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BANKS VS. CREDIT UNIONS: DYNAMIC COMPETITION IN LOCAL MARKETS

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ABSTRACT

One interesting aspect of the financial services industry is that for-profit institutions such as commercial banks compete directly with not-for-profit financial intermediaries such as credit unions. In this article, we analyze competition among banks and between banks and credit unions using a dynamic model of spatial competition. The model allows for the co-existence of (for-profit) banks and (not-for-profit) credit unions. Using annual county-level data on banking market concentration and credit-union participation rates for the period 1989-96, we find empirical evidence of two-way competitive interactions between banks and credit unions.

Keywords: Bank; Concentration; Credit Unions; Not-for-profit institutions
JEL classifications: G21, L11, L31

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‡ The views expressed in this paper are those of the authors and do not necessarily reflect official positions of Federal Reserve Bank of St. Louis or the Federal Reserve System.

BANKS VS. CREDIT UNIONS: DYNAMIC COMPETITION IN LOCAL MARKETS

Competition between banks and credit unions is interesting in part because it entails a for-profit sector (commercial banks) competing with a not-for-profit sector (credit unions). In addition, the distribution of market shares across financial intermediaries with different clienteles, service offerings, and cost structures may have implications for efficiency. This article shows that, although this competition is potentially unstable in the sense that it could lead to a complete crowding-out of one sector or the other, this has not occurred in the United States. We develop a dynamic model of competition that includes both types of financial intermediary and we analyze the model’s stability properties. Then, using annual county-level observations of the concentration of commercial-bank deposit market shares and household participation in occupational credit unions for the period 1989-96, we find empirical support for the model's predictions about competitive interactions between the two sectors. In particular, we find evidence that banks and credit unions directly affect each other’s performance in local deposit markets that are dominated by commercial banks and thrifts.

Previous research (Emmons and Schmid, 1999a) finds a link between the concentration of local commercial-bank deposit shares (as measured by the Herfindahl index [1]) and indicators of credit-union efficiency (wage expense) and risk-taking (loan-loss allowances). In addition, higher bank-deposit concentration is associated with higher household participation rates (the fraction of those who are eligible who choose to join) at credit unions (Emmons and Schmid, 1999b). Thus, competitive conditions among banks appear to influence the behavior of credit unions and their (potential) members.

The purpose of this article is to investigate the two-way interaction between banks and credit unions in local markets. We develop and analyze a dynamic model of spatial competition between banks and credit unions. We derive and empirically estimate dynamic response functions of the local banking sector and of households (who comprise the credit unions).

The article is organized as follows: The first section briefly discusses our approach in light of the existing literature on banking market structure and competition. The second section develops a model of spatial competition between credit unions and banks. The model allows for two countervailing influences on occupational credit-union participation rates (previously discussed and documented by Emmons and Schmid, 1999b). These are, first, economies of scale as the credit union's operations grow, and second, decreasing within-group affinity (i.e., strength of the "common bond") as the membership increases. The banking sector is modeled as a homogenous Cournot oligopoly. The third section describes the dataset and the econometric methods we employ. The fourth section presents our empirical results, and the fifth section draws conclusions. Two appendixes appear at the end of the article with details on the econometric methodology and on the construction of the dataset, respectively. The dynamics of the model are demonstrated in Appendix 3.

CREDIT UNIONS AND COMPETITIVE ANALYSIS OF LOCAL BANKING MARKETS

There is a large and growing literature investigating competition among banks (for example, see Berger, Demsetz, and Strahan, 1999; Amel and Hannan, 1998; or Prager and Hannan, 1998). Much less research focuses on the interactions between credit unions (or thrift institutions) and banks (Hannan and Liang, 1995). This section briefly discusses the bank-competition literature and asks whether credit unions are important for banking competition.

Banking Market Concentration, Prices, and Profits

Mergers among depository institutions and steady expansion of credit unions have been two hallmarks of the U.S. financial landscape in the 1980s and 1990s, but simply acknowledging these trends is not sufficient to characterize the evolution of competitive conditions in local deposit markets. The number of commercial-bank charters in existence has declined by between three and five percent annually since 1988, resulting in a nine-year (1988-97) cumulative disappearance of 33 percent of all bank charters (Berger, Demsetz, and Strahan, 1999, Tables 1 and 2). Mergers accounted for about 84 percent of disappearances and failures for 16 percent. Deposit-market concentration, however, actually declined slightly over this period. Average commercial-bank deposit Herfindahl indexes in metropolitan statistical areas fell from 0.2020 to 0.1949, and those in non-metropolitan counties fell from 0.4317 to 0.4114 (Berger, Demsetz, and Strahan, 1999, Table 1). Meanwhile, credit-union membership grew more than 38 percent in the decade ending in 1996, while the country's population grew about 10 percent (U.S. Treasury, 1997, p. 24).

Does market concentration matter for prices and profits? In a non-banking context, Tirole notes that, "Most cross-sectional analyses find a weak but statistically significant link between concentration and profitability (1988, p. 222)." With regard to banking markets, Gilbert (1984) concluded in an early review of the empirical literature that the economic significance of market concentration by banks before deregulation was very difficult to assess, not least because of the poor quality of much of the empirical research. More recently, Shaffer (1992) summarized the (lack of) current consensus by stating that the degree to which banking market structure matters for competition and performance is "a hotly debated topic."

The predominant empirical approach to banking competition in the last several decades has been the so-called SCP (Structure-Conduct-Performance) paradigm. Also, the SCP approach remains the dominant approach in the regulatory analysis of antitrust issues in banking (Kwast, Starr-McCluer, and Wolken, 1997). This approach presumes that measures of banking market structure, including measures of market concentration, are good indicators of the intensity of competition that occurs ("conduct") (Scherer and Ross, 1990, pp. 4-7). The intensity of competition influences the price for financial services, which are, in turn, assumed to determine firms' profits ("performance"). In a nutshell, the higher the concentration in the local banking market, the higher are the prices for financial services, and consequently the higher are banks' profits. The policy implication is that higher market concentration is associated with lower social welfare and, therefore, higher concentration is undesirable. However, the sources of differential levels of market concentration are left unexplained.

Numerous banking studies demonstrate statistical relationships that are consistent with some aspects of the SCP paradigm, at least in a static context (Berger and Hannan, 1989; Neumark and Sharpe, 1992; Hannan, 1997; Prager and Hannan, 1998; Amel and Hannan, 1998). On the other hand, some evidence is inconsistent with the predictions of this framework (Shaffer, 1989). Moreover, some of the evidence that is consistent with the predictions of the SCP paradigm is subject to a different interpretation. For example, the link between market structure and profitability may be spurious in the sense that an important variable is omitted. Firms of all kinds that are more efficient may have larger market shares simply because their costs are lower (Demsetz, 1973; Berger, 1995). This is a desirable outcome, not one that reduces social welfare.

It is not the purpose of this paper to test the SCP paradigm. We do not observe financial-services prices directly (corresponding to the "conduct" element), for example, and our measure of performance is not bank profits. The SCP framework is related to our theoretical model, however, so it is useful to outline what our model has in common with the SCP paradigm as well as to clarify those respects in which they differ.

The SCP paradigm is predominantly an empirical approach and has been applied most often without explicit reference to a theoretical model of competition. Consequently, the link between higher prices and higher profitability is usually viewed as an empirical question. It is this link between prices and profitability, however, that is questionable on theoretical grounds. Even if higher concentration leads to higher prices for financial services (and consequently to a decline in demand), this does not necessarily translate into higher bank profits. With free entry and exit, the zero- (i.e., "normal") profit condition may hold even in a world of oligopolistic competition. If oligopolistic competitors operate with fixed costs—a reasonable assumption in most cases—then higher concentration may co-exist with higher prices without violating the zero-profit condition. In such a case, the higher price would be caused by higher average costs due to lower market output.[2] Thus, while users of the SCP paradigm typically draw the conclusion that higher prices (caused by higher concentration) lead to higher profitability, this conclusion is not warranted on purely theoretical grounds. It remains an empirical question whether higher concentration actually leads to higher profitability—that is, whether barriers to entry or exit matter.

Credit Unions in the Analysis of Banking Competition: Do They Belong?

The primary focus of bank antitrust enforcement in the merger-review process carried out by the Antitrust Division of the Department of Justice and by federal bank regulators (the Federal Reserve, the Office of the Comptroller of the Currency, and the Federal Deposit Insurance Corporation) is "the availability of banking services, including loans and credit, to small and medium-sized businesses" (Nanni, 1998, p. 193). This focus on the availability of small-business credit means that credit unions, which serve the household sector almost exclusively, are routinely ignored for purposes of regulatory analysis of banking market competition. Compounding the problem of defining and measuring the relevant market, the only comprehensive local-market data available are annual observations of commercial-bank and thrift deposits. Deposit data do not differentiate between commercial and retail ownership (i.e., business and household deposits), so comparisons of depository institutions with different mixes of deposit customers are necessarily imprecise.[3]

Direct empirical evidence that credit unions matter in banking markets has been difficult to find. Amel and Hannan (1998) conclude on the basis of empirically estimated residual deposit supply elasticities that commercial banks in a local market continue to be their own most relevant competitors. That is, it does not appear that non-local or non-bank financial institutions such as out-of-market banks, credit unions, other thrift institutions, finance companies, or providers of money-market mutual funds are important determinants of deposit rate-setting behavior by banks.

There is at least some indirect evidence that credit unions matter to banks, however. Bankers themselves frequently complain about (possibly unfair) competition from credit unions.[4] Banks have collectively spent large sums of money lobbying Congress to inhibit credit-union expansion. Thus, there is at least a reasonable presumption that banks view credit unions as competitors—actual or potential—in at least some of their market segments, such as the market for household deposits.

A MODEL OF COMPETITION BETWEEN BANKS AND CREDIT UNIONS

In this section we present a model of local market competition between banks and credit unions. Credit unions operate as not-for-profit institutions in a spatial setting. Banks, which are for-profit firms, interact strategically with each other, forming a homogenous oligopoly. All households are potential members of occupational credit unions.[5] Each household may join its credit union or do business with a bank, but not both.

In our model, a decrease (increase) in the price of bank services that follows decreased (increased) bank concentration is caused by a rising (falling) market share in the presence of economies of scale and zero (i.e., normal) profits. The zero-profit condition distinguishes our model from the SCP paradigm. In contrast to the SCP paradigm, our model does not assume that higher concentration necessarily results in abnormally high profits even though prices may increase.

The Model

We examine a circular-city economy similar to the one first described by Salop (1979) and subsequently analyzed by many other researchers interested in markets with diverse preferences of consumers. The perimeter of a circle of length L (representing the city) is covered by a continuum of households. All households are identical except for their location on the circle. We interpret the location, or address, of each household as the reading on an index that describes its unique preferences for financial services. A given household's most-preferred bundle of financial services would be provided precisely at the household's location. Financial services provided anywhere else on the circle are inferior. Following Salop, we will say that a household's consumption of financial services at any location other than its own address involves costly "traveling."

Providers of financial services are of two types, both of which also lie along the circle. We assume that credit unions exist only at discrete points along the circle while bank branches are ubiquitous. Metaphorically, one could think of banks as providing electronic banking to every home, while credit unions only offer physical branches at selected locations. Since banks are able to deliver their bundles of financial services to every household's living room, households do not need to travel when doing business with banks. Households must travel if they buy from credit unions, however. This represents the fact that the credit union's financial services may not suit any given household's preferences exactly. Thus, the travel cost is meant to reflect the restricted nature of financial services credit unions typically offer.

Credit Unions. All credit unions are identical except for their location. For simplicity, we assume that all credit unions have a single-common-bond charter.[6] A credit union's costs of production are

(1) $$C^{CU}(mx^{CU}) = f + vmx^{CU}$$

where f is a fixed cost of operation (net of any subsidy from the employer), v is the variable cost of providing one unit of financial services, x^{CU} is the number of bundles of financial services this credit unions sells per customer (i.e., member), and m is the number of members. Thus, credit unions exhibit strictly decreasing average costs with respect to the amount of financial services they provide, consistent with empirical evidence on credit unions (Emmons and Schmid, 1999b). This feature of credit unions, along with their policy of setting price equal to average cost, renders the nature of competition between banks and credit unions particularly interesting.[7]

FIGURE 1 HERE

Households. The households located on the circle are split into N contiguous segments, each of length L / N (see Figure 1). Each segment i $(i = 1,\ldots,N)$ comprises the households of employer i. If employer i $(i = 1,\ldots,N)$ sponsors a credit union (credit union i), it will be located in the center of the arc that forms household segment i, as we argue below. Employment defines the "common bond." That is, all households located in segment i are eligible to join credit union i but not any other credit union.

Each household has a linear price-elastic demand for (bundles of) financial services, which may be purchased either from a bank or from a credit union. Households face marginal costs of $t \times r$ per unit of distance r when traveling to the credit union, where t is a travel-cost parameter. A household located r_j away from credit union i, $i = 1,\ldots,N$, incurs total travel costs to the credit union of $r_j^2 t / 2$. Since travel costs are quadratic in the distance to the credit union, the management team that locates the credit union in the center of the household segment will win an election contest over any other competing management team that suggests locating the credit union somewhere else.[8]

A household's demand for financial services reads

$$(2) \qquad x^j(p^j) = \frac{\alpha}{\beta} - \frac{1}{\beta} p^j$$

where $j = CU$ (Credit Union) if the household is a member of the credit union (and thus buys there) and $j = CB$ (Commercial Bank) otherwise (i.e., when the households does business with a bank). The variable p^j is the price the household expects to pay for a unit of financial services with the credit union (p^{CU}) and the bank (p^{CB}), respectively. The parameters $\alpha > 0$ and $\beta > 0$ are fixed demand parameters, with $\alpha > p^j$ for positive amounts x^j ($j = CU, CB$).[9]

Each household maximizes its net benefit, B, from buying financial services. The net benefit is its consumer surplus minus travel costs, if any. When doing business with the credit union, household *j*'s net benefit amounts to

$$(3) \qquad B^{CU} = \frac{\beta\left(x^{CU}\right)^2}{2} - \frac{r_j^2 t}{2} \equiv \frac{(\alpha - p^{CU})^2}{2\beta} - \frac{r_j^2 t}{2} \ .$$

The net benefit from doing business with a bank is

$$(4) \qquad B^{CB} = \frac{(\alpha - p^{CB})^2}{2\beta} \ .$$

FIGURE 2 HERE

We assume that when the households make their credit-union membership decisions, they take the price of the banks' financial services, p^{CB}, as given. The membership of credit union i, $i = 1,\dots,N$, will comprise all households *j* within the potential membership for which the following inequality holds (see Figure 2):

$$(5) \qquad B^{CU} - B^{CB} \equiv \frac{\beta\left(x^{CU}\right)^2}{2} - \frac{r_j^2 t}{2} - \frac{(\alpha - p^{CB})^2}{2\beta} \geq 0 \ .$$

Credit-Union Equilibrium. As Figure 2 illustrates, not every potential member joins the credit union. Membership in a credit union is worthwhile only if the price of financial services it offers is sufficiently lower than those of a bank to offset the travel costs households face to do business with a credit union. A household relatively far from the credit union (at a distance greater than r^*) buys financial services from a commercial bank. The marginal households, those located at the distance r^* on either side of the credit union, are indifferent between joining their credit union and doing business with a bank:

(6) $$B^{CU} - B^{CB} \equiv \frac{\beta\left(x^{CU}\right)^2}{2} - \frac{\left(m^*\right)^2 t}{8} - \frac{(\alpha - p^{CB})^2}{2\beta} = 0$$

where $m^* \equiv 2r^*$ is the number of members credit union *i* attracts.

For $m^* \equiv 2r^* \geq L / N$, a corner solution obtains. All households will join their (occupational) credit union, and no banks will exist. A necessary condition for an interior solution, i.e., $0 < m^* < L / N$, is that the benefit from being with the credit union, B^{CU}, is positive for intra-marginal households *j*, $r_j < r^*$, and negative for extra-marginal households, $r_j > r^*$. Thus a necessary condition for an interior equilibrium is that the benefit from credit-union membership, B^{CU}, decreases for the marginal households with an increase in the number of members, *m*. On the one hand, the higher *m*, the higher are the travel costs for the marginal households. The higher *m*, the lower the price for a unit bundle of financial services offered by the credit union, p^{CU}. The price decrease is due to two factors: new demand from new members and induced demand from incumbent members. Induced demand is caused by the decrease in average costs (and thus price) that the new demand brings about. We assume that an interior equilibrium obtains, i.e., that the benefit B^{CU} of credit-union membership decreases for the marginal households as their distance to the credit union, $m / 2$, increases.

Another corner solution obtains for $m^* \leq 0$. In this case the average cost of the credit union, i.e., its price for a unit bundle of financial services, is too high to attract members. Thus, the employer in question will not sponsor a credit union.

The price of credit unions' financial services is determined by average costs, which is obtained from (1) by dividing through by mx^{CU}, the amount of services provided:

(7) $$p^{CU} = \frac{C^{CU}(mx^{CU})}{mx^{CU}} = \frac{f}{mx^{CU}} + v \ .$$

Price and quantity of the financial services sold by credit union j, $j = 1,\ldots,N$, are related via the demand aggregated over its members:

$$(8) \qquad m^* x^{CU}(p^{CU}) = m^* \times \left(\frac{\alpha}{\beta} - \frac{1}{\beta} p^{CU} \right).$$

For any given price charged by the banks, p^{CB}, equations (6), (7), and (8) allow us to solve for the equilibrium values of the price, p^{CU*}, the per-member quantity, x^{CU*}, and the number of credit-union members, m^*.

We define the participation rate, $part$, as the fraction of the household continuum on the circle that joins a credit union:

$$(9) \qquad part \equiv \frac{mN}{L}$$

where L/N is the length of each household continuum j, $j = 1,\ldots,N$.

The following Lemma describes the relationship between the credit-union participation rate, $part$, and the price at which banks offer their services, p^{CB}.

Lemma 1: An increase in the price for a unit bundle of bank services leads to higher credit-union participation.

Proof: Starting with equation (6), assume that p^{CB} increases marginally. Joining the credit union will now become worthwhile for previously extra-marginal households. This increases the equilibrium number of members, m^*. The new demand decreases average costs and thus induces additional demand from previously intra-marginal members, increasing the per-member quantity, x^{CU}. Under the previously stated assumption that the benefit from credit-union membership declines for the marginal households with an increase in *m*, the adjustment process will stop at an interior solution (i.e., before a corner solution at a participation rate of $m = L/N$ is reached), bringing equation (6) back into equilibrium.

Banks. There are K identical banks. These banks face the residual demand for financial services, i.e., the demand that is not served by credit unions. Total demand amounts to $L \times x(p)$, while the demand served by credit unions equals $Nm \times x(p^{CU})$. Thus the following (indirect) demand function defines the residual market:

(10) $$p^{CB}(x^{CB}) = \alpha - \tilde{\beta}\, x^{CB}$$

where $\tilde{\beta} \equiv \beta / (L - Nm)$.

Because banks are ubiquitous, their market is not segmented due to travel costs. This means that each bank faces the total (residual) market. Also, since banks are homogenous, there is a uniform price for bank services.

We assume that each bank follows a Cournot conjecture (i.e., each bank *i* assumes that any variation of its own output will not affect the output chosen by any other bank $j,\ i \neq j;\ i, j = 1,\ldots,K$):

(11) $$\frac{dx_i^{CB}}{dx_j^{CB}} = \begin{cases} 1 & \text{for } i = j \\ 0 & \text{otherwise} \end{cases} .$$

Similarly, we assume that each bank takes the residual market as given, i.e., each bank assumes that any variation of its own output leaves the credit-union participation rate (and thus *m*), unchanged:

(12) $$\frac{dm}{dx_j^{CB}} = 0 \ , \quad j = 1,\ldots,K \ .$$

The profit of bank $j,\ \ j = 1,\ldots,K$ is:

(13) $$\pi_j(x_j^{CB}, x_{i\neq j}^{CB}) = p^{CB}(x^{CB})x_j^{CB} - wx_j^{CB} - e$$

where x_j^{CB} is the output of bank *j*, and $x_{i\neq j}^{CB}$ is a vector of bank *j*'s expectations about the output levels of its competitors. The parameters *w* and *e*, $0 < w < \alpha$ and $e > 0$, are the marginal and the fixed costs of production of financial services, respectively.

Bank j, $j = 1,...,K$, maximizes its profits by choosing the optimal amount of output, x_j^{CB}, subject to assumptions (11) and (12), which implies a given value for $L - Nm$. The bank's first-order condition reads:

$$(14)\qquad 0 = \frac{d\pi_j\left(x_j^{CB}, x_{i \neq j}^{CB}\right)}{dx_j^{CB}} = -\tilde{\beta}\; x_j^{CB} + p^{CB} - w = 0\ .$$

Thus, the supply of financial services by bank j, $j = 1,...,K$, equals

$$(15)\qquad x_j^{CB} = \frac{(p^{CB} - w)}{\tilde{\beta}}\ .$$

Aggregated over all banks j, $j = 1,...,K$, the total supply of bank services becomes

$$(16)\qquad x^{CB} = K\frac{(p^{CB} - w)}{\tilde{\beta}}\ .$$

In equilibrium, demand equals supply. Substituting the (indirect) demand function, (10), into supply, (16), we obtain

$$(17)\qquad x^{CB} = \frac{K}{K+1}\frac{(\alpha - w)}{\tilde{\beta}}\ .$$

We now impose the banks' zero-profit condition, $\pi_j \equiv x_j^{CB} \times (p^{CB} - w) - e = 0$, $j = 1,...,K$, which follows from the assumption that the costs of entry and exit are zero. Solving equation (15) for p^{CB} and inserting it into the zero-profit condition yields

$$(18)\qquad \left(\frac{x^{CB}}{K}\right)^2 \tilde{\beta} - e = 0.$$

Only in special cases, the number of banks, K^*, that fulfills equation (18) (and thus makes the zero-profit condition binding) is an integer. Because banks exit the market if they make losses, their profits must be nonnegative. Consequently, the equilibrium number of banks in the market, K_{int}^*, equals

(19) $$K^*_{int} = floor\left(x^{CB}\sqrt{\frac{\tilde{\beta}}{e}}\right)$$

where *floor* is an operator that rounds down to the nearest integer. The equilibrium number of banks, K^*_{int}, is an increasing step function in K^*.

By assumption, all banks are the same size. Therefore, the Herfindahl index, $herf$, (calculated as the sum of the squared market shares of all banks) equals $1/K^*_{int}$.

We can now state Proposition 1, which establishes a relationship between the degree of concentration in the local banking market, $herf = 1/K^*_{int}$, and the credit-union participation rate, $part$.

Proposition 1: An increase in credit-union participation leads to higher banking-market concentration.

Proof: Inserting equation (17) into equation (18), we have

(20) $$\frac{(\alpha - w)^2}{e} = \tilde{\beta}\ (1+K)^2 \equiv \frac{\beta}{L - Nm}(1+K)^2 \ .$$

By starting with equation (20), assume that there is an increase in the participation rate, $part$, which corresponds to an increase in *m*. This decreases $L - Nm$, which requires a decline in *K*, to bring equation (20) back into equilibrium. For a sufficiently big decrease in *K*, the number of competitors, K^*_{int}, will drop to the next lower integer.

Lemma 2 relates the price of a unit bundle of financial services offered by banks, p^{CB}, to the degree of concentration in the local banking market, $herf = 1/K^*_{int}$.

Lemma 2: Higher banking-market concentration causes a higher price for a unit bundle of bank services.

Proof: Solve equation (20) for $\tilde{\beta}$ and insert the resulting expression into equations (10) and (17). Subsequently inserting equation (17) into equation (10) leads to:

(21) $$p^{CB} = \alpha - (\alpha - w)\frac{K}{K+1} \;.$$

With $w < \alpha$ (as assumed), this shows that the price for bank services, p^{CB}, increases with a decrease in the number of competitors, K^{*}_{int}, which proves the Proposition.

Finally, we are able to link the credit-union participation rate, $part$, to the degree of banking-market concentration, $herf = 1 / K^{*}_{int}$.

Proposition 2: Higher banking-market concentration leads to higher credit-union participation.

Proof: Follows from Lemmas 1 and 2.

Thus, our model implies that the interaction between credit unions and commercial banks operates in both directions. An exogenous increase in credit-union participation rates causes bank concentration to increase (Proposition 1). In addition, an exogenous increase in bank concentration causes credit-union participation rates to increase (Proposition 2).

The Dynamic Model. In this section, we cast the model's hypotheses as a pair of response functions. These response functions allow for a one-period response lag by both households (who comprise credit unions) and commercial banks. The first response function characterizes the behavior of households (HH), who may join and participate in the governance of occupational credit unions. The second response function characterizes the behavior of the commercial-banking sector, members of which are governed by shareholders:

(22a) $$part_t = r^{HH}(\, herf_{t-1}\,)$$

(22b) $$herf_t = r^{CB}(\, part_{t-1}\,)$$

where r^{HH} is the response function of households and r^{CB} is the response function of commercial banks, and *part* and *herf* are the household participation rate (number joining divided by number eligible) in occupational credit unions and the commercial-bank Herfindahl index, respectively.

Equation (22a) is a dynamic extension of Proposition 2. It says that the level of household participation in credit unions reacts with a one-period lag to the degree of concentration in the banking sector. As described above, the link between higher (lower) concentration and higher (lower) credit-union participation is a higher (lower) price for financial services offered by banks. Equation (22b) extends Proposition 1. It states that the concentration of the banking sector responds with a one-period lag to the level of household participation in credit unions.

FIGURE 3 HERE

Figure 3 plots the two response functions. Their slopes are positive, as shown in Propositions 2 and 1. If the intersection of the two response functions occurs in the plane spanned by the set {0<*herf*<1, 0<*part*<1}, an interior equilibrium obtains. Otherwise, a corner solution applies with either $part = 0$ (i.e., there is no credit-union sector) or $part = 1$ (i.e., there is no commercial-banking sector).

FIGURE 4 HERE

Solution to the Dynamic Model and Stability Conditions. We now solve the dynamic model (22) and check for stability of the (interior) equilibrium. Appendix 3 provides the details, while Figures 3 and 4 present a graphical illustration of the stability problem. As shown in the appendix, an interior equilibrium is stable if (and only if) the Herfindahl-index response function, r^{CB}, is steeper than the participation-rate response function, r^{HH}. Figure 3 displays a stable equilibrium. If the local financial-services industry finds itself off the equilibrium value, M, say, at points A or A', an adjustment process back toward the equilibrium M will occur (see arrows). If, however, the participation-rate response function, r^{HH}, is steeper than the Herfindahl-index response function, r^{CB}, then the parties' mutual responses will lead the local financial services industry even further away from the equilibrium. Either the banking sector will completely crowd out the credit unions, leading to a participation rate , or the credit unions will drive the banks out of the market, which leads to a participation rate *part = 1.*

DATA AND EMPIRICAL METHODS

This section describes our data briefly and outlines the empirical methods we employ. Further details on both topics are given in separate appendixes at the end of this paper.

TABLE 1 HERE

Data

The observations in our empirical analysis are annual credit-union participation rates and bank Herfindahl-index levels in individual counties (and independent cities). In other words, we aggregate information about credit unions and banks to the county level. We use these aggregated observations to estimate the response functions specified in equations (22).

To construct our county-level variables, we begin with all federally chartered and federally insured occupational credit unions in the period 1989-96. Commuity, associational, and corporate credit unions also exist, but occupational credit unions are by far the dominant type of retail credit union, holding almost 85 percent of member deposits (U.S. Treasury, 1997, p. 19). Table 1 shows the Type of Membership (TOM) codes that apply to the occupational credit unions in our dataset. These include credit unions with single and with multiple common bonds. We do not examine community, associational, or corporate credit unions because they are less numerous than occupational credit unions and because their members and purposes may differ substantially from those of occupational credit unions. Including these diverse types of credit unions would make interpretation of our results more difficult.

TABLE 2 HERE

Table 2 provides a breakdown of our credit-union dataset in each year according to the type of membership (TOM) group. The most common type (single and multiple-common bonds) is manufacturing, followed by government authorities. The number of credit unions covered remains fairly constant at just over 5,000 in each year between 1989 and 1996, with a pooled total of 41,329 before screens are applied on the county level that reduce this number slightly.

TABLE 3 HERE

Table 3 summarizes the bank-concentration measure we use, the Herfindahl index of bank deposit shares calculated at county level, for each of the years in our sample. The median and mean values of the annual county Herfindahl index levels decline over our sample period, which is consistent with the index values reported in Berger, Demsetz, and Strahan (1999, Table 1). The number of counties (and independent cities) for which values of the Herfindahl index are calculated varies slightly from year to year because we discard counties that do not have any banks or credit unions in a given year.

TABLE 4 HERE

Table 4 gives an overview of the credit-union participation rates at county level in our sample in each of the years we examine. As noted above, we discarded counties in which the participation rate was zero (i.e., there were no credit unions). We also discarded counties with participation rates equal to one for econometric reasons, as discussed in Appendix 2. The range of participation rates in the remaining counties is large, with a low of around 1 percent and a high of nearly 100 percent. The median and mean participation rates show a tendency to increase until they peak in the year 1994 at around 59 percent. They then decline slightly, ending at values in the year 1996 that are still notably higher than in the year 1989.

We use control variables in the regression to account for exogenous differences in the competitive environment in cross section and over time. We control for county-level differences in urbanization patterns by using each county's population density in each year as an exogenous regressor. We also use state indicator variables in our regressions to proxy for cross-sectional differences in state law and regulation, and in economic activity. We use year indicator variables to control for changes in economic activity and legislation at the national level over time. Further details are provided in Appendix 2.

Empirical Methods

We estimate an econometric model that pools annual observations from 1,062 counties (and independent cities) for the years 1989 through 1996. The dataset is unbalanced, which means that the number of counties included is not the same in each year. The empirical model is a system of two "seemingly unrelated" equations, which are the response functions we derived in the theoretical model. These are, first, the participation-rate response function and, second, the Herfindahl-index response function. Both variables are bounded on the (0,1) interval. We applied a logit transformation to ensure that these variables are unbounded, making a normal distribution of errors possible. Details on the econometric approach are provided in Appendix 1.

EMPIRICAL RESULTS

Our empirical results are contained in Tables 5, 6, and 7 and are illustrated in Figure 7. Our two main sets of findings can be summarized as follows. First, we find empirical support for the predictions of our dynamic model of competition between banks and credit unions. In particular, we find a statistically significant positive relationships between the current credit-union participation rate and the lagged value of the commercial-bank sector Herfindahl index in a particular county. In addition, we document a positive relationship between the current value of the banking-sector Herfindahl index and the lagged value of the credit-union participation rate in a given county. Second, we detect shifts in both the credit-union participation rate response function (upward) and the commercial-bank Herfindahl index response function (leftward) over our sample period (1989-96).

Model-Specification Tests

Before presenting and discussing the main results of interest, this section reports values for goodness-of-fit measures and results of model-specification tests.

Explanatory Power of the Model. We calculated two sets of measures of fit for the empirical model. The first measures were an R^2 and an adjusted R^2 for the system of equations, which were 0.150 and 0.143, respectively. The second type of goodness-of-fit measures was a pseudo R^2 and an adjusted pseudo R^2 for each equation, both of which are reported in Tables 5 and 6. The values for the pseudo R^2 (and the adjusted pseudo R^2) for the two equations were relatively close to each other which indicates that the model provides about the same amount of explanatory power for each equation.[10]

TABLES 5&6 HERE

We report three Wald tests in Tables 5 and 6. They compare the full model versus alternative specifications that exclude one set of variables at a time: all nonconstant regressors; the set of state indicator variables; and the set of year indicator variables. All Wald test values are statistically significant.

Test for Stability. The dynamic model of competition between banks and credit unions that we derived above rests on two response functions that relate current-period behavior to lagged indicators of the other sector's behavior. Appendix 3 derived the stability conditions for this model, which must hold if the model is to result in a stable equilibrium. As described in Appendix 1, we used the delta method to test the equilibrium condition for our second-order difference equation. The value of this test under the hypothesis that the two response functions have the same slope was -9.59, which is statistically different from zero at the one-percent level.

Participation Rate Response Function. Table 5 presents the participation-rate response function results. As predicted, for given values of the control variables, credit-union participation rates are significantly higher in local banking markets that have more concentrated banking sectors in the previous year. Taking into account the logit transformation that was performed on both variables, it turns out that an increase in the commercial-bank Herfindahl index of 0.01 increases the credit-union participation rate in a county by 0.0100.

The results for the control variables used in the estimation are also interesting. Table 5 shows that participation rates are higher in counties with higher population densities. This indicates that urban and suburban areas have proportionately more credit-union members than rural areas have on average. It also means that credit-union participation tends to be associated with higher average per-capita income levels because metropolitan areas generally have higher average per-capita incomes than rural areas. As the year indicator variables show, participation rates generally increased during our sample period, holding banking market structure and population densities constant (subsequent years are measured against the baseline of 1989). This also constitutes evidence that credit unions are becoming more important competitors in the retail (i.e., household) financial-services market. A complete set of state indicator variables was used in the regressions but are not reported here.

TABLE 6 HERE

Herfindahl-Index Response Function. Table 6 presents the Herfindahl-index response function results. As our theoretical model predicts, for given values of the control variables, concentration in the commercial-banking sector increases if the corresponding county has shown an increase in the credit-union participation rate in the previous year. Banks respond to losses in market shares with consolidation (for example, in-market mergers that result in the closing of branches), or simply with exit (for example, bank failures). Taking into account the logit transformation of the variables, we find that an increase in a county's credit-union participation rate by 0.01 increases the Herfindahl index in the county by 0.0035. Thus, our evidence is consistent with the notion that credit unions are important competitors in the retail financial-services market in many localities.

Parameter estimates for the control variables given in Table 6 are consistent with findings from other studies of local banking markets. We find that bank concentration is significantly lower in more densely populated counties, as is well known from previous research. For example, the average Herfindahl index in Metropolitan Statistical Areas (MSA) was 0.1949 in 1997, while it was 0.4114 in non-metropolitan areas (Berger, Demsetz, and Strahan, 1999, Table 1). We also identify a significantly declining level of local-market bank concentration over our sample period after controlling for credit-union participation rates and population density, which is consistent with other findings as well (Berger, Demsetz, and Strahan, 1999, Table 1). State indicator variables were used in this equation, too, but are not reported here.

TABLE 7 HERE

Shifts in the Response Functions. Tables 5 and 6 shed some light on the reaction of participation rates and commercial-bank Herfindahl index values to changes in lagged values of the other variable—i.e., movements along the response functions—, as well as movements of the response functions themselves. Shifts of the response functions over time are reflected in changes in the population density and in the year indicator variables. We define an autonomous change as one that is due to the year indicator variables, which measure changes relative to the base period, the year 1989.

Table 7 provides estimates of the (cumulative) autonomous changes of the participation rate and the commercial-bank Herfindahl index. The autonomous changes in both the credit-union participation rate and the Herfindahl index were statistically significant for the later years of the 1990-96 period. After taking into account the logistic transformation we performed on the dependent variable, we find that participation rates increased by 0.0589 for reasons unrelated to the dynamic interaction of banks and credit unions or to changes in county population densities over time. Similarly, county-level bank Herfindahl index levels decreased by 0.0158 due to autonomous factors over the period 1989-96.

FIGURE 4 HERE

We summarize the quantitative effects discussed above in Figure 4. The figure shows the shifts of the response functions in the directions indicated by the data displayed in Table 7. The original equilibrium lies at point M. Depending on the relative strengths of the two shifts, the new equilibrium may exhibit a higher participation rate and a lower Herfindahl index, as shown in the new equilibrium, point M'. This is consistent with the 1.8 percentage-point increase actually observed in the median county participation rate over the 1989-96 period (see Table 4) and the 2.6 percentage-point decrease in the median county Herfindahl index (see Table 3) over the same period.

In the case of the credit-union participation rate, the dynamic response highlighted by our model was in the same direction as the autonomous effect, so they reinforced each other. In the case of the commercial-bank Herfindahl index, however, the effect from the response function worked in the opposite direction from the autonomous changes. Therefore, the autonomous effect of declining values of the commercial-bank Herfindahl index dominated the endogenous effect of increases in the Herfindahl index due to credit-union competition over the period 1989-96.

CONCLUSIONS

Credit unions are a growing part of the retail financial landscape in the United States. Credit-union expansion is a controversial issue, particularly among bank and thrift owners and managers. The strength of opposition by these interested parties alone provides some indirect evidence that credit unions are relevant competitors to banks in some markets segments of retail financial services. This article provides more direct evidence consistent with the notion that credit unions and banks compete directly.

Our dynamic theoretical model of competition between banks and credit unions makes two predictions. First, an increase in the concentration of the commercial-banking sector of a county in a given year will lead to an increase in the participation rate at occupational credit unions in this county in the next year, all else being equal. In other words, some households respond to increased concentration among local banks by moving accounts to credit unions. Second, an increase in the rate of participation at credit unions in a given year will cause more concentration in the commercial-banking market of the respective county in the next year, all else held constant. Both theoretical predictions are supported by our empirical results, suggesting that commercial banks and credit unions are direct competitors in the local household deposit market.

[1] The (commercial-bank) Herfindahl indexes reported and employed in this article are based on the commercial-bank and thrift deposits only, i.e., the exclude the credit unions.

[2] It is important to differentiate between fixed costs and sunk costs (Tirole, 1988, pp. 307-8). Fixed costs that are recoverable are not sunk and do not function as a barrier to entry. Only sunk costs confer market power.

[3] This shortcoming of the data means that evidence of interaction between commercial banks and other types of limited-purpose depository institutions such as credit unions may be obscured by the inclusion of markets in which they do not compete directly. The evidence reported in this article therefore understates the true extent of credit unions' effects on banks and may overstate the effects of banks on credit unions.

[4] Credit unions are exempt from federal income taxes, which allows them to operate with low costs of capital.

[5] Employers sponsor occupational credit unions for the benefit of their employees. There is good reason to believe that at least some employers use credit unions to deliver tax-favored fringe benefits to employees. Thus, occupational credit unions may be quite attractive to many households and may become formidable competitors for banks.

[6] See Emmons and Schmid (1999b) for a model of credit unions with multiple common bonds.

[7] See Hart and Moore (1996) for an analysis of the competitive differences between shareholder-owned for-profit firms and mutually owned firms.

[8] This holds independently of whether the one man – one vote rule applies or whether bribing is allowed (that is, side payments to assure that the most efficient decision overall is made).

[9] We implicitly assume that there is no wealth effect arising from a household's travel costs on its demand for financial services. This makes the parameters of the demand function for financial services independent of the household's credit-union-membership decision.

[10] A traditional R^2 is not well-defined in a multi-equation framework. Therefore, we calculated a pseudo R^2, which is the squared coefficient of correlation between the dependent variable and the "predicted value" for the independent variables (i.e., the sum of the dependent variable and the residuals). The adjusted pseudo R^2 was obtained from the pseudo R^2 in the usual manner.

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APPENDIX 1

ECONOMETRIC METHODOLOGY

We estimate an econometric model that pools observations from $N = 1{,}061$ U.S. counties and $T = 8$ (consecutive) years. The dataset is unbalanced, which means that for at least one year t the number of counties, n_t, falls short of $N = max\{n_1, n_2, ..., n_T\}$. The total number of observations (per equation) equals S, $S < N \times T$.

The model consists of $J = 2$ equations that were derived in the theoretical model. These are the participation-rate response function and the Herfindahl-index response function in their structural form. Let $(y_{1,t}, y_{2,t} c_{1,t}, c_{2,t}, ..., c_{k-1,t})$ be the $n_t \times (J + k - 1)$ matrix of the time-t observations of the J endogenous variables and the $k - 1$ exogenous regressors, and let $u_{i,t}$ $(i = 1, ..., J)$ be the $n_t \times 1$ vectors of the time-t disturbances, then the model reads:

(A1a) $$y_{1,t} = y_{2,t-1}\beta_{11} + \sum_{i=2}^{k} c_{i,t}\beta_{1i} + u_{1,t}$$

(A1b) $$y_{2,t} = y_{1,t-1}\beta_{21} + \sum_{i=2}^{k} c_{i,t}\beta_{2i} + u_{2,t}.$$

The variables y_1 and y_2 are the logit values of the (weighted) credit-union participation rate of the counties and the Herfindahl indexes, respectively. In each equation, the lagged value of the left-hand-side (lhs) variable of the other equation serves as a regressor. The set of exogenous variables is the same across equations. See Appendix 2 for the definitions of the variables.

The logit transformation was applied to overcome the problem that the participation rate and the Herfindahl index are constrained to the (0;1] interval. Such a constraint conflicts with the assumption of normally distributed disturbances (which requires an unbounded support). After removing a few observations of unit values—which give rise to a Tobit problem—, the original variables, y_1^{orig} and y_2^{orig}, were transformed into their logit values,

$ln(y_i^{orig} / (1 - y_i^{orig})),\ i = 1,2,$ where ln is the natural logarithm. For details on the construction of the variables y_1 and y_2, see Appendix 2.

We allow the error terms for each of the n counties in each of T years $(m = 1,\dots,n_t;\, t = 1,\dots,T),\ u_{1,m,t}$ and $u_{1,m,t}$, to be serially correlated within equations. We allow the disturbances to be contemporaneously correlated (between the same counties) across equations, i.e., we allow for nonzero values of $\sigma_{ij} \equiv cov\left(u_{i,m,t}, u_{j,m,t}\right), \forall\ i,j,m,t; i,j = 1,\dots,J; m = 1,\dots,n_t;\ t = 1,\dots,T.$ We impose zero values on contemporaneous correlation across (different) counties within equations, contemporaneous correlation among different counties across equations, and serial correlation across counties within and across equations. In summary, this means that we allow for $cov\left(u_{i,m,s}, u_{j,p,t}\right) \neq 0,\ i,j = 1,\dots,J; m,p = 1,\dots,n_t; s,t = 1,\dots,T,$ if (and only if) either $i \neq j \wedge m = p \wedge s = t$ or $i = j \wedge m = p.$

We do not allow for county-specific random or fixed effects. Instead, we control for state-specific (fixed) effects because the main cross-sectional differences in the environment of the credit unions are likely to be legal differences across states. [1]

If there were no serial correlation in either equation, the model could be estimated with a standard Generalized Least Squares (GLS) Seemingly Unrelated Regression (SUR) procedure. In the presence of serial correlation, however, these GLS-SUR estimators of the $2k \times 1$ vector of parameters, $\beta = (\beta_1', \beta_2')'$, are inconsistent because the lagged endogenous right-hand-side (rhs) variable in each equation is correlated with the error term. [2] Assume that the disturbances in equation (A1a) are serially correlated, causing $u_{1,m,t-1}$ to be correlated with $u_{1,m,t}, \forall\ t, t = 2,\dots,T$. Since $y_{1,m,t-1}$ is correlated with $u_{1,m,t-1}$, this implies that $u_{1,m,t}$ is correlated with $y_{1,m,t-1}$. In the presence of contemporaneous correlation (i.e., $u_{2,m,t}$ is correlated with $u_{1,m,t}$) this means that $u_{2,m,t}$

and $y_{1,m,t-1}$ (which is a rhs variable in equation (A1b)) are correlated. Similarly, serial correlation in equation (A1b) causes $y_{2,m,t-1}$ (which is a regressor in equation (A1a)) and $u_{1,m,t}$ to be correlated. To solve this problem, we use an Instrumental Variables (IV) approach. As instruments for the regressors $y_{1,t-1}$ and $y_{1,t-1}$, we use their ranks. The ranks are likely to be strongly correlated with the actual variables but only weakly correlated with the error term.[3]

We control for heteroskedasticity (i.e, differences in the variances across counties) and serial correlation (within counties in the same equation) using the Newey-West (1987) procedure. We obtain estimators, $\hat{\sigma}_{ij}$, for the (scalar) covariances, $\sigma_{ij}, i, j = 1,\ldots, J$, by applying the iteration procedure proposed by Oberhofer and Kmenta (1974).

Let $X_{j,t}$ be the time-t $n_t \times k$ matrix $(y_{j,t-1}, c_{1,t}, c_{2,t}, \ldots, c_{k-1,t}),\ i, j = 1,\ldots J; i \neq j$. In the absence of serial correlation (and consequently without the use of instrumental variables), the estimator of the parameter vector β in the standard GLS-SUR framework reads:[4]

(A2a) $\hat{\beta} = \left(\mathrm{X}'\hat{V}^{-1}\mathrm{X}\right)^{-1}\mathrm{X}'\hat{V}^{-1}y$

with

$$y = \begin{pmatrix} y_1 \\ y_2 \end{pmatrix}$$

$$\mathrm{X} = \begin{pmatrix} \mathrm{X}_1 & 0_{S\times k} \\ 0_{S\times k} & \mathrm{X}_2 \end{pmatrix}$$

$$^{-1} = \hat{\Omega}^{-1} \otimes I_S$$

$$\hat{\Omega} = \begin{pmatrix} \hat{\sigma}_{11} & \hat{\sigma}_{12} \\ \hat{\sigma}_{21} & \hat{\sigma}_{22} \end{pmatrix}$$

where I_S is an $S \times S$ identity matrix. The vector y_i is of size $S \times 1$, whereas the matrix X_i is of size $S \times k,\ i = 1,\ldots, J$.

Under the assumption of no heteroskedasticity, the GLS-SUR estimator of the covariance matrix of $\hat{\beta}$ reads:

(A2b) $\hat{\Sigma} = \left(X'\hat{V}^{-1}X \right)^{-1}$.

With an IV approach, these estimators are:[5]

(A3a) $\hat{\beta}^{IV} = \left(Z'\hat{V}^{-1}X \right)^{-1} Z'\hat{V}^{-1}y$.

Still assuming spherical disturbances for both equations, the estimator for the covariance matrix of $\hat{\beta}^{IV}$ is given by:

(A3b) $\hat{\Sigma}^{IV} = \left(Z'\hat{V}^{-1}X \right)^{-1} Z'\hat{V}^{-1}Z\left(X'\hat{V}^{-1}Z \right)^{-1}$

where Z is the matrix of instrumental variables.

We now consider the case in which the disturbances of each equation are nonspherical, i.e., $E\left(u_i u_i'\right) = \psi_i \ \left(i = 1,\ldots,J\right)$. The error terms may be serially correlated (within the same county in the same equation) or heteroskedastic (or both). The assumption of serial correlation necessitates the use of the estimator $\hat{\beta}^{IV}$, displayed in equation (A3a). If the $S \times S$ matrices $\psi_i\left(i = 1,2\right)$ were known, the GLS-SUR estimator for the covariance matrix of $\hat{\beta}^{IV}$ would read:

(A4a) $\hat{\Sigma}_{\psi}^{IV} = \left(Z'\hat{V}^{-1}X \right)^{-1} Z'\hat{V}^{-1}\psi\hat{V}^{-1}Z\left(X'\hat{V}^{-1}Z \right)^{-1}$,

with

$$\psi = \begin{pmatrix} \psi_1 & 0_{S\times S} \\ 0_{S\times S} & \psi_2 \end{pmatrix}.$$

Even though the GLS-SUR estimator, $\hat{\beta}^{IV}$, does not account for heteroskedasticity and serial correlation, it is a consistent estimator. Consequently, the residuals are "point-wise" consistent estimators of the error terms, $u_{1,m,t}$.[6] Based on these results, the Newey-West (1987) estimator of the covariance matrix of $\hat{\beta}^{IV}$, which is robust to heteroskedasticity and serial correlation, reads:

(A4b) $\hat{\Sigma}^{IV}_{NW} = \left(Z'\hat{V}^{-1}X\right)^{-1} W \left(X'\hat{V}^{-1}Z\right)^{-1}$

with

$$W = Z'\hat{V}^{-1}\hat{\psi}\hat{V}^{-1}Z \equiv Q + R$$

$$Q = \sum_{j=1}^{J}\sum_{m=1}^{n_t}\sum_{t=1}^{T} e^2_{j,m,t}\, g'_{j,m,t}\, g_{j,m,t}$$

$$R = \sum_{j=1}^{J}\sum_{m=1}^{n_t}\sum_{l=1}^{L}\sum_{t=l+1}^{T} w_l e_t e_{t-l}\left(g'_{j,m,t}\, g_{j,m,t-l} + g'_{j,m,t-l}\, g_{j,m,t}\right)$$

$$G = \hat{V}^{-1}Z$$

$$w_l = \frac{l}{L+1}$$

$$L = floor\left(4\left(\frac{T}{100}\right)^{\frac{2}{9}}\right).$$

The matrix W is of size $2k \times 2k$. The variable $e_{i,m,t}$ is the county-m, time-t residual of equation i $(i = 1,2)$. The row vector $g_{j,m,t}$ is the $j \times m \times t - th$ row of the $2S \times 2k$ matrix $G = \hat{V}^{-1}Z$. The operator *floor* rounds down a real number to the next smallest integer.

The covariance estimator in equation (A4b), $\hat{\Sigma}^{IV}_{NW}$, implicitly was derived for a balanced panel. We have an unbalanced panel, however, even though only a few observations were missing. (See Appendix 2 for details.) If observations were missing, they were missing in both equation (A1a) and (A1b). Technically, we calculated the estimator $\hat{\Sigma}^{IV}_{NW}$ after "inflating" the vectors of the residuals, e_i $(i = 1,...,J)$ and the matrix $\hat{}^{-1}Z$ to form a balanced panel. We did this by inserting (scalar values or row vectors of) zeroes into the vectors e_i $(i = 1,...,J)$, and the matrix $\hat{\ }^{-1}Z$ in the places where the corresponding observation (by county and year) would have been located, had it not been missing.

The equilibrium condition for the second-order difference equation derived in our theoretical model, $a_1 a_2 < 1$, translates into the condition $\beta_{11}\beta_{21} < 1$ in the terminology of the econometric model outlined above. We test this condition using the Delta method.[7] The value of the corresponding nonlinear function, $f(\beta^{IV})$, can be consistently estimated by $f(\hat{\beta}^{IV})$. The estimator for the asymptotic covariance matrix of $f(\hat{\beta}^{IV})$ reads:

(A5) $$\hat{\Sigma}^{IV}_{NW,f} = C\hat{\Sigma}^{IV}_{NW}C'$$

with $C = \dfrac{\delta f(\beta^{IV})}{\delta \beta^{IV\prime}}$.

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[1] Accounting for county-specific effects would require 1,225 county indicator variables, which would affect the efficiency of the estimation adversely.

[2] See Greene (1997, p. 435-36) on the inconsistency of Least-Squares estimators when a regressor is correlated with the error term.

[3] All other regressors are their own instruments. See Greene (1997, pp. 440-42) on the use of ranks as instrumental variables.

[4] See Kmenta (1986, pp. 635-48).

[5] See Greene (1997, pp. 291-93).

[6] See Greene (1997, p. 504).

[7] See Greene (1997, p. 278-79).

APPENDIX 2

DATASET AND VARIABLES

The Dataset

We analyze a dataset comprising all federally chartered and federally insured credit unions during the years 1989-96. Before eliminating some counties for some years due to reasons mentioned below, the credit-union dataset comprised 41,329 observations. The dataset was obtained from the Report of Condition and Income for Credit Unions (NCUA 5300, 5300S), produced by the National Credit Union Administration (NCUA). These reports are issued semiannually in June and December. We used the December data from each year. The business years of the credit unions are on a calendar basis. The flows in the December income statements cover the entire year, whereas the stocks in the balance sheet are end-of-year values.

We concentrate on the following Types Of Membership (TOM) groups among occupationally based credit unions: educational; military; federal, state, and local government; manufacturing; and services. This means that we do not include community credit unions, associational credit unions, or corporate credit unions. Lists of TOM classification codes are from the NCUA (Instruction No. 6010.2, July 28, 1995).

We excluded observations for any of the following reasons:

- Missing TOM codes
- Activity codes other than "active"
- Number of members or of potential members not greater than one; applies to actual and to lagged values
- Nonpositive values for total assets or lagged total assets
- No banks operating in the respective county (or independent city) in the year in question or in the preceding year.

Total assets, number of members, and potential number of members are end-of-year values.

We calculated county-specific Herfindahl indexes as measures of concentration of the local banking market. A Herfindahl index is defined as the sum of squared market shares. We measured market shares by the fraction of total bank deposits (as of June 30) within a county (or independent city) based on FDIC Summary of Deposits data. These data are available online at <http://www2.fdic.gov/sod/>.

We used population density to control for cross-sectional differences across counties. Population density was calculated by dividing the total county population by the total land area of the county (or independent city). Both the county population and land area data were obtained from the U.S. Census Bureau <http://www.census.gov>. The population data are Census Bureau estimates as of July 1 of the corresponding year. The land area measurements are from the 1990 census.

Definition of Variables

The county-specific credit-union participation rate was calculated as a weighted average over the participation rates of all credit unions in our dataset. We used the number of potential members to weight the participation rates.

The county-specific participation rates and Herfindahl indexes, which serve as the two endogenous variables, are bounded on the [0,1] and the (0,1] intervals, respectively. This conflicts with the assumption of normally distributed error terms. This is why we applied the logit transformation, $ln(y / (1 - y))$, where y is the participation rate and the Herfindahl index, respectively, and where ln is the natural logarithm. To keep the econometric results easy to interpret, we applied this transformation also to the lagged values of these two variables which served as explanatory variables.

The logit transformation requires the values of the variables to lie within the (0,1) interval. Whereas zero values for the Herfindahl index are not defined, zero values for the participation rate can occur. This happens when there are counties without occupational credit unions (of the TOM codes covered by this analysis). This poses the problem of data truncation, for which we did not control explicitly. In the period 1989-96, about 64 percent of the U.S. counties (independent cities included) did not have credit unions of the type we analyzed. There was one observation of a county (FIPS 8047) that did not have a bank in the year 1988 only (and thus we could not calculate a lagged Herfindahl index for 1989).

When the participation rate or the Herfindahl index are used as dependent variables, unit values pose the problem of data censoring (the so-called Tobit problem). Observations of unit values for the Herfindahl index did not occur, but there were 34 cases of unit values for the participation rate. We did not control for this Tobit problem because the number of censored observations was low. Instead, we eliminated these observations from the dataset. We also eliminated two observations of unit values for the lagged Herfindahl index and 35 cases of unit values for the lagged participation rate because these observations represent corner solutions. Because there was some overlap in these 72 cases of unit values, we only needed to eliminate 62 observations (out of a total of 8,496 county-years).

There was a total of 3,141 counties (and independent cities) in the United States in 1996, including 43 independent cities (40 in Virginia) and the District of Columbia. The corresponding number for 1989 was 3,139 (*The World Almanac and Book of Facts*, 1989, 1996). Overall, our study covers 1,061 counties (and independent cities), which is about 34 percent of the total. There were 117 observations (out of a total of 8,434 county-years), for which there was either no bank in the county (one such observation) or no occupational credit union in a given year or in the preceding year.

Definitions of variables and underlying data sources are listed below. For data taken from the Report of Condition and Income for Credit Unions, produced by the National Credit Union Administration, the relevant item numbers are in brackets.

Dependent Variables. We employed two dependent variables in the regressions:

1) Participation Rate (PARTICIPATION): Number of actual credit-union members [CUSA6091] divided by the number of potential members [CUSA6092]. In the regressions, we use the logit transformation, $ln(y / (1 - y))$.
2) HERF: Sum of squared market shares of commercial banks within a county based on total bank deposits. By definition, the Herfindahl index is greater than zero; its maximum value is one.

Independent Variables. In addition to lagged values of the dependent variables, we used three types of control variables:

1) POPDENS: Population Density, population per square mile in the respective county.
2) State indicator variables; one of the state-indicator variables was eliminated. We do not report parameter estimates for these variables because of their number and because we do not wish to place any particular interpretation on them.
3) Year indicator variables; the indicator variable for 1989 was eliminated.

APPENDIX 3

SOLUTION AND STABILITY CONDITIONS IN THE DYNAMIC MODEL

This appendix solves the dynamic model of competition outlined in the main text and then derives stability conditions. The dynamic model consists of two sector-response functions relating the extent of credit-union activity and banking market structure. When linearized using a first-order Taylor-approximation, the response functions read:

(a) $$part_t = a_2 herf_{t-1} + d_t$$

(b) $$herf_t = a_1 part_{t-1} + c_t \ .$$

Equation (a) is the participation-rate response function, which summarizes the optimal response of households to the level of bank prices for financial services, which are directly related to the concentration of the banking sector. Equation (b) is the Herfindahl-index response function, which summarizes the response of the banking sector to innovations in the credit-union sector's activity level. For statistical reasons (see Appendix 1), $part_t$ and $herf_t$ are the logit values of the participation rates of counties and the Herfindahl indexes, respectively.

Solving this system of equations for $herf$, for example, leads to the following second-order difference equation:

(c) $$herf_{t+2} - a_1 a_2 herf_t = a_1 d_{t+1} + c_{t+2} \ .$$

The solution of equation (c) consists of two parts: a particular integral, $herf_{p,t}$, that represents the (moving) equilibrium level of $herf$, and a complementary function, $herf_c$, that gives the deviation of $herf_t$ from the equilibrium value, $herf_{p,t}$.

For $a_1 a_2 \neq 1$, the particular integral reads:

(d) $$herf_{p,t} = \frac{a_1 d_{t+1} + c_{t+2}}{1 - a_1 a_2} \ .$$

The characteristic equation is $Ab^{t+2} - a_1 a_2 Ab^t = 0$, with roots $b_{1,2} = \pm\sqrt{a_1 a_2}$, where A is an arbitrary constant. Since our theoretical model is based on the assumptions that the parameters a_1 and a_2 are greater than zero, we focus on the case in which the characteristic roots, $b_{1,2}$, are distinct and real. The complementary function can thus be written as:

(e) $$herf_c = A_1 b_1^t + A_2 b_2^t$$

with $A_1 A_2$ being two arbitrary constants that are defined by two initial conditions.

The general solution to the difference equation (c) is consequently:

(f) $$herf_t \equiv herf_{p,t} + herf_c = \frac{a_1 d_{t+1} + c_{t+2}}{1 - a_1 a_2} + A_1 (a_1 a_2)^{0.5t} + A_2 (-(a_1 a_2)^{0.5})^t \ .$$

The time path of $herf_t$ will converge to the (moving) intertemporal equilibrium value, $herf_{p,t}$, if (and only if) the (positive) value of the product $a_1 a_2$ is smaller than one.[1] For $a_1 a_2 > 1$, the system will explode.

[1] In absolute value, the two roots are identical, i.e, there is no dominant root.

TABLE 1
Credit Union Type Of Membership (TOM) Codes

TOM Code [1]	Type of Membership
4	Educational
5	Military
6	Federal, state, local government
10-15	Manufacturing
20-23	Services
34	Multiple group – primarily educational
35	Multiple group – primarily military
36	Multiple group – primarily federal, state, local government
40-49	Multiple group – primarily manufacturing
50-53	Multiple group – primarily services

[1] National Credit Union Association (NCUA), Instruction No. 6010.2, July 28, 1995.

TABLE 2
Distribution of Credit Unions by TOM Code [1]

TOM Code(s) / Year	4	5	6	10-15	20-23	34	35	36	40-49	50-53	Number of Observations (Credit Unions)
1989	326	40	410	481	126	601	948	603	848	740	5,123
1990	327	40	412	482	126	603	949	604	848	742	5,133
1991	329	39	426	484	127	620	942	614	848	745	5,174
1992	329	39	427	485	127	624	944	616	848	749	5,188
1993	331	40	429	484	128	627	952	624	851	753	5,219
1994	332	41	432	489	128	634	951	628	853	756	5,244
1995	332	41	432	489	129	637	936	621	857	757	5,231
1996	322	38	415	486	125	631	846	582	834	738	5,017

[1] Before eliminating 62 county-years for the reasons mentioned in Appendix 2.

TABLE 3
Descriptive Statistics for the Herfindahl Index [1]

Herfindahl Index / Year	Minimum	Median	Mean	Maximum	Standard Deviation	Number of Observations (Counties)
1989	0.0344	0.2524	0.2707	0.9051	0.1280	1,035
1990	0.0358	0.2510	0.2699	0.8806	0.1279	1,039
1991	0.0360	0.2466	0.2661	0.8584	0.1259	1,040
1992	0.0341	0.2507	0.2654	0.8489	0.1221	1,042
1993	0.0444	0.2376	0.2611	0.8506	0.1218	1,037
1994	0.0519	0.2304	0.2572	0.8500	0.1205	1,039
1995	0.0512	0.2287	0.2532	0.8537	0.1178	1,047
1996	0.0535	0.2263	0.2519	0.8357	0.1150	1,038

[1] After eliminating 62 county-years for the reasons mentioned in Appendix 2.

TABLE 4
Descriptive Statistics for the County Participation Rate [1]

Participation Rate / Year	Minimum	Median	Mean	Maximum	Standard Deviation
1989	0.0423	0.5600	0.5696	0.9997	0.1913
1990	0.0119	0.5567	0.5733	0.9959	0.1928
1991	0.0411	0.5623	0.5726	0.9974	0.1908
1992	0.0246	0.5787	0.5846	0.9964	0.1915
1993	0.0375	0.5893	0.5937	0.9985	0.1895
1994	0.0602	0.5897	0.5909	0.9998	0.1881
1995	0.0367	0.5833	0.5824	0.9983	0.1889
1996	0.0330	0.5776	0.5822	0.9987	0.1859

[1] After eliminating 62 county-years for the reasons mentioned in Appendix 2.

TABLE 5
Participation-Rate Response Function

Independent variable	Coefficient	*t*-statistic
HERF (lagged)	2.005	22.446 ***
POPDENS	2.648×10^{-5}	2.868 ***
YEAR: 1990	7.935×10^{-3}	0.193
YEAR: 1991	2.300×10^{-2}	0.434
YEAR: 1992	1.231×10^{-1}	1.984 **
YEAR: 1993	1.716×10^{-1}	2.774 ***
YEAR: 1994	1.974×10^{-1}	3.177 ***
YEAR: 1995	1.814×10^{-1}	2.910 ***
YEAR: 1996	2.443×10^{-1}	3.912 ***
CONSTANT	2.572	14.358 ***
Wald-test (nonconstant regressors)	27,679 ***	
Wald-test (state indicator variables)	448.8 ***	
Wald-test (year indicator variables)	21.57 ***	
Pseudo-R^2	0.725	
Pseudo-R^2 adj.	0.723	
Number of observations	8,317	

/* Significant at the 5/1 percent level (*t*-tests are two-tailed).

TABLE 6
Herfindahl-Index Response Function

Independent variable	Coefficient	*t*-statistic
PART (lagged)	3.428×10^{-1}	19.072 ***
POPDENS	-1.827×10^{-5}	-4.216 ***
YEAR: 1990	1.762×10^{-2}	1.000
YEAR: 1991	-5.196×10^{-3}	-0.224
YEAR: 1992	-2.196×10^{-3}	-0.081
YEAR: 1993	-4.058×10^{-2}	-1.491
YEAR: 1994	-7.905×10^{-2}	-2.922 ***
YEAR: 1995	-9.305×10^{-2}	-3.476 ***
YEAR: 1996	-8.716×10^{-2}	-3.237 ***
CONSTANT	-1.271	-19.181 ***
Wald-test (nonconstant regressors)	3,858 ***	
Wald-test (state indicator variables)	1,443 ***	
Wald-test (year indicator variables)	22.86 ***	
Pseudo-R^2	0.860	
Pseudo-R^2 adj.	0.859	
Number of observations	8,317	

*** Significant at the 1 percent level (*t*-tests are two-tailed).

TABLE 7
Autonomous Changes in Concentration

Year	Autonomous changes relative to the median values in 1989	
	Participation rate	Herfindahl index
1990	0.0020	0.0033
1991	0.0057	-0.0010
1992	0.0300 **	-0.0004
1993	0.0417 ***	-0.0075
1994	0.0479 ***	-0.0144 ***
1995	0.0440 ***	-0.0169 ***
1996	0.0589 ***	-0.0158 ***

/* Significant at the 5/1 percent level in two-tailed tests (based on Tables 5 and 6).

Figure 1: The Circular Economy

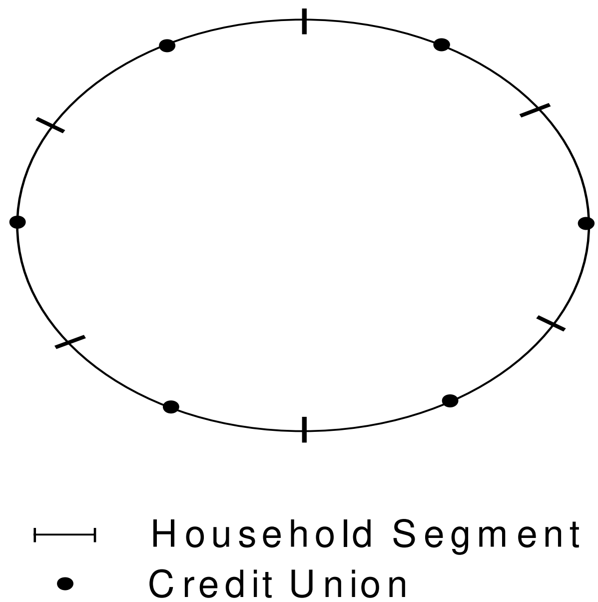

Figure 2: Credit Union Participation

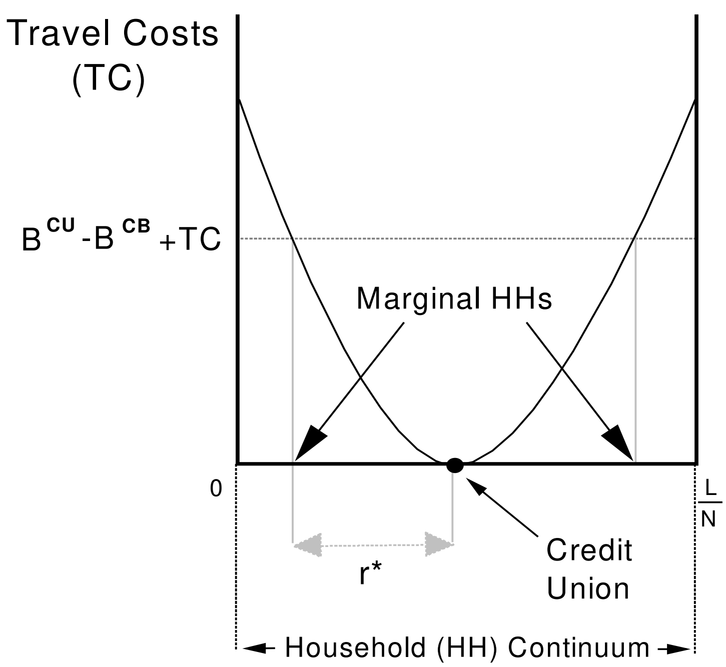

Figure 3: Dynamic Response Model

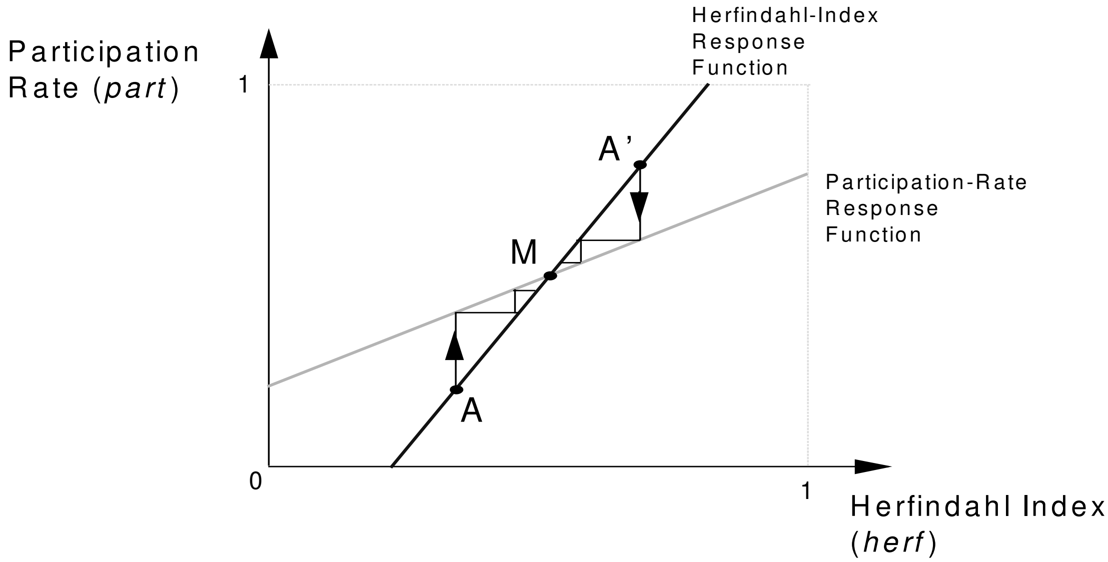

Figure 4: Autonomous Changes in Participation and Concentration