at 1%). Column 3 further describes municipal bond holding effects of column 2 by accounting for banks' financial health, as proxied by tier 1 capital ratios. There is now a reversal of the Wisconsin and Ohio bank municipal bond holding effects in proportion to how well banks are capitalized (interactions of tier 1 capital with Wisconsin and Ohio bank municipal bond holdings are −0.0940 and −0.0912, respectively, both significant at 1%). In other words, pension cut legislation passage is more value relevant for Wisconsin and Ohio banks that hold more municipal securities, particularly if they are more capital constrained. Column 4 includes Wisconsin and Ohio indicator variable direct effects to the specification in column 3 and shows very similar results. The Wisconsin indicator coefficient is 0.0080 (significant at 10%), while the Ohio indicator coefficient is 0.0023 (not significant). Both coefficients fall relative to their observed values in column 1 of Table 5. Overall, column 4 of Table 6 provides supporting evidence that municipal bond holdings are a value-relevant driver of bank value adjustments to the likelihood of pension cut legislation passage, beyond potential valuation effects that may come from broad economic conditions associated with pension reform.

Columns 5–8 in Table 6 report the regression model results of columns 1–4 when pooling across the 8 major event dates in Table 4. Overall, the regression coefficients in columns 5–8 are similar in significance and magnitudes to their counterparts in columns 1–4. We note that the Wisconsin and Ohio direct effects in column 8 are, again, similar in magnitude to their initial estimate values in column 2 of Table 5. However, they are no longer statistically significant.

3. Municipal Bond Spreads

If pension cut legislation affects bank values directly through changing default risk exposures in banks' municipal bond portfolios (as a result of shifts in investor expectations of states' budget solvency), then we would also expect spreads on debt issued by Wisconsin or Ohio to decrease more (less), on average, for event days, signaling an increased (decreased) probability of passage of pension cut legislation relative to spread changes on nonevent days, and relative to municipal spread changes on debt issued by other states, on average. To test this, we collect daily historical municipal bond yields from Thomson Reuters Municipal Market data and form spreads on states' AAA-rated 10-year general obligation municipal bond rates over 10-year Treasury yields from 2010 through 2011.[17] We then examine abnormal municipal bond spread changes for debt issued by Wisconsin and Ohio on event days relative to spread changes on nonevent days and spread changes on debt issued by other control states.[18]

Specifically, we model daily variation in spreads, $\Delta\text{SPREAD}_{i,t}$, using 1-day lead minus 1-day lagged spreads on AAA 10-year general obligation municipal

[17] Available state-level data are limited to AAA-rated municipal bond spreads.

[18] Preferential tax treatment of municipal bonds causes spreads to Treasury yields to sometimes be negative. However, examining daily changes in spreads differences out the tax effect.

bond rates benchmarked against 10-year Treasury rates, as:[19]

$$(5) \quad \Delta \text{SPREAD}_{i,t} = \alpha + \beta_1 \text{EVENT\_INDICATOR}_t + \beta_2 \text{WI\_INDICATOR}_i + \beta_3 \text{OH\_INDICATOR}_i + \beta_4 \text{EVENT\_INDICATOR}_t \times \text{WI\_INDICATOR}_i + \beta_3 \text{EVENT\_INDICATOR}_t \times \text{OH\_INDICATOR}_i + \varepsilon_{i,t}.$$

EVENT_INDICATOR_t equals 1 when spread changes are centered around event days, and 0 otherwise. Each indicator variable equals 1 when spread changes come from debt issued by their respective states, and 0 otherwise. In the panel underlying equation (5), we again pool events by multiplying a value of -1 to the dependent variable when observations correspond to the negative events in Table 3. In equation (5), α reports average spread changes in control states on nonevent days, β_1 reports average spread changes in control states on event days, β_2 and β_3 report average spread changes in the treatment states (Wisconsin and Ohio, respectively) on nonevent days, and β_4 and β_5 report the average incremental abnormal spread changes in the treatment states on event days.

Table 7 presents regression results. Column 1 shows that daily spreads across control states on nonevent days increase by 0.21% (significant at 1%), on average, during the 2010–2011 sample period, and the average increase is greater, by 0.38% (significant at 1%), across the 20 event days in Table 3. Over nonevent days, average daily municipal spread changes are lower by 0.05% (significant at 1%) for debt issued by Wisconsin and by 0.03% (significant at 5%) for debt issued by Ohio. However, consistent with Hypothesis 2, on event days, municipal spreads decline by 0.74% (significant at 1%) on debt issued by Wisconsin and decline by 0.12% (significant at 5%) on debt issued by Ohio. Column 2 shows a broadly similar pattern with respect to the 8 major event days. Column 2 reports an event indicator estimate of -0.79%, down from 0.38% in column 1 (both significant at 1%). This drop indicates the relatively important influence the 8 major events of Table 4 had on lowering municipal spreads across all states, on average. Adding the average event day effect to the incremental event day interaction effects for Wisconsin and Ohio, the overall impact of strengthened state solvency through shifts in pension cut probabilities decreases Wisconsin municipal spreads by 1.03% in column 2 (-0.79% -0.24%, significant at 1%), which exceeds the overall Wisconsin effect in column 1 by 0.67%. Similarly, the total effect of event day spread tightening for Ohio municipal bonds in column 2 ($0.96\% = -0.79\% - 0.17\%$) exceeds that of column 1 by 1.22%. Columns 3 and 4 of Table 7 show the results of columns 1 and 2, respectively, when including state fixed effects to control for fixed economic policies, regulations, and other conditions within states. The overall effects in columns 3 and 4 are virtually identical to those in columns 1 and 2, respectively. Importantly, the state-level spread adjustments, although economically small, are still surprising given that AAA municipal bonds are typically insured securities.

[19]Since spread change volatility varies substantially across states, we winsorize spread changes within each state at 5% to prevent oversampling states in which spread changes are generally more volatile.

TABLE 7

Reaction of Municipal Bond Spreads to Major News Announcements Associated with the Passage of Pension Cut Legislation

Table 7 presents regression results for daily state-level changes in municipal bond spreads on announcement days associated with the passage of pension cut legislation in Wisconsin and Ohio, relative to spread changes on nonannouncement days and spread changes in other states, from 2010 to 2011. For each day t and each state i, we run various forms of the following regression:

$$\begin{aligned}\Delta\text{SPREAD}_{i,t} \quad = \quad & \alpha + \beta_1 \text{EVENT\_INDICATOR}_t + \beta_2 \text{STATE\_INDICATOR}_i \\ & + \beta_3 \text{EVENT\_INDICATOR}_t \times \text{STATE\_INDICATOR}_i + \varepsilon_{i,t},\end{aligned}$$

where $\Delta\text{SPREAD}_{i,t}$ is 1-day lead minus 1-day lagged spreads using AAA 10-year general obligation municipal bond yields benchmarked against 10-year Treasury yields; EVENT_INDICATOR_t equals 1 when spread changes are centered around event days, and 0 otherwise; and STATE_INDICATOR_i equals 1 when spread changes come from debt issued by Wisconsin or Ohio, and 0 otherwise. Since different events associate with varying signed predictions, we pool events by multiplying a value of −1 to the dependent variable when observation dates correspond to the negative events in Table 3. The set of events underlying columns 1 and 3 comprises all events listed in Table 3. The set of events underlying columns 2 and 4 comprises the subset of 8 major events for which the individual event regressions reported in Table 4 identify statistically significant effects. Standard errors clustered by state. t-statistics are reported in parentheses. *, **, and *** indicate significance at the 10%, 5%, and 1% levels, respectively.

	Dependent Variable: Daily Spread Changes ($\text{SPREAD}_{i,t+1} - \text{SPREAD}_{i,t-1}$)			
	Sample: Daily Municipal Spreads			
	All 20 Events	8 Major Events	All 20 Events	8 Major Events
Variable	1	2	3	4
EVENT_INDICATOR	0.0038***	−0.0079***	0.0038***	−0.0079***
	(8.605)	(−17.649)	(8.689)	(−17.819)
WI_INDICATOR	−0.0005***	−0.0008***	−0.0004***	−0.0006***
	(−4.701)	(−7.249)	(−20.553)	(−87.176)
OH_INDICATOR	−0.0003**	−0.0003***	−0.0001***	−0.0001***
	(−2.380)	(−2.684)	(−5.224)	(−15.880)
EVENT_INDICATOR × WI_INDICATOR	−0.0074***	−0.0024***	−0.0074***	−0.0024***
	(−16.638)	(−5.377)	(−16.801)	(−5.429)
EVENT_INDICATOR × OH_INDICATOR	−0.0012**	−0.0017***	−0.0012***	−0.0017***
	(−2.673)	(−3.727)	(−2.699)	(−3.763)
Intercept	0.0021***	0.0024***	—	—
	(18.267)	(21.675)	—	—
State fixed effects	No	No	Yes	Yes
R^2	0.0001	0.0002	0.0001	0.0002
No. of obs.	25,992	25,992	25,992	25,992

4. Results for Industrial Firms

A general question is: Why look at banks as opposed to other firms doing business in Wisconsin and Ohio? Although banks' exposures to municipal bonds may narrowly drive valuation effects as expectations about pension cut legislation change, one alternative is bank valuation effects occur because of broader changes in expectations about future taxation. This, like asset shocks from banks' exposures to municipal bonds, would effect changes in expectations about lending and deposit-taking opportunities in the state, but, more generally, it would affect expected economic conditions for other firms doing business in Wisconsin and Ohio as well. To examine the effect of pension cut legislation on banks versus other firms in pension cut states, we collect stock returns of all industrial firms over the period of analysis. We also collect information on the state in which each industrial firm is headquartered and each firm's size (measured as the natural log of book value of total assets) and leverage ratio (total debt to total assets) as control variables. Using a regression framework similar to equation (4), but applied to industrial firms, we test the impact of industrial firms' state headquarter indicators on abnormal returns.[20]

Regression results are reported in Table 8. Column 1 shows the impact of pension cut legislation on industrial firms headquartered in Wisconsin and Ohio relative to other publicly traded industrial firms, pooling, as before, across the 20 events in Table 3; column 2 shows the relative impact of pension cut legislation on Wisconsin and Ohio firms, pooling on the 8 major events in Table 4; columns 3 and 4 repeat the tests in columns 1 and 2, respectively, by including state fixed effects to further control for fixed state characteristics. In all regressions, we see no evidence that the increased probability of pension cut legislation being passed in Wisconsin and Ohio impacts stock prices of industrial firms headquartered in these two states.

TABLE 8

Pooled Regression Results of Stock Price Reactions for Industrial Firms to Major News Announcements Associated with the Passage of Pension Cut Legislation

Table 8 examines the effect of pension cut legislation on industrial firms in pension cut states by pooling cross-sectional event data. We estimate 3-day cumulative abnormal returns (CARs) for each event from a market model. For Mar. 30, 2011, we use a 5-day market CAR to span 3 consecutive event days. Since not all industrial firms trade regularly, we add additional lead and lag market excess return factors to control for nonsynchronous trading effects (Dimson (1979)). Since different events associate with varying signed predictions, we sign the dependent variable in the pooled regressions, multiplying it by a value of −1 for each fully negative hypothesized event date within the sample, as shown in Table 3. We collect information on each industrial firm's total assets and leverage ratio as well as the state in which each firm is headquartered from Compustat. The set of events underlying columns 1 and 3 comprises all events listed in Table 3. The set of events underlying columns 2 and 4 comprises the subset of 8 major events for which the individual event regressions reported in Table 4 identify statistically significant effects. Standard errors are clustered by event and headquarter state. *t*-statistics are reported in parentheses. *, **, and *** indicate significance at the 10%, 5%, and 1% levels, respectively.

	Dependent Variable: 3-Day CAR (market model)			
	Sample: Industrial Firms over Event Dates			
	All 20 Events	8 Major Events	All 20 Events	8 Major Events
Variable	1	2	3	4
ln(ASSETS)	0.0001	−0.0000	0.0000	−0.0001
	(0.501)	(−0.128)	(0.143)	(−0.280)
LEVERAGE	0.0000	0.0002	0.0001	0.0002
	(0.133)	(0.407)	(0.166)	(0.395)
WI_INDICATOR	0.0003	0.0009	0.0052	0.0033
	(0.329)	(0.870)	(0.981)	(0.422)
OH_INDICATOR	−0.0009	−0.0001	0.0040	0.0023
	(−0.936)	(−0.044)	(0.759)	(0.337)
Intercept	−0.0010	0.0000	—	—
	(−0.588)	(0.012)	—	—
State fixed effects	No	No	Yes	Yes
R^2	0.0001	0.0000	0.0015	0.0019
No. of obs.	73,513	29,488	73,513	29,488

B. Bank Lending Effects

As a last set of tests, we consider the impact of pension cut legislation in Wisconsin and Ohio on bank lending. The successful passage of pension cut legislation in Wisconsin plausibly lowers the state's municipal debt default

[20]Although larger, publicly traded firms typically generate revenues beyond their headquarter states, recent literature indicates, on average, 41.4% of publicly traded firms' economic interests are driven from within firms' headquarter states, with a standard deviation of 28.3% (Bernile, Kumar, and Sulaeman (2015)), Panel A in Table 2).

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probabilities, thereby increasing asset values and freeing up lending capital of Wisconsin banks. Thus, we expect lending by Wisconsin banks to ultimately grow more for capital-constrained banks with greater municipal bond holdings. In Ohio, pension cut legislation was ultimately overturned, rendering Ohio bank municipal bond holdings an unlikely source of additional capital to promote loan growth. Thus, we expect Ohio banks' lending growth to decline, if at all, particularly for the more capital-constrained banks.

We examine cross-sectional growth in bank lending as a function of i) banks' extent of operations in the treatment states (Wisconsin and Ohio); ii) banks' exposure to state solvency outcomes of pension cut legislation through municipal bond holdings; and iii) the extent to which banks are capital constrained. In particular, we model bank lending growth with various forms of the following regression:

$$
\begin{aligned}
(6) \qquad \ln(1+\mathrm{LOAN\_GROWTH}_i) = {} & \alpha + \beta_1 \mathrm{WI}_i + \beta_2 \mathrm{OH}_i \\
& + \beta_3 \mathrm{TRADABLE\_MUNICIPAL\_BOND\_HOLDINGS}_i \\
& + \beta_4 \mathrm{LOAN\_LOSS\_RESERVES}_i \\
& + \beta_5 \mathrm{TRADABLE\_MUNICIPAL\_BOND\_HOLDINGS}_i \times \mathrm{WI}_i \\
& + \beta_6 \mathrm{TRADABLE\_MUNICIPAL\_BOND\_HOLDINGS}_i \times \mathrm{OH}_i \\
& + \beta_7 \mathrm{LOAN\_LOSS\_RESERVES}_i \times \mathrm{WI}_i \\
& + \beta_8 \mathrm{LOAN\_LOSS\_RESERVES}_i \times \mathrm{OH}_i + \varepsilon_i,
\end{aligned}
$$

where $\mathrm{LOAN\_GROWTH}_i$ is $(\mathrm{TOTAL\_LOANS}_t - \mathrm{TOTAL\_LOANS}_{t-1})/\mathrm{TOTAL\_LOANS}_{t-1}$; WI_i equals 1 for Wisconsin banks (those with Wisconsin deposit shares at least 1 standard deviation above the mean), and 0 otherwise; and OH_i similarly equals 1 for Ohio banks, and 0 otherwise. In equation (6), $\mathrm{TRADABLE\_MUNICIPAL\_BOND\_HOLDINGS}_i$ is, as in equations (3) and (4), the fair value of available-for-sale securities issued by states and political subdivisions in the United States, and $\mathrm{LOAN\_LOSS\_RESERVES}_i$ is the 3-year sum of annual allowances and provisions for loan losses provided by banks over total assets.[21] Independent variables in equation (6) are measured in 2010, and we examine loan growth over a variety of horizons spanning 2008–2012. In equation (6), we use all bank holding companies since variables in equation (6) are available in call reports data files for both publicly traded and private banks. Appendix B gives summary statistics for variables in equation (6) over the full sample of public and private bank holding companies.

Table 9 presents the regression results. Panel A examines lending growth over several horizons: 2008–2011 (column 1), 2008–2012 (column 2), 2009–2011 (column 3), and 2009–2012 (column 4). In column 1, the WI_INDICATOR coefficient indicates a 16.9% increase (significant at 1%) in loan growth for Wisconsin banks, on average, relative to other banks from 2008 to 2011. Column 2 indicates the loan growth continued through 2012, as the WI_INDICATOR coefficient

[21] In particular, we compute each year's loan loss reserves using data item number 3123 in call reports, which is the sum of the following: i) allowance for loan and lease losses (item number 3124); ii) recoveries in allowance for loan and lease losses (item number 4605); iii) provision for allowance in loan and lease losses (item number 4230); and iv) adjustments due to amended reports (item number 4815); minus charge offs for allowance in loan and lease losses (item number 4635).

TABLE 9

Municipal Bond Holdings, Capital Constraints, and Bank Loan Growth around Pension Cut Legislation

Table 9 presents regression results comparing the dependency of bank lending across bank-states on ex ante municipal holdings and the extent of loan loss reserve accumulations in bank loan portfolios in the 3 years leading up to 2011. For the set of private and publicly traded banks, we model the cross section of total bank loan growth as a function of banks' presence in Wisconsin and/or Ohio relative to changes in loan growth by other banks. Bank-state presences in Wisconsin and Ohio are further interacted with levels of bank municipal bond holdings and accumulated loan loss reserves. In Panels A–C, we test variations of the regression

$$\begin{aligned}\ln(1+\text{LOAN\_GROWTH}_i) \quad = \quad & \alpha+\beta_1\,\text{WI}_i+\beta_2\,\text{OH}_i+\beta_3\,\text{TRADABLE\_MUNICIPAL\_BOND\_HOLDINGS}_i+\beta_4\,\text{LOAN\_lOSS\_RESERVES}_i \\ & +\beta_5\,\text{TRADABLE\_MUNICIPAL\_BOND\_HOLDINGS}_i\times\text{WI}_i+\beta_6\,\text{TRADABLE\_MUNICIPAL\_BOND\_HOLDINGS}_i\times\text{OH}_i \\ & +\beta_7\,\text{LOAN\_lOSS\_RESERVES}_i\times\text{WI}_i+\beta_8\,\text{LOAN\_lOSS\_RESERVES}_i\times\text{OH}_i+\varepsilon_i,\end{aligned}$$

where LOAN_GROWTH_i is the change in total loans for bank i across a given horizon over the initial total loan level; WI_i as an indicator equals 1 for bank holding companies with 2010 deposit shares falling at least 1 standard deviation above the mean deposit shares in those states relative to the general population of private and publicly traded bank holding companies, and OH_i as an indicator is defined similarly. Independent variable levels are lagged by 1 year (determined at year end 2010). Data on deposit holdings of banks by state are from the FDIC Summary of Deposits data. Data on other bank characteristics are from call reports provided by the Federal Reserve Bank of Chicago. Standard errors are clustered by bank headquarter state. t-statistics are reported in parentheses. *, **, and *** indicate significance at the 10%, 5%, and 1% levels, respectively.

	Dependent Variable: ln(1 + LOAN_GROWTH)							
	Sample: Public and Private Banks							
	Panel A. Extreme Deposit Ratio Indicator Models				*Panel B. Continuous Deposit Ratio Models*			
	Test Horizon							
	2008–2011	2008–2012	2009–2011	2009–2012	2008–2011	2008–2012	2009–2011	2009–2012
Variable	1	2	3	4	5	6	7	8
WI	0.1690***	0.2098***	0.0709**	0.0960*	0.1703***	0.2021***	0.1102***	0.1381**
	(3.263)	(3.632)	(2.036)	(1.838)	(2.783)	(2.949)	(3.118)	(2.317)
OH	−0.1316*	−0.1135**	−0.1905***	−0.1884***	−0.1809***	−0.1922***	−0.2192***	−0.2359***
	(−1.812)	(−2.026)	(−4.737)	(−3.915)	(−2.741)	(−2.813)	(−6.758)	(−4.878)
TRADABLE_MUNICIPAL_BOND_HOLDINGS × WI	0.2760***	0.4776***	0.3027***	0.4984***	0.2672***	0.5158***	0.4225***	0.6728***
	(4.301)	(5.387)	(5.935)	(6.845)	(7.691)	(11.863)	(11.455)	(17.802)
TRADABLE_MUNICIPAL_BOND_HOLDINGS × OH	−0.0208	−0.0324*	0.0129	−0.0161	−0.0284	−0.0402	0.2838	0.1326
	(−1.210)	(−1.914)	(0.446)	(−0.751)	(−0.153)	(−0.169)	(1.292)	(0.497)
LOAN_LOSS_RESERVES × WI	−4.6753***	−5.9340***	−2.2699***	−3.2861***	−4.6208***	−5.8707***	−3.3120***	−4.6294***
	(−3.956)	(−4.454)	(−2.792)	(−2.728)	(−3.249)	(−3.554)	(−4.046)	(−3.168)
LOAN_LOSS_RESERVES × OH	2.8512*	2.7722**	4.2790***	4.4273***	4.0188**	4.3436**	4.8115***	5.3043***
	(1.976)	(2.127)	(5.508)	(4.229)	(2.452)	(2.316)	(6.062)	(3.842)
TRADABLE_MUNICIPAL_BOND_HOLDINGS	0.0594***	0.0823***	0.0109	0.0527**	0.0577***	0.0784***	0.0045	0.0470**
	(3.508)	(5.139)	(0.381)	(2.630)	(3.404)	(5.035)	(0.188)	(2.559)
LOAN_LOSS_RESERVES	−6.4978***	−6.9975***	−3.6865***	−4.7178***	−6.5250***	−7.0592***	−3.6884***	−4.7328***
	(−5.872)	(−5.516)	(−7.336)	(−4.948)	(−6.001)	(−5.629)	(−7.508)	(−5.034)
Intercept	0.2405***	0.2622***	0.1170***	0.1642***	0.2414***	0.2654***	0.1170***	0.1653***
	(4.858)	(4.878)	(4.759)	(4.022)	(4.990)	(4.989)	(4.902)	(4.065)
R^2	0.176	0.121	0.151	0.0726	0.175	0.120	0.150	0.0718
No. of obs.	637	609	656	631	637	609	656	631

(continued on next page)

TABLE 9 (continued)

Municipal Bond Holdings, Capital Constraints, and Bank Loan Growth around Pension Cut Legislation

Panel C. Continuous Deposit Ratios, Loan Growth, and the Impact of Bank Municipal Bond Holdings under Capital Constraints

	Dependent Variable: ln(1 + LOAN_GROWTH)			
	Sample: Public and Private Banks			
	Test Horizon			
	2008–2011	2008–2012	2009–2011	2009–2012
Variable	1	2	3	4
WI	0.1149**	0.0692	−0.0220	−0.0320
	(2.191)	(1.170)	(−0.561)	(−0.571)
OH	−0.1825***	−0.1906***	−0.2452***	−0.2368***
	(−2.737)	(−2.797)	(−5.782)	(−4.898)
TRADABLE_MUNICIPAL_BOND_HOLDINGS × WI	2.1828**	5.3858***	3.9366***	6.9740***
	(2.541)	(4.155)	(5.626)	(6.145)
TRADABLE_MUNICIPAL_BOND_HOLDINGS × OH	0.0008	−0.0755	−0.3343	0.1391
	(0.004)	(−0.329)	(−0.591)	(0.546)
LOAN_LOSS_RESERVES × WI	−3.7178***	−3.7933**	−0.9567	−1.9847
	(−2.992)	(−2.651)	(−1.038)	(−1.475)
LOAN_LOSS_RESERVES× OH	3.9940**	4.3431**	6.2438***	5.3097***
	(2.470)	(2.336)	(5.134)	(3.846)
TRADABLE_MUNICIPAL_BOND_HOLDINGS × LOAN_LOSS_RESERVES × WI	−28.2579**	−71.8804***	−53.9475***	−92.6511***
	(−2.290)	(−3.863)	(−5.205)	(−5.674)
TRADABLE_MUNICIPAL_BOND_HOLDINGS × LOAN_LOSS_RESERVES	0.7190	1.0147**	2.9832*	0.3981
	(1.246)	(2.543)	(1.930)	(0.862)
TRADABLE_MUNICIPAL_BOND_HOLDINGS	0.0125	0.0255	−0.0753**	0.0234
	(0.353)	(1.066)	(−2.552)	(0.665)
LOAN_LOSS_RESERVES	−6.6379***	−7.1486***	−4.6271***	−4.7862***
	(−6.065)	(−5.664)	(−7.023)	(−4.997)
Intercept	0.2473***	0.2700***	0.1566***	0.1681***
	(5.046)	(5.012)	(5.071)	(4.024)
R^2	0.177	0.122	0.139	0.0737
No. of obs.	637	609	656	631

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increases to 0.2098 (significant at 1%). Starting from 2009, we see a more moderate but similar continued loan growth pattern across columns 3 and 4 related to the 2009–2011 and 2009–2012 horizons, respectively. Across columns 1–4, the interaction coefficients between the WI_INDICATOR and tradable municipal bond holdings show, over each horizon, loan growth intensifies in proportion to Wisconsin bank municipal bond holdings. For example, from Panel A of Table B1 in Appendix B, we see total loans for Wisconsin banks averaged \$17.3 billion in 2008. Given that we scale municipal bond holdings data in units of billions of dollars for ease of presentation, the interaction coefficient of 0.2760 (significant at 1%) in column 1 suggests an increase of \$100 million in Wisconsin bank municipal bond holdings increases lending by \$4.775 billion (0.2760 $\times$ \$17.3 billion). This is consistent with greater municipal bond holdings by Wisconsin banks supporting capital adequacy for additional lending over the 2008–2011 horizon. Next, we note the relatively larger Wisconsin bank municipal bond holding interaction coefficient in column 2 (0.4776) versus column 1 (0.2760) and, similarly, in column 4 (0.4984) versus column 3 (0.3027), which shows the relative economic importance of the impact of Wisconsin bank municipal bond holdings during 2012. For example, the relatively larger interaction coefficient in column 2 (0.4776, significant at 1%) suggests a \$100 million increase in Wisconsin bank municipal bond holdings lending by \$8.262 billion (0.4776 $\times$ \$17.3 billion), on average, over the 2008–2012 horizon, which is 73% greater than the \$4.77 billion estimate in column 1 over 2008–2011. Altogether, the results suggest greater municipal bond holdings by Wisconsin banks provide liquidity for additional lending by year end 2011 and particularly during 2012.

Columns 1–4 report negative coefficients on the Wisconsin bank interaction with loan loss reserves (e.g., −4.6753 in column 1). These results signify that Wisconsin bank loan growth is concentrated in more capital-constrained banks with lower reserve accumulations over the 3 years leading up to 2011. In particular, the transmission of lending capital from additional Wisconsin municipal bond holdings offsets capital adequacy constraints on lending from lower loan loss reserves. For example, with total assets of Wisconsin banks averaging \$24.4 billion in 2008, a 1% increase in loan loss reserves over total assets results in \$244 million toward satisfying new-loan capital adequacy requirements. In Panel A, the same dollar-for-dollar increase of \$244 million in Wisconsin municipal bond holdings results in \$11.651 billion in additional lending (the \$4.775 billion computed previously for column 1 times 2.44, since that calculation is based on a \$100 million increase in municipal bond holdings). Moreover, if an increase in liquid asset values due to municipal bond holdings relaxes capital adequacy constraints on lending for banks with lower loan loss reserve accumulations, we expect a liquidity substitution effect to allow lending to *increase* more for constrained Wisconsin banks with lower loan loss reserve accumulations relative to Wisconsin banks with higher loan loss reserve accumulations. Consistent with the substitution effect, the column 1 interaction coefficient on loan loss reserves and Wisconsin banks, −4.6753 (significant at 1%), means that each 1% *decline* in Wisconsin bank loan loss reserves *increases* total lending by \$808 million (−4.6753 $\times$ \$17.3 billion, the mean of total loans for Wisconsin banks in 2008) over the 2008–2011 horizon. Finally, from the Ohio indicator coefficients, across columns 1–4, we observe that lending decreases, on average, for Ohio banks and for Ohio banks there is no lending growth sensitivity to municipal bond holdings. In contrast to results for Wisconsin banks, in the absence of loan growth sensitivity to municipal bond holdings, interactions between loan loss reserves and the Ohio indicator show that the declines in Ohio bank lending are greater for Ohio banks with lower initial loan loss reserves. We consider this to be falsification evidence consistent with removed access to the municipal bond lending channel for Ohio banks, which, in the absence of liquidity substitution effects, delivers a more typical pattern of reduced lending by more capital-constrained Ohio banks.

In Panel B of Table 9, we repeat the specifications from Panel A. However, rather than using discrete indicator variables based on extreme deposit ratios to designate Wisconsin and Ohio banks, we use banks' Wisconsin and Ohio deposit ratios as continuous test variables. We find the patterns of Panel A carry over virtually unchanged to Panel B.[22]

The evidence for the liquidity substitution effect we find in Panels A and B would be more conclusive if loan growth proportionality to municipal bond holdings for Wisconsin banks were more than coincident with the increased lending found in more capital-constrained Wisconsin banks. In particular, to examine whether the municipal bond lending effects for Wisconsin banks are greater, particularly within the more constrained Wisconsin banks, Panel C reports regression results similar to the those in Panel B but including the triple interaction $\text{TRADABLE\_MUNICIPAL\_BOND\_HOLDINGS}_i \times \text{LOAN\_LOSS\_RESERVES}_i \times \text{WI}_i$.[23] Across the four specifications, the triple interactions are consistently negative and statistically significant, while the $\text{TRADABLE\_MUNICIPAL\_BOND\_HOLDINGS}_i \times \text{WI}_i$ interactions remain positive and significant, but with much higher coefficients than in Panel B. In this specification, the greater $\text{TRADABLE\_MUNICIPAL\_BOND\_HOLDINGS}_i \times \text{WI}_i$ interaction coefficients show the more extreme impact of municipal bond holdings on loan growth for the most constrained Wisconsin banks, those with no loan loss reserve accumulations. Further, the negative loadings on the triple interactions show that the impact of Wisconsin bank municipal bond holdings on lending growth declines for the banks with greater loan loss reserve accumulations. Thus, the two interactions in Panel C provide more direct evidence of municipal bond liquidity substitution effects in Wisconsin banks. Finally, the Wisconsin indicator direct effect coefficients in Panel C are greatly reduced

[22]We further examine the source of loan growth across standard bank-lending reporting categories (namely, real estate loans, commercial and industrial loans, loans to depository institutions, agriculture loans, consumer loans, and loans to foreign governments), and find growth in real estate lending (primarily) and agriculture lending (secondarily) delivers economic and statistical patterns highly similar to those of Table 9. We interpret these results more as corroborating but not generalizable, as particular sources of loan growth across lending categories are likely a function of the local economy. The results by lending category are available from the authors.

[23]We also include the interaction $\text{TRADABLE\_MUNICIPAL\_BOND\_HOLDINGS}_i \times \text{LOAN\_LOSS\_RESERVES}_i$ as a control. The results of Panel C in Table 9 are fully robust to the alternative specification in Panel A, which uses Wisconsin and Ohio indicator variables instead of continuous deposit ratios. The results of Panel C are also fully robust to the inclusion of the triple interaction $\text{TRADABLE\_MUNICIPAL\_BOND\_HOLDINGS}_i \times \text{LOAN\_LOSS\_RESERVES}_i \times \text{OH}_i$, which is not significant and is consistent with there being no Ohio bank lending dependency on liquidity substitution effects.

relative to those in Panel B, with only the coefficient in column 1 of Panel C (0.1149) remaining statistically significant. When we condition the impact of Wisconsin bank municipal bond holdings on bank capital constraints, the reduced direct effect coefficients suggest that Wisconsin bank loan growth around pension cut legislation is primarily explained by the expansion of credit supply from more capital-constrained banks through their municipal bond holdings.

V. Conclusions

This study provides an empirical analysis of the impact of Wisconsin and Ohio pension cut legislation on the stock values and operations of banks concentrated in Wisconsin and Ohio. We find that banks doing business in Wisconsin and Ohio experience positive (negative) stock price reactions to announcements that indicate an increased (a decreased) probability of pension cut legislation. The stock price reactions are further explained by the extent of Wisconsin and Ohio bank municipal bond holdings, particularly for banks with lower capital adequacy. To further corroborate municipal bond valuation effects, we find that Wisconsin and Ohio municipal bond spreads decrease (increase) significantly in reaction to announcements indicating an increased (decreased) probability of pension cut legislation. Examining effects of pension cut legislation on industrial firms, we find no evidence that the increased probability of pension legislation passed in Wisconsin and Ohio impacts stock prices of industrial firms headquartered in these two states. Thus, stock price reactions are unique to banks operating in pension cut states.

We find that total lending by banks operating in Wisconsin increases over and after the period in which pension cut legislation is enacted, while total lending by banks operating in Ohio decreases over and after the period in which pension cut legislation is enacted and, eventually, overturned. We find that these different lending patterns are largely a function of bank municipal bond holdings. Finally, we find that the increased loan growth from Wisconsin banks with greater municipal bond holdings comes from the riskier portfolios of capital-constrained banks, where lower reserves have been set aside for loan losses. This suggests a new lending channel linked to state solvency, whereby capital-constrained banks disproportionately supply credit as municipal bond appreciations free up capital.

Appendix A. Supplemental Discussion of the Events Leading to Enactment of Pension Cut Legislation in Wisconsin and Ohio, and Its Subsequent Repeal in Ohio

Wisconsin's pension cut bill took a tumultuous and highly publicized path to passage. On Dec. 7, 2010 (event 1 listed in Table 3), Governor-elect Walker first released details of a proposal to save millions of dollars by having state employee union members pay more into pension funds and pay more for health insurance. To get such concessions, Governor-elect Walker raised the possibility of changing state law to decertify the unions. On Feb. 2, 2011 (event 3), Governor Walker affirmed in his State of the State address that state of Wisconsin employees needed to contribute to their pension costs and pay more for health insurance to help balance the Wisconsin state budget. These comments quickly became legislation, as a bill eliminating most collective bargaining rights from nearly all Wisconsin public

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employees passed the legislature's budget writing committee on Feb. 17 (event 4). After this vote, Democratic lawmakers left the state in an attempt to stop a vote on the bill. Despite this, on Feb. 24 (event 5), the Assembly reached an agreement on the contents of the bill; on Mar. 9 (event 7), the Assembly passed the bill (despite the absence of 14 Democratic senators); and on Mar. 11 (event 8), Governor Walker signed the bill into law. This was not the end, however, as just 5 days later (event 9), the Dane County district attorney filed a legal challenge to the bill, stating that Republican lawmakers violated Wisconsin's open meetings law (by not giving the proper public notice that the committee planned to meet) when they amended the plan. The challenge requested that a judge void the law and issue an emergency order blocking the secretary of state from publishing the law. On Mar. 18 (event 10), the judge temporarily blocked the law from taking effect. In an unexpected and confusing move, on Mar. 25 (event 11), the Legislative Reference Bureau published the law (the last step before the law goes into effect), despite the court order blocking its publication, while challenges to the law were being considered. However, on Mar. 31 (event 14), the law was put on indefinite hold by the same Dane County Circuit Court judge until the case could be heard by the Wisconsin Supreme Court. Thus, at this point, the bill would not go into effect. Further, the Supreme Court had not indicated whether it would even take the case. While the Wisconsin Supreme Court deliberated, on May 27, 2011, Dane County Judge MaryAnn Sumi issued a permanent injunction that effectively threw out the pension cut bill (event 16). Judge Sumi concluded that the Republicans passed the bill by violating the state's strong "open meeting" law that requires 24 hours' notice of official meetings and that the law was, thus, invalid. This decision was reversed on June 15, 2011 (event 17), when the Wisconsin Supreme Court rejected the ruling from the county court that invalidated the pension cut bill. The unusually quick decision was decided by a vote of 4 to 3. Two weeks later, on June 29, 2011, the pension cut bill took effect (event 18).

Ohio's pension cut bill was passed in a more typically smooth manner than was Wisconsin's bill. On Feb. 1, 2011, Bill 5 (the pension cut bill) was introduced in the Ohio Senate (event 2 in Table 3). This was followed by Senate approval (on Mar. 2 (event 6)), approval by a House panel (on Mar. 29 (event 12)), and approval by the full House and Senate (on Mar. 30 (event 13)); the bill was signed into law on Mar. 31, 2011 (event 14). After this relatively swift passage, however, opponents of the bill started the repeal process. On Apr. 15, the Ohio attorney general certified summary language for a referendum seeking repeal of Senate Bill 5 (event 15). This language needed to be certified in order for opponents of the bill to start collecting the 231,150 signatures (within 90 days of the passage of the bill) needed to get it on the November ballot. On June 29 (event 18), opponents marched to the secretary of state's office to hand over the petitions with 1,298,301 signatures, and on July 21, the Ohio secretary of state announced that sufficient signatures had been certified to put a repeal of Bill 5 on the November ballot as a veto referendum (event 19). On Nov. 8, 2011, the bill was overwhelmingly repealed: by a vote of 61% in favor of repeal to 39% against repeal (event 20). Table A1 presents regression results on the impact of pension cut legislation events on bank values for event dates in Table 3 that are not featured in Table 4.

TABLE A1

Regression Results of Stock Price Reactions to Other News Announcements Associated with the Passage of Pension Cut Legislation

Table A1 presents regression results for the announcements associated with passage of pension cut legislation in Wisconsin and Ohio found in Table 3 but not in Table 4. Cumulative abnormal returns (CARs) for each event are estimated from a market model. Since not all public banks trade regularly, additional lead and lag market excess return factors are added to control for nonsynchronous trading effects (Dimson (1979)). For each event t, the following regressions are run across all publicly traded bank holding companies (with each bank denoted by subscript i):

$$\begin{aligned}\text{CAR}_i = \ & \alpha + \delta_1 \ln(\text{ASSETS}_i) + \delta_2 \ln(1 + \text{TRADABLE\_MUNICIPAL\_BOND\_HOLDINGS}_i) + \delta_3 \text{TIER\_I\_CAPITAL\_RATIO}_i \\ & + \delta_4 \text{DEPOSITS\_HELD\_IN\_WI\_TO\_TOTAL\_DEPOSITS}_i \\ & + \delta_5 \text{DEPOSITS\_HELD\_IN\_OH\_TO\_TOTAL\_DEPOSITS}_i + \varepsilon_i.\end{aligned}$$

Independent variable levels are lagged by 1 year (determined at year end 2010). Data on deposit holdings of banks by state are from the FDIC Summary of Deposits data. Data on other bank characteristics are from call reports provided by the Federal Reserve Bank of Chicago. Standard errors are clustered by bank headquarter state. t-statistics are reported in parentheses. *, **, and *** indicate significance at the 10%, 5%, and 1% levels, respectively.

	Dependent Variable: 3-Day CAR (market model)					
	Sample: Publicly Traded Bank Holding Companies					
Variable	Dec. 7, 2010	Feb. 2, 2011	Feb. 24, 2011	Mar. 11, 2011	Mar. 16 2011	Mar. 18, 2011
ln(ASSETS)	0.0059***	−0.0032	−0.0033	0.0002	0.0039**	−0.0020
	(3.830)	(−1.659)	(−0.986)	(0.222)	(2.128)	(−1.667)
TIER_I_CAPITAL_RATIO	0.1907	0.3466**	−0.2539	−0.1238	0.1761**	0.1074
	(1.363)	(2.107)	(−1.309)	(−1.464)	(2.038)	(1.163)
ln(1 + TRADABLE_MUNICIPAL_BOND_HOLDINGS)	0.0003	0.0014*	0.0003	0.0000	0.0010	−0.0000
	(0.815)	(1.785)	(0.323)	(0.013)	(1.565)	(−0.043)
DEPOSITS_HELD_IN_WI_TO_TOTAL_DEPOSITS	−0.0040	0.0045	−0.0025	−0.0223	0.0143**	0.0416***
	(−0.518)	(0.606)	(−0.427)	(−1.649)	(2.366)	(7.765)
DEPOSITS_HELD_IN_OH_TO_TOTAL_DEPOSITS	0.0015	−0.0044	−0.0202*	−0.0002	0.0026	−0.0100**
	(0.365)	(−0.692)	(−1.794)	(−0.073)	(0.544)	(−2.233)
Intercept	−0.1104***	−0.0116	0.0857	0.0050	−0.0882**	0.0107
	(−3.127)	(−0.246)	(1.203)	(0.230)	(−2.630)	(0.398)
R^2	0.0706	0.0383	0.0155	0.0140	0.0412	0.0181
No. of obs.	317	319	319	318	318	318

	Dependent Variable: 3-Day CAR (market model)					
	Sample: Publicly Traded Bank Holding Companies					
Variable	Mar. 29, 2011	Mar. 31, 2011	Apr. 15, 2011	June 15, 2011	July 21, 2011	Nov. 9, 2011
ln(ASSETS)	0.0024	0.0001	0.0025	0.0032	0.0023*	0.0025*
	(1.648)	(0.093)	(1.623)	(1.394)	(1.930)	(1.697)
TIER_I_CAPITAL_RATIO	−0.0192	0.1713	0.3929**	0.1810	−0.1114	0.2715***
	(−0.122)	(1.579)	(2.162)	(1.036)	(−1.164)	(2.708)
ln(1 + TRADABLE_MUNICIPAL_BOND_HOLDINGS)	−0.0006	0.0011**	0.0005	0.0012	−0.0004	0.0000
	(−0.488)	(2.256)	(0.647)	(1.380)	(−0.642)	(0.042)
DEPOSITS_HELD_IN_WI_TO_TOTAL_DEPOSITS	0.0001	0.0066	0.0259***	−0.0176	0.0032	−0.0127
	(0.013)	(0.799)	(4.747)	(−1.623)	(0.563)	(−1.045)
DEPOSITS_HELD_IN_OH_TO_TOTAL_DEPOSITS	−0.0011	0.0083*	0.0073	0.0031	−0.0027	0.0202***
	(−0.213)	(1.932)	(1.402)	(0.462)	(−0.551)	(3.797)
Intercept	−0.0332	−0.0287	−0.0971**	−0.0769	−0.0188	−0.0803**
	(−0.925)	(−0.838)	(−2.033)	(−1.351)	(−0.680)	(−2.574)
R^2	0.00566	0.0358	0.0741	0.0413	0.0121	0.0487
No. of obs.	318	318	316	310	307	300

Appendix B. Descriptive Statistics for Public and Private Bank Holding Companies

Appendix B presents descriptive statistics on financial characteristics for the sample banks. Data are collected at year end 2010 for Panel A of Table B1 and according to the

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noted year in Panel B. Wisconsin and Ohio banks are defined as publicly traded bank holding companies with 2010 deposit shares falling at least 1 standard deviation above the mean deposit shares in those states relative to the general population of private and publicly traded bank holding companies. Data on deposit holdings of banks by state are from the FDIC Summary of Deposits data. Data on other bank characteristics are from call reports provided by the Federal Reserve Bank of Chicago.

TABLE B1

Descriptive Statistics for Public and Private Bank Holding Companies

	Bank Holding Company Categories		
Variable	Wisconsin Banks	Ohio Banks	Other Public and Private Banks
Panel A. Deposit Shares, Municipal Bond Holdings, and 3-Year Loan Loss Reserve Accumulations for Public and Private Banks at Year End 2010			
No. of banks	28	22	607
DEPOSITS_IN_WI/TOTAL_DEPOSITS (%)			
Mean	73.56	0.39	0.01
Minimum	7.76	0	0
Median	100	0	0
Maximum	100	8.49	6.03
DEPOSITS_IN_OH/TOTAL_DEPOSITS (%)			
Mean	0.62	73.86	0.02
Minimum	0	10.74	0
Median	0	87.10	0
Maximum	12.85	100	4.31
TRADABLE_MUNICIPAL_BOND_HOLDINGS (in USD millions)			
Mean	292.92	362.14	73.20
Minimum	0	0	0
Median	26.35	39.80	22.62
Maximum	6,303	6,303	6,604
LOAN_LOSS_RESERVES (%)			
Mean	4.75	4.32	4.56
Minimum	2.14	1.43	1.43
Median	4.63	4.06	4.20
Maximum	9.77	8.17	13.26
Panel B. Total Loan Levels for Public and Private Banks from 2008 to 2012			
No. of banks in 2010	28	22	607
TOTAL_LOANS in 2008 (in USD billions)			
Mean	17.3	32.90	3.39
Minimum	0.55	0.49	0
Median	1.31	1.57	1.28
Maximum	376	376	192
TOTAL_LOANS in 2009 (in USD billions)			
Mean	17.0	30.9	3.77
Minimum	0.54	0.50	0
Median	1.30	1.58	1.28
Maximum	391	391	237
TOTAL_LOANS in 2010 (in USD billions)			
Mean	8.61	16.2	2.14
Minimum	0.30	0.27	0.05
Median	0.63	0.73	0.69
Maximum	202	202	119
TOTAL_LOANS in 2011 (in USD billions)			
Mean	18.1	32.4	3.75
Minimum	0.54	0.58	0
Median	1.21	1.57	1.24
Maximum	429	429	249
TOTAL_LOANS in 2012 (in USD billions)			
Mean	19.9	34.3	3.78
Minimum	0.55	0.61	0
Median	1.19	1.63	1.22
Maximum	460	460	250

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