A spatial analysis of state banking regulation*

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Abstract. We use a spatial model to investigate a state's choice of branch banking and interstate banking regimes as a function of the regime choices made by other states and other variables suggested in the literature. We extend the basic spatial econometric model by allowing spatial dependence to vary by geographic region. Our findings reveal that spatial effects have a large, statistically significant impact on state regulatory regime decisions. The importance of spatial correlation in the setting of state banking policies suggests the need to consider spatial effects in empirical models of state policies in general.

JEL classification: G28, C23, R10

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1 Introduction

The U.S. banking industry has consolidated rapidly over the past two decades. From a postwar peak of 14,496 banks in 1984, the number of U.S. commercial banks had fallen to 7,789 banks by the end of 2003. Over the same period the average size of banks, measured in terms of total assets, increased from $307 million in 1984 (in 2003 dollars) to $979 million in 2003.

The consolidation and increased average size of U.S. banks have coincided with a substantial relaxation of geographic restrictions on the location of bank

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branches and bank holding company subsidiaries (Rhoades 2000). The Riegle-Neal Interstate Banking and Branching Efficiency Act of 1994 eliminated federal restrictions on interstate banking and branching by U.S. commercial banks.[1] The legislation also eliminated most of the remaining state barriers to interstate banking; it was implemented after many states had already relaxed their restrictions on branching within their borders. In 1970 only twelve states permitted statewide branching, and none allowed entry by bank holding companies headquartered in other states.[2] But by 1994, all states except Iowa permitted statewide branching through the acquisition of existing bank offices, and many allowed for branching through the establishment of entirely new offices. Also by 1994 all states except Hawaii permitted some entry by out-of-state holding companies. Whereas less than half of all U.S. commercial banks operated any branch offices in 1984, 71% of banks had multiple offices in 2003.

Prior studies have concluded that advances in information-processing technology explain why states began to relax geographic restrictions on banking in the 1970s and 1980s. Such advances facilitated the development of ATM networks, cash management accounts and other alternatives to bank deposits, and enabled outside lenders to penetrate local banking markets, all of which reduced incentives to defend legal restrictions on the location of bank branches and bank holding company subsidiaries.[3]

The pattern and timing of deregulation varied considerably across states, however, and this article presents new evidence on why states chose particular interstate banking and intrastate branching regulations. Prior studies such as Kane (1996) and Kroszner and Strahan (1999), attribute differences in the timing of deregulation to the relative power of interest groups that benefited from the status quo versus those that would benefit from expanded geographic powers for banks. Banks historically have located in small communities, therefore rural areas lobbied against legislation to permit branch banking, whereas large city banks generally favoured branching. Consumers of banking services were often similarly divided. Farming and other small town interests often opposed branch banking, hoping to ensure that their local banks would continue to supply credit during economic downturns (Calomiris 1992). Other consumers of banking services favoured branching, however, through their desire for more convenient and stable banking systems.[4]

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[1] Interstate banking refers to the location of bank subsidiaries of bank holding companies in different states. Interstate branching refers to the location of bank branches in different states.

[2] When states enacted laws prohibiting entry by out-of state holding companies, they typically did not force out-of-state holding companies to give up their existing banks.

[3] Petersen and Rajan (2002) argue that advances in computing and communications technology made quantifiable information about potential borrowers more readily available and thereby reduced the value of "soft" information in small business lending. Thus close proximity between borrowers and lenders became less important than in the past. Kroszner and Strahan (1999) also discuss the relationship between advances in information technology and branching deregulation.

[4] Historically, larger branching banks have fared better during banking crises. An earlier wave of branching deregulation occurred during the Great Depression. Abrams and Settle (1993) find that deregulation in that era was more likely to occur where interests favourable to branching had relatively more political strength, and in states that experienced higher bank failure rates, which were more numerous among small, unit banks than among large, branching banks.

In addition to the influence of interest groups, we believe that a state's choice of bank regulatory regime could, in part, reflect the choices made by other states. Several studies have noted regional differences in state banking laws. For example, states in the Midwest and South historically had the most restrictive branching laws, likely reflecting a relatively strong influence exerted by small banks and rural interests upon state legislatures. Such states also were among the last to deregulate.[5] The first form of interstate banking deregulation consisted of regional compacts that permitted holding companies headquartered in one member state to locate subsidiary banks in the other member states. Almost by definition, one state's decision to enter such a compact was dependent on the decisions of other member states.

A state's decision to adopt a new branching regime within its borders might also have been influenced by the branching regulations adopted by its neighbours. Kroszner and Strahan (1999) argue that the possibility of participating in a regional compact could have influenced a state's decision to permit intrastate branching, because states typically relaxed branching restrictions before entering compacts. Moreover, states could have been influenced by the effects of deregulation on banking markets, access to banking services, or economic growth in neighbouring states that deregulated first.[6]

Whereas anecdotal evidence suggests the possibility of interstate dependence in state branching and interstate banking policies, prior studies have not tested explicitly for spatial patterns or dependence in the choice of regulatory regime. Strong evidence of spatial dependence has been found in the analysis of other state policies, however, such as lotteries (Alm et al. 1993; Garrett and Marsh 2002), budgeted expenditures (Case et al. 1993), and tax rates (Brueckner and Saavedra 2001; Hernandez 2003). Failure to account for spatial dependence or spatially-correlated errors in an empirical model of regime choice can lead to erroneous conclusions about why states chose particular policies.

In the present article we test for spatial dependence on bank regulatory decisions by incorporating spatial effects directly into an empirical model of regime choice. Specifically, we estimate probit models of the choices between permitting state-wide branch banking ("intrastate branching") or not, and of permitting entry by out-of-state bank holding companies ("interstate banking") or not. The spatial probit model is a flexible and established framework for relaxing the assumption of cross-sectional independence.[7] Further, our model allows spatial dependence to vary across geographic regions. We find strong evidence that a state's choice

[5] On the historical differences in bank regulation across states, see White (1983).

[6] The experiences of nearby states might have more influence on a state's decisions because similar employment patterns or industries might make the experiences of nearby states seem more relevant than those of distant states, or simply because rivalries are stronger among nearby states.

[7] By contrast, spatial interdependencies cannot be modeled adequately with a hazard model, because in the hazard model, observations on individual states are no longer influential once the event of interest (e.g., deregulation) has occurred. The hazard model also ignores the possibility that a state could change regime more than once. Although no state tightened its branching laws during the period of our analysis, some states have done so historically (White 1983). Conversely, discrete choice models, such as the probit, cannot make use of information about the timing of events as well as can the hazard model.

of regulatory regime was influenced by decisions made by other states, but that the size of this influence varied across regions.[8] We find, moreover, that certain results others have obtained about the determinants of deregulation are not robust to the inclusion of spatial effects in our model. For example, in contrast to Kroszner and Strahan (1999), we find no evidence that the size of a state's small business sector has affected the choice of banking regime, and only weak evidence of a relationship between state branching or interstate banking policies and state regulation of insurance sales by banks. That said, however, our results strongly support the widely held view that a state was less likely to adopt a liberal branching regime when its banking system was dominated by small banks.

2 Hypotheses about the choice of regulatory regime

In their empirical study of the removal of state branching restrictions, Kroszner and Strahan (1999) test various hypotheses associated with private-interest, public-interest, and political-institutional models of regulation. We include many of their variables in our model of regime choice. In addition, we examine the influence of spatial effects on regime choice and include an expanded set of variables to capture the influence of partisan politics on the regulatory decision.

Patterns of state deregulation of intrastate branching and interstate banking are illustrated in Figs. 1 and 2, respectively. Twelve states located primarily on the West Coast, New England, and the Carolinas permitted statewide branching before 1970.[9] Beginning in 1970, deregulation spread westward, beginning in the Northeast, then moving to the South, and finally to the Midwest and Great Plains. The opening of states to interstate banking followed a similar pattern, with states in the East and Far West generally deregulating before those in the Midwest and Plains. While these spatial patterns do not necessarily indicate that state decisions about banking regulations were interdependent, they are suggestive of the need for further study.

The Bank Holding Company Act of 1956 prohibited interstate banking except in states that explicitly permitted the acquisition of their state banks by out-of-state holding companies. No state enacted such legislation until 1975, when Maine became the first state to permit out-of-state holding companies to acquire its banks.

Other states gradually followed suit, often enacting laws that required reciprocity from states whose holding companies wished to enter their markets. Regional compacts were established in New England and the Southeast in which each member state permitted entry by holding companies based in any other

[8] Although spatial discrete choice models, such as the spatial probit model, have had several applications in the literature (Case 1992; Marsh et al. 2000; Murdoch et al. 2003), only Marsh et al. 2000 tested for regional differences in patterns of spatial autocorrelation.

[9] Other states permitted limited branching within market areas, contiguous counties, etc., or prohibited branching altogether (Spong 1994).

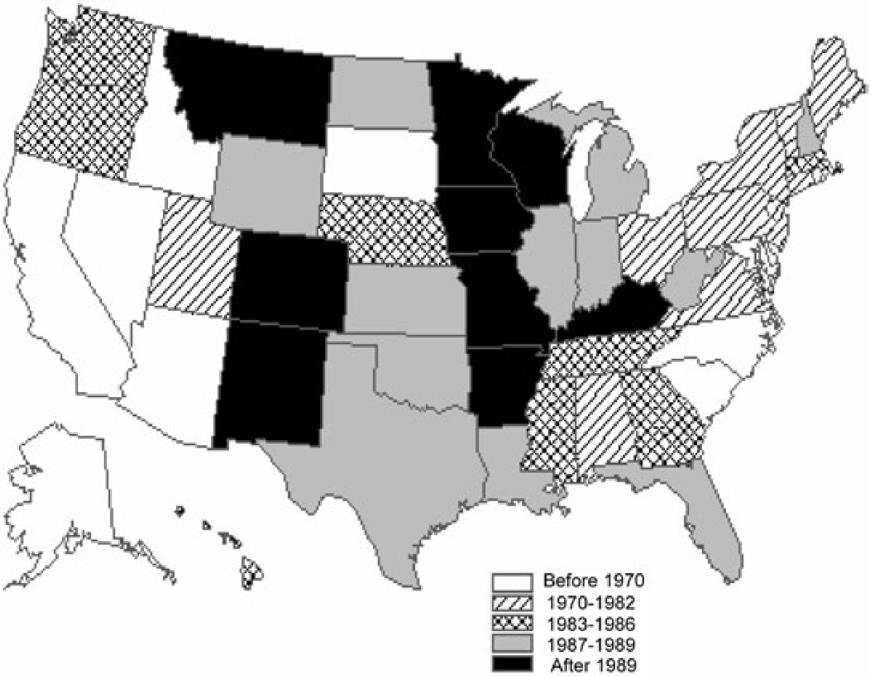

Fig. 1. When states first permitted intrastate branching

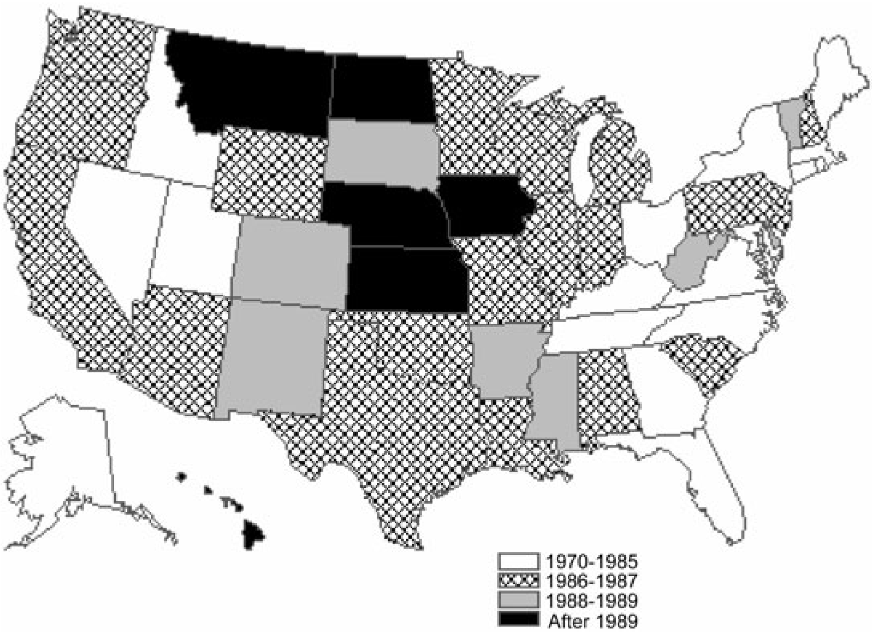

Fig. 2. When states first permitted interstate banking

member state. Elsewhere, individual states enacted laws that permitted entry by holding companies headquartered in contiguous states, usually with reciprocity.[10]

Agreements between nearby states to allow entry by each other's holding companies are suggestive of spatial dependence in the choice of interstate banking regime. Spatial dependence in the choice of intrastate branching regime is suggested by the fact that states often relaxed their restrictions on intrastate branching as a precursor to entering interstate banking agreements with other states. Evidence reported in Jayaratne and Strahan (1996) further suggests that deregulation had a large impact on the performance of banks and state economies.[11] Such dramatic effects probably would not have gone unnoticed in other states.

Whereas Kroszner and Strahan (1999) point to advances in communications and information processing technology and financial innovation as the fundamental reasons why geographic restrictions on banks were relaxed beginning in the 1970s, they find that differences in the relative power of interest groups, as well as political-institutional differences, explain differences in the timing of deregulation across states. We include the variables that Kroszner and Strahan (1999) find to be important determinants of the timing of deregulation in our empirical model of regime choice.

The relative political influence of small and large banks has often been cited as an important determinant of a state's choice of branch banking regulations. Traditionally, small banks located in small markets favoured restrictive branching laws, presumably to limit competition from large, urban banks. Following Kroszner and Strahan (1999), we include the fraction of a state's banking assets controlled by banks smaller than the state median to test this hypothesis. We also include the difference between the average capital-to-asset ratios of small and large banks to test whether the relative financial strength of small banks influenced state regulatory decisions, where "small" and "large" are determined relative to the median bank in terms of total assets. A state with financially weak small banks might have viewed the adoption of liberal branching or interstate banking rules as ways of increasing the supply of credit, and weak banks might not have had the resources to fight such changes in regulation.

Kroszner and Strahan (1999) argue that rivalry between banks and insurance companies also affected the timing of branching deregulation. Hence, we include an indicator variable for whether or not a state permits banks to sell insurance under the hypothesis that insurance companies have a stronger incentive to oppose relaxation of branching laws in states that permit banks to sell insurance. We also include the ratio of total insurance sector assets to the sum of insurance and banking assets within a state to test the hypothesis that a relatively large insurance sector made the adoption of liberal branching and interstate banking laws less likely.

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[10] See Spong (1994).

[11] Jayaratne and Strahan (1996) argue that deregulation enabled better performing banks to grow faster than weaker banks, which caused average operating costs and loan losses to decline sharply. They also estimate that state per capita income growth increased by as much as 33% after a state eliminated its restrictions on branch banking. See Freeman (2002), however, for evidence suggesting that deregulation had a much smaller impact on state growth rates.

We also follow Kroszner and Strahan (1999) and test whether the relative size of a state's small business sector was an important determinant of regime choice. Although small firms might view small banks as a more reliable source of credit than large banks, branching deregulation also tends to reduce local market power to the benefit of bank customers. Like Kroszner and Strahan (1999), we construct this variable as the ratio of firms with less than 20 employees to the total number of all firms in a state.

We also include state real per capita income and the federal funds interest rate in our model to capture other possible bank customer-related influences on the choice of regulatory regime. Presumably, the demand for banking services is positively associated with income levels. Thus, the consumers of banking services might have a greater incentive to press state governments for efficient banking markets in higher income states. Also, income may proxy for business cycle effects. We include the federal funds rate to control for the possible influence of the level of market interest rates on banking markets and, thus, regime choice.[12]

Finally, we include variables to test whether the regulatory regime was affected by the political party affiliation of state legislatures or governor. We include dummy variables for the party affiliation of the state governor and whether the same party controlled both houses of a state's legislature. One legislature dummy is set equal to "1" if both houses have a Democratic Party majority, and equal to "0" if not, and the other dummy is set equal to "1" if both houses have Republican majority, and to "0" if not.[13]

3 Data and empirical model

We use data on the 48 contiguous states in our empirical models of the determinants of intrastate branching and interstate banking regime during the 28-year period 1970 to 1997. The Interstate Banking and Branching Efficiency Act of 1994 took full effect in 1997.[14] Under this Act, bank holding companies are permitted to acquire banks in any state, merge banks across state lines and operate the merged banks as branches. Although state restrictions on intrastate branching remained, we end our study in 1997 because the change in federal law governing interstate banking operations introduced a substantially new regime.

Table 1 lists the years in which each state first permitted intrastate branching and interstate banking. For states that adopted intrastate branching or interstate

[12] Kroszner and Strahan (1999) include the average yield on bank loans in the state minus the federal funds rate as an independent variable in one specification to test the hypothesis that pressure for deregulation might be more intense in states where interest rates on bank loans were relatively high. The coefficient on this variable is not significant or large in their model, however, and the data needed to construct it are not available over the entire sample period. We therefore do not include it here.

[13] We set the dummy variables for party control of the state legislature equal to "0" for Nebraska, which has a unicameral legislature. Our choice of variables to capture political influence differs from those specified by Kroszner and Strahan (1999). They specify two variables: i) a dummy set equal to "1" if the same party controls the governor's office and has majorities in both legislative chambers, and ii) the fraction of the three bodies (governorship, house of representatives and senate) controlled by Democrats.

[14] The Act permitted interstate acquisitions by bank holding companies in 1995.

Table 1. Years when states first permitted intrastate branching and interstate banking

State	Intrastate branching (through mergers and acquisitions)	Full interstate banking permitted
Alabama	1981	1987
Alaska	<1970	1982
Arizona	<1970	1986
Arkansas	1994	1989
California	<1970	1987
Colorado	1991	1988
Connecticut	1980	1983
Delaware	<1970	1988
District of Columbia	<1970	1985
Florida	1988	1985
Georgia	1983	1985
Hawaii	1986	**
Idaho	<1970	1985
Illinois	1988	1986
Indiana	1989	1986
Iowa	**	1991
Kansas	1987	1992
Kentucky	1990	1984
Louisiana	1988	1987
Maine	1975	1978
Maryland	<1970	1985
Massachusetts	1984	1983
Michigan	1987	1986
Minnesota	1993	1986
Mississippi	1986	1988
Missouri	1990	1986
Montana	1990	1993
Nebraska	1985	1990
Nevada	<1970	1985
New Hampshire	1987	1987
New Jersey	1977	1986
New Mexico	1991	1989
New York	1976	1982
North Carolina	<1970	1985
North Dakota	1987	1991
Ohio	1979	1985
Oklahoma	1988	1987
Oregon	1985	1986
Pennsylvania	1982	1986
Rhode Island	<1970	1984
South Carolina	<1970	1986
South Dakota	<1970	1988
Tennessee	1985	1985
Texas	1988	1987
Utah	1981	1984
Vermont	1970	1988
Virginia	1978	1985
Washington	1985	1987
West Virginia	1987	1988
Wisconsin	1990	1987
Wyoming	1988	1987

** States not yet deregulated. Source: Kroszner and Strahan (1999).

Table 2. Descriptive statistics ($N = 1{,}344$)

Variable	Mean	Standard error	Minimum	Maximum
Democrats control state legislature	0.556	0.497	0	1
Republicans control state legislature	0.227	0.419	0	1
Democratic governor	0.579	0.494	0	1
Republican governor	0.412	0.492	0	1
Small bank asset share	0.089	0.050	0	0.210
Small firm ratio	0.878	0.017	0.790	0.9233
Small/large bank capital ratio	0.033	0.026	–0.060	0.150
Insurance sector size	0.154	0.215	0.050	0.780
Federal funds rate	7.556	3.089	3.020	16.380
Per capita income	18,807	3,951	9,432	34,097
Bank insurance sales	0.362	0.481	0	1

Variable definitions and data sources: The political control variables were obtained from the *Statistical Abstract of the United States* and *The Book of States*, various years. Small Bank Asset Share is the proportion of total assets in banks with assets of less than the state median. Small/Large Bank Capital Ratio is the aggregate equity/asset ratio of small banks minus the aggregate equity/asset ratio of large banks in a state, where small and large are defined in terms of median bank assets. Data on bank assets and capital ratios were obtained from *Reports of Income and Condition* ("Call Reports"). Small Firm Ratio is the proportion of firms in a state with less than 20 employees. Data on firms were obtained from the Bureau of the Census, *County Business Patterns*, various years. Insurance Sector Size is the ratio of total insurance assets in a state to the sum of insurance and banking assets. Information on the size of the insurance sector was obtained from the U.S. Bureau of Economic Analysis and the Rand Institute. Bank Insurance Sales is an indicator variable set equal to "1" for states that permit banks to sell insurance, and to "0" otherwise. Information on insurance sales was obtained from Conference of State Bank Supervisors, *A Profile of State Chartered Banking*, and individual state banking departments. Federal Funds Rate is the annual average market federal funds interest rate, obtained from the Federal Reserve. Data on state Per Capita Income were obtained from the Department of Commerce, Bureau of Economic Analysis.

banking between 1970 and 1997, our dependent variables are set to "1" in the year of adoption and all subsequent years of our sample period. We report descriptive statistics and data source information for all independent variables in Table 2.

3.1 Empirical model: the spatial probit

The basic model of spatial correlation developed by Cliff and Ord (1981) and Anselin (1988) allows for spatial dependence in the dependent variable (termed a "spatial lag" or "spatial autoregression") or in the error component (termed a "spatial error lag" or "spatial autocorrelation"). The dependent variable and the error terms are correlated across space in a consistent manner. Spatial correlation in cross-sectional data is multi-dimensional in that it depends upon all contiguous or influential units of observation (in this case, states). Just as one corrects for autocorrelation in time series analysis, accurate cross-sectional analysis requires correcting for spatial autocorrelation. Ignoring spatial dependence in the dependent variable can result in biased and inconsistent coefficient estimates, and a failure to control for spatial autocorrelation can result in inefficient coefficient estimates (Anselin 1988).

The framework we adopt is similar to the standard spatial econometric model, although our specification is modified in the spirit of Case (1992) and Marsh et al. (2000) to account for the discrete nature of our dependent variable and the panel structure of the data. Maximum likelihood estimation traditionally produces consistent estimates of spatial models with continuous dependent variables. However, unless corrected for, spatial correlation in probit models introduces heteroskedasticity (Case 1992; Marsh et al. 2000).

The regime status for a state is derived, as in the usual binary choice model, from a latent variable y_{it}^* and the rule $y_{it} = 1$ if $y_{it}^* > 0$ and $y_{it} = 0$ if y_{it}^* ″ 0. Our first-order spatial lag probit model can be expressed as:

$$\mathbf{y}^* = \rho \cdot \mathbf{W} \cdot \mathbf{y}^* + \mathbf{X}\boldsymbol{\beta} + \boldsymbol{\varepsilon} \tag{1}$$

where $\mathbf{X}$ is a $(TN \times K)$ matrix of exogenous variables, and $\boldsymbol{\varepsilon}$ is a $(TN \times 1)$ vector of i.i.d. error terms. $\mathbf{W}$ is a $(TN \times TN)$ block diagonal matrix having $(N \times N)$ spatial weights matrices $\mathbf{w}$ along T block diagonal elements. Individual elements of $\mathbf{w} = \{\omega_{ij}\}$. The scalar ρ is the spatial lag coefficient and reflects positive spatial correlation in the dependent variable if $\rho > 0$, negative spatial correlation if $\rho < 0$, and no spatial correlation if $\rho = 0$.[15] The estimated ρ can be interpreted as follows: For any state i, an increase/decrease in the average of other states' spatially weighted regime choice ($\mathbf{Wy}^*$) results in an increased/decreased probability that state i will deregulate. Performing OLS on (1) will result in biased and inconsistent coefficients because corr[$\mathbf{Wy}^*$, $\boldsymbol{\varepsilon}$] $\neq 0$, and a failure to account for the spatial lag in (1) if $\rho \neq 0$ will bias the elements of $\boldsymbol{\beta}$ (via omitted variable bias). See Anselin (1988, p. 58) for details.

Spatial correlation can also occur in the error term, $\boldsymbol{\varepsilon}$. Spatially correlated errors may occur due to spatial correlation among the independent variables, spatial heterogeneity in functional form, omitted variables, or spatial correlation in the dependent variable when a spatially lagged dependent variable is not included in the model (Anselin 1988, chapter 8). The first-order spatial error lag model is given as:

$$\boldsymbol{\varepsilon} = \lambda \mathbf{W} \boldsymbol{\varepsilon} + \boldsymbol{\upsilon} = (\mathbf{I} - \lambda \mathbf{W})^{-1} \boldsymbol{\upsilon} \tag{2}$$

where ε is the $(TN \times 1)$ vector of error terms, $\boldsymbol{\upsilon}$ is a $(TN \times 1)$ component of the error terms made up of i.i.d. random variables, $\mathbf{W}$ is the $(TN \times TN)$ matrix described earlier, and λ is a scalar that measures spatial error correlation. The errors are positively correlated if $\lambda > 0$, negatively correlated if $\lambda < 0$, and not

[15] Unlike the standard first-order autoregressive model in time series, the spatial correlation coefficients do not necessarily have to lie between −1 and 1 in the first-order spatial autoregressive model. Generally, when a binary weights matrix is used the values for the spatial correlation coefficients are between the inverse of the largest and smallest eigenvalues of the weights matrix. See Anselin (1995).

correlated if $\lambda = 0$. As with autocorrelation in time series, a failure to account for spatial error correlation when $\lambda \neq 0$ will render the parameter estimates inefficient because of the non-diagonal structure of the error covariance matrix (see Anselin 1988, p. 59).

Many alternative weighting schemes for **w** have been used in the literature. Perhaps the most common is the binary joins matrix (Cliff and Ord 1981; Anselin 1988; Case 1992) in which $\omega_{ij} = 1$ if observations i and j ($i \neq j$) share a common border, and $\omega_{ij} = 0$ otherwise. In this specification, the elements of matrix **w** are row-standardised by dividing each ω_{ij} by the sum of each row i. A limitation of the binary joins matrix is that it assumes equal weights across all bordering spatial neighbours and does not allow the effective capture of spatial distances across all cross-sectional units. Thus, we also consider various measures of spatial distance (d) that have been discussed in the literature (Bodson and Peters 1975; Dubin 1988; Garrett and Marsh 2002; Hernandez 2003), including inverse distance where $\omega_{ij} = 1/d_{ij}$, inverse distance squared, and exponential distance decay where $\omega_{ij} = \exp(-d_{ij})$. As the distance between states i and j increases (decreases), ω_{ij} decreases (increases), thus giving less (more) spatial weight to the state pair when $i \neq j$. In all cases, $\omega_{ij} = 0$ for $i = j$. We follow Hernandez (2003) in using the distance between state population centres as our measure of distance.[16]

We found the inverse distance measure to outperform the alternatives based on the maximum likelihood principle and, hence, we report model estimates based on this measure. For comparison, we also report one specification based on the more common binary joins weights matrix. Furthermore, we test whether the influence of spatial dependence varies across the nine Census regions. Regional differences in bank regulation patterns, as well as differences in state land areas, suggest that the coefficients on spatial terms could differ across regions.[17] Allowing for regional spatial correlation coefficients gives the following structure:

$$\mathbf{y}^* = \sum_{k=1}^{R} \rho_k \cdot \mathbf{W_k} \cdot \mathbf{y}^* + \mathbf{X\beta} + \boldsymbol{\varepsilon} \tag{3a}$$

where

$$\boldsymbol{\varepsilon} = \sum_{k=1}^{R} \lambda_k \mathbf{W_k} \boldsymbol{\varepsilon} + \boldsymbol{\upsilon} = \left(\mathbf{I} - \sum_{k=1}^{R} \lambda_k \mathbf{W_k} \right)^{-1} \boldsymbol{\upsilon} \tag{3b}$$

Here R denotes the total number of regions (nine), and ρ_k and λ_k denote the spatial lag and spatial error lag coefficients, respectively, for region k. $\mathbf{W_k}$ remains the $(TN \times TN)$ block diagonal matrix having $(N \times N)$ spatial weights matrix $\mathbf{w_k}$ along T block diagonal elements. Now, however, we construct the elements of each

[16] We use the geographic coordinates for the population centroids computed by the Bureau of the Census for the year 2000. Population centroids did not differ significantly in early decades. They also appear to reasonably approximate most state financial centres.

[17] The regions are: New England, Mid-Atlantic, East North Central, West North Central, South Atlantic, East South Central, West South Central, Mountain, and Pacific.

matrix $\mathbf{w_k}$ to capture spatial correlation between each state in region k and the remaining 47 states.[18] Thus, for each state i in region k, row i of $\mathbf{w_k}$ contains some measure of distance between state i and all remaining 47 states. If state i is not in region k, then row i of $\mathbf{w_k}$ contains all zeros. In essence, we construct each matrix $\mathbf{w_k}$ by pre-multiplying each $\mathbf{w_k}$ by a dummy variable that has a value of "1" if state i is in region k and "0" otherwise.

Rewriting the full spatial autoregressive model and incorporating the structure in (3a) and (3b) gives:

$$\mathbf{y}^* = \left(\mathbf{I} - \sum_{k=1}^{R} \rho_k \mathbf{W_k}\right)^{-1} \mathbf{X}\boldsymbol{\beta} + \left(\mathbf{I} - \sum_{k=1}^{R} \rho_k \mathbf{W_k}\right)^{-1} \left(\mathbf{I} - \sum_{k=1}^{R} \lambda_k \mathbf{W_k}\right)^{-1} \boldsymbol{\upsilon} \tag{4}$$

The above structure induces heteroskedasticity (Case 1992).[19] The covariance matrix is:

$$\mathrm{E}[\boldsymbol{\varepsilon}\boldsymbol{\varepsilon}'] = \sigma_\upsilon^2 \left[\left(\mathbf{I} - \sum_{k=1}^{R} \rho_k \mathbf{W_k}\right)' \left(\mathbf{I} - \sum_{k=1}^{R} \lambda_k \mathbf{W_k}\right)' \times \left(\mathbf{I} - \sum_{k=1}^{R} \lambda_k \mathbf{W_k}\right)\left(\mathbf{I} - \sum_{k=1}^{R} \rho_k \mathbf{W_k}\right)\right]^{-1} \tag{5}$$

where σ_υ^2 is the common variance of the υ_{it}'s and:

$$\boldsymbol{\varepsilon} = \left(\mathbf{I} - \sum_{k=1}^{R} \rho_k \mathbf{W_k}\right)^{-1} \left(\mathbf{I} - \sum_{k=1}^{R} \lambda_k \mathbf{W_k}\right)^{-1} \boldsymbol{\upsilon}$$

We correct for heteroskedasticity using the method of Case (1992) and Marsh et al. (2000). We premultiply the full spatial model in (4) by the variance normalising transformation $\mathbf{Z} = (\mathrm{diag}(\mathrm{E}[\boldsymbol{\varepsilon}\, \boldsymbol{\varepsilon}']))^{-1/2}$. The transformed model is:

$$\mathbf{Z} \cdot \mathbf{y}^* = \mathbf{Z} \cdot \left(\mathbf{I} - \sum_{k=1}^{R} \rho_k \mathbf{W_k}\right)^{-1} \mathbf{X}\boldsymbol{\beta} + \mathbf{Z} \cdot \left(\mathbf{I} - \sum_{k=1}^{R} \rho_k \mathbf{W_k}\right)^{-1} \left(\mathbf{I} - \sum_{k=1}^{R} \lambda_k \mathbf{W_k}\right)^{-1} \boldsymbol{\upsilon} \tag{6}$$

Because $y_{it}^* > 0$ is the same as the event $y_{it} > 0$, we set $y_{it} = 1$ if $y_{it}^* > 0$, indicating that the state has adopted a liberal branching (or interstate banking) regime, or $y_{it} = 0$ if the state does not permit intrastate branching (or interstate banking). The log-likelihood function for the spatial probit model is then expressed as:

[18] Note that this specification allows for asymmetry in spatial correlation between two states each located in a different region. That is, if states i and j are in different regions, then the spatial effect of i on j could be different than the spatial effect of j on i.

[19] Our model makes the assumption that the off-diagonal elements of the covariance matrix are zero. Relaxing this assumption, while potentially increasing efficiency, greatly complicates the estimation procedure. Research has explored several alternative methods for estimating the spatial probit models that use information in the off-diagonal elements (Anselin 2002; Fleming 2004). However, the literature has not established a consistently reliable estimation technique. We also assume that the error structure is not subject to temporal autocorrelation. To our knowledge there is no established framework to correct for temporal autocorrelation in a spatial panel probit model.

$$\ln \mathrm{L} = \sum_{i=1}^{N} \sum_{i=1}^{T} \left\{ y_{it} \ln F\left[X_{it}^{*}\boldsymbol{\beta}\right] + (1 - y_{it}) \ln\left(1 - F\left[X_{it}^{*}\boldsymbol{\beta}\right]\right) \right\} \tag{7}$$

where $\mathbf{X}^* = \mathbf{Z} \cdot (\mathbf{I} - \Sigma\rho_k \mathbf{W_k})\mathbf{X}$. Setting either all $\rho_k = 0$ or all $\lambda_k = 0$ allows estimation of the spatial error lag or spatial lag model, respectively, and setting all ρ_k and λ_k to zero gives the standard probit log-likelihood.

4 Estimation results and discussion

We estimate various specifications of the spatial probit model using both the binary spatial weights matrix and the inverse distance spatial weights matrix described above.[20] We report the results for models of intrastate branching and interstate banking regime choice in Tables 3 and 4, respectively.

For comparison we report estimates of a non-spatial probit model ($\lambda = \rho = 0$) in the first column of each table. We find that including spatial lag and/or spatial error terms significantly enhances the explanatory power of the model and affects the magnitude and significance of the coefficient estimates of some of the independent variables. Based on likelihood ratio tests, we found that the spatial lag model consistently outperformed the spatial error lag model. Hence, we report estimates of the basic spatial lag model, which assumes that the coefficients on the spatial term are equal across all regions, in the second and third columns of each table. We use the binary joins weights matrix in the estimation reported in column 2, and the inverse distance weights matrix in the estimation reported in column 3. The specification reported in column 4 allows the coefficients on the spatial lag term to vary across regions and is estimated using the inverse distance weights matrix. That specification also includes a spatial error term (λ) and generates the best fit of all the models we estimated.

We find strong evidence of spatial dependence and spatial autocorrelation in our models of regime choice. Regardless of which weights matrix we use, the estimate of ρ is statistically significant at $\alpha = 0.01$. As expected, all estimates of ρ are positive, which is consistent with the hypothesis that a state is more likely to deregulate if nearby states have also chosen to deregulate. Except models (2) and (3) in Table 4, the log-likelihood is larger for the inverse distance weights matrix specifications compared to the binary weights matrix specification.[21]

The economic significance of the various estimates of ρ are quite reasonable. From column (2) in Table 3, the probability that a state will permit intrastate

[20] We estimated several other models, but do not present them here for the sake of brevity and clarity in presentation. Several specifications permitting regional differences in the spatial lag coefficients dominated specifications that assumed no such differences, regardless of whether a spatial error term was included or not. Also, we estimated a spatial error model using both the binary and distance weighting matrix. The results from these models will gladly be provided upon request. The log-likelihoods from these alternative models were significantly lower than for the spatial lag models presented in Tables 3 and 4.

[21] This is not surprising because the binary matrix assumes that only contiguous states are influential on a state's regime choice, whereas the inverse distance matrix assumes that all states have some influence, albeit decreasing with distance.

Table 3. Spatial probit results for intrastate branching regime

Variable	(1) No spatial effects $\rho = \lambda = 0$	(2) Binary weights matrix $\lambda = 0$, ρ	(3) Distance weights matrix $\lambda = 0$, ρ	(4) Distance weights matrix λ, regional ρ's
Constant	–1.762** (0.724)	0.397 (0.269)	0.616*** (0.202)	0.563*** (0.135)
Legislature – D	–0.016 (0.027)	–0.029 (0.024)	–0.048* (0.026)	–0.062*** (0.021)
Legislature – R	0.009 (0.032)	–0.046 (0.029)	0.031 (0.030)	–0.005 (0.025)
Governor – D	–0.308** (0.134)	–0.160 (0.226)	–0.411*** (0.138)	–0.282*** (0.101)
Governor – R	–0.358*** (0.135)	–0.196 (0.226)	–0.487*** (0.136)	–0.331*** (0.099)
Small bank asset share	–3.437*** (0.255)	–3.134*** (0.329)	–4.202*** (0.258)	–3.961*** (0.207)
Small firm ratio	2.601*** (0.757)	0.029 (0.149)	0.167 (0.151)	0.086 (0.094)
Small/large bank capital ratio	1.080* (0.624)	–0.689* (0.400)	–1.942*** (0.457)	–1.931*** (0.334)
Insurance sector size	–0.946*** (0.102)	–0.138** (0.062)	–0.210*** (0.060)	–0.165*** (0.046)
Federal funds rate	–0.022*** (0.003)	–0.002 (0.001)	–0.001 (0.0008)	–0.001 (0.001)
Per capita income	0.037*** (0.003)	0.008***(0.002)	0.013*** (0.002)	0.010*** (0.001)
Bank insurance sales	0.028 (0.022)	0.074*** (0.022)	–0.010 (0.022)	–0.022 (0.018)
ρ	———	0.172*** (0.011)	1.817*** (0.096)	———
λ	———	———	———	–2.913*** (0.870)
$\rho_{\text{New England}}$	———	———	———	1.068*** (0.209)
$\rho_{\text{Mid-Atlantic}}$	———	———	———	3.458*** (0.448)
$\rho_{\text{East North Central}}$	———	———	———	2.483*** (0.443)
$\rho_{\text{West North Central}}$	———	———	———	1.010*** (0.181)
$\rho_{\text{South Atlantic}}$	———	———	———	0.812*** (0.159)
$\rho_{\text{East South Central}}$	———	———	———	2.174*** (0.492)
$\rho_{\text{West South Central}}$	———	———	———	2.265*** (0.385)
ρ_{Mountain}	———	———	———	1.906*** (0.310)
ρ_{Pacific}	———	———	———	7.058*** (1.261)
Log likelihood	–558.48	–503.17	–479.56	–434.59

Note: *** denotes significance at 1%, ** at 5%, and * at 10%. Standard errors are in parentheses. Dependent variable has a value of "1" if state *i* in year *t* had deregulated intrastate branching, and a value of "0" otherwise. Sample period is 1970 to 1997. $N = 1{,}344$.

Table 4. Spatial probit results for interstate banking regime

Variable	(1) No spatial effects $\rho = \lambda = 0$	(2) Binary weights matrix $\lambda = 0$, ρ	(3) Distance weights matrix $\lambda = 0$, ρ	(4) Distance weights matrix λ, regional ρ's
Constant	0.987 (0.604)	–0.378 (0.313)	0.012 (0.156)	0.074 (0.127)
Legislature – D	0.032 (0.023)	0.031 (0.030)	–0.004 (0.025)	0.002 (0.014)
Legislature – R	–0.005 (0.026)	–0.025 (0.045)	0.007 (0.032)	–0.004 (0.016)
Governor – D	–0.206* (0.108)	0.522* (0.307)	0.062 (0.144)	0.023 (0.085)
Governor – R	–0.222** (0.109)	0.559* (0.316)	0.062 (0.139)	0.028 (0.084)
Small bank asset share	0.280 (0.221)	–1.929** (0.866)	–1.558*** (0.382)	–0.999*** (0.325)
Small firm ratio	–1.404** (0.633)	–0.136 (0.165)	–0.104 (0.108)	–0.113 (0.098)
Small/large bank capital ratio	1.900*** (0.441)	0.252 (0.265)	–0.465 (0.292)	–0.048 (0.170)
Insurance sector size	–0.688*** (0.079)	0.051 (0.039)	–0.004 (0.034)	0.010 (0.023)
Federal funds rate	–0.045*** (0.003)	–0.001* (0.0007)	–0.001*** (0.0005)	–0.001*** (0.0003)
Per capita income	0.049*** (0.003)	0.005** (0.002)	0.010*** (0.002)	0.005*** (0.001)
Bank insurance sales	0.045**(0.018)	0.047*** (0.016)	–0.051** (0.025)	–0.033** (0.015)
ρ	———	0.074*** (0.011)	0.827*** (0.077)	———
λ	———	———	———	–0.923*** (0.291)
$\rho_{\text{New England}}$	———	———	———	0.316** (0.139)
$\rho_{\text{Mid-Atlantic}}$	———	———	———	0.551* (0.289)
$\rho_{\text{East North Central}}$	———	———	———	1.530** (0.700)
$\rho_{\text{West North Central}}$	———	———	———	0.331* (0.169)
$\rho_{\text{South Atlantic}}$	———	———	———	0.404** (0.204)
$\rho_{\text{East South Central}}$	———	———	———	0.782 (0.500)
$\rho_{\text{West South Central}}$	———	———	———	1.044* (0.551)
ρ_{Mountain}	———	———	———	0.673** (0.299)
ρ_{Pacific}	———	———	———	4.050*** (1.533)
Log likelihood	–393.53	–186.59	–194.00	–160.33

Note: *** denotes significance at 1%, ** at 5%, and * at 10%. Standard errors are in parentheses. Dependent variable has a value of "1" if state *i* in year *t* had deregulated interstate banking, and a value of "0" otherwise. Sample period is 1970 to 1997. $N = 1{,}344$.

branching increases by 8.9% at the mean value of **Wy*** (0.519). From the distance weights matrix specification in column (3) of Table 3, the probability that a state will permit intrastate branching increases by 7.4% at the mean value of **Wy*** (0.041). Considering the estimates for ρ in Table 4, similar computations reveal increases of 3.1% (column 2) and 2.6% (column 3) at the mean values of **Wy***, respectively. Interestingly, the spatial lag effects are larger for the binary matrix, on average, then for the distance weights matrix, suggesting that direct neighbours had the greatest influence on a state's regime choice. Furthermore, the impact of spatial dependence appears to have been larger for the choice of intrastate branching regime than for the interstate banking regime.

When we estimate individual spatial lag coefficients for each Census region, we find that all of the regional coefficients are positive and statistically significant in the intrastate branching model (Column 4, Table 3), and all but one of the coefficients is significant in the interstate banking model (Column 4, Table 4). Although we find that spatial dependence was important throughout the country, we reject the hypothesis that the coefficients are equal in several instances. There are several reasons why the impact of spatial dependence might vary across regions, including differences in the prevalence of regional banking compacts, other aspects of banking market structure, and regional differences in average state size. Tables 5 and 6 contain *p*-values for pair wise equality tests of all ρ_κ for the intrastate branching and interstate banking models, respectively.

4.1 Other determinants of intrastate branching regime

In addition to supporting our hypothesis of spatial dependence in the choice of banking regimes, our estimates reveal several differences in the size and significance of the coefficients on other independent variables between the spatial and non-spatial models.

One difference concerns the influence of a state's small business sector on its choice of branching regime. The non-spatial probit model estimates indicate that a 1 percentage point increase in the small firm ratio increases the probability of adopting a liberal branching regime by 2.6%. The coefficient on the small firm variable is much smaller and not statistically significant when spatial dependence is controlled for, however, regardless of which weights matrix is used. This casts doubt on the hypothesis that pressure from small business interests had an important effect on the choice of state branching regulations.[22]

A second difference between the spatial and non-spatial models concerns the influence of the relative financial strength of small and large banks on the choice of regime. The coefficient estimate on the relative capital ratios of small versus

[22] Our results are not directly comparable to those of Kroszner and Strahan (1999) because of their use of a non-spatial hazard model, differences in our specifications (e.g., we include per capita income as an independent variable and use different political variables), and because our sample period, 1970–1997, differs from theirs. However, we re-estimated our models over their 1970–1992 sample period and obtained results qualitatively similar to our original estimates.

Table 5. Regional spatial correlation coefficient equality – p-values: Intrastate branching

	New England	Mid Atlantic	East north central	West north central	South Atlantic	East south central	West south central	Mountain	Pacific
New England	———	**0.000**	**0.001**	0.817	0.273	**0.028**	**0.004**	**0.013**	**0.000**
Mid Atlantic	**0.000**	———	0.168	**0.000**	**0.000**	**0.073**	**0.060**	**0.006**	**0.004**
East north central	**0.001**	0.168	———	**0.001**	**0.000**	0.629	0.681	0.238	**0.000**
West north central	0.817	**0.000**	**0.001**	———	0.374	**0.018**	**0.001**	**0.007**	**0.000**
South Atlantic	0.273	**0.000**	**0.000**	0.374	———	**0.008**	**0.000**	**0.001**	**0.000**
East south central	**0.028**	**0.073**	0.629	**0.018**	**0.008**	———	0.882	0.625	**0.000**
West south central	**0.004**	**0.060**	0.681	**0.001**	**0.000**	0.882	———	0.436	**0.000**
Mountain	**0.013**	**0.006**	0.238	**0.007**	**0.001**	0.625	0.436	———	**0.000**
Pacific	**0.000**	**0.004**	**0.000**	**0.000**	**0.000**	**0.000**	**0.000**	**0.000**	———

Note: p-values are from joint significance t-tests on regional spatial coefficients from Table 3, column (4). $N = 1{,}344$. Bold values for pairs significantly different at 10% or better.

Table 6. Regional spatial correlation coefficient equality – p-values: Interstate banking

	New England	Mid Atlantic	East north central	West north central	South Atlantic	East south central	West south central	Mountain	Pacific
New England	———	0.221	0.125	0.861	0.356	0.286	**0.096**	**0.024**	**0.004**
Mid Atlantic	0.221	———	0.275	0.212	0.424	0.604	0.269	0.548	**0.008**
East north central	0.125	0.275	———	0.149	0.178	0.462	0.641	0.329	0.103
West north central	0.861	0.212	0.149	———	0.451	0.301	**0.092**	**0.016**	**0.005**
South Atlantic	0.356	0.424	0.178	0.451	———	0.393	0.137	**0.069**	**0.005**
East south central	0.286	0.614	0.462	0.301	0.393	———	0.694	0.810	**0.019**
West south central	**0.096**	0.269	0.641	**0.092**	0.137	0.694	———	0.390	**0.036**
Mountain	**0.024**	0.548	0.329	**0.016**	**0.069**	0.810	0.390	———	**0.016**
Pacific	**0.004**	**0.008**	0.103	**0.005**	**0.006**	**0.019**	**0.036**	**0.016**	———

Note: p-values are from joint significance t-tests on regional spatial coefficients from Table 4, column (4). $N = 1{,}344$. Bold values for pairs significantly different at 10% or better.

large banks is positive in our non-spatial probit model, suggesting that a 1 percentage point increase raises the probability of adopting intrastate branching by slightly more than 1%. However, from model (4) in Table 3, we find that a 1 percentage point increase in the bank capital ratio results in a 2% *decrease* in the probability of adopting intrastate branching. Finally, both our spatial and non-spatial model estimates reveal that a larger share of banking assets in small banks reduced the probability of adopting intrastate branching, although the coefficient estimate is lower for the non-spatial model. States were more willing to protect local banking markets from the competitive effects of intrastate branching when their small banks were relatively strong financially or held relatively large shares of state banking assets. Our results are thus consistent with Kroszner and Strahan (1999), who find evidence that deregulation occurred later when states had relatively large or strong small banks, and with Abrams and Settle (1993) and Kane (1996), who argue that geographic restrictions on banks reflected the relative strength of small, non-branching banks.

Both our spatial and non-spatial models also indicate that the probability of adopting a liberal branching regime was lower, the larger the share of a state's combined banking and insurance assets held by insurance companies. The coefficient on this variable is, however, much smaller in the spatial models. The coefficient on per capita income is also different between the non-spatial and spatial models. Specifically, a $1,000 dollar increase in per capita income increased the probability of adopting intrastate branching by 3.7% in the non-spatial model but by just 1% in the spatial model shown in column (4) of Table 3. Finally, the coefficients on our political variables are broadly similar across our spatial and non-spatial models; though only in the last two specifications do we find evidence that control of a state's legislature by the Democratic Party reduced the probability of adopting intrastate branching.

4.2 Other determinants of interstate banking regime

With regard to the choice of interstate banking regime, we again find that several variables with statistically significant coefficients in the non-spatial probit model are not significant or are much smaller in the spatial lag models.[23] As with intrastate branching, once spatial dependence is accounted for, we find no support for the hypothesis that the size of a state's small business sector affected the choice of interstate banking regime. Other variables that have significant coefficients in the basic probit but insignificant coefficients in the spatial lag probit models include insurance sector size and the difference between small and large bank capital ratios. By contrast, the coefficient on small bank asset share is significant only in the spatial lag model, which supports the hypothesis that the probability of adopting interstate banking was lower the larger the share of state banking assets held by small banks.

[23] Kroszner and Strahan (1999) do not estimate a separate model for the deregulation of interstate banking.

5 Conclusion

Scholars have long noted regional patterns in bank regulation, market structure and performance. Recently, researchers have exploited the differences in bank regulation at the state level to study the effects of banking policies on economic growth (Jayaratne and Strahan 1996; Freeman 2002), and have considered the effects of banking industry consolidation on the cost and availability of credit to small firms (Petersen and Rajan 2002; Avery and Samolyk 2004). Other studies have sought to explain differences in bank regulation across states, particularly with regard to their choice of branching and interstate banking laws (Abrams and Settle 1993; Kroszner and Strahan 1999).

The present article extends this literature by modeling spatial dependence in the choice of intrastate branching and interstate banking regimes. Obvious regional patterns in bank regulation and the formation of regional banking compacts beginning in the 1970s suggest that states' decisions to adopt particular regulatory regimes were influenced by the decisions made by neighbouring states. Our estimation results strongly indicate such dependence. We find that proximity to states that had liberal branching or interstate banking laws increased the probability that a given state would also adopt liberal laws. We find, however, significant quantitative differences in the impact of spatial dependence across regions.

Our study also provides new evidence on the importance of the political, interest group, and public benefit explanations of banking regulation. We find strong support for the hypothesis that the probability of permitting either interstate banking or intrastate branching was lower the more that a state's banking assets were held by small banks, or the stronger a state's small banks were financially relative to large banks. Our results are thus consistent with prior research that finds a strong association between the relative dominance of small banks within a state and the state's choices of branching and interstate banking regimes. Furthermore, we find that the larger a state's insurance industry was relative to its banking industry, the lower the probability that the state would adopt liberal branching or interstate banking regulations. However, contrary to previous work, we find no evidence that the size of a state's small business sector influenced bank regulation once we control for spatial effects. Similarly, controlling for spatial effects greatly reduces the estimated impacts of state per capita income and of whether banks are permitted to sell insurance.

Although state branching and interstate banking regulations have now largely been supplanted by changes in federal law, states continue to set a variety of banking regulations, such as limits on market share. State governments also remain heavily involved in regulating insurance and other financial services, and engage actively in various economic development policies. The importance of spatial effects on the choice of interstate banking and intrastate branching regime from 1970 to 1997 suggests that such effects should be considered when investigating the determinants of other state economic policies.

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