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Bank Market Structure and Local Employment Growth

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Abstract

The relationship between financial structure and job growth is both an unexplored issue and a possible channel through which financial structure impacts income growth. We explore these issues using both longrun and shortrun models. Our shortrun model provides evidence of a robust relationship between local employment growth and geographic deregulation of bank activity in the United States. We also found that U.S. nonmetropolitan employment grew faster in 1973-96 where there were fewer locally owned bank offices and a more concentrated initial banking market structure; these linkages were less stable in metropolitan areas. Overall, however, we found only weak evidence in support of an employment growth channel linking bank structure to subsequent economic growth. Our findings suggest that job creation is not consistently a major channel by which banking structure stimulates income growth. A corollary is that the macroeconomic benefits of banking structure accrue primarily to those already working, rather than new workers.

Keywords: commercial banking, employment growth, geographic liberalization, bank ownership.

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Summary

A growing body of empirical literature has established a linkage between the market structure of the financial intermediation sector and economic growth rates. The general conclusion of this literature has been that a larger, deeper, or more efficient financial intermediation sector tends to be associated with more rapid growth rates of per capita income. Previous ERS research suggests that mergers or acquisitions of local banks by nonlocal banks need not impair local economic growth and may even have beneficial effects in rural markets, with the possible exception of farm-dependent areas. However, possible associations between banking structure and other aspects of macroeconomic activity including employment growth were neglected in earlier studies.

This report explores the empirical linkages between growth rates in total local employment and the structure of local bank markets (ownership, deposit control, and concentration) and their level of geographic deregulation. These linkages are of interest for two reasons. First, local communities, their leaders, and policymakers are all sensitive to employment growth and contraction. These constituencies want to know if nonlocal ownership of bank offices or nonlocal control of the deposit base (and, therefore, geographic deregulation of banking) is associated with faster or slower rates of employment growth. Second, the identification of plausible mechanisms by which banking structure may influence the growth rates of per capita income might further strengthen our confidence in the causal nature of the empirical association between banking structure and economic growth found by previous research.

With these ideas in mind, we explored both longrun and shortrun empirical linkages between banking structure and subsequent employment growth in local geographic markets in a nationwide sample spanning 1973-96. Our findings suggest that the initial number of bank offices, the relative market shares of banks, and the ownership structure of bank offices (local versus out-of-market) tend to be associated with subsequent longrun but not shortrun growth rates of local employment. Further, some of these linkages have shifted over time, and differ systematically in metropolitan versus nonmetropolitan markets. Consistent with previous research, we found no evidence that, on average, geographic deregulation, nonlocal bank office ownership, or nonlocal deposit control adversely impacted nonmetropolitan income or employment growth rates.

The observed linkages between the bank market structure and employment growth were in some cases quite different from those reported between bank market structure and income growth. In addition, the inclusion of contemporaneous employment growth rates did not substantially change the linkages between banking structure and either longrun or shortrun income growth rates. These findings suggest that job creation, while responsive to banking structure and important in its own right, is not a major channel by which banking structure stimulates per capita income growth.

The longrun regressions provided evidence that the growth rates of local employment tend to be associated with the initial numbers of bank offices, initial relative market shares of banks, and the initial ownership structure of banks (local versus out-of-market). Shortrun regressions generally failed to find a significant association between bank ownership and deposit control variables and employment growth in either metropolitan or nonmetropolitan markets. Mirroring the longrun results, employment grew faster in more concentrated nonmetropolitan markets but more slowly in more concentrated metropolitan markets. The shortrun regressions also showed a positive and

economically significant association between geographic deregulation and employment growth. However, this association differed between metropolitan and nonmetropolitan markets. In nonmetropolitan markets, employment growth was more sensitive to the first stage of geographic deregulation (allowing nonlocal entry through mergers and acquisitions of existing banks), while the second stage of geographic deregulation (allowing nonlocal entry through *de novo* branching) was more important in metropolitan markets.

The strongest evidence for an employment growth channel linking banking to economic growth was from the shortrun model results related to geographic deregulation. Overall, however, this linkage appeared to explain only a small portion of the relationship between banking and economic growth. Otherwise, we found only weak evidence in support of an employment growth channel linking bank structure to subsequent economic growth. These findings suggest that job creation, while responsive to banking structure and important in its own right, is not consistently a major channel by which banking structure stimulates income growth. A corollary is that the macroeconomic benefits of banking structure accrue primarily to those already working, rather than to new workers. Thus, the stimulus to growth provided by financial structure has diverse distributional effects, a previously overlooked point that may warrant further study.

Introduction

A growing body of empirical literature has established a linkage between the market structure of the financial intermediation sector and local economic growth rates. This linkage has been studied across countries (King and Levine, 1993a and b; Levine, 1998; Rajan and Zingales, 1998; Cetorelli and Gambera, 2001), across States within the U.S. (Jayaratne and Strahan, 1996, hereafter JS; Krol and Svorny, 1996), and at the level of local markets (Collender and Shaffer, 2000, hereafter CS). The general conclusion of this literature has been that a larger, deeper, or more efficient financial intermediation sector tends to be associated with more rapid subsequent growth rates of per capita income. However, though JS provided evidence that growth in their sample was fostered by improved screening and monitoring of loan quality, the causal mechanism relating banking structure to income growth is not completely understood. Moreover, possible associations between banking structure and other aspects of macroeconomic activity have been neglected in earlier studies.

To the extent that a plausible mechanism can be identified by which banking structure may influence the growth rates of per capita income, the recognition of such a mechanism might further strengthen our confidence in the causal nature of the empirical association between banking structure and economic growth. In the absence of an explicit mechanism, the question of causality must rely more heavily on purely statistical tests.

With these ideas in mind, we explored both longrun and shortrun empirical linkages between banking structure and subsequent employment growth in local geographic markets in a nationwide sample spanning 1973-96. We also examined how the impact of banking structure on subsequent income growth was affected by correcting for contemporaneous employment growth. Our findings suggest that the initial number of bank offices, the relative market shares of banks, and the ownership structure of bank offices (local versus out-of-market) tend to be associated with subsequent longrun but not shortrun growth rates of local employment. Further, some of these linkages have shifted over time, and differ systematically in metropolitan versus nonmetropolitan markets.[2] However, the observed linkages between the structural regressors and employment growth were in some cases quite different from those reported between similar regressors and income growth by CS. In addition, the inclusion of contemporaneous employment growth rates did not substantially change the linkages between banking structure and either longrun or shortrun income growth rates. These findings suggest that job creation, while responsive to banking structure and important in its own right, is not a major channel by which banking structure stimulates per capita income growth. In contrast to these negative results, our shortrun results provide modest evidence consistent with the notion that employment growth constitutes a linkage between geographic deregulation of bank activity and subsequent economic growth as documented by JS and CS.

A few previous studies have explored determinants of employment growth. O hUallachain and Satterthwaite (1992) investigate this issue for metropolitan areas by sector, but omit any measure of financial structure. Henderson and McDaniel (2000) present empirical estimates of employment growth in rural counties, and likewise do not include banking market structure among the regressors.

[2] DeYoung et al. (1999) disaggregated urban versus rural lending markets and found that small business lending is positively related to concentration rates in urban areas but negatively related to concentration in rural areas. Avery et al. (1999) found no differences in the effect of within-ZIP-code banking consolidation on branch density between urban and rural markets.

Background

A separate body of research has explored the impact of bank consolidation on lending to small businesses.[3] The rapid decline in the number of U.S. banks via merger and acquisition since 1986 has stimulated practical interest in this question. Though there is some evidence that bank consolidation reduces lending to small businesses, the findings overall have been mixed.[4] Conflicting theoretical predictions have failed to shed light on this issue. On the one hand, the venerable structure-conduct-performance (SCP) paradigm predicts that more banks would generate stronger competition in lending, driving down the interest rates and fees charged for loans and generating a larger aggregate amount of lending. If only projects with a positive expected net present value are funded in equilibrium, a further implication would be that a larger number of banks should be associated with higher macroeconomic growth rates. Indeed, one might argue that the importance of any linkage between banking structure and lending to small businesses ultimately derives from its impact on economic growth.

On the other hand, Petersen and Rajan (1995) postulated that small businesses in more concentrated banking markets may find it easier to obtain bank loans because market power makes loan contracts easier to enforce. Thus, the linkage between bank structure and economic growth is theoretically ambiguous. Similarly, a number of observers have postulated that U.S. banking exhibited excess capacity through the 1970s and 1980s (see Frydl, 1993; Edwards, 1996). This notion, sometimes called the "overbanking hypothesis," implies that excess resources may have been channeled into banking. Consistent with this hypothesis, Shaffer (1996) presented evidence that a small amount of unsustainable excess capacity in loans existed in U.S. banking from the 1930s until the mid-1980s. Overbanking suggests that macroeconomic growth rates might potentially benefit from some degree of rational consolidation among banks.

Recent studies of adverse borrower selection have identified a further possibility. Broecker (1990), Riordan (1993), and Shaffer (1998) all noted a phenomenon known as "the winner's curse." The winner's curse refers to the possibility that high-risk borrowers rejected by one bank can apply at additional banks and eventually obtain a loan if there are enough banks and if banks' credit-screening mechanisms are both noisy and imperfectly correlated. Thus, loan loss rate may be an increasing function of the number of banks in a market, a pattern empirically supported in Shaffer (1998). The winner's curse could conceivably offset the favorable effect of atomistic competition, if too many banks fund projects with negative expected value and thereby reduce macroeconomic growth. The limited empirical evidence on this point suggests that greater numbers of banks are associated with higher subsequent growth rates of income per household in metropolitan areas (ibid.; see also CS) and possibly also in rural markets (CS).

The historical linkage between firm size and job growth strengthens both the relevance and the complexity of this issue and its policy implications. Historically, smaller firms have created the most new jobs (see, for example, Davis and Haltiwanger, 1992; Davis et al., 1996). This pattern might seem to suggest that reduced lending to small firms could impair job growth and thereby restrict overall economic growth. However, jobs at smaller firms have tended to be less permanent than those at larger firms, in part because of the higher failure rates of smaller firms (ibid.; Rob, 1995). Therefore, even if the restructuring of the banking industry shifted lending away from smaller businesses toward larger firms overall, the net macroeconomic impact on employment and growth would remain an open empirical question.

Characterizing the empirical linkages between banking structure and subsequent employment growth can potentially shed light on this issue, as well as on one possible mechanism by which financial structure may influence the growth rate of income. The present study

[3] See, for example, Whalen, 1995; Berger and Udell, 1996; Peek and Rosengren, 1996; Avery et al., 1999; DeYoung et al., 1999; Jayaratne and Wolken et al., 1999.

[4] While larger banks typically make a smaller percentage of small business loans than smaller banks (Berger and Udell, 1996; Peek and Rosengren, 1996; Jayaratne and Wolken, 1999), small businesses do not appear to be more credit constrained in the absence of smaller banks (Jayaratne and Wolken, 1999). Berger et al. (1998) found that banks overall originate fewer small business loans following their participation in a merger; however, if the merging banks were both small, their merger was associated with a subsequent increase in small business lending (Peek and Rosengren, 1998; Strahan and Weston, 1998).

focuses on market-level linkages. It follows previous research and regulatory practice in defining individual banking markets as metropolitan statistical areas (MSAs) or nonmetropolitan counties (Whitehead, 1990; Jackson, 1992; Kwast et al., 1997). This concept of localized banking markets is empirically supported for small-business borrowers and retail depositors in particular (Elliehausen and Wolken, 1990; Jackson, 1992). The sample comprises more than 2,200 nonmetropolitan counties and more than 260 MSAs. The measures of banking structure include the market-wide Herfindahl-Hirschman index (HHI) computed for deposits (as in several previous studies); and various measures of ownership and banking growth described below (as in CS).

Conceptual Issues and Empirical Models

Previous studies of the impact of bank consolidation on lending to small businesses have tended to focus on the experience of individual banks that have undergone a merger or acquisition. Such studies, besides providing mixed results, are intrinsically unable to distinguish among various possible *causes* of the observed relationships. If consolidation is followed by a reduced share of bank assets devoted to small business lending, is that reduction a function of the larger size of the post-merger bank, or of a shift from local to remote ownership and decisionmaking with a concomitant neglect of viable local credit demand, or of a reduction in the number of local lenders and consequently a slackening of competition?[5]

Likewise, prior studies do not fully address the *consequences* of the observed relationships. If one consequence of too many lenders is more loans to negative-value projects because of adverse selection or ruinous competition, and if consolidation reduces the number of lenders in a market, then consolidation among banks could simultaneously reduce aggregate lending and increase overall economic activity, welfare, and growth. If reduced lending is targeted specifically against small businesses, and if small businesses typically generate most new jobs, then the growth of aggregate employment may also be affected by bank consolidation, concentration, and ownership structure.

The nexus between small business as a source of employment growth and de novo entry as a source of small business lending may also be important. Small firms create roughly 75 percent of all net new jobs in the economy (Small Business Administration, 2000). Several studies have found evidence that *de novo* banks lend disproportionately more of their assets to small businesses than do mature small banks, other things held equal, and that this difference tends to persist for as along as 20 years after entry (DeYoung, 1998; Goldberg and White, 1998; DeYoung, Goldberg, and White, 1999). These studies examined *de novo* banks (newly chartered banks) as distinguished from *de novo* branches (branches built, rather than acquired from other banks, by previously chartered banks), for which data are generally not available. Within these limitations, such studies do raise the possibility that easing new entry into banking markets may spur both economic growth and employment by improving performance in the market for small business loans.

Therefore, an important set of unexplored issues is the linkage between bank concentration and ownership structure versus growth rates in local employment. The empirical tests below distinguish the effects on growth from the raw number of bank offices, from the market concentration of banks, and from the mix between locally owned and remotely owned bank offices. The tests further distinguish between the linkages observed in metropolitan markets and those in nonmetropolitan markets, and explore the stability of these linkages over time. And, in conjunction with previous studies linking banking structure to income growth (such as JS; Krol and Svorny, 1996; Shaffer, 1998; and CS), they provide the first evidence as to whether job growth may be a significant channel by which financial intermediation is associated with per capita income growth.

Our goal is to test whether employment growth is an avenue through which geographic liberalization of banking increases local economic growth. Previous research has indicated that economic growth effects may stem from improved loan portfolio efficiency (King and Levine, 1993a and 1993b; JS) and has failed to find evidence that the locus of bank office ownership or deposit control or changes in HHI are explanatory mechanisms (CS). To explore these issues we examined both longrun and shortrun models, using two strategies to test for a linkage between bank market structure and local employment growth in models based on those used by CS.

Our first strategy used the CS models of banking structure and growth rates in local real per capita personal income, but replaced income growth rates with employment growth rates as the dependent variable. This approach allowed us to examine whether or not employment growth is linked to banking structure. Comparing the signs, magnitudes, and statistical significance of the coefficients from these regressions with those found in CS provides some insight as to the potential role of employment growth as a channel through which banking structure influences income growth.

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[5] Perhaps the most relevant prior study for these issues is Whalen (1995), who explored the impact on small business lending when an acquiring bank holding company was headquartered in a different State from the target bank.

The second strategy augmented the CS models by adding employment growth rates to the explanatory variables. This strategy afforded a more direct test of employment growth as a channel through which banking structure may influence income growth. If the magnitudes and/or statistical significance of various coefficients on bank structure variables in an income growth regression are sensitive to the inclusion of employment growth, then employment growth is more likely to be an important channel for the linkage between bank deregulation and bank structure and income growth documented by J&S and CS. However, such a conclusion must be drawn with caution since spurious correlation, reverse causality, or joint causality may also explain such observations. To the extent that evidence of a linkage was found, we explored those possibilities as well.

Longrun Models

Besides the market-wide Herfindahl-Hirschman index (HHI) of deposits that is commonly used both in empirical banking research and by Federal regulators in assessing the degree of banking competition, we used a variety of additional measures of market structure and bank ownership: numbers of offices of banks headquartered in the market at the beginning of the sample period (NIB); numbers of local branches of banks headquartered outside the market at the beginning of the sample period (NXB); the ratio of remotely owned to locally owned branches at the beginning of the sample period (XTB); the growth rate in the number of locally owned bank offices during the sample period (DIB); and the growth rate in the number of remotely owned bank offices during the sample period (DXB). These variables permit a decomposition of the effects of raw numbers of bank offices, relative sizes of banks, local versus remote bank ownership, and trends in each of these factors. The locus of ownership is potentially relevant to credit patterns because many multi-market banks centralize their lending decisions for larger loans, with the final decision being made outside the borrower's market, and because size is related to technology and larger banks are more likely to be nonlocally owned (Cole, Goldberg, and White, 1999; Haynes, Ou, and Berney, 1999).

The time period spans 1973-96 and is fitted as two contiguous periods, 1973-84 and 1984-96.[6] The use of a single growth rate measured over a period of 12 or 13 years in each regression parallels that of Levine (1998) and others, and provides the advantages of smoothing out high-frequency intertemporal noise and mitigating the impact of outlier years on growth rates. Several factors suggest that the empirical linkages may be different in the first half of the period than in the second. The structure of U.S. banking remained fairly stable during the first half with more than 14,000 banks nationwide from 1970 through 1986, followed by an almost linear decline to fewer than 10,000 banks by the end of 1996. Most of the decline was the result of mergers and acquisitions, though a precipitous rise in the number of bank failures (peaking in the years 1985-92) also contributed to the trend in the mid-1980s. A major wave of banking deregulation began in 1980 with the Depository Institutions Deregulation and Monetary Control Act, many provisions of which (such as the removal of ceilings on deposit interest rates) were phased in over a subsequent multi-year period. Other Federal laws that further deregulated various aspects of banking were passed during the 1980s. At the same time, many States relaxed their restrictions on bank branching, opening the door toward consolidation across local banking markets and permitting aggressive competition from more distant banks.

The model also includes a vector of control variables as follows. The inclusion of the deposits per capita as of the initial year of the regression period (DPC) controls for the relative supply of funds and intensity of intermediation in the market, similar to King and Levine (1993b). The change in the ratio of deposits in nonlocally owned branches to deposits in locally owned bank offices over the sample period (DDEP) controls for any shift in the aggregate market share of remotely owned bank offices. We do not attach a causal interpretation to this variable because it will reflect any structural response by the banking industry to contemporaneous local economic conditions and trends. The natural logarithm of the percentage of the total adult population having completed at least 4 years of college (LEDU) as of 1970—or, for the later regressions, 1980—controls for the average level of education, a proxy for human capital and work force quality. The natural logarithm of market population (LPOP) as of the first year of the regression period controls for market size in demographic terms, and may also be interpreted as a measure of urbanization economies (similar to the total labor force variable used in O hUallachain and Satterthwaite, 1992; Henderson and McDaniel, 2000). The natural logarithm of real per capita personal income in the market as of the initial

[6] The time periods are not overlapping in that the endpoint of the first 1984 is the starting point of the second. That is, the data from the 1984 calendar year are not included in both periods.

year of each regression (LRPC) controls for initial market wealth. These demographic and income variables were obtained from BEA data. Table 1 summarizes the definitions and sources of the regressors.[7]

The regression relating longrun employment growth rates to these factors is therefore:

$$
\begin{aligned}
(1)\quad EG_{0,T} = {} & \alpha_0 + \alpha_1 NIB_0 + \alpha_2 NXB_0 + \alpha_3 XTB_0 \\
& + \alpha_4 DIB_{0,T} + \alpha_5 DXB_{0,T} + \alpha_6 DDEP_{0,T} \\
& + \beta_1 DPC_0 + \beta_2 LEDU_0 + \beta_3 LPOP_0 \\
& + \beta_4 LRPC_0 + \beta_5 HHI_0 + \varepsilon
\end{aligned}
$$

where $EG_{0,T}$ is the geometric mean of the annual growth rates of total employment in the market from the initial year, 0, to the end of the period, T, and the other variables are defined above. In the nonmetropolitan regressions, USDA county typology dummies are also included to indicate farming-dependent (FM) and mining-dependent (MI) counties, measured as of 1989 (see Cook and Mizer, 1994, for further details). Although other typologies are also assigned to counties by the USDA, systemic shocks to agriculture and mining during the 1980s made it essential to control for these two characteristics in particular.

A second regression relates longrun economic growth rates to these same variables, controlling for employment growth rates:

$$
\begin{aligned}
(2)\quad YG_{0,T} = {} & \gamma_0 + \gamma_1 NIB_0 + \gamma_2 NXB_0 + \gamma_3 XTB_0 \\
& + \gamma_4 DIB_{0,T} + \gamma_5 DXB_{0,T} + \gamma_6 DDEP_{0,T} \\
& + \delta_1 DPC_0 + \delta_2 LEDU_0 + \delta_3 LPOP_0 \\
& + \delta_4 LRPC_0 + \delta_5 HHI_0 + \varphi EG_{0,T} + \eta
\end{aligned}
$$

[7] Some previous studies, such as O hUallachain and Satterthwaite (1992) and Henderson and McDaniel (2000), have included different regressors such as measures of per capita income growth, localization economies, government fiscal policies, and hedonic indices (climate, recreation, or scenic amenities). We did not include per capita income growth here because it is considered endogenous, as rendered explicit in our income growth regressions. We omitted localization economies (industry employment) because our model is market-wide rather than sector-specific. We omitted government fiscal variables because O hUallachain and Satterthwaite (1992) found no significant linkage with employment growth in metropolitan areas, while Henderson and McDaniel (2000) similarly found no significant linkage with employment growth among rural counties.

where $YG_{0,T}$ is the geometric mean of the annual growth rates of real per capita personal income from the initial year, 0, to the end of the period, T, and the other variables are defined above. This equation corresponds to the longrun economic growth equation (equation 4 estimated in CS), augmented by the inclusion of $EG_{0,T}$. This equation allows us to determine whether employment growth rates are significantly associated with per capita income growth rates after controlling for banking structure and other factors. In conjunction with the results reported in CS, equation 2 also indicates whether controlling for employment growth rates alters the associations between the remaining variables and per capita income growth rates. To the extent that aggregate employment growth is both determined by prior financial structure and a significant channel of per capita income growth, its inclusion here should diminish the absolute magnitudes and statistical significance of the coefficients on the measures of financial structure. For both equation 1 and equation 2, separate regressions were fitted for nonmetropolitan counties alone and for MSAs alone.

Shortrun Models

Following CS, we also estimated a set of shortrun regressions. Corresponding to the longrun regressions, equations 3 and 4 model shortrun growth in local employment and local per capita real income, respectively. Equation 3, like equation 1 for the longrun case, estimates the relationship between local employment growth rates and the structure of the local bank industry and regulation. As with equation 2, equation 4 relates local economic growth rates to geographic deregulation and to various measures of local bank market structure controlling for employment growth rates. Thus, our shortrun models are:

$$
\begin{aligned}
(3)\quad EG_{t,i} = {} & \alpha_t + \beta_i + \gamma_1 DMA_{t,i} + \gamma_2 DNOVO_{t,i} \\
& + \delta_1 HHI_{t,i} + \delta_2 NIB_{t,i} + \delta_3 NXB_{t,i} \\
& + \delta_4 IDEPS_{t,i} + \delta_5 XDEPS_{t,i} + e_{t,i}
\end{aligned}
$$

$$
\begin{aligned}
(4)\quad Y_{t,i}/Y_{t,i} = {} & \alpha_t + \beta_i + \gamma_1 DMA_{t,i} + \gamma_2 DNOVO_{t,i} \\
& + \delta_1 HHI_{t,i} + \delta_2 NIB_{t,i} + \delta_3 NXB_{t,i} \\
& + \delta_4 IDEPS_{t,i} + \delta_5 XDEPS_{t,i} + e_{t,i},
\end{aligned}
$$

where β_i represents the cross-section-specific—or local market—component of longrun economic growth; α_t represents the common, economywide shock to growth at time t; $DMA_{t,i}$ is a binary variable equal to 1 for

Table 1—Regressors and their sources

Variable	Description
DMA	Binary variable equal to 1 if market entry allowed through mergers and acquisitions. Source: Amel, no date.
DNOVO	Binary variable equal to 1 if market entry allowed through establishing new branches. Source: Amel, no date.
NIB	Initial number of in-market owned bank offices. Source: FDIC Summary of Deposits.
NXB	Initial number of out-of-market owned bank offices. Source: FDIC Summary of Deposits.
XTB	Initial ratio of out-of-market owned bank offices to total bank offices. This ratio is undefined for markets with zero bank offices. For those markets, we set XTB equal to 1 under the presumption that such markets are more like those whose banks are controlled outside the local market than those whose banks are controlled in-market. Computed from FDIC Summary of Deposits.
DIB	Ratio of the number of in-market owned bank offices at end of period to that at beginning of period. This ratio is undefined for markets with zero in-market owned bank offices in the base year. For those markets, we set the initial level at 0.01. Computed from FDIC Summary of Deposits.
DXB	Ratio of the number of out-of-market owned bank offices at end of period to that at beginning of period. This ratio is undefined for markets with zero out-of-market owned bank offices in the base year. For those markets, we set the initial level at 0.01. Computed from FDIC Summary of Deposits.
DDEP	Change in the ratio of deposits held at out-of-market owned bank offices to total deposits at bank offices from beginning of period to end of period. This ratio is undefined for markets with zero deposits in bank offices in the base year. For those markets, we set the initial level to 0. If, for example, the market has no deposits in bank offices in either the initial or final year, then DDEP is set to 0. Computed from FDIC Summary of Deposits.
DPC	Initial level of deposits per capita held at all bank offices in market. Computed from FDIC and BEA data.
LEDU	Log of the percent of total adult population with at least 4 years of college at the beginning of the decade in which t0 falls. Source: U.S. Census 1970, 1980.
LPOP	Log of market population (in millions). Source: BEA
LRPC	Log of real per capita personal income (in thousands) in market. Source: BEA
HHI	Initial market level (MSA or rural county) Herfindahl-Hirschman Index (divided by 10000) computed with banks consolidated to the holding company level. For markets with zero banks, this is set equal to 1 under the presumption that consumers in these markets will have no more choices than those in markets served by only one bank. Computed from FDIC Summary of Deposits.
IDEPS	Initial amount of deposits controlled by in-market owned banks. Source: FDIC Summary of Deposits.
XDEPS	Initial amount of deposits controlled by out-of-market owned banks. Source: FDIC Summary of Deposits.
Nonmetropolitan county types (source: Economic Research Service/USDA computation based on BEA data):	
FM	Farming-dependent, 1989 (farm income > 20% of total income, 1987-89)
MI	Mining-dependent, 1989 (mining income > 15% of total, 1987-89)

markets in States that allow unrestricted branching through mergers and acquisitions in year t, DNOVO is a binary variable equal to one for markets in States that allow unrestricted *de novo* branching in year t, HHI is the Herfindahl-Hirschman index of bank deposits (the sum of squared market shares for all market participants), IDEPS is the inflation-adjusted amount of local deposits controlled by in-market owned banks, XDEPS is the inflation-adjusted amount of local deposits controlled by out-of-market owned banks, and NIB and NXB are as defined earlier. DMA and DNOVO help account for the typical two-stage process of geographic liberalization as documented by Amel (no date). In the first stage, multibank holding companies (MBHC's) may convert subsidiary banks into branches and may expand geographically through acquisition and conversion of existing banks. In the second stage, banks are allowed to expand geographically by establishing new (*de novo*) branches anywhere in the State. NIB, NXB, IDEPS, and XDEPS provide information on the impact of nonlocal ownership of bank offices and control of deposits.

Equation 3 allows us to test hypotheses relating local employment growth to geographic liberalization, local market growth, and the loci of bank office ownership and of control of local deposits (in-market and out-of-market). First, we tested for a statistically significant relationship between our explanatory variables and local economic growth. Then, we tested whether the coefficients on each pair of variables related to local and nonlocal control were the same. That is, we tested whether the relationship of growth to nonlocally owned offices or nonlocally owned deposits was the same as the relationship of growth to locally owned bank offices or locally owned deposits. Equation 4 provides evidence about the importance of employment growth as a significant channel through which bank structure and regulation affect local economic growth.

Sample and Estimation

Table 1 lists the variables and their sources. Following previous research and regulatory practice, we define local markets as metropolitan statistical areas (MSAs) or nonmetropolitan counties (Whitehead, 1990; Jackson, 1992; Kwast et al., 1997). Different agencies define U.S. counties somewhat differently because of anomalies among States and changes over time. To ensure consistency across data sets and over time, we imposed the following standards on the data. We defined urban banking markets based on 1993 definitions of MSA's and held this definition constant over the sample period to abstract from local changes over time. We defined rural banking markets as counties not included in MSA's. For consistency with previous research, we excluded Alaska and Hawaii from our shortrun models but not our longrun model. We aggregated Virginia's independent cities with the county that surrounds them, and aggregated certain counties in Montana and Wisconsin for which treatment is not uniform across agencies. We used data from years 1981-96 to estimate our shortrun models and from 1973, 1984, and 1996 for our longrun model.

Our measures of market concentration and deposit control were derived from the FDIC's annual Summary of Deposits report. We defined the locus of ownership and control as either in-market or out-of-market based on the location of a bank's headquarters office at the bank charter level, not at the holding company level. We eliminated banks with nonpositive aggregate deposits across all offices, but included offices that reported zero deposits at the county level.

Local employment growth rates were calculated from Bureau of Economic Analysis (BEA) estimates of county-level employment. Similarly, per capita personal income was calculated from BEA estimates of county populations and personal incomes adjusted for inflation using the national consumer price index. To control for educational attainment, we used data from the Bureau of the Census on the percentage of adult population in each county with at least 4 years of college at the start of the relevant decade.

Sample Statistics and Correlations

Table 2 reports descriptive statistics for the data set, decomposed by subperiod and by metropolitan versus nonmetropolitan markets. The table shows that metropolitan markets experienced more rapid growth than nonmetropolitan markets in both employment and real per capita personal income. The gap in employment growth rates, as well as the mean rates of employment growth, remained similar between 1973-84 and 1984-96. The gap in income growth rates narrowed in the later period, even as mean income growth rates rose in both sets of markets.

Structurally, rural and urban markets were quite different. The nonmetropolitan markets in our sample had an average of only eight banking offices (versus 152 for MSAs), a correspondingly higher HHI (0.42 versus 0.18), and much lower levels of aggregate deposits ($159 million versus $6 billion). Standard deviations and coefficients of variation (ratios of the standard deviation to the mean) on these variables indicated that nonmetropolitan markets were structurally more homogeneous in both absolute and relative terms than metropolitan markets, the latter being skewed by a few outliers such as New York, Los Angeles, and Chicago.

Nonmetropolitan markets have experienced geographic liberalization at a slower pace and entry by nonlocal firms has been less likely after liberalization. The relatively slow rate of entry into nonmetropolitan markets has previously been documented by Amel and Liang (1992 and 1997) and by Berger, Bonime, Goldberg, and White (2000). Despite these observations, control of local banking markets by out-of-market banks is surprisingly similar in nonmetropolitan and metropolitan markets: out-of-market banks controlled 27 percent of nonmetropolitan bank offices (versus 29 percent of metropolitan) and 26 percent of nonmetropolitan bank deposits (versus 28 percent of metropolitan).

Some striking differences between pairwise correlations in the metropolitan and nonmetropolitan samples should be noted. The correlation between the numbers of in-market and out-of-market owned bank offices is 0.01 in nonmetropolitan areas but 0.48 in metropolitan markets. That is, in-market and out-of-market office numbers often exhibit similar structures in metropolitan markets but not in nonmetropolitan markets. A corresponding contrast arises in in-market vs. out-of-market controlled deposits. Finally, the correlation

Table 2a—Descriptive statistics for longrun models

	1973-84				1984-96			
Variable	Mean	St. Dev.	Min.	Max.	Mean	St. Dev.	Min.	Max.
				Nonmetropolitan counties				
IncGth	0.00248	0.0164	-0.1238	0.085	0.0108	0.0112	-0.0771	0.0563
EmpGth	0.0129	0.0193	-0.0452	0.2184	0.0127	0.0160	-0.0543	-0.1522
NIB	4.710	4.178	0	42	5.822	5.1179	0	46
NXB	0.983	2.4917	0	22	1.795	3.884	0	31
XTB	0.167	0.3259	0	1	0.206	0.3365	0	1
HHI	0.4727	0.261	0.0799	1	0.4404	0.2407	0.0784	1
DIB	1.853	3.1724	0	40	1.130	1.1289	0	18
DXB	3.369	11.5616	0	230	12.242	20.9597	0	210
DDEP	0.042	0.183	-1	1	0.191	0.2892	-1	1
DPC	2.32	1.0421	0	7.2305	6.450	3.2101	0	30.5946
LEDU	-2.8255	0.4136	-4.5254	-1.0188	-2.3462	0.3538	-3.467	-0.773
LPOP	9.5292	0.9214	5.6699	11.9975	9.6248	0.9424	4.4659	11.9893
LRPC	2.2593	0.2701	1.4106	3.3028	2.2862	0.2015	1.2832	3.3722
FM	0.245	0.4302	0	1	0.245	0.4299	0	1
MI	0.064	0.2448	0	1	0.064	0.2448	0	1
				Metropolitan statistical areas				
IncGth	0.00982	0.00657	-0.00798	0.034	0.0141	0.00589	-0.00946	0.0282
EmpGth	0.0214	0.0163	-0.0158	0.0735	0.0218	0.0108	-0.0073	0.0699
NIB	54.404	73.244	0	526	97.693	181.89	0	1365
NXB	8.173	19.416	0	139	20.655	38.638	0	261
XTB	0.151	0.274	0	1	0.222	0.289	0	1
HHI	0.2203	0.0935	0.0456	0.5646	0.1957	0.0789	0.0403	0.4872
DIB	3.429	9.546	0	100	4.218	8.704	0	69.667
DXB	30.754	99.440	0	820	60.345	142.852	0	1190
DDEP	0.0723	0.1568	-0.1396	0.7123	0.2179	0.245	-0.296	0.872
DPC	2.365	0.6608	0.8858	4.5108	5.6113	2.162	1.790	23.736
LEDU	-2.2715	0.3263	-2.9786	-1.177	-1.8785	0.2988	-2.5582	-0.953
LPOP	12.3689	0.9331	10.5125	14.9283	12.5807	1.0172	11.0938	15.9106
LRPC	2.3563	0.1484	1.7523	2.7274	2.4663	0.1547	1.8062	3.0278

Table 2b—Descriptive statistics for shortrun models

	Metro (4,272 observations)				Nonmetro (36,128 observations)			
Variable	Mean	St. Dev.	Min.	Max.	Mean	St. Dev.	Min.	Max.
Yt/Yt-1	1.0143	0.024	0.866	1.163	1.0158	0.074	0.453	4.097
EmpGth	1.0181	0.035	0.803	1.293	1.0107	0.077	0.127	4.440
NIB	118.02	281.899	0	3532	5.52	5.048	0	55
NXB	34.30	75.040	0	1113	2.36	4.175	0	49
IDEPS (in millions)	4,046	14,081	0	225,109	94	98	0	3,974
XDEPS (in millions)	781	2,452	0	45,721	34	68	0	806
DMA	0.688	0.463	0	1	0.583	0.493	0	1
DNOVO	0.520	0.500	0	1	0.369	0.483	0	1
HHI	0.1779	0.0793	0.0265	0.8199	0.4190	0.2378	0.0737	1
Ratio of bank offices owned out-of-market	0.294	0.287	0	1	0.275	0.348	0	1
Ratio of local bank deposits controlled out-of-market	0.284	0.307	0	1	0.258	0.354	0	1

between employment growth and per capita real income growth, while consistently positive and statistically significant, was substantially higher in metropolitan (0.38) than in nonmetropolitan (0.09) markets.

Model Estimation

We estimated the models 1 through 4 separately for metropolitan markets and for nonmetropolitan counties. Up to 10 metropolitan areas and 10 nonmetropolitan counties were deleted due to missing values of one or more variables, leaving a sample that ranged from 2,260 to 2,264 nonmetropolitan observations and from 259 to 263 metropolitan observations in each period.

There are reasons to expect violations of OLS assumptions in these data sets, especially with respect to multicollinearity and heteroskedasticity. Correlation coefficients are quite high between several pairs of variables. Of concern in the longrun data in both the metropolitan and nonmetropolitan subsamples are correlations between NXB and XTB (ranging from 0.51 to 0.65), and NIB and LPOP (ranging from 0.61 to 0.81), and, in the nonmetropolitan subsamples, NIB and HHI (-0.62 and –0.56), and HHI and LPOP (-0.63 and –0.64). Of particular concern in the shortrun data are the correlations between NIB and IDEPS (0.82 in nonmetropolitan markets and 0.94 in metropolitan markets), NXB and XDEPS (0.90 and 0.93), and DNOVO and DMA (0.65 and 0.70). We therefore tested for multicollinearity using the condition index. For the longrun models, standardizing the data to mean zero and unit variance brought all condition indices below 10, indicating no major problem with statistical dependencies[8], and *F* tests (not reported here) also indicated little impact of collinearity on the statistical significance of coefficients testing our hypotheses. For the shortrun models, only condition indices related to cross-sectional and time-series fixed effects indicate significant statistical dependencies. Thus, we concluded that multicollinearity did not substantially affect the hypothesis tests that are the primary focus of this report.

Both CS and JS found heteroskedasticity related to the size of economies and used weighted least squares to correct it. JS gave the following three reasons why such weighting makes sense econometrically: (1) Measurement errors may be relatively larger for small economies, (2) measurement problems related to interstate commerce are likely to be relatively larger for smaller States, and (3) small economies are more likely to be dominated by specific industries and suffer from industry-specific shocks that would make their growth rates more variable. We, too, found that using weighted least squares substantially improved the fit of our models and, in the interest of brevity, we report only the WLS results.

Given the level of disaggregation of our data, we were also concerned about outliers and influential observations. We tested for influential observations using Cook's D statistic (Cook, 1977). We removed a small number of observations from each regression because Cook's D statistic identified them as influential (Cook, 1977). In addition, a few other observations were deleted as outliers, using the criterion that the observation's regression residual was at least 50 percent larger than the next largest in absolute value.[9]

Division by zero occurred in some of the ratios, which were treated as follows. Where a market had zero locally owned bank offices in the base year, DIB was set to 0.01; where a market had zero remotely owned bank offices in the base year, DXB was set to 0.01.[10] For markets with zero total bank offices, we set XTB equal to 1 under the presumption that such markets are more like those whose banks are controlled outside the local market than those whose banks are locally controlled. In such markets, we set HHI to 1 under the presumption that consumers in these markets have no more choices than in a monopoly market. For markets with zero bank deposits in the base year, DDEP was calculated based on a zero initial value.

[8] Belsley, Kuh, and Welsch (1980) suggest the following relationship between the condition index and multicollinearity: A condition index around 10 indicates that weak dependencies may be starting to affect the regression estimates. A condition index of 30 to 100 indicates moderate to strong collinearity. A condition index larger than 100 indicates that estimates may have a fair amount of numerical error. In this case, the statistical standard error is almost always much greater than the numerical error.

[9] The following markets were influential: Chicago; Los Angeles; New York City; Houston; Sioux Falls, SD; and Summit County, CO. The following counties were outliers: Tunica, MS; Eureka, NV; Somervell, TX; Dillingham, AK.

[10] This procedure has been previously used in a number of banking cost studies in adjusting zero quantities of one or more outputs to avoid undefined values when taking logarithms (see Shaffer, 1993). It modifies the interpretation of the estimated coefficients but does not bias them. Here we use 0.01 rather than a smaller number to avoid the creation of large outlier values in the adjusted ratios.

Longrun Models

The estimates from models 1 and 2 as well as the CS results are shown in table 3 for nonmetropolitan markets and table 4 for metropolitan markets. F-tests for hypotheses tests involving more than one coefficient are shown in table 5.

Employment Growth Regressions (Model 1)

Results from model 1 in tables 3 (nonmetropolitan markets) and 4 (metropolitan markets) and the top panel of table 5 (F-tests for both markets) indicate that initial bank structure and contemporaneous changes in bank structure were statistically significantly related to employment growth in both nonmetropolitan and metropolitan markets in the earlier period and in nonmetropolitan markets in the later period. However, no statistically significant relationship persisted in metropolitan markets in the period of deregulatory activity from 1984 to 1996. Goodness-of-fit statistics (F statistics and adjusted R-squared) reinforced this conclusion, falling much less for nonmetropolitan regressions from the earlier to the later period. For example, the adjusted R-squared fell from 0.243 to 0.209 for the nonmetropolitan regressions and from 0.561 to 0.150 for the metropolitan regressions. Compared with the CS regressions, the goodness-of-fit statistics indicated that with the exception of metropolitan markets in the earlier period, the model explained a significantly lower percentage of variation in employment growth than in per capita income growth.

Initial bank market conditions and employment growth. The hypotheses that longrun average employment growth is independent of initial bank market structure (NIB=NXB=XTB=HHI= 0 and NIB=NXB=0) were rejected for nonmetropolitan markets in both periods and for metropolitan markets in the earlier period, with greater statistical significance for both markets in the earlier period. The initial number of in-market owned bank offices (NIB) was negatively associated with subsequent employment growth at the 0.01 level in the earlier period for both metropolitan and nonmetropolitan markets, indicating that employment grew faster in markets that initially had fewer locally owned bank offices. However, the magnitude and significance of this effect declined in the later period, especially in metropolitan markets. In contrast, the initial number of out-of-market owned bank offices (NXB) was positively and significantly associated with subsequent employment growth only in the earlier nonmetropolitan regressions, but not elsewhere. This finding suggests that remotely owned bank offices, unlike locally owned bank offices, may have provided a stimulus to local employment in nonmetropolitan markets in the 1970s and early 1980s. The hypothesis that these two variables have the same coefficients (NIB=NXB) was rejected only in nonmetropolitan markets in the earlier period.

The coefficient on the initial share of total bank offices owned by out-of-market banks (XTB) was significantly negative in the earlier nonmetropolitan regressions, but significantly positive in the earlier metropolitan regressions and the later nonmetropolitan regressions, weakening that interpretation. The final measure of initial market structure, HHI, is significantly positive in both nonmetropolitan regressions. This result says that employment grew faster in more concentrated nonmetropolitan bank markets throughout the sample period. The magnitude and significance of this linkage were both greater in the more recent period. However, metropolitan markets exhibited the opposite pattern: a negative association in the earlier period, and no significant association in the later period.

The coefficients on NIB and NXB must be interpreted jointly with the initial mix of local versus nonlocal bank offices (XTB) in this model, since XTB represents a nonlinear interaction between NIB and NXB. The coefficients on XTB should be interpreted as the association between employment growth and the share of out-of-market bank offices, holding the total number of bank offices constant. For example, a joint calculation involving the estimated coefficients on NIB, NXB, and XTB indicates that, at the sample mean values of these variables, the point estimate of the subsequent average change in employment growth associated with a change in the initial number of bank offices owned out-of-market (in-market) in nonmetropolitan markets in 1973 was -0.038 (-0.021) percentage points per year, or -3.0 (-1.6) percent of the expected average annual employment growth over the subsequent 13 years. The corresponding effects in nonmetropolitan markets for the 1984-96 period were 0.043 (-0.029) percentage points per year, or 3.4 (-2.3) percent of the expected average annual employment growth over the next 12 years. The magnitude of these estimates for metropolitan markets was smaller, especially in the later period.

Table 3—Nonmetropolitan longrun regression results (t-statistics in parentheses)

Dep. Var	1973-84 EG	1973-84 YG CS	1973-84 YG Aug.	1984-96 EG	1984-96 YG CS	1984-96 YG Aug.
Intercept	0.1016	0.0721	0.04913	0.0293	0.0829	0.0788
	(9.94)*	(11.51)*	(8.26)*	(3.06)*	(13.95)*	(13.58)*
EG	--	--	0.2241	--	--	0.1396
			(18.85)*			(10.95)*
NIB	-3.60E-4	-4.28E-5	3.42E-5	-1.25E-4	2.33E-4	2.51E-4
	(-4.42)*	(-0.86)	(0.74)	(-1.95)***	(5.85)*	(6.45)*
NXB	3.49E-4	1.26E-4	5.40E-5	-1.01E-4	2.09E-4	2.23E-4
	(2.91)*	(1.46)	(0.67)	(-1.31)	(4.36)*	(4.77)*
XTB	**-0.00502**	-0.0035	**-0.0025**	0.00534	0.0011	3.40E-4
	(-3.02)*	(-3.21)*	**(-2.43)****	(3.62)*	(1.18)	(0.38)
HHI	0.00452	-0.0008	-0.00191	0.00717	-0.0011	-0.00206
	(2.09)**	(-0.64)	(-1.54)	(3.31)*	(-0.79)	(-1.57)
DIB	**6.72E-4**	8.64E-5	**-6.38E-5**	**0.00152**	5.74E-4	**3.63E-4**
	(8.42)*	(1.76)**	**(-1.38)**	**(7.44)***	(4.54)*	**(2.91)***
DXB	**5.98E-5**	2.65E-5	**1.33E-5**	**4.79E-5**	2.06E-5	**1.01E-5**
	(3.04)*	(2.19)**	**(1.18)**	**(5.70)***	(2.52)**	**(1.26)**
DDEP	-0.00738	-7.34E-4	8.92E-4	-0.00352	2.20E-4	7.10E-4
	(-3.66)*	(-0.59)	(0.77)	(-2.76)*	(0.28)	(0.92)
DPC	-2.83E-4	0.00125	0.00133	-7.09E-4	-9.77E-5	1.30E-6
	(-0.58)	(4.14)*	(4.72)*	(-4.77)*	(-1.06)	(0.01)
LEDU	0.0142	0.0041	9.03E-4	0.00803	0.0039	0.00281
	(17.04)*	(8.01)*	(1.80)a	(9.70)*	(7.65)*	(5.50)*
LPOP	-0.00270	0.0025	0.0031	0.00118	-9.91E-4	-0.00116
	(-3.63)*	(5.43)*	(7.34)*	(1.65)***	(-2.23)**	(-2.67)*
LRPCI	-0.00944	-0.0367	-0.0347	-0.00305	-0.0234	-0.0230
	(-4.97)*	(-31.44)*	(-31.79)*	(-1.59)	(-19.69)*	(-19.82)*
FM	-0.0101	-0.00659	-0.00426	-0.00596	-0.0022	-0.00140
	(-9.46)*	(-10.05)*	(-6.86)*	(-6.01)*	(-3.63)*	(-2.32)**
MI	0.00819	0.00296	0.00112	-0.01162	-0.0066	-0.00502
	(8.19)*	(3.64)*	(1.48)	(-10.47)*	(-9.65)*	(-7.30)*
n	2260	2265	2265	2264	2265	2265
Adj. R^2	0.243	0.522	0.587	0.209	0.243	0.281
F value	56.85*	191.45*	231.12*	46.97*	57.02*	64.31*

Two-tailed significance levels: *0.01 (t > 2.550); **0.05 (2.550 > t > 1.960); ***0.10 (1.960 > t > 1.645).
Bold indicates possible support for an employment channel of bank structure influence on economic growth.

Table 4—Metropolitan longrun regression results (t-statistics in parentheses)

Dep. Var	1973-84 EG	1973-84 YG CS	1973-84 YG Aug.	1984-96 EG	1984-96 YG CS	1984-96 YG Aug.
Intercept	0.1740	0.0514	0.0250	0.1023	0.0442	0.0352
	(7.70)*	(4.58)*	(2.27)**	(4.72)*	(4.36)*	(3.42)*
EG	--	--	0.1665	--	--	0.1050
			(6.85)*			(3.44)*
NIB	-1.89E-5	-1.54E-6	2.24E-6	-8.23E-7	4.21E-6	4.66E-6
	(-4.70)*	(-0.30)	(0.48)	(-0.22)	(3.09)*	(3.47)*
NXB	1.94E-6	8.36E-6	7.78E-6	-7.22E-6	-9.58E-6	-8.82E-4
	(0.06)	(0.48)	(0.48)	(-0.54)	(-1.53)	(-1.44)
XTB	0.01778	-0.0037	-0.0066	3.71E-4	-0.0044	-0.00438
	(3.48)*	(-1.56)	(-2.99)*	(0.09)	(-2.29)**	(-2.33)**
HHI	**-0.02054**	-0.0112	**-0.00827**	0.00728	0.0133	0.0119
	(-2.31)**	(-2.81)*	**(-2.25)****	(0.78)	(3.09)*	(2.81)*
DIB	-1.54E-4	-6.1E-5	-3.61E-5	-8.95E-5	1.11E-5	2.62E-5
	(-3.32)*	(-2.63)*	(-1.66)***	(-1.27)	(0.56)	(1.30)
DXB	2.57E-5	4.53E-7	-4.42E-6	-5.43E-7	-4.27E-6	-4.70E-6
	(5.16)*	(0.18)	(-1.86)***	(-0.12)	(-3.08)*	(-3.45)*
DDEP	0.00271	0.0056	0.00524	-0.00412	-0.0028	**-0.00221**
	(0.50)	(2.43)**	(2.50)**	(-1.34)	(-1.85)***	**(-1.50)**
DPC	0.00732	0.0014	1.54E-4	-9.51E-4	-0.0004	-2.83E-4
	(5.60)*	(2.36)**	(0.26)	(-2.29)**	(-2.29)**	(-1.62)
LEDU	0.0401	0.0088	0.00219	0.0157	0.0032	0.00150
	(13.01)*	(6.62)*	(1.42)	(4.96)*	(2.15)**	(0.97)
LPOP	0.00175	0.0017	0.00137	0.00153	0.0011	8.99E-4
	(1.56)	(2.78)*	(2.39)**	(1.37)	(2.35)**	(1.88)***
LRPCI	-0.0420	-0.0187	-0.0122	-0.0260	-0.0150	-0.0126
	(-5.48)*	(-5.70)*	(-3.86)*	(-4.09)*	(-4.46)*	(-3.75)*
n	263	260	260	261	264	264
Adj. R^2	0.561	0.311	0.418	0.150	0.306	0.334
F value	31.59*	11.61*	16.53*	5.18*	11.53*	12.01*

Two-tailed significance levels: *0.01 (t > 2.550); **0.05 percent (2.550 > t > 1.960); ***0.10 (1.960 > t > 1.645).
Bold indicates possible support for an employment channel of bank structure influence on economic growth.

Table 5—F-tests on longrun effects of banking structure

Period	Sample	Hypothesis: NIB=NXB=XTB=HHI=0	NIB=NXB=0	NIB=NXB	DIB=DXB
Employment growth regressions					
1973-84	Nonmetro	12.10*	15.26*	26.05*	57.32*
	Metro	10.11*	11.55*	0.38	15.17*
1984-96	Nonmetro	5.62*	2.56***	0.06	53.40*
	Metro	0.28	0.17	0.21	1.67
Real per capita income growth regressions (CS)					
1973-84	Nonmetro	3.24**	1.58	3.14***	1.46
	Metro	3.77*	0.17	0.30	6.95*
1984-96	Nonmetro	13.56*	24.68*	0.16	20.48*
	Metro	8.96*	4.95*	4.09**	0.61
Augmented income growth regressions					
1973-84	Nonmetro	1.19	0.45	0.05	2.73***
	Metro	1.81	0.22	0.11	2.07
1984-96	Nonmetro	21.02*	29.82*	0.22	8.54*
	Metro	6.12*	6.11*	4.08**	2.43

Significant at the following levels: *0.01, **0.05, ***0.10.

Contemporaneous changes. To this point, we have examined results relating initial conditions to subsequent longrun average growth. Now, we turn to contemporaneous associations between bank ownership structure and deposit control and growth. The model contains two types of contemporaneous measures. The first is the ratio of bank offices owned in-market (DIB) or out-of-market (DXB) at the end of the period to that at the beginning of the period. The second is the change in the local deposit market share controlled by banks owned out-of-market (DDEP).

The contemporaneous change in the deposit market share of nonlocally owned bank offices (DDEP) is negatively and significantly associated with employment growth in both nonmetropolitan regressions, in contrast to the coefficient on NXB. This says that, for a given number of nonmetropolitan banks, subsequent employment grew more slowly where remotely owned bank offices had a growing market share. Because this variable is measured contemporaneously with employment growth, its coefficient likely reflects a pattern that locally owned bank offices expand relatively more than remotely owned bank offices in response to a growing labor market. This might be due to an informational advantage of locally owned bank offices in identifying growth conditions, or to more limited flexibility of remotely owned bank offices with respect to conditions in any one market (given that multi-market banks typically must balance competing opportunities for expansion).

Consistent with the notion that local banks are more responsive to job growth, the contemporaneous growth rate of locally owned bank offices (DIB) was positively and significantly associated with employment growth in both nonmetropolitan regressions, and its coefficient was an order of magnitude larger than that of the contemporaneous growth rate of remotely owned bank offices (DXB) in those regressions. F-tests reported in table 5 rejected the equality of the coefficients on DIB and DXB in the nonmetropolitan markets in both periods and, for metropolitan markets, in the earlier period but not in the later period. These results suggest that locally owned bank offices responded more strongly to local job growth, or perhaps contributed more to it, than did remotely owned bank offices in nonmetropolitan markets.

The point estimate of the coefficient on DIB in the later period is more than twice that in the earlier period for the nonmetropolitan sample, suggesting that either locally owned bank offices have responded more flexibly to changing market conditions in recent years or changes in locally owned banking structure had a greater impact on contemporaneous job growth in recent years.

In the metropolitan regressions, DIB showed a negative effect in 1973-84 but an insignificant effect in the later period. An F-test rejected the equality of the coefficients on DIB and DXB for the earlier metropolitan data but not the later (table 5). DIB had a significantly negative coefficient in the earlier metropolitan regression but not in the later metropolitan regression. DXB was significantly positive in both nonmetropolitan regressions and in the earlier metropolitan regression. DDEP was not significant in either metropolitan regression.

Control variables. DPC exhibited a significantly negative coefficient in both later regressions. However, its coefficient was significantly positive for metropolitan markets from 1973-84, suggesting that those markets were formerly dependent on local funding for a significant measure of their job growth. These results do not parallel the coefficients found in income growth regressions by CS, and thus do not support the hypothesis of an employment channel for the linkage between banking structure and income growth.

LEDU showed a significantly positive coefficient in every employment regression, suggesting that job creation is associated with greater human capital. This result is consistent with the findings of Henderson and McDaniel (2000) for rural counties, and with the findings of O hUallachain and Satterthwaite (1992) for urban manufacturing. CS found the same pattern of coefficients for income growth regressions. However, the point estimate was smaller in the later period than in the earlier period for both metropolitan and nonmetropolitan samples, and this difference was significant at the 0.0001 level ($F = 44.4$ and 38.1, respectively).

This trend suggests a declining marginal social productivity of education as measured by aggregate job creation, a potentially important question for future research. CS found similar declines of the marginal impact of education over time in income growth regressions for metropolitan markets and farm-dependent counties. Negative trends in private returns to education have also been found for other countries (see Psacharopoulos, 1989; Lam and Levison, 1991) but contrast with previous findings for the U.S. over this time period (see, for example, Goldin, 1986; Kane, 1994; Topel, 1997). Among the possible explanations besides simple diminishing marginal returns are an increasing need for postgraduate education in many jobs, a diminished effectiveness of a college education in screening workers once a higher percentage of the work force attains such education, or a decline in the average quality of education obtained through the first 4 years of a college education. Jaeger and Page (1996) found that estimated returns to education differed as a function of years completed versus degrees earned, because many students do not complete their degrees in the standard number of years; therefore, if the average time to finish a degree has lengthened in recent years, this factor also might contribute to our results. However, any such explanation must be reconciled with previous findings of increasing private returns to education in recent years. We defer all such questions to future research as they lie outside the primary focus of this study.

LPOP shows mixed effects: significantly negative in the early nonmetropolitan regression, significantly positive in the later nonmetropolitan regressions, and not significant in either metropolitan regression. The implication is that job growth was most rapid in the less populous nonmetropolitan counties during 1973-84 but in the more populous nonmetropolitan counties in the later period. The later result is consistent with the significantly positive effect of total labor force on rural employment growth found by Henderson and McDaniel (2000), and is also consistent with economies of urbanization. The nonmetropolitan results are exactly opposite those found in income growth regressions by CS, who moreover found significantly positive coefficients in both longrun metropolitan income growth regressions.

LRPC takes negative coefficients in each regression, significantly so in all but the later nonmetropolitan regressions. This pattern is consistent with the negative wage coefficient found by Henderson and McDaniel (2000) and generally consistent with CS. As in CS's income growth regressions, farming-dependent nonmetropolitan counties (FM) exhibited slower job growth. Mining-dependent nonmetropolitan counties (MI) likewise showed slower job growth in the later period but more rapid job growth in the earlier period.

Augmented Income Regressions (Model 2)

Tables 3 and 4 also report estimates of equation 2, with and without the employment growth variable

(EG). These regressions test whether the structure-income growth linkage is robust to the inclusion of employment changes, and whether employment growth exhibits an association with income growth that is independent of local banking structure. As noted above, if employment growth is a primary channel through which banking structure influences income growth, then the inclusion of contemporaneous employment growth rates ($EG_{0,T}$) should dilute the measured effect of ex ante structural variables in the regression.

EG itself exhibits consistently positive coefficients, with t-statistics ranging from 3.44 to 18.85, though the magnitude and significance of this coefficient were smaller in the later regressions. This result demonstrates that job growth is associated with growth in per capita income, though the direction of causality is unclear. More jobs in a community may increase per capita income to the extent that they are associated with an increase in the labor force participation rate. In addition, more jobs may enhance the productivity of a typical worker if there are agglomeration economies in production.[11] Under either of these mechanisms, causality would flow from jobs to income. Alternatively, both employment and income may respond simultaneously to changes in macroeconomic activity, rising in expansions and falling in recessions. Such explanations need not be mutually exclusive nor exhaustive. We do not pursue these issues here.

The inclusion of EG does not materially alter the measured association between any of the initial banking structure variables and per capita income growth. Therefore, these results do not provide evidence of a significant employment channel by which initial banking structure affects income growth. By contrast, several contemporaneous associations between bank ownership structure and per capita income growth are affected by the inclusion of EG. In particular, the inclusion of EG results in a loss of significance of DIB in the earlier nonmetropolitan regressions, DXB in both nonmetropolitan regressions, DDEP in the later metropolitan regression, and DPC in both metropolitan regressions. Conversely, XTB and DXB both gain significance in the earlier metropolitan regression with the inclusion of EG.

[11] Previous studies have found agglomeration economies in costs; see Eberts and McMillen (1998) for a review.

Evidence regarding the employment growth channel. We looked for evidence of employment growth as a channel through which bank structure may spur per capita income growth. Our strategy was to look for two types of evidence consistent with this hypothesis. For model 1, such evidence includes statistically significant coefficients consistent in sign with those found by CS. For model 2, such evidence consists of coefficients that are lower in absolute magnitude and/or less significant statistically than those found by CS.

In both instances, we found at best weak support for an employment growth channel. For metropolitan markets (table 4), only the coefficients on initial deposit market concentration (HHI) and on contemporaneous growth in the numbers of in-market owned bank offices (DIB) in the earlier period were consistent with such a channel in both models. In addition, the coefficient on the contemporaneous change in the share of deposits held at out-of-market owned bank offices (DDEP) was consistent with an employment growth channel, but only in model 2 and only for the later period. For nonmetropolitan markets (table 3), the evidence was a little stronger. For both periods, the coefficients on contemporaneous growth in both DIB and in the numbers of out-of-market owned bank offices (DXB) were consistent with such a channel in both models. In addition, the coefficient on initial ratio of out-of-market owned bank offices to total bank offices (XTB) was consistent with an employment growth channel in both models for the earlier period.

Two results provide countervailing evidence to the hypothesis that employment growth serves as a major channel through which financial structure affects income growth. First, no statistically significant relationship persisted between employment growth and either initial bank structure or contemporaneous changes in bank structure in metropolitan markets in the 1984 to 1996 period, despite a fairly robust statistical relationship between the regressors and per capita income growth. Second, in nonmetropolitan markets in the later period, the relationship between NIB and employment growth was statistically significant and of the opposite sign as their relationship to real per capita income growth. That is, the initial number of in-market owned bank offices at the outset of the later period was negatively related to subsequent employment growth but positively related to subsequent per capita income growth in nonmetropolitan counties.

Shortrun Models

Table 6 presents regression results for models 3 and 4, again weighted by total personal income in the local market. The CS results are also included for comparison. As is generally the case for the longrun models, the goodness-of-fit statistics were considerably weaker for the employment growth regressions than for the real per capita income growth regressions. Consistent with this result was the general insignificance of regressors related to bank structure, the exceptions being market concentration (HHI) in both metropolitan and nonmetropolitan markets and the amounts of deposits held by out-of-market owned banks (XDEPS) in metropolitan markets. Note that the coefficient on HHI was negative and significant in metropolitan markets and positive and significant in nonmetropolitan markets.

Evidence Regarding the Employment Growth Channel

Also consistent with the longrun results was the generally weak evidence that employment growth may serve as a channel through which bank structure affects income growth. The evidence we are seeking from models 3 and 4 is analogous to that sought from models 1 and 2. This evidence includes statistically significant coefficients consistent in sign with those found by CS for model 3 and coefficients that are lower in absolute magnitude and/or less significant statistically than those found by CS for model 4. The only evidence consistent with our hypothesis involves the coefficients on HHI in both the employment growth regression (model 3) and in the augmented real per capita income growth regression (model 4) for metropolitan markets. In the employment growth regression, this coefficient was statistically significant and had the same sign as in the CS regres-

Table 6—Shortrun regression results (t-statistics in parentheses)

	Metro			Nonmetro		
Dep. Var	EG	YG (CS)	YG (augmented)	EG	YG (CS)	YG (augmented)
Intercept	1.01824 (446.29)*	1.00781 (663.11)*	0.79141 (78.97)*	1.00384 (460.32)*	1.00699 (651.10)*	0.90230 (224.92)*
EG	--	--	0.21252 (21.82)*	--	--	0.10429 (28.23)*
NIB	-1.9E-6 (-0.72)	2.5E-7 (0.14)	6.6E-7 (0.39)	7.7E-5 (0.92)	1.8E-4 (3.12)*	1.8E-4 (3.01)*
NXB	9.4E-6 (1.29)	-9.0E-6 (-1.84)***	-1.1E-5 (-2.38)**	1.9E-4 (1.46)	2.1E-4 (2.24)**	**1.9E-4** **(2.05)****
IDEPS	-3.0E-8 (-0.57)	6.6E-8 (1.89)***	7.3E-8 (2.18)**	-2.2E-6 (-0.61)	-7.5E-6 (-2.94)*	**-7.3E-6** **(-2.89)***
XDEPS	-4.1E-7 (-2.3)**	-1.3E-7 (-1.06)	-4.1E-8 (-0.36)	-3.7E-8 (-0.01)	-1.5E-5 (-2.95)*	-1.5E-5 (2.98)*
DNOVO	**0.0076** **(5.19)***	0.0020 (2.06)**	**4.0E-4** **(0.43)**	2.2E-4 (0.19)	0.0014 (1.70)***	0.0014 (1.69)***
DMA	**0.0031** **(2.00)****	0.0102 (9.78)*	**0.009**5 **(9.64)***	**0.0036** **(3.09)***	0.0074 (8.92)*	**0.0070** **(8.55)***
HHI	**-0.0309** **(-4.01)***	-0.0120 (-2.34)**	**-0.0054** **(-1.12)**	0.0063 (2.59)*	-0.0024 (-1.37)	-0.0030 (-1.77)***
n	4,272	4,272	4,272	36,128	36,128	36,128
Adj. R^2	0.3415	0.5705	0.6141	0.0434	0.1405	0.1591
F value	33.10*	83.22*	98.11*	24.44*	85.38*	97.26*

Two-tailed significance levels: *0.01 (t > 2.550), **0.05 (2.550 > t > 1.960), ***0.10 (1.960 > t > 1.645).
Bold indicates possible support for an employment channel of bank structure influence on economic growth.

sion. In the augmented income growth regression, this coefficient had a smaller absolute magnitude and lost its statistical significance. In contrast, the coefficient on HHI in nonmetropolitan markets displayed the opposite behavior. In the CS results, this coefficient was negative and statistically insignificant. However, in the model 3 results it was positive and statistically significant, and in the model 4 results it was negative and statistically significant. Also note, that while the coefficients on NXB and IDEPS were smaller in absolute value in the nonmetropolitan regression of model 4, the evidence from model 3 was not consistent with our hypothesis.

In contrast to the results with respect to bank structure, the shortrun results provide stronger evidence that increased employment growth has been associated with geographic deregulation of bank activity. This evidence is also consistent with an employment growth channel for the relationship between geographic deregulation and economic growth documented by JS at the State level and by CS at the local market level. Evidence of such a link can be seen in the magnitudes and signs of the coefficients on the binary variables related to the two major stages of geographic deregulation (DMA and DNOVO).

As is the case for per capita income growth, there was a statistically and economically significant relationship between employment growth and geographic deregulation in both metropolitan and nonmetropolitan markets. The first stage of geographic deregulation, allowing out-of-market entry through mergers and acquisitions of existing banks, was associated with a 0.31-percentage-point increase in the expected annual rate of employment growth in metropolitan markets and a 0.36-percentage-point increase in nonmetropolitan markets. Given the average annual employment growth rates for the sample period of 1.81 percent in metropolitan markets and 1.07 percent in nonmetropolitan markets, these results suggest an average increase in expected employment growth of 17 percent in metropolitan markets and 34 percent in nonmetropolitan markets. The second stage of geographic deregulation, allowing out-of-market entry through de novo branching, was associated with a 0.76-percentage-point increase in annual employment growth rates in metropolitan areas, but no increase in nonmetropolitan areas. Thus, the results from both stages of geographic deregulation suggest a total average increase of 59 percent of the expected rate of employment growth in metropolitan markets and 34 percent in nonmetropolitan markets. For comparison, CS found that geographic liberalization was associated with an average increase of 59 to 87 percent of the expected real per capita income growth in metropolitan markets and of 28 to 53 percent in nonmetropolitan markets. Thus, our estimate of the proportional relationship between geographic deregulation and employment growth is in the lower range of that found by CS between geographic deregulation and economic growth.

The augmented income regressions (model 4) also yielded evidence consistent with an employment growth linkage between geographic deregulation of banking and real per capita income growth rates. In both metropolitan and nonmetropolitan markets, the coefficients on DMA fell between 5 and 7 percent in absolute value

Table 7—F-tests on shortrun effects of banking structure

Sample	*Hypothesis:* NIB=NXB=IDEPS=XDEPS=HHI=0	NIB=NXB=0	IDEPS=XDEPS=0	NIB=NXB	IDEPS=XDEPS
Employment growth regressions					
Nonmetro	2.46*	1.41	0.18	0.58	0.07
Metro	7.09*	0.96	2.95**	1.92	3.79**
Real per capita income growth regressions (CS)					
Nonmetro	5.05*	6.98*	9.01*	0.06	1.78
Metro	23.23*	1.71	2.11	2.87***	2.22
Augmented income growth regressions					
Nonmetro	5.45*	6.29*	8.93*	0.02	1.92
Metro	22.28*	2.82***	2.39***	5.07**	0.84

Significant at the following levels: *0.01, **0.05, ***0.10.

compared with the CS results while retaining their statistical significance. In addition, the coefficient on DNOVO fell in absolute value and lost its statistical significance in the metropolitan regression.

Besides providing plausible evidence of an employment channel linking geographic deregulation of banking to per capita income growth, results from models 3 and 4 also suggest the importance of market contestability and of the role of *de novo* banks in small business lending. Both CS and Berger, Bonime, Goldberg, and White documented differences in the rates of entry after deregulation in metropolitan and nonmetropolitan markets. Not only was out-of-market entry slower and less frequent after deregulation in nonmetropolitan markets, but many nonmetropolitan markets were simply too small to attract entry from enough banking firms to create competitive market conditions. Results from our shortrun models suggest the plausible threat of entry after the *de novo* stage of deregulation coupled with a market size barrier to actual entry as measured by NXB and XDEPS may influence the relationships between banking and economic and/or employment growth. The likelihood of *de novo* entry may be particularly important for employment growth if *de novo* branches behave like *de novo* banks, which lend disproportionately to the small business sector, an important engine of job growth.

Robustness Issues

Results from our investigation consist of two types of conclusions. First, statistically significant associations appear to exist between bank structure (model 1) or geographic deregulation (model 3) and employment growth rates. Second, some evidence points to employment growth as a possible channel through which bank structure and deregulation may spur local economic growth as measured by real per capita income. However, the empirical models in this report are susceptible to several criticisms related to spurious causality or omitted variables.

The possibility of reverse causality is usually addressed by considering initial as opposed to contemporaneous independent variables in the longrun context or lagged independent variables in the shortrun context. Thus in our longrun models, the interpretation of relationships related to variables measured at the start of the long sample periods is less ambiguous. The possibility of reverse causality between these variables (NIB, NXB, XTB, and HHI) and employment or income growth rates (i.e., the possibility that subsequent income growth rates influence the initial banking structure) is remote. Although banks, like other businesses, have a financial incentive to try to predict and adapt to future market conditions, accurate forecasts are very difficult and rarely attained, particularly over horizons in excess of 10 years as measured by our growth variables. Moreover, the economic growth rates exhibit virtually no persistence from one decade to another for the average market in our sample. The Pearson correlation coefficients between the growth rate of income over 1973-84 and that over 1984-96 are not significantly different from zero and are actually slightly negative: -0.021 and -0.101 for the nonmetropolitan and metropolitan samples, respectively.[12] Thus, simple extrapolation from historical economic growth rates would not have permitted banks to foresee accurately the future growth rates in the average local U.S. banking market. Nonetheless, the causal interpretation of results related to contemporaneous changes in bank structure remains ambiguous.

The ambiguity of results from shortrun models 3 and 4 related to geographic deregulation relates primarily to the possibility of omitted variables. Further light may be shed by controlling for plausible contemporaneous changes or business cycle effects. In this context, JS presented evidence that geographic deregulation did not coincide with growth-enhancing policy changes at the State level and that States tended to liberalize at the trough of a recession. This evidence is relevant to the current research, as many important macro policies are determined at the State level. Unfortunately, similar information for local growth policies is not readily available. To control for business cycle effects, we reestimated models 3 and 4 with three lags of the growth rate in real per capita income. These results (table 8) reinforce the link between the first stage of deregulation (DMA) and both employment and per capita income growth in nonmetropolitan markets, leaving support for the employment growth channel intact in this context. The waters are muddied, however, with respect to metropolitan markets. The link between *de novo* deregulation and employment growth remained robust, but the evidence in support of an employment growth channel was undermined by the loss of statistical significance of the coefficient on DNOVO in the income growth models.

A further relevant point is that growth in *per capita* income does not necessarily indicate overall market growth or an attractive market for bank entry; it is quite possible to experience growing per capita income even in a market with declining population and declining aggregate economic activity. It is also possible for *per capita* income to fall in a market with increasing population and increasing economic activity. Finally, as described above, changes in bank structure over the sample period are controlled as separate regressors that should capture any response by the banking industry to local market conditions.

[12] The corresponding values for the growth rates in employment are somewhat higher at 0.268 and 0.508, respectively.

Table 8—Estimates from shortrun models with lagged real per capita income growth, 1982-96

	Metro			Nonmetro		
Dep. Var	EG	YG (CS)	YG (augmented)	EG	YG (CS)	YG (augmented)
Intercept	0.5691 (16.71)*	0.83277 (37.08)*	0.7336 (32.64)*	0.8331 (52.12)*	1.2301 (116.86)*	1.1444 (105.62)*
EG	--	--	0.1743 (16.60)*	--	--	0.1028 (28.06)*
NIB	-9.09E-8 (-0.03)	3.75E-7 (0.20)	3.91E-7 (0.22)	1.45E-4 (1.63)	2.42E-4 (4.13)*	**2.27E-4** **(3.92)***
NXB	1.39E-5 (1.65)***	-9.37E-6 (-1.68)***	-1.18E-5 (-2.19)**	1.43E-4 (1.02)	2.49E-4 (2.70)*	**2.34E-4** **(2.57)***
IDEPS	-1.22E-7 (-2.13)**	3.60E-8 (0.95)	5.73E-8 (1.57)	-4.93E-6 (-1.31)	-7.29E-6 (-2.93)*	**-6.78E-6** **(-2.76)***
XDEPS	-3.36E-7 (-1.62)	-6.12E-8 (-0.45)	-2.64E-9 (-0.02)	1.33E-6 (0.17)	-1.87E-5 (-3.69)*	-1.88E-5 (-3.76)*
DNOVO	0.0061 (3.98)*	4.16E-4 (0.41)	-6.48E-4 (-0.66)	-0.0018 (-1.50)	-5.52E-4 (-0.68)	-3.62E-4 (-0.45)
DMA	-5.34E-4 (-0.31)	0.0099 (8.77)*	0.0100 (9.18)*	**0.0038** **(2.96)***	0.0133 (15.73)*	**0.0129** **(15.46)***
HHI	-0.0247 (-3.04)*	-0.0095 (-1.78)***	-0.0052 (-1.01)	0.0067 (2.58)*	-3.45E-4 (-0.20)	-0.0010 (-0.61)
YG(t-1)	0.3261 (13.38)*	0.2001 (12.44)*	0.1432 (9.02)*	0.0790 (9.60)*	-0.2251 (-41.54)*	-0.2332 (-43.51)*
YG(t-2)	0.0917 (3.86)*	0.0289 (1.85)***	0.0129 (0.85)	0.0649 (8.31)*	-0.0389 (-7.56)*	-0.0456 (-8.96)*
YG(t-3)	0.0235 (1.04)	-0.0586 (-3.94)*	-0.0627 (-4.37)*	0.0230 (3.11)*	0.0360 (7.39)*	0.0336 (6.99)*
n	3,738	3,738	3,738	31,612	31,612	31,612
Adj. R^2	0.3731	0.5911	0.6195	0.0457	0.1915	0.2112
F value	32.77*	78.16*	86.71*	22.33*	106.47*	118.54*

t-statistics in parentheses. Two-tailed significance levels: *0.01 percent (t > 2.550), **0.05 (2.550 > t > 1.960), ***0.10 percent (1.960 > t > 1.645).
Bold indicates possible support for an employment channel of bank structure influence on economic growth.

Conclusion and Policy Implications

This report has explored the empirical linkages between growth rates in total local employment and the structure of local bank markets (ownership, deposit control, and concentration) and their level of geographic deregulation. These linkages are of interest for two reasons. First, local communities, their leaders, and policymakers are all sensitive to employment growth and contraction. These constituencies want to know if nonlocal ownership of bank offices or nonlocal control of the deposit base (and, therefore, geographic deregulation of banking) is associated with faster or slower rates of employment growth. Second, the identification of plausible mechanisms by which banking structure may influence the growth rates of per capita income might further strengthen our confidence in the causal nature of the empirical association between banking structure and economic growth found by previous research.

Our findings shed new light on the macroeconomic impact of structural and regulatory changes in banking. Several significant differences were found between metropolitan and nonmetropolitan markets, and, in the longrun context, between the earlier (1973-84) and later (1984-96) periods studied. Consistent with previous research, we found no evidence that, on average, geographic deregulation, nonlocal bank office ownership, or nonlocal deposit control adversely impacted nonmetropolitan income or employment growth rates.

The longrun regressions provided evidence that the growth rates of local employment tend to be associated with the initial numbers of bank offices, initial relative market shares of banks, and the initial ownership structure of banks (local versus out-of-market). Employment grew faster in markets that initially had fewer locally owned bank offices, though the significance of this linkage waned into the 1990s. By contrast, employment in nonmetropolitan markets grew faster in markets that had more remotely owned bank offices in 1973, but this relationship did not persist into the 1990s. Overall, employment grew faster in markets with initially concentrated banking structures from 1973-84, but this linkage was dissipated for metropolitan markets after 1984.

Shortrun regressions generally failed to find a significant association between bank ownership and deposit control variables and employment growth in either metropolitan or nonmetropolitan markets. The negative coefficient on out-of-market deposit control in metropolitan markets was the exception to this pattern, but the magnitude suggests little economic significance. Mirroring the longrun results, employment grew faster in more concentrated nonmetropolitan markets but slower in more concentrated metropolitan markets. The shortrun regressions also showed a positive and economically significant association between geographic deregulation and employment growth. However, this association differed between metropolitan and nonmetropolitan markets. In nonmetropolitan markets, employment growth was more sensitive to the first stage of geographic deregulation (allowing nonlocal entry through mergers and acquisitions of existing banks), while the second stage of geographic deregulation (allowing nonlocal entry through *de novo* branching) was more important in metropolitan markets.

The strongest evidence for an employment growth channel linking banking to economic growth was from the shortrun model results related to geographic deregulation. Overall, however, this linkage appeared to explain a small portion of the relationship between banking and economic growth. Otherwise, we found only weak evidence in support of an employment growth channel linking bank structure to subsequent economic growth. In the longrun model, relationships between the structural regressors and employment growth were in some cases quite different from linkages reported by CS between the same regressors and income growth. In addition, the inclusion of contemporaneous employment growth rates did not substantially change the linkages between banking structure and income growth rates. These findings suggest that job creation, while responsive to banking structure and important in its own right, is not consistently a major channel by which banking structure stimulates income growth. A corollary is that the macroeconomic benefits of banking structure accrue primarily to those already working, rather than to new workers. Thus, the stimulus to growth provided by financial structure has diverse distributional effects, a previously overlooked point that may warrant further study. These findings also extend the literature on empirical determinants of employment growth by identifying the structure of financial intermediation as a relevant set of previously overlooked factors.

We addressed the possibility of reverse causality in several ways. In the longrun models, the measurement of subsequent economic growth over long time horizons (more than 10 years), besides smoothing out high-frequency intertemporal noise and mitigating the impact of outlier years, reduced the likelihood that the estimated coefficients could reflect reverse causality. Intertemporal persistence in income growth rates was virtually zero in our sample (indeed, the Pearson correlation coefficients were negative). Moreover, growth in *per capita* income need not imply aggregate market growth or financial incentives for banks' entry.

Finally, controlling for the contemporaneous change in banking structure should capture any endogeneity in structure and leave the coefficient on initial structure to reflect exogenous effects alone. We also reestimated the shortrun models with lags of real per capita income growth rates to control for business cycle effects. General conclusions about the importance of geographic deregulation for employment and per capita income growth remained robust. However, the evidence that employment growth may serve as a channel linking deregulation to income growth in metropolitan markets lost its statistical significance.

References

Amel, Dean F., no date, "State Laws Affecting the Geographic Expansion of Commercial Banks." Unpublished paper. Board of Governors of the Federal Reserve System. Washington, DC.

Amel, Dean F., and J. Nellie Liang, 1992, "The relationship between entry into banking markets and changes in legal restriction on entry," *The Antitrust Bulletin* 37, pp. 631-49.

Amel, Dean F., and J. Nellie Liang, 1997, "Determinants of Entry and Profits in Local Banking Markets," *Review of Industrial Organization*, Vol. 12, No. 1, pp. 59-78.

Avery, Robert B., Raphael W. Bostic, Paul S. Calem, and Glenn B. Canner, 1999, "Consolidation and Bank Branching Patterns," *Journal of Banking and Finance* 23 (2-4), February, 497-532.

Belsley, D. A., E. Kuh, and R. E. Welsch, 1980, *Regression Diagnostics* (New York: John Wiley and Sons, Inc.).

Berger, Allen N., Seth D. Bonime, Lawrence G. Goldberg, Lawrence J. White, 2000, "The Dynamics of Market Entry: The Effects of Mergers and Acquisitions on De Novo Entry and Small Business Lending in the Banking Industry," working paper, Board of Governors of the Federal Reserve System. Washington, DC.

Berger, Allen N., Anthony Saunders, J. Scalise, and Gregory F. Udell, 1998, "The Effects of Bank Mergers and Acquisitions on Small Business Lending," *Journal of Financial Economics* 50, 187-229.

Berger, Allen N., and Gregory F. Udell, 1996, "Universal Banking and the Future of Small Business Lending," in *Universal Banking: Financial System Design Reconsidered*, ed. by Anthony Saunders and Ingo Walter (Burr Ridge, Ill.: Irwin Publishing).

Broecker, Thorsten, 1990, "Credit-Worthiness Tests and Interbank Competition," *Econometrica* 58 (2), March, 429-452.

Cetorelli, Nicola, and Michele Gambera, 2001, "Banking Market Structure, Financial Dependence and Growth: International Evidence from Industry Data," *Journal of Finance* 56 (2), April, 617-648.

Cole, Rebel A., Lawrence G. Goldberg, and Lawrence J. White, 1999, "Cookie-cutter versus character: The micro structure of small business lending by large and small banks," in Jackson L. Blanton, Alicia Williams, and Sherrie L.W. Rhine, eds., *Business Access to Capital and Credit: A Federal Reserve System Research Conference*. Federal Reserve Bank of Chicago, pp. 362-89.

Collender, Robert N., and Sherrill Shaffer, May 2000, *Local Bank Office Ownership, Deposit Control, Market Structure, and Economic Growth*, U.S. Department of Agriculture, Economic Research Service, Technical Bulletin Number 1886.

Cook, R. D., 1977, "Detection of Influential Observations in Linear Regression," *Technometrics* 19, 15-18.

Cook, Peggy J., and Karen L. Mizer, 1994, *The Revised ERS County Typology*, U.S. Department of Agriculture, Economic Research Service, Rural Development Research Report Number 89.

Davis, Steven, and John Haltiwanger, 1992, "Gross Job Creation, Gross Job Destruction, and Employment Reallocation," *Quarterly Journal of Economics* 107 (3), August, 819-863.

Davis, Steven, John Haltiwanger, and Scott Schuh, 1996, "Small Business and Job Creation: Dissecting the Myth and Reassessing Facts," *Small Business Economics* 8 (4), August, 297-315.

DeYoung, Robert, 1998, "Comment on Goldberg and White," *Journal of Banking and Finance* 22, pp. 868-72.

DeYoung, Robert, Lawrence G. Goldberg, and Lawrence J. White, 1999, "Youth, Adolescence and Maturity at Banks: Credit Availability to Small Business in an Era of Banking Consolidation," *Journal of Banking and Finance* 23 (2-4), February, 463-492.

Eberts, Randall W., and Daniel P. McMillen, 1998, “Agglomeration Economies and Urban Public Infrastructure,” in *Handbook of Regional and Urban Economics, Volume 3: Applied Urban Economics*, Paul Cheshire and Edwin S. Mills, eds. (Amsterdam: North-Holland).

Edwards, Franklin R., 1996, *The New Finance: Regulation and Financial Stability* (Washington, DC: AEI Press).

Elliehausen, Gregory E., and John D. Wolken, 1990, “Banking Markets and the Use of Financial Services by Small and Medium-Sized Businesses,” Federal Reserve Board of Governors, Staff Study No. 160, September.

Frydl, Edward J., 1993, “Studies on Excess Capacity in the Financial Sector,” Federal Reserve Bank of New York Research Paper 9310-9323, June.

Goldberg, Lawrence G., and L.J. White, 1998, “De novo banks and lending to small businesses,” *Journal of Banking and Finance* 22, pp. 851-867.

Goldin, Claudia, 1986, “The Earnings Gap Between Male and Female Workers: An Historical Perspective,” National Bureau of Economic Research Working Paper No. 1888, April, Cambridge, MA.

Haynes, George W., Charles Ou, and Robert Berney, 1999, “Small business borrowing from large and small banks,” in Jackson L. Blanton, Alicia Williams, and Sherrie L.W. Rhine, eds., *Business Access to Capital and Credit: A Federal Reserve System Research Conference*. Federal Reserve Bank of Chicago, pp. 287-327.

Henderson, Jason R., and Kendall McDaniel, 2000, “The Impact of Scenic Amenities on Rural Employment Growth,” mimeo, Purdue University, August.

Jackson, William E., 1992, “Is the Market Well Defined in Bank Merger and Acquisition Analysis?” *The Review of Economics and Statistics* 74, November, 655-661.

Jaeger, David A., and Marianne E. Page, 1996, “Degrees Matter: New Evidence on Sheepskin Effects in the Returns to Education,” *Review of Economics and Statistics* 78 (4), November, 733-740.

Jayaratne, Jith, and Philip E. Strahan, 1996, “The Finance-Growth Nexus: Evidence from Bank Branch Deregulation,” *Quarterly Journal of Economics* 111 (3), August, 639-670.

Jayaratne, Jith, and John Wolken, 1999, “How Important are Small Banks to Small Business Lending? New Evidence from a Survey of Small Firms,” *Journal of Banking and Finance* 23 (2-4), February, 427-258.

Kane, Thomas J., 1994, “College Entry by Blacks Since 1970: The Role of College Costs, Family Background and the Returns to Education,” *Journal of Political Economy* 102 (5), October, 878-911.

King, Robert, and Ross Levine, 1993a, “Finance and Growth: Schumpeter Might Be Right,” *Quarterly Journal of Economics* 108 (3), August, 717-738.

King, Robert, and Ross Levine, 1993b, “Finance, Entrepreneurship, and Growth: Theory and Evidence,” *Journal of Monetary Economics* 32, pp. 513-542.

Krol, Robert, and Shirley Svorny, 1996, “The Effect of the Bank Regulatory Environment on State Economic Activity,” *Regional Science and Urban Economics* 26 (5), August, 531-541.

Kwast, Myron, Martha Starr-McCluer, and John D. Wolken, 1997, “Market Definition and the Analysis of Antitrust in Banking,” *Antitrust Bulletin* 42 (4), Winter, 973-995.

Lam, David, and Deborah Levison, 1991, “Declining Inequality in Schooling in Brazil and Its Effects on Inequality in Earnings,” *Journal of Development Economics* 37 (1-2), November, 199-225.

Levine, Ross, 1998, “The Legal Environment, Banks, and Long-Run Economic Growth,” *Journal of Money, Credit, and Banking* 30 (3), August, Part 2, 596-613.

O hUallachain, Breandan, and Mark A. Satterthwaite, 1992, “Sectoral Growth Patterns at the Metropolitan Level: An Evaluation of Economic Development Incentives,” *Journal of Urban Economics* 31 (1), January, 25-58.

Peek, Joe, and Eric S. Rosengren, 1996, “Small Business Credit Availability: How Important is Size of Lender?” in *Universal Banking: Financial System Design Reconsidered*, ed. by Anthony Saunders and Ingo Walter (Burr Ridge, Ill.: Irwin Publishing).

Peek, Joe, and Eric S. Rosengren, 1998, “Bank Consolidation and Small Business Lending: It’s Not Just Bank Size that Matters,” *Journal of Banking and Finance* 22 (6-8), August, 799-819.

Petersen, Mitchell, and Raghuram G. Rajan, 1995, “The Effects of Credit Market Competition on Lending Relationships,” *Quarterly Journal of Economics* 60, 407-444.

Psacharopoulos, George, 1989, “Time Trends of the Returns to Education: Cross-National Evidence,” *Economics of Education Review* 8 (3), 225-231.

Rajan, Raghuram G., and Luigi Zingales, 1998, “Financial Dependence and Growth,” *American Economic Review* 88 (3), June, 559-586.

Riordan, Michael H., 1993, “Competition and Bank Performance: A Theoretical Perspective,” in *Capital Markets and Financial Intermediation*, ed. by Colin Mayer and Xavier Vives, (Cambridge: Cambridge University Press), 328-343.

Rob, Rafael, 1995, “Entry and Exit of Firms and the Turnover of Jobs in U.S. Manufacturing,” Federal Reserve Bank of Philadelphia *Business Review*, March/April, 3-18.

Shaffer, Sherrill, 1993, “Can Megamergers Improve Bank Efficiency?” *Journal of Banking and Finance* 17 (2-3), April, 423-436.

Shaffer, Sherrill, 1996, “Viability of Traditional Banking Activities: Evidence from Shifts in Conduct and Excess Capacity,” *International Journal of the Economics of Business* 3 (2), 125-143.

Shaffer, Sherrill, 1998, “The Winner’s Curse in Banking,” *Journal of Financial Intermediation* 7 (4), October, 359-392.

Small Business Administration, 2000, “Small Business Frequently Asked Questions,” <http://www.sba.gov/advo/stats/sbfaq.html#q5>.

Strahan, Philip, and J. Weston, 1998, “Business Lending and the Changing Structure of the Banking Industry,” *Journal of Banking and Finance* 22 (6-8), August, 821-845.

Topel, Robert H., 1997, “Factor Proportions and Relative Wages: The Supply-Side Determinants of Wage Inequality,” Journal of Economic Perspectives 11 (2), Spring, 55-74.

Whalen, Gary, 1995, “Out-of-state Holding Company Affiliation and Small Business Lending,” Economic and Policy Analysis Working Paper 95-4, Comptroller of the Currency, Washington, DC.

Whitehead, David D., 1990, “Relevant Geographic Banking Markets: How Should They Be Defined?” Federal Reserve Bank of Atlanta Economic Review, January/February, 20-28.